Financial Market Effects of Macroeconomic Policies

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Zusammenfassung

Die Finanzkrise und darauffolgende Weltwirtschaftskrise, welche das vergangene Jahrzehnt maßgeblich prägten, zeigten Ökonomen wie auch politischen Entscheidungsträgern die Bedeutung der Finanzmärkte für das Verständnis makroökonomischer Schwankungen deutlich auf. Um bewusst und gezielt Effekte auf die Realwirtschaft mithilfe von maßgeschneiderten fiskal- und geldpolitischen Maßnahmen ausüben zu können, spielt folglich die Untersuchung von Transmissionseffekten über die Finanzmärkte eine bedeutende Rolle.

Vor diesem Hintergrund untersucht die vorliegende Dissertation die Effekte und Wechselwirkungen verschiedener makroökonomischer Politikmaßnahmen an den Finanzmärkten empirisch. Dabei setzt sich diese Arbeit insbesondere mit den Effekten unkonventioneller Geldpolitik in Form von Anleiheankaufprogrammen durch die Europäische Zentralbank auf die Staatsanleihenmärkte und mit den direkten Auswirkungen einer solchen Geldpolitik auf das öffentliche Schuldenmanagement in Deutschland auseinander. Des Weiteren liefert sie anhand von U.S. Wirtschaftsdaten neuartige Belege für den Kreditmarkt als relevanten Transmissionskanal staatlicher Ausgabenpolitik in der Anregung des privaten Konsums.

Nach einer kurzen Einleitung in Zielsetzung, Methodik und Aufbau der Arbeit im ersten Kapitel geht das zweite Kapitel dieser Dissertation aus einer modelltheoretischen und empirischen Perspektive auf die Finanzmarkteffekte der Staatsanleihekäufe der Europäischen Zentralbank ein. Hierzu wird das komparativ-statische preistheoretische Geldangebotsmodell nach Bofinger et al. (2017) genutzt, um die Effekte und Transmissionskanäle unkonventioneller Ankaufprogramme der Europäischen Zentralbank auf dem Anleihenmarkt in geeigneter Form darzustellen. Als Folge einer Ankündigung von Anleihekäufen durch

die Zentralbank sind im Modell (i) negative Zinseffekte am Anleihenmarkt durch ein (ii) rückläufiges Kreditrisiko und (iii) sinkende Laufzeitprämien zu verzeichnen. Diese Beobachtungen werden im Rahmen eines Fehler-Korrekturmodells und in Ereignisstudien empirisch für fünfjährige deutsche, französische, irische, italienische, portugiesische und spanische Staatsanleihen und für eine ebenfalls fünfjährige Eurozonen Benchmark Anleihe überprüft. Daraus bestätigen sich insbesondere Hypothesen (i) und (ii) für die Zielländer, d.h. Irland, Italien, Spanien, Portugal, im Rahmen der gezielten Käufe notleidender Staatsanleihen. Weiterhin zeigt sich infolge der Ankündigung massiver Ankäufe von Staatsanleihen aller Mitgliedsstaaten der Europäischen Währungsunion im Januar 2015 ein signifikanter Zinssenkungseffekt für die Anleihen aller Länder der Stichprobe. Diese Ergebnisse bestätigen zum einen die Vorhersagekraft des Modells im Anwendungsfall und zum anderen die Wirksamkeit dieser unkonventionellen Maßnahmen in der Erfüllung ihrer jeweiligen Zielsetzung. Während letztere bei den Anleiheankaufprogrammen vor 2014 insbesondere in der Wiederherstellung des Vertrauens an den Finanzmärkten und der Reduktion von internationalen Zins-Spreads bestand, zielte das massive Anleiheankaufprogramm, für das ein Ende der Nettoankäufe zum Zeitpunkt der Einreichung dieser Dissertation noch nicht absehbar ist, auf eine flächendeckende Senkung langfristiger Zinsen ab.

Das dritte Kapitel dieser Dissertation erforscht einen kausalen Zusammenhang zwischen den massiven Anleihekäufen der Europäischen Zentralbank und der aktiven Gestaltung der Fälligkeitsstruktur am Primärmarkt durch das öffentliche Schuldenmanagement am Beispiel Deutschlands. Mit einem aus Pressemitteilungen der Deutschen Finanzagentur GmbH gewonnenen Planungsdatensatz zur Neuemission von Staatsanleihen werden Diskrepanzen zwischen geplanter und tatsächlicher Fälligkeitsstruktur festverzinslicher Wertpapiere des Bundes ermittelt und im zeitlichen Zusammenhang mit Ankündigungen unkonventioneller geldpolitischer Maßnahmen auf den Staatsanleihenmärkten deskriptiv und mit ökonometrischen Methoden analysiert. Hierbei werden die getesteten geldpolitischen Ereignisse einerseits qualitativ nach ihrem Bezug zu Staatsanleihenmärkten selektiert und zu binären Event-Zeitreihen zusammengefasst sowie andererseits quantitativ mit dem Zinseffekt auf dem Anleihenmarkt in engem zeitlichem Zusammenhang gewichtet. Unter Anwendung der lokalen Projektions-Methode nach Jordà (2005) kann hierbei ein robuster, um einige Monate verzögerter, signifikanter Kausalzusammenhang zwischen expansiver unkonventioneller Geldpolitik und ungeplanter Verlängerung der Fälligkeitsstruktur identifiziert werden. Die Ergebnisse legen dabei nahe, dass das öffentliche Schuldenmanagement ein Kostensenkungsmotiv verfolgt, indem es entsprechend der Zinsanreize der Geldpolitik kurzfristig von der geplanten Fälligkeitsstruktur abweicht. Eine Berücksichtigung dieses Zusammenhangs ist vor dem Hintergrund der sich daraus potenziell ergebenden gegenläufigen Effekte auf die Anleihezinsen am Sekundärmarkt folglich für eine effektive und effiziente Gestaltung unkonventioneller Ankaufprogramme durch die Zentralbank von entscheidender Bedeutung.

Das vierte und letzte Kapitel dieser Dissertation setzt sich anhand von U.S. Wirtschaftsdaten mit den Wirkungskanälen staatlicher Ausgabenpolitik auf private Konsumausgaben auseinander und liefert Anhaltspunkte für die Relevanz privater Kreditmärkte in der Übertragung fiskalpolitischer Impulse. Unter Anwendung einer rekursiven Vektor-Autoregression (VAR) lässt sich infolge einer Staatsausgabenerhöhung ein Anstieg des privaten Konsums beobachten, welcher – im Einklang mit dem traditionellen keynesianischen Paradigma – mit einem anhaltenden Anstieg des gesamtwirtschaftlich verfügbaren Einkommens einhergeht. Durch Fixierung des Einkommens nach Steuern auf den Pfad der ohne einen staatlichen Eingriff zu erwarten wäre, wird anhand des weiterhin signifikanten positiven Anstieg des privaten Konsums erkennbar, dass ein gestiegenes Einkommen nicht der einzige Wirkungskanal für die Übermittlung öffentlicher Ausgabenimpulse hin zu den Haushalten sein kann. Infolge einer Erweiterung des VAR-Modells um Kreditvolumen und -preise und einer entsprechenden Anpassung der Identifikationsstrategie zur Berücksichtigung vorausschauender Finanzmärkte wird ein Anstieg der privaten Kreditaufnahme als weiterer Wirkungskanal der staatlichen Ausgabenpolitik identifiziert. Begleitet von sinkenden Zinsen, rückläufigen Kreditaufschlägen und steigenden Preisen für Kreditsicherheiten wie beispielsweise Immobilien, zeigt sich, dass eine Lockerung der Kreditmarktbedingungen die makroökonomischen Effekte staatlicher Ausgabenpolitik weiter verstärken.

Acknowledgment

The dissertation at hand is the product of my doctorate studies, which I pursued during my employment as a research and teaching assistant at the Chair for Monetary Policy and International Economics at the University of Würzburg and during a six month traineeship at the Directorate General Microprudential Supervision 1 at the European Central Bank. My time at the Chair for Monetary Policy and International Economics at the University of Würzburg and the European Central Bank constituted a stimulating environment to conduct my research and to learn more about the exciting field of financial markets, monetary policy, fiscal policies, and macroeconomics, in general.

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List of Abbreviations

1y	1-year
3m	3-month
3y	3-year
6m	6-month
a	saturation amount of credit demand
ABSPP	Asset-Backed Securities Purchase Programme
APP	Expanded Asset Purchase Programme
AR	autoregression
b	interest rate elasticity of credit demand
bp	basis point
B	amount of refinancing in the bond market
BoE	Bank of England
BoJ	Bank of Japan
C	cash
$\mathbf{c}_{\mathrm{B}}$	credit default costs of banks
$\mathbf{c}_{\mathrm{NB}}$	credit default costs of non-bank suppliers of funds
$\mathbf{Cr}_{\mathrm{B}}$	bank credit
$\mathbf{Cr}_{\mathrm{CB/B}}$	credit from central bank to banks
$\mathbf{Cr}_{\mathrm{B}}^{\mathrm{D}}$	credit demand in the bank credit market
$\mathbf{Cr}_{\mathrm{NB}}^{\mathrm{D}}$	credit demand in the bond market
$\mathbf{Cr}_{\mathrm{B}}^{\mathrm{S}}$	credit supply by banks
$\mathbf{Cr}_{\mathrm{NB}}^{\mathrm{S}}$	credit supply by non-banks

CBPP	Covered Bond Purchase Programme
CDS	Credit Default Swap
CSPP	Corporate Sector Purchase Programme
d	substitution elasticity between bank credit and bond market credit
D	deposits
DF	Deposit Facility
DFR	Deposit Facility Rate
DSGE	dynamic stochastic general equilibrium
E	refinancing in the equity market
EA-MPD	Euro Area Monetary Policy Event-Study Database
ECB	European Central Bank
ECM	error correction model
EFSF	European Financial Stabilisation Facility
e.g.	exempli gratia
EONIA	Euro Overnight Index Average
ESM	European Stability Mechanism
Fed	Federal Reserve Bank
FRFA	Fixed Rate Full Allotment Procedure
FTO	Fixed Term Operation
FX-Swap	Foreign Exchange Swap
GDP	gross domestic product
h	cash holding coefficient of the private sector
H	high-powered money
$\mathbf{H}^{\mathrm{D}}$	high-powered money demand
$\boldsymbol{\eta}^{\mathrm{B}}$	ratio of bond refinancing to bank credit
$\boldsymbol{\eta}^{\mathrm{E}}$	ratio of equity refinancing to bank credit
I	investment
$\mathbf{I}^{\mathrm{k}}$	information cost of a non-bank k
i.a.	inter alia
i.e.	id est
$\mathbf{i}_{\mathrm{B}}$	interest rate for bank credit
$\mathbf{i}_{\mathrm{D}}$	interest rate for deposits
$\mathbf{i}_{\mathrm{E}}$	interest rate for equity
$\mathbf{i}_{\mathrm{e}}^{\mathrm{t+1}}$	expected interest rate for $t+1$

$\mathbf{i}_{\mathrm{NB}}$	interest rate for bonds
$\mathbf{i}_{\mathrm{R}}$	interest rate for central bank credit
LIBOR	London Interbank Offered Rate
LTRO	Long-Term Refinancing Operation
m	number of non-bank suppliers of funds
M	maturity
MLF	Marginal Lending Facility
MR	Minimum Reserve
MRO	Main Refinancing Operation
$\mathbf{m}_{\mathrm{B}}$	bank credit multiplier
$\mathbf{M}^{\mathrm{D}}$	money demand
$\mathbf{M}^{\mathrm{S}}$	money supply
MPE	monetary policy event
n	number of banks
non-stat.	non-stationary
O	operational costs
OECD	Organisation for Economic Co-operation and Development
OMT	Outright Monetary Transactions
ov	outstanding volume
p	p-value
PCE	personal consumption expenditures
PDM	public debt management
PSPP	Public Sector Purchase Programme
$\mathbf{\Pi}_{\mathrm{B}}^{\mathrm{j}}$	profit of bank j
$\mathbf{\Pi}_{\mathrm{NB}}^{\mathrm{k}}$	profit of non-bank supplier of funds k
SMP	Securities Markets Programme
q	nominal value of a portfolio
QE	quantitative easing
QSB	quantified sovereign bonds proxy
r	reserve requirement
R	reserves
RBC	real business cycle
SBN	sovereign bond news proxy
SBP	sovereign bond purchase proxy

stat.	stationary
STRO	Special-Term Refinancing Operation
SVAR-IV	instrumental variable identified structural VAR
TED	Treasury Bill Eurodollar Difference
TLTRO	Targeted Longer-Term Refinancing Operation
tp	term premium
U.S.	United States of America
UK	United Kingdom
V	credit risk costs
VAR	vector autoregression
w.	with
wo.	without
WAM	weighted average maturity
WAMD	deviation of weighted average maturity from its planned value
Y	aggregate income

Contents

List of Figures

List of Tables

Chapter 1

Introduction

"The central idea is that government fiscal policy, its spending and taxing, its borrowing and repayment of loans, its issue of new money and its withdrawal of money shall all be undertaken with an eye only to the results of these actions on the economy and not to any established traditional doctrine about what is sound or unsound."

(see Lerner, 1943)

Formulated almost 80 years ago, the prepended quote by Abba P. Lerner remains of great importance for a scientific approach to macroeconomic policy, encompassing various fiscal and monetary policy measures, today. In line with their goal of improving the overall performance of the economy, it is Lerner's conviction that macroeconomic policy decisions should be taken along *functional* criteria, i.e., according to their *empirical* repercussions on macroeconomic aggregates, instead of following the dictate of economic schools of thought that are currently en vogue. As a step toward the required pragmatic approach to policy, the financial crisis and the subsequent Great Recession have taught academics and policy makers that financial markets are of crucial importance for the overall understanding of economic fluctuations. This further implies that if policy makers want to achieve the desired effects on the real economy through tailored fiscal and monetary policy measures, they have to internalize transmission effects via financial markets.

While the relevance of money and financial markets for a correct understanding of the workings of an economy was already recognized in theory by great economists, including Keynes (1936), Schumpeter (1954), and Tobin (1963), this theoretical concept appears to have gone out of fashion with the rise of market liberalization and globalization after the 1960s. Theoretical underpinnings on the *irrelevance* of finance for the real economy are, i.a., motivated by Barro (1974) with regard to government taxation and deficit financing, and by Modigliani and Miller (1958) regarding the debt and equity mix of firms. As a result, standard quantitative models that have dominated macroeconomic research prior to the financial crisis, including real business cycle (RBC) models, initially introduced by Kydland and Prescott (1982), and New Keynesian dynamic stochastic general equilibrium (DSGE) models, as for instance Smets and Wouters (2007), either fully abstract from a financial sector or assign only a minuscule role to it. While the traditional monetary paradigm has been revitalized and gained momentum in economic models since the Great Recession (see, e.g., Disyatat, 2011; Werner, 2014; Bofinger et al., 2017), further empirical research is necessary to inform us about the effects of both, traditional as well as innovative macroeconomic policy tools in practice.

Against this backdrop and following the appeal by Lerner (1943), within three self-contained studies, this dissertation aims to further improve our understanding on the repercussions of fiscal and unconventional monetary policies (i) by *empirically* and *quantitatively* evaluating the impact and interactions of different macroeconomic policy measures, while (ii) taking the role of *financial markets* seriously. In particular, this dissertation sheds light on the financial market effects of unconventional central bank asset purchase programs in the Eurozone and their potential interactions with national public debt management policies. In addition, it provides novel evidence on the role of private credit markets in the propagation of public spending toward private consumption in the U.S. economy.

From a methodological perspective, this dissertation draws on a set of state-of-the-art time-series-econometric methods to study the transmission mechanisms of macroeconomic policies. To analyze contemporaneous interactions in financial and macroeconomic data in the context of unconventional monetary policy, it applies an event study framework along the lines of Altavilla et al. (2015) and De Santis (2020). To trace the dynamic consequences of macroeconomic policies over time, the dissertation makes use of (i) an error correction model framework, which was initially applied by Sargan (1964), (ii) local projections, which were developed by Jordà (2005), and (iii) vector autoregressions (VARs), which were introduced by Sims (1980) into macroeconometric research.[1]

An important challenge in time-series analysis of macroeconomic policies lies in the identification of exogenous movements in the respective policy instrument (see Ramey, 2016). In this vein, a successful isolation of the marginal effects of macroeconomic policies necessitates to take into account policy announcements and foresight issues, which are particularly pronounced in the context of financial markets. Therefore, this dissertation carefully deals with the challenge of identification by applying narrative information from outside of the time-series models in each included study.[2] More specifically, for the analysis of unconventional monetary policy actions, this thesis relies on press releases and central bank announcements in the construction of monetary policy proxies, explores planning data and announcements of revisions to these plans published by the public debt management office to identify sudden shifts in the maturity structure of public debt issuance, and controls for news about selected components of government spending to account for fiscal foresight considerations, when studying the pass-through of government spending programs. As a result, changes to the analyzed policy variables can be understood as plausibly exogenous, enabling a clean interpretation of the empirical results as being directly attributable, i.e., a causal response, to the policy change.

[1] In particular, the VAR methodology applied in this dissertation encompasses structural VARs (SVARs) identified via the recursive approach (Blanchard and Perotti, 2002) as well as via the use of external instruments in the so-called proxy-SVAR framework (Stock and Watson, 2012; Mertens and Ravn, 2013; Montiel Olea et al., forthcoming).

[2] This narrative approach is widespread in macroeconomic research and extends to different applications, including Romer and Romer (2004) on conventional monetary policy, Fieldhouse et al. (2018) on government asset purchases, Ramey (2011) on public spending, and Mertens and Ravn (2013) on tax policy.

In the subsequent Chapter 2, financial market effects of unconventional monetary policy measures of the European Central Bank (ECB), in particular sovereign bond asset purchase programs, are dissected both theoretically and empirically. For this purpose, a tractable financial market model, consisting of a bank credit and bond market is applied to formalize the effects of central bank asset purchase programs. In a second step, the model predictions on bond yields and underlying transmission channels are empirically validated for multiple unconventional ECB asset purchase programs.

As a theoretical framework, this Chapter applies the price-theoretic model of the money supply, which was initially developed by Bofinger and Schächter (1995) and recently extended to feature a bond market by Bofinger et al. (2017). This model is particularly suited to explain interactions between bank credit and bond markets and the impact of central bank actions. Furthermore, it allows to derive closed-form analytical solutions and a graphical illustration of results due to its comparative-static nature. Based on the assumption that forward looking market participants tend to price in central bank actions upon mere announcement, three hypotheses on the effect and transmission of central bank asset purchase programs on sovereign bond yields are derived: (i) asset purchase programs are effective in lowering bond yields, while this effect is achieved through (ii) decreasing (perceived) credit risk, and (iii) declining term-premia in the bond market. These hypotheses are subsequently tested empirically in an error-correction and event-study framework for 5-year sovereign bonds issued by Germany, France, Ireland, Italy, Spain, Portugal and a 5-year Euro area benchmark bond. Contrasting empirical studies that focus either on the effect on inter-country yield spreads for the ECB's asset purchase programs that were targeted to distressed sovereign markets (Szczerbowicz, 2015; Falagiarda and Reitz, 2015) or on the effects of the ECB's large-scale asset purchase program launched in 2015 on yield levels (Altavilla et al., 2015; De Santis, 2020), the empirical approach taken in this Chapter consists of studying the effects on sovereign bond *yield levels* for the *entire universe* of asset purchase programs by the ECB.

Overall, the results from the empirical analysis largely confirm the hypotheses derived from the model and support the ECB's success in fulfilling the respective objectives of different types of asset purchase programs. In accordance with hypothesis (i), the error correction model, on the one hand, reveals a negative yield effect upon the announcement of a targeted and large-scale sovereign-bond asset purchase program for assets within the purchased spectrum. In addition, decreases in sovereign bond yields following announcements of non-sovereign bond asset purchase programs hint to the presence of a portfolio rebalancing channel across different asset classes. The performed event based regressions, on the other hand, reveal a significant reduction of country-specific credit default spreads and in parts also for term-premia, in particular for announcements of targeted sovereign bond market programs, which indicates that the ECB may have been successful by acting as a "lender of confidence" and by inducing a convergence of the previously diverging sovereign bond yields in the Euro area.

Over the course of the ECB's large-scale asset purchase program, which was launched in 2015 and is still active at the time of submission of this dissertation, the ECB accumulated more than €2.5 billion of sovereign bonds on its balance sheet. Notably, for the case of Germany, the ECB's holding represent more than 30% of the outstanding volume of marketable sovereign debt, which again constitutes as much as 35% of the total national marketable debt volume (Bundesbank, 2018). As a result, this policy measure may not only have triggered behavioral adjustments by private investors that hold positions in these markets, but also by policy makers that decide upon the level and structure of debt instruments to be issued.

Against this backdrop, Chapter 3 of this dissertation studies and quantifies the direct implications of unconventional monetary policy on decisions by German public debt management regarding the maturity structure of gross issuance. Since central bank asset purchase programs are effective in lowering long-term yields, i.e, the cost of borrowing long term, public debt managers may extend issuance volumes of long-term bonds and hence the maturity structure of gross issuance to exploit a cost saving motive. Related studies on this topic either focus on the interactions with conventional monetary policy, such as Blommestein and Turner (2012) and Hoogduin et al. (2011), or on potential counteractions by public debt management (Greenwood et al., 2016a). Yet, none of these studies allows for a clear *causal* interpretation of results.

By compiling a unique data set based on press releases of the German debt management office, this study traces deviations in actual from weighted average maturity of gross issuance in the temporal context of announcements by the ECB involving sovereign-bond markets. Following the qualitative selection based on their market reverence, these announcements are proxied (i) by binary event-series (as in, e.g., Dedola and Georgiadis, 2018; Enders et al., 2019; Urbschat and Watzka, forthcoming) and (ii) by quantified event-series, in which the binary event dummies are weighted by high-frequency movements in financial market variables measured within a narrow time window around the announcement.[3] A joint descriptive analysis of both macroeconomic policy time series is suggestive of a link between deviations from maturity plans and unconventional monetary policy announcements that involve central bank sovereign bond market actions. Yet, to uncover a potential causal link and to capture dynamic interrelations between monetary policy announcements and shifts in the maturity structure of public debt, a regression-setup is required.

Local projections in the style of Jordà (2005) reveal a statistically significant maturity extension in gross issuance by public debt management. In particular, it is shown that German public debt managers extend the maturity of gross issuance in response to unconventional monetary policy announcements of central bank sovereign bond market actions in an economically meaningful and dynamic manner. These results are suggestive that public debt management in fact exploits a cost saving motive by reacting to the expected yield depressing effects that central bank asset purchase announcements trigger as shown in Chapter 2. Taken together, these findings support the view of a meaningful intertwining of monetary policy and public debt management in the Euro area. If unaccounted for by policy makers, this behavior could impair the transmission of monetary policy and necessitate readjustments of monetary stimulus, since the maturity extension and consequent re-raising of the relative supply of long-term public debt instruments may counteract the intended yield depressing effects of large-scale asset purchase programs (Greenwood et al., 2016a; Swanson, 2011).

As Chapter 3 establishes, a clear cut distinction between the spheres of fiscal and monetary policy becomes increasingly difficult when taking a financial markets view to economic research. In this vein, studies on the financial market effects of fiscal policy and public spending, which argue that changes in the relative supply of long-term sovereign bonds might lead to lower public bond spreads in financial markets (see, for instance, Krishnamurthy and Vissing-Jorgensen, 2012; Greenwood and Vayanos, 2010) represent the fiscal policy counterpart to the plethora of empirical findings on sovereign yield depressing effects of sovereign bond asset purchase program. Yet, while the debate on how fiscal policy affects financial market conditions on sovereign bond markets is receiving increasing attention in economic research, far less is known about the consequences of fiscal policy for *private* credit markets.

[3] For this purpose, the data is taken from the novel Euro Area Monetary Policy Event-Study Database developed by Altavilla et al. (2019).

In an attempt to make progress on this front, Chapter 4 of this dissertation thus studies the macroeconomic consequences of shifts in fiscal policy and identifies substantial interactions with credit markets in the overall transmission of public stimulus. In particular, the Chapter tests the role of household indebtedness and household income in propagating increases in fiscal spending to private consumption expenditures and, thereby, to the broader economy. With private consumption representing the largest component of aggregate demand, its reaction to public stimulus is vital for the overall effectiveness of fiscal policy in stabilizing the economy. From a theoretical perspective, however, the effects of public on private spending are rather ambiguous. While the traditional Keynesian paradigm predicts an *income-induced rise* of private consumption, prototype New Keynesian and RBC models suggest that public spending crowds out private spending by inducing households to substitute from consumption to labor supply. Moreover, to better understand the causal link of public and private spending, it is also conceivable that household debt may be a part of the transmission mechanism to private consumption.

As an empirical laboratory, the Chapter employs aggregate time series for postwar U.S. data and recovers exogenous government spending surprises via a recursive VAR, e.g., by assuming that government spending—due to implementation lags—is predetermined with respect to aggregate activity (Galí et al., 2007). To take into account potential contamination by anticipation effects, the identification scheme additionally controls for simultaneous fluctuations in news on military spending (Ramey, 2011). The fiscal policy VAR is further augmented by credit market volumes and prices, and the identification strategy is consequently modified to account for forward-looking financial variables in this case.[4] In particular, the adjusted SVAR uses an external instrument, military spending, to recover plausibly exogenous shifts in government spending in the presence of credit market spreads and interest rates (see Miyamoto et al., 2019, for a related approach using local projections).

The benchmark SVAR reveals consumption crowding-in for surges in public spending and suggests this consumption response to be accompanied by a persistent increase in disposable income. Endogenously reacting income, however, is insufficient to entirely rationalize conditional comovement of private and public spending: once after-tax income is hypothetically forced to its pre-shock path, consumption still rises. Moreover, the results

[4] See Gertler and Karadi (2015), who caution against the recursive approach in VARs for models that include both financial and macroeconomic time series.

of the credit market augmented SVAR-IV model provide corroborating causal evidence of fiscal stimulus prompting households to take on more credit. This favorable debt cycle is paralleled by dropping interest rates, narrowing credit spreads, and inflating collateral prices, e.g., real estate prices. As a result, this Chapter of the thesis supports the notion that to understand the macroeconomic effects of fiscal policy actions, it is indispensable to take into account potential credit market interactions. For public spending, the Chapter shows that a loosening of credit market conditions reinforces the macroeconomic effects of fiscal stimulus, on average.

As a whole, the results of this dissertation give new insight on the transmission of macroeconomic policies through financial markets. As a result of a more realistic and complex setting within the monetary paradigm, the findings may foster a better informed application and enhanced calibration of unconventional monetary, public debt management, and government spending policies. Since central bank asset purchase programs are on the verge of becoming the new normal in the Eurozone and discretionary government spending across the globe is at unprecedented levels in response to the current pandemic, a thorough understanding of macroeconomic policy effects is imperative to make informed decisions.

Nevertheless, this dissertation represents merely a small, yet important step toward a truly functional approach to macroeconomic policies (see Lerner, 1943). Subject to the availability of comparable data, the analyses of this dissertation may be extended as follows. First, in addition to the empirically tested effects of asset purchase programs on bond markets, the theoretical model from Chapter 2 suggests that these unconventional measures transmit to bank credit markets, too; further empirical studies may thus test the model predictions on this end. Second, the approach in Chapter 3 could be extended to analyze the relationship of public debt management and central bank sovereign bond market actions for other (Eurozone) countries; further research may also shed light on the extent to which the identified public debt management reactions to unconventional monetary policy counteract the stimulus brought about by monetary policy. Third, potential fiscal policy and credit market linkages similarly could be explored for countries other than the U.S. economy in follow up research; moreover, apart from the identified credit market consequences of fiscal spending for households, complementary future research may center on the repercussions on financial conditions for firms.

Chapter 2

A Theoretical and Empirical Assessment of Asset Purchase Programs in the Eurozone[1]

[1]This Chapter is based on joint work with Mathias Ries. Due to the finalisation of this project and subsequent publication in Ries (2018), this paper reflects the state of research up to May 2017. An earlier version of this paper appeared as Ries and Simon (2017).

2.1 Introduction

"The problem with QE is it works in practice, but it doesn't work in theory", Bernanke replied when asked about the theoretic foundation of quantitative easing (QE) in 2014 (see Bernanke, 2014). In fact, a plethora of empirical studies, mainly dealing with unconventional monetary policy programs by the Federal Reserve (Fed), the Bank of England (BoE), and the Bank of Japan (BoJ), have shown that large-scale QE programs do indeed have the desired effects. In particular, a large number of studies identify and disentangle various transmission channels of QE and find support for a decline of sovereign bond *yields* following from QE.[2] For the Euro area, where the majority of the European Central Bank's (ECB) unconventional measures during early stages were aimed at reducing inter-country financial market distortions in response to the financial crisis, the sovereign-banking nexus, and the sovereign debt crisis, a substantial focus of literature centers on the impact of these measures on bond *spreads*. For instance, a significant impact of asset purchase programs, in terms of decreasing sovereign bond spreads relative to the German Bund, has been found for the Eurozone by Falagiarda and Reitz (2015), Szczerbowicz (2015), and Eser and Schwaab (2016) in the context of programs launched prior to 2014.[3] However, from 2015 onward, the ECB followed the role-model of other central banks in conducting large-scale asset purchase programs, and hence meeting the narrower definition of QE.[4] Consequently, empirical studies concluding that the ECB's QE is effectively lowering bond yields in the Eurozone emerged, as well.[5]

This paper aims to contribute to the rapidly-growing literature on unconventional monetary policy in the Euro area along the following dimensions: first, we formulate a tractable theoretical model assigning quantities and prices in the markets for bonds as well as bank credit a meaningful role. Second, we show that the ECB's various asset purchase programs, in fact, are capable of lowering long-term bond yields and further transmit to interest rates on the banking market in our comparative-static framework. Third, we

[2]These studies include, i.a., Gagnon et al. (2011), Joyce et al. (2011) and Bauer and Rudebusch (2014). For a comprehensive overview of transmission channels see Krishnamurthy and Vissing-Jorgensen.

[3]Note, however, that since sovereign bond spreads are typically computed relative to German Bund yields for Eurozone countries, these studies are unable to make statements about the effects of asset purchase programs on the *German* sovereign bond market.

[4]Since the ECB's Expanded Asset Purchase Programme (APP) is characterized by a large regular purchase volume and consequently entails a massive extension of the central bank balance sheet, it can be referred to as QE.

[5]In this vein, e.g., Altavilla et al. (2015) and De Santis (2020) identify an overall negative reaction of sovereign bond yields resulting from the announcement of the ECB's APP in January 2015 and further dissect the transmission channels of QE in the Euro area.

evaluate the hypotheses we derive from the model application in empirically testing the effect of the entire range of ECB asset purchase programs, including QE, in country-specific time-series regressions. Overall, we find empirical support for the effectiveness of these measures, but at the same time, reveal substantial cross-country heterogeneity as well as differences in the relevance of transmission across countries and programs.

We begin our paper by providing a historical track record of the monetary policy tools that the ECB introduced since the outbreak of the financial crisis. This exposition is meant to illustrate the variety of unconventional instruments of the ECB and the shift of its monetary policy focus toward the bond market. We proceed by carefully discussing existing empirical literature on the effects of unconventional monetary policy. Based on this literature review, we conclude that existing evidence on asset purchase programs paints a consistently favorable picture regarding the effectiveness of QE on financial markets, i.e., QE indeed appears to work in practice.

Based on this empirical motivation, we contribute to the existing literature by proposing a tractable model suited for the analysis of asset purchase programs. For our analysis we build on the price-theoretic model of the money supply developed by Bofinger and Schächter (1995) and recently extended by Bofinger et al. (2017) and apply it to the analysis of the ECB's asset purchase programs, including QE. This comparative-static model is particularly suitable to reconcile the behavior of the central bank, banks, and non-banks in the process of credit supply, as it provides closed-form analytical solutions and a graphical representation of results. The model set-up consist of two interacting financial markets. First, the market for bank credit, where banks supply credit and in this way create money, is modelled similarly to the banking market in the model by Disyatat (2011). Since the supply of bank credit is based on banks' profit maximization in our model, refinancing conditions of banks influence the equilibrium on the bank credit market. Second, a bond market, in which non-banks redistribute the money created by the banking sector by purchasing bonds and in doing so implicitly grant loans to banks and non-banks, complements the theoretical framework. In evaluating the effects of asset purchase programs, we focus on the bond market as the effective arena of asset purchases, as qualify non-bank suppliers of credit as a counterparty for the ECB's large-scale asset purchases. This leads us to the main upshot of our theoretical model: by acting as an additional supplier of money in the bond market, the central bank is able to lower bond yields. This effect can be observed upon the mere announcement of asset purchase programs and transmits through decreasing credit risk and long-term interest rate expectations in our model, since agents on the bond market tend to price in actions of monetary policy as soon as they can be anticipated. Taken together, our theoretical model is supportive of the notion that (i) QE is effective in lowering bond yields, and operates through (ii) lowering (perceived) credit risk and (iii) impacting on term-premia in the bond market.

In the second part of our paper we empirically test these hypotheses by estimating the effect of the ECB's asset purchase programs on 5-year sovereign bond yields, credit risk spreads, and term premia for Germany, France, Portugal, Spain, Italy, Ireland, and the European benchmark bond. That is, we do not only quantify the overall effect of asset purchase programs in the Euro area, but also explore cross-country heterogeneity in individual-country time-series regressions.[6] For this purpose, we apply an error correction model in order to further distinguish between long and short-run effects on the bond market equilibrium of our model. In addition, we test our hypotheses on the transmission channels of asset purchase programs of reducing credit risk and long-term interest rate expectations, i.e. the term premium, by deploying event-based regressions.

Our empirical results lend overall support to the hypotheses formalized in the theoretical model. Remarkably, we find a negative yield effect on sovereign bond yields for most countries. However, we also document a yield-increasing effect on German and French bond yields, which could be rationalized by the fact that early ECB measures focused on European periphery countries as the epicenter of the Euro area crisis. In addition, we identify a clear-cut effect of unconventional monetary policy measures on lowering credit risk in the bond market, leading us to conclude that the ECB was successful in rebuilding trust between financial actors and can therefore be viewed as a "lender of confidence". Lastly, our findings regarding the effect of asset purchase programs on interest rate expectations, which we measure by term premia, reveal a rather diverse picture and may speak in favor of a heterogeneous portfolio rebalancing effect across countries.

The paper is organized as follows: After giving an overview on the existing monetary policy tools of the ECB in Section 2.2, we analyze the empirical literature on QE effects in Section 2.3. In Section 2.4, we propose our model and derive testable predictions from

[6]In doing so, on the one hand our study extends the studies on Eurozone QE by Altavilla et al. (2015) and De Santis (2020) by including the ECB asset purchase programs that were targeted to distressed markets and performed prior to 2015 into the analysis. On the other hand, in contrast to studies on early asset purchase programs by the ECB not fulfilling the criteria of QE, including Szczerbowicz (2015) and Falagiarda and Reitz (2015), our approach explicitly allows for spill-over effects to sovereign bonds outside the scope of these purchases, i.e., French and German sovereign bonds.

it, before estimating the effects of various asset purchase programs in the Euro area upon their announcement and dissecting transmission channels in Section 2.5. Ultimately, we conclude in Section 2.6.

2.2 Historical overview of monetary policy instruments in the Eurozone

Similar to other major central banks, such as the Fed, BoE and the BoJ, the ECB draws on a set of monetary policy tools to influence the economy. Under "normal" conditions, central banks provide liquidity to the banking system by using highly standardized instruments in their interaction with banks. Since the financial crisis, however, unconventional measures have been added to their toolboxes. In the following, we characterize the ECB's monetary policy instruments by categorizing them along the dimensions of (i) conventionality and (ii) targeted market, before discussing the individual measures taken between January 2008 and September 2017 in more detail:[7]

1. Conventional instruments
    a) Banking market
        i. Main Refinancing Operations
        ii. Fine Tuning Reverse Operations
        iii. Structural Reverse Operations
        iv. Longer-Term Refinancing Operations
2. Unconventional instruments
    a) Banking market: liquidity support measures
        i. Longer-Term Refinancing Operations with a term > 3 months (LTRO)
        ii. Targeted Longer-Term Refinancing Operations (TLTRO)

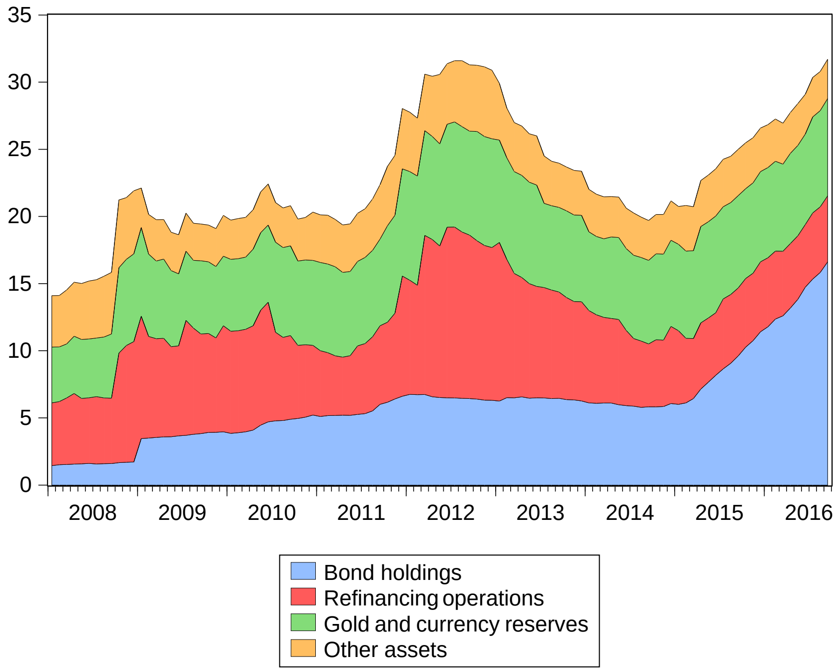
[7] The conventional instruments listed under item 1a form the operational framework of the Eurosystem, with the ECB's interaction with the banking sector mainly consisting of setting the price for banks' short and longer-term refinancing at the central bank. Depending on the market in which the central bank becomes active, the ECB's unconventional measures can be divided into liquidity support measures (banking market) and asset purchase programs (bond market).

b) Bond market: asset purchase programs

   i. Covered Bond Purchase Programme (CBPP)
   ii. Securities Markets Programme (SMP)
   iii. Outright Monetary Transactions (OMT)
   iv. Public Sector Purchase Programme (PSPP)
   v. Corporate Sector Purchase Programme (CSPP)

Although the ECB initially attempted to counteract the loss of confidence among banks and the resulting dry-up of the interbank market through conventional instruments, i.e., by lowering the refinancing rate, it soon engaged in unconventional measures that extended its balance sheet significantly over time. The size and composition of the ECB's assets in relation to Eurozone GDP is depicted in Figure 2.1.

Figure 2.1: Composition of the ECB's assets in percent of GDP



Notes: Cumulated balance sheet positions from the ECB's weekly financial statements are plotted for a data sample covering 2008M1 to 2016M9. The horizontal axis measures time in months, while the vertical axis measures the volume relative to the quarterly Eurozone GDP. The colored areas represent the sums of individual positions belonging to each category specified in the legend.

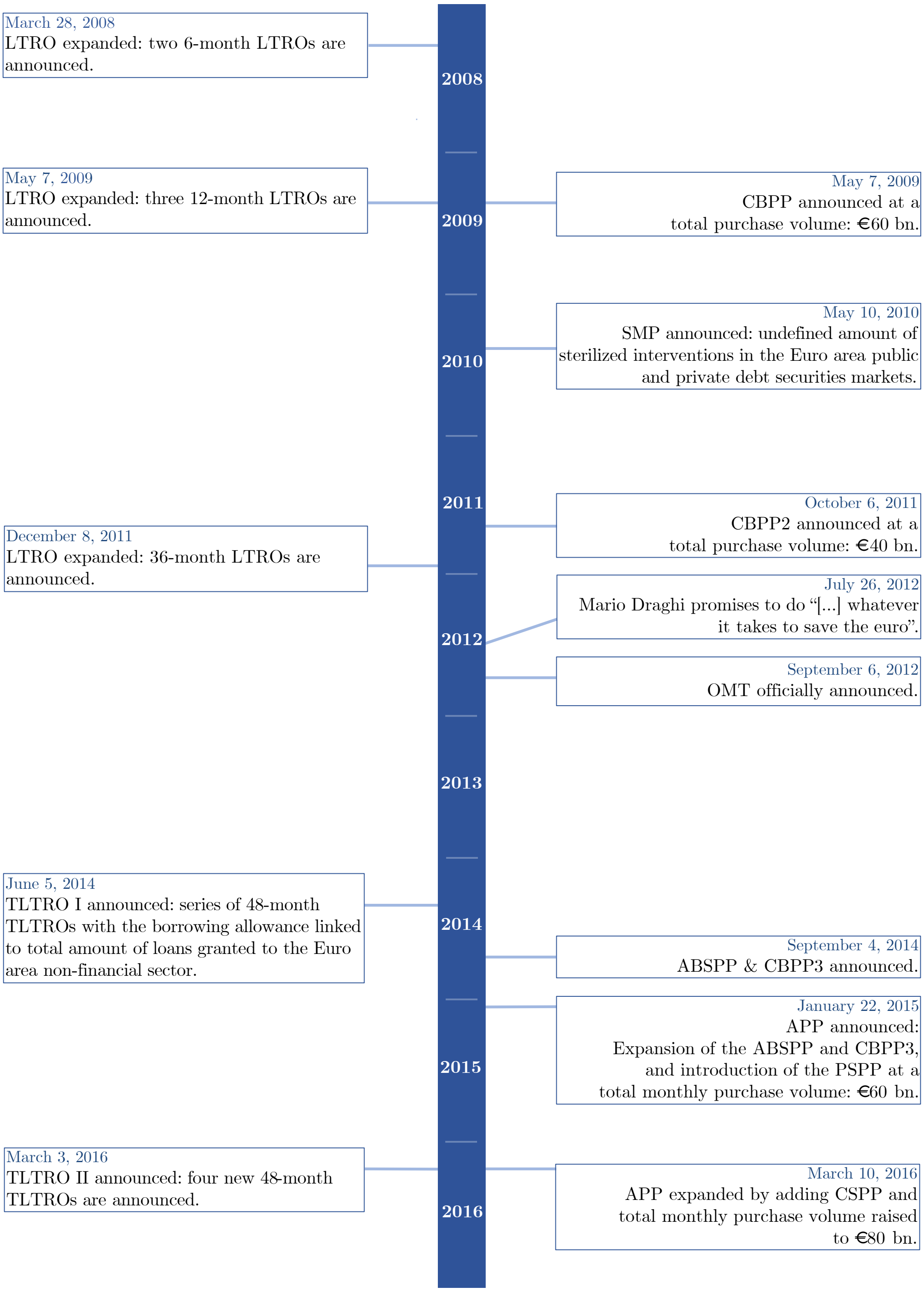
Within our sample from January 2008 to September 2016, sharp rises in the relative volume of refinancing operations in 2008 and 2012 (red area) show that the balance sheet expansion in the early years after the financial crisis and the sovereign debt crisis are mainly accounted for by liquidity support measures on the banking market. At the same time, the ECB's early asset purchase programs, which targeted distressed financial market segments, only gradually increased the amount of central bank bond holdings (blue area), before the introduction of regular, unsterilized, and large-scale asset purchases in 2015 caused a gradual and massive expansion of central bank assets, which marks the beginning of QE in the Eurozone.

In what follows, we present a detailed chronology of the ECB's announcements on the respective unconventional monetary policy programs driving these balance sheet effects within our sample (see Figure 2.2 for a graphical illustration).

Shortly after the collapse of Lehman Brothers and the spillover of the financial crisis of 2008–2009 to Europe, the ECB initially attempted to counteract the loss of confidence among banks and the resulting dry-up of the interbank market by extending the term of conventional Longer-Term Refinancing Operations. In particular, the ECB granted full allotment and extended the maturity of LTROs gradually from three months up to three years until the end of 2011, in order to close the funding gap in the banking sector, which had arisen as a result of the dysfunctioning interbank market. After the prolongation of two LTROs to a term of six months that had been announced on March 28, 2008 and the consequent increase of refinancing operations on the asset side of the central bank balance sheet, Jean-Claude Trichet offered three 12-month LTROs to provide even longer-term liquidity to banks, on May 7, 2009.

On the same day, he further announced the ECB's first asset purchase program, the CBPP. Collateralized by a pool of loans that continue to be obligations of the issuer, covered bonds represent an important funding instrument of banks in the medium and long term. Accordingly, the ECB's motivation in purchasing covered bonds was (i) to ease the funding conditions of banks, and (ii) to exert positive effects on funding conditions of non-financial corporations and households. Beyond the ongoing problems in the interbank market, the emergence of the sovereign debt crisis in Greece in 2010 induced an increase in default risk and fire sales of Eurozone government bonds. With the objective of preventing this development from getting out of hand and in order to "ensure the sustainability of [their] public finances" (see ECB, 2010), the ECB announced to purchase bonds in distressed sovereign bond market segments under the SMP on May 10, 2010. After the Greek debt crisis had somewhat stabilized in the beginning of 2011, concerns arised about spillovers to Italy and Spain. This led Mario Draghi to affirm the ECB's subsequent willingness

Figure 2.2: Timeline of the ECB's unconventional monetary policy measures



Notes: The timeline shows the date and content of unconventional monetary policy announcements made by the ECB between 2008M1 and 2016M9. Events on the right hand side of the timeline depict announcements of asset purchase programs, whereas the events on the left hand side relate to announcements of liquidity support measures.

to continue the SMP in August 2011. Over the course of the SMP, the ECB conducted *sterilized* purchases on public and private debt securities markets and purchased €219.5 billion in Irish, Greek, Spanish, Italian, and Portuguese sovereign bonds at an average maturity of approximately 5 years. Furthermore, the ECB reintroduced the CBPP on October 6, 2011, at a volume of €40 billion in response to the persistently stressed banking sector and the negative feedback loop of government bond yields on banks in the European periphery countries. At the end of October 2012, however, the ECB ended the CBPP2 unexploited, after purchasing a volume of only €16.4 billion. Despite recurring criticism that the ECB was overstepping its mandate, these financial market stability focused asset purchases did only drive the balance sheet expansion to a minor extent, whereas the major expansion of the central bank balance sheet during this time can be attributed to the 3-year LTROs. To counteract the banks' ongoing fire sales of government bonds and to further stabilize the lending of the banking sector, on December 8, 2011, the ECB announced to extend LTROs to an exceptionally long period of 36 months in order to enable cheap long-term funding to combat the continual deleveraging of the banking sector.

Since concerns about the stability of the Eurozone increased due to the sovereign-banking nexus and the continuous accumulation of sovereign debt, Mario Draghi promised to do "whatever it takes to save the euro" on July 26, 2012. This vague statement was interpreted by the markets as an unofficial announcement of another asset purchase program. The statement was further substantiated when the Governing Council revealed the replacement of the SMP by the OMT on September 6, 2012, in order to smooth the monetary transmission and to harmonize credit conditions in the Eurozone. In contrast to the SMP, the OMT required governments to comply with the adjustment programs of the European Financial Stability Facility (EFSF) or the European Stability Mechanism (ESM), as a precondition to qualify for central bank purchases of sovereign bonds with a shorter maturity of between 1 and 3 years. However, no actual purchases were conducted under the OMT.

A period of regeneration, underlined by the gradual repayment of the 3-year LTROs and a shrinking central bank balance sheet, followed in 2013 and early 2014, before stress tests of the European Banking Authority again put pressure on European banks. In order to support the banking sector, while encouraging its provision of credit to the private sector, in June 2014, the ECB extended the LTROs once more to a maturity of 48 months and set the borrowing allowance for banks contingent upon the total amount of loans granted to the Euro area non-financial sector (TLTROs). This recurrent easing of funding for banks was followed by the introduction of additional asset purchase programs. On September 4, 2014, the ECB announced another CBPP and introduced purchases of asset-backed securities (ABSPP) to relieve stress in the banking sector. As the underlying assets consist of claims against the non-financial private sector, the ABSPP was aimed at facilitating new credit flows to the non-financial sector. Both the ABSPP and the CBPP3 were introduced without a predefined end date and are still ongoing with current holding volumes of €31 and €278 billion, respectively, as of May, 2020.

When the weak economic situation in the Eurozone was exacerbated further by low inflation rates and restrained inflation expectations, the ECB announced the addition of the PSPP to its current purchase programs in January 2015. Amounting to €60 billion, the monthly purchases of combined assets under the CBPP3, ABSPP, and PSPP were designed to counteract deflationary pressure and second-round deflationary effects on wages and prices. In contrast to earlier asset purchase programs, the employment of this so called Expanded Asset Purchase Programme (APP), which was later augmented by the CSPP, led to unsterilized monthly purchases that caused a gradual expansion of central bank assets after 2015. Due to this prolongation of the ECB's balance sheet, the APP is often referred to as the actual beginning of QE in the Eurozone. Soon after the first purchases were made under the PSPP, the ECB expanded the total monthly purchase volume and added investment-grade bonds of non-financial corporations to its portfolio on March, 10 2016 and at the same time announced further conditional long-term liquidity provision for European banks (TLRTO II). Being the first ECB program to directly purchase corporate bonds, the aim of the CSPP is to bypass the banking sector and to strengthen the credit conditions for business financing more directly.

2.3 Literature overview

Due to a previous lack of data on and experience with unconventional monetary policy in practice, the majority of empirical literature in this field has mainly evolved over the course of the last decade, but has been growing dynamically since. In the following we aim to give a structured overview of existing empirical studies on unconventional monetary policy and to summarize their main results. For this purpose, we organize the wealth of scientific research along (i) the type of the program, i.e., either asset purchase programs or liquidity support measures, and (ii) the implementing central bank covered and further touch upon the methodologies applied.

2.3.1 Literature on liquidity support measures

The literature on liquidity support measures almost exclusively analyzes financial market effects, due to the aim of such measures to restore the function of monetary transmission rather than to affect inflation and growth directly (Rieth and Gehrt, 2016). A comprehensive overview on the estimation methods and results of studies on the effectiveness of liquidity support measures can be found in a survey of the literature by Borio and Zabai (2018).

With regard to the central banks considered, the literature is limited to examination of programs by the ECB and the Fed.[8] In particular, the Fed's eased refinancing conditions of the Term Auction Facility and the Term Securities Lending Facility, whereas the ECB extended the term of existing long-term refinancing operations for the banking sector via LTROs and TLTROs. With regard to the Fed, both event studies on Term Auction Facility focusing on LIBOR spreads (see, e.g., Wu, 2008; Thornton, 2011; McAndrews et al., 2017), and time-series regressions on the Term Securities Lending Facility (see, e.g. Hrung and Seligman, 2011; Fleming et al., 2010) generally find these programs to be effective in mitigating shortages of liquid Treasury collateral.

In addition, the literature also consistently supports the view that the ECB's liquidity support measures achieved their goal of additional liquidity provision. For instance, Abbassi and Linzert (2012) find a sizeable reduction in Euribor rates of more than 100 bp in their time-series estimations, attributable the increase in the aggregate amount of outstanding open market operations, whereas Angelini et al. (2011) detect a significant spread reduction of 10 to 15 basis points between secured and unsecured interbank loans for the announcement of LTROs after the Lehman shock. Similarly, Szczerbowicz (2015) further underlines that the announcement and implementation of 3-year LTROs reduced the spread between the Euribor and Overnight Index Swap (OIS) and consequently eased interbank lending significantly in an event-based regression.

2.3.2 Literature on asset purchase programs

The most frequently covered type of unconventional monetary policy in the literature are represented by large-scale asset purchase programs, in particular QE. After the implemen-

[8]While the BoE did not introduce a special liquidity provision program for banks at all, the Stimulating Bank Lending Facility introduced by the BoJ in 2012 did not receive considerable attention in academic literature.

tation of large-scale asset purchases by the BoJ shortly after the turn of the millennium and once the Fed and BoE adopted similar measures in 2008 and 2009, the number of empirical studies in this field started to proliferate (see, i.a., Bernanke et al., 2004; Gagnon et al., 2011; Joyce et al., 2011). However, as the implementation of equivalent large-scale asset purchase programs in the Eurozone took time until 2015, empirical studies on the ECB's QE measures were first conducted and published with some delay (see, e.g., Altavilla et al., 2015; De Santis, 2020). With regard to the effect of QE, two strands of literature focusing on different parts of the transmission mechanism can be identified. Studies on the *macroeconomic effects* of asset purchase programs most commonly apply VAR methodology and find positive lagged effects on output growth and inflation (see, e.g., Baumeister and Benati, 2013; Gambacorta et al., 2014).[9] In testing the effects to the real economy these studies often model QE shocks based on the findings from studies on the *financial market* effects of asset purchase programs (Baumeister and Benati, 2013).

Studies on *financial market effects* of QE are typically based on the assumption that financial market participants are forward looking and tend to price in information on monetary policy actions as soon as their expectations are affected. As a result, asset purchase programs and QE are generally found to be effective upon announcement (Altavilla et al., 2015; Fratzscher et al., 2018). The most commonly chosen empirical approach to study these announcement effects is that of an event study aka. event-based regression (see, i.a., Krishnamurthy and Vissing-Jorgensen; Swanson, 2011; Joyce et al., 2011). Following this approach, yield changes are observed around the time—typically within an event window of one to two days—of an unconventional monetary policy announcement by integrating the time of announcement into the model as a dummy variable.

Beyond proving the existence and identifying the magnitude of a decreasing effect on long-term yields, many event studies additionally try to disentangle and examine the distinct channels through which asset purchases affect long-term yields and financial conditions. In accordance with term structure theory, the majority of these event studies identify the signaling and portfolio rebalancing channel as the main transmission channels of QE.[10] For instance, Joyce et al. (2011) and Bauer and Rudebusch (2014) attribute changes in the OIS rate to the signalling channel and changes in the UK Gilt or the U.S. Treasury to OIS spreads

[9] In addition, Gambacorta et al. (2014) further document that these effects are non-persistent at the zero lower bound for the Eurozone.

[10] Via the signaling channel, announcements of asset purchases underpin expectations of future low short-term yields and thus lower long-term yields. In combination with the assumption of market segmentation, QE further decreases the risk premium on the purchased assets through reducing local supply.

to the portfolio rebalancing channel. In addition Krishnamurthy and Vissing-Jorgensen and D'Amico et al. (2012) empirically uncover additional (sub-)channels through which QE affects financial markets; among others, various safety premium channels and a duration risk channel.[11]

Considering the respective central bank addressed by each event study, the most commonly cited event studies are conducted on the data of unconventional programs in the U.S. (see, e.g., Gagnon et al., 2011; D'Amico et al., 2012; Bauer and Rudebusch, 2014) and the BoE, (see, e.g., Joyce et al., 2011; McLaren et al., 2014), or both Meaning and Zhu (2011); Christensen and Rudebusch (2012). Other event studies were performed on the asset purchase programs of the BoJ by Bernanke et al. (2004), Ugai (2007), and Ueda (2012). Overall, these studies unanimously report significant diminishing effects on long-term yields for all large-scale asset purchase programs and some even indicate significant, albeit small international spillover effects (see, e.g., Glick and Leduc, 2012; Neely, 2015). With regard to the Eurozone, event studies identifying the impact of QE on long-term yields of asset classes purchased in the course of the more recent large-scale asset purchase programs were performed, for instance by Altavilla et al. (2015) and De Santis (2020), who both find a negative yield effect on asset classes purchased in the course of the APP.

Beyond the analysis of the overall effects on yields, empirical studies on the effectiveness of asset purchase programs by observing inter-country yield spreads have been performed for the Eurozone (see, e.g., Szczerbowicz, 2015; Falagiarda and Reitz, 2015; Eser and Schwaab, 2016). In contrast to QE measures, the ECB's unique unconventional monetary policy measures prior to 2015, in particular targeted purchases of sovereign bonds, were aimed to solve Eurozone-individual challenges, such as increasing bond-spreads between member countries. In line with the goal of these targeted asset purchase programs, these empirical studies measure the country individual spread effects toward the German Bund rate as the risk free basis. In this regard, Eser and Schwaab (2016) found that the yield spread of periphery countries decreased significantly for the SMP, whereas Falagiarda and Reitz (2015) and Szczerbowicz (2015) extend this finding for both the SMP and OMT. While the main focus of most contributions to this area of research lies on the price and yield of a purchased asset, some papers additionally analyze the spillover effect on other asset classes as well. In this vein, Szczerbowicz (2015) documents that the CBPP caused a spillover effect on sovereign bond spreads, and conversely, SMP and OMT produced a similar effect

[11] The duration risk channel arises from the fact that central banks primarily purchase long-term bonds and hence reduce the level of duration risk in secondary markets. As a result, the term premium for all long-term securities may decrease (see, e.g. Vayanos and Vila, 2009)

on covered bond yields.

As event studies are primarily suited for identifying the significance of an initial yield drop around the announcement date of an asset purchase program, further econometric studies are often additionally applied in some of the previously mentioned papers, for statements on the long-run impact of asset purchase programs on bond rates and spreads. Overall, these studies find quantitatively smaller effects of asset purchase programs than event studies over the long-run. This result may be attributed to a strong initial announcement effect of purchase programs, which then subsides over time (Martin and Milas, 2012). Abstracting from event dummies, QE can be modelled either as a stock or as a flow variable to study changes in yields in time-series regressions. In this vein, Gagnon et al. (2011) and Joyce et al. (2011) base their estimates for yield changes on the *stock* of publicly held bonds, while Meaning and Zhu (2011) regress the yield curve on the size of the regular asset purchase *flows*. Lastly, across econometric methods applied, empirical studies can be further distinguished by type of data used (Martin and Milas, 2012). Specifically, empirical studies using a "historical data approach" (see, e.g. Joyce et al., 2011; Gagnon et al., 2011) assess the yield effect based on data from periods prior to the implementation of QE and additionally control for inflation and output movements, but only show the overall effect of various QE measures. By contrast, studies applying the "contemporary data approach" (see, e.g. Meaning and Zhu, 2011; Glick and Leduc, 2012; D'Amico and King, 2013) use data at a daily or even higher frequency within the period in which asset purchase programs took place to assess the effect of individual asset purchase programs by studying changes in the relationship between monetary policy and bond rates in times of financial distress. Taken together, across contributions studying different central banks and using different estimation techniques, there is a strong consensus on the effectiveness of asset purchase programs in lowering bond yields and spreads.

2.4 The model

The keynote of empirical literature is that central banks are able to influence long-term interest rates on bonds by purchasing them in the (secondary) bond market. In order to analytically capture and graphically depict the effects of asset purchase programs in a theoretical framework, we therefore need a model, which is capable of distinguishing the banking market from the bond market within the financial system. For our theoretical analysis we revisit the theoretical framework developed by Bofinger and Schächter (1995) and recently extended by Bofinger et al. (2017), which we set out shortly before applying it to the analysis of recent central bank asset purchase programs by the ECB, including QE.

The two most important insights of the model are the illustration of endogenous credit creation in the banking market (Palley, 1996; Disyatat, 2011; McLeay et al., 2014; Werner, 2014) and the development of the bond market where the created money is redistributed. In the banking market, banks are the suppliers of credit, while borrowers represent the demand side. After credit provision, banks can choose between a mixture of central bank credit, deposits, equity and bonds to refinance their businesses. In this environment, the central bank is able to influence the banking sector by controlling the refinancing rate, making it a key determinant of banks' credit supply. In the process of credit creation, banks create money, defined as the sum of cash and deposits, by creating additional deposits. In this vein, the difference between money and credit lies in its maturity. Money is a short-term concept on the liability side of the banking sector, whereas credit is recorded on the asset side of banks' balance sheets and refinanced with deposits, high-powered money, and longer-term refinancing sources, such as bonds and equity. Money holders have the option of holding money either in liquid (cash and deposits) or illiquid (bonds) form. In buying bonds they implicitly provide money to counterparts who have a liquidity shortage.[12] Thus, when credit is granted in the bond market, money is merely changing hands. In a financial system consisting of these two markets, borrowers have the option of demanding bank credit or demanding credit on the bond market. Beyond the interconnection of the two demand sides, the supply side of the banking market is linked to the bond market as well. Banks are able to refinance their credit business by issuing bonds in the bond market. Thus the cost of the banking sector also depends on the interest rate for bonds.

The model is described as follows: we first derive the equilibrium interest rate and credit volume on the banking market. After granting credit, banks demand a fixed proportion of credit of high-powered money for refinancing purposes (credit multiplier relation). In line with the equilibrium amount of credit, we derive the demand for high-powered money, which is fully satisfied by the central bank as the monopolistic supplier. In a final step, the equilibrium in the bond market is derived similarly to the banking market equilibrium.

[12] Note that, what we refer to as credit supply in the bond market is often called bond demand in the literature.

2.4.1 Banking market

To derive the equilibrium for the banking market, we need to set up the respective supply and demand functions. The market is in equilibrium when the supply of loans is equal to their demand.

Supply side Banks seek to maximize their profit. While for the banking and bond markets the revenue generated by granting credit depends on the interest rate spread between the interest rate for lending and that for borrowing (see for banks Spahn, 2013; Friedman, 2013, 2015), the cost structure differs for the banking and bond markets, with the banking sector facing higher costs. The reasons for the higher costs of the banking sector are that banks face higher credit risk due to the higher risk profile of its borrowers, and ultimately, specified capital requirements due to banks' higher risk profile.

We specify the profit function for one representative bank j as

$$\Pi_B^j = i_B Cr^j_{B/NB} - i_D D^j - i_R(Cr^j_{CB/B} - R^j) - i_E E^j - i_{NB} B^j - O^j - V_B^j, \quad \text{with } V_B^j = c_B (Cr^j_{B/NB})^2, \tag{2.1}$$

where revenues are determined by credit granted to non-banks $Cr_{B/NB}$ at the price of credit i_B. The costs for the banking sector consist of the interest paid on deposits $i_D D$, on the net refinancing costs arising from central bank refinancing $i_R(Cr_{CB/B} - R)$, on equity refinancing $i_E E$, and on the funds borrowed from the bond market $i_{NB} B$, plus operational costs O and credit risk costs V_B, whereby the latter are assumed to increase disproportionately with an increase in the credit volume (Fuhrmann, 1987).

Table 2.1: Simplified balance sheet of a bank j

Assets	Liabilites
Credit from bank to non-banks $Cr_{B/NB}$	Equity E
Reserves R	Bonds B
	Deposits D
	Credit from central bank to bank $Cr_{CB/B}$

Using the balance sheet identity according to the following balance sheet of a bank j (see Table 2.1), we can further derive

$$Cr^j_{CB/B} - R^j = Cr^j_{B/NB} - D^j - E^j - B^j. \tag{2.2}$$

In addition, we assume that a fixed proportion of credit granted to the non-banking sector $\eta^E = \frac{E^j}{Cr^j_{B/NB}}$ is held as equity according to the Basel Regulatory framework, and another proportion $\eta^B = \frac{B^j}{Cr^j_{B/NB}}$ is held as bonds to reduce interest rate risk (according to the Liquidity Coverage Ratio and Net Stable Funding Ratio declared in Basel III). This leads us to the following profit function (Equation 2.3). By maximizing Equation 2.3 with respect to credit volume and solving for $Cr^j_{B/NB}$, we receive the credit supply for a single bank j, which leads us to the credit supply for the banking sector (Equation 2.4) by summing over n homogeneous banks.

$$\Pi^j_B = (i_B - i_R) - \eta^E(i_E - i_R) - \eta^B(i_{NB} - i_R)Cr^j_{B/NB} - (i_D - i_R)D^j - O^j - c_B(Cr^j_{B/NB})^2 \quad (2.3)$$

$$Cr^S_{B/NB} = n\sum_{j=1}^{n} Cr^j_{B/NB} = n\left(\frac{(i_B - i_R) - \eta^E(i_E - i_R) - \eta^B(i_{NB} - i_R)}{2c_B}\right) \quad (2.4)$$

Demand side The demand for credit stems from borrowers (sovereigns, non-financial corporations, and households) that are usually driven by the desire to invest and/or consume (Minsky et al., 1993). Because of high entrance costs and the lack of opportunity to trade small volumes of credit on the bond market, the two types of credit (bank credit and bonds) represent imperfect substitutes and the cost of credit is different for each market. Consequently, apart from the economy's aggregate income, the determinants of credit demand are the spread between the interest rate for credit in the respective market and the credit interest rate in the substitution market.

The amount of credit demanded depends negatively on the respective price, where the saturation amount a is a linear function of income. Furthermore, the demand for bank loans depends positively on the price for the substitute loan type and the substitution elasticity d, ranging from values of 0 (fully independent loans) to ∞ (perfect substitutes).[13] This yields the following demand function for bank loans:

$$Cr^D_B = a - bi_B + d(i_B - i_{NB}), \quad (2.5)$$

Equilibrium Assuming $n = 1$ and solving the equilibrium condition for the banking market, we get

$$Cr^*_{B/NB} = \frac{a - (b+d)(i_R + \eta^E(i_E - i_R) + \eta^B(i_{NB} - i_R)}{1 + 2c_B(b+d)} \quad (2.6)$$

$$i^*_B = \frac{2c_B(a + di_{NB}) + (i_R + \eta^E(i_E - i_R) + \eta^B(i_{NB} - i_R))}{1 + 2c_B(b+d)}. \quad (2.7)$$

[13] The demand function with respect to the substitutability is derived by Singh and Vives (1984), Wied-Nebbeling (2013), and Ledvina and Sircar (2011).

Bank credit multiplier In granting credit, banks simultaneously demand high-powered money in accordance with their liability structure. In order to derive the fraction of credit refinanced by high-powered money, we first need to define a bank credit multiplier m_B, which is the ratio of credit from banks to non-banks $Cr_{B/NB}$ to high-powered money H. As money consists of cash and deposits, high-powered money consists of cash and reserves, and $Cr_{B/NB}$ can be rewritten as $\frac{M}{1-\eta^E-\eta^B}$, the money multiplier can be redefined as follows:

$$m_B = \frac{Cr_{B/NB}}{H} = \left(\frac{1+h}{h+r}\right)\left(\frac{1}{1-\eta^E-\eta^B}\right), \tag{2.8}$$

where h represents the cash holding coefficients of the public and r the minimum reserve requirements, both of which are calculated as fractions of deposits. Assuming $\eta^E+\eta^B<1$, $h<1$ and $r<1$, the bank credit multiplier is always greater than one.

Market for high-powered money The demand for high-powered money is determined by the volume of bank credit at a given refinancing rate. For the derivation of the high-powered money demand function, we need to obtain two points, which suffice to pin down the function due to its linearity. First, we use the equilibrium amount of credit granted ($Cr^*_{B/NB}$) to obtain the demanded volume of high-powered money (H^*) over the multiplier relation at the respective refinancing rate (i_{R_0}). Second, we determine the refinancing rate at which the demand for high-powered money equals zero. By subtracting the spread for equity and bond refinancing from the prohibitive price of credit demand, we obtain the refinancing rate at which the volume of granted credit is equal to zero and consequently the demand for high-powered money is equal to zero as well. Analytically, the demand function for high-powered money is defined as:

$$H^D = e\frac{m_B}{Cr^*_{B/NB}}(e-1) - \frac{m_B}{Cr^*_{B/NB}}(e-i_{R_0})i_R$$
$$\text{with } e = (\frac{a+di_{NB}}{b+d}) - \eta^E(i_E-i_R) - \eta^B(i_{NB}-i_R).$$

Since the central bank serves as a monopolistic supplier of high-powered money, it meets the full demand for high-powered money at the fixed price of the refinancing interest rate.

2.4.2 Bond market

Once money is created in the process of bank lending, it can be used for buying bonds in the bond market.[14] The bond market functions similarly to the banking market. But in contrast to the banking market's role as the platform for money creation, the bond market is the platform for money circulation, where money is reused multiple times in order to create credit.

Supply side The revenues of the bond market suppliers are determined by the spread between the interest rate for long-term lending and the deposit rate because investors can only choose between either holding money as deposits or lending it. In contrast to the banking sector, non-banks do not face any cost due to capital requirements and their cost due to interest rate risk arise from opportunity costs of holding money as deposits. Consequently, the profit function of a non-bank k appears as follows:

$$\Pi^k_{NB} = i_{NB}Cr^k_{NB} - i_D Cr^k_{NB} + (\frac{i_{NB}}{i^e_{t+1}} - \frac{i_{NB}}{i_t})Cr^k_{NB} - I^k - V^k_{NB}, \quad \text{with } V^k_{NB} = c_{NB}(Cr^k_{NB})^2. \tag{2.9}$$

Revenues are determined by inflows emerging from the credit business $i_{NB}Cr^k_{NB}$. The costs stemming from granting credit are opportunity costs $i_D Cr^k_{NB}$, and those from potential bond price losses, which determine the so called *term premium*, are depicted in the term $(\frac{i_{NB}}{i^e_{t+1}} - \frac{i_{NB}}{i_t})Cr^k_{NB}$, according to which an increase in the expected interest rate i^e_{t+1} results in losses on bonds. Furthermore, information cost I^k and credit risk costs V^k_{NB} add to the costs faced by non-banks. For the purpose of simplicity, we assume that $i_D = i_R$ and bonds are priced at par, yielding to $i_{NB} = i_t$. After maximizing the resulting profit function (Equation 2.10) with respect to credit volume and solving for Cr^k_{NB}, we receive the credit supply for a single non-bank k, which we convert to the credit supply for the non-banking sector by summing it up for m homogeneous non-banks (Equation 2.11):

$$\Pi^k_{NB} = (i_{NB} - i_D)Cr^k_{NB} + (\frac{i_{NB}}{i^e_{t+1}} - 1)Cr^k_{NB} - I^k - c_{NB}(Cr^k_{NB})^2 \tag{2.10}$$

$$Cr^S_{NB} = m\sum_{k=1}^{m} Cr^k_{NB} = m\left(\frac{(i_{NB} - i_R) + (\frac{i_{NB}}{i^e_{t+1}} - 1)}{2c_{NB}}\right). \tag{2.11}$$

[14] For the derivation of the bond market we assume that no additional funds from the banking market are created only to directly flow into the bond market.

Demand side Alongside sovereigns and non-financial corporations, banks are a major borrower in the bond market. In line with regulatory requirements, banks demand credit on the bond market in order to reduce the maturity mismatch in the balance sheet, which results from their business model of lending long and borrowing short. The determinants of credit demand in the bond market are the given economy's income, the cost of credit, and cost of credit of the substitute loan type, similar to those in the banking market. This yields the following demand function:

$$\begin{aligned} Cr^D_{NB} &= a - bi_{NB} + d(i_B - i_{NB}), \\ &\text{with } a = \mu + \gamma Y. \end{aligned} \tag{2.12}$$

Equilibrium After equating credit demand with supply, we obtain the equilibrium amount of credit and interest rate in the bond market:

$$Cr^*_{NB} = \frac{(a + di_B)\left(\frac{i^e_{t+1}+1}{i^e_{t+1}}\right) - (b+d)(i_R + 1)}{\frac{i^e_{t+1}+1}{i^e_{t+1}} + 2c_{NB}(b+d)}, \tag{2.13}$$

$$i^*_{NB} = \frac{2c_{NB}(a + di_B) + i_R + 1}{\frac{i^e_{t+1}+1}{i^e_{t+1}} + 2c_{NB}(b+d)}. \tag{2.14}$$

Comparing the equilibria in the banking and the bond market, we detect asymmetry with regard to interest rates and credit volumes, which is a result of differing costs on the supply sides. However, bank loans and bonds coexist in equilibrium due to institutional factors.[15]

2.4.3 Graphical illustration of the model equilibrium

In Figure 2.3 we graphically illustrate the bond market. In contrast to the intercept of the loan supply in the banking market, which is determined by the refinancing rate, the cost of equity, and the cost of bonds, the intercept in the bond market is set by the refinancing rate and the interest rate expectations in the bond market. The equilibrium amount of non-bank credit $Cr^*_{NB/NB}$ and the interest rate i^*_{NB} in the bond market lies at the intersection of the—in comparison with the banking market—similarly shaped demand curve and the flatter supply curve.[16]

[15] Banks' money creation is a prerequisite for the functioning of the bond market, while regulatory requirements underline the necessity of the bond market for the money creation by banks.

[16] As aforementioned, this difference in slope is due to non-bank suppliers facing lower costs than the banking sector.

Figure 2.3: Bond market in the model

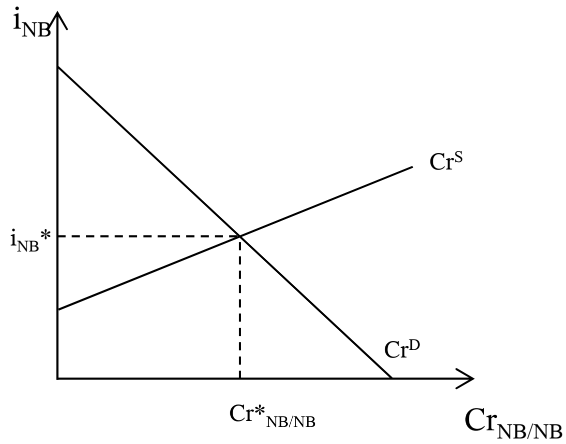
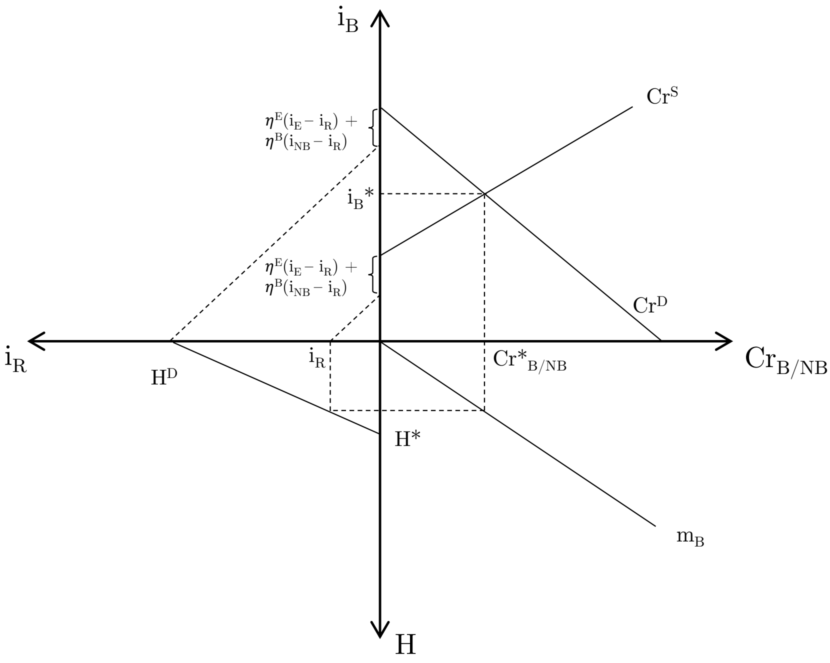


Figure 2.4: Bank credit market in the model



With regard to the market for bank credit, depicted in Figure 2.4, the equilibrium amount of credit $Cr^*_{B/NB}$ and the interest rate i^*_B lie at the intersection of the negatively sloped loan demand curve and the positively sloped loan supply curve. By inserting the equilibrium amount of bank credit from the credit market (first quadrant) into the bank credit multiplier relation m_B with a slope of > 1, we obtain the demanded amount of high-powered money H (second quadrant). This demand for high-powered money H is visualized in a negatively sloped demand function for high-powered money at the price of i_R (third quadrant).

2.4.4 Application to unconventional monetary policy

Due to the careful distinction of the intertwined banking and bond market, this model is particularly suited to illustrate the differing effects of announcements on liquidity support measures and asset purchase programs on the financial system.

Liquidity support measures target the liability side of the *banking sector's* balance sheet. In the context of a distressed financial market, these measures offer the banking system an opportunity to ameliorate the maturity mismatch by refinancing with lower-yield central bank loans instead of high-yield bonds. As a result, the proportion of borrowing conducted by the banking sector in the bond market, η^B, declines.

Asset purchase programs, by contrast, target the asset side of the balance sheets of suppliers of financing on the *bond market.* In line with empirical literature on the effects of central bank asset purchases, in particular QE measures, are characterized by three effects in our model. These are identified along the following hypotheses:

Hypothesis 1: decline in bond yields

In line with empirical literature described in Section 2.3, we expect a decline in bond yields due to the intervention of the ECB on the bond market. The ECB acts as additional supplier of liquidity that is able to shift the supply curve to the right, which ceteris paribus leads to a decrease in bond yields. Assuming forward looking agents on the bond market, these effects are already taken into account upon announcements and influence the behavior of the bond market supply side in our model, which leads us to the second and third hypothesis.

Hypothesis 2 : decrease in credit risk

Since the quality of outstanding credit deteriorates in times of financial turmoil, the credit risk of these assets increases. By acting as a "lender of confidence", the ECB helps to *decrease the (perceived) credit risk* of bonds on secondary markets issued by sovereigns (SMP, OMT and PSPP), banks (CBPP), and non-financial corporations (CSPP) by purchasing the respective assets. The consequent decline in credit risk costs c_{NB} ceteris paribus lowers interest rates in the bond market i_{NB}. This effect becomes visible upon the announcement of asset purchase programs by the ECB, as forward looking agents on the bond market price this effect in as soon as the impact on the central bank balance sheet of a program is foreseeable.

Hypothesis 3: decrease in long-term interest rate expectations

Additionally, the central bank's interventions influence the *expectations on medium to long-term interest rates*, i.e. term premium. However, the effect on term premia could be ambiguous over time. If bond market participants expect an ongoing decline in interest rates due to further asset purchase programs based on the previous hypothesis, this will be reflected in decreasing term premia. At the same time, bondholders may expect a rise in long-term bond rates over time, as soon as asset purchase programs are tapered, due to the fact that the central bank is not able to lower the short-term interest rate further at the zero lower bound. Hence, it may be possible that QE programs lower term premia in the short-run, whereas in the long-run this effect diminishes.

Graphical application for the Eurozone Within our model, the effects of the ECB's liquidity support measures and asset purchase programs are twofold. First, asset purchases lead to a clock-wise rotation of the credit supply curve on the bond market due to reduced credit risk costs ($c_{NB_0} \rightarrow c_{NB_1}$), and a parallel downward shift due to lower interest rate expectations in the short-run ($i^e_{t+1_0} \rightarrow i^e_{1+1_1}$), resulting in a lower equilibrium bond market interest rate ($i^*_{NB_0} \rightarrow i^*_{NB_1}$).[17] In the banking market, the lower interest rate on bonds ($i_{NB_0} \rightarrow i_{NB_1}$) and the reduction in the proportion of lending in the bond market ($\eta^B_0 \rightarrow \eta^B_1$) due to the provision of liquidity support measures lead to a parallel downward shift in the credit supply curve of the banking sector. This results in a lower interest rate and an increase in credit volume in the banking market. Due to the shift in bank's financing structure away from refinancing in the bond market and toward central bank refinancing, e.g., via liquidity support measures, the bank credit multiplier declines. As a result the demand for high-powered money increases.

[17]We assume that at the new equilibrium $d(i_{B_0} - i_{NB_0}) = d(i_{B_1} - i_{NB_1})$ in order to abstract from demand side effects.

Figure 2.5: Bond market effect of unconventional monetary policy in the model

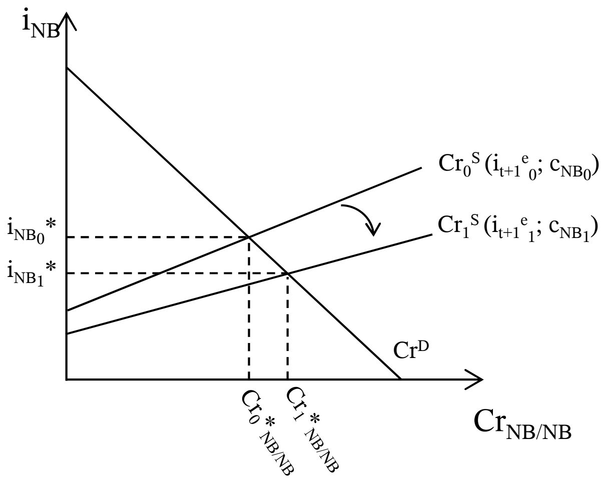
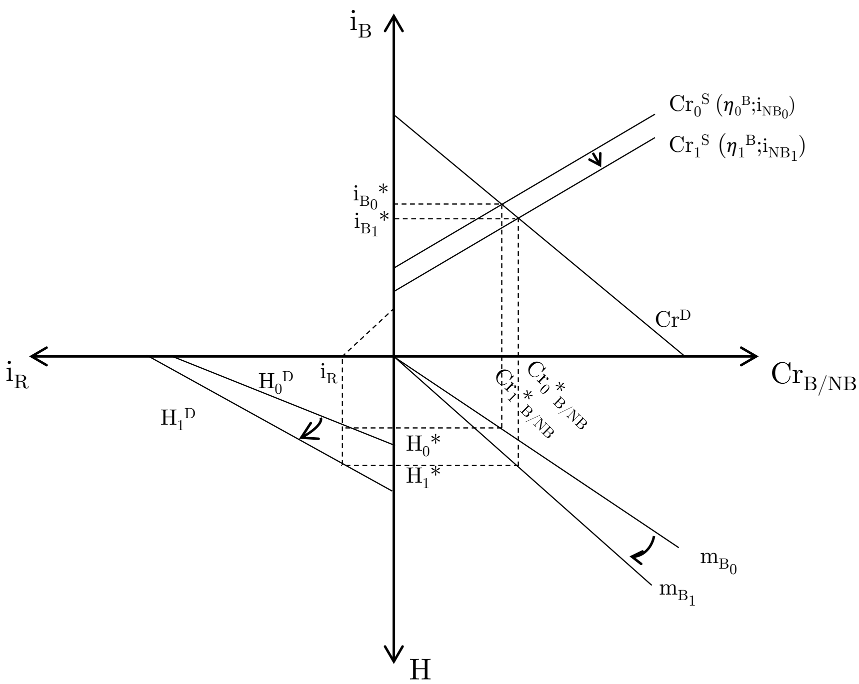


Figure 2.6: Bank credit market effect of unconventional monetary policy in the model



2.5 Empirical evidence

In the following, we test the previously presented hypotheses on the effect and transmission of central bank asset purchases on the bond market, in particular on the sovereign bond market in daily data from January 1, 2008 to September 30, 2016.[18] First, we test the hypothesis that the announcements of asset purchases lead to a decline in bond yields (hypothesis 1) by applying an error correction model, which provides the advantage of addressing both the long and the short-run effects of our theoretical model. In a second step, we apply an event-based regression to isolate the transmission of the yield-depressing effect asset purchase programs exert through reducing credit risk (hypothesis 2) and medium to long-term interest rate expectations (hypothesis 3).

A possible issue of empirical evaluations in the context of the ECB's asset purchase programs is that they might be endogenous as soon as the ECB reacts to market developments such as, e.g., a rise in credit spreads. We deal with this issue by following Fratzscher et al. (2018) in assuming that the announcements were of the "leaning-against-the-wind" type and by employing announcement event-dummies rather than the actual policy instrument as well as appropriate controls that proxy for market conditions.

2.5.1 Error correction model

The methodology of an error correction model was first applied by Sargan (1964) in the context of wage and price adjustments in the UK. Particularly within the framework of financial markets, many authors have estimated the long-run money demand equation or interest rate adjustments using an error correction model (Mehra, 1993; Heffernan, 1997; Winker, 1999; Dreger and Wolters, 2015).

In the previous Section, we derived the long-run equilibrium for the bond market (see Equations 2.13 and 2.14). When estimating this equilibrium in levels, we face the problem of spurious regression results due to non-stationary time series (see Appendix, Tables 2.8–2.13). In order to solve this problem, we apply an error correction model, which is based on the assumption that a long-run equilibrium relationship exists, while in the short run we observe disturbances that lead to a divergence from the equilibrium. Based on this distinction

[18] We do not consider the effects in the banking market, which arose via liquidity support measures due to the non-availability of banking data at a sufficiently high-frequency.

between long and short-run effects, we now present the two parts of the error correction model following Sargan (1964) and Davidson et al. (1978).

First, we identify the long-run relationship, which is explained by our theoretical model. Using daily data for our estimation and excluding bank interest rates, due to their non-availability on a daily basis, we define

$$i_{NB_t} = \alpha_0 + \alpha_1 i_{R_t} + \alpha_2 log(Y_t) + \alpha_3 c_{NB_t} + \alpha_4 i_t^e + u_t, \tag{2.15}$$

consisting of the sovereign bond yield i_{NB_t}; the refinancing rate of banks i_{R_t}; the log of income in the current period $log(Y_t)$; the credit risk costs c_{NB_t}; interest rate expectations for bonds within the same maturity segment of the respective purchased government bond i_t^e; and the error term of the long-run model u_t. All variables are specified as levels at time t except for income, which is indicated in log-levels.

Second, the short-run relationship is established by applying the variables of the long-run model in first differences. Accordingly, we obtain the following equation:

$$\begin{aligned}\Delta i_{NB_t} = \beta_0 + \beta_1 \sum_{n=0}^{N_1} \Delta i_{R_{t-n}} + \beta_2 \sum_{n=0}^{N_2} \Delta log(Y_{t-n}) + \beta_3 \sum_{n=0}^{N_3} \Delta c_{NB_{t-n}} \\ + \beta_4 \sum_{n=0}^{N_4} \Delta i_{t-n}^e + \beta_5 \sum_{n=0}^{N_5} \Delta i_{NB_{t-n}} + \beta_6 u_{t-1} + \epsilon_t.\end{aligned} \tag{2.16}$$

For the short-run equation the variables are similar to those of the long-run equation, but are defined in first differences with current and past lags for $N_i = \{1, 2, 3, ...\}$. u_{t-1} is the lagged disturbance term of the long-run equation and ϵ_t is the short-run error term. Its corresponding coefficient β_6 is the adjustment term of the short-run equation. It states that the interest rate of government bonds deviating from the equilibrium converges toward it.

To maintain the validity of the error correction model, the interest rate of government bonds must not diverge from the long-run equilibrium, requiring u_t in the long-run equation to be stationary and the coefficient of u_{t-1} in the short-run equation to be negative. Based on this assumption, there are two possible ways to estimate the error correction model. One method is to estimate Equation 2.15 and plug the obtained error term into Equation 2.16, while alternatively Equation 2.15 on the long-run relationship can be carried back by one period and inserted for u_{t-1} in Equation 2.16 on the short-run relationship (Stock, 1987).

Applying the latter method, we obtain

$$\begin{aligned}
\Delta i_{NB_t} = \theta_0 + \beta_1 \sum_{n=0}^{N_1} \Delta i_{R_{t-n}} + \beta_2 \sum_{n=0}^{N_2} \Delta log(Y_{t-n}) \\
+ \beta_3 \sum_{n=0}^{N_3} \Delta c_{NB_{t-n}} + \beta_4 \sum_{n=0}^{N_4} \Delta i^e_{t-n} + \beta_5 \sum_{n=0}^{N_5} \Delta i_{NB_{t-n}} \\
+ \theta_1 i_{NB_{t-1}} + \theta_2 i_{R_{t-1}} + \theta_3 log(Y_{t-1}) + \theta_4 c_{NB_{t-1}} + \theta_5 i^e_{t-1} + \epsilon_t,
\end{aligned} \tag{2.17}$$

where the coefficients are defined as follows:

$$\begin{aligned}
\theta_0 &= \beta_0 - \beta_6 \alpha_0; \\
\theta_1 &= \beta_6; \\
\theta_2 &= -\beta_6 \alpha_1; \\
\theta_3 &= -\beta_6 \alpha_2; \\
\theta_4 &= -\beta_6 \alpha_3; \\
\theta_5 &= -\beta_6 \alpha_4.
\end{aligned}$$

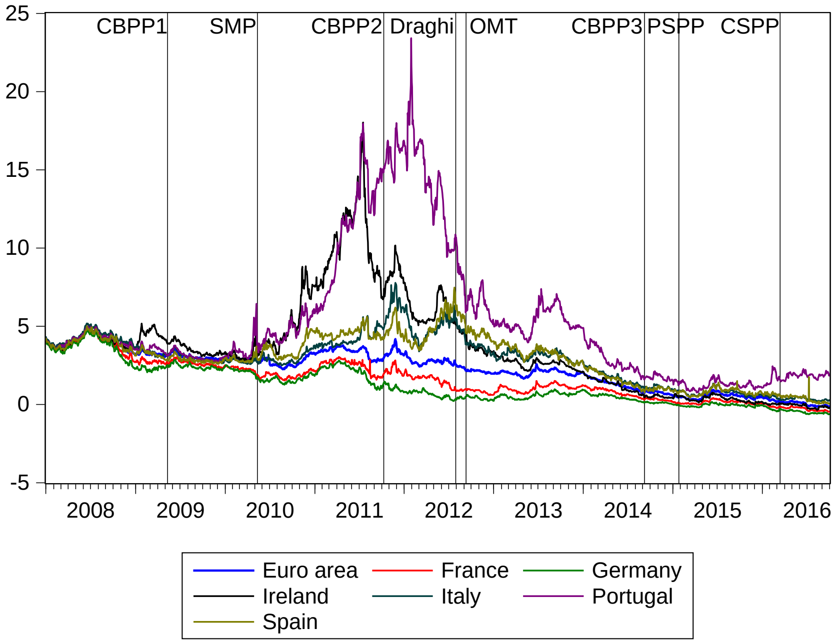
The short-run coefficients (β_1 to β_5), as well as (θ_1 to θ_5) are directly obtained from estimating Equation 2.17. To obtain the long-run effects (α_1 to α_5) from Equation 2.15 we reverse the previous calculation for the short-run effects from Equation 2.17. In doing so, we calculate the long-run coefficient of the refinancing rate $i_{R,t}$, that is α_1, for instance, by dividing θ_2 by $-\theta_1$, which is equal to $-\beta_6$.

Data

For our estimations we use daily data from January 1, 2008 through September 30, 2016 in order to evaluate the effect of the ECB's different asset purchase programs on 5-year sovereign bond yields of France, Germany, Ireland, Italy, Portugal, Spain, and the Euro area. We choose to focus on a maturity of five years due to the focus of the ECB's purchases on bonds of this maturity in the SMP and since it lies between the shorter maturities targeted by the OMT of 1 to 3 years and the longer maturities targeted by the PSPP of mostly 10 years. The evolution of sovereign bond rates for the selected countries and the Euro area within our sample are shown in Figure 2.7 along with unconventional monetary policy announcements represented by vertical lines.

Prior to the financial crisis the sovereign bond yields for all Euro area countries coincided, except for small deviations, occurring in early 2009. Since the start of the sovereign debt

Figure 2.7: Sovereign bond yields



Notes: Yields for French, German, Irish, Italian, Portuguese and Spanish 5-year sovereign bonds and the 5-year Euro area benchmark bond are plotted in basis points for a daily data sample from January 1, 2008 to September 30, 2016. Vertical lines mark the days on which announcements on asset purchase programs were made by the ECB.

crisis in 2010, however, the sovereign bond rates of Spain, Portugal, Ireland, and Italy have started to increase and to diverge considerably from the government bond rates of France and Germany. Particularly Ireland and Portugal, which received financial support from the EFSF and ESM, experienced very high interest rates from 2010 until the end of 2012. Since the end of 2012, interest rates have declined and reached even lower levels than in pre-crisis data by 2014. Over the entire time horizon, the sovereign bond yields for Germany and France stayed the lowest, which underlines their status as a "safe haven" for investors in the Euro area. Graphically, the yield on the Euro area benchmark bond separates the countries, which suffered from the sovereign debt crisis (Italy, Ireland, Portugal, and Spain) from these "safe haven" countries (Germany and France).

Beyond the course of sovereign bond yields, Figure 2.7 also displays the announcement days of the ECB's asset purchase programs. Following the announcements with regard to sovereign bond markets, i.e., SMP, Draghi's speech[19], OMT and PSPP, were followed by a decline in sovereign bond yields. In detail, the PSPP appears to have contributed

[19] In a speech at the Global Investment Conference in London on July 26, 2012, Mario Draghi stated that "Within our mandate, the ECB is ready to do whatever it takes to preserve the euro. And believe me, it will be enough."(Draghi, 2012). Despite not containing information on specific measures these words are often interpreted as an implicit announcement of further asset purchase programs by the ECB.

to the convergence of sovereign bond yields, whereas programs for banks (CBPP1-3) and non-financial corporations (CSPP) seem to have left interest rates of sovereign bonds largely unaffected.

As previously implied by our estimation equation, a key determinant for sovereign bond yield is the short-term refinancing rate. As a daily measure of this variable we use the EONIA across the country-individual and Euro area estimations. The course of the EONIA over the selected sample is depicted in Panel (a) of Figure 2.8 and suggests a link to the course of sovereign bond yields depicted in Figure 2.7. For instance, the temporary increase in the EONIA in 2011 was followed by an increase in government bond yields, whereas its decline into negative territory in 2016 is accompanied by low sovereign bond yields and even negative ones in the case of German sovereign bonds.

Another factor for the course of sovereign bond yields, which we derive from our theoretical model, are credit risks. As a daily measure of credit risk we use country-individual CDS spreads, which reflect the price for credit insurance, and thus the perceived default risk of each borrower. Due to the unavailability of CDS spreads for the Euro area benchmark bond, we proxy credit risk by the bond spread relative to the German sovereign bond.[20] The respective time series are depicted in Panel (b) of Figure 2.8 and show a quite similar qualitative development of credit risk spreads to that of sovereign bond yields: after the SMP, Draghi's speech and the OMT, CDS spreads for Ireland, Italy, Spain, and Portugal fell severely, while German and French CDS spreads—already at low levels—experienced a smaller reduction for these announcements.[21] From the set of unconventional monetary policy announcements, Draghi's speech seems to have reduced the CDS spreads for all observed Euro area countries most visibly.

To additionally capture the effect of long-term interest rate expectations, and implicitly of expected capital losses on long-term bonds, we compute the term premium for each country.[22] In absence of a daily series for term premia, we proxy interest rate expectations by the difference between the current bond rate with a maturity of 5 years and the mean of the EONIA forward rates of 1,2,3,4, and 5 years and plot these values in Panel (c) of Figure 2.8. During the financial crisis investors have perceived the risk of capital losses associated

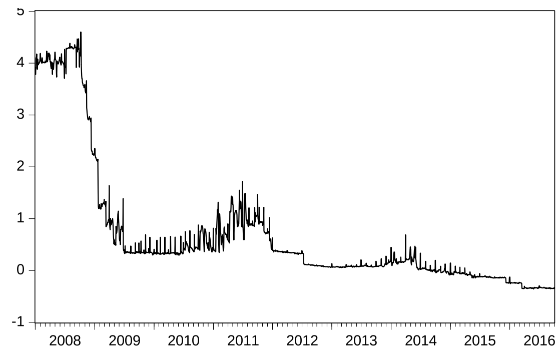
[20] The bond spread is calculated as the difference between the yield on the Euro area benchmark bond and the German sovereign bond yield.

[21] Note, that these qualitative observations also align with the findings by Gerlach-Kristen (2015).

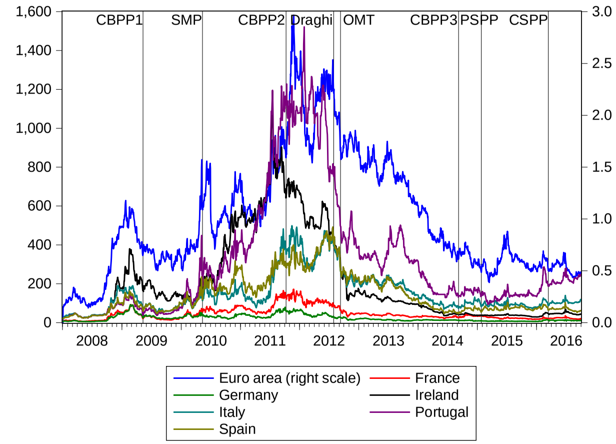
[22] The notion behind term premia is that risk averse lenders want to be compensated for the risk of capital losses throughout their holding period, i.e., they demand term premia (see Gürkaynak and Wright, 2012).

Figure 2.8: Time series entering the model

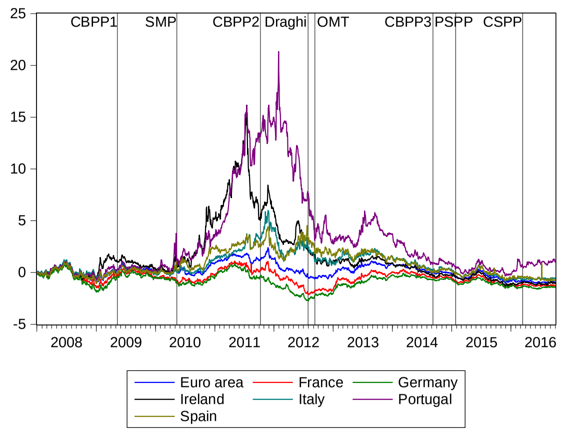
(a) EONIA



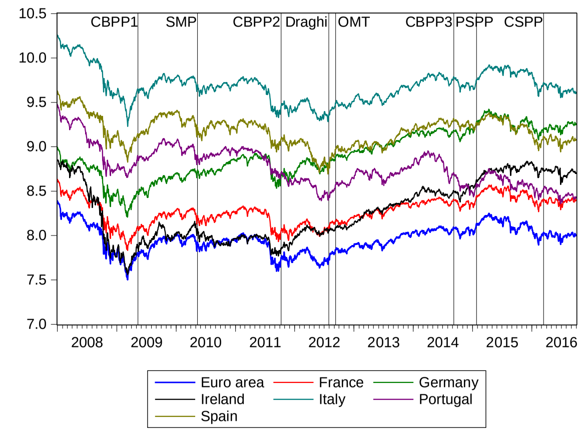
(b) Credit risk spread



(c) Term premium



(d) Equity indexes



Notes: The time series entering our model are plotted for a daily data sample from January 1, 2008 to September 30, 2016. Panel (a) shows the EONIA in percent, Panel (b) depicts credit risk spreads of 5-year French, German, Irish, Italian, Portuguese, and Spanish sovereign bonds and the bond spread of the 5-year European benchmark bond in basis points, Panel (c) displays the respective term premium for each bond in percent and Panel (d) the respective country equity indexes entering the model in logarithmic values. Vertical lines mark the days on which announcements on asset purchase programs were made by the ECB.

with longer maturities as high, in particular for periphery countries. During the Euro area debt crisis, however, we even observe negative term premia for Germany and France, which suggest that investors prefer fixed interest over the entire investment horizon to fluctuating interest rates of shorter-term investments. As for the credit risk spreads, after Draghi's speech the term premia across observed countries have converged to pre-crisis levels.

Lastly, to proxy for aggregate income on a daily basis, we use country-individual and Euro area equity indexes within our sample. As a quantity variable income enters our estimation equation in log levels. Panel (d) of Figure 2.8 reflects the course of the logged equity indexes, which is suggestive of a positive influence of the APP on equity markets. While in 2009 after the burst of the housing bubble, Euro area equity markets were at their lowest levels, the equity indexes have increased for all observed countries after the announcement of the PSPP.

Unit roots and cointegration

To examine our time series on the presence of unit roots, we use the augmented Dickey-Fuller test, the Phillips-Perron test, and the KPSS test.[23] For almost every time series we apply the tests to, the results indicate non-stationarity, with the exception of the equity indexes, for which the augmented Dickey-Fuller and the Phillips-Perron tests suggest stationarity, while the KPSS test indicates non-stationarity. Consequently, we assume that each equity index has a unit root for our long-run model. When we regress a non-stationary variable on other non-stationary variables, cointegration of these variables should lead to stationary results. If this holds true, the linear combination of the variables is stationary as well. In order to test this assumption, we apply the Johansen cointegration method. For each country, the test results of the trace and maximum eigenvalue tests reveal at least one cointegrated equation at the 5% significance level.

Regression setup

Building on these preliminary test results, we estimate Equation 2.17 by appyling the previously specified error correction model methodology. To explicitly capture the effects of asset purchase programs, we implement event dummies that equal 1 for the announcement days of the respective asset purchase program.[24] In line with Szczerbowicz (2015), we

[23]For test results, see Appendix, Tables 2.8–2.13.
[24]An overview of these events can be found on the right hand side of Figure 2.2

expect spillover effects on sovereign bond yields to prevail for asset purchase programs that are targeted towards securities issued by banks and non-financial corporations. To allow for such effects, we thus include announcements of the CBPP, CBPP2, and CBPP3 and the CSPP. Following Gerlach-Kristen (2015), who studies the effects on CDS-spreads attributable to the ECB's unconventional monetary policy measures, we additionally control for the effects from bail-outs for Greece (May 5, 2010 and July 22, 2011), Portugal (May 16, 2011), and Ireland (November 22, 2010) and from the sovereign default of Greece (February 21, 2012). By including all relevant ECB purchase programs as well as the control events as dummy variables as well as in Equation 2.17, we receive the following equation:

$$\begin{aligned}
\Delta i_{NB_t} &= \theta_0 + \theta_1 i_{NB_{t-1}} + \theta_2 i_{R_{t-1}} + \theta_3 log(Y_{t-1}) + \theta_4 c_{NB_{t-1}} + \theta_5 i^e_{t-1} \\
&+ \beta_1 \sum_{n=0}^{N_1} \Delta i_{R_t - n} + \beta_2 \sum_{n=0}^{N_2} \Delta log(Y_{t-n}) + \beta_3 \sum_{n=0}^{N_3} \Delta c_{NB_{t-n}} \\
&+ \beta_4 \sum_{n=0}^{N_4} \Delta i^e_{t-n} + \beta_5 \sum_{n=0}^{N_5} \Delta i_{NB_{t-n}} \\
&+ \beta_6 \text{CBPP}_t + \beta_7 \text{OMT}_t + \beta_8 \text{SMP}_t + \beta_9 \text{PSPP}_t + \beta_{10} \text{CSPP}_t + \beta_{11} \text{Draghi's speech}_t \\
&+ \beta_{12} \text{Greece}_t + \beta_{13} \text{GreeceDefault}_t + \beta_{14} \text{Portugal}_t + \beta_{15} \text{Ireland}_t + \epsilon_t,
\end{aligned} \tag{2.18}$$

where the coefficients β_6 to β_{11} measure the effect of the respective announcement on the change in sovereign yields Δi_{NB_t}.

Regression results

We estimate our model for the full sample (see Table 2.2) and for three subsamples, i.e., for the pre-crisis, the crisis, and post-crisis period (see Tables 2.3, 2.4, and 2.5).[25]

For the full sample, the error correction model seems appropriate by delivers significant results with a negative sign for the lagged bond yield for France, Italy, Ireland, Portugal and Spain. By contrast, the coefficients for Germany and the Euro area are insignificant despite showing the expected sign. For the pre-crises sample the error correction model seems appropriate by showing highly significant negative coefficients across all selected countries. In the case of France, Ireland, Spain, and the Euro area this holds for the crisis and post-crisis subsample, as well, whereas for Germany and Italy we observe indications that the error correction model does not reveal the expected effects in at least one sub-sample.

[25] The sample split is suggested by multiple breakpoint tests pointing towards structural breakpoints in the time series around the days at the beginning (April 22, 2010) and end (August 1, 2012) of the crisis.

With regard to the long-run coefficients (lagged credit risk spread, lagged EONIA, lagged equity, and lagged term premium) the full-sample estimation delivers the expected positive sign for the coefficients of the Euro area, France, Italy, Spain, and Portugal, but only some coefficients are significant. Interestingly, we find a significant negative effect of the EONIA on bond yields during the crisis, again with the exception of Germany and the Euro area. This finding may point to the fact that transmission of the ECB's conventional monetary policy tool was impaired, due to the distorting factors (as, e.g., uncertainty) during this period. While the lagged level of the CDS spread appears to be of minor importance for changes in sovereign bond yields, the sign of the coefficient for the term premium indicates its significant positive influence on bond yields in most of the subsamples and for the full sample, which is underlined by the consistently positive and significant coefficients of *changes* in the term premium for all individual countries. With regard to the change in the term premium for the yield in the Euro area benchmark bond, the influence appears to be more ambiguous. With regard to the announcements of asset purchase programs not directly related to sovereign bond markets, in line with our expectations, the CBPP1, CBPP2, CBPP3 and the CSPP show significant coefficients, albeit heterogeneous in sign and size across countries and announced programs. This finding could be interpreted in favor of asset substitution across asset classes in the course of portfolio balancing effects.

The coefficients of asset purchase programs targeting the sovereign bond market show a different profile. While PSPP has a significant negative effect on the bond yields for every observed country, bond yields for Germany and France increased on the announcement days of the SMP and the OMT. A plausible explanation for this mixed result may be that prior to the PSPP the ECB only acted as a "lender of confidence" for the countries most heavily affected by the Euro area crisis. A different interpretation may be that investors of German and French sovereign bonds, substituted their bond holdings for the bonds of periphery countries to sell them to the ECB. The finding of a rise in the 5-year bond yields of Spain, Italy, and Ireland with the announcement of the OMT can in turn be explained by the fact that this program was designed to purchase bonds with a shorter maturity, ranging between 1 and 3 years. In line with the qualitative course of sovereign bond yields from Figure 2.7, our estimation results show a negative significant effect on the bond yields of Germany, France, Italy, Portugal and the Euro area for Draghi's "Whatever it takes".

Table 2.2: Regressions for ECM—full sample

	France	Germany	Ireland	Italy	Portugal	Spain	Euro area
Constant	-0.042	0.055	0.082	-0.150	-0.077	-0.225***	-0.081
Lagged variables							
Bond yield, $(i_{NB_{t-1}})$	-0.003*	-0.004	-0.005**	-0.002*	-0.005***	-0.005***	-0.001
Credit risk spread, $c_{NB_{t-1}}$	3.62E-05	-4.18E-05	1.24E-05	2.43E-05	1.62E-05	8.17E-05***	0.002
EONIA, $i_{R_{t-1}}$	0.002	0.002	-0.001	0.001	0.003**	0.003**	0.001
log(Equity index), $log(Y_{t-1})$	0.005	-0.006	-0.009	0.015	0.009	0.024***	0.010
Term premium, i^e_{t-1}	0.003*	0.003	0.004**	0.003*	0.004***	0.002	0.001
First differences							
ΔCredit risk spread , Δc_{NB_t}	0.001	-0.002***	0.001***	0.001***	2.40E-04***	0.001***	0.342***
ΔEONIA, Δi_{R_t}	-0.013	-0.003	-0.017	-0.015	-0.018*	-0.017*	-0.006
Δlog(Equity index), $\Delta log(Y_t)$	0.262***	0.602***	-0.066	-0.035	-0.086	0.023***	0.550***
ΔTerm premium, Δi^e_t	0.660***	0.491***	0.919***	0.763***	0.950***	0.859***	0.461
Asset purchase announcements							
CBPP1	0.111***	0.058***	0.008**	0.032***	0.017***	0.022***	0.064***
CBPP2	0.053***	0.020***	0.043***	0.025***	0.030***	0.027***	0.041***
CBPP3	-0.051***	-0.049***	-0.030***	-0.038***	-0.028***	-0.029***	-0.056***
SMP	0.029***	0.035***	-0.069***	-0.009	-0.008	-0.046***	-0.059***
Draghi's speech	-0.032***	-0.005**	-0.006	-0.106***	0.002	-0.005	-0.046***
OMT	0.005*	0.029***	0.011***	0.013	-0.015***	0.026***	0.010***
PSPP	-0.025***	-0.029***	-0.020***	-0.027***	-0.017***	-0.015***	-0.036***
CSPP	-0.015	0.030***	-0.066***	-0.048***	-0.074***	-0.070***	0.021***
Sovereign bail-out and default announcements							
Portugal	0.013***	0.005***	0.024***	0.002	-0.004	0.011***	0.007***
Ireland	-0.015***	-0.020***	-0.013***	-0.018***	-0.013***	-0.008***	-0.019***
Greece	-0.016***	-0.014	0.010	0.011***	0.001	-0.016	-0.015***
Greece Default	0.032***	0.029***	0.039***	0.016***	0.052***	0.015***	0.021***
Adjusted R^2	0.684	0.611	0.944	0.849	0.972	0.890	0.635
Number of observations	2283	2283	2082	2283	2253	2275	2283

Dependent variable: Δbond yield, i_{NB_t}.
Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Two-sided test.

Table 2.3: Regressions for ECM—pre-crisis sample

	France	Germany	Ireland	Italy	Portugal	Spain	Euro area
Constant	0.322	-0.034	0.338	0.417	0.496**	-0.050	0.106
Lagged variables							
Bond yield, $(i_{NB_{t-1}})$	-0.076***	-0.060***	-0.050***	-0.071***	-0.073***	-0.066***	-0.064***
Credit risk spread [1], $c_{NB_{t-1}}$	2.36E-05	1.09E-04	2.09E-04**	1.02E-04	2.58E-04***	2.37E-04**	0.022
EONIA, $i_{R_{t-1}}$	0.028***	0.021***	0.014**	0.029***	0.025***	0.025***	0.020***
log(Equity index), $log(Y_{t-1})$	-0.017	0.020	-0.028	-0.025	-0.035	0.022	0.007
Term premium, i^e_{t-1}	0.033***	0.028**	0.020**	0.033***	0.035***	0.022*	0.021**
First differences							
ΔCredit risk spread , Δc_{NB_t}	-0.001	-0.002*	0.001***	0.001**	0.001***	-3.43E-04	0.119
ΔEONIA, Δi_{R_t}	-0.020	-0.001	-0.026	-0.020	-0.009	-0.024	-0.008
Δlog(Equity index), $\Delta log(Y_t)$	0.349****	0.598***	0.041	0.097	0.337**	0.047	0.545***
ΔTerm premium, Δi^e_t	0.503***	0.457***	0.624***	0.543***	0.551***	0.412***	0.461***
Asset purchase announcements							
CBPP1	0.141***	0.067***	0.014**	0.037***	0.017***	0.028***	0.060***
Adjusted R^2	0.633	0.626	0.713	0.606	0.650	0.670	0.564
Number of observations	418	418	400	418	418	418	418

Dependent variable: Δbond yield, i_{NB_t}.
Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Two-sided test.

Table 2.4: Regressions for ECM—intra-crisis sample

	France	Germany	Ireland	Italy	Portugal	Spain	Euro area
Constant	-0.180	-0.140	-0.307	0.062	0.481**	0.020	-0.440**
Lagged variables							
bond yield, ($i_{NB_{t-1}}$)	-0.040***	-0.010	-0.023***	-0.011	-0.021**	-0.012*	-0.017**
Credit risk spread , $c_{NB_{t-1}}$	-3.92E-05	-5.03E-05	-5.10E-06	5.67E-08	-3.37E-05	2.88E-05	0.012**
EONIA, $i_{R_{t-1}}$	-0.014*	-0.003	-0.014*	-0.017**	-0.019***	-0.017**	-0.002
log(Equity index), $log(Y_{t-1})$	0.034	0.018	0.047*	-0.003	-0.046*	0.001	0.060**
Term premium, i^e_{t-1}	0.024***	0.005	0.022***	0.012*	0.022***	0.012*	0.007
First differences							
ΔCredit risk spread , Δc_{NB_t}	1.41E-04	-0.001**	0.001	3.8E-04**	7.98E-05	3.51E-04**	0.357***
ΔEONIA, Δi_{R_t}	-0.024**	-0.010	-0.021*	-0.026**	-0.029**	-0.026**	-0.011
Δlog(Equity index), $\Delta log(Y_t)$	0.264**	0.828***	0.092	0.067	0.120	0.068	0.804***
ΔTerm premium, Δi^e_t	0.847***	0.566***	0.976***	0.913***	0.982***	0.919***	0.514***
Asset purchase announcements							
CBPP2	0.037***	0.020***	0.041***	0.032***	0.036***	0.030***	0.034***
SMP	-0.001	0.019***	-0.031***	-0.017	-0.021	-0.036***	-0.057***
Draghi's speech	-0.026***	-0.020***	-0.006	-0.053***	-0.018***	-0.021***	-0.048***
Sovereign bail-out and default announcements							
Portugal	0.009**	0.006**	0.007	0.005	0.002	0.012***	0.009***
Ireland	-0.007**	-0.017***	4.42E-04	-0.005*	4.53E-04	-0.001	-0.015***
Greece	-0.004	-0.011	0.005	0.009	0.005	-0.005	-0.012***
Greece Default	0.034***	0.027***	0.036***	0.030***	0.044***	0.026***	0.020***
Adjusted R^2	0.847	0.728	0.985	0.944	0.993	0.944	0.795
Number of observations	595	595	595	595	595	595	595

Dependent variable: Δbond yield, i_{NB_t}.
Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Two-sided test.

Table 2.5: Regressions for ECM—post-crisis sample

	France	Germany	Ireland	Italy	Portugal	Spain	Euro area
Constant	-0.002	0.105	0.070	-0.063	0.128	0.003	-0.153*
Lagged variables							
bond yield, ($i_{NB_{t-1}}$)	-0.005*	-0.009**	-0.010***	-0.007*	-0.007	-0.017***	-0.018**
Credit risk spread , $c_{NB_{t-1}}$	6.25E-05	-1.86E-04	7.36E-06	4.26E-05	-1.72E-05	1.28E-04**	0.025***
EONIA, $i_{R_{t-1}}$	0.005	0.013	-0.016*	-0.003	0.002	0.006	0.007
log(Equity index), $log(Y_{t-1})$	3.29E-04	-0.011	-0.007	0.007	-0.014	4.31E-04	0.019*
Term premium, i^e_{t-1}	0.002*	0.002	0.012***	0.007**	0.008**	0.013***	0.007***
First differences							
ΔCredit risk spread , Δc_{NB_t}	2.40E-04	-0.001	0.001*	0.003***	0.001**	0.001	0.396***
ΔEONIA, Δi_{R_t}	-0.010	-0.017	-0.022	-0.009	-0.020	-0.033	-0.022
Δlog(Equity index), $\Delta log(Y_t)$	0.020*	0.242***	-0.236***	-0.234**	-0.248**	-0.052	0.0915
ΔTerm premium, Δi^e_t	0.435***	0.376***	0.651***	0.614***	0.882***	0.886***	0.358***
Asset purchase announcements							
CBPP3	-0.059***	-0.048***	-0.034***	-0.038***	-0.027***	-0.025***	-0.050***
OMT	0.015***	0.048***	0.021***	0.048***	-0.035***	0.038*	0.032***
PSPP	-0.023***	-0.025***	-0.023***	-0.027***	-0.018***	-0.014***	-0.030***
CSPP	-0.018***	0.042***	-0.029***	-0.025***	-0.071***	-0.072***	0.029***
Adjusted R^2	0.442	0.394	0.688	0.791	0.934	0.904	0.528
Number of observations	1087	1087	1087	1087	1087	1087	1087

Dependent variable: Δbond yield, i_{NB_t}.
Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Two-sided test.

2.5.2 Event based regression

After capturing the total effect of asset purchase programs on sovereign bond yields, we examine the two channels via which asset purchases lead to the bond yield reduction within our theoretical model, namely credit risk and term premium channel. As already stated in Section 2.4.4, the credit risk of sovereign bond yields decreased with the ECB acting as a "lender of confidence". Furthermore, the ECB's purchase programs affected the expectations of sovereign bond investors beyond expected short-term rates, which is reflected in a reduced term premium. We test these hypotheses on the transmission channels of asset purchase programs by performing an event based regression in which we capture the effect of asset purchase program announcements on the CDS spreads and the term premium (Szczerbowicz, 2015; Falagiarda and Reitz, 2015).

Regression setup

We apply a standard linear regression and estimate it using OLS with Newey-West standard errors, regressing the change in CDS spreads on its lagged change, the announcement day dummies for asset purchase programs, and control variables:

$$\begin{aligned}\Delta \mathrm{cds}_t = \alpha &+ \beta_1 \Delta \mathrm{cds}_{t-1} + \beta_2 \mathrm{CBPP}_t + \beta_3 \mathrm{OMT}_t + \beta_4 \mathrm{SMP}_t + \beta_5 \mathrm{PSPP}_t + \beta_6 \mathrm{CSPP}_t \\ &+ \beta_7 \text{Draghi's speech}_t + \beta_8 \mathrm{EFSM/ESM}_t + \beta_9 \text{zero lower bound}_t \\ &+ \beta_{10} \Delta \mathrm{VStoxx}_t + \beta_{11} \Delta \mathrm{ted}_t + \beta_{12} \Delta \mathrm{EuroStoxx50}_t + \epsilon_t.\end{aligned}$$

We control for financial turmoil using the volatility stock index VStoxx and for market-wide business climate changes with a stock market index for the EU (Euro Stoxx 50) as well as with information on credit risk in the global economy drawn from the Treasury Bill Eurodollar Difference (TED) spread (see Falagiarda and Reitz, 2015). Additionally, we control for dates of news releases on the European rescue programs EFSM and ESM, and for the dates on which the ECB reached the zero lower bound (see Szczerbowicz, 2015). Taking the same approach as for CDS spreads, we estimate the effects of the ECB's asset purchase programs on the term premium (tp):

$$\begin{aligned}\Delta \mathrm{tp}_t = \alpha &+ \beta_1 \Delta \mathrm{tp}_{t-1} + \beta_2 \mathrm{CBPP} + \beta_3 \mathrm{OMT} + \beta_4 \mathrm{SMP} + \beta_5 \mathrm{PSPP} + \beta_6 \mathrm{CSPP} \\ &+ \beta_7 \text{Draghi's speech} + \beta_8 \mathrm{EFSM/ESM} + \beta_9 \text{zero lower bound} \\ &+ \beta_{10} \Delta \mathrm{VStoxx}_t + \beta_{11} \Delta \mathrm{ted}_t + \beta_{12} \Delta \mathrm{EuroStoxx50}_t + \epsilon_t.\end{aligned}$$

Regression results

With regard to CDS spreads (see Table 2.6), it is apparent that the CBPP1 and CBPP3 were effective in all countries of the sample. The SMP, OMT, Draghi's speech and CSPP were especially powerful in reducing the CDS spreads of the countries that were most severely hit by the European sovereign debt crisis (Spain, Portugal, Italy, and Ireland). The results for the SMP and OMT correspond to the results of Szczerbowicz (2015), who analyzes the spread of Eurozone sovereign bonds compared to that of German sovereign bonds. Nevertheless, we find also a negative significant effect of Draghi's speech on CDS spreads of Germany and France. An increase in CDS spreads was triggered by the CBPP2 for Spain, Portugal, Italy, Ireland, and France, and by the PSPP for Ireland, Germany, and France. Concerning the term premia, for the SMP we detect the expected decrease (see Table 2.7) for Spain, Portugal, Ireland, Germany, and the Euro area. For the OMT we observe a significant negative effect on Spain, Portugal, and Italy, whereas we identify a significant positive effect for Ireland, Germany, and France. Draghi's speech contributed to a decline in term premia of each member state with the exception of Germany. With respect to the PSPP, we report a significant negative impact on the term premia of Portugal and Italy. By contrast, with the introduction of the PSPP, the effect on the term premium for France was positive. The results for the CSPP indicate a positive impact on term premia for Portugal, Spain, Ireland, Germany, and France, whereas for Italy a negative effect was observed. Taken together, we document a decrease in CDS spreads as a result of the majority of asset purchase program announcements. The effects on term premia across the Euro area, however, are not as distinct, which can be explained by the fact that the effect of asset purchases on term premia is likely to diminish over the course of each announcement day and we cannot capture this short-run effect without intra-day data.

Table 2.6: Event based regression for credit risk spreads

	France	Germany	Ireland	Italy	Portugal	Spain	Euro area
Constant	0.002	0.004	0.043	0.063	0.116	0.046	3.08E-04
ΔCredit risk spread , $\Delta c_{NB_{t-1}}$	0.477***	0.477***	0.477*	0.442***	0.465***	0.441***	0.442***
Asset purchase announcements							
CBPP1	-3.513***	-4.090***	-10.898***	-7.542***	-8.485***	-6.727***	-0.054***
CBPP2	2.168***	-3.555***	6.429***	5.316***	33.109***	4.540***	-0.036***
CBPP3	0.220	0.188***	1.003**	-0.545	-3.300***	-1.020***	0.006***
SMP	0.446	-0.662	-14.647**	-8.447	-70.503***	-8.376	-0.091***
Draghi's speech	-0.892***	-0.856***	-4.325***	-7.259***	-8.521***	-7.959***	-0.040***
OMT	-1.787***	-0.117	-14.006***	-34.630***	-13.403***	-38.343***	-0.068***
PSPP	0.930***	0.919***	1.981***	-2.140***	-4.887***	0.023	-0.005***
CSPP	-0.893***	-0.121**	-2.187***	-6.318***	-3.174***	-5.300***	-0.041***
Control variables							
EFSM/ESM	-0.630	0.201	-7.412	-10.102*	-12.482***	-12.366**	-0.049
Zero lower bound	3.911***	-0.872***	-5.459***	7.629***	16.486***	9.394***	0.016***
ΔVSTOXX_t	0.096	0.064*	-0.013	0.132	0.347	0.219	0.001*
ΔTED_t	0.001	0.003	0.143**	-3.19E-05	0.008	-0.010	1.69E-05
$\Delta \text{EuroStoxx}_t$	-0.007***	-0.003**	-0.033***	-0.036***	-0.100***	-0.030***	-1.08E-04***
Adjusted R^2	0.528	0.560	0.469	0.572	0.522	0.577	0.541
Number of observations	2281	2281	2080	2281	2251	2281	2281

Dependent variable: ΔCredit risk spread , Δc_{NB_t}.
Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Two-sided test.

Table 2.7: Event based regression for term premia

	France	Germany	Ireland	Italy	Portugal	Spain	Euro area
Constant	-0.001	-0.001	9.35E-05	4.57E-05	0.001	-2.28E-04	-2.38E-04
ΔTerm premium, Δi^e_{t-1}	0.492***	0.468***	0.491***	0.489***	0.495***	0.494***	0.494***
Asset purchase announcements							
CBPP1	0.221***	0.128***	0.033***	0.032***	0.030***	0.102***	0.075***
CBPP2	0.058***	0.073***	0.021***	-0.034***	0.147***	0.035***	0.036***
CBPP3	-0.017***	-0.023***	-0.008	-0.022***	-0.004	0.032***	-0.015***
SMP	0.003	-0.054***	-0.346**	0.014	-0.299*	-0.143*	-0.127***
Draghi's speech	-0.046***	0.005	-0.124***	-0.113***	-0.027**	-0.003*	-0.022***
OMT	0.042***	0.062***	0.044***	-0.196***	-0.524***	-0.134***	-0.005
PSPP	0.003	-0.008	-4.67E-04	-0.033***	-0.030***	-0.003	-0.011***
CSPP	0.041***	0.073***	0.037***	-0.010**	0.016***	0.089**	0.027***
Control variables							
EFSM/ESM	-0.005	-0.007	-0.187	-0.122	-0.136	-0.094	-0.037
Zero lower bound	-0.158***	-0.088***	-0.055***	-0.041***	-0.147***	0.151***	-0.083***
ΔVSTOXX$_t$	-0.003*	-0.001	0.002	0.002	0.005	-0.001	6.61E-04
ΔTED$_t$	3.23E-04	2.63E-04	0.001	1.67E-04	0.001	1.25E-04	2.83E-04
ΔEuroStoxx$_t$	3.09E-05	2.26E-04***	6.29E-05	-8.95E-05	1.17E-04	-6.61E-05	1.19E-04***
Adjusted R^2	0.520	0.540	0.483	0.528	0.483	0.497	0.504
Number of observations	2281	2281	2281	2281	2281	2281	2281

Dependent variable: ΔTerm premium, Δi^e_t.
Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Two-sided test.

2.6 Conclusion

In this paper we present a tractable theoretical model, which is suited to evaluate the effects of unconventional monetary policy measures by the ECB on the financial system before we derive and empirically test hypotheses on the effects and transmission process of asset purchase programs on the bond market for a set of European sovereign bonds.

The proposed model is based on the extension of Bofinger et al. (2017) to the price-theoretic model of the money supply developed by Bofinger and Schächter (1995) and consists of two interconnected markets, the bank credit market and the bond market. Once money is created by banks in the process of credit creation on the bank credit market, non-bank holders of money may redistribute the money by purchasing bonds, i.e., implicitly granting loans to banks and non-banks, on the bond market. With the central bank taking action on the bond market in the course of an asset purchase program, it is able to exert an expansionary stimulus on credit supply in the bond market through decreasing credit risk and interest rate expectations. Due to forward looking market participants, the effects of asset purchase programs, in particular QE, are anticipated and thus take place upon the mere announcement of asset purchase programs. From our theoretical model, we derive three hypotheses that we test empirically: (i) asset purchase programs are effective in lowering bond yields through (ii) lowering (perceived) credit risk and (iii) impacting on term-premia in the bond market.

In the second part of our paper we empirically test the hypotheses derived from the model by estimating country-individual effects of the ECB's asset purchase programs on 5-year sovereign bond yields in an error correction model and credit risk spreads and term premia in event-based regressions for France, Germany, Ireland, Italy, Portugal, Spain and the European benchmark bond. Despite showing substantial heterogeneity in coefficients across countries and programs, our results from the error correction model provide (i) tentative evidence of a portfolio rebalancing effect across asset classes for non-sovereign bond asset purchase programs (CBPP, CBPP2, CBPP3 and CSPP) and (ii) underline the effectiveness of asset purchase programs on targeted sovereign bond markets. In particular for the announcement of the large-scale asset purchases under the PSPP, representing the official launch of QE in the Eurozone, we find significant yield-diminishing effects across all observed countries.

The results from our event based regressions, further suggest a decrease in credit risk for most announcements on central bank asset purchases, leading us to the conclusion that by carrying out these programs, the ECB effectively acted as "lender of confidence". Using the same methodology to analyze the impact on term premia, we are unable to identify a clear-cut effect neither for the individual programs nor for the respective countries. This result aligns with the notion of mixed signals on future interest rates resulting from asset purchase programs. Overall, the empirical assessment supports the results of our model and legitimates its use for the understanding of the effects of asset purchase on bond market interest rates and sovereign bond yields, in particular.

Taking this study as a reference, future research might successfully be directed at theoretically and empirically analyzing the transmission channels and the effects our model implies on the interest rate for bank credit, as a result of the decrease of the interest rate on the non-banking market. That is, a decrease in bank credit interest rates caused by lower bond market interest rates, which we have shown to be an effect of central bank asset purchases in this paper. In addition, further extensions of the model could include an equity market to endogenously determine the interest rate on equity, which again determines the bank credit interest rate as a key factor of banks' credit supply.

2.7 Appendix to Chapter 2

Table 2.8: Data sources

Variable	Source
Sovereign bond yields	Datastream
CDS spreads	Datastream
Equity indexes	Datastream
EONIA	Datastream
VSTOXX	Datastream
TED spread	Datastream
Forward rates	Bloomberg

Table 2.9: Unit root tests for EONIA

Variable	Test	P-val.	Test-stat.	Critical-val.: 5%	Decision
EONIA	ADF (w. trend)	0.443			not stat.
	ADF (wo. trend)	0.052			not stat.
	Phillips-Perron (w. trend)	0.395			not stat.
	Phillips-Perron (wo. trend)	0.111			not stat.
	KPSS (w. trend)		0.643	0.146	not stat.
	KPSS (wo. trend)		3.155	0.463	not stat.

ADF = Augmented Dickey Fuller test.
KPSS = Kwiatkowski-Phillips-Schmidt-Shin test.
w.= with; wo.=without.

Table 2.10: Unit root tests for sovereign bond yields of each country

Country	Test	P-val.	Test-stat.	Critical-val.: 5%	Decision
Spain	ADF (w. trend)	0.758			not stat.
	ADF (wo. trend)	0.836			not stat.
	Phillips-Perron (w. trend)	0.690			not stat.
	Phillips-Perron (wo. trend)	0.787			not stat.
	KPSS (w. trend)		1.055	0.146	not stat.
	KPSS (wo. trend)		3.202	0.463	not stat.
France	ADF (w. trend)	0.080			not stat.
	ADF (wo. trend)	0.606			not stat.
	Phillips-Perron (w. trend)	0.070			not stat.
	Phillips-Perron (wo. trend)	0.572			not stat.
	KPSS (w. trend)		0.187	0.146	not stat.
	KPSS (wo. trend)		5.249	0.463	not stat.
Germany	ADF (w. trend)	0.191			not stat.
	ADF (wo. trend)	0.368			not stat.
	Phillips-Perron (w. trend)	0.253			not stat.
	Phillips-Perron (wo. trend)	0.387			not stat.
	KPSS (w. trend)		0.479	0.146	not stat.
	KPSS (wo. trend)		5.100	0.463	not stat.
Ireland	ADF (w. trend)	0.488			not stat.
	ADF (wo. trend)	0.525			not stat.
	Phillips-Perron (w. trend)	0.630			not stat.
	Phillips-Perron (wo. trend)	0.634			not stat.
	KPSS (w. trend)		0.807	0.146	not stat.
	KPSS (wo. trend)		2.582	0.463	not stat.
Italy	ADF (w. trend)	0.663			not stat.
	ADF (wo. trend)	0.766			not stat.
	Phillips-Perron (wi. trend)	0.600			not stat.
	Phillips-Perron (wo. trend)	0.718			not stat.
	KPSS (w. trend)		0.919	0.146	not stat.
	KPSS (wo. trend)		3.366	0.463	not stat.
Portugal	ADF (w. trend)	0.789			not stat.
	ADF (wo. trend)	0.580			not stat.
	Phillips-Perron (w. trend)	0.822			not stat.
	Phillips-Perron (wo. trend)	0.620			not stat.
	KPSS (w. trend)		0.914	0.146	not stat.
	KPSS (wo. trend)		1.255	0.463	not stat.
Euro area	ADF (w. trend)	0.273			not stat.
	ADF (wo. trend)	0.877			not stat.
	Phillips-Perron (w. trend)	0.297			not stat.
	Phillips-Perron (wo. trend)	0.859			not stat.
	KPSS (w. trend)		0.644	0.146	not stat.
	KPSS (wo. trend)		5.154	0.463	not stat.

ADF = Augmented Dickey Fuller test.
KPSS = Kwiatkowski-Phillips-Schmidt-Shin test.
w.= with; wo.=without.

Table 2.11: Unit root tests for credit risk spreads of each country

Country	Test	P-val.	Test-stat.	Critical-val.: 5%	Decision
Spain	ADF (w. trend)	0.658			not stat.
	ADF (wo. trend)	0.332			not stat.
	Phillips-Perron (w. trend)	0.665			not stat.
	Phillips-Perron (wo. trend)	0.365			not stat.
	KPSS (w. trend)		1.121	0.146	not stat.
	KPSS (wo. trend)		1.145	0.463	not stat.
France	ADF (w. trend)	0.493			not stat.
	ADF (wo. trend)	0.269			not stat.
	Phillips-Perron (w. trend)	0.431			not stat.
	Phillips-Perron (wo. trend)	0.218			not stat.
	KPSS (w. trend)		0.861	0.146	not stat.
	KPSS (wo. trend)		1.021	0.463	not stat.
Germany	ADF (w. trend)	0.094			not stat.
	ADF (wo. trend)	0.102			not stat.
	Phillips-Perron (w. trend)	0.102			not stat.
	Phillips-Perron (wo. trend)	0.113			not stat.
	KPSS (w. trend)		0.506	0.146	not stat.
	KPSS (wo. trend)		0.810	0.463	not stat.
Ireland	ADF (w. trend)	0.512			not stat.
	ADF (wo. trend)	0.522			not stat.
	Phillips-Perron (w. trend)	0.617			not stat.
	Phillips-Perron (wo. trend)	0.598			not stat.
	KPSS (w. trend)		0.776	0.146	not stat.
	KPSS (wo. trend)		2.149	0.463	not stat.
Italy	ADF (w. trend)	0.508			not stat.
	ADF (wo. trend)	0.207			not stat.
	Phillips-Perron (w. trend)	0.573			not stat.
	Phillips-Perron (wo. trend)	0.249			not stat.
	KPSS (w. trend)		0.923	0.146	not stat.
	KPSS (wo. trend)		0.927	0.463	not stat.
Portugal	ADF (w. trend)	0.795			not stat.
	ADF (wo. trend)	0.471			not stat.
	Phillips-Perron (w. trend)	0.777			not stat.
	Phillips-Perron (wo. trend)	0.448			not stat.
	KPSS (w. trend)		0.948	0.146	not stat.
	KPSS (wo. trend)		0.945	0.463	not stat.
Euro area	ADF (w. trend)	0.498			not stat.
	ADF (wo. trend)	0.199			not stat.
	Phillips-Perron (w. trend)	0.674			not stat.
	Phillips-Perron (wo. trend)	0.332			not stat.
	KPSS (w. trend)		1.180	0.146	not stat.
	KPSS (wo. trend)		1.182	0.463	not stat.

ADF = Augmented Dickey Fuller test.
KPSS = Kwiatkowski-Phillips-Schmidt-Shin test.
w.= with; wo.=without.

Table 2.12: Unit root tests for term premia of each country

Country	Test	P-val.	Test-stat.	Critical-val.: 5%	Decision
Spain	ADF (w. trend)	0.723			not stat.
	ADF (wo. trend)	0.440			not stat.
	Phillips-Perron (w. trend)	0.679			not stat.
	Phillips-Perron (wo. trend)	0.387			not stat.
	KPSS (w. trend)		1.232	0.146	not stat.
	KPSS (wo. trend)		1.253	0.463	not stat.
France	ADF (w. trend)	0.245			not stat.
	ADF (wo. trend)	0.140			not stat.
	Phillips-Perron (w. trend)	0.204			not stat.
	Phillips-Perron (wo. trend)	0.092			not stat.
	KPSS (w. trend)		0.152	0.146	not stat.
	KPSS (wo. trend)		0.910	0.463	not stat.
Germany	ADF (w. trend)	0.412			not stat.
	ADF (wo. trend)	0.203			not stat.
	Phillips-Perron (w. trend)	0.400			not stat.
	Phillips-Perron (wo. trend)	0.196			not stat.
	KPSS (w. trend)		0.199	0.146	not stat.
	KPSS (wo. trend)		0.827	0.463	not stat.
Ireland	ADF (w. trend)	0.535			not stat.
	ADF (wo. trend)	0.346			not stat.
	Phillips-Perron (w. trend)	0.644			not stat.
	Phillips-Perron (wo. trend)	0.449			not stat.
	KPSS (w. trend)		0.819	0.146	not stat.
	KPSS (wo. trend)		1.252	0.463	not stat.
Italy	ADF (w. trend)	0.662			not stat.
	ADF (wo. trend)	0.390			not stat.
	Phillips-Perron (w. trend)	0.614			not stat.
	Phillips-Perron (wo. trend)	0.341			not stat.
	KPSS (w. trend)		1.022	0.146	not stat.
	KPSS (wo. trend)		1.069	0.463	not stat.
Portugal	ADF (w. trend)	0.865			not stat.
	ADF (wo. trend)	0.474			not stat.
	Phillips-Perron (w. trend)	0.816			not stat.
	Phillips-Perron (wo. trend)	0.517			not stat.
	KPSS (w. trend)		0.940	0.146	not stat.
	KPSS (wo. trend)		0.943	0.463	not stat.
Euro area	ADF (w. trend)	0.452			not stat.
	ADF (wo. trend)	0.333			not stat.
	Phillips-Perron (w. trend)	0.408			not stat.
	Phillips-Perron (wo. trend)	0.294			not stat.
	KPSS (w. trend)		0.649	0.146	not stat.
	KPSS (wo. trend)		1.180	0.463	not stat.

ADF = Augmented Dickey Fuller test.
KPSS = Kwiatkowski-Phillips-Schmidt-Shin test.
w.= with; wo.=without.

Table 2.13: Unit root tests for logarithm of equity indexes of each country

Country	Test	P-val.	Test-stat.	Critical-val.: 5%	Decision
Spain	ADF (w. trend)	0.092			not stat.
	ADF (wo. trend)	0.017			stat.
	Phillips-Perron (w. trend)	0.141			not stat.
	Phillips-Perron (wo. trend)	0.028			stat.
	KPSS (w. trend)		0.531	0.146	not stat.
	KPSS (wo. trend)		0,.876	0.463	not stat.
France	ADF (w. trend)	0.016			stat.
	ADF (wo. trend)	0.041			stat.
	Phillips-Perron (w. trend)	0.022			stat.
	Phillips-Perron (wo. trend)	0.056			not stat.
	KPSS (w. trend)		0.578	0.146	not stat.
	KPSS (wo. trend)		1.817	0.463	not stat.
Germany	ADF (w. trend)	0.009			stat.
	ADF (wo. trend)	0.747			not stat.
	Phillips-Perron (w. trend)	0.010			stat.
	Phillips-Perron (wo. trend)	0.778			not stat.
	KPSS (w. trend)		0.323	0.146	not stat.
	KPSS (wo. trend)		4.835	0,463	not stat.
Ireland	ADF (w.trend)	0.041			stat.
	ADF (wo. trend)	0.565			not stat.
	Phillips-Perron (w. trend)	0.033			stat.
	Phillips-Perron (wo. trend)	0.588			not stat.
	KPSS (w. trend)		0.858	0146	not stat.
	KPSS (wo. trend)		3.008	0.463	not stat.
Italy	ADF (w. trend)	0.105			not stat.
	ADF (wo. trend)	0.018			stat.
	Phillips-Perron (w. trend)	0.102			not stat.
	Phillips-Perron (wo. trend)	0.017			stat.
	KPSS (w. trend)		0.727	0.146	not stat.
	KPSS (wo. trend)		0.761	0.463	not stat.
Portugal	ADF (w. trend)	0.057			not stat.
	ADF (wo. trend)	0.046			stat.
	Phillips-Perron (w. trend)	0.063			not stat.
	Phillips-Perron (wo. trend)	0.044			stat.
	KPSS (w. trend)		0.247	0.146	not stat.
	KPSS (wo. trend)		3.379	0.463	not stat.
Euro area	ADF (w. trend)	0.021			stat.
	ADF (wo. trend)	0.013			stat.
	Phillips-Perron (w. trend)	0.029			stat.
	Phillips-Perron (wo. trend)	0.020			stat.
	KPSS (w. trend)		0.590	0.146	not stat.
	KPSS (wo. trend)		1.048	0.463	not stat.

ADF = Augmented Dickey Fuller test.
KPSS = Kwiatkowski-Phillips-Schmidt-Shin test.
w.= with; wo.=without.

Chapter 3

Intertwined, After All? A German Perspective on Unconventional Monetary Policy and Public Debt Management

3.1 Introduction

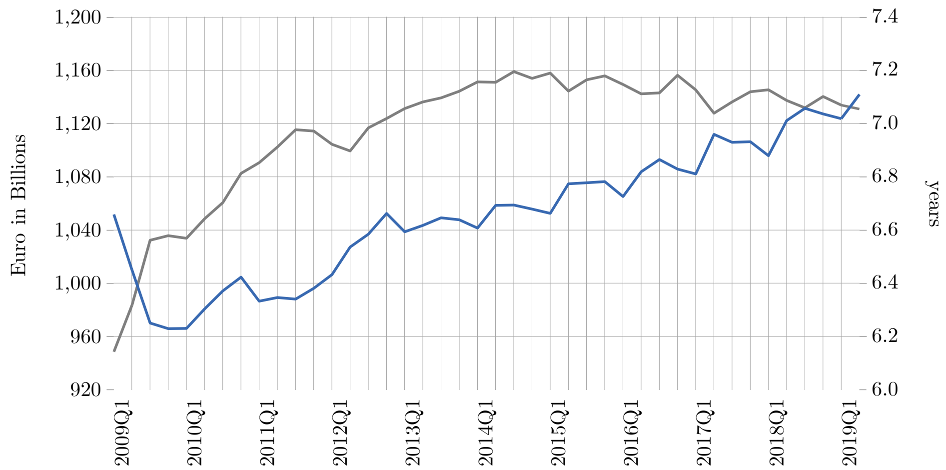
The past decade has been a turbulent time for monetary policy makers across the globe. The arrival at the zero lower bound of short-term interest rates and the subsequent rise of unconventional monetary policy measures have directed unprecedented attention toward major central banks, including the European Central Bank (ECB). Under the Public Sector Purchase Programme (PSPP)—the main component of the Expanded Asset Purchase Programme (APP)—the ECB has purchased medium and long-term national and supranational debt instruments in large quantities over a period of more than four years. In addition, the recent reactivation and intensification of monthly net asset purchases under the PSPP have brought quantitative easing (QE) closer to becoming the new normal in the Eurozone.[1]

Among the ECB's accumulated PSPP holdings, German sovereign debt instruments represent by far the largest component, amounting to approximately 25%. At the same time, the ECB has become the prime investor in German sovereign debt, with PSPP holdings summing up to more than 30% of the total outstanding volume of German marketable sovereign debt instruments. In parallel to these intensified large-scale asset purchases on the secondary market for German sovereign debt, the maturity structure of outstanding volumes and gross issuance of German public debt instruments have changed considerably: since the ECB's first indications of a large-scale asset purchase program in 2014, the weighted average maturity (WAM) of the *outstanding stock* of German marketable public debt has increased by 16% from 6.6 years to 7.1 years, despite its volume having decreased by 3% at the same time. Accordingly, the *flow of gross issuance* has experienced an even stronger maturity extension of approximately 34%, while the average issuance volume decreased by 29% from 2014 to 2019, compared to its average structure and size over the years 2009 to 2013 (see Figure 3.1).

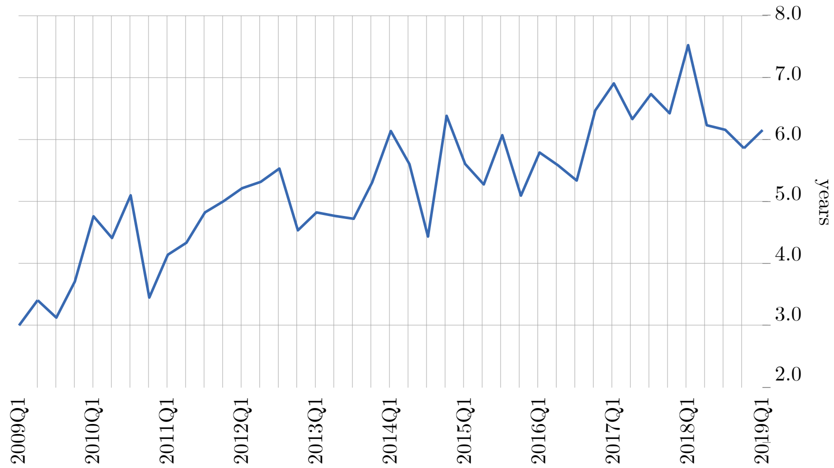
The coincidence of central bank actions on sovereign bond markets with active maturity extensions in gross issuance brought about by the German debt management office is suggestive to the existence of a nexus between monetary policy and public debt management

[1]Since the start of the PSPP in March 2015 until its end in January 2020, the ECB has increased its combined asset holdings from €55 to €2,579 billion. As of January 2020, purchases within the PSPP constitute approximately 60% of monthly net purchases and around 75% of total central bank asset holdings accumulated in the course of the APP. Prior to the temporary suspension of net asset purchases between January and November 2019, monthly volumes of net purchases have ranged between €15 to €80 billion per month. Since December 2019, net purchases are resumed at a regular pace of €20 billion per month.

Figure 3.1: Development of German public debt volume and maturity structure since 2009



(a) Volume and WAM of German public debt outstanding



(b) WAM of German public debt gross issuance

Notes: Panel (a) depicts volume (grey line) in Euro Billions and weighted average maturity (WAM) (blue line) in years of the outstanding stock of German marketable central government debt over the period of 2009Q1 to 2019Q1 on a quarterly basis. Panel (b) displays the WAM in years of the quarterly flow of gross issuance adding towards the stock of German marketable central government debt for the same time period.

(PDM). Institutional specifics of PDM in combination with theoretically and empirically documented effects of monetary policy on the yield curve nourish the idea of a direct link. First, from an institutional perspective, actively managing the maturity structure is a crucial instrument of PDM in fulfilling its mandate.[2] In this vein, the shape of the yield curve

[2] According to the definition by the International Monetary Fund and World Bank (2014), PDM is responsible for choosing and implementing the maturity, interest rate, and currency *composition* of government debt along its goals of minimizing cost and maintaining a prudent level of risk. By contrast, deciding upon the *level* of government debt to be issued is in the responsibility of fiscal authorities and thus outside the scope of PDM.

is of first-order importance for PDM: it reflects the *costs*—i.e., interest to be paid across maturities—and to some degree also the *risks*—i.e., interest rate risk and rollover risk—that PDM faces in deciding upon the optimal maturity structure of public debt. Second, a vast theoretical and empirical literature has shown that both conventional and unconventional monetary policy actions considerably affect the yield curve; thus, it is conceivable that monetary policy actions may also have repercussions on PDM decisions. While expansionary conventional measures, i.e., policy rate cuts, predominantly reduce short-term rates and only imperfectly pass through to long-term rates, recent studies unanimously find overwhelming support for large-scale asset purchases directly lowering the longer end of the yield curve, while leaving the short end largely unaffected (Altavilla et al., 2019; Eser et al., 2019). In this context, there is furthermore a broad consensus that the strongest yield effects of purchase programs occur upon their announcements due to so called stock effects, i.e., changes in the projected stock of assets on the central bank balance sheet.[3]

Against this backdrop of well-documented effects of monetary policy on bond yields and the practical relevance of these yields for PDM, the lack of empirical studies on the interaction of both macroeconomic policies is startling. Some first-path guidance is given by Greenwood et al. (2016a), who study potentially counteracting effects of simultaneous maturity extensions by the Treasury and QE programs by the Federal Reserve between 2008 and 2014. Their findings align with documented positive long term yield effects of Treasury-induced maturity extensions by, i.a., Swanson (2011) for the 1962–1964 Operation Twist and by Greenwood and Vayanos (2010) for the 2000–2002 buybacks. Importantly, these contributions analyze effects induced by PDM as isolated phenomena, i.e., as an exogenous shift in the conduct of PDM. In light of the theoretical effects of monetary policy on the yield curve, however, the explored maturity extensions may—at least to some extent—represent a direct reaction to the stimulus provided by monetary policy. Yet, not many studies deal with a possible *causal dependency* of maturity decisions by PDM on monetary policy. As one of the few exceptions, Blommestein and Turner (2012) review the relationship between monetary policy and PDM and find tentative evidence on a direct link between expansive *conventional* monetary policy and a shortening of the maturity structure of outstanding debt for *U.S. data*. With regard to the Eurozone, Hoogduin et al. (2011) find similar evidence for conventional monetary policy actions. Still, the question whether

[3]The documented size of the long-term yield effect of QE announcements, however, varies substantially across papers (see, e.g., Altavilla et al., 2015; Andrade et al., 2016; Koijen et al., 2017; De Santis and Holm-Hadulla, 2017; Urbschat and Watzka, forthcoming, as well as Chapter 2 of this dissertation).

unconventional monetary policy actions by the ECB may have caused adjustments in the maturity of primary market issuance of European sovereign bonds, remains unanswered.[4]

This paper aims to fill this void by—to the best of my knowledge—being the first to provide causal estimates on the sign and size of unplanned maturity shifts by the German public debt issuer in direct response to unconventional monetary policy actions involving sovereign bond markets by the ECB.

In a first step, I derive narrative evidence on a potential nexus between QE announcements and PDM decisions in an event study inspired analysis. In this vein, I measure sudden shifts in the maturity structure of monthly gross issuance in public auctions around ECB announcements of unconventional policy actions. To identify relevant announcements, I qualitatively evaluate data on press releases and press conferences from the ECB website over the sample period 2014M1 to 2019M3 with respect to their sovereign bond market related news value.[5] Following this qualitative approach, I construct two alternative event series of announcements that directly or indirectly affect sovereign bond markets. First, I identify a narrow binary event series (SBP_t), which solely includes announcements on sovereign bond purchase programs by the ECB. Second, I build a wider binary event series (SBN_t), which nests the SBP_t series, but further comprises news on changes in accepted sovereign bond collateral by the ECB. A drawback of this binary approach is that it merely captures the occurrence of an announcement while ignoring the information given on direction—extension or tapering—and relative size of the unconventional measure. To account for these dimensions, I, third, build a quantified event series (QSB_t) that allows for different sign and size of events according to the way they were perceived by financial markets. For this purpose, I weigh the binary announcement indicators (event dummies) with the immediate response in a financial market variable that is observed within a narrow time window around the respective announcement on a Governing Council monetary meeting day. Specifically, I use the change in the 10-year Bund rate within the "monetary policy window" from the novel Euro Area Monetary Policy Event-Study Database (EA-MPD) developed by Altavilla et al. (2019).

[4] Related discussions on primary market effects of QE in the Eurozone are mostly led by Euro-critics and typically center on the question whether the ECB incites an overall increase of public debt by potentially exceeding its mandate in purchasing sovereign bonds on secondary markets (see, e.g., Heinemann, 2017; Sinn, 2018; Afflatet, 2019). In this context, the effects on the *structure* of public debt are mostly overlooked.

[5] Other recent studies on unconventional monetary policy events identify announcement events by the degree of news coverage they trigger (see, e.g., Cieslak and Schrimpf, 2019; Urbschat and Watzka, forthcoming). Since this strategy includes announcements beyond sovereign bond market actions, such as bank lending programs, it is not appropriate for the purpose of this study.

To capture active PDM decisions, I focus on the maturity structure of gross issuance, i.e., on the flows toward the stock of debt. With buy-backs and early redemption by the issuer being legally forbidden in Germany (Bundesministerium der Finanzen, 2012a), active changes by German PDM are limited to maturity adjustments in gross issuance. Yet, a potential complication in the analysis of the maturity structure of gross issuance is that it may be driven by longer-term strategical factors rather than reflecting short-term reactions to, e.g., monetary policy announcements. To overcome this challenge, I use auction and planning data provided by the "Bundesrepublik Deutschland—Finanzagentur GmbH" (German Finance Agency) to construct a novel time series of *unplanned adjustments in maturity structure* of monthly primary market bond emissions. These changes in the maturity structure are persuasively not reflective of long-term PDM strategies; they occur through deviations in the actual issuance volumes across different maturity classes from the respective planned volumes, and thus result in unplanned WAM deviations in monthly gross issuance. As a consequence of the uniqueness of these information from the German Finance Agency, I will exclusively focus on German public debt data in what follows.[6]

A joint descriptive analysis of the data unearths the following inference. An inspection of the timing and the size of unplanned maturity adjustments in gross issuance reveals a regular but delayed lengthening of the maturity structure by PDM following announcements of central bank actions in sovereign bond markets.[7] Overall, the dynamic correlations documented in this narrative analysis are suggestive to a direct link between QE and sudden maturity extensions in German sovereign debt gross issuance. However, such correlations do not necessarily imply a causal relationship to be statistically present in the data.

Therefore, as a second contribution of this paper, I use time-series regression techniques to estimate the causal effect of unconventional monetary policy announcements in monthly data. Within my regression setups I control for heteroscedasticity and autocorrelation. Importantly, the applied proxies for QE announcements—in particular the measure of surprise movements in asset prices measured within narrow announcement windows—can be considered as convincingly exogenous regressors. By design and due to the abovementioned institutional

[6]In contrast to other debt management offices in the Eurozone, which announce details beyond the auction date on planned issuance merely days ahead, the German Finance Agency provides a detailed annual plan that precisely describes the timing, volume, and maturities of scheduled primary market issuance. Deviations from this annual plan are rare and only occur after prior announcement in one of the quarterly updates, published in March, June, and September.

[7]Note that such an apparently protracted maturity extension is fully in line with the transparency convention of the German Finance Agency; that is, the prerequisite of a prior announcement of potential plan deviations by PDM in one of the quarterly press releases.

characteristics, these proxies are orthogonal to PDM decisions; that is, endogeneity concerns do not arise. In addition, these tightly measured asset price fluctuations are plausibly not induced by news about macroeconomic or financial conditions and should instead be considered as primarily being explained by QE announcements.[8] Nevertheless, I test the effect of potentially confounding variables, but do not find evidence of any other drivers of unanticipated deviations in WAM. To track *dynamic* causal effects of QE announcements on unplanned deviations in the maturity structure over time, I apply the local projections framework of Jordà (2005).

Corroborating the findings from the first-stage narrative analysis, the contemporaneous coefficients in my estimations are insignificant, i.e., unconventional monetary policy does not impact on maturity decisions of PDM within the same month. However, this is not the case for dynamic causal effects: local projections reveal a delayed and statistically significant effect of QE announcements on unplanned adjustments in the maturity of gross issuance of German sovereign debt. Following expansionary QE announcements, the maturity structure indeed increases. The strongest unplanned maturity extension occurs approximately 8 months after the unconventional monetary policy announcement and ranges between 98 and 151 days, depending on the empirical specification. These results are robust to various sensitivity checks.[9] Overall, my empirical findings provide striking evidence in favor of an extension in the maturity of primary sovereign bond market issuance by German PDM *in direct response* to unconventional monetary policy announcements by the ECB.

Taken together, this paper contributes to the scarce empirical literature on the overlap between PDM and monetary policy in several dimensions. It complements existing studies, which are typically reduced form, with causal effects of unconventional monetary policy on the maturity structure of newly issued public debt in a Eurozone country. The careful selection of sovereign bond market relevant events and their quantification based on high-frequency asset price data, provide several improvements in comparison to the literature: the proxy series (i) explicitly focus on central bank sovereign bond market actions, (ii) cover an extended time sample until March 2019, and (iii) account for different quality and weight

[8]By contrast, the empirical estimations on the PDM effect of conventional monetary policy by Hoogduin et al. (2011) and Blommestein and Turner (2012) fail to take the reverse causality effect of maturity structure on the interest rate (see Swanson, 2011; Greenwood and Vayanos, 2010; Greenwood et al., 2016a) into account.

[9]For instance, I employ alternative unconventional monetary policy measures and perform further modifications to the model specifications. Notably, applying the model to expansive *conventional* monetary policy events in earlier data, I demonstrate the external validity of my empirical approach. In line with theoretical predictions, the announcements of conventional short-term interest reductions induce an unplanned shortening of the maturity structure.

of announcements. Furthermore, the newly constructed time series of plan deviations in the conduct of German PDM poses an innovation that may serve as a starting point for future research. Conditional on the provision of similar planning data by the respective debt management offices, a corresponding empirical approach may be used to extend the analysis to further member states of the Eurozone. Ultimately, the quantitative evidence for a maturity extension conditional on central bank actions on sovereign bond markets from this study could enhance term-structure models that serve to inform monetary policy makers. By accounting for the documented endogenous reactions of PDM, an unintended weakening in the transmission and the need for readjustments of QE could be mitigated.

The remainder of this paper is structured as follows: Section 2 reviews the theoretical foundation and existing empirical work on the respective financial market effects and potential interaction of monetary policy and public debt management. Section 3 describes the construction of the data series and provides a narrative analysis of the data, before Section 4 quantifies the causal influence of QE on maturity shifts in gross issuance of public debt. Section 5 discusses extensions and sensitivity checks to the dynamic model, and Section 6 concludes.

3.2 Theoretical background

In understanding the relationship between public debt management and monetary policy, i.e., their respective financial market implications and potentially resulting feedback effects, the yield curve is the critical common touchstone: it graphically represents the interest rate of securities with comparable credit quality—e.g., bills and bonds of similar issuers—across the spectrum of maturities. Furthermore, the yield curve serves both as an indicator as well as a transmission element of monetary policy and public debt management. In the following, I thus set the base for the subsequent empirical discussion by initially touching upon the most important theories on the term structure of interest rate, before highlighting the individual and potentially overlapping influence of both policies on the yield curve.

3.2.1 Theories on the term structure of interest rates

When it comes to the theoretical explanations of the typically upward-sloping yield curve, various hypotheses based on partially conflicting assumptions and predictions have been developed on the so-called term structure of interest rates. Since a detailed discussion is beyond the scope of this paper, I follow Cox et al. (1985) and give a condensed overview over the most prominent theories.[10]

The **expectations theory** is based on the ideas of Fisher (1930) and can be viewed as the cornerstone of theories on the term structure of interest rates. It is built on the assumption of *complete substitutability* between assets of different maturities. For the no-arbitrage condition to be met under this premise, the expected return of investing in a long-term bond at an annual interest rate of $i_{long,t}$ at time t for n periods needs to be equal to a repeated investment in a series of short-term bonds at the current short-term rate $i_{short,t+0}$ in period $i = 0$ and at expected short-term rates $i^e_{short,t+i}$ for the subsequent periods $i = 1, ..., n$. Consequently, the expected path of short-term rates plays a crucial role for the pricing of bonds and the associated long-term interest rate. This connection can be expressed as follows:

$$(1 + i_{long,t})^n = \prod_{i=0}^{n}(1 + i^{(e)}_{short,t+i}). \tag{3.1}$$

Still, the pure expectations theory fails to explain why long-term interest rates usually exceed the geometrical mean of current and expected short-term rates for the typically upward-sloping term structure of interest rates. Hence, additional theories originated over time to offer alternative explanations for the existence of such a term premium, based on the assumption of *limited asset substitutability.*

The **liquidity preference theory** developed by Hicks (1946) explains the existence of such a term premium with the risk aversion of investors: to compensate for increasing risk levels and liquidity constraints associated with longer-term capital lock in, risk averse investors demand a positive term premium that is strictly increasing with maturity.

By contrast, the **preferred habitat theory** established by Modigliani and Sutch (1967) is based on the assumption that market participants have differing planning horizons and a preferred maturity segment to invest or to borrow in. Within these individual preferred habitats, supply and demand of an asset determine the size of its risk premium. Spill-over effects to other, primarily neighboring, maturity segments can occur, once risk premia are attractive enough to incite market participants to shift away from their preferred habitat.

A more extreme version of this theory, completely abstracting from the idea of expected short-term rates and term-premia shaping long-term yields is the **market segmentation theory** introduced by Culbertson (1957). Based on the assumption of extremely strong

[10]For a textbook treatment, see e.g., Bofinger (2001) and Fabozzi (2012).

maturity preferences of market participants and hence complete *insubstitutability of assets* of different maturities, this hypothesis predicts that bond prices and yields are solely determined by the supply and demand within the individual market segment and entirely disconnected from assets of neighboring maturities.

Since the strong assumptions of strict market segmentation theory are hard to reconcile with the empirically smooth term structure, recent research on term-structure models mostly supports a mix of expectations theory, liquidity preference theory, and preferred habitat theory. In this vein, recent theoretical approaches to support the assumption of limited asset substitutability—consistent with preferred habitat theory—were developed, e.g., by Daines et al. (2012), who explain preferred asset classes by regulation incentives and market convention; and by Hanson and Stein (2015), who find specialized long-term investors, such as insurance companies, to be more yield-sensitive than other market participants.

3.2.2 Effects of monetary policy on the yield curve

Besides explaining the slope of the yield curve, theories on the term structure of interest rates furthermore feed into the explanation of effects that monetary policy exerts on the yield curve. Due to their relevance for investment and consumption decisions, financial market effects represent an important intermediate step in the transmission of monetary policy. Depending on the respective instrument applied by monetary policy, however, the size and the horizon of effects on the yield curve may differ substantially.

In conventional times, decisions on the level of the official interest rate directly impact the short end of the yield curve. Given a high degree of asset substitutability, the monetary stimulus further translates to medium and long-term yields by changing the expected path of future short-term rates, in line with expectations theory. In this vein, the central role of the expectations hypothesis is, inter alia, confirmed by a recent empirical study of ECB staff presented by Lane (2019). Based on a shadow-rate term-structure model by Geiger and Schupp (2018), this study reveals that for a sample ending prior to the launch of QE, the expectations component indeed explains more than 75% of the yield curve impact of conventional monetary policy measures by the ECB. This quantitative evidence and similar findings by Altavilla et al. (2019) underline that the transmission via the expectations channel is not one-for-one.[11]

[11]Recent studies relying on high-frequency movements in financial market variables partially explain this observation by stressing the relevance of central bank communication, i.e., on the economic outlook (Jarociński and Karadi, 2020) or the future path of interest rates (Altavilla et al., 2019; Cieslak and Schrimpf, 2019) in additionally shaping the yield curve through expectations of market participants.

While the transmission mechanism of policy rate changes along the yield curve is well understood in conventional times, it proves to be more than insufficient once the zero-lower bound of interest rates is reached and the main conventional monetary policy instrument of setting the short-term policy rate is de facto exhausted. In addition, increased uncertainty on financial markets can lead to limited asset substitutability and thereby further hamper the transmission of monetary policy stimulus along the yield curve. On these grounds, central banks across the globe decided to target the long-end of the yield curve directly by purchasing large amounts of medium to long-term securities with newly created money base in the aftermath of the financial crisis. At the same time, a large body of empirical literature has developed that proves QE to be effective in lowering long-term yields while leaving short-term rates largely unaffected for the U.S. (Gagnon et al., 2011; Bauer and Rudebusch, 2014; Krishnamurthy and Vissing-Jorgensen), the UK Joyce et al. (2011); McLaren et al. (2014), and the Eurozone (Altavilla et al., 2015; Andrade et al., 2016; Koijen et al., 2017; De Santis and Holm-Hadulla, 2017; Urbschat and Watzka, forthcoming; Altavilla et al., 2019, as well as Chapter 2 of this dissertation). Depending on the methodology applied, the documented quantitative effects of QE on long-term yields vary across studies.[12]

Nevertheless, there is some consensus among scholars on the theoretical transmission channels through which the stimulus of QE translates to the yield curve. The two mechanisms that are proposed most frequently to explain the reduction of long-term yields upon the announcement of QE and upon implementation are the *portfolio rebalancing* and the *signaling channel*. Under the *signalling channel*, QE programs confirm the central bank's commitment to an expansionary path and to keep short-term rates low for an extended period of time (see, e.g., Krishnamurthy and Vissing-Jorgensen, 2013; Bauer and Rudebusch, 2014). In line with the expectations hypothesis on the term structure of interest rates, the resulting expectations of lower future short-term rates affect long-term yields through the reduced expectations component. However, this effect is found to be of minor importance for the impact of large-scale asset purchase programs in the Eurozone according to recent research by Lemke and Werner (2020). The *portfolio rebalancing channel*, on the other hand, offers two potential explanations for reductions in term premia of long-term yields, based on the theories of market segmentation and preferred habitat: first, according to the *local supply effect*, central bank large-scale asset purchases lead to a reduction of the

[12] For an overview of literature on unconventional policies see, e.g., Andrade et al. (2016) and Borio and Zabai (2018).

bond supply and increased bond prices for those assets that are purchased by the central bank. Consequently, investors with a preference for these assets move to other assets of close substitutability to readjust their portfolio thus bidding up bond prices within the same maturity segment (D'Amico and King, 2013). As a result, the associated long-term yields decline. Second, the *duration extraction effect* formalized by Vayanos and Vila (2009), offers an alternative explanation via the transfer of aggregate duration risk from secondary markets onto the central bank balance sheet in the course of QE. The reduced free float of long-term bond supply and hence duration risk in the hands of risk-averse arbitrageurs manifests itself in declining risk premia across maturity segments, which lead price-sensitive investors to reshuffle their portfolios in line with their risk-bearing capacity (see, e.g., Gagnon et al., 2011; McLaren et al., 2014; Eser et al., 2019).

In addition to identifying these theoretical mechanisms, most studies support the view that these channels become operative upon announcement, as financial market participants immediately price in new information on the stock of assets to be held by the central bank ("stock effects") (see, e.g., Altavilla et al., 2015; Koijen et al., 2017; De Santis and Holm-Hadulla, 2017; De Santis, 2020). By contrast, "flow effects" from the actual execution of asset purchases through the same channels are found to be of minor importance (see, e.g., De Santis and Holm-Hadulla, 2017).

3.2.3 Effects of public debt management on the yield curve

The mechanisms explaining the financial market transmission of monetary policy, especially in the context of asset purchase programs, help to understand the interaction of public debt management with the yield curve, as well. Primarily, the yield curve serves as an indicator for the optimal design of PDM as "part of the overall macroeconomic policy framework which encompasses monetary, fiscal and macro-prudential policies." (OECD, 2019). At the same time, the free-float of duration risk on secondary markets is just as well influenced by active debt management policy of the public issuer as it is by large-scale asset purchase programs of the central bank. In the following, a closer look is taken on the interaction of PDM with the yield curve in line with its institutional setup and put existing studies on PDM and monetary policy into perspective.

As a macroeconomic policy tool, PDM is closely related to fiscal policy. The Revised Guidelines For Public Debt Management define PDM as "the process of establishing and executing a strategy for managing the government's debt in order to raise the required amount of funding at the lowest possible cost over the medium to long run, consistent with a prudent degree of risk"(International Monetary Fund and World Bank, 2014). In line with its mandate, PDM is responsible for deciding and implementing the maturity, interest rate, and currency composition of government debt, whereas the level of government debt is prescribed by the fiscal authority and therefore outside the scope of PDM. In this vein, the maturity structure of outstanding debt is often viewed as the core policy variable of PDM (Scherf, 2011; Blommestein and Turner, 2012).

From a theoretical point of view, the volume and structure of public debt and taxes should not matter for household welfare according to the debt neutrality theorem of Barro (1974). However, the failure of his strong underlying assumptions—in particular, the assumption of frictionless financial markets—imply the relevance of PDM and its optimal execution (see, e.g. Greenwood et al., 2016b; De Fontenay et al., 1995). Most literature on the optimal maturity structure of public debt is based on the trade-off between cost and risk that PDM faces for a given maturity, which is largely reflected by the yield curve. Given the typically positive slope of the latter, suggesting the existence of a positive term premium, the cost of servicing debt of longer maturities is usually higher compared to the interest due for short-term debt. However, the risks of issuing short-term debt are looming larger, since the consequent necessity to roll over public debt more frequently increases both interest rate risk and the chances of a rollover crisis (Greenwood et al., 2014).[13] Against the backdrop of such counteracting incentives for PDM, it is not surprising that the literature on optimal maturity structure of public debt substantially varies in its conclusions. While one part of the literature recommends a longer maturity structure of public debt to limit the risks entailed with frequent rollovers (Alesina et al., 1990; Missale, 2012), more recent empirical studies have a tendency to highlight on the cost aspect of sovereign debt and hence recommend issuing short (Greenwood et al., 2016b; Maravalle and Rawdanowicz, 2018).

Beyond using the yield curve as an *indicator* to decide upon the optimal maturity structure, implementing the maturity structure also *exerts* an influence on the yield curve itself, since changes to the flow of public debt slowly transmit to the stock, as well. Similar to the yield depressing effect that unconventional monetary policy exerts on the yield curve by *increasing the demand* of long-term sovereign bonds—known as the duration extraction

[13]In this context, Alesina et al. (1990) point out that the risk of a rollover crisis, as a result of fading investors' confidence in the government's ability to service public debt, is especially present for high-debt countries.

channel—, changes to the relative supply of long-term bonds caused by adjustments in the maturity structure of gross issuance by PDM could influence the long-term yield accordingly. Notably, these strong parallels between monetary policy and PDM were already acknowledged by Keynes (1936), Tobin (1963), and Friedman (1960).

Overall, yield raising effects of maturity extensions by PDM through increases in the long-term sovereign bond supply are widely supported by empirical studies.[14] Hence, to limit potential counteracting effects of both policies in conventional times, when monetary policy is usually operative at the short end of the yield curve, many scholars advise PDM to mainly issue at the long end of the yield curve (Tobin, 1963; Bofinger, 1997; Hoogduin et al., 2011). This separation is further formalized in an institutional separation of PDM and monetary policy. While in most European countries both tasks have historically been executed by the central bank, many countries—including Germany—founded independent public debt agencies in the late 90s (see, e.g. Goodhart, 2010).

With the advent of QE, the effective separation between PDM and monetary policy via operating at opposing ends of the yield curve became less clear. Since both policies are recently active at the long-end of the yield curve, potential interactions are conceivable. In this vein, Greenwood et al. (2016a) study potentially counteracting effects of monetary policy and PDM for the simultaneous occurence of QE programs and maturity extensions by PDM in the U.S. between 2008 and 2014. According to their back of the envelope calculation, the cumulative impact of the Treasury's maturity extensions on long-term interest rates cancels out the yield-reducing effect of the Federal Reserve's QE measures by approximately one third in size.

Importantly, it shall be noted that the simultaneous occurrence of maturity extensions and unconventional monetary policy may not be isolated phenomena and that PDM decisions may partly be reflective of the (expected) impact of unconventional monetary policy on the yield curve. In this vein, for conventional times, an endogenous reaction of PDM to policy rate cuts in the form of shortening maturity of outstanding debt is shown by Blommestein and Turner (2012) in U.S. data. Similarly, Hoogduin et al. (2011) find a direct link between both policies for the Eurozone.

[14] See, for instance, evidence from term structure models (Greenwood and Vayanos, 2014), error correction models (Meaning and Zhu, 2012), bond-spread regressions (Krishnamurthy and Vissing-Jorgensen, 2012), as well as event studies for the 2000–2002 Treasury buy-backs (Greenwood and Vayanos, 2010) and for the 1962 Operation Twist (Swanson, 2011). Related, Greenwood et al. (2016a) find that the size of the effect on the term premium induced by the Treasury's announcements is only roughly one half of the effect triggered by announcements on comparably sized large-scale asset-purchases by the Fed. The authors rationalize this observation by the high transparency of Treasury actions and thus a smaller surprise component of Treasury announcements as compared to Fed announcements.

In line with these findings for conventional times, the shift in the conduct of monetary policy toward targeting the long-end of the yield curve during QE is likely to entail feedback effects on the conduct of PDM, as well. Specifically, against the backdrop of the large body of existing literature showing that the long end of the yield curve—reflecting the cost of servicing newly issued long-term public debt— is lowered by announcements of large-scale asset purchase programs, an extension of maturity in public debt following expansionary QE announcements is conceivable. This is exactly where this paper steps in by providing qualitative and quantitative evidence for the causal effect of QE in the Eurozone on the maturity structure in gross issuance of German public debt.

3.3 Data and descriptive analysis

As a first step of my empirical analysis, I establish narrative evidence on the interplay between unconventional monetary policy and public debt management. For this purpose, I describe the construction of three main specifications of event series that serve as proxies for unconventional monetary policy actions, as well as a series of short-term maturity adjustments by PDM. Beyond that, I also discuss underlying trends in the data and infer stylized facts on potential interactions in a joint descriptive analysis of both series.

3.3.1 Establishing the data and stylized facts

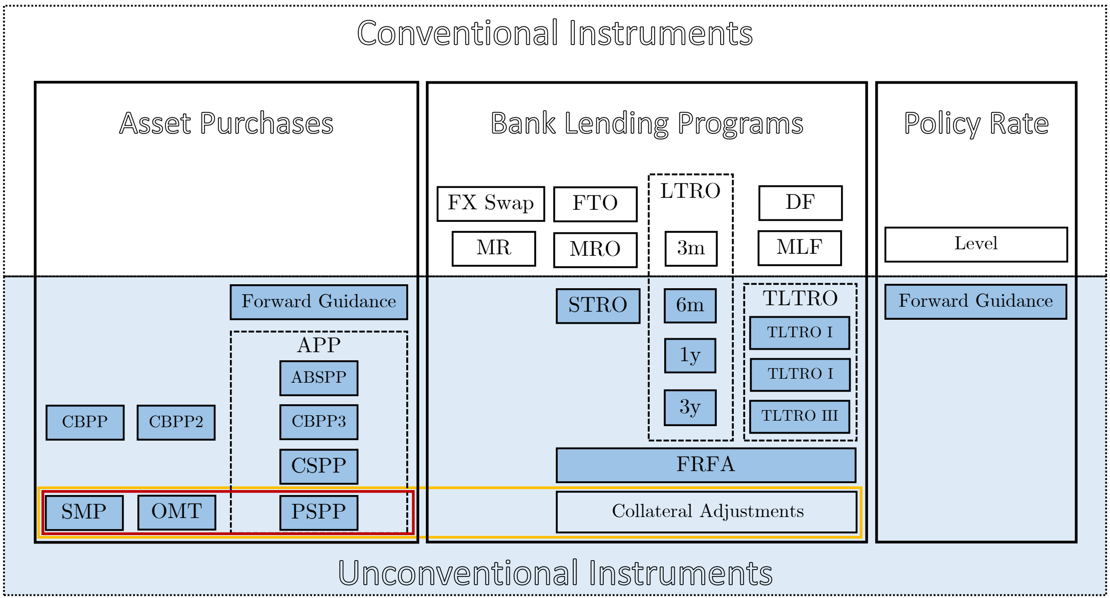
Similar to other studies on interactions between monetary policy and PDM, my empirical analysis focuses on the direct influence of the monetary policy variable on the maturity structure, as the core policy variable of PDM. Other studies in this field by Hoogduin et al. (2011) and Blommestein and Turner (2012) estimate the effect of the level of interest rates as the main instrument of *conventional* monetary policy on the maturity structure of *debt outstanding on secondary markets*, which they consider as an appropriate measure for maturity decisions by PDM. Although closely related, this study deviates from their approach in two important dimensions: first, it centers on the repercussions of *unconventional* monetary policy actions on PDM and, second, i employs *short-term maturity adjustments in gross issuance on the primary market* as a more sensitive measure of active policy choices made by PDM.

Monetary policy proxy In contrast to the times prior to the Financial Crisis, when European monetary policy mainly consisted of setting short-term interest rates, the presence of the zero lower bound necessitated the launch of various unconventional monetary policy measures for the ECB to maintain its expansionary path. The categorization of all monetary policy measures taken in the history of the ECB is depicted in Figure 3.2 and underlines the complication in identifying one single unconventional monetary policy tool. To overcome this difficulty, I follow event-study literature by building a binary dummy time series that takes a value of one for those dates, when unconventional monetary policy impulses, i.e., an announcement about future unconventional programs, take place. Similar to Fratzscher et al. (2016); Dedola and Georgiadis (2018); Enders et al. (2019), I limit the set of events to announcements that were made by the ECB in press releases and press conferences. Other studies by Cieslak and Schrimpf (2019) and Urbschat and Watzka (forthcoming) limit their selection of announcements further by the degree of news coverage they triggered.[15] However, for the question at hand of my study, I do not deem the level of media presence of an announcement—which captures several dimensions of policy measures—but its content with regard to its potential relevance for PDM as decisive. Similar to the approach taken by Dedola and Georgiadis (2018), who limit their set of events to those announcement with a sizeable quantitative effect on the ECB's balance sheet, I restrict my selection of announcements to those events that are likely to alter the stock and composition of *sovereign bonds* on the central bank balance sheet. For this purpose, I qualitatively evaluate the content of 92 press releases and press conferences on monetary policy decisions from the ECB website over the sample period 2014M1 to 2019M3 to identify sovereign bond market related announcements. Specifically, out of the universe of the ECB's various policy measures, I make a selection for the construction of the policy proxies along the following criteria: first, in a narrow definition, I include announcements on large-scale asset purchase programs involving purchases on sovereign bond markets that affect the projected stock of sovereign bonds held by the central bank (Sovereign Bond Purchase proxy: SBP_t). Second, in a broader definition, I extend this measure by adding announcements that affect the eligibility of certain sovereign bonds to be accepted as collateral for refinancing at the ECB to the set of events (Sovereign Bond News proxy: SBN_t).[16]

[15] Cieslak and Schrimpf (2019), for instance, base their selection of events on their mentioning in the Bloomberg economic calendar, whereas Urbschat and Watzka (forthcoming) limit their set of announcement events to those monetary policy meetings that have been discussed on the first three pages of the Financial Times the following day.

[16] The exclusion of certain high-risk sovereign bonds from the list of accepted collateral for refinancing operations may lead to substitution effects to other sovereign bonds accepted as collateral, especially those regarded as "safe havens". Vice versa, the readmission of higher-risk sovereign bonds to the ECB's collateral framework could entail negative substitution effects for sovereign bonds of issuers with lower risk of default.

Figure 3.2: Categorization matrix of ECB monetary policy tools.

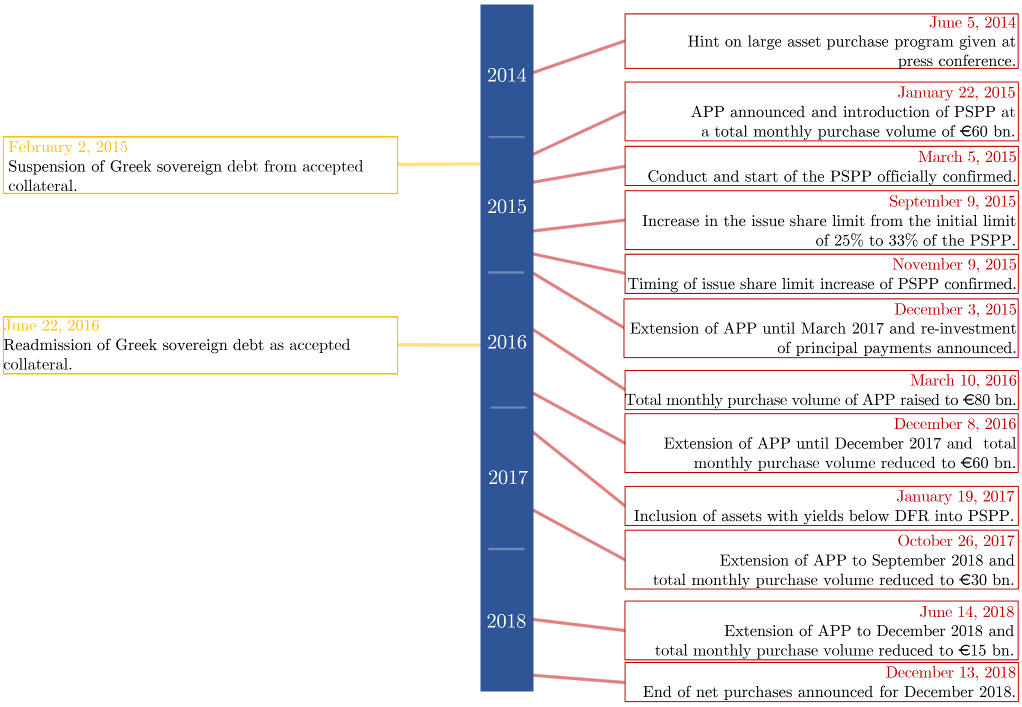


Notes: The two horizontally divided areas distinguish conventional (white upper area) and unconventional monetary policy instruments (lower light blue area), whereas the three vertically separated boxes differentiate categories, to which the individual instruments can be attributed, such as asset purchases, bank lending programs, and policy rate decisions. The colored frames in the lower part of the Figure show the programs included in the narrow SBP_t and wider SBN_t proxy definitions of unconventional monetary policy. The abbreviations used stand for the following programs: Foreign Exchange Swaps (FX Swap), Fixed Term Operations (FTO), Longer-Term Refinancing Operations (LTRO, with terms from 3 month to 3 years), Deposit Facility (DF), Minimum Reserve (MR), Main Refinancing Operations (MRO), Marginal Lending Facility (MLF), Special-Term Refinancing Operations (STRO), Targeted Longer-Term Refinancing Operations (TLTRO, in three editions), Fixed Rate Full Allotment Procedure (FRFA), Expanded Asset Purchase Programme (APP), Assed Backed Securities Purchase Programme (ABSPP), Covered Bond Purchase Programmes (CBPP, in three editions), Corporate Sector Purchase Programme (CSPP), Securities Markets Programme (SMP), Outright Monetary Transactions (OMT), and Public Sector Purchase Programme (PSPP).

Figure 3.3 depicts the exact event dates included in the two binary event series. Notably, since my empirical analysis focuses on the era of QE, which is commonly viewed to have started with the first indications of large-scale asset purchases for the Eurozone in 2014 (Altavilla et al., 2019), the narrow proxy series SBP_t de facto only includes announcements on the PSPP that provide news on the stock of sovereign debt on the central bank balance-sheet. In the wider SBN_t series, this set of events is extended by two announcements on the ECB's acceptance of Greek sovereign bonds as collateral within the sample.[17] Consequently, despite their potential relevance for public debt managers, announcements regarding earlier asset purchase programs, such as the Securities Markets Programme (SMP) and Outright Monetary Transactions (OMT), are excluded from both proxies due to their

[17]In particular, the announcement on the exclusion and readmission of Greek sovereign bonds from the collateral framework allows to draw inferences on the future stock of German sovereign bonds on the central bank balance sheet, as these may serve as substitute collateral in the meantime.

Figure 3.3: Timeline of announcements included in binary event series



Notes: The timeline shows the date and content of the unconventional monetary policy announcements that are likely to affect PDM decisions. Events on the right hand side of the timeline are included in $\boxed{SBP_t}$, whereas events on the left hand side are additionally included for the construction of $\boxed{SBN_t}$. The abbreviations used stand for the following programs: Expanded Asset Purchase Programme (APP), Deposit Facility Rate (DFR), Public Sector Purchase Programme (PSPP).

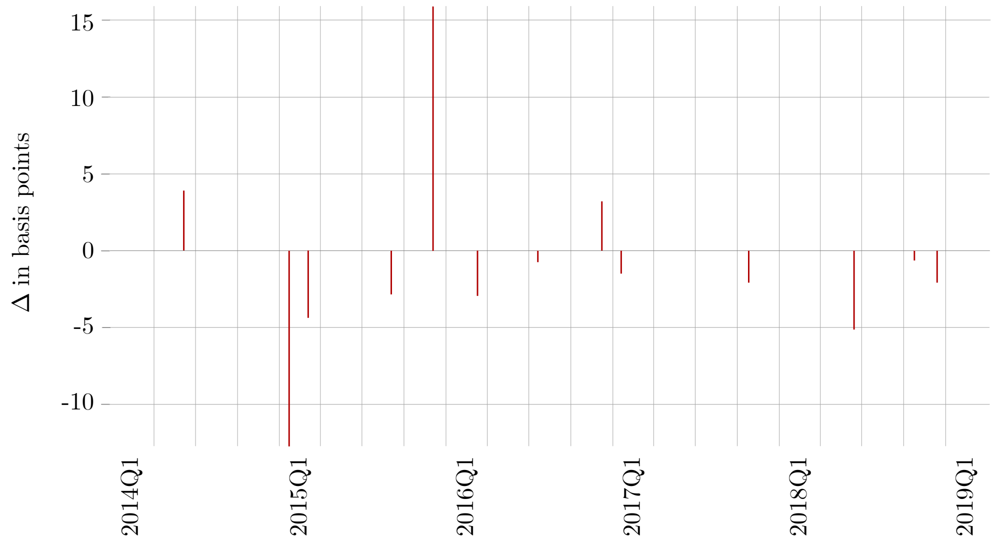
earlier occurrence. With the majority of events in the selected sample referring to asset purchase programs meeting the definition of QE, the study sometimes loosely refers to SBP_t and SBN_t as proxies for QE.

A potential downside of this binary approach is that all announcement dates are weighted equally regardless of their stock effect. This is especially relevant, as the announcements in included into the event series are heterogeneous in the direction of monetary policy stimulus on German sovereign bond markets in the sample. While in the narrow SBP_t series, the announcements until March 10, 2016 increased the projected stock of German sovereign bonds held by the central bank, an opposite effect resulted from the announcements on the phasing out and end of net asset purchases thereafter. Furthermore, the announcements on the suspension and subsequent readmission of Greek sovereign bonds, additionally included in the wider SBN_t, pose further divergent effects on German sovereign bond markets. Yet, the different quality of announcements can not be distinguished in these pure binary event series. Therefore, as an alternative to the binary dummy series, I apply changes in financial market variables—such as asset prices, yields, and spreads—in narrow time windows identified by high-frequency literature (i.a., Gürkaynak et al., 2005; Cieslak and Schrimpf, 2019; Altavilla et al., 2019; Jarociński and Karadi, 2020) to infer the quantitative effect of monetary policy actions. In this vein, I make use of the innovative Euro-Area Monetary Policy Event-Study Database (EA-MPD) developed by Altavilla et al. (2019), which provides high-frequency movements in selected financial market variables measured in different narrow time windows on ECB monetary policy meeting days.[18] From this data set, I take the reported changes in 10-year Bund yields within the 2 hour 25 minute long monetary event window to quantify the *magnitude and direction* of surprise shifts in yields that selected policy announcements on monetary policy meeting days induced on the German Bund market. In this vein, the reported change in basis points within the monetary event window indicates whether investors understand an announcement relating to an unconventional monetary policy measure as a loosening—marked by a drop in long-term yields—or tightening—marked by a rise in long-term yields—of the policy stance compared to their expectations. In particular, I use these short-run yield adjustments to weigh the individual events in two alternative proxy definitions, i.e., SBP_t and SBN_t, which results in quantified versions of these proxies. Note, however, that since data on high-frequency yield adjustments in the EA-MPD is only available for those events that took place on a monetary policy meeting day, the quantified series omit some of the previously specified events. This results in an even narrower selection of quantified events, which is identical for the previously differing binary event dummy series SBP_t and SBN_t for the relevant sample. Figure 3.4 plots this quantified event series of unconventional central bank sovereign bond market actions (Quantified Sovereign Bond proxy: QSB_t).

As opposed to the two pure dummy series, which merely indicate the *occurrence* of an unconventional announcement according to the aforementioned selection criteria, the quantified QSB_t series conveys additional information on the direction of subsequent yield adjustments. In addition, the varying size of yield changes allows us to take the potentially differing impact of the individual announcements into account.

[18] In constructing their database, Altavilla et al. (2019) quote changes in financial market variables for the *monetary event window* between 13:25–15:50 on a Governing Council meeting day as well as two enclosed event windows around the press release (13:25–14:10) and press conference (14:15–15:50). The changes in financial market variables reported for each window are computed as the difference in the mean quote of the first 10 minutes and the last 10 minutes of the window.

Figure 3.4: Changes in 10-year Bund yields on selected event days



Notes: The figure displays the values of quantified event series of unconventional monetary policy actions on sovereign bond markets QSB_t over the sample. The red bars represent the change in basis points of the 10-year Bund yield within the monetary event window according to Altavilla et al. (2019) for those dates of the SBP_t and SBN_t series that coincide with a General Council policy meeting during 2014Q1 to 2019Q1.

The depiction of the QSB_t series in Figure 3.4 over time shows that the majority of selected unconventional monetary policy announcements with regard to sovereign bond markets were initially perceived as expansive by financial markets. In particular, only three event dates surprised the markets by being less accommodating than expected. The largest yield reduction of 12.7 basis points within the monetary event window, i.e., the event that was perceived as most expansive by financial markets, was the initial announcement of the PSPP that took place in January 2015. By contrast, the largest yield increase of 15.9 basis points is associated with the extension of net asset purchases under the PSPP until March 2017 that was announced in December 2015. According to this adverse yield reaction, markets appear to have expected a stronger extension of the PSPP to be announced on this day. Overall, the declining size of surprise yield-shifts over time reported in the QSB_t series are qualitatively in line with evidence on a gradually diminishing impact of repeated QE announcements on the yield curve provided by Eser et al. (2019).

Public debt management instrument Out of its responsibilities of deciding and implementing the maturity and interest rate structure as well as the currency composition of government debt, the choice of the appropriate maturity structure can be viewed as the main policy instrument of PDM (Scherf, 2011). In line with proponents of the duration extraction channel of unconventional monetary policy, such as Greenwood and Vayanos

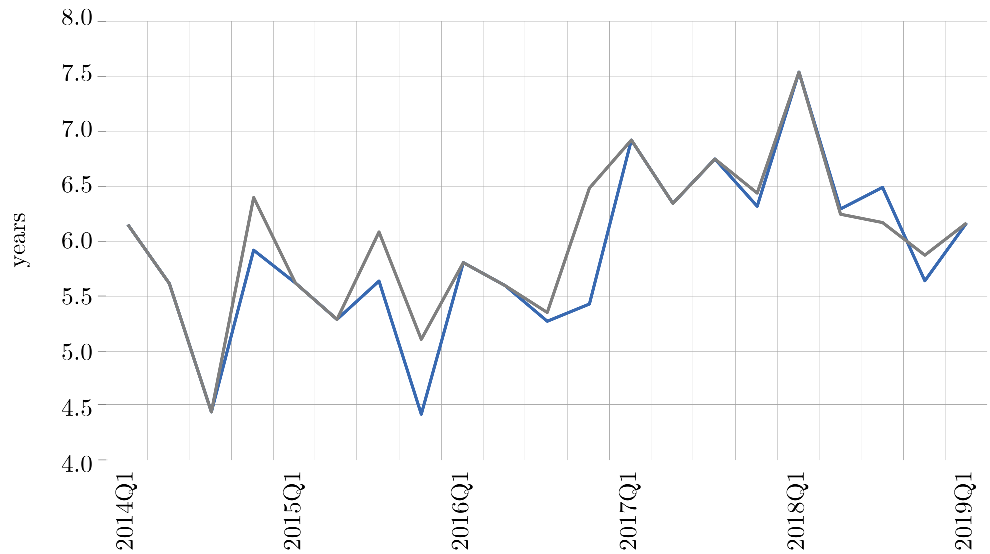
(2014), Greenwood et al. (2014), and Li and Wei (2013), the yield curve impact of maturity structure decisions by central banks and debt management offices can be expected to transmit through its implied changes in the Macaulay duration of outstanding debt in the hands of the investors outside the fiscal and central banking sector. In this vein, many scholars analyze the *duration* of the outstanding *stock* of debt as the target variable for PDM. In the absence of a continuous time series on the Macaulay duration, the WAM, i.e., the volume weighted average maturity, of outstanding debt can serve as a proxy for the duration risk in the market, as suggested by Eser et al. (2019). However, the only way how PDM can *actively* influence the maturity structure of outstanding debt is via changes in the *flow* of debt. Since early redemption and buy-backs on secondary public debt markets by the issuer are legally forbidden for German sovereign debt instruments (Bundesministerium der Finanzen, 2012a,b), it is essentially the *maturity structure of gross issuance*—besides the regular ageing of the outstanding portfolio—that shapes the duration in the hands of the public from a PDM perspective. I thus measure the maturity structure of gross issuance by computing the WAM of a newly issued portfolio within a period. The WAM for a respective period t can be calculated by weighing the maturity M of each portfolio component j with its nominal value q and by dividing it by the nominal volume of the entire newly-issued portfolio, consisting of n elements as follows:

$$WAM_t = \sum_{j=1}^{n} M_j \cdot \frac{q_j}{Q}, \text{ with } Q = \sum_{j=1}^{n} q_j, \tag{3.2}$$

Depending on the point in time at which maturity is measured, different WAM figures can be distinguished. Specifically, the maturity used for the computation of the WAM can be either the *original*, i.e., the term to maturity measured on the issuance date of the corresponding security, or the *residual* term to maturity measured at the end of the respective period observed. Since for gross issuance the components j change for each period t, inserting the residual maturity into the calculation of the WAM would lead to a smaller weight of securities issued at the beginning as compared to those issued at the end of period t. To avoid such a distortion, I calculate the WAM based on the original maturity M, i.e., the difference between the maturity date and issuance date, for each issued component j within the respective period t.

Taking stock of data availability for WAM measures of German sovereign debt, I do not find an adequate time series in public databases. In the few cases, in which WAM data are provided, the series either reflects the WAM of debt outstanding, or is available only at low frequencies or a small number of observations.[19] Yet, to analyze the possible causal effects of monetary policy communication on primary market maturity decisions of PDM, the availability of the maturity structure of issuance flows for a larger number of observations and at a higher frequency of data is imperative. To obtain a time series for the WAM of new issuance of German government debt at a sufficiently high frequency, I thus proceed to construct data in a bottom-up approach, involving individual bond-level data and aggregating them up to produce consistent and comprehensive quarterly and monthly time series. For this purpose, a detailed data set, containing tick data including types, volumes, and maturities for each issuance of all publicly auctioned gross issuance of German sovereign debt on the primary market since 1999 was kindly provided upon request by the German Finance Agency.

Figure 3.5: Planned and actual WAM of German public debt gross issuance



Notes: The *planned* WAM (blue line) compared to the *actual* WAM (grey line) of gross issuance adding towards the stock of German marketable government debt issued in public auctions are plotted for a data sample covering 2009Q1 to 2019Q1. The horizontal axis measures time in months, whereas the vertical axis measures WAM in years. For illustration purposes, I plot quarterly averages of the underlying monthly time series for the original WAM of gross issuance.

Nevertheless, this data needs to be handled with caution when analyzing monetary policy effects. A potential drawback of the WAM of new issuance is that it may be driven by a structural component, e.g., by a long-run trend and changes in the target value for

[19]The two only publicly available time series on the WAM of German sovereign debt reside in the BIS Debt Securities Statistics and the OECD Central Government Debt Database. Both sources contain the WAM of *outstanding* debt, with the BIS providing 12 data points covering annual data from 2007 to 2018. The corresponding series from the OECD offers 14 annual data points ending in 2010.

the WAM of outstanding stock as well as a seasonal component, e.g., a default issuance rhythm for bonds of certain maturities. The good news is, however, that these long-term adjustments are usually gradually introduced and communicated well in advance by the German Finance Agency. Unconventional monetary policy news, on the other hand, might entail a more spontaneous reaction of German PDM. Therefore, to observe such short-run policy adjustments, I build a unique time series of *unplanned deviations* in the WAM by subtracting the planned WAM of gross issuance from the WAM emerging from the actual auction for each period.

While a data set on auction data from which the maturity structure of gross issuance can be inferred, is provided by the German Finance Agency, a complication is the lack of a counterpart data set for planning data on gross issuance, which is required for the calculation of deviations from plans. To overcome this challenge, I laboriously retrieve the required data from annual and quarterly press releases in the media archive on the German Finance Agency's website. To the best of my knowledge, I am the first to exploit this valuable source of information for research purposes. For the construction of the time series, I proceed as follows. First, I build a data set reflecting the German Finance Agency's plans for monthly gross issuance, from which I calculate the WAM according to Equation 3.2. Second, I collect information from annual issuance outlooks published in December on planned issuance for the entire upcoming year. Third, I compare the WAM from annual plans with the one actual auction data. Note that the German Finance Agency releases three quarterly press releases "Issues planned by the Federal Government" as updates to the annual issuance plan, each March, June, and September that announce short-run deviations from the original plan for the subsequent quarter. In anticipation of the specification of my econometric framework in Section 4, I would like to highlight that due to the German Finance Agency's commitment to communicate any unplanned changes to their emissions one quarter ahead of their implementation, (i) no deviations in WAM are possible in the first quarter of each year and (ii) any surprise in unconventional monetary policy can not simultaneously cause adjustments in the conduct of PDM. Due to the necessity of prior notifications of any deviations from planned issuance by the German Finance Agency, there should be at least one quarter of a time lag in any PDM reactions to QE events.

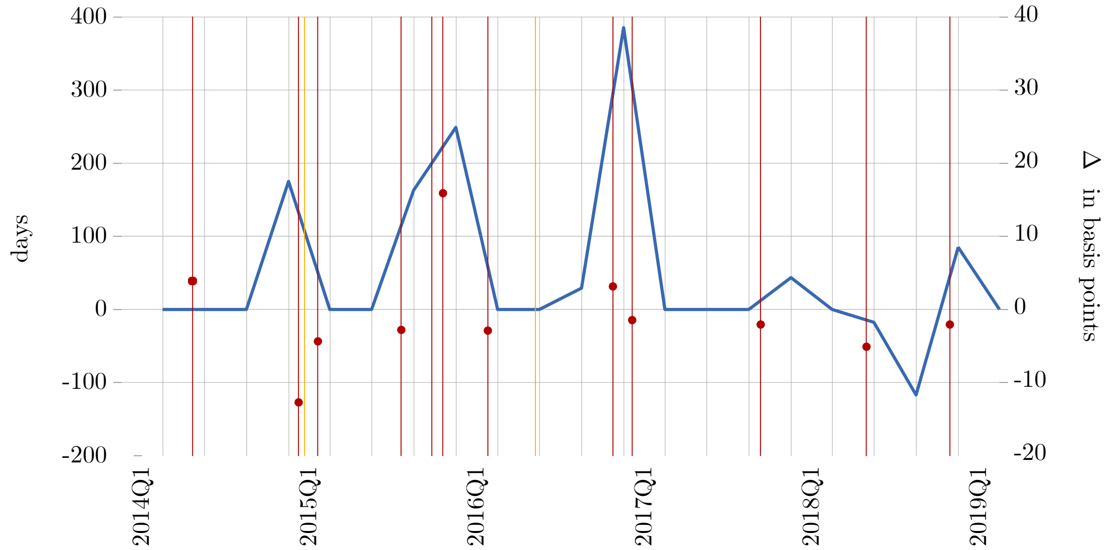
Figure 3.5 plots the resulting evolution of the *planned* (blue line) and of the *actual* WAM of gross issuance (grey line), over time. Both depicted WAMs show some oscillation that may be attributed to issuance "seasonality", i.e., the issuance frequency is regular but not recurring in a fixed quarterly rhythm. As previously shown in Panel (b) of Figure 3.1, a clear positive trend of the actual WAM of gross issuance is present since 2009. For the QE sample starting in 2014, depicted in Figure 3.5, this trend appears to be slightly muted, yet still visible in the actual WAM and also consistently of actual issuance in the planned WAM. In this sample, the shortest average maturity actually issued within a quarter amounts to 4.5 years measured in 2014Q4, after which a strong increase that peaks in 2018Q1 at an average maturity in gross issuance of 7.5 years can be observed. For the WAM according to the annual issuance plans, the picture is qualitatively similar. Yet, when comparing both WAMs in more detail, considerable deviations in several quarters of the sample catch the eye. These tend to be positive on average implying unplanned extensions of maturity. The only negative deviation in the WAM of actual gross issuance as compared to the annual issuance plan, i.e., an unplanned shortening of maturity, occurs in 2018Q3. On average these deviations amount to an unplanned maturity extension of 139 days. The range of variation lies between a minimum of 117 days of unexpected shortening (2018Q3) and 385 days of unexpected lengthening (2016Q4) of maturity.

3.3.2 Joint descriptive analysis

In what follows, I establish a narrative example to support the argument that unconventional monetary policy may change the preferred maturity structure of newly issued debt by PDM. For this purpose, I relate the observed deviations in the maturity structure of actual versus planned gross issuance to the previously identified unconventional monetary policy events. The results from this joint descriptive analysis provide tentative evidence for an unplanned extension in the maturity of gross issuance by the German Finance Agency in response to unconventional monetary policy announcements by the ECB with a delay between two and three quarters.

For a sample ranging from 2014Q1 to 2019Q1, Figure 3.6 shows deviations in actual from planned WAM of gross issuance within a quarter (blue line graph) and the binary event dummies (red and yellow vertical lines), and the quantified event series (red dots) representing changes in Bund yields on event dates that coincide with a General Council meeting. In 2014Q2, the ECB hinted a potential large-scale asset purchase program during a press conference. Although the positive basis point change within the monetary event window suggests that this announcement was perceived as less expansive than expected on the German sovereign bond market, a positive WAM deviation of 175 days followed

Figure 3.6: WAM deviations by PDM and unconventional monetary policy proxies



Notes: I plot unplanned changes in the policy instrument of PDM together with the various unconventional monetary policy proxies at quarterly frequency for a data sample covering 2014Q1 to 2019Q1. The blue line chart shows deviations of actual versus planned WAM in gross issuance in days on the left vertical axis. The vertical lines represent event-dates included in SBP_t (red lines) and SBN_t (red an yellow lines). The red dots on some of these vertical lines show the changes in the 10-year Bund yield in basis points from the quantified event series QSB_t on the right vertical axis for those event dates that coincide with a monetary policy meeting.

after two quarters in 2014Q4. Soon thereafter, the ECB announced the launch of the PSPP, excluded Greek sovereign bonds from accepted collateral, and provided additional details on the PSPP in 2015Q1. From these announcements, the two relating to the PSPP took place on monetary policy meeting days and were perceived as strongly expansionary according to the yield surges of more than 12bp and 5bp within the monetary policy event window. In 2015Q3 an increase in the WAM of 163 days followed. Shortly after, another expansive announcement of an increase in the issue share limit from 20 to 33% for the PSPP was indicated, which potentially may have fueled an additional positive deviation in the WAM of 249 days in 2015Q4. A prolongation of the PSPP until March 2017 was announced in 2015Q4, yet perceived as less expansive compared to market expectations and followed by only a small increase in WAM of 29 days in 2016Q3. Similarly, an increase in the WAM by 385 days could be observed in 2016Q4. This substantial increase in maturity followed (i) after the ECB had announced to raise the monthly purchase volume of the PSPP to €80 billion, which was perceived as expansionary by financial markets, in 2016Q1, and (ii) after the readmission of Greek sovereign bonds to the set of accepted collateral was announced in 2016Q2. Since the prolongation of the APP at a lower monthly purchase volume until

December 2017 and the acceptance of sovereign bonds with negative yields in late 2016Q4 and 2017Q1 induced comparably small adjustments to expectations in German Bund markets, they were consequently followed by a smaller maturity extension of gross issuance of 44 days in 2017Q4. The tapering of the APP, i.e., the prolongation of the program at a reduced monthly purchase volume (€30 billion) in September 2018, was perceived as slightly more expansive than expected, and was followed by an unplanned reduction in the WAM of 18 days in 2018Q2. This sign-flipped effect of tapering announcements on the deviation of actual WAM from its plan was repeated in 2018Q3, when the deviation in WAM was negative by 117 days, potentially triggered by the announcement of a further reduction of the monthly purchase volume by 50% in 2018Q4. Conversely and most recently sample, the announcement of ongoing re-investment of principals in 2018Q2 led to a positive deviation in the WAM in 2018Q4 by 85 days. Overall, this narrative analysis is suggestive to a *dynamic direct link* between unconventional monetary policy events and PDM.

3.4 Econometric framework and results

Since the approach from the previous Section is purely narrative in nature, more in-depth empirical methods are required to make statements on causality. Therefore, in the following, I describe the econometric methods used to study the causal effects of unconventional monetary policy announcements on the maturity structure of newly-issued German sovereign bonds. In contemporaneous and dynamic regression setups at monthly data frequency (i) the established QE announcement dummies are used to proxy for shifts in the monetary policy stance and (ii) the newly constructed series of WAM deviations from plan serve to proxy for endogenous PDM reactions. In line with the institutional characteristics of PDM plan adjustments as sketched above, these estimations unearth novel evidence suggesting that QE does typically not affect PDM decisions in the very short-run. Notably, the application of local-projection techniques for dynamic regressions as in Jordà (2005), reveals an economically meaningful link between QE and the maturity structure of newly issued German government debt with a delay of several months.

3.4.1 Time-series regressions: contemporaneous effects

In a first step of my quantitative analysis, I rely on time-series regressions to analyze the contemporaneous effect of QE announcements on the maturity structure of sovereign bonds measured by the WAM of new issuance in German sovereign bonds. Specifically, I estimate the following equation by ordinary least squares and construct autocorrelation- and heteroscedasticity-robust standard errors according to Newey and West (1987) for the period of 2014M1 to 2019M3 in monthly data:

$$WAMD_t = c + \beta \cdot MPE_t + \boldsymbol{\gamma} \cdot \mathbf{z}_{t-1} + u_t. \tag{3.3}$$

In testing influence of unconventional monetary policy events by the ECB, MPE_t, on the deviation of actual to planned WAM of monthly gross issuance in German fixed rate debt instruments, $WAMD_t$ the parameter c serves an intercept and β as the coefficient of interest. In addition, $\mathbf{z}_{t-1}$ denotes a vector of lagged controls, $\boldsymbol{\gamma}$ is the corresponding vector of coefficients, and u_t represents regression residuals. I run the estimation for different specifications of the contemporaneous monetary policy event variable MPE_t by inserting one of the unconventional monetary policy proxies SBP_t, SBN_t, or QSB_t, one at a time.[20] Since the proxy series represent the timing of a specific set of events or short-term financial market reactions that—by design—are exogenous to the deviation in WAM, the estimation theoretically does not require further control variables. Nevertheless, I account for the lagged influence of macroeconomic and financial conditions, by including in $\mathbf{z}_{t-1}$: bond yields i_{t-1}, the growth rate in industrial production y_{t-1}, inflation π_{t-1}, the growth rate of the outstanding volume of public debt ov_{t-1}, and the planned WAM of gross issuance WAM^P_{t-1}. In addition, I include the first lag of WAM deviations, i.e., $WAMD_{t-1}$ to control for autocorrelation.[21] I run the time series regressions (i) for all three unconventional monetary policy proxies, and (ii) by abstracting from macroeconomic and financial controls and by alternatively controlling for them; the first lag of $WAMD_{t-1}$, however, always enters $\mathbf{z}_{t-1}$. Taken together, this strategy yields a total amount of six empirical specifications.

The corresponding point estimates, p-values, and Newey-West corrected standard errors for these six specifications are depicted in columns (1) to (6) of Table 3.1. Note that, in Section 3.3, I have argued already that institutional specifics of PDM, such as decision and implementation lags within the German Finance Agency, suggest that these

[20] The events are transformed from an irregular daily frequency into a monthly frequency. For the binary dummies, the proxy takes a value of 1 in case that an announcement occurred within a month, and 0 otherwise. For the quantified series, I accumulate the yield surprises that occur within a month to obtain a monthly series.

[21] I retreive the outstanding volume of public debt and the planned WAM according to the annual issuance plans at a monthly frequency from my unique Finance Agency dataset established in Section 3.3.1. To receive the German 10-year Bund yield, industrial production, and the consumer price index at a monthly frequency, I utilize the FRED database. The control variables are included in first lags to account for potential endogeneity.

Table 3.1: Contemporaneous relations between QE and WAM deviations

	Unconventional Monetary Policy Proxies					
$WAMD_t$	SBP_t		SBN_t		QSB_t	
	(1)	(2)	(3)	(4)	(5)	(6)
MPE_t	p=0.66 0.06 (0.13)	p=0.64 0.07 (0.15)	p=0.75 0.04 (0.12)	p=0.86 0.02 (0.14)	p=0.23 − 0.02 (0.01)	p=0.24 − 0.02 (0.02)
$WAMD_{t-1}$	p=0.09 0.26 (0.15)	p=0.08 0.27 (0.15)	p=0.08 0.27 (0.15)	p=0.06 0.28 (0.15)	p=0.06 0.31 (0.17)	p=0.05 0.33 (0.16)
i_{t-1}		p=0.40 − 0.08 (0.10)		p=0.41 − 0.09 (0.10)		p=0.37 − 0.08 (0.09)
y_{t-1}		p=0.97 − 0.00 (0.02)		p=0.94 − 0.00 (0.02)		p=0.65 − 0.01 (0.02)
π_{t-1}		p=0.17 − 0.38 (0.27)		p=0.16 − 0.37 (0.26)		p=0.19 − 0.35 (0.27)
ov_{t-1}		p=0.70 0.01 (0.04)		p=0.63 0.02 (0.04)		p=0.44 0.03 (0.03)
WAM^P_{t-1}		p=0.26 0.03 (0.02)		p=0.29 0.02 (0.02)		p=0.28 0.03 (0.02)
R^2	0.08	0.13	0.08	0.13	0.09	0.14
$\bar{R}^2$	0.05	0.02	0.05	0.02	0.06	0.04

Notes: The table shows OLS estimation results across different specifications of the monetary policy event MPE_t with confidence intervals constructed as proposed by Newey and West (1987). For the sake of brevity, I refrain from reporting the point estimates of the constant. The dependent variable is the deviation of the *actual* weighted average maturity of gross issuance on primary markets from the one *planned* according to the German Finance Agency's gross issuance plan, $WAMD_t$. Different specifications for the unconventional monetary policy proxy MPE_t are displayed in column (1) to (6). The three different specifications tested include two binary dummy event series, SBP_t in columns (1) and (2) and SBN_t in columns (3) and (4), as well as the quantified event series QSB_t in column (5) and (6). An extended set of control variables included in the estimation in even column numbers consists of the interest rate, i_{t-1}, output, y_{t-1}, inflation π_{t-1}, the outstanding volume of interest bearing public debt ov_{t-1} and the WAM according to the original plan WAM^P_{t-1}. AR(1) effects are controlled for by the inclusion of $WAMD_{t-1}$. I perform all estimations using data from 2014M1 to 2019M3 and present p-values above and Newey-West standard errors in parentheses below the point estimate.

contemporaneous regressions may severely mask the true impact of unconventional monetary policy shocks on the WAM of new issuance. The results from Table 3.1 confirm such notion: the contemporaneous response of the maturity deviation from gross issuance plans in all three specifications of unconventional monetary policy announcements within the same period are small, show mixed signs, and are, above all, insignificant. This finding further corresponds to the delayed PDM reaction suggested in the joint descriptive analysis in Section 3.3.2. When comparing the results in columns (1), (3), and (5) with those of columns (2),(4), and (6) in pairs, i.e., when comparing the specifications with and without controls, I observe that (i) the point estimates for the β coefficients are remarkably insensitive to the inclusion of controls, and (ii) the point estimates for SBN_t and QSB_t become less significant when including additional controls. These findings underline and add further credibility to the exogeneity assumption of the unconventional monetary policy proxies. Across different proxy specifications, the explanatory power of the regressions, measured by the adjusted $\bar{R}^2$, typically declines in the estimations that include additional controls. Remarkably, all the estimated coefficients in these settings turn out to be statistically not distinguishable from zero; the only exemption are the AR-parameters of WAM deviations, with point estimates ranging between 0.26 and 0.33 and thus pointing to a modest degree of autocorrelation in the deviation of actual WAM of gross issuance from annual plans. Overall, these results point to the need for a dynamic regression setup to infer the true causal effect of QE events on maturity decisions by PDM.

3.4.2 Local projections framework: dynamic effects

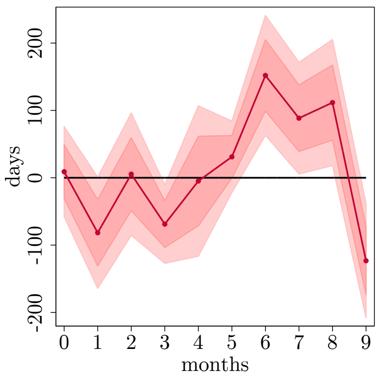
To establish causal evidence for a dynamic unplanned adjustment of maturity structure in gross issuance as a consequence of unconventional monetary policy announcements, I need to slightly adjust my regression setup. Specifically, I establish a set of dynamic single equation regression counterparts to Equation 3.3, so-called local projections as in Jordà (2005). For each of the three baseline specifications of monetary policy events captured by MPE_t, I specify the set of regressions as follows:

$$WAMD_{t+h} = c_h + \beta_h \cdot MPE_t + \sum_{i=1}^{l} \gamma_{i,h} \cdot \mathbf{z}_{t-i} + u'_{t+h-1} + u_{t+h}, \text{ for } h \in \{0, ..., 9\}. \quad (3.4)$$

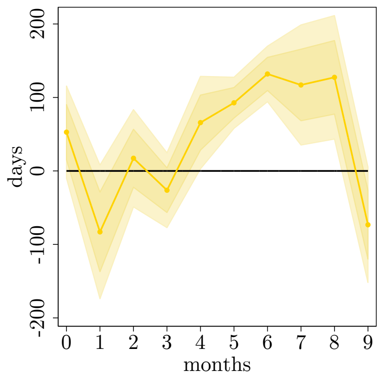
I estimate the deviations in WAM from its plan at horizon h in monthly data by running a series of regressions, where the sequence of β_h captures the coefficients of interest. For each horizon h each estimation includes a constant c_h, the monetary policy event MPE_t at time

t, and a vector of lagged controls $\mathbf{z}_{t-i}$, with the corresponding vector of coefficients $\boldsymbol{\gamma}_{i,h}$. In my baseline setting, I control for lags of monetary policy shocks and lags of deviations in WAM, by setting the lag length l to six months. Despite the plausible exogeneity of the selected monetary policy events, I include these dynamic regressors for WAM deviations and lagged shocks to absorb any serial correlation, as proposed in Ramey (2016). Note however, that Ramey (2016) further states that the Jordà (2005)-method does not necessitate the inclusion of control variables if the independent variable is credibly exogenous, as is the case in my application.[22] Since the effective degrees of freedom are already limited given the rather short sample and due to the inclusion of six lags in the dynamic regressions, I choose to omit additional control variables in the local projections to prevent a further loss in degrees of freedom. To increase the efficiency of my estimation, I follow Jordà (2005) and include u'_{t+h-1}, i.e., the residuals recovered from the previous estimation at $h-1$. I set h to a maximum of nine months in order to be able to reflect the maximum possible horizon of deviations from annual PDM plans (see Section 3.3.1 for a detailed description of institutional specifics). Due to the already short time dimension of the data, I refrain from further increasing the amount of horizons.

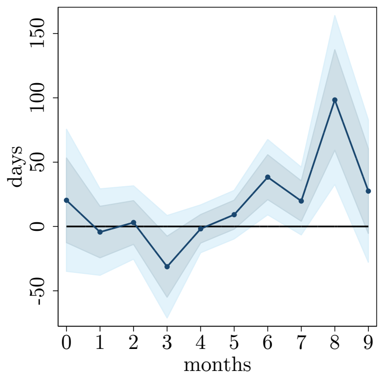
Figure 3.7: Dynamic response of PDM to unconventional monetary policy announcements



(a) SBP



(b) SBN



(c) Quantified

Notes: Dynamics obtained from local projections are plotted for a data sample covering 2014M1 to 2019M3. The horizontal axis measures time in months, whereas the vertical axis measures the deviation in actual from planned weighted average maturity $WAMD_t$ in days. The solid line represents the reaction of $WAMD_t$ over the horizon of 9 months after an unconventional monetary policy announcement in t took place. The three panels show the results for three different specifications of the unconventional monetary policy proxies, with the response to the narrower binary event series SBP_t in Panel (a), the wider binary event series SBN_t in Panel (b), and the quantified dummy series QSB_t in Panel (c). The dark to light shaded areas depict 68 and 90 percent confidence intervals.

[22] Two recent contributions that study conventional monetary policy surprises using local projections without additional controls are Coibion et al. (2017) and Lewis et al. (2019). Even more related, Meier and Reinelt (2020) show that the alternative inclusion and exclusion of control variables do not change results for local projections when studying unconventional monetary policy surprises.

Figure 3.7 plots the dynamic response of the deviation in the actual WAM of gross issuance as compared to the maturity structure set out in the annual plan for a horizon of 9 months after a standard unconventional monetary policy announcement takes place in $t = 0$, that is I study an expansionary QE event typical for the chosen sample. Due to the majority of events in the sample representing extensions to the PSPP, an expansive unconventional monetary policy shock in $t = 0$ is represented by a value of 1 in the binary series SBP_t in Panel (a) and SBN_t in Panel (b), whereas for the quantified event series QSB_t in Panel (c) I normalize the monetary policy surprise for to the negative value of the absolute mean of non-zero observations. The solid lines represent point estimates, and the dark to light shaded areas depict the 68 and 90 percent confidence intervals for each specification.

As expected and foreshadowed by the contemporaneous regression results in Section 3.4.1, the response in h = 0 is rather muted and stays insignificant for six months in the case of the SBP_t and QSB_t series, whereas for the SBN_t specification the estimation results become significant after four months already. The largest maturity extensions occur at $h = 6$ for both binary unconventional monetary policy proxies, and amount to 151 days for SBP_t and to 131 days for SBN_t. The QSB_t specification, by contrast, shows the most remarkable response in $WAMD_t$ eight months after a QE announcement took place; the WAM deviation in this case suggests a maturity extension of 98 days. To calculate the cumulative deviations in WAM for the three specifications, I set the month of the maximum response of QSB_t as the maximum horizon, i.e., $h = 8$, and obtain accumulated maturity extensions ranging between 150 and 450 days.

Overall, the causal evidence from the local projection setup fully aligns with the institutional characteristics of the German Finance Agency and the joint descriptive analysis of the data from Section 3.3 in pointing to a delayed maturity extension after stimulative unconventional monetary policy announcements.

3.5 Extensions and sensitivity

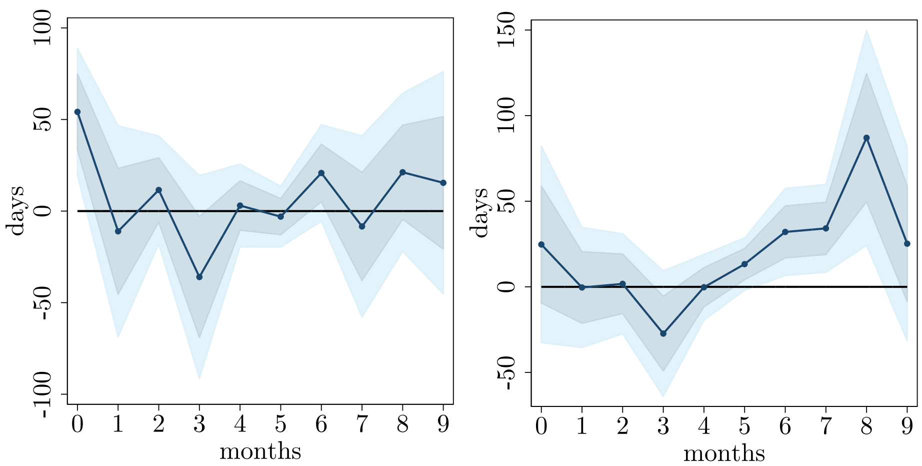
To verify the viability of the dynamic regression results, I confront my baseline estimations to a battery of sensitivity checks regarding the empirical specification and the identification strategy set out below.

3.5.1 Alternative unconventional monetary policy measures

In my three baseline specifications, I test the causal reaction of the WAM deviation to QE announcements specified by two binary SBP_t and SBN_t series and a quantified QSB_t event series. The unconventional monetary policy announcements reflected by the benchmark binary series are selected by means of qualitative evaluation of their content and their implied effect on the stock of German government bonds. Based on this setting, the quantified series is obtained by weighing the event dates in the binary series with the change in the 10-year Bund yield within the monetary policy event window as reported in the EA-MPD. To validate the robustness of my results to the chosen specifications of unconventional monetary policy announcements, I test alternative specifications of binary and quantified proxies in this Section.

First, I test my results by using the change in an alternative financial market variable on monetary policy meeting days provided in the EA-MPD to quantify the SBP_t and SBN_t series. In particular, I replace the 10-year Bund yield with the respective changes in 2-year and 10-year Overnight Index Swap (OIS) rates.[23]

Figure 3.8: Response to OIS rate-weighted unconventional monetary policy events.



(a) 2-year OIS rate (b) 10-year OIS rate

Notes: Dynamics obtained from local projections are plotted for a data sample covering 2014M1 to 2019M3. The horizontal axis measures time in months, whereas the vertical axis measures the deviation of actual from planned weighted average maturity $WAMD_t$ in days. The two panels show the results for two additional specifications of the quantified unconventional monetary policy proxy using changes in 2-year OIS rates in Panel (a) and using 10-year OIS rates in Panel (b) to weigh binary event series. The solid line represents the reaction of $WAMD_t$ over the horizon of 9 months after an unconventional monetary policy announcement. The dark to light shaded areas depict 68 and 90 percent confidence intervals.

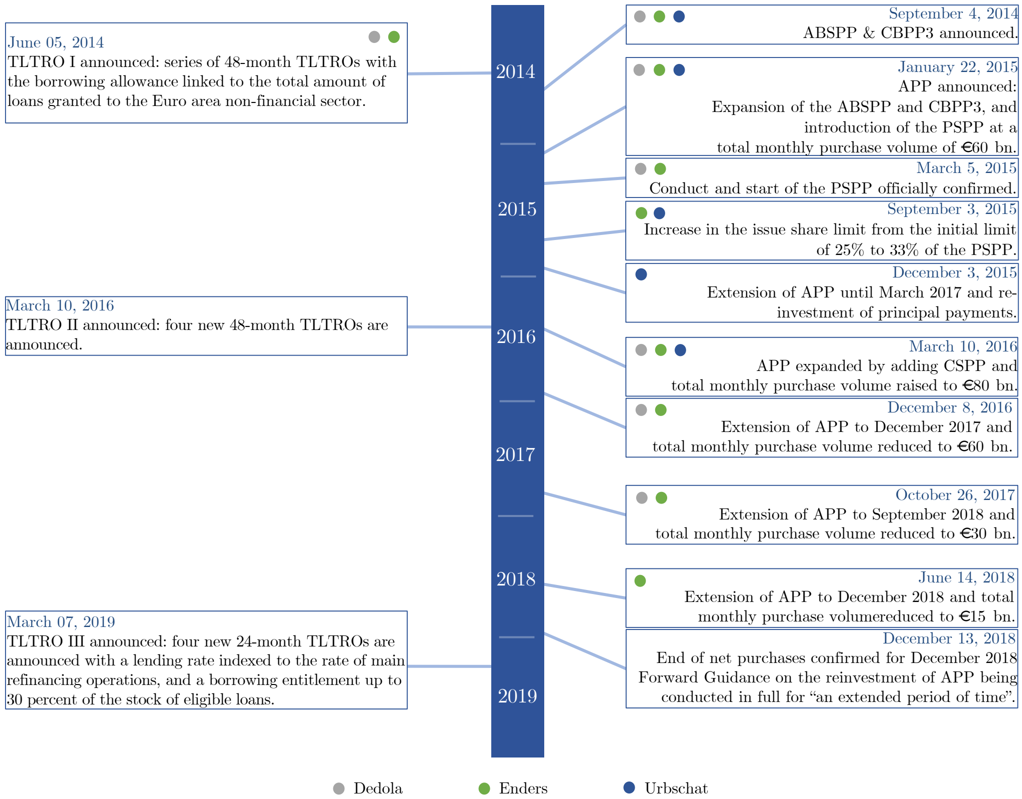
[23]Specifically, the OIS rate reflects the fixed rate offered in exchange for variable Euro Overnight Index Average (EONIA) rate payments over the specified maturity and hence reflects long-term interest rate expectations for a given maturity. Compared to Eurozone sovereign bond yields it thus abstracts from country-specific default risk and thus can be interpreted as the sum of expected short-term rates and the maturity-specific term premium.

Figure 3.8 shows deviations in maturity structure in response to a QE event quantified by the 2-year OIS rate and the 10-year OIS rate. While the results for the 2-year OIS rate are insignificant, the deviation in WAM in response to the event series weighted by changes in the 10-year OIS rate is qualitatively and quantitatively close to the benchmark evidence for the QSB_t series. The difference in results across OIS rates is not surprising, since QE announcements should primarily affect long-term yields (Altavilla et al., 2019) and hence may leave 2-year OIS rates largely unaffected. Importantly, the similarity in the response to the proxies weighted by the 10-year OIS and Bund yield suggests that effects on country-specific default risks of the selected unconventional monetary policy announcements are of second order, at least for maturity shifts in the German primary market. Instead, deviations in maturity appear to follow the immediate effect of QE on long-term interest rate expectations, in line with the signaling and portfolio rebalancing channel.

Second, I scrutinize the causal evidence by using a different selection strategy of unconventional monetary policy announcements to rule out that the specific choice of events is driving my results. In line with the empirically documented dominance of stock effects for the effectiveness of QE, the benchmark binary series reflect those announcements that plausibly affect the projected stock of sovereign bonds on the central bank balance sheet. As an alternative, I build a much more broadly defined event series, which contains all announcements conveying new information on the entire spectrum of unconventional monetary policy actions; that is, I include news about asset purchase programs, bank lending programs, and forward guidance. Figure 3.9 details the timing and content of the included announcements within the sample period between 2014M1 and 2019M3. As highlighted by the colored dots, this selection largely corresponds to the event series developed by Enders et al. (2019), Dedola and Georgiadis (2018), and Urbschat and Watzka (forthcoming) and extends their sample by a minimum of two observations.

From this broader set of events, I build a monthly binary event series that equals unity for those months, in which an announcement took place. Similar to the approach in specifying the benchmark QSB_t series, I obtain a quantified version of this extended event series by weighting the announcements for monetary policy meeting days with the immediate change in 10-year Bund yields and by accumulating them in case of multiple announcements within one month. Figure 3.10 plots the dynamic estimation results for these two alternative event series.

Figure 3.9: Timeline of events included in unconventional monetary policy proxies



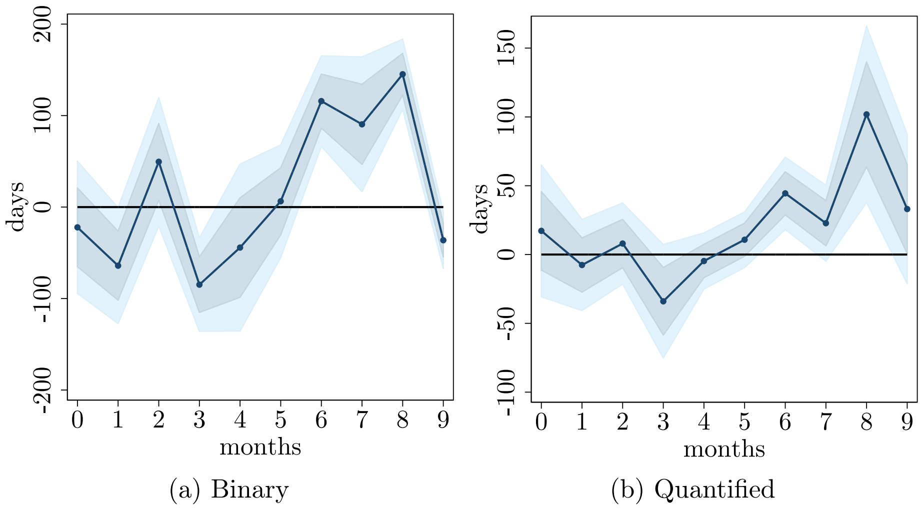
Notes: The events depicted are included to the broadly defined unconventional monetary policy proxy. The events on the right hand side of the timeline depict asset purchase programs and forward guidance events, whereas the events on the left hand side represent bank lending programs. The colored dots label those events that are included in recent event studies on unconventional monetary policy in the Eurozone by Dedola and Georgiadis (2018) (grey dot), Enders et al. (2019) (green dot) and Urbschat and Watzka (forthcoming) (blue dot). The abbreviations used stand for the following programs: Asset-Backed Security Purchase Programme (ABSPP), Covered Bond Purchase Programme 3 (CBPP3), Targeted Longer Term Refinancing Operation (TLTRO), Expanded Asset Purchase Programme (APP), and Public Sector Purchase Programme (PSPP).

In comparison to the benchmark specification, the results from the estimation using the binary and quantified series for the extended selection of unconventional monetary policy events shows a qualitatively similar propagation in deviations of WAM. However, due to the inclusion of announcements that are unrelated to sovereign bond markets, it is not surprising that the response in short-term maturity adjustments is more ambiguous in the binary specification and less significant for the quantified specification relative to the benchmark counterparts.

3.5.2 Modifications to the model specification

In the benchmark specifications, the vector of lagged controls $\mathbf{z}_{t-i}$, is computed for $i = 1, ..., l$ with $l = 6$ lags of both the unconventional monetary policy proxy MPE_t, and the WAM

Figure 3.10: Evidence from extended unconventional monetary policy proxies



Notes: Dynamics obtained from local projections are plotted for a data sample covering 2014M1 to 2019M3. The horizontal axis measures time in months, whereas the vertical axis measures the deviation of actual from planned weighted average maturity $WAMD_t$ in days. The two panels show the results for two extended specifications of the unconventional monetary policy proxies for event dates depicted in Figure 3.9. The response in $WAMD_t$ to innovations in the binary event series is depicted in Panel (a), whereas the quantified version is depicted in Panel (b). The solid lines represent $WAMD_t$ reactions over the horizon of 9 months after an unconventional monetary policy announcement took place. The dark to light shaded areas depict 68 and 90 percent confidence intervals.

deviation $WAMD_t$. In the following, I test the sensitivity of my results to changes in the lag order. For this purpose, I first re-estimate the results by reducing the maximum lag length l to 4 months and by increasing it to 8 months and report the outcome in Figure 3.11.

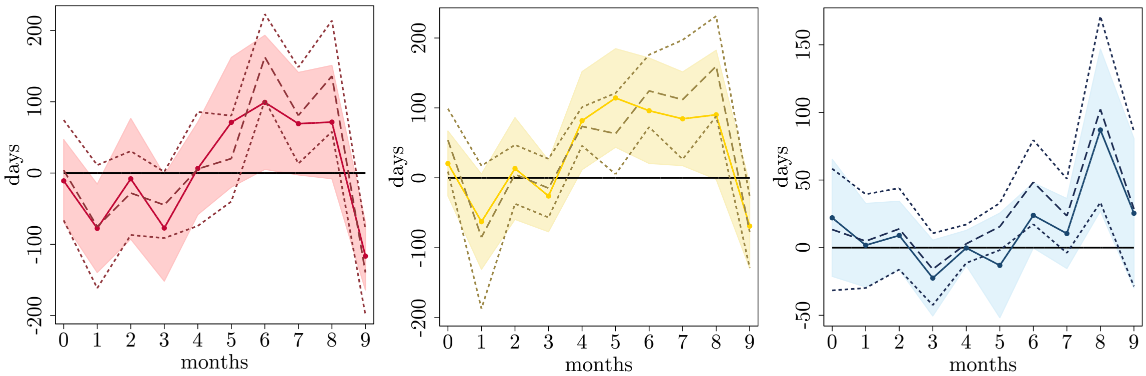
Second, I test the sensitivity to excluding lags of the event series and for the WAM deviations both separately and altogether. Figure 3.12 shows the results for specifications, in which I alternatively omit (i) lagged deviations of WAM, (ii) lagged unconventional monetary policy proxies, and (iii) lags for both variables altogether.

In total, the results of these extensions qualitatively match my benchmark results of a lagged positive deviation in the WAM after an unconventional monetary policy announcement of central bank actions on sovereign bond markets. Notably, among the alternative unconventional monetary policy proxies the QSB_t specification shows the least sensitivity to the performed extensions.

3.5.3 WAM responses to conventional monetary policy

Ultimately, to provide external validity for my empirical model, I test the reaction of unplanned deviations in the actual WAM of gross issuance to (i) conventional monetary

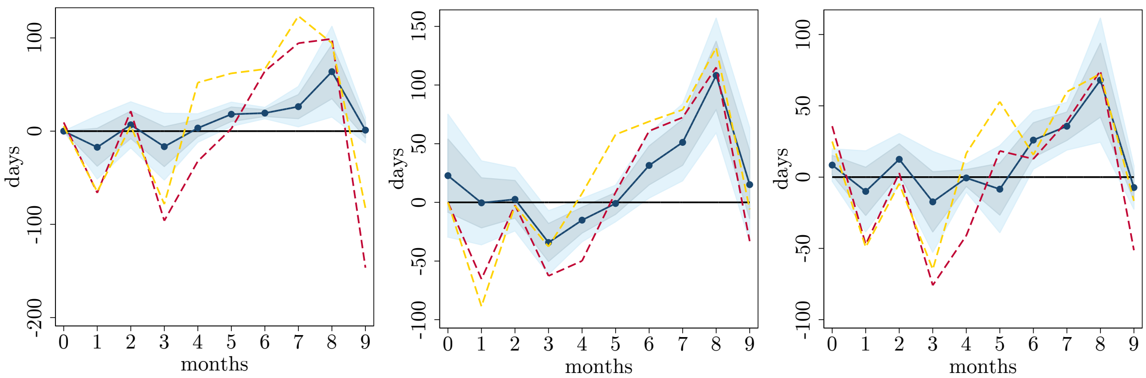
Figure 3.11: Evidence from changing the lag order



(a) SBP (b) SBN (c) QSB

Notes: Dynamics obtained from local projections are plotted for a data sample covering 2014M1 to 2019M3. The horizontal axis measures time in months, whereas the vertical axis measures the deviation of actual from planned weighted average maturity $WAMD_t$ in days. The three panels show the results for three different specifications of the unconventional monetary policy proxies, with the response to the narrower binary event series SBP_t in Panel (a), the wider binary event series SBN_t in Panel (b), and the quantified dummy series QSB_t in Panel (c). The point estimates and 90 percent confidence intervals for the deviation in actual from planned maturity structure $WAMD_t$ are represented by solid lines and shaded areas for the lag length $l = 4$ and by the darker dashed lines for $l = 8$.

Figure 3.12: Evidence from excluding lags entirely or exlusively for selected regressors

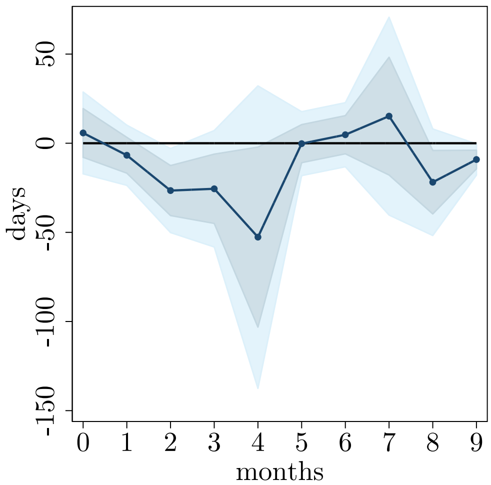


(a) no lags of $WAMD$ (b) no lags of MPE (c) no lags at all

Notes: Dynamics obtained from local projections are plotted for a data sample covering 2014M1 to 2019M3. The horizontal axis measures time in months, whereas the vertical axis measures the deviation of actual from planned weighted average maturity $WAMD_t$ in days. The three Panels show the results for the exclusion of lags from the dynamic model regression with respect to $WAMD$ in Panel (a) and to MPE in Panel (b). Panel (c) omits lags in the regression altogether. The point estimates as well as the 68 and 90 percent confidence intervals for the deviation in actual from planned maturity structure $WAMD_t$ in response to the quantified proxy for unconventional monetary policy QSB_t are represented by solid lines and dark to light shaded areas. The dashed lines represent the point estimates for the binary event series SBP_t (red dashed line) and SBN_t (yellow dashed line).

policy, (ii) in an earlier sample period. In line with the approach previously specified for unconventional monetary policy, I set a binary event series to 1 for those days on which a conventional monetary policy announcement implied changes to the ECB's key interest rate in a sample from 1999M1 to 2007M6.[24] In addition, I weigh these announcement days by the adjustment in the monetary event window for the shortest yield available, i.e., the 1-week OIS rate, obtained from the EA-MPD.[25]

Figure 3.13: Response to conventional monetary policy



Notes: The dynamic response in $WAMD_t$ obtained from local projections to an expansionary conventional monetary policy announcement weighted by the change in the 1-week OIS rate in the monetary event window is plotted for a data sample covering 1999M1 to 2007M6. The horizontal axis measures time in months, whereas the vertical axis measures the deviation of actual from planned weighted average maturity $WAMD_t$ in days. The solid lines represent $WAMD_t$ reactions over the horizon of 9 months after an unconventional monetary policy announcement took place. The dark to light shaded areas depict 68 and 90 percent confidence intervals.

The dynamic estimation results depicted in Figure 3.13 show the response in the deviation of maturity to an expansionary conventional monetary policy announcement. In line with previous specifications, the monetary policy surprise in $t = 0$ is normalized to the negative absolute mean of non-zero observations. Overall, Figure 3.13 shows that the causal effect on deviations in bond maturity is less significant and the range of variation is smaller compared to the benchmark specification for unconventional monetary policy events. The only significant point estimate is a negative unplanned reduction of the WAM in gross issuance of 25 days two months after the conventional monetary policy announcement. The peak response is a negative deviation of approximately 50 days taking place 4 months after the announcement. These results are in line with theoretical predictions

[24] Note, that for the purpose of comparability to my benchmark specification I consciously deviate from the common identification of exogenous conventional monetary policy decisions in the style of Romer and Romer (2004).

[25] I refrain from performing estimations for binary specifications of conventional monetary policy announcements since these can not reliably differentiate between expansionary or contractionary conventional adjustments to the policy stance in the sample. By contrast, the domination of expansionary shocks in the QE sample make the distortion in the binary setting less severe.

for expansionary conventional monetary policy, which is known to work predominantly at the short end of the yield curve (Altavilla et al., 2019). That is, by reducing the short end of the yield curve, monetary policy makes it more attractive for PDM to issue short-term debt. Beyond providing external validity to the previously obtained evidence on QE, the maturity shortening observed for conventional monetary policy easing further underlines the importance of the cost motive in the conduct of PDM, i.e., unplanned adjustments in the maturity structure appear to follow yield incentives in the respective maturity segment that conventional and unconventional monetary policy announcements typically affect.

3.6 Conclusion

Since the first indications of quantitative easing in the Eurozone in 2014, the maturity structure of German sovereign debt has extended notably, both in outstanding volumes and gross issuance. In this paper, I study the direct effects of unconventional monetary policy announcements in the Eurozone on decisions by German public debt management on the maturity structure. By using binary event series in line with event-study literature as well as quantified event series inspired by high-frequency literature to proxy for unconventional monetary policy announcements, I trace deviations in the actual from planned WAM of gross issuance in response to QE policies. For this purpose I, first, examine the course of QE events and PDM plan deviations in a joint descriptive analysis. This narrative results are suggestive to positive deviations in the maturity of gross issuance some quarters after an unconventional monetary policy announcement.

To provide insight on the causal influence of unconventional monetary policy on public debt management, I, second, perform local projections as proposed by Jordà (2005). Contrasting preceding studies that focus on the effects of *conventional* interest rate changes on PDM (Hoogduin et al., 2011; Blommestein and Turner, 2012), my empirical framework carefully deals with endogeneity issues by deploying plausibly exogenous monetary policy event proxies. In addition, I complement existing studies by showing that German PDM reacts to *unconventional* monetary policy announcements of central bank sovereign bond market actions in a significant and dynamic manner, in line with the expected yield curve effect. Across different empirical specifications, the maximum maturity extensions range between 98 and 151 days. In cumulative terms over a horizon of 8 months, the causal evidence even suggest maturity extensions between 150 and 450 days. These results are robust to model extensions and sensitivity tests. This evidence may serve to inform monetary policy makers in order to avoid an unintended weakening of unconventional monetary policy transmission via counteracting effects emerging from endogenous PDM reactions. Moreover, conditional on the provision of similar plan data by the respective debt management offices, a similar approach may be applied to study the connection of PDM and unconventional monetary policy for further member states of the European Monetary Union. With respect to Germany, the results of this study allow to ultimately conclude that, in times of QE, sovereign bond market related monetary policy actions and public debt management decisions are substantially intertwined, after all.

Chapter 4

How Do Income and the Debt Position of Households Propagate Public into Private Spending?[1]

[1] This Chapter is based on joint work with Sebastian Rüth. A recent version can be downloaded from the authors' webpages. A recent version of this paper appeared as (see Rüth and Simon, 2020).

4.1 Introduction

Is impacting after-tax income an important transmission mechanism by which increases in public spending stimulate private consumption expenditures and, thereby, the broader economy? A vast amount of macroeconomic time-series contributions have established the finding of systematic crowding-in of household spending—induced by unexpected variation in public spending—to be a salient feature in postwar U.S. data.[2] So far, however, the literature has not converged to a consensus when it comes to the underlying drivers of the conditional comovement of private and public spending. While the traditional Keynesian paradigm exactly predicts an *income-induced rise* of private consumption, the empirical regularity of consumption crowding-in poses a serious challenge for plain-vanilla New Keynesian and RBC models in which "throw-in-the-ocean" public spending induces negative wealth effects and causes optimizing households to substitute from consumption to labor supply.

Several routes have been taken to align these models with the conflicting evidence by introducing mechanisms that are designed to weaken the negative wealth effect, i.e., the reduction of present-value after-tax income, that public stimulus brings about for households.[3] Galí et al. (2007) propose a direct approach to make dynamics in sticky price (wage) models dependent on *current income* by adding Non-Ricardians, that is, households that do not hold physical capital and who consume their earnings in each period. Versions of this "limited asset market participation" or "Two-Agent New Keynesian" model have been commonly adopted in the literature.[4] Notably, "Heterogeneous-Agent New Keynesian" models (Kaplan et al., 2018) rely on richer representations of household finance, but generally feature consumers that, in equilibrium, behave in an income-constrained fashion (Bilbiie, 2019). Overall, the common ingredient to solve the consumption "crowding-in puzzle" across large parts of the fiscal policy literature can be summarized as: impacting income.

[2] See, e.g., Rotemberg and Woodford (1992), Fatás and Mihov (2001), Blanchard and Perotti (2002), Mountford and Uhlig (2009), Ravn et al. (2012), and Perotti (2014). Among contributions studying news shocks on military spending, i.e., *anticipated* variation in government spending, Fisher and Peters (2010), Forni and Gambetti (2016), and Ben Zeev and Pappa (2017) document crowding-in of consumption, while Ramey (2011) reports contractionary effects. Caldara and Kamps (2008) show that postwar U.S. data favor a conditional public-private-spending comovement, across identification schemes, once model specifications are harmonized.

[3] An incomplete selection includes, inter alia, deep-habits in consumption (Ravn et al., 2012), non-separable preferences between consumption and labor (Monacelli and Perotti, 2008), or useful and productive spending (Leeper et al., 2017; Sims and Wolff, 2018).

[4] See Corsetti et al. (2012), Drautzburg and Uhlig (2015), and McKay and Reis (2016) for academic approaches or the FRB/US model for a central bank adoption.

In this paper, we take one step back and revisit the role of disposable income as a propagation mechanism for public spending, directly in the data. Our approach consists of explicitly modeling post-tax income in structural VARs, thus imposing only few assumptions and structure to recover public spending shocks. As a point of departure, we center around an updated version of the recursively-identified (Blanchard and Perotti, 2002) VAR proposed by Galí et al. (2007), covering data from 1954Q1 to 2015Q4.[5] We identify plausibly unexpected variation in government spending by contemporaneously conditioning on a proxy for fiscal foresight and report that debt-financed fiscal stimulus jointly raises consumption along with post-tax income.

Nevertheless, this public spending induced consumption-income-comovement is at best suggestive to an income channel that may causally rationalize why consumption is not crowded out, contrary to Neoclassical models. Essentially, it is the *extra effect* that income adds to the response of consumption, which matters for this narrative. We approximate this extra effect of the income channel via a statistical decomposition. Following the procedure in Bachmann and Sims (2012), we neutralize the endogenous income response with counteracting, exogenous surprises in after-tax income. In the absence of income dynamics, the macroeconomic repercussions of fiscal stimulus are muted, but notably, household absorption still reveals an inverted hump-shaped adjustment pattern. In our preferred model specification, the income channel explains around one third of the reaction of consumption; yet, *crowding out* does not appear to be an empirical regularity, even when post-tax income remains hypothetically fixed.

The finding of households being capable of expanding consumption volumes—without supporting income—suggests that in addition to an intact income channel at work, households' debt position may adjust to finance the additional private spending; that is, *consumption crowding in may be reinforced by debt accumulation.* In this vein, Fernández-Villaverde (2010) shows how financial frictions may amplify the consequences of public stimulus. His model predicts that crowding-out of investment is counteracted once government spending propagates via imperfect financial markets (see Eggertsson and Krugman, 2012). While he emphasizes a (reverse) mechanism à la Fisher's debt deflation, Carrillo and Poilly (2013) stress that the propagation of stimulus through the value of firms' collateral may directly narrow credit spreads and support equilibrium debt. To the extent that households face

[5] We focus on postwar data and thus mainly on *civilian spending shocks.* This focus is consistent with Perotti (2014) and Nakamura and Steinsson (2014) who stress that samples covering the Korean war or WWII may cause identification problems due to, e.g., price controls or rationing.

similar credit market imperfections, and borrowing constraints are conditionally mitigated, an expansion of household debt is equally conceivable.[6]

We provide empirical aid to this proposition by carefully modeling the joint dynamics of households' credit conditions and the fiscal policy stance. First, we explore the conditional interplay of prices and quantities in credit markets, i.e., private interest rates, spreads and debt volumes, in fiscal policy VARs. Second, we note that the recursiveness assumption to recover the VAR's structural form is no longer warranted, once fast-moving financial variables are present (see Gertler and Karadi, 2015).[7] To separate fiscal policy surprises from systematic reactions to, e.g., the state of the financial system, we augment the VAR by an external instrument (SVAR-IV) for identification (Stock and Watson, 2012; Mertens and Ravn, 2013), allowing for simultaneous feedback to public spending from all variables in the system. Following Barro and Redlick (2011), we use changes in military spending as an instrument. Military spending exhibits substantial swings over time and plausibly exhibits comovement with the unobserved policy innovations of interest (instrument relevance). In addition, military expenditures regularly reflect developments abroad, e.g., geopolitical instability, such that variation induced by domestic economic or financial conditions appears to be meager, in relative terms (exclusion restriction).[8]

Our main observations for debt-augmented fiscal SVAR-IVs are twofold. First, unexpected variation of public spending causally induces proxies for household credit such as consumer credit, mortgage debt, or total household indebtedness to expand significantly and persistently—in excess to a continuing rise in post-tax income and consumption. This striking interaction of household debt and the fiscal policy stance aligns with Cloyne and Surico (2017), who track variation in U.K. tax rates, and with Bernardini and Peersman (2018), who stress private debt as a *state variable* for fiscal output multipliers.[9] Yet, both papers are silent on the role of private debt as a *transmission mechanism for public spending*. Second, the conditional debt cycle we document is paralleled by declining interest rates in credit markets. This finding corroborates Ramey (2011), who reports falling bond

[6]Miyamoto et al. (2019) propose that fiscal stimulus may relax credit market conditions via redistribution of income toward saving agents.

[7]Caldara and Kamps (2017) make the case that even with respect to aggregate activity there may exist some degree of within-quarter endogeneity of recursively-recovered fiscal policy innovations. For forward-looking financial time series, which are strongly correlated with and typically lead the business cycle, such concerns are likely to apply even more so.

[8]In fact, our instrument reveals close-to-zero and insignificant contemporaneous, unconditional correlation coefficients regarding changes in U.S. GDP or interest rates and spreads. To formally test instrument relevance, we rely on methodological progress of Montiel Olea et al. (forthcoming).

[9]Demyanyk et al. (2019) study consumer debt as a state variable for open-economy relative fiscal multipliers using geographical variation in U.S. defense spending during the Great Recession.

rates conditional on positive *news shocks* about military spending. However, neither does Ramey (2016) relate her finding to the debt position of households, as we do, nor does her VAR address the simultaneity problem between financial and macroeconomic variables.[10] Moreover, our findings reinforce Auerbach et al. (2020), who exploit geographical variation in U.S. federal contracts across U.S. cities and find positive effects on local credit markets in data starting at the millennium. Our time series approach puts their evidence into perspective and provides external validity by exploring the entire U.S. postwar history and by explicitly capturing general equilibrium effects.

Overall, the *conditional divergence of credit market volumes and prices* we observe supports the perception of reinforcing financial conditions strengthening the expansion of debt and, ultimately, the crowding in of consumption. To further rationalize our findings, we provide tentative evidence on the transmission mechanisms underlying our results. First, as we do not reveal inflationary pressure to be unleashed by the surge in public spending, we conjecture that the conditional debt cycle does not appear to be induced by Fisher effects (Fernández-Villaverde, 2010). Second and aligning with the documented price dynamics, we find no tightening of (real) policy or long-term risk-free rates; this absence of counteracting risk-free rates may thus contribute to the debt accumulation of households. Third, our SVAR-IV model reveals a significant compression of interest spreads in credit markets suggesting a softening of borrowing constraints, that is, easier access to credit for households (Auerbach et al., 2020). Fourth, this loosening of borrowing conditions is likely related to inflating asset prices: we document that public stimulus boosts collateral values such as real estate prices, which should positively impact on households' balance sheets and may reduce their (perceived) default probabilities (see Bernanke et al., 1999, for the related financial accelerator mechanism).

Taken together, our findings imply that income dynamics of households as suggested by the Keynesian framework are likely an important empirical moment to help expand our knowledge about the underlying drivers of the propagation of public into private spending. However, our causal evidence on credit-augmented fiscal policy VARs prompts the view that this transmission mechanism is complemented by the pass-through of stimulus into households' debt position as another vital mechanism to make progress toward that direction.

[10]This concern regarding narrative identification in general is also expressed by Ramey (2016). D'Alessandro et al. (2019) and Miranda-Pinto et al. (2019) also report declining interest rates conditional on surges in public spending, relying on exclusion restrictions that identify their VARs.

The remainder of the paper is as follows: Section 4.2 presents the empirical setting and evidence on income. Section 4.3 explores debt-augmented VARs, Section 4.4 provides tentative insight into the transmission mechanisms, and Section 4.5 concludes.

4.2 Empirical framework

In this Section, we provide a structural VAR framework that we use to study the propagation of government spending to the broader economy, in particular, to consumption and disposable income of households. First, we describe the data and the identification strategy; second, we propose a method to isolate the contribution of post-tax income in the shock pass-through; and third, we present empirical results.

4.2.1 Structural VAR representation

We postulate that the variables of interest can be cast in a finite-order linear VAR representation of the form:

$$\boldsymbol{A_0 x}_t = \sum_{i=1}^{l} \boldsymbol{A}_i \boldsymbol{x}_{t-i} + \boldsymbol{\varepsilon}_t, \text{ with } \mathbb{E}\{\boldsymbol{\varepsilon_t}\} = 0 \text{ and } \mathbb{E}\{\boldsymbol{\varepsilon_t \varepsilon_t}'\} = \boldsymbol{\Sigma_\varepsilon}, \tag{4.1}$$

abstracting from the intercept for notational convenience. At lag, $i = 1, ..., l$, the $n \times n$ matrix $\boldsymbol{A}_i$ comprises the model's dynamics, and $\boldsymbol{A}_0$ captures contemporaneous relations. $\boldsymbol{\varepsilon}_t$ represents mutually uncorrelated innovations, i.e., $\boldsymbol{\Sigma_\varepsilon}$ is diagonal. We identify a government spending shock by a Cholesky-factorization of the VAR's reduced-form variance-covariance matrix, $\boldsymbol{\Omega}$, assuming government spending to be pre-determined with respect to within-quarter macroeconomic conditions (Blanchard and Perotti, 2002). For direct comparability with the VAR evidence in Galí et al. (2007), we stick to this strategy as the basic structure for identification, but relax the identification assumption in Section 4.3. However, such a purely recursive approach to recover the VAR's structural form may be subject to fiscal foresight concerns, i.e., the structural innovations may be forecastable, to some extent. Therefore, we recover plausibly unanticipated innovations in the fiscal policy stance by simultaneously conditioning public spending in the structural VAR on measures of fiscal news. As a baseline, we follow Bachmann and Sims (2012) and augment the VAR by military news, g_t^m, as proposed in Ramey (2011), ordered first in $\boldsymbol{x}_t$, followed by government spending (consumption plus investment), g_t, ordered second. Results are similar when conditioning on proxies as provided in Fisher and Peters (2010) or Ben Zeev and Pappa (2017), or when omitting news alltogether.

The remaining $j = 3, ..., n$ entries in vector $\boldsymbol{x}_t$ comprise the variables of the Galí et al. (2007) VAR, in this order: personal disposable income, y_t^{dis}, GDP, y_t, hours worked, h_t, consumption (non-durables and services), c_t, non-residential investment, i_t, wages, w_t, and the budget deficit, d_t.[11] The after-tax income series we explore is extensive, comprising the following sources of income: U.S. residents' labor income, employer-provided supplements such as insurance, income from owning a business or from rental property, benefits from social security, interest income, and dividends; the series explicitly excludes valuation effects stemming from asset price movements.

We estimate the VAR over a sample beginning in 1954Q1 and extending to 2015Q4, including four lags of the vector of quarterly observables, i.e., $l = 4$. The start of the sample is motivated by, inter alia, Perotti (2014) and Nakamura and Steinsson (2014), who argue that including the WWII era may cause identification problems as the influence of interfering factors like price and production controls, rationing, or the military draft are hard to assess. Similarly, during the Korean war, results may be contaminated by new Fed regulations at the time. The impact of these events is particularly hard to gauge in military spending news identification approaches as, e.g., in Ramey (2011), in which results are mainly driven by the defense spending shocks during wars, whereas in the sample starting in 1954Q1 civilian spending shocks prevail. For the benchmark model, we end the sample in 2015Q4 due to the availability of proxies for fiscal foresight. We find throughout consistent results in estimations that omit the Great Recession episode; and we explore additional sub-sample sensitivity in Section 4.2.3.

4.2.2 A systematic analysis of the disposable income channel in a VAR

In this Section, we propose a procedure to statistically single out and *approximate the marginal effect* that impacting post-tax income adds to the response of consumption, within the structural VAR framework above. In this vein, the original argument put forth in, e.g., Galí et al. (2007) to call for an immediate income mechanism in New Keynesian models was based, among others, on the empirical comovement of private consumption and income after fiscal stimulus. Since this justification builds upon a conditional consumption income

[11]The budget deficit enters as a ratio to trend GDP, proxied by lagged potential output; all measures enter at the quarterly frequency in real terms. Quantity series are population normalized. Except for the budget deficit, which we measure in percent, time series enter in log levels.

correlation, we attempt to support this view by giving perspective on a causal relation, in a statistical sense.

Our strategy consists of statistically decomposing the repercussions of public spending into those effects arising from the endogenous response of disposable income and those observed after fixing this transmission variable to its pre-shock path.[12] Specifically, we quantify the so-defined marginal effect of income by positing hypothetical sequences of exogenous, Cholesky-factor-orthogonalized innovations in disposable income. We calculate the latter innovations such as to neutralize the endogenous response of post-tax income, conditional on the initial government spending surprise. Contrasting the benchmark response of consumption with the corresponding response absent movements in income, allows to infer the quantitative importance of the income channel for the projected path of private consumption. By exactly canceling out the endogenous response of post-tax income, the exercise is capable of capturing rich effects, since the entire spectrum of directly and indirectly operating effects stemming from the income variable are shut down. Consequently, we view the corresponding results as an upper bound approximation of an independently operating income channel.

Importantly, this decomposition is purely statistical in nature. For this reason, we do not assign any economic interpretation to the disposable income surprises we generate. By ordering personal disposable income after public spending in the VAR model, income surprises are allowed to contemporaneously pass-through to all variables in the system, except for government spending (and fiscal news).[13]

In a recursive setting, we compute the innovations to variable $y_t^{dis} \equiv \eta$, that are necessary to force the respective endogenous response to zero, as follows:

$$\boldsymbol{\varepsilon}_{\eta,h} = -\sum_{j=1}^{n} \boldsymbol{\Theta}_{\eta,j} \boldsymbol{y}_{j,h} - \sum_{m=1}^{\min(l,h)} \sum_{j=1}^{n} \boldsymbol{\Theta}_{\eta,mn+j} \boldsymbol{z}_{j,h-m}. \tag{4.2}$$

$\boldsymbol{y}_{j,0}$ denotes the $t = 0$ effect of a spending shock on variable j, whereas the same effect sans endogenous response of income reads: $\boldsymbol{z}_{j,0} = \boldsymbol{y}_{j,0} + \boldsymbol{\Phi}_{j,\eta,0} \boldsymbol{\varepsilon}_{\eta,0}/\sigma_\eta$, where $\boldsymbol{\Phi}_{j,\eta,0}$ is the

[12] Bachmann and Sims (2012) revitalize this method, which was pioneered by Bernanke et al. (1997 (1), and apply it to the reaction of consumer confidence in their Cholesky-identified fiscal VAR. Recent applications include Bachmann and Rüth (2020).

[13] In the model of Galí et al. (2007), disposable income was ordered last in the VAR. Note however, that the position of income does only matter for the reduced-form disposable income surprises, but does not have any statistical impact on the dynamics triggered by the fiscal policy shock. As a matter of fact, the ordering within the subset of variables $j = 3, ..., n$ is orthogonal to the results for the public stimulus shock, with public spending ordered second, i.e., $j = 2$.

$\{j, \eta\}$ element of the impulse response matrix for horizon $h = 0$. The standard deviation of income disturbances is σ_η, and for horizons $h > 0$ we calculate:

$$\boldsymbol{y}_{j,h} = \sum_{m=1}^{\min(l,h)} \sum_{i=1}^{n} \boldsymbol{\Theta}_{j,mn+i} \boldsymbol{z}_{j,h-m} + \sum_{i<j}^{n} \boldsymbol{\Theta}_{j,i} \boldsymbol{y}_{i,h}, \tag{4.3}$$

and ultimately:

$$\boldsymbol{z}_{j,h} = \boldsymbol{y}_{j,h} + \frac{\boldsymbol{\Phi}_{j,\eta,0} \varepsilon_{\eta,h}}{\sigma_\eta}. \tag{4.4}$$

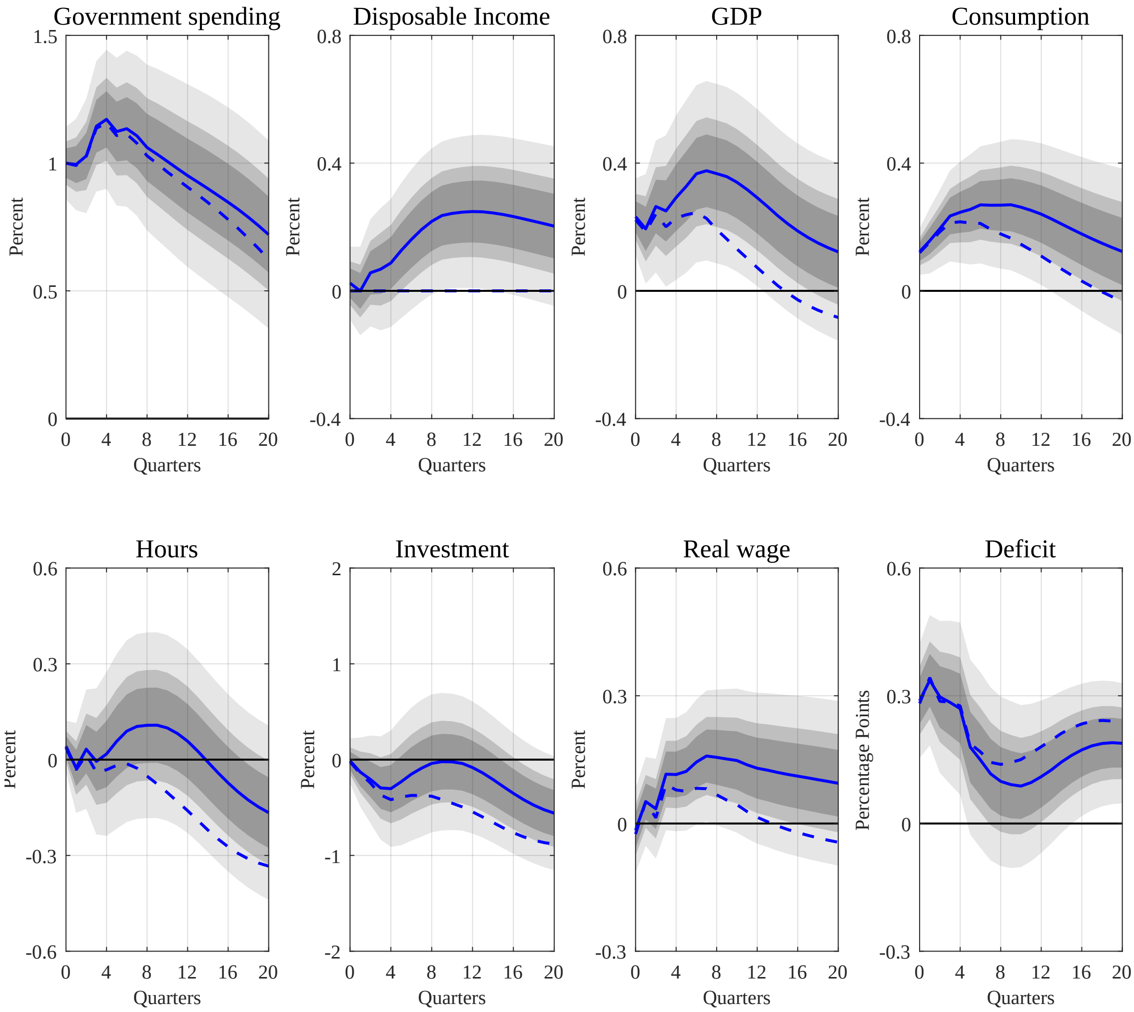
4.2.3 An empirical perspective on the income channel of public stimulus

Ignoring the dashed lines for the moment, Figure 4.1 traces VAR dynamics (solid blue lines) conditional on an expansionary government spending surprise, for quarterly U.S. data ranging from 1954Q1 to 2015Q4. Dark to light shaded areas depict 50, 68, and 90 percent confidence intervals, obtained from a bootstrapping procedure (Goncalves and Kilian, 2004). We normalize the shock size such as to move public spending by one percent away from its pre-shock path, following convention and for comparability across different specifications. Despite the fact that we account for fiscal foresight and use a longer sample relative to Galí et al. (2007), whose data end in 2003Q4, we report fully consistent dynamics. Consumption rises sluggishly for two years, before slowly abating and the response is significantly positive for more than three years; that is, we observe the consumption crowding-in puzzle. GDP mimics the consumption response qualitatively and peaks at almost 0.4 percent. The budget deficit increases on impact and reaches the maximum response of 0.3 percentage points shortly after the impact period, i.e., we study a deficit-financed public spending innovation. In addition, we reveal Keynesian dynamics by documenting procyclically responding wages and hours, over the medium run.[14] Investment responds negatively to stimulus, but insignificantly so.[15] In line with the labor market dynamics, post-tax income is sticky in the first year after the shock, but subsequently reveals a protracted, inverted hump-shaped impulse response, which deviates by more than 0.2 percent from its conditional mean, and which remains different from zero in a statistically significant sense for roughly two years.

[14]Note that the positively reacting real wage that we document empirically is typically also the key ingredient within several New Keynesian approaches that aim to strengthen the income channel and, ultimately, attempt to rationalize consumption crowding-in.

[15]Blanchard and Perotti (2002) and Mountford and Uhlig (2009) report an investment decline, while Fatás and Mihov (2001) document an increase.

Figure 4.1: Conditional dynamics of government spending shocks



Notes: We plot dynamics at the quarterly frequency. Solid blue lines represent point estimates of impulse response functions. Dark to light shaded areas display 50, 68, and 90 percent confidence intervals, which we obtain from 1,000 replications of a recursive-design wild bootstrap procedure. Dashed blue lines denote results for a fixed disposable income experiment, along the lines of Bachmann and Sims (2012).

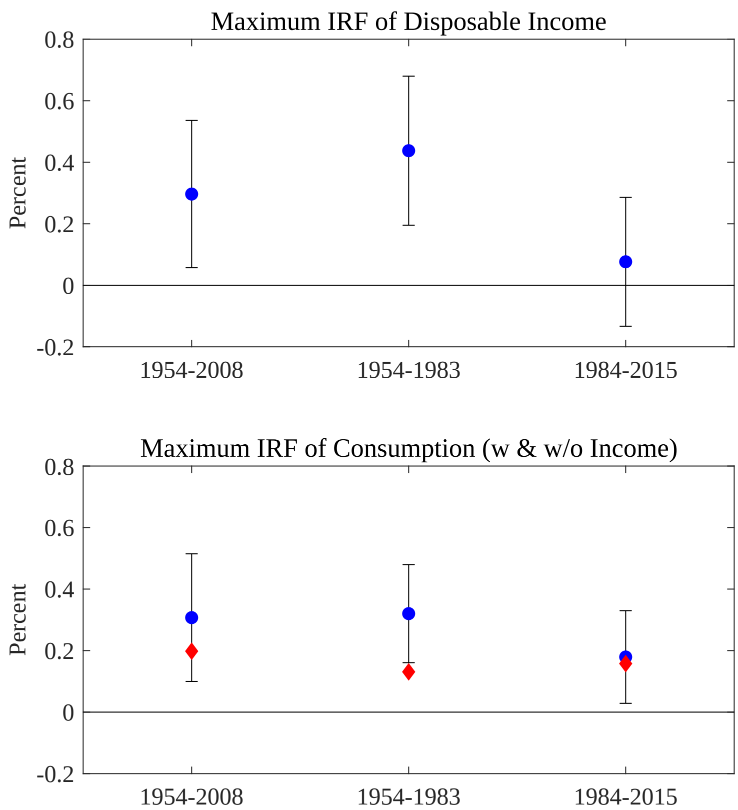
The dashed blue lines in Figure 4.1 contrast the benchmark VAR's impulse responses (solid blue lines) with corresponding dynamics observed for a fixed-income scenario. As the solid and dashed lines in the government spending panel are very similar, zeroing out the disposable income response does barely affect the systematic reaction of public spending to the exogenous shock, i.e., we essentially study the same fiscal stimulus hitting the economy. Post-tax income does, per definition, not react in the fixed-income experiment. The remaining variables' impulse responses in the fixed-income scenario closely track the responses of the benchmark model at short horizons, which is consistent with the protracted reaction of post-tax income in the unrestricted case. Over medium horizons, however, the income channel appears to become operative. The zeroing out of income makes the GDP reaction more short-lived and mutes its maximum response by roughly one third. Wages become less procyclical in the shock propagation and investment as well as hours worked process the shock via declining impulse responses. The public deficit is somewhat amplified, which—given an almost unchanged path of public spending—aligns with the documented lower aggregate activity and thus lower tax revenues. Most importantly, household absorption behaves qualitatively similar to the GDP response; that is, the maximum increase is mitigated and the impulse response returns faster toward its conditional mean. In accumulated terms, the fixed-income scenario predicts a surge in consumption that is roughly one third smaller relative to the unrestricted VAR case. Yet, even when we shut down the income channel, there is no evidence of consumption crowding out, in contrast to what Neoclassical theory would predict. Put differently, we infer that the conditional dynamics of disposable income are not sufficient to rationalize why consumption is crowded in by public spending surprises.

In what follows, we scrutinize this finding for different samples and for disaggregate measures of post-tax income of households in order to learn more about the structural properties of our result.

The income channel during different postwar episodes Given the substantial time series dimension we have exploited so far, it is instructive to inspect the stability of our results across different postwar episodes. For the different sub-samples we consider in what follows, Figure 4.2 summarizes the *maximum* impulse response coefficients (blue circles), along with 90 percent confidence intervals (black lines), for income (upper panel) and consumption (lower panel) to a one percent government spending expansion. In the Figure's lower panel, the red diamonds further present the corresponding reaction of consumption in a fixed income scenario.

A natural sample modification involves the omission of the Great Recession during which, inter alia, the effective lower bound on short-term nominal interest rates became binding. We thus follow the dating convention in Ramey and Zubairy (2018), who however do not document significant non-linearities stemming from the zero lower bound, and end the sample in 2008Q3. Our result of a positive and significant shock procession of disposable income and consumption appears to be insensitive to the exclusion of the Great Recession period (utmost left element of Figure 4.2); in addition, the consumption response in the fixed income scenario is muted by roughly one third, but still positive as in the benchmark sample. Another potential source of parameter instability may be the transition from macroeconomic turbulence witnessed in the U.S. during, e.g., the '70s, to a more tranquil episode starting in the '80s that was subsequently coined the Great Moderation episode. Following, among others, Gambetti and Galí (2009) we split the sample into pre- and post-1984 data. The second and third element in Figure 4.2 depict the corresponding results. For pre-Great-Moderation data, the reaction of household consumption is similar to the baseline model, yet estimated with somewhat higher precision, while the income response is magnified. Interestingly, the marginal effect that after-tax income adds to the consumption response is substantial such that the consumption reactions with and without hypothetically fixed income are statistically different from each other.

Figure 4.2: Consumption and income dynamics in sub-samples



Notes: Blue circles denote the maximum reaction of, respectively, disposable income (upper panel) and private consumption (lower panel) to a one percent increase in government spending; the black lines depict 90 percent confidence intervals. The red diamonds in the lower panel measure the consumption reaction for a scenario in which the income response remains hypothetically fixed. The first four estimates in each panel report results for sample splits as denoted on the abscissa.

For the episode starting in 1984 the consumption reaction is somewhat muted, but still significantly positive, whereas this is no longer true for disposable income; its maximum impulse response is statistically not different from zero, and the extra effect that income adds to the consumption response small. This latter finding could loosely be interpreted as constituting an *unrestricted counterfactual* answering the question on how consumption responds to fiscal stimulus without supporting income dynamics since the Great Moderation. The answer is: while the consumption response is much smaller, it is still positive and consumption crowding in appears to remain an empirical regularity. Consequently, the question on the existence of an additional channel to rationalize the empirical puzzle of consumption crowding in arises, since the income channel does not seem to explain crowding in alone.

4.3 Fiscal stimulus, consumption, and household debt

How is it possible that consumption responds positively to stimulus programs if income can not (fully) explain this reaction? Although our VAR framework, by including the budget deficit, accounts for the prominently discussed role of *public* sector debt (e.g., Reinhart and Rogoff, 2010; Reinhart et al., 2012) the model is, up to this point, silent on the role of *private* sector leverage. In the following, we test whether conditional variation in households' indebtedness constitutes a further transmission mechanism of fiscal policy in a structural VAR, which may offer a path to structurally corroborate our findings. The notion is that in excess to supporting post-tax income, consumers raise their debt position to finance the observed expansion in household absorption.

By explicitly testing the hypothesis of stimulus propagating via household debt, we link our evidence on consumption crowding in to the literature documenting large macroeconomic repercussions emerging from changes in household balance sheets or bank credit growth (e.g., Schularick and Taylor, 2012; Mian et al., 2017). Quite striking movements of private debt can be identified during the U.S. mortgage cycle, which marked its peak prior to the Great Recession. We explicitly capture this boom and bust episode, which was paralleled by substantial swings in the fiscal policy stance, in our VAR. In doing so, we put the literature on autonomous credit variation into perspective by analyzing systematic reactions of private debt, conditional on shifts in public spending. Our approach is therefore also closely related to Bernardini and Peersman (2018), who analyze how deviations of domestic non-financial private sector debt-to-GDP ratios from their trend path affect the fiscal output-multiplier in historical U.S. data as a state—yet not as a transmission—variable (see Demyanyk et al., 2019). Cloyne and Surico (2017) use U.K. survey data and document the consumption response to variation in income taxes to be more pronounced for U.K. mortgagors relative to outright home owners; thus stressing the role of private debt as a propagation mechanism for fiscal policy, as we do.

We proceed by (i) proposing extensions to the reduced-form fiscal VAR that allow us to study credit conditions, (ii) discussing adjustments to recover the model's structural form in an attempt to make identification more credible in a macro-financial setting, and (iii) presenting the corresponding findings, before (iv) providing sensitivity analysis within the SVAR-IV framework.

4.3.1 Modeling the interplay of credit conditions and the fiscal policy stance

To model the joint dynamics of the fiscal policy stance and of fluctuations in private debt markets, we add to the fiscal VAR representation measures of prices and quantities of credit that are relevant for households. In terms of credit volumes, we rely on the subsequent debt stocks from the U.S. Flow of Funds database, which enter the VAR in logged, GDP deflator-normalized, seasonally-adjusted, per-capita terms: overall consumer credit granted by banks, the volume of outstanding home mortgage contracts, and overall household indebtedness. In terms of prices for credit, we are not aware of a consistent and consecutive series on lending rates for U.S. households over our sample period that is available at the quarterly frequency. Thus we proxy overall household borrowing conditions by Moody's Baa corporate bond yields, as in Bachmann and Rüth (2020), and further study, more household-specific, Federal-Housing-Agency-provided mortgage interest rates, as observed in secondary markets. Mortgage interest rates are, however, only available from 1964.

Against the backdrop of our VAR's rich specification, comprising four lags and nine variables (including fiscal news for which we had not reported IRFs), we rotate one pair of credit market quantities and prices jointly into the VAR, once at a time. Given the typically insignificant nature of the response of hours worked in our VAR, we only proxy labor market conditions in these credit-augmented fiscal VARs by the real wage and abstract from dynamics in hours worked in the estimations for parsimony; that is, the credit-augmented VARs consist of ten variables, measured at the quarterly frequency.[16]

[16]Note that results are insensitive to maintaining hours in the VAR as an eleventh variable.

4.3.2 Identifying fiscal policy shocks in the presence of financial variables

To obtain the unobservable government spending shocks from Equation 4.1, i.e., to make a structural analysis feasible in our baseline model, we have recovered the parameters in $\mathbf{A_0}$ by a Cholesky factorization of the reduced-form variance-covariance matrix. While this approach of imposing timing restrictions appears to be plausible as long as our aim is to orthogonalize shifts in the fiscal policy stance from systematic reactions to the macroeconomic environment, such contemporaneous zero restrictions are hard to defend in the presence of fast-moving and forward-looking financial time series. This concern of simultaneity has been acknowledged and addressed by, e.g., Gertler and Karadi (2015) and Caldara and Herbst (2019), for the case of *monetary* policy shocks.[17] We borrow from this strand of literature and tackle the identification challenge, for the case of *fiscal* policy shocks, by employing information from outside of the VAR, i.e., an external instrument. This SVAR-IV methodology allows to recover the unexpected innovations in public spending without necessitating exclusion restrictions on the contemporaneous relations in the model. By contrast, the identifying information can be obtained from an external proxy series that correlates with the government spending shock, but is *contemporaneously* uncorrelated with the remaining shocks in the system (see Caldara and Kamps, 2017, for a related strategy involving non-fiscal proxies). Conditional on the discretionary selection by the researcher of such a proxy, the identification is data-determined and the parameters in $\mathbf{A_0}$ can be recovered, even if the VAR comprises macro-financial linkages. In what follows, we characterize the instrument we will consider, and refer to Stock and Watson (2012) and Mertens and Ravn (2013) for details on the implementation of the SVAR-IV.

Following Barro and Redlick (2011), we employ changes in *actual military spending* relative to lagged real GDP as an external instrument.[18] In our application, by contrast, we use this external information to recover the structural parameters of a fully-specified VAR

[17]Moreover, Caldara and Kamps (2017) argue that recursively-recovered fiscal policy shocks may suffer from similar simultaneity problems regarding contemporaneous fluctuations in economic activity. Such reservation may be particularly valid for financial variables, which typically reveal strong leading properties for business cycle movements.

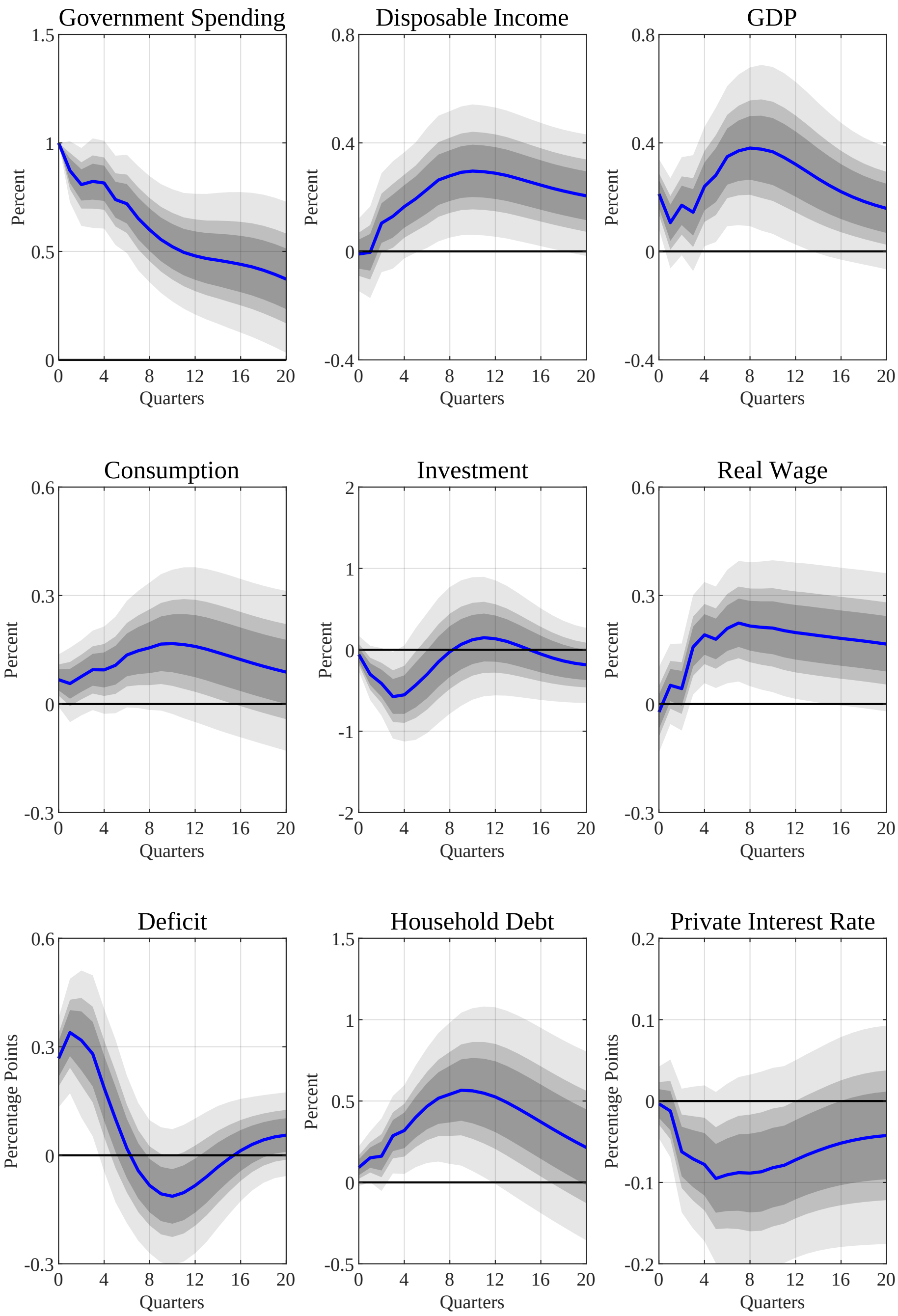
[18]Using *news* about military spending, which foreshadow public spending materializing in the future, is not a viable strategy since we are interested in unexpected shifts in the fiscal policy stance. Moreover, news series are known to be a weak instrument for unexpected short-run variation in public spending (Ramey and Zubairy, 2018). Note, however, that while this variable will not serve as the primary source of information for identification, our VAR will still account for news on military spending in an attempt to tackle fiscal foresight concerns.

model and to trace dynamic responses over time (see Miyamoto et al., 2019, for a local projections approach using military spending as an instrument). Why should such variation in military expenditures serve as an external instrument? The answer is: because it plausibly meets the so-called relevance and exclusion restrictions of a proper instrument. With regard to instrument relevance, military spending displays substantial swings in the sample and contributes substantially to the variability in overall fiscal spending; more importantly, it plausibly exhibits comovement with the unobserved policy innovations of interest in $\boldsymbol{\varepsilon_t}$. In fact, when we calculate the correlation of the military spending instrument with the Cholesky-identified shocks from Section 4.2, we obtain significantly positive coefficients. To test the relevance of the instrument more formally and to construct consistent confidence intervals in the SVAR-IV setting, we follow Montiel Olea et al. (forthcoming). In terms of the exclusion restriction, military expenditures are known to be regularly driven by conditions abroad, especially geopolitical instability such as events in Middle East. Consequently, shifts in military expenditures that reflect domestic economic or financial conditions are plausibly not an important driver. This proposition can be corroborated by calculating correlations of the instrument with changes in GDP or interest rates/spreads. Throughout, these coefficients are estimated close to zero and statistically insignificantly different from zero. In addition, we propose alternatives to the baseline instrument for robustness in Section 4.3.4.

4.3.3 The interplay of the fiscal policy stance and household credit

In a first step, we formally test the strength of changes in military spending as an instrument. To do so, we calculate a Wald statistic under the null hypothesis that the instrument is irrelevant, i.e., that it does not correlate with the unobserved government spending innovations. Following the methodology of Montiel Olea et al. (forthcoming), we estimate a Wald statistic of 54.3, which remarkably crosses critical values to reject the Null of a weak instrument, at conventional significance levels. Thus, changes in military spending represent a “strong” instrument for identification. Figure 4.3 traces adjustment patterns for the credit-augmented fiscal VAR, conditional on an unexpected (one percent) surge of public spending that we identify by changes in military spending as an external instrument; the selected measures of credit market volumes and prices are overall household debt and non-mortgage bond yields.

Figure 4.3: SVAR-IV public spending shock and credit markets



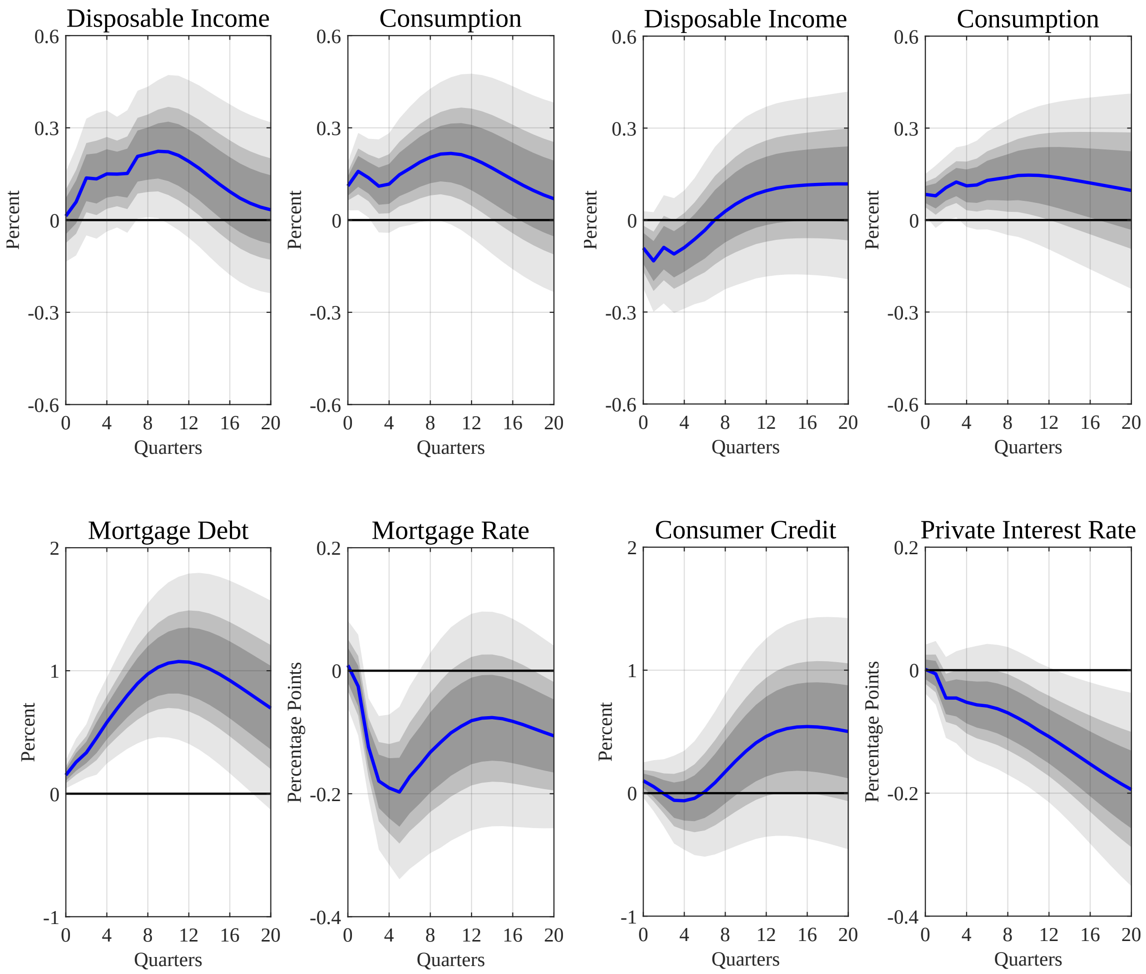
Notes: We plot dynamics at the quarterly frequency. Solid blue lines represent point estimates of impulse response functions obtained from the SVAR-IV model. Dark to light shaded areas display 50, 68, and 90 percent confidence intervals, which we obtain from the weak-instrument-robust procedure in Montiel Olea et al. (forthcoming).

Despite the fact that government spending processes the exogenous shock in a more immediate way, i.e., the impulse response moves faster toward the conditional mean relative to Figure 4.1 in which the response was hump-shaped, the overall adjustment patterns of the core set of variables are consistent with our main results up to this point. In particular, public stimulus induces a surge in private consumption expenditures, which is paralleled by persistent improvements in post-tax income. Reinforcing the empirical literature on consumption crowding in, we thus provide evidence that the conditional rise in private spending is a robust empirical regularity, even in a setting that does not require contemporaneous zero restrictions. In addition, this observation validates our former results from a Cholesky scheme that had disregarded the critique of Caldara and Kamps (2017) of existing within quarter feedback from economic activity to public spending. As an additional novel finding, we add to the literature by reporting that fiscal spending significantly propagates through the debt position of households: overall indebtedness increases on impact, slowly builds up until the third post-shock year—reaching its peak response around 0.6 percent above the pre-shock trend path—before slowly reverting back toward zero.

At the same time, borrowing conditions for households, roughly proxied by long-term corporate bond yields due to the lack of continuous and sufficiently long data on household credit rates, appear to soften. Private interest rates, namely, mirror the reaction of household debt qualitatively, i.e., we observe a hump-shaped decline in credit rates, which trough around roughly minus ten basis points, one year after the shock has hit. Related, Ramey (2011) reveals consistent findings in a recursive VAR that she uses to recover military news shocks. Our aggregate evidence on the conditional response of credit markets further aligns with Auerbach et al. (2020), who study local credit markets employing geographical variation in U.S. federal government contracts across U.S. cities.

In Figure 4.4, we zoom into the components driving these results in more detail, first, by re-estimating the SVAR-IV model using the sub-component of mortgage debt along with mortgage interest rates and, second, by including the consumer credit component along with the benchmark interest rate series. We restrict the presentation to the core set of variables of interest for the sake of a more parsimonious illustration. Panel (a) reveals that the importance of household leverage as an endogenous propagation mechanism of fiscal stimulus is even more sizable when focusing on *mortgage* indebtedness, which is the major component of overall household debt (accounting for roughly 75 percent of household debt

Figure 4.4: Mortgage debt and mortgage interest & consumer credit and interest rate



(a) Mortgage credit volumes/prices

(b) Consumer credit volumes/prices

Notes: We plot dynamics at the quarterly frequency. Solid blue lines represent point estimates of impulse response functions obtained from the SVAR-IV model. Dark to light shaded areas display 50, 68, and 90 percent confidence intervals, which we obtain from the weak-instrument robust procedure in Montiel Olea et al. (forthcoming). Panel (a) on the left displays IRFs for a fiscal VAR that is augmented by mortgage debt and mortgage interest rates; Panel (b) on the right presents IRFs for a fiscal VAR that is augmented by consumer credit and corporate bond yields.

during the last ten years of our sample). The magnitude of the mortgage debt reaction exceeds the counterpart reaction of overall debt by increasing around 1.1 percent. In addition, the impulse response is statistically different from zero almost throughout the entire forecast horizon. Correspondingly, mortgage interest rates process the shock more strongly, as well, and decline by roughly 20 basis points. In Panel (b), the according adjustment patterns for consumer credit are consistent, albeit less pronounced. The impulse response is sticky at short horizons, smaller in absolute magnitude, and estimated with less precision. The interest rate response appears more protracted in this scenario, too.

The impulse response function of private consumption, however, is fairly insensitive to the inclusion of consumer credit, revealing throughout positive coefficients.

These observations corroborate our inference that increases in disposable income of households are not the only transmission channel for the observed consumption crowding in following public stimulus. In particular, we show that there exists a leverage that public spending exerts on the debt position of households, raising the latter between 0.5 and one percent, depending on the empirical specification. This finding is remarkable as it implies that, contrary to state-of-the-art limited asset participation models, consumers without access to capital markets can not be the only explanation to rationalize crowding in of private spending. By contrast, the result of surges in household debt implies that also intertemporally optimizing consumers *with* access to credit markets are prompted to take on *more* debt, presumably reinforcing consumption crowding in. In particular, the dynamics revealed by our VAR stress that this mechanism is likely to be operative particularly via the mortgage debt component of private indebtedness.

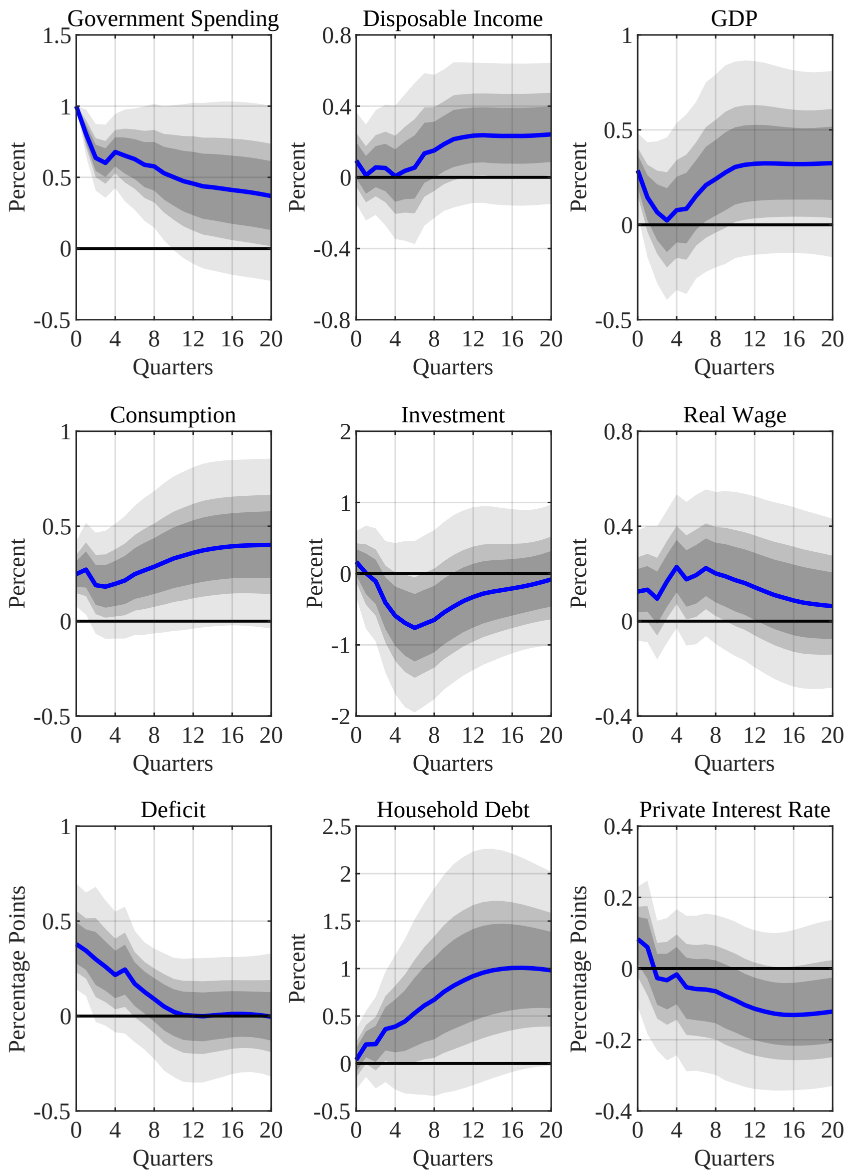
Specifically for the U.S. economy, the link of mortgage debt and consumption expenditures—non-durable goods included—is known to be strong, among others, due to the common practice of home equity extraction (see, e.g., Mian and Sufi, 2011, for evidence on household borrowing via the so-called home equity lines of credit).[19] We add to this line of research the finding that the debt channel also kicks in endogenously, when conditioning on exogenous variation in public spending. The prominent role of *mortgage*—and thus long-term—debt in the propagation of the shock further aligns with results from an additional exercise: once we include the durable expenditures component in the baseline consumption variable, crowding in of private spending is reinforced. The maximum deviation of private spending from its trend increases by approximately 30 percent, relative to the counterpart that excludes durable consumption items.

4.3.4 Sensitivity analysis for the SVAR-IV setting

Ultimately, we scrutinize our SVAR-IV results along the following dimensions. First, we perform several modifications to the baseline instrument. Instead of expressing changes in military spending relative to lagged real GDP (Barro and Redlick, 2011), we (i) use

[19] There is ample empirical evidence documenting that mortgage financing for households has become a driving factor in commercial banks' lending to the household sector, with the share of mortgage loans on banks' balance sheets having doubled in advanced economies over the course of the twentieth century (Jordà et al., 2016).

Figure 4.5: SVAR-IV public spending shock using professional forecast errors



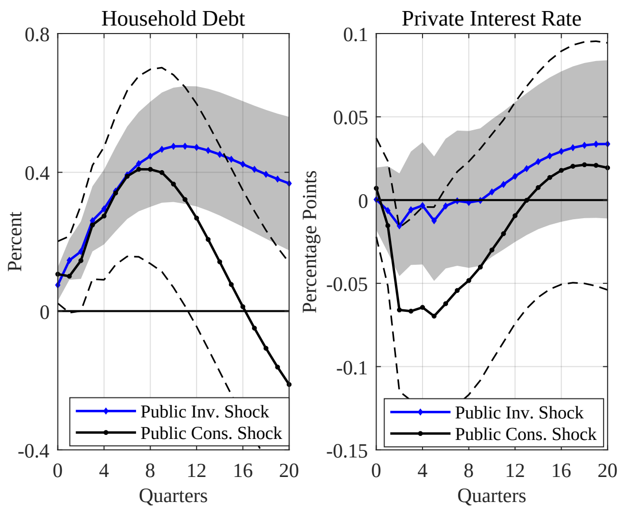
Notes: We plot dynamics at the quarterly frequency. Solid blue lines represent point estimates of impulse response functions obtained from the SVAR-IV model. Dark to light shaded areas display 50, 68, and 90 percent confidence intervals, which we obtain from the weak-instrument-robust procedure in Montiel Olea et al. (forthcoming). Due to data availability, the sample covers data ranging from 1966Q3 to 2008Q4, as in Auerbach and Gorodnichenko (2012).

the civilian population as the normalizing series instead; (ii) we study raw changes in military spending; and (iii) we employ the residual from an AR(1)-regression applied to the benchmark instrument. These adjustments barely affect the impulse response dynamics; the corresponding results are available upon request from the authors. Second, as an alternative to using military spending, we introduce professional forecast errors, g_t^{fe}, on *overall* public spending as an instrument into the VAR model, as proposed in Auerbach and Gorodnichenko

(2012). While these authors identify unanticipated variation in public spending by modeling forecast errors in a recursive VAR with g_t^{fe} ordered first, we, by contrast, deploy this information as an instrument. As explained above, the SVAR-IV strategy appears to be more appropriate in our setting, as it is insensitive to the ordering of variables and relaxes contemporaneous exclusion restrictions, thus capturing the simultaneous interplay of macroeconomic and financial variables that enter our VAR.

In a first step, we calculate the Wald statistic to test the strength of this alternative instrument, which amounts to 29.2, i.e., forecast errors are likely not subject to weak instrument concerns.[20] In a second step, we track in Figure 4.5 the dynamics of a one percent increase in public spending that is identified by forecast errors of professional forecaster. Overall, our inference does not change in this alternative specification, although the impulse response coefficients are estimated with somewhat less precision. Interestingly, the conditional cycle of household debt is quantitatively more pronounced relative to Figure 4.3; the maximum deviation from the conditional mean exceeds 1 percent. After an initial spike, credit rates ease by more than 10 basis points over the medium run.

Figure 4.6: Disaggregate public spending shocks: investment versus consumption



Notes: We plot dynamics at the quarterly frequency for impulse response functions from the SVAR-IV that is identified by forecast errors on government spending (Auerbach and Gorodnichenko, 2012). Solid blue lines (with diamonds) represent point estimates for a government investment shock and shaded areas display the corresponding 68 percent confidence intervals (Montiel Olea et al., forthcoming). Solid black lines (with circles) represent point estimates for a government consumption shock, and confidence intervals are given as dashed black lines.

[20]Note that the forecast error data is only available from 1966Q3. The corresponding Wald statistic for our benchmark military spending instrument over the same starting in 1966Q3 is 44.5.

As a final modification to our baseline SVAR-IV setting, we analyze to what extent the military proxy can be used to recover innovations in more disaggregate government spending data, i.e., we separately identify government consumption and government investment shocks. The respective Wald statistics amount to 36.3 and 33.7, respectively; that is both disaggregate surprises in public spending can be recovered via a strong military spending instrument. The triggered dynamics of these shocks with respect to household debt and credit rates are depicted in Figure 4.6. The expansion in credit appears to be similar in size for both shocks, albeit the shock procession is more persistent in the case of the innovation in government investment. At the same time, interest rates ease significantly for the government consumption shock, whereas the corresponding impulse response is rather flat in the case of a surprise in government investment.

4.4 The transmission mechanism of stimulus to private debt

In this Section, we ultimately provide some tentative insight into the transmission mechanisms that drive our novel aggregate results, i.e., we offer some first path guidance on what actually underlies the conditional *surge* of households' debt position that helps to sustain consumption. To study such propagation channels, we add to the SVAR-IV model from Figure 4.3 one additional time series at a time and report the dynamics for this variable in isolation.

Fisher effects For instance, Fernández-Villaverde (2010) argues that in the presence of financial frictions and multi-period nominal debt contracts, "Fisher effects" may kick in; that is, boosts in inflation may reduce finance premia for borrowing money and thus stimulate debt accumulation and amplify the macroeconomic repercussions of fiscal stimulus. We scrutinize this proposition by incorporating the log of the GDP deflator and, alternatively, the PCE deflator into the SVAR-IV model. Panel (a) of Figure 4.7 illustrates the corresponding impulse response functions, which both reveal a negative, hump-shaped procession of the shock. A surge in public spending hence unleashes disinflationary dynamics in our setting, which aligns with empirical findings of, e.g., D'Alessandro et al. (2019). However, while these authors rely on a Cholesky identification, our external instruments approach—allowing for simultaneous feedback from prices to government spending—still establishes the disinflation result (see Zubairy, 2014, for an estimated DSGE model). Put together, it is unlikely that the conditional debt cycle we observe is driven by Fisher effects.

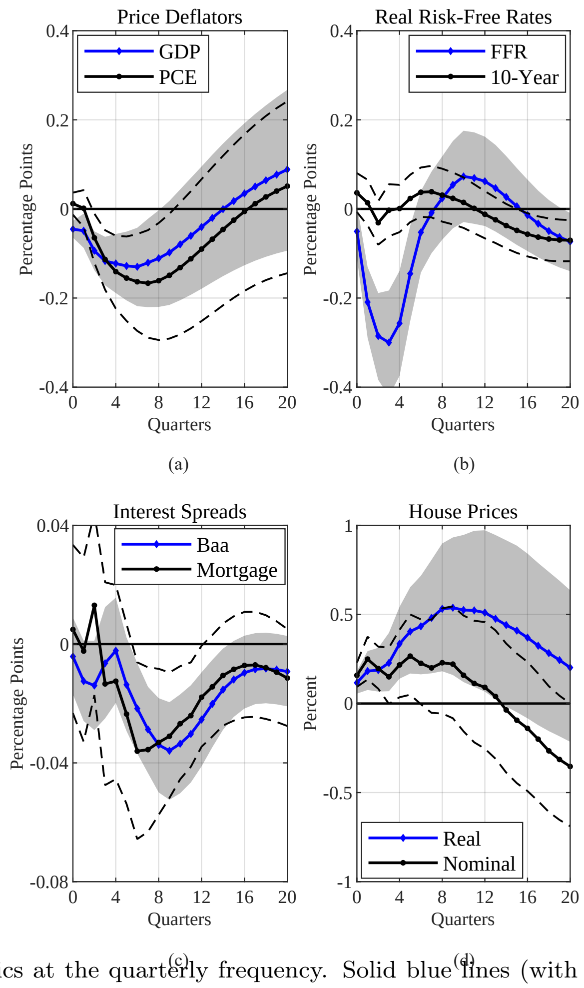
Passive monetary policy Another related mechanism that is typically emphasized in theoretical strands of the literature is that active monetary policy dampens the effects of fiscal stimulus (e.g., Christiano et al., 2011), which may particularly extend to its influence on private credit markets. Considering the declining price levels we have established above in conjunction with conventional central bank reaction functions, such counteracting factors appear to be unlikely to apply in our setting, ex ante. To test such a narrative, we study the conditional dynamics of risk-free rates, i.e., the Federal funds rate and 10-year Treasury yields, in real terms.[21] Panel (b) of Figure 4.7 shows that in line with the first panel of the Figure, monetary policy softens and interest rates at the longer end of the yield curve do not reveal significant cost pressure for credit markets as well. These results, in addition, hold for *nominal* interest rates (not reported). Overall, we do not observe tightening financial conditions, as proxied by risk-free rates, that may depress equilibrium debt. The evidence is instead consistent with the expansion in credit markets that we document.

Softening of credit market constraints Closely related to the mechanism stressed by Fernández-Villaverde (2010), Carrillo and Poilly (2013) argue that fiscal policy, if propagated via imperfect financial markets, may more directly support credit conditions. By stimulating economic activity and by supporting asset prices, public spending may inflate collateral values of borrowers (firms in their case). As a consequence and due to improved balance sheets of borrowers, their access to credit eases, which precipitates in a compression of credit spreads. We inspect such a mechanism by, first, studying to what extent fiscal stimulus widens or narrows interest rate spreads in credit markets. We do so, by analyzing long-term interest rates in relative terms to, e.g., the Treasury yield following convention (in the case of the mortgage rate) or Moody's Aaa bond yield (in the case of the Baa yield as in Gilchrist and Zakrajšek, 2012, to alleviate cash-flow or duration mismatch). Second, we track the dynamics of house prices, in real and nominal terms, as provided by Shiller (2016), to evaluate how collateral values of households absorb the surge in public spending.

Panels (c) and (d) of Figure 4.7 present the corresponding results. Panel (c) is suggestive to easier lending conditions being part of the story, as we observe a hump-shaped drop in credit spreads. Beyond potential shifts in the sovereign yield curve—as depicted in Panel (b) of Figure 4.7—a financial accelerator mechanism thus appears to be at work (Bernanke et al., 1999). Baa corporate bond yields measured relative to their Aaa corporate bond yield

[21] We follow Gilchrist and Zakrajšek (2012) and calculate nominal interest rates relative to the lagged growth rate of the PCE deflator. Our results are robust to using the shadow rate of Wu and Xia (2016) instead of the nominal Federal funds rate; alternatively, restricting the analysis to pre-Great Recession data does not affect the result.

Figure 4.7: Inspecting the transmission channel of fiscal spending to credit markets



Notes: We plot dynamics at the quarterly frequency. Solid blue lines (with diamonds) represent point estimates of IRFs for the variable denoted in the respective panel legend, and shaded areas display 68 percent confidence intervals. In each panel, we study an alternative measure (see legend in each panel) for which the point estimates are represented by solid black lines (with circles), and confidence intervals are given as dashed black lines.

counterpart, deviate negatively from their conditional mean, reaching their trough in the third year after the shock, before leveling off. The result of narrowing spreads also extends to spreads for mortgage credit.[22] The last panel of Figure 4.7 further reveals significant asset price inflation; the level of real estate prices rises on impact, both in nominal and real terms. Real house prices subsequently rise by more than 0.5 percent, two years after the shock. These dynamics make a case for a collateral channel through which fiscal stimulus compresses credit spreads and impacts on households' debt position, in the presence of financial frictions (see, e.g., Carrillo and Poilly, 2013).

[22]These findings put the results of Born et al. (2020) into perspective, who document a widening of the *sovereign* default premium in response to a *cut* in public spending in a sample of 38 countries, on average over the business cycle. They establish their findings using exclusion restrictions that identify their VAR.

4.5 Conclusion

Can public spending stimulate the economy and if so, how? These questions are some of the oldest and certainly most important ones around which large parts of the history of macroeconomics have centered and, which received renewed attention during the rapidly unfolding economic disruptions at the onset of the Great Recession. A crucial mechanism that policymakers often seek to activate in order to boost economic activity, is triggering private via public spending. For instance, fiscal stimulus *payments*, such as the tax rebates ranging between \$ 500 and \$ 1,000 that U.S. Congress authorized during the economic slowdowns of 2001 and the Great Recession, can be viewed as a type of public intervention that was directly intended to raise household absorption. Of course, the success of stimulating private spending—the largest component of aggregate demand—through public stimulus critically hinges on the specific calibration and composition of the public spending program under consideration. Unfortunately, such interventions may not always cause the behavioral adaptions policymakers intend to induce; in this vein, Hoekstra et al. (2017) provide evidence that the 2009 \$ 3 billion Cash for Clunkers scrappage program, which was—apart from the idea to put safer and more fuel-efficient vehicles on U.S. roadways—explicitly tailored to promote private spending, might actually have reduced net total vehicle spending by \$ 5 billion.

The good news is, however, that empirical evidence by a vast number of time series contributions supports the notion of *unexpected* shifts in fiscal spending significantly *raising* private consumption expenditures for aggregate data and on average across programs, in postwar U.S. data. The bad news is, however, we structurally still do not satisfactorily understand why consumption reacts in this way. While crowding in is at odds with the predictions of plain-vanilla New-Keynesian models, there is a fast growing literature that tries to rationalize this empirical regularity. Yet, with all contributions offering "*one* solution to a fiscal policy puzzle" (Bilbiie, 2011), the transmission mechanism of public to private spending is still not well-understood. Consequently, rigorous empirical testing of alternative theoretical approaches is key to a better understanding of the propagation of fiscal stimulus.

In this paper, we provide comprehensive empirical evidence that the most widely-adopted modeling device of rationalizing consumption crowding in by giving disposable income of households a meaningful role, may be accompanied by further transmission channels. In fact, we observe consumption crowding in effects, even in the absence of movements in disaggregate and aggregate measures of post-tax income, in postwar U.S. data. Complementing this finding, we test the hypothesis, whether variation in household indebtedness may reinforce the pass-through of public to private spending. We do so, by carefully modeling the simultaneous interplay of the fiscal policy stance, household consumption, after-tax income, and private credit markets in a structural VAR model that is identified by an external instrument.[23] Indeed, we observe a striking role for public spending to prompt surges in the debt position of households; this leverage that fiscal policy exerts on credit markets appears to be particularly strong for the mortgage component of household indebtedness. In addition, the significant household debt cycle and the crowding in of household absorption are paralleled by declining interest rates in credit markets. This conditional divergence of prices and quantities in credit markets is suggestive to accommodating financial conditions underlying our results. To better understand this mechanism, we provide some first-path guidance on the underlying transmission channels: First, we do not find counteracting effects stemming from risk-free rates, such as the Fed's policy instrument or 10-year Treasury yields. Second, since we observe declining price levels, our results are unlikely to be driven by Fisher effects. Third, we reveal a narrowing of interest spreads in credit markets and, fourth, public stimulus significantly improves real estate prices. The latter two results prompt the view of looser collateral constraints—brought about by rising collateral prices—and thus easier access to credit markets for households, reinforcing the conditional comovement of private spending and the debt position of the household sector.

To put our paper into perspective, we emphasize that for an analysis of the macroeconomic repercussions of public stimulus, more generally, it is vital to carefully address the question of how public spending propagates into private credit markets. Future research should explicitly take into account the dynamic interactions we have identified and should attempt to improve our knowledge on how credit supply and demand conditions react to public stimulus. Making progress toward that direction is crucial to better inform the calibration of and modeling strategies for theoretical approaches that aim to inform policymakers. In addition, further work on the transmission channel of public spending into private credit markets may provide a clearer picture on how discretionary fiscal policy can be used as a tool for macroeconomic stabilization. Specifically, in the presence of a private debt

[23] We thus also add to the literature the observation that consumption crowding in prevails in a time series setting that abstracts from zero or sign restrictions imposed to recover the VAR's structural form.

channel, countercyclical fiscal policy may be desirable during economic downturns not only due to conventional mechanisms, but because of the stabilizing effect it may exert on private credit markets. In the case of recessions that are triggered or accompanied by pronounced private sector deleveraging, such considerations may be of first-order importance.

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