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„The effect of the Eurosystem's asset purchase programme on Euro Area sovereign bond yields“

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Tobias Himmelbauer, B.A. (Econ.)

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Abstract

This master thesis examines the interplay between unconventional central bank policy and government bond yields in the Euro Area. I focus on the Eurosystem's asset purchase programme (APP), and in particular its core constituent, the public sector purchase programme (PSPP). Following a strain of economic literature that attempts to estimate the "flow" effect of these programs, I compile a dataset consisting of monthly data on 9 Euro Area countries over 2003-2018 and model the impact of various macro-financial regressors on 10-year government bond yields. In a static panel context (fixed effects model), I find that bid-ask spreads, short-term interbank rates and economic sentiment have highly significant and economically plausible effects on yields. Furthermore, I find a highly significant yield-decreasing effect for increases in i) PSPP purchases and ii) total Eurosystem assets. While I can partly corroborate these findings in a dynamic context (difference and system GMM models), I cannot confirm a clear-cut flow effect of unconventional central bank policy on yields.

Kurzfassung

Diese Masterarbeit untersucht die Wechselwirkungen zwischen unkonventionellen Zentralbankmaßnahmen und Staatsanleihenrenditen in der Eurozone. Der Fokus liegt auf den Wertpapierankaufprogrammen (APP) des Eurosystems und insbesondere deren zentralem Bestandteil, dem Programm zum Ankauf öffentlicher Wertpapiere (PSPP). Ich baue auf einem Strang der ökonomischen Literatur auf, der versucht, die Effekte der laufenden Implementierung dieser Programme zu schätzen. Ich erstelle ein Datenset mit monatlichen Daten zu 9 Eurozone-Ländern über 2003-2018 und modelliere den Einfluss verschiedener makro-finanzieller Regressoren auf die Renditen von Staatsanleihen mit 10-jähriger Laufzeit. Im statischen Panel Kontext (fixed effects Modell) finde ich hochsignifikante und ökonomisch plausible Effekte von bid-ask spreads, kurzfristigen Interbankenzinsen und der Konjunkturstimmung auf Staatsanleihenrenditen. Außerdem finde ich einen hochsignifikanten renditenverringenden Effekt für Erhöhungen von i) PSPP Ankäufen sowie ii) aggregierten Eurosystem Aktiva. Während ich diese Ergebnisse zum Teil in einem dynamischen Panel Kontext (difference und system GMM Modelle) bestätigen kann, finde ich in diesem Kontext keine Bestätigung für einen eindeutigen Einfluss der laufenden Implementierung unkonventioneller Zentralbankmaßnahmen auf Staatsanleihenrenditen.

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List of abbreviations in alphabetical order

ABSPP	Asset-backed securities purchase programme
APP	Asset Purchase Programme
bn.....	billion (American convention = 10^9)
BoE.....	Bank of England
BoJ	Bank of Japan
bps	basis points (1 basis point = 1/100 of one percentage point)
CBBP3	3 rd covered bond purchase programme
CBOE	Chicago Board Options Exchange
CJEU	Court of Justice of the European Union
CSPP	Corporate sector purchase programme
EA.....	Euro Area
EC.....	European Commission
ECB.....	European Central Bank
ESCB.....	European System of Central Banks
ESFS	European Financial Stability Facility
ESM.....	European Stability Mechanism
EU	European Union
EUR	Euro
Fed	Federal Reserve System
GDP.....	Gross Domestic Product
HICP	Harmonized index of consumer prices
ISIN	International Securities Identification Number
ISO	International Organization for Standardization
LTRO	Longer-term refinancing operations
mn.....	million (10^6)
NCB.....	National Central Bank
OMT.....	Outright Monetary Transactions
PEPP.....	Pandemic Emergency Purchase Programme
PSPP.....	Public Sector Purchase Programme
pp.....	percentage points
SMP.....	Securities Markets Programme
S&P	Standard and Poor's
TLTRO.....	Targeted longer-term refinancing operations
QE	Quantitative Easing
VIX.....	Chicago Board Options Exchange Volatility Index
WAM.....	weighted average maturity

0. Thesis structure

This thesis is structured as follows:

Section 1 - Introduction gives an overview of the context of unconventional monetary policy in the aftermath of the global financial crisis (1a), describes the ECB's unconventional monetary policy measures adopted in general (1b) and focusses on a particular unconventional policy measure, the public sector purchase programme or PSPP in particular (1c) before outlining the research questions of the thesis (1d).

Section 2 – Literature overview gives a brief overview of relevant papers that have attempted to analyse and quantify the influence that macroeconomic and financial market indicators in general (2a) and the Eurosystem's PSPP in particular have had on sovereign bond yield spread dynamics (2b).

Section 3 – Data motivates the choice of sample period and sample countries and describes in detail the variables used in the regressions to follow, discusses their economic relevance and their use in literature (3a-3i) and provides various sample statistics (3j).

Section 4 – Methodology and results, the main empirical part of the thesis, starts by univariate analysis of the individual time series in the dataset (4a) before moving on to static panel modelling (4b), which includes estimating pooled OLS, random effects and fixed effects models, conducting various statistical tests and economic interpretation of the main estimation results. Similarly, the section dynamical panel modelling (4c) discusses the results of difference and system GMM model estimation and related statistical tests.

Section 5 – Conclusions wraps up the main insights from this master thesis, discusses potential shortcomings of the chosen approach and hints at potential avenues for further research.

Section 6 – Appendix includes additional information referred to in the main text, but shifted to the appendix for the sake of brevity.

Section 7 – References lists all sources used throughout the thesis in alphabetical order.

The econometric software programs used for this thesis are Stata and R Studio. Original graphs and tables have mostly been produced using Stata and R, and in a few cases using Microsoft Excel or PowerPoint.

1. Introduction

a. Unconventional monetary policy in the wake of the global financial crisis

Central banks in major economies of the world, ranging from the Federal Reserve System (Fed), Bank of England (BoE) and Bank of Japan (BoJ) to the European Central Bank (ECB), initially reacted with conventional monetary policy to the challenges that arose after the start of the global financial crisis. Responding to falling inflation and increasing risks of economic and financial instability, central banks cut their key interest rates (see Figure 1 for illustrative purposes) with the aim of, inter alia, easing credit conditions on the interbank market and in turn stimulating lending from commercial banks to businesses and consumers, thereby countering deflationary tendencies and supporting economic growth and employment.

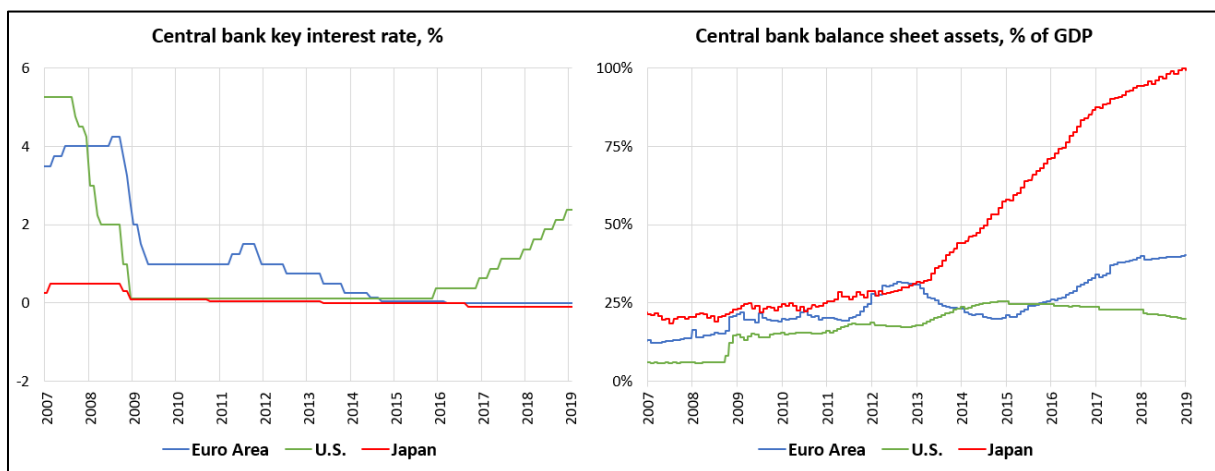


Figure 1 –Key rates and central bank assets/GDP. Data sources: BIS, Refinitiv.

However, these measures did not suffice to bring the desired results as swiftly as it was hoped for. Unemployment rose as economic activity decreased sharply while inflation fell below central banks' targets (BIS, 2019a). Nominal interest rates approached levels that were previously commonly held to be their lower limit: The **effective lower bound**, i.e. the concept that short-term nominal market interest rates are effectively bounded from below by zero or a number close to zero. This concept is based on the proposition that short-term nominal interest rates cannot decrease without limit since eventually such a decrease would lead to a situation where economic agents would prefer holding cash to tying up their wealth in assets that carry negative real interest. In the challenging situation caused by the economic fallout of the global financial crisis, major central banks at various times initiated newly designed monetary policy measures. In the public and academic discussion, these measures were labelled "unconventional" in order to distinguish them from "conventional" measures applied before the global financial crisis, even though some of these unconventional tools had historic precedent or were related to "conventional" liquidity provision by central banks (BIS, 2019a).

In a retrospective overview (ibid.), the Bank of International Settlements describes central banks' "conventional" policy as the pursuit of a primary target (usually a certain inflation level) by influencing short-term interest rates, thereby steering financial conditions and the

macroeconomy through the monetary transmission mechanism. Important **conventional** tools included directly setting a key policy rate or using central bank reserve scarcity to steer the policy rate to a desired level, conducting open market operations (e.g. lending to banks against collateral) and communicating how the central bank would react to the macroeconomic environment (i.e. communicating a policy reaction function). After the onset of the global financial crisis, new **unconventional** policy measures were adopted, including negative interest rate policies (de-facto showing that the effective lower bound was lower than previously expected), expanded lending operations (e.g. by widening the scope of existing facilities or creating new facilities for liquidity provision, loosening collateral requirements, lengthening lending maturities etc.), forward guidance (communicating future policy action intentions) and asset purchase programmes or quantitative easing (QE) (ibid.).

The latter - QE - shall be the focus of this thesis. QE can be generally defined as *“central bank purchases of securities from financial market participants funded by an increase in central bank reserves”* (Joyce et al., 2012). QE was eventually deployed by many central banks including Fed, BoE, BoJ as well as ECB and subsequently led to massive increases in central banks’ balance-sheet assets as a share of GDP of the respective economic area.

The various programs subsumed under the umbrella term “QE” initiated by major central banks, differ widely in their scope, size, technicalities and goals. In line with their respective monetary policy objectives, central banks have stated different headline objectives of their QE programs. Objectives included the Federal Reserve System’s *“to improve overall conditions in private credit markets”* and *“to promote economic recovery and price stability”* (Fed, 2010), Bank of England’s *“boosting economic activity, keeping the UK on track to meet the Government’s 2% inflation target”* (BoE, 2016), Bank of Japan’s *“aiming to achieve the price stability target of 2 percent, as long as it is necessary for maintaining that target in a stable manner”* (BoJ, 2013), and European Central Bank’s *“addressing the risks of a too prolonged period of low inflation”* (ECB, 2015a). To carefully define each of these programs, to analyse what they have in common and what separates them from each other is beyond the scope of this thesis. I will restrict myself to describing the ECB’s program in detail in section 1.c.

The international dimension and sequencing of the introduction of QE is nevertheless noteworthy. The introduction of the ECB’s main QE measure in the aftermath of the global financial crisis, the asset-purchase program (APP), came relatively late as compared to other major central banks’ QE policies. The Fed announced the start of QE in November 2008 and significantly widened its program’s scope in 2009 (Fed, 2008 and 2009), the BoE introduced QE in March 2009 and the BoJ initiated QE in October 2010, significantly expanding its program after 2013, while the ECB’s large-scale asset purchase program was started only in March 2015 (Dell’Ariccia et al., 2018). Thus, when starting its program, the ECB could build on years of real-world policy experience of other major central banks and a growing body of academic literature on the effects of the aforementioned programs. While “unconventional” monetary policy tools in general and QE in particular have been in global use on a substantial scale for over a decade at the time of writing this thesis, the long-run impact of the usage of these tools

on socio-economic circumstances, market incentives and financial stability cannot yet be finally evaluated (BIS, 2019a). As the world economy is faced with new challenges and uncertainties sparked by the COVID-19 pandemic and Russia's war against Ukraine, the question which policy tools are feasible and effective responses to economic crises has gained a new dramatic urgency while the legacy effects of past measures continue to linger. The study of "unconventional" monetary policy measures, which have been widely adopted into central banks' toolkits will thus remain relevant for years and decades to come. The following subsections aim to lay a foundation for understanding

- which unconventional tools the ECB deployed from 2007-2018 under which macroeconomic circumstances (subsection 1b) and
- the key characteristics of the ECB's main QE program (subsection 1c).

b. The ECB's unconventional monetary policy – chronology and context

In its own account of monetary policy decisions and non-standard measures taken since the global financial crisis began in 2007, the ECB differentiates between **three phases**.

In the **first phase**, characterized by a lack of interbank market liquidity, the main aim of the ECB's non-standard measures was *"to provide liquidity to banks and to keep financial markets functioning."* (ECB, 2019d). To that end, the ECB inter alia substantially increased the volume of liquidity provision to banks, extended the maturity of credit provided and widened the list of assets eligible as collateral in refinancing operations (ibid.). (see appendix 6h for a visualization of the Eurosystem's liquidity provisioning).

In the **second phase**, the ECB's non-standard measures *"aimed to address markets' malfunctioning and to reduce differences in financing conditions faced by businesses and households"* (ECB, 2019d). This second phase was a period commonly dubbed the "European sovereign debt crisis" (Lane, 2012). Many sovereigns experienced a spike in bond yields from 2009 onwards (see Figure 2). Rising yields coincided with fiscal austerity measures, increasing unemployment, depressed consumption and the threat of banking sector instability. In this environment, institutional shortcomings of the Euro Area's unique setup, consisting of a common monetary policy authority, but lacking a common fiscal authority or budget of substantial size relative to GDP, became evident (Lane, 2012).

Substantially expanding fiscal stimulus across the EA on national levels to counter deflationary tendencies and boosting economic growth or introducing potentially alleviating intergovernmental fiscal transfers was constrained by a combination of lack of fiscal space in many crisis-hit countries (worsened by government bailout packages for private banks) and political reservations in fiscally "sound" countries with relatively low debt levels and sufficient fiscal space. Monetary stimulus via increasingly accommodative policies emerged as a tool deemed crucial to counter the Euro Area's economic woes and preventing further contagion of financial instability.

At the time, the question to which extent the ECB should intervene to stop sovereign borrowing costs from rising was subject to heavy debate. The announcement of the Securities Markets Programme (SMP) on 10 May 2010 represented a significant shift in the ECB's approach. The SMP consisted of secondary market purchases of sovereign bonds of selected EA governments. Details (total amounts, timeframe, which securities would be targeted) on these purchases were not made public ex ante. Purchases focused at first on Greek, Irish, and Portuguese securities and were later extended to include Italian and Spanish securities (Eser and Schwaab, 2016). In a certain way, SMP was a precursor to later asset purchase programs, but it differed from the latter in several important aspects. Crucially, liquidity injected into the market through SMP operations was to be fully sterilized, i.e. reabsorbed by the ECB, thereby leaving the monetary policy stance unaffected (ECB, 2010). Purchases were made between May 2010 and March 2011 and between August 2011 and February 2012. At the peak, the Eurosystem (comprising the ECB and the national central banks (NCB) of those countries that had adopted the Euro) held securities for worth of EUR 219.5bn under the programme (ECB, 2013a). Eser and Schwaab (2016) find that both the announcement of SMP and the implemented purchases had yield-decreasing effects on affected government bonds. In late 2011 and through 2012, the ECB also extended further the maturity of the liquidity it provided to banks and massively increased the volume of loans to banks in so-called longer-term refinancing operations or LTROs (ECB, 2011). Further information on the ECB's liquidity provision to banks can be found in appendix 6h.

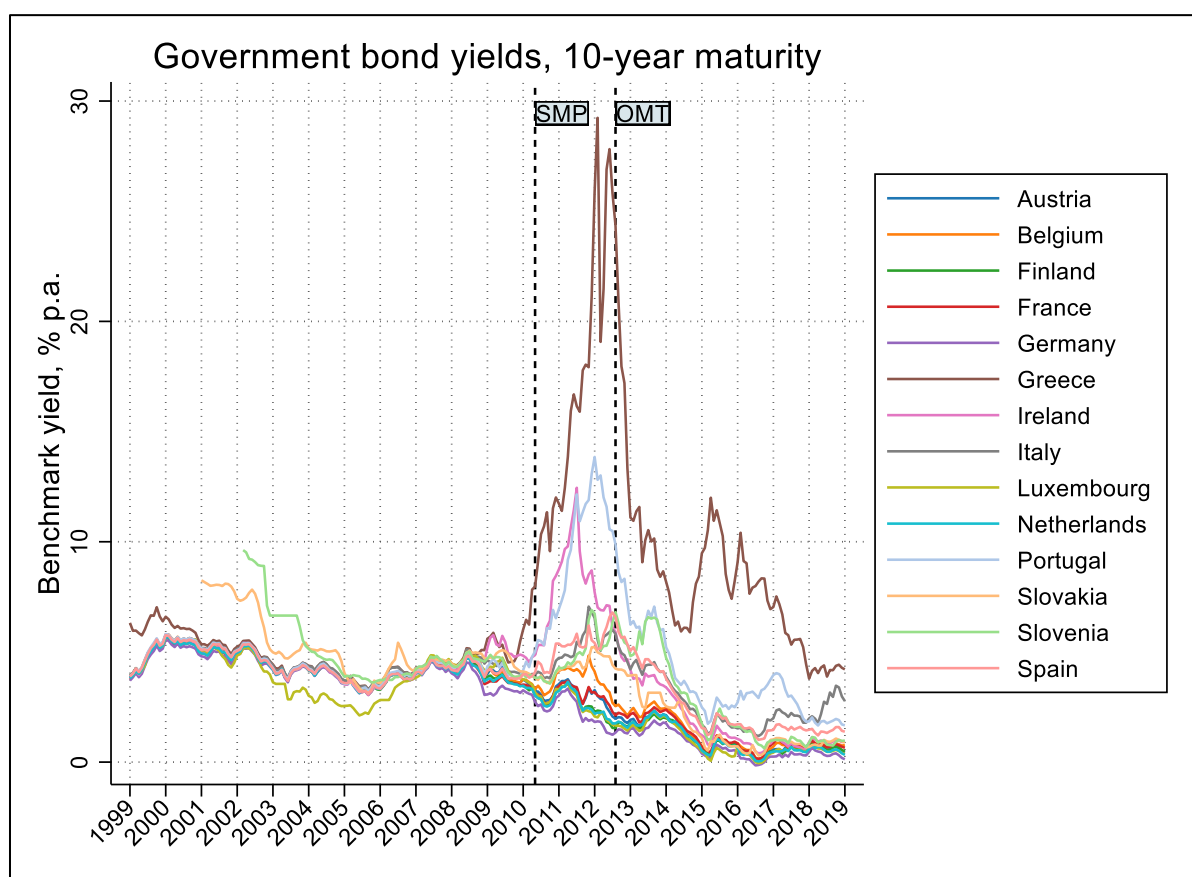


Figure 2 - Government bond yields, SMP, OMT

The ECB's measures coincided with substantial shifts in the multilateral lending architecture of the Euro Area. Following an initial request by Greek authorities, the "Greek Loan Facility" was agreed in May 2010. Under this facility, EA countries jointly with the IMF provided loans worth EUR 73bn to Greece (ESM 2019a). In June 2010, in an attempt to institutionalize the process of loans extension within the EA, the first new multilateral EA lending institution was created: the European Financial Stability Facility (EFSF), set up as a *"temporary backstop"* for countries facing difficulties refinancing on financial markets. EFSF started loan disbursements to Ireland in February 2011, to Portugal in June 2011 and to Greece in March 2012 (ESM, 2019b).

Despite the newly created EA multilateral lending framework on top of the SMP implementation and other ECB measures, yields on sovereign bonds of certain Euro Area countries remained substantially higher than prior to the start of the crisis (see Figure 2). This was particularly true for Greece, Portugal, Ireland, Italy and Spain. The ECB's readiness to further intervene in bond markets was uncertain until then-president of the ECB Mario Draghi made his infamous remark as part of a speech on 26 July 2012: *"Within our mandate, the ECB is ready to do whatever it takes to preserve the euro. And believe me, it will be enough."* (ECB, 2012a).

A week later, the ECB announced its readiness to conduct *"outright transactions in secondary sovereign bond markets that aim at safeguarding an appropriate monetary policy transmission and the singleness of the monetary policy"* (ECB, 2012b). The so-called Outright Monetary Transactions (OMT) framework, which replaced the SMP, was never actually used until the end of 2018 to conduct any purchases. Yet, the announcement of the existence of OMT alone led to significant reductions in short-term government bond yields across the Euro Area (see e.g. Altavilla et al., 2014 or Szczerbowicz, 2015). According to the ECB, OMT *"acted as a powerful circuit breaker against self-reinforcing fears in sovereign bond markets"* (ECB, 2019d).

The announcement of OMT in August 2012 was followed by another significant change in the multilateral lending architecture of the Euro Area. The successor of the ESFS, the revamped European Stability Mechanism (ESM) was officially established in October 2012 as a *"permanent solution for [...] the lack of a backstop for euro area countries no longer able to tap the markets"* (ESM, 2019b). The newly set up ESM started loan tranche disbursements to Spain in December 2012 to and Cyprus in March 2013, later followed by another programme for Greece in August 2015 (ESM, 2019b).

In the **third phase**, the ECB's non-standard measures *"addressed the onset of a credit crunch and the risk of deflation"*, targeting to *"influence the whole constellation of interest rates that are relevant for financing conditions in the euro area"* (ECB, 2019d).

Entering this phase, ECB had already cut its key interest rates (deposit facility rate, main refinancing operations rate [= fixed rate from October 2008 onwards] and marginal lending facility rate) to previously unseen record lows as visualised in Figure 3 (ECB, 2019f).

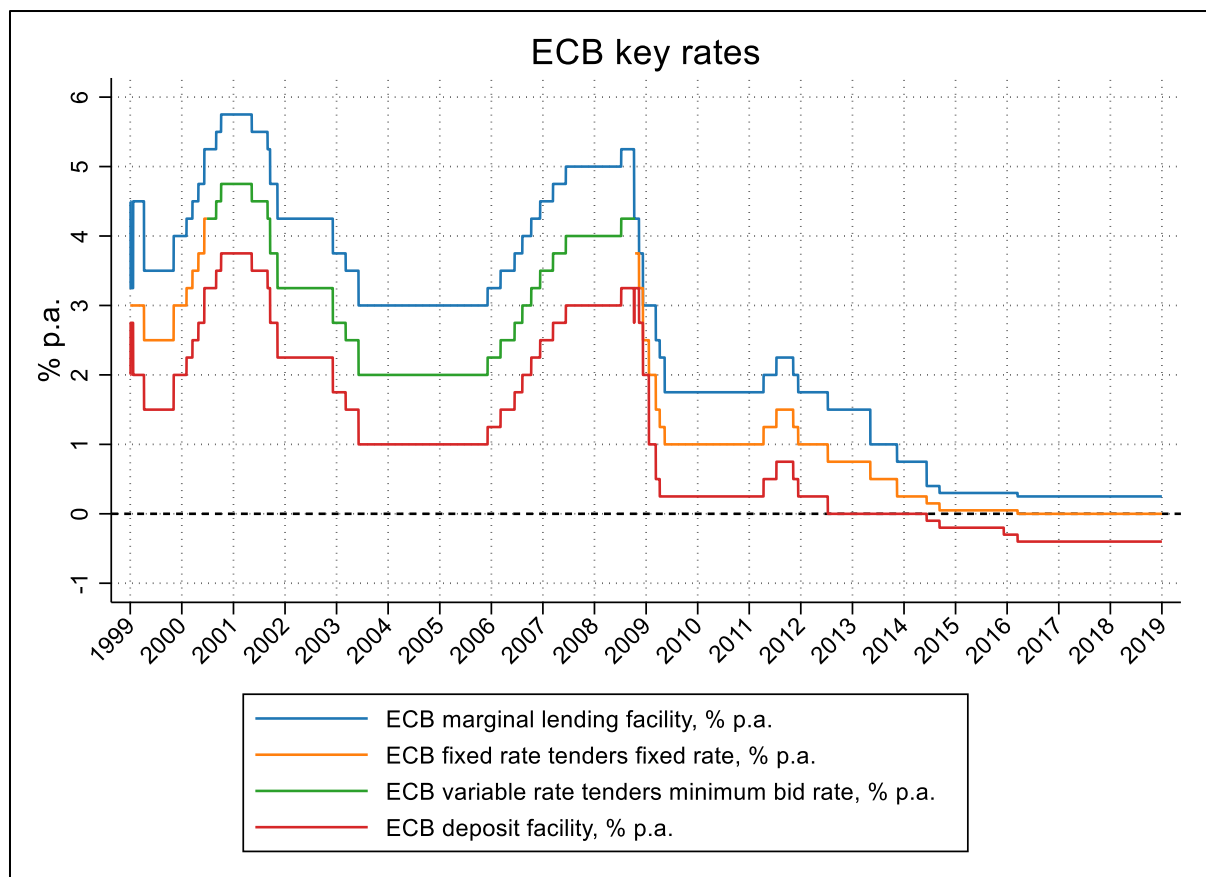


Figure 3 - ECB key rates. Data source: ECB

However, despite its more accommodative monetary policy stance, the ECB during that phase arguably did not fulfil its set primary objective of maintaining price stability: From a peak of 3% year-on-year in October 2011, inflation declined, falling below the ECB's target rate of "below, but close to, 2%" for the EA as a whole after February 2013 as visualised in Figure 4. The ECB responded by cutting key interest rates through 2011-2013 and introducing its own version of forward guidance in July 2013 with the initial formulation "*The Governing Council expects the key ECB interest rates to remain at present or lower levels for an extended period of time*" (ECB, 2013b). The ECB's decision to introduce forward guidance was motivated by the desire to anchor market participants' expectations of future monetary policy intentions (ECB, 2019d).

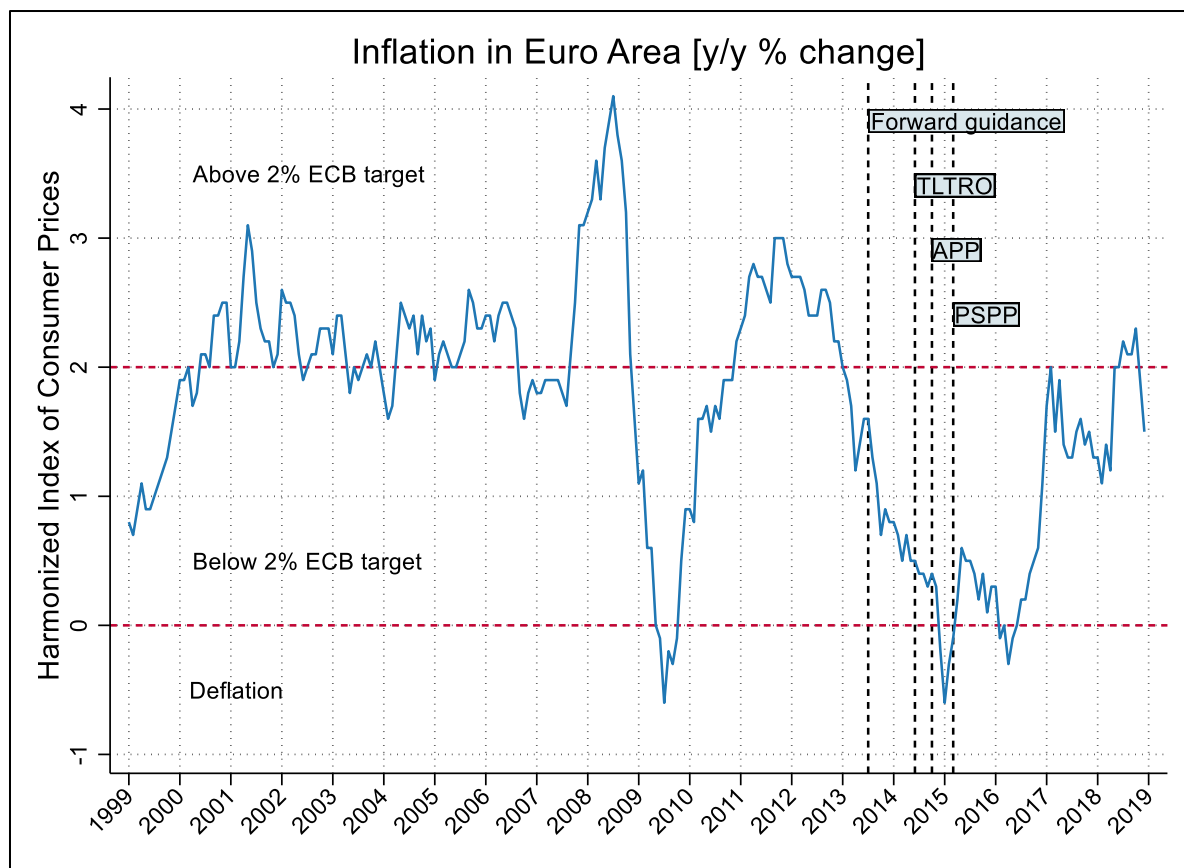


Figure 4 - Inflation in Euro Area. Data source: ECB

However, despite the rate cuts and forward guidance statements, inflation declined further. In this context, the ECB took the unprecedented decision to cut its deposit facility rate, i.e. the rate which “banks may use to make overnight deposits with the Eurosystem” (ECB, 2019f) into negative territory to -0.1% in June 2014. Simultaneously, the ECB introduced so-called targeted longer-term refinancing operations (TLTROs). TLTROs were designed to “support bank lending to businesses and households” (ECB, 2019d) by linking volume and later also the interest rates of loans from the ECB to banks to these banks’ loans to non-financial corporations and households (ECB, 2019g). The idea was to incentivize banks to increase lending to the “real economy” by giving such banks that focus on lending to non-financial sectors easier access to ECB liquidity vis-à-vis banks that focus on lending to the financial sector. Rates were further cut in November 2014. Nevertheless, inflation continued to decline through the second half of 2014.

Having taken a number of new and unprecedented monetary policy measures as outlined above while still being faced with significant challenges in achieving its primary monetary policy objective of maintaining price stability as defined by inflation “below, but close to, 2%”, the ECB decided to set in motion yet another new and unprecedented unconventional policy measure. In late-2014, the Eurosystem started its asset purchase program (APP). Within the APP framework, the first two programs were the third covered bond purchase programme

(CBBP3) and the asset-backed securities purchase programme (ABSPP). Purchases under these programs were at first conducted on a relatively small scale (Eisenschmidt et al., 2017).

On 22 January of 2015, the ECB announced that it would significantly expand the APP to include its by far largest component, the Public Sector Purchase Programme (PSPP) (ECB, 2015a). The ECB then argued that *“the ECB’s interventions [note: referring to the APP] will reduce yields on government bonds, which will set in motion a more conventional chain of propagation channels that will support the economic recovery and help bring inflation back to levels below, but close to, 2%.”* (ECB, 2015b). The PSPP programme will be explained in detail in subsection 1c.

c. The public sector purchase programme (PSPP) – chronology and technicalities

Henceforth, the Eurosystem will be referred to as the main enactor of the APP in the EA. This is because the operational implementation of the APP in the EA is decentralized, i.e. NCB conduct purchases under the ECB governing council’s design, control and coordination (ECB, 2015b).

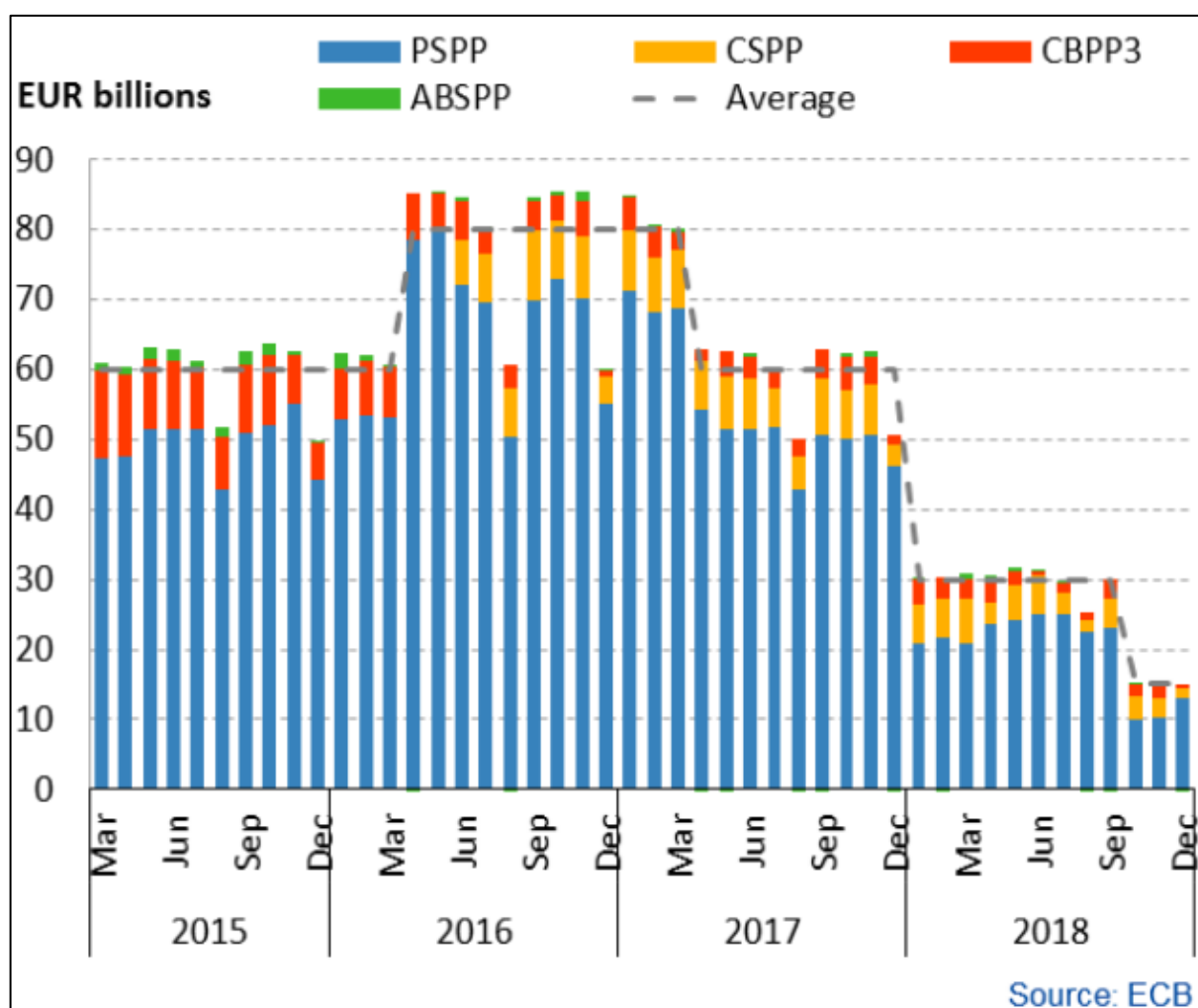


Figure 5 - The Eurosystem's APP – Monthly net purchases. Source: ECB, 2019a.

On 9 March of 2015 the Eurosystem started to buy public sector securities under the PSPP. Later, in June 2016 the corporate sector purchase programme (CSPP) complemented the three other sub-programs under the APP, which saw net purchases until the end of 2018 (ECB 2019a). CBBP3, ABSPP and CSPP are mentioned here for the sake of completeness; however, since this master thesis focuses on public sector bonds, the other-sub-programs will not be analysed or explained in further detail (See appendix 6g for a further visual overview).

The total volume of monthly net APP purchases was capped to EUR 60bn from March 2015 - March 2016, raised to EUR 80bn from April 2016 until March 2017 and subsequently downsized to EUR 60bn from April 2017 to December 2017, EUR 30bn from January 2018 to September 2018 and finally EUR 15bn from October 2018 to December 2018, after which net purchases of the first PSPP phase (covered by this thesis) were terminated.

Securities covered by the PSPP include nominal and inflation-linked central government bonds as well as bonds issued by recognised agencies, regional and local governments, international organisations and multilateral development banks located in the euro area. The Eurosystem allocated 90% (88% from March 2015 until March 2016) of the total purchases to government bonds and recognised agencies, and 10% (12% from March 2015 until March 2016) to securities issued by international organisations and multilateral development banks. The Eurosystem had an “issuer limit” for the PSPP, i.e. the maximum share of any individual issuer’s outstanding securities that it can buy. This issuer limit was set at 33% of total eligible assets of Euro area governments and at 50% of total eligible assets of certain international organizations (ECB, 2019a).

Allocation of PSPP purchases across eligible jurisdictions is *“guided, on a stock basis, by the respective national central banks’ subscription to the ECB capital key”* (ECB 2019a). The national shares in the ECB’s capital, in turn, are calculated using *“a key which reflects the respective country’s share in the total population and gross domestic product of the EU [...] with equal weighting”*. (ECB 2019b). Thus, countries with a relatively high population and share of EU GDP have received higher allocations of monthly PSPP purchases than “smaller” ones (see table 1).

With the establishment of these rules, the size of purchases of securities of a given country in relation to the total permitted amount of purchases was roughly defined. However, the wording of these rules still gave the ECB and national banks conducting purchases on behalf of the Eurosystem some leeway as to the exact amount of monthly purchases. Hence, actual monthly purchases varied over time even though the capital key remained unchanged between 1 January 2014 and 1 January 2019.

Column C in table 1 shows whether a country’s assets were over-or underrepresented in cumulative net PSPP purchases from 2015-2018 by comparing these purchases to the respective NCB’s share in paid up ECB capital. Assets issued by the Austrian, Belgian, French, Italian and Spanish authorities were overrepresented in cumulative net PSPP purchases up until 31.12.2018 (Data sources: ECB, 2019d and ECB, 2019b; author’s calculations).

Country	A: Share of country's assets in cumulative net PSPP purchases as of 31.12.2018, %.	B: Share of NCB in ECB capital as of 31.12.2018, %.	C: Difference of columns [A-B], percentage points.
Austria	3.00	2.99	0.0038
Belgium	3.78	3.72	0.0558
Cyprus	0.03	0.22	-0.1865
Estonia	0.00	0.29	-0.2895
Finland	1.71	1.87	-0.1650
France	21.65	20.93	0.7241
Germany	26.71	27.05	-0.3410
Ireland	1.55	1.73	-0.1790
Italy	18.82	17.38	1.4365
Latvia	0.11	0.40	-0.2954
Lithuania	0.16	0.60	-0.4349
Luxembourg	0.13	0.33	-0.1998
Malta	0.06	0.11	-0.0484
Netherlands	5.93	5.99	-0.0580
Portugal	1.90	2.41	-0.5127
Slovakia	0.60	1.18	-0.5770
Slovenia	0.41	0.50	-0.0857
Spain	13.44	12.28	1.1528
Notes	100% = sum of all cumulative net PSPP purchases for eligible countries (i.e. EA members excluding Greece and supranationals).	100% = sum of paid-up capital from NCBs for which country is eligible for PSPP (i.e. EA members excluding Greece)	Positive: Country's assets overrepresented. Negative: Country's assets underrepresented.

Table 1 - PSPP purchases vs ECB capital key

It is also worth noting that the importance of the Eurosystem as an actor in secondary bond markets of various countries varies, since PSPP purchases made up different shares of the secondary bond market turnover across countries. This is because the allotment of purchases via the ECB's capital key is only an imperfect proxy for the relative size of these markets. The widening of the Eurosystem's balance sheet as a consequence of the APP is unprecedented. Eurosystem holdings of securities accumulated via the APP amounted to EUR 2,102.04bn at the end of December 2018, roughly corresponding to 18.1% of 2018 annual EA GDP (ECB 2019a, Eurostat 2019). This number dwarfed the EUR 295bn figure corresponding to the total volume of loans disbursed by ESM/EFS to date (ESM, 2019c). Securities of EA residents denominated in EUR made up just 2.9% of the annual consolidated balance sheet assets of the Eurosystem in 1999. By end-2007, the share had increased to 9.5% of the total and by end-2014 (before the start of the PSPP in March 2015) to 26.7%. Since then, the share has surged to an astonishing 61.7% of total assets (end-2018). Over the same timeframe, the ratio of total Eurosystem assets as a share of EA GDP also increased drastically; it reached just 16.6% in 2007, rose to 22.2% in 2014 and almost doubled to 41% in 2018 (ECB 2019c, Eurostat 2019).

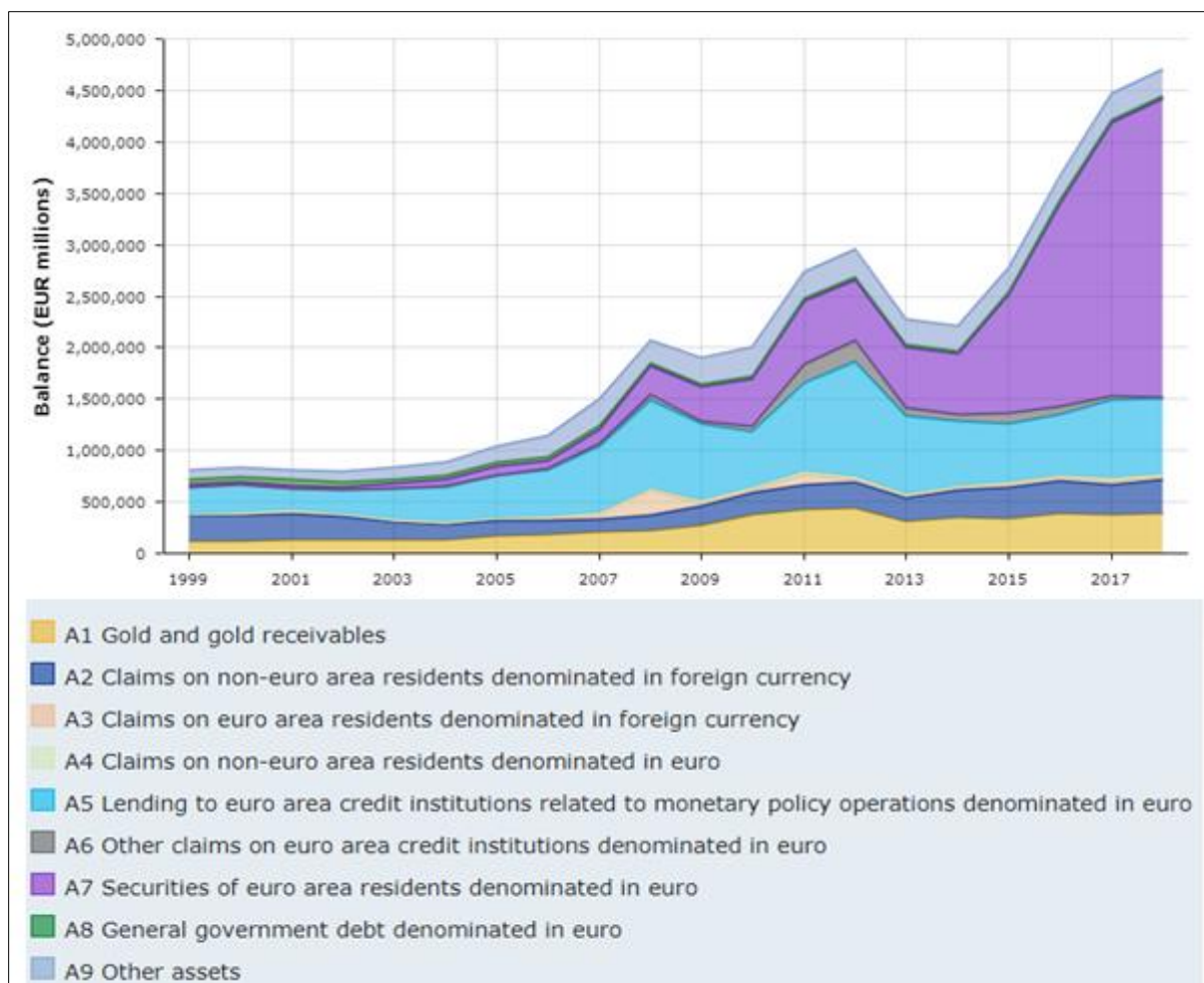


Figure 6 - Eurosystem annual balance sheet. (Source: ECB 2019c)

The Eurosystem entered the secondary market for bonds issued by EA governments and related sub- or supranational issuers on a massive scale while altering drastically its balance-sheet composition as well as widening its arsenal of monetary policy measures. At end-2014, the ECB had 4.1% of all outstanding debt securities issued by EA residents on its balance sheet (by nominal value). At end-2018, this ratio reached 20%; the lion share of the increase can be attributed to PSPP holdings. Hence, the Eurosystem has emerged as a major holder of sovereign and sovereign-related bonds. While the first period of APP net asset purchases ended in December 2018, net purchases were restarted from November 2019 (ECB 2019e) and a new, massive asset purchase program (Pandemic Emergency Purchase Programme or PEPP) was announced in March 2020 (ECB 2020) in reaction to the economic effects of the COVID-19 pandemic. Even before the restart of net purchases and the drastic expansion of QE under new operational guidelines due to COVID-19, the Eurosystem was planning to continue to reinvest principal payments from maturing securities purchased under the APP “for an extended period of time past the date when it [the ECB governing council] starts raising the key ECB interest rates” (ECB 2018b, 2019a and 2019c). Furthermore, the Eurosystem was seeking to maintain the size of cumulative net purchases under each APP sub-programme at

their end-2018 levels (ECB 2019c). In any case, the Eurosystem is set to remain a crucial actor on EA securities markets for the foreseeable future.

Although the ECB with regards to PSPP ascertained that by “*strictly adhering to the prohibition on monetary financing by not buying in the primary market*” and only buying bonds “*after a market price has formed*” it was ensuring that the ECB did not “*distort the market pricing of risk*” during the first phase of net purchases (ECB 2016), such a deep structural monetary policy shift with long-term implications for sovereign bond markets deserves a comprehensive analysis. The remainder of this master thesis will be devoted to investigating the link between PSPP purchases and long-run changes in EA sovereign bond yield spreads, with subsection 1.d. defining the research questions to be tackled.

d. Research questions

From a political point of view, the announcement of OMT in 2012 and the implementation of the PSPP from 2015-2018 were controversial. Cases were brought to the Court of Justice of the European Union (CJEU) against both OMT and PSPP. While the two programs differed substantially in scope and technical criteria, the argument made against them was similar in essence: in both cases, it was alleged that the programs violated the prohibition of financial assistance from the European System of Central Banks to a member state of the EU as stipulated by treaties on the ECB’s mandate and the functioning of the EU. The CJEU ruled that both OMT and PSPP were in accordance with EU treaties as well as the ECB’s mandate (CJEU, 2015 and 2018). While these rulings might have temporarily side-lined the legal dispute, the political debate on how large the Eurosystem’s role in sovereign bond markets- and by extension indirectly in governments’ financing- should be is far from over.

On top of the legal and political disputes, the policy discussion over the future of the Eurosystem’s unconventional monetary measures is ongoing. The general discussion on QE is as relevant as ever, and many questions remain open. What impact, if any, would the phasing out of net purchases have on sovereign bond yields, on the affordability of government debt and in turn on fiscal space and general economic developments? For how long should redemptions be reinvested? Given that the volume of principal reinvested will be subject to the volume of maturing securities, would reinvestments add to volatility in bond markets as total purchases during a reinvestment phase could be more unevenly spread over time than during the net purchase phase of the PSPP? How will changes in the implementation of the asset purchases programme affect interest rates across other asset classes, bank lending and overall financial stability? Should the Eurosystem use similar instruments in future crises, and what impact will the Pandemic Emergency Purchase Programme (PEPP) have? It is essential to understand the impact that the Eurosystem’s QE has had in order to assess its longer-term implications and its potential future. To narrow down the scope of this master thesis, I will focus on the following two research questions:

- 1) What are the main factors driving the changes in Euro Area sovereign bond yields?
- 2) Did the Eurosystem's public sector purchase programme (PSPP) actual purchases or, more broadly, the overall expansion of the Eurosystem balance sheet have a statistically measurable impact on 10-year Euro Area sovereign bond yields?

2. Literature review

This section contains a brief overview of relevant papers that have attempted to analyse and quantify the influence that macroeconomic and financial market indicators in general and the Eurosystem's PSPP in particular have had on sovereign bond yield spread dynamics.

a. Macroeconomic and financial market drivers of sovereign bond yield spreads

Poghosyan (2014) analyses driving factors of sovereign bond yield spreads across countries in a panel covering annual data for 22 advanced economies over the years 1980–2010. In Poghosyan's estimations, the dependent variable is not the spread of bond yields (difference of yields of country A vis-à-vis country B), but rather the level of bond yields by country. The yield spreads across countries are explained by the differences in the fundamental factors that influence the level of yields. The author finds two main long-term drivers of increases in real yields of benchmark government bond with a 10-year maturity: increases in the debt-to-GDP ratio and increases in the potential growth rate. Furthermore, Poghosyan's results indicate that increases in real short-term money market rates (influenced by monetary policy of the respective NCB) and increases in global economic uncertainty (using the Chicago Board Options Exchange's Volatility Index VIX as a global proxy) increase sovereign borrowing costs in the short run.

Costantini et al (2014) use monthly data on 9 EA countries covering the period from January 2001 to December 2011 to analyse the driving forces of sovereign bond yield spreads of these countries vis-à-vis Germany. In their model specification, all variables are expressed as relative values (differentials) vis-à-vis Germany, which is taken as a benchmark country in the framework of an optimal currency area perspective. The authors find that increases in the values of expected public debt/GDP differentials (using European Commission Forecasts), increases in market liquidity risk differentials (bid-ask spreads on secondary government bond markets as a proxy) as well as increases in accumulated inflation differentials are the most important drivers of increases in sovereign bond yields spreads. Furthermore, according to the authors' estimations, the relative importance of the independent variables varies across countries: Differences in expected public debt/GDP ratios and accumulated inflation have a stronger effect on yield spreads for "peripheral" EA member countries. The authors argue that investors might perceive these "peripheral" countries as not belonging to a "core" group of EA members. The latter group, investors might think, would be better suited to form a monetary union among itself than the EA as a whole, including peripheral and core countries. The authors' line of reasoning suggests that large cumulated inflation differences vis-à-vis Germany, which are used to identify "peripheral" EA countries, could in the longer run lead to

a loss of competitiveness of the periphery vis-à-vis core economies as rising wages in the periphery are not accompanied by rising productivity of the same magnitude. As periphery states' lower competitiveness is associated with a higher likelihood of future debt repayment difficulties, the authors argue, investors perceive further increases of public debt/GDP in the periphery as more harmful than for core economies.

Kinateder and Wagner (2017) use monthly data covering the period from January 2003 to December 2013 on sovereign yield spread changes for the 10 biggest EA economies to analyse the driving forces of yield spreads of these countries vis-à-vis Germany. The dependent variable are yield spread changes (i.e. the first difference of yield spreads). Explanatory variables include a group of country-specific variables (bid-ask spreads, S&P credit ratings, public debt/GDP, local equity market returns, lagged yield spreads) as well as common variables (inter alia market liquidity proxies, short-term interest rates and total ECB balance sheet assets as a proxy for unconventional monetary policy). Among the most important drivers, the authors find that deteriorations in sovereign credit ratings and increases in bid-ask spread have highly significant spread-increasing effects. Furthermore, deterioration in overall market liquidity is found to have a spread-increasing effect while expansions of the ECB's total assets are found to have a spread-decreasing effect (in particular during the "crisis period" from March 2009 - December 2013). The authors further note that common unobservable time-varying factors play a big role in driving yield spread changes.

Summing up, economic literature has identified several important long-term drivers of long-term sovereign bond yield spreads across countries. Commonly named important driving factors include differences in public debt/GDP, short-term interest rates influenced by monetary policy, differences in bond market liquidity, GDP growth developments, and international financial risk sentiment.

b. PSPP effect on EA sovereign bond yields spreads

The literature on estimation of the effects of large-scale asset purchases by central banks on financial markets in general can be loosely grouped into two sometimes overlapping strains: one strain relies on event study methodology and another strain on time-series analysis (BIS, 2019a). This subsection contains a brief discussion of the merits and downsides of the two strains as well as an overview of literature that has attempted to estimate the effects of the APP/PSPP on EA bond yields.

- The **event study methodology** usually focuses on a specific timeframe around the announcement of new policy measures by a central bank and then estimates the immediate effect of this announcement on asset prices. In this context, a key assumption is that market participants did not anticipate the announcement prior to the point in time when it was made. In the case of the ECB, the assumption that market participants did not assume the start of an EA APP prior to the announcement of the PSPP might not be sensible as the ECB's APP was preceded by APPs of other major central banks including Fed, BoE and BoJ as outlined in subsection 1a. Thus, one could

argue that in view of falling inflation in the EA after October 2011 (subsection 1b) and APP implementation by other central banks at the time, market participants might have expected an EA APP to be announced sooner or later and that such an announcement might have been at least partially priced in ahead of the actual announcement. This methodological challenge is noted and tackled by De Santis (2020). Another potential shortcoming of the event study methodology is that short-run results obtained through this method cannot generally be used to judge the long-run effects of central bank APPs over these program's lifetime (BIS, 2019a). For a nation-state, it is arguably more important whether a) yields on government bonds are reduced by an APP and remain at a lower level over time than whether b) an APP has short-lived decrease in government bond yield on the secondary market or not. In case a), interest rates locked in at future primary government bond auctions are also likely to be lowered in comparison to a hypothetical world without APP, thereby lowering longer-term refinancing and debt servicing costs for the sovereign.

- The **time-series analysis methodology** generally relies on longer periods of observation than event studies, incorporating trends of macro control variables by using data that extends into a period before the global financial crisis when effective lower bound-constraints on interest rates did not play an immediate role and indices of financial stress stood at lower levels both for specific countries and the EA as a whole. However, by doing so, the implicit assumption that model parameters are stable over time is often taken (BIS, 2019a). This might also be problematic as policy measures taken by governments and central banks in the aftermath of the global financial crisis were fundamentally different from previously used instruments and were implemented in a changed macroeconomic context, as discussed in subsection 1b. This change in macroeconomic context could represent a “structural break”, which might effect estimation results. Furthermore, the ability of researchers to quantify the effect PSPP has had on government bond yields is made inherently difficult by the impossibility of observing the counterfactual, i.e. how government bond yields across the EA might have evolved in the absence of the PSPP.

In an early event study analysis of the effects of the Eurosystem's APP, **Altavilla et al. (2015)** find that the **APP had a yield-reducing effect by about 30-50 basis points (depending on the approach) for 10-year maturity EA sovereign bond yields** and a bigger effect for “higher-yield” member countries Italy and Spain. The authors focus on the “stock effects”, i.e. on the effects that the announcement/events shaping market anticipation of the APP has caused, rather than on the “flow effects”, i.e. the effects of the actual purchases conducted under PSPP. The authors identify key event dates over September 2014 to March 2015, including events that positively “*affected expectations about the APP*” (using an index on news coverage constructed from the Dow Jones “Factiva” newswire) as well as the official announcements of PSPP and its details in January and March 2015. Then, the authors analyse intraday yield data for these event dates on the four largest member countries as well as the EA sovereign term

structure as published by the ECB and find the aforementioned overall effect of 30-50bps, using the assumption that the overall effect of the APP can be obtained by adding up the changes in yields which occurred at each event. The cumulative yield decreases across selected event periods, which is compared to a hypothetical baseline scenario where these yield changes would not have materialized as the underlying events causing these yield decreases did not occur, is thus used to estimate the APP's effects.

De Santis and Holm-Hadulla (2019) focus on flow effects, i.e. the effect of net APP purchases instead of the effects caused by APP announcement/newsflow. More specifically, they model the price impact on PSPP-eligible securities as a function of the withdrawal of net bond supply via Eurosystem PSPP purchases. Their data set consists of daily observations on the bond universe eligible for PSPP (3,061 securities on ISIN-level) over the period from 9 March 2015 to 21 June 2016. The authors identify an important endogeneity problem arising from simultaneity: the Eurosystem's agents who are conducting the actual bond purchases on the secondary market, while being guided by an overarching framework, in their decisions on which specific bonds to buy on a daily basis also consider the prevailing secondary market prices of securities eligible for purchase. It follows that volume and prices of bonds traded in the secondary market are jointly determined. To control for this endogeneity, the authors use one of the PSPP's technical features: the "blackout" period. According to the PSPP framework, under certain condition for a predetermined time window around the primary market issuance of new sovereign bonds, existing PSPP-eligible assets with a remaining maturity close to the maturity of the newly issued bond must not be purchased. The resulting daily variation in quantities of purchase-eligible subgroups of the total PSPP-eligible bond universe is assumed to be an exogenous instrument variable. Within their model framework, the authors find that that PSPP purchases cause significant price increases in the bonds being purchased as well as on close substitute securities: **reducing the net supply of outstanding bonds in the relevant market segment by 1% through PSPP purchases leads to an increase in the return of the security being purchased by around 0.55%** (the return is defined as the percentage change in the price level from day $t-1$ to day t). The effects are found to fade out after one trading day and are *"concentrated in securities issued by higher yield jurisdictions, located at the longer end of the maturity spectrum, and characterized by low liquidity"*.

Eser et al (2019) combine event study methodology with a theoretical model and use daily data on bond yields from December 2009 to June 2018 in a panel of the four biggest EA economies to estimate the overall effect of the APP on the term structure of EA bond yields. In a simplifying step, granular data on government bonds issued by these four countries and weighted by GDP is used together with data on PSPP to estimate APP effects on EA sovereign bond yields. The authors' approach to model the effects of QE onto sovereign bond yields is mainly based on the duration risk channel. The main idea of their model is that the term premium of bonds (yield differentials for different bond maturities) is an increasing function of bond holdings by bond-price-sensitive arbitrageurs. The proposition is that yields of bonds increase when a larger share of them is held by bond-price sensitive market investors (as

opposed to investors that prefer to hold the bond until maturity) which might want to sell the bond on the secondary market prior to the bond's maturity. By conducting large-scale asset purchases, central banks can reduce the exposure of price-sensitive investors to duration risk and thereby increase the risk-bearing capacity of such market participants. This should induce a decline in sovereign bond yields. In order to estimate the effects of APP, the authors construct a measure of the ratio of outstanding sovereign bonds held by price-sensitive arbitrageurs out of the total. They also construct an estimate of the anticipated future development of this ratio as a function of the Eurozone's APP announcements. They find that **the APP has cumulatively reduced 10-year EA sovereign term premia by around 95 bps (estimate range: 65-130 bps), with the bulk of the effect (50bps) created by the initial APP announcement in early 2015.** The authors also find that the effects are stronger for longer-maturity bonds than for shorter-maturity bonds.

An interesting combination of time series analysis with event identification is performed by **De Santis (2020)**. The author uses monthly data (financial data: end of month) covering the period from February 2004 to October 2015 in a panel consisting of the 10 largest EA countries by GDP to analyse the driving forces of sovereign bond yields over that timeframe. De Santis identifies a significant yield-increasing effect for increases in the following macro-financial indicators: market liquidity risk (bid-ask spreads on secondary government bond markets as a proxy), short-term (3-month) market interest rates, expected future inflation, macroeconomic uncertainty (proxied by real GDP growth/inflation forecast dispersion). Conversely, a significant yield-decreasing effect is found for increases in expected real GDP growth and the fiscal balance/GDP. On top of controlling for these factors, the author outlines as a central problem the identification of the point in time from which on market participants started to "price in" the introduction of an APP in the EA. De Santis tackles this challenge by using a keyword search applied to financial news and data provider Bloomberg to construct an index measuring the intensity of the discussion around introduction of an APP in the EA. This index is then used as a regressor and is found to have a significant sovereign yield-decreasing effect, i.e. when more media reports on a potential APP in the EA were published, yields declined. The author finds that the **Eurosystem's APP reduced yields on EA sovereign bonds with a 10-year maturity by cumulatively 72bps up until October 2015** (compared with a counterfactual world without the Eurosystem's APP). The 72bps correspond to the effect estimated for the average yield across the analysed group of EA countries weighted by nominal GDP; the estimated effect is found to be higher for "vulnerable" countries (e.g. Italy, Spain, Portugal).

To summarize, economic literature has employed event study methodology and time series methodology to investigate the effects of the APP/PSPP on EA bond yields. Authors have often focussed on yield data on bonds with a 10-year maturity as benchmark cases and have sometime simplified by using data on a subgroup of bigger countries within the EA to make general statements on APP/PSPP effects on EA government bonds. The effects created by the announcement of such programs ("stock effects") have generally been found to be bigger than the impact of the actual purchases of sovereign bonds ("flow effects").

In this master thesis, I want to focus on the flow effects of PSPP purchases as done by De Santis and Holm-Hadulla (2019) instead of the announcement effects as e.g. in Altavilla et al. (2015). Due to lack of availability of higher-frequency and granular bond data in sources which I can access, I am going to use aggregated monthly data. I want to employ time series methodology and combine data on government bond yields with data on other macro-financial indicators, as done by De Santis (2020) or the papers described in section 2a. Both data and methodology used for this thesis will be described in more detail in the following sections 3 (Data) and 4 (Methodology and estimation).

3. Data

In the literature on modelling the changes of sovereign bond yield levels or spreads over time, the frequency of data used differs considerably. While some papers use annual data to analyse long-run dynamics (Poghosyan, 2014), most use higher-frequency data, like monthly data (e.g. Costantini et al, 2014 or Kinatader and Wagner, 2017 or De Santis, 2020), daily data (e.g. De Santis and Holm-Hadulla, 2019 or Eser et al., 2019) or even intra-day data (e.g. Altavilla et al., 2015). I am using monthly data, since data on a higher frequency is hard to obtain using only open sources. The data I use covers the period from 01/2003 until including 12/2018 (192 months). Out of this timeframe, PSPP purchases were made in 46 consecutive months (03/2015-12/2018). The data covers the following 9 Euro Area members: Austria, Belgium, Finland, France, Italy, Germany, the Netherlands, Portugal, and Spain. All these countries irrevocably fixed their exchange rates via the intra-EU exchange rate mechanism (ERM II), on January 1st, 1999 (ECB 2018). They were the biggest Eurozone members in terms of annual nominal GDP for most of the timeframe under consideration and accounted for the biggest shares in Eurosystem's net cumulative PSPP purchases (see Figure 7; data source: ECB).

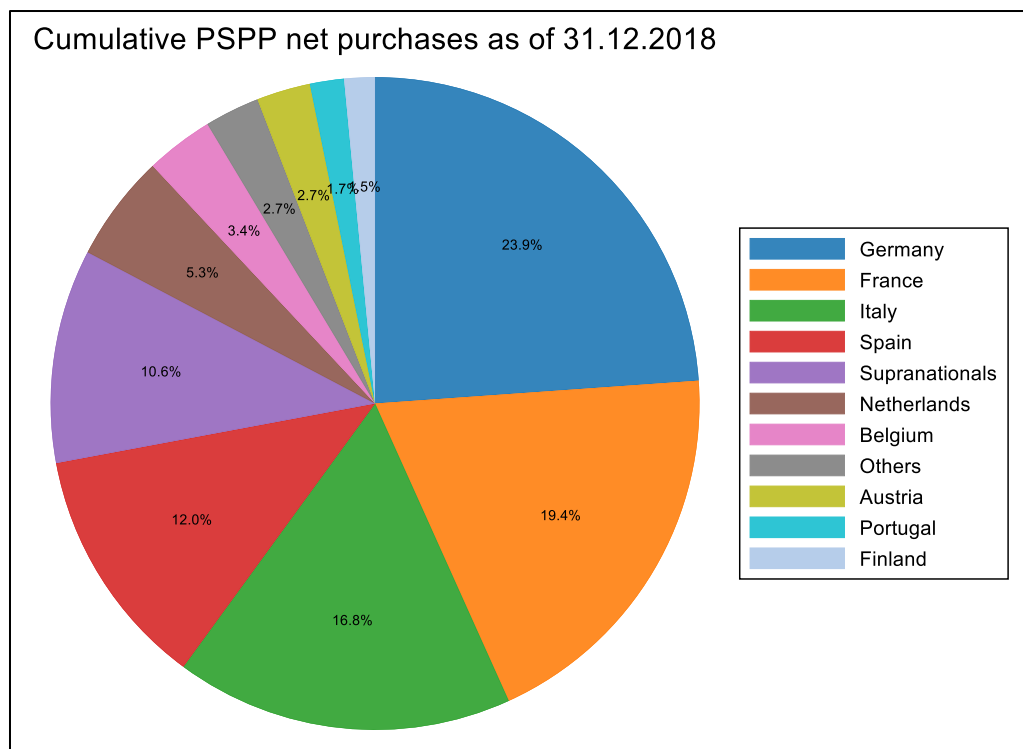


Figure 7 - Cumulative PSPP net purchases as of 31.12.2018. Data source: ECB

Greece is not included in the analysis despite its economy being bigger than Finland's or Portugal's over a sizeable timeframe within the sample. The reason for Greece's exclusion is that the country's government bonds were never eligible for the PSPP timeframe under consideration. However, Greece was indirectly affected by the Eurosystem's APP programme. The European Stability Mechanism (ESM) and its predecessor, the European Financial Stability Facility (EFSF), are key supranational institutions whose bonds are eligible for the PSPP program (ECB, 2019a). During Europe's sovereign debt crisis, ESM and EFSF lent money to Greece, Ireland, Portugal, Spain, and Cyprus; as of 30 June 2019, 53.6% of Greek public debt was still held by EFSF/ESM (ESM, 2019c). ESM, in turn, raises its funding needs on financial markets (ESM 2019a). While acknowledging this indirect linkage, for my modelling purposes it is more clear-cut to exclude data on Greece from the analysis. Furthermore, smaller Eurozone members in terms of both nominal GDP and PSPP purchases were also not included in the analysis due to (partial) lack of data and to keep in line with literature practice of focussing on a smaller set of "bigger" countries.

The following subsections describe in detail the variables used in the regressions to follow, discussing their economic significance and their use in literature. First, the dependent variable is described, followed by explanatory variables. In this section, the variables are presented in levels, i.e. without any transformation (which might be applied later on).

A table summarizing all data sources and containing data definitions and links to data sources can be found in the appendix 6a.

a. *yield*: Government bond yields

The key variable that I want to model are Euro Area sovereign bond yields using ECB's monthly data on long-term interest rates. Variable "***yield***" refers to secondary market yields of government bonds denominated in Euro with maturities "close to ten years" (see Figure 8). I am following relevant literature (e.g. Altavilla et al., 2015; Eser et al., 2019; De Santis, 2020 and many more) in focussing on bonds with a 10-year maturity due to the wide use within the financial industry of such securities as a benchmark, the relatively high liquidity of the secondary market for bonds with this maturity as well as due to data availability.

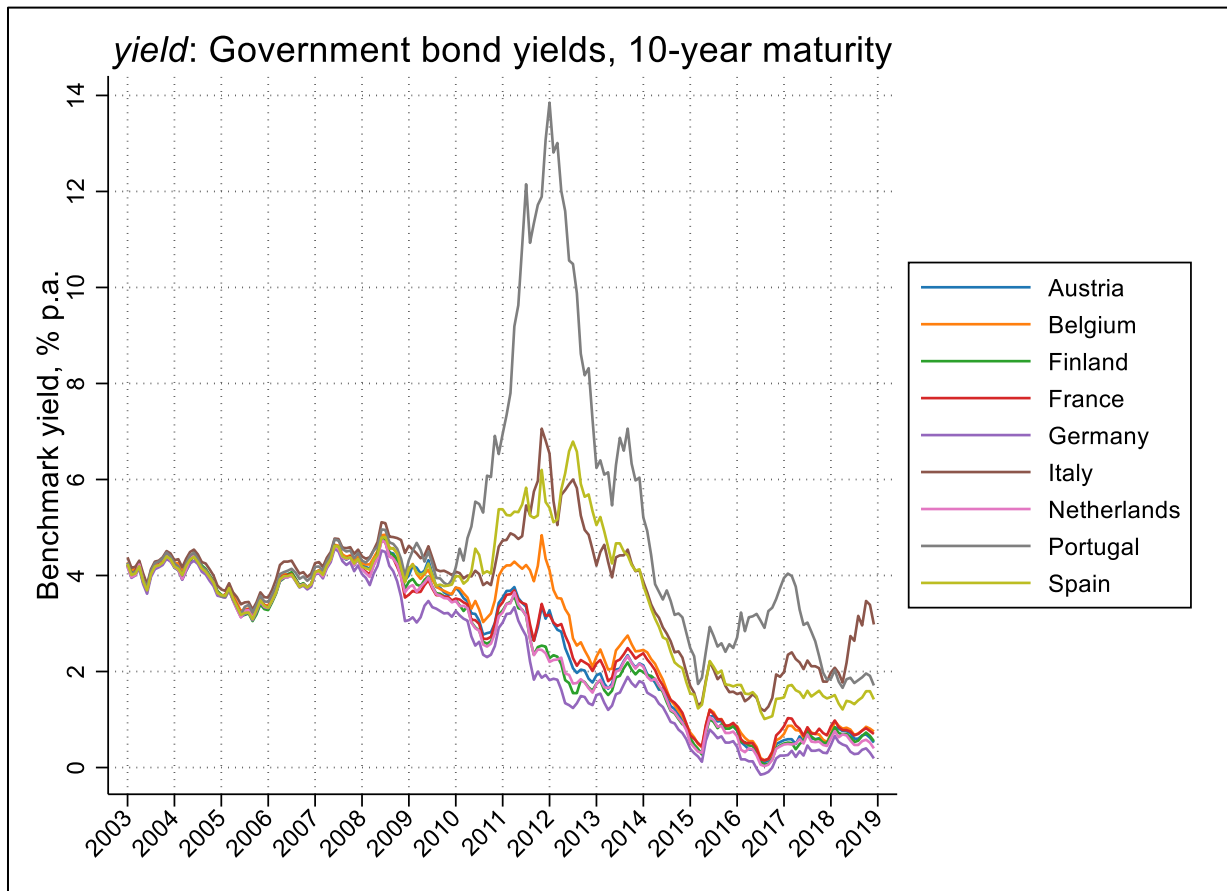


Figure 8 - *yield*: Government bond yields, 10-year maturity

Data is complete for the entire sample. As Figure 8 shows, yields vary considerably across time and countries. After a period of very low yield spreads between Euro Area economies in the early 2000s, yields started to diverge substantially after 2008. Overall, the lowest mean yields as well as minimum value are observed in Germany and the highest mean yield as well as maximum value in Portugal. Monthly averages of the daily yields are collected by individual national central banks and reported to the ECB. Data provided by the ECB is using the "redemption yields" definition (ECB 2022b). The formula used to calculate these yields from observable bond prices is documented in appendix 6c.

b. *bas*: Bid-Ask spreads for benchmark government bonds

As Fabozzi (2010) discusses, one of the risks bond investors face is liquidity risk, i.e. the risk of not being able to sell a bond close to its value. As discussed in section 2, economic literature has frequently used bid-ask spreads as a proxy for liquidity risk. The *bid price* is the price at which a buyer is willing to buy a bond, while the *ask price* is the price at which a bond holder is willing to sell at a given point in time. The larger the difference between these two prices (i.e. the *bid-ask price spread*) in a secondary bond market, the greater the liquidity risk the bond holder faces. A high bid-ask spread might in turn lead to investors expecting higher yields to compensate for that liquidity risk.

The relation between bond prices and bond yields is well known to be inverse: When bond prices increase, yields decrease and when bond prices decrease, yields increase. As Fabozzi (2010) discusses, the fundamental reason is that the current price of a bond represents the present value of all future cash flows associated with the bond until maturity, including coupon payments as well as principal repayment. Fabozzi defines the yield-to-maturity as “*the interest rate that will make the present value of the cash flows equal to the price*”. Hence, when bond prices fall, yields – the interest rate with which fixed future cash flows are discounted to the present value – rise, and conversely.

Typically, ask prices are higher than bid prices. Hence, bid yields are typically higher than ask yields as each side of the bond market tries to maximize its profits – sellers do so by asking for a higher price, while buyers attempt to secure the future cash flows associated with a given bond at a lower price.

I follow De Santis (2020) in using bid & ask bond yield-to-maturity data rather than bid & ask bond price data. One advantage of this approach as discussed by De Santis (2020) is that in this setup the basic units of both yields (the variable to be modelled) and bid-ask spreads is the same, namely % p.a. I use daily bid/ask yield data provided by “*Refinitiv Datastream*” on 10-year benchmark Euro Area government bonds (yield-to-maturity based on daily closing prices as per *Refinitiv* emailed comments). In order to use a comparable aggregation approach as for yields (*yield* data: monthly averages of daily yields), I calculate monthly averages of daily bid & ask yields and calculate the monthly bid-ask-spread variable “*bas*” defined as the difference between average monthly bid & ask yields.

	Observations	Mean	Std. deviation	Minimum value	Maximum value
Country					
Austria	189	0.0237	0.0174	0.0015	0.1056
Belgium	192	0.0161	0.0176	-0.0105	0.1045
Finland	191	0.0157	0.0160	-0.0546	0.0920
France	192	0.0130	0.0108	0.0037	0.0629
Germany	192	0.0091	0.0124	0.0024	0.0979
Italy	192	0.0112	0.0112	-0.0083	0.1002
Netherlands	192	0.0105	0.0110	0.0002	0.0915
Portugal	192	0.1196	0.2492	-0.0081	1.3073
Spain	192	0.0233	0.0269	-0.0072	0.1382
Total	1724	0.0269	0.0906	-0.0546	1.3073

Table 2 - Sample statistics for *bas*

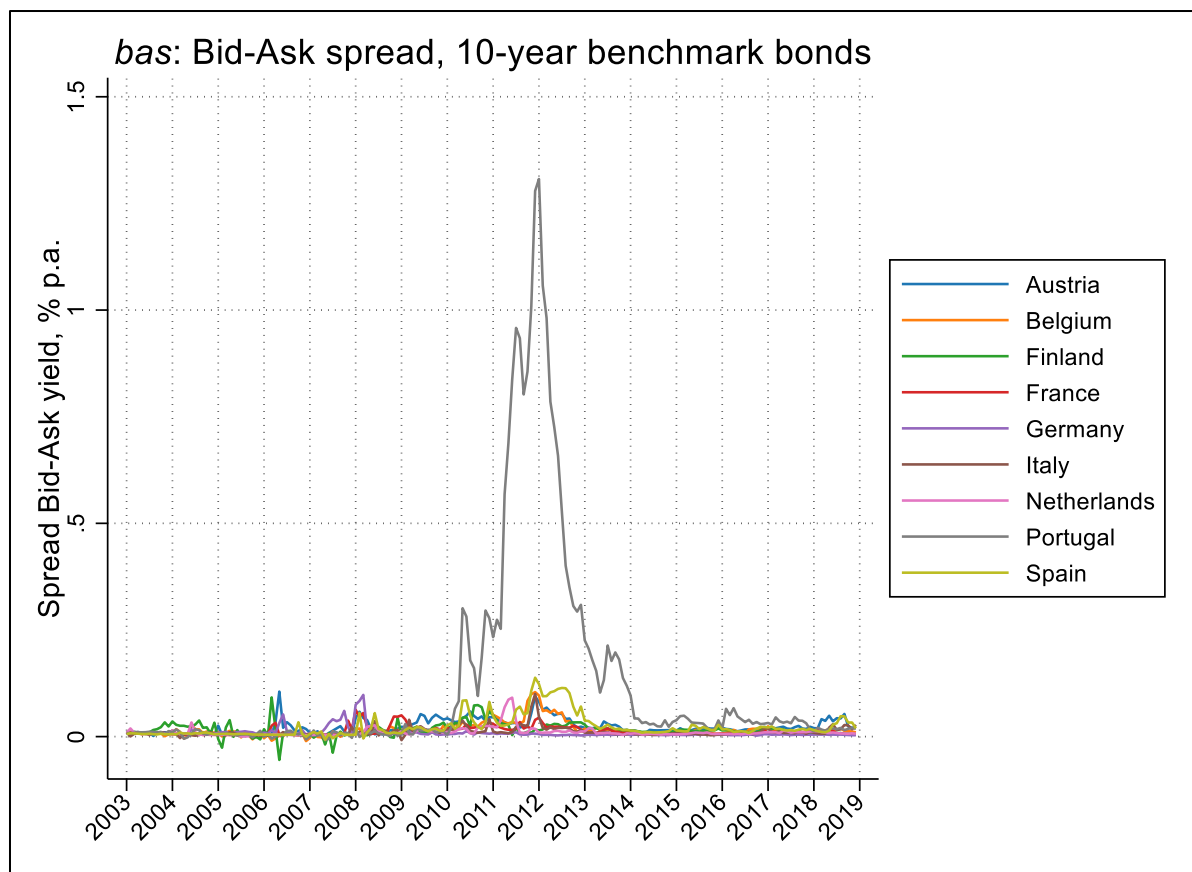


Figure 9 - *bas: Bid-Ask spreads of benchmark bonds*

Data is almost complete: there are 3 missing months for Austria (June-August 2007) and 1 missing month for Finland (April 2006). The lowest mean value is observed for Germany, while the bid-ask yield spread is highest throughout the sample for Portugal in early 2012. The data indicates that Portuguese government bond markets experienced substantial periods of illiquidity compared to EA peers during the European sovereign debt crisis, while - on the other end of the spectrum - the market for Germany's government bonds was characterized by high liquidity throughout the sample.

c. *rating*: Standard & Poor's ratings

I follow Kinateder & Wagner (2017) in using Standard and Poor's (S&P) sovereign credit ratings as proxy for market assessment of creditworthiness of sovereigns/default risk. As e.g. Fabozzi (2010) discusses, in market practice, three rating providers (Moody's Investor Service, Fitch Ratings and Standard and Poor's) dominate the global market for sovereign bond ratings. These ratings are designed to reflect the risk that bond holders will not receive in time and in full all cash flows owed to them by the issuer of the bond (i.e. here the governments of EA countries). The importance and relevance of the credit ratings produced by these agencies is also underlined by the fact that according to the eligibility criteria for PSPP, only bonds issued by governments with a *"credit quality assessment of at least Credit Quality Step 3 in the Eurosystem's harmonised rating scale expressed in the form of at least one public credit rating provided by an external credit assessment institution (ECAI) accepted within the Eurosystem credit assessment framework"* (EU Journal, 2015) were eligible for PSPP purchases. All three aforementioned credit rating agencies were registered as ECAI in 2011 (EBA, 2022).

The wide use of credit ratings within the financial industry motivates and warrants the inclusion of rating information into this thesis' model framework. If ratings do indeed reflect a sensible assessment of credit risk, sovereigns with lower ratings are expected to face higher refinancing costs and thus higher yields. I use Standard & Poor's long-term foreign currency sovereign issuer ratings for the Euro Area governments in my sample as published on the S&P website. As per S&P, their rating incorporates *"the factors that we believe affect a sovereign government's willingness and ability to service its financial obligations to nonofficial creditors on time and in full"*, and include assessments of various institutional, economic, external, fiscal and monetary factors, which result in sovereign ratings (S&P, 2022a).

I follow Kinateder & Wagner (2017) in assigning numerical values to S&P sovereign credit ratings and outlooks. Numerical values range from 0 to 21, where 21 represents the best attainable rating level (AAA) while 0 is the worst attainable level (sovereign default). The mapping procedure is described in appendix 6d. In order to use a comparable aggregation approach as for yields (yield data: monthly averages of daily yields), I define the variable ***rating*** as calculated monthly averages of daily values of numerical transformations of rating levels and outlooks. Figure 10 shows the levels of *rating* over the sample. Data is complete for the entire sample. Levels and number of changes of *rating* values vary significantly across countries; while Germany maintained the best attainable *rating* value for almost the entire sample period, the rating levels for Italy, Spain and Portugal changed particularly frequently, sharply deteriorating after the onset of the global financial crisis and subsequent European sovereign debt crisis and partly recovering later on.

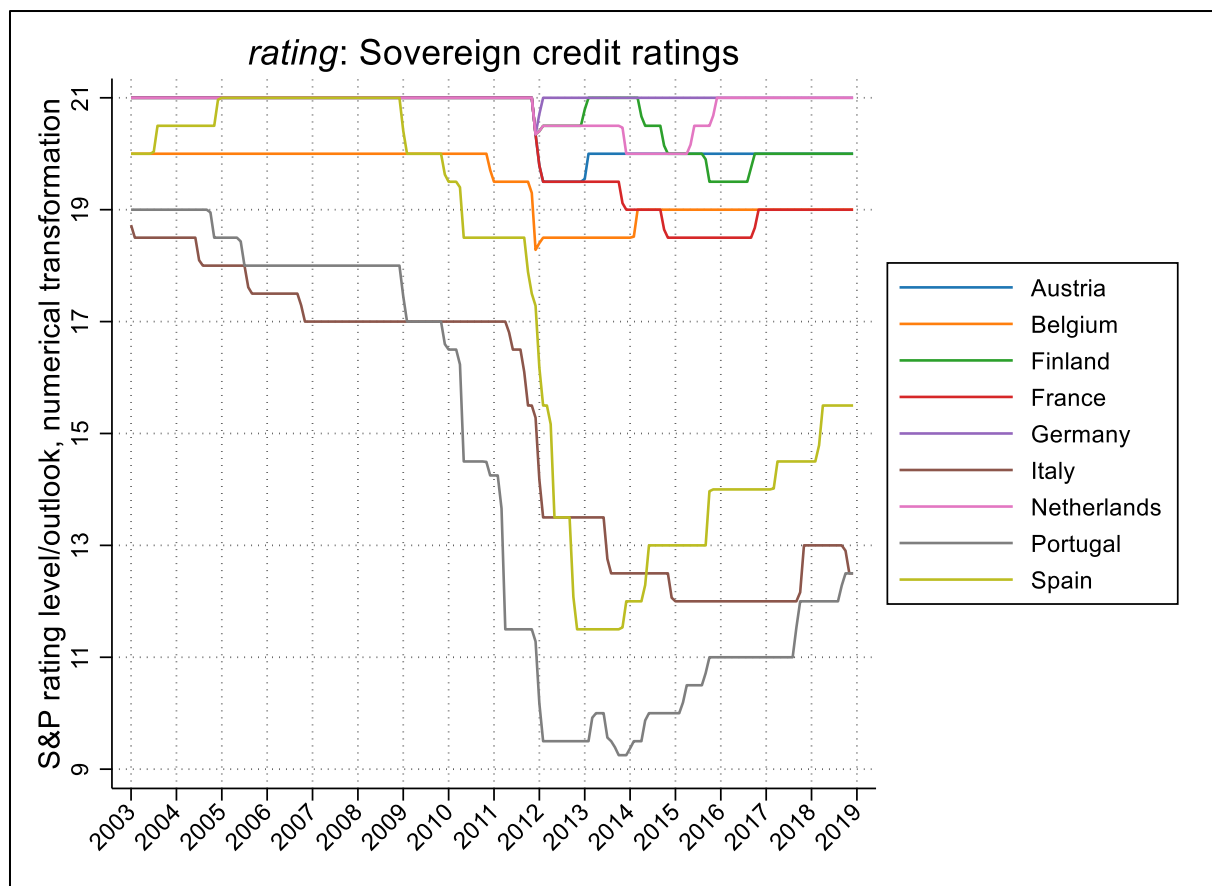


Figure 10 - rating: S&P sovereign credit ratings

d. *debt*: Public debt/GDP

As discussed in section 2, economic literature investigating macroeconomic and financial market drivers of sovereign bond yield spreads frequently use public debt/GDP as a measure of a sovereign's fiscal policy and debt sustainability. I follow this practice and use public debt/GDP data as published by the ECB (based on data transmitted by Eurostat) for my variable *debt*. Ceteris paribus, countries with higher public debt/GDP might face higher refinancing costs as investors might demand higher yield to compensate for longer-term debt sustainability concerns.

Public debt/GDP is reported as consolidated government debt at face value as a ratio to the annual moving sum of GDP. Data is reported for the end of each quarter. I map the quarterly data to monthly values in the following way: end-of-quarter data is used for the following 3 months, i.e. end of Q1 data is reported for April, May and June, end of Q2 data for July, August, September, end of Q3 data for October, November, December and end of Q4 data for January, February and March. This methodology takes into account the fact that public finance as well as GDP estimates are usually reported with a time lag. Hence, it makes sense to assume that market knowledge about public debt/GDP is updated once new estimates are published, but market participants do not have perfect knowledge of public debt/GDP in Q1 at the beginning or during Q1. Data is complete for the entire sample and variable *debt* is depicted in figure 11. Most Euro area

sovereigns' public debt trended upwards after the global financial crisis. Italy had the highest public debt/GDP throughout most of the sample, while Finland and the Netherlands had comparably low public debt/GDP throughout the sample; Spain's public debt/GDP shot up notably after 2008, stabilizing only after around 2013 while Portugal's public debt/GDP exhibited a similar dynamic, but at much higher levels.

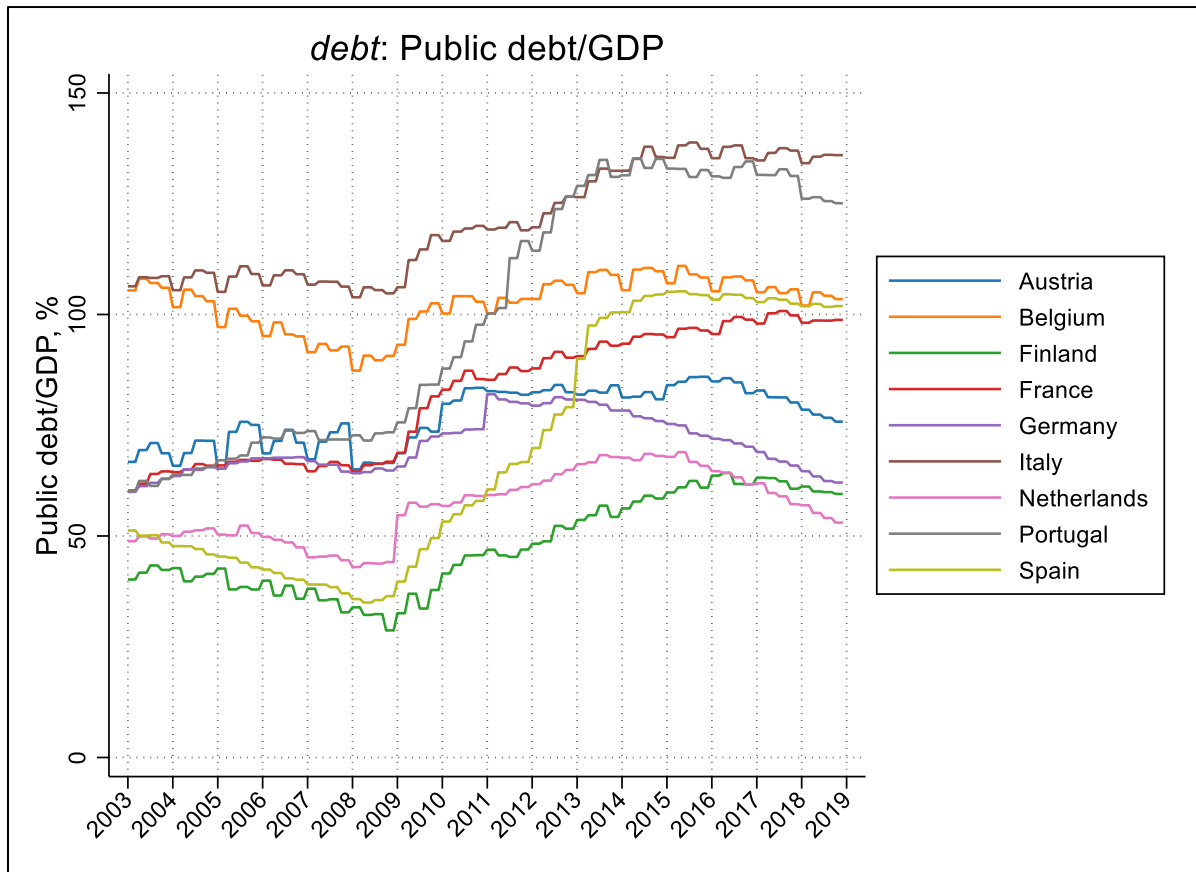


Figure 11 – debt: Public debt/GDP

e. *hicp*: Inflation

In economic literature investigating the macro-financial drivers of yield levels, authors have used various inflation-related measures, including cumulated inflation differentials vs. a benchmark economy to control for competitiveness between countries (Costantini et al, 2014) as well as inflation forecasts (De Santis, 2020) and actual inflation rates (Poghosyan, 2013) to control for a Fisher-type effect (i.e. to control for the extent to which yields react to changes in inflation). As Fabozzi (2010) discusses, bond investors are exposed to inflation risk (unless they buy inflation-linked bonds) as the real yields of bonds are dependent on both nominal coupon rates/bond prices as well as inflation. This relation warrants the inclusion of inflation data into the econometric model. Due to data availability, I follow Poghosyan (2013) in using actual inflation rates. I use data from Eurostat harmonized index of consumer prices (HICP) for individual countries (variable *hicp*) as well as the Euro Area in total (changing composition; variable *hicp_ea*). Eurostat provides monthly rates of change relative to the same month in the previous year.

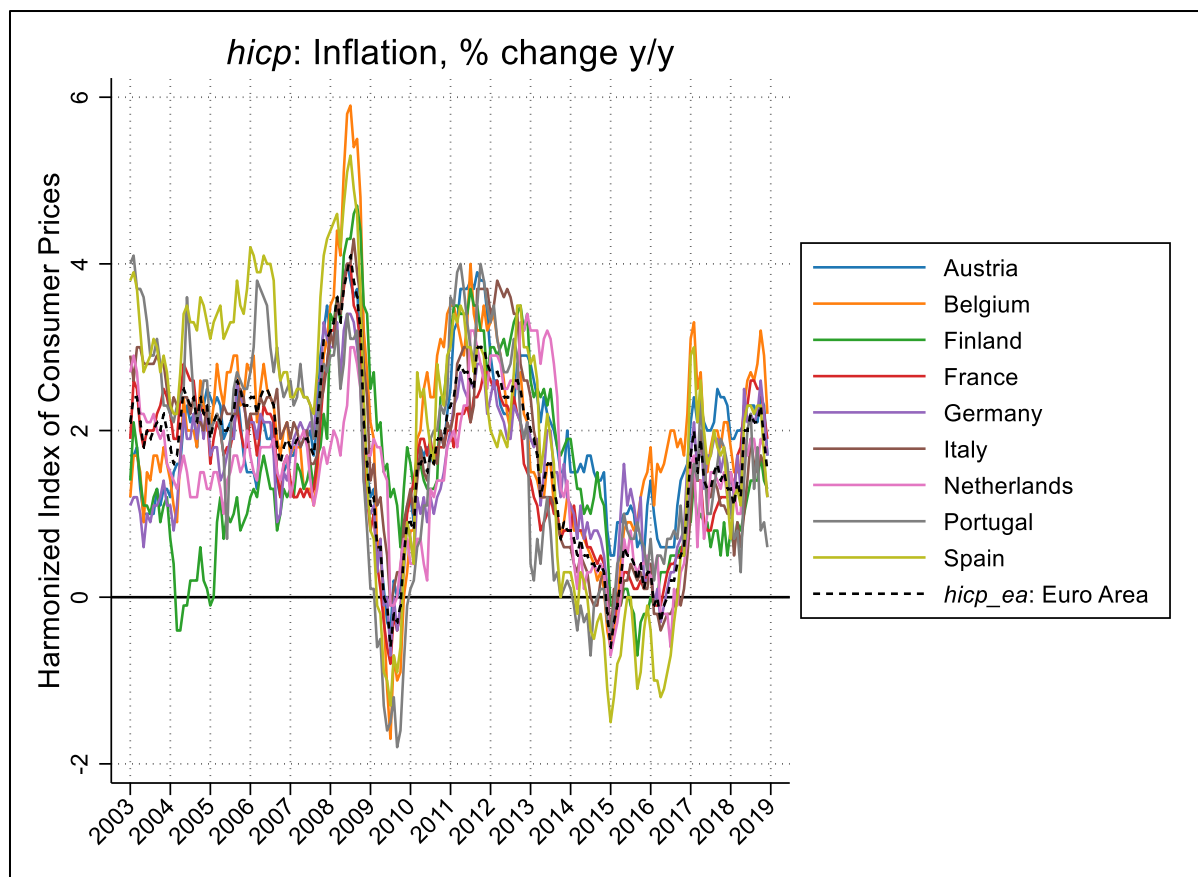


Figure 12 – hicp and hicp_ea: Inflation rates

Data is complete for the entire sample. The lowest mean inflation was observed in the Netherlands while the highest mean inflation was observed in Spain. As figure 12 suggests, inflation levels across countries differed substantially across various sample periods, complicating ECB monetary policy targeting the entire Euro Area.

f. esi: European Commission's Economic Sentiment Indicator

Following economic literature, I want to control for macroeconomic developments/business cycles in a given country. Poghosyan (2014) uses both changes in output growth (GDP growth rates) and potential growth as control variables, arguing that higher GDP growth rates could drive down yields as a bigger economy offers broader taxation possibilities to the sovereign, hence lowering default risk. Similarly, Afonso et al (2015) use the industrial production growth rates as a “proxy for the state of business cycle and captures the effect of economic growth on spreads”, arguing that sovereign risk increases when business cycles deteriorate. De Santis (2020) uses expected real GDP growth rates (forecast values for subsequent year) to capture the feedback loop between developments in the real economy and market expectations of future monetary policy decisions (economic downturns might lead to more expansionary monetary policy).

In order to use data that is readily available on a monthly frequency, I use the European Commission's Economic Sentiment Indicator (ESI) as a proxy for macroeconomic developments in a given country. As described in the ESI documentation (EC 2022), the

indicator for each EU country is based on monthly surveys across five different sectors (industry, services, construction, retail trade and consumers). Survey sample sizes vary across EU countries; all sample countries are covered. The formula and processes used to calculate the ESI from monthly survey data (15 questions covering the aforementioned 5 sectors) is documented in appendix 6e. The purpose of ESI is “to track GDP growth” by way of summarizing “the attitudes and judgements of a large number of economic actors”. ESI is “scaled to have a long-term mean of 100” where values “greater than 100 indicate an above-average economic sentiment, whereas values below 100 indicate a below-average position” (EC 2022). Data is complete for the entire sample. Figure 13 shows the evolution of variable *esi* across countries in the sample period. While there is some variety among countries, there is a clearly synchronized downturn amidst the global financial crisis in 2008/2009 and a generally positive sentiment towards the end of the sample period.

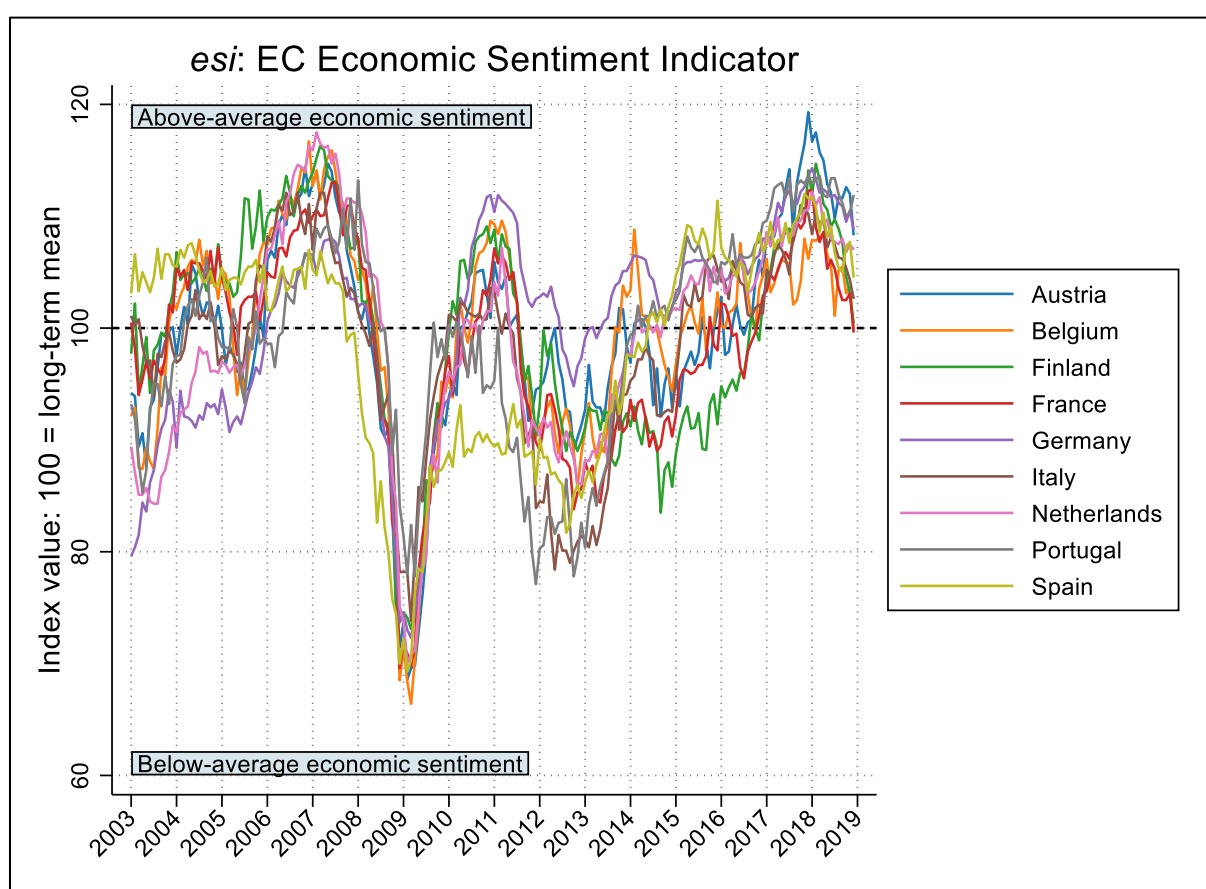


Figure 13 - esi: Economic Sentiment Indicator

[g. euribor_3m](#): 3-month Euro Interbank Offered Rate

Following economic literature, I want to use short-term market interest rates to control for effects of monetary policy and banks' funding conditions on sovereign long-term yields. As banks are major intermediaries and participants of government bond markets, their costs of accessing liquidity might have an influence on bond yields. In the relevant literature, Poghosyan (2014) uses real short-term rates in various countries to “measure the influence of monetary policy on long-term bond rates”. Kinatader and Wagner (2017) use the 3-month

Euribor as a proxy for short-term interest rates to “account for funding conditions”, pointing to earlier findings that “yield spreads are significantly driven by short-term interest rates.” De Santis (2020) uses “3-month Overnight Index Swap rate as a proxy of conventional monetary policy stance provided by the ECB”. I follow Kinateder and Wagner in using monthly nominal 3-month Euribor rate as published by the ECB (data provided by Refinitiv, which in turn uses data from the European Money Markets Institute or EMMI).

According to the current Euribor methodology (EMMI 2019), Euribor rates are based on data reported daily by a panel of 18 large European banks and aim to represent “the rate at which wholesale funds in euro could be obtained by credit institutions in current and former EU and EFTA countries in the unsecured money market.” The panel banks report rates for 5 different tenors (maturity: 1 week to 12 months) based on a three-tiered system, where rates from real wholesale unsecured money market transactions for the respective tenor (borrowing transactions made by reporting bank on the previous day) are seen as the most important tier of data to be used with priority for calculating Euribor. Following reporting by the panel banks, the reported rate data for each Euribor tenor is truncated by removing the highest 15% and lowest 15% of reported rates. Then, Euribor is calculated as the arithmetic average of the truncated data for a given tenor. Figure 14 shows the changes of variable *euribor_3m* over the sample period. Monthly averages of daily ECB key rates are also shown in the figure to motivate the idea that Euribor rates can be used as a rough proxy for ECB’s conventional monetary policy stance. *euribor_3m* fell sharply after 2008 in parallel with the substantial easing of the ECB’s policy stance (key rate cuts) and fell into negative nominal territory in 2015, following the ECB’s negative interest rate policy adapted in 2014.

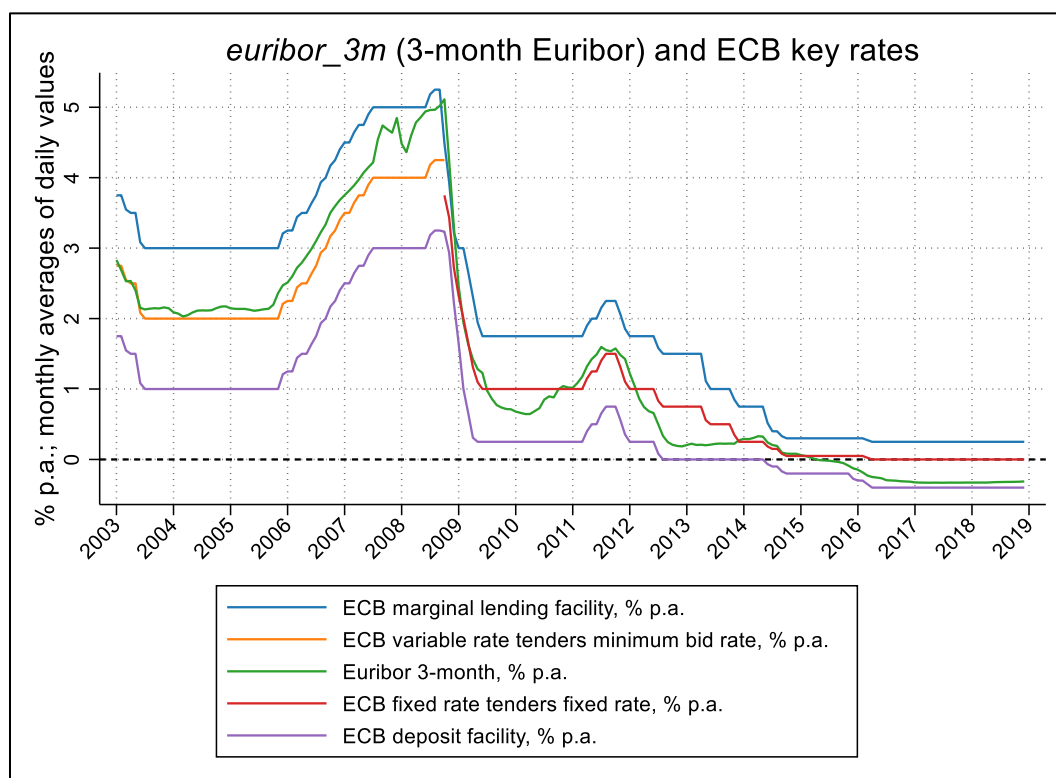


Figure 14 - *euribor_3m*: 3-month Euro Interbank Offered Rate and ECB key rates

h. *vix*: CBOE Volatility Index

Following economic literature, I want to control for global financial market perception of uncertainty. High uncertainty might influence the risk appetite of investors and cause them to seek assets perceived as safe, which could lead to higher demand for certain government bonds. Kinatder and Wagner (2017) discuss that many previous studies have used the Chicago Board Options Exchange Volatility Index (CBOE VIX) to “*measure global uncertainty*”; i.e. Afonso et al (2015) and Poghosyan (2014) use this index in their papers investigating the drivers of sovereign bond yields to “*capture the international risk factor*” and as a “*proxy of global economic uncertainty*”, respectively. VIX uses observable information on options listed for trading on the CBOE on the US “S&P 500” stock market index to calculate a proxy measure of expected volatility of the “S&P 500” index over the next 30 days (CBOE 2019). Higher VIX values can be interpreted as a proxy for larger expected stock index volatility in this time period. High stock index volatility, in turn, can be interpreted as a sign of high uncertainty. Given the global connectedness of financial markets, the inclusion of this widely used index based on US stock market as a proxy for global uncertainty sentiment seems warranted.

The formula used to calculate the VIX from observable option prices is documented in appendix 6f. Data is complete for the entire sample. Monthly values are calculated as the average of daily index values as obtained from the CBOE website. Figure 15 shows the evolution of variable *vix* over time; in the sample period, *vix* peaked in late 2008 amidst the global financial crisis and exhibits further notable spikes in 2010 and 2011.

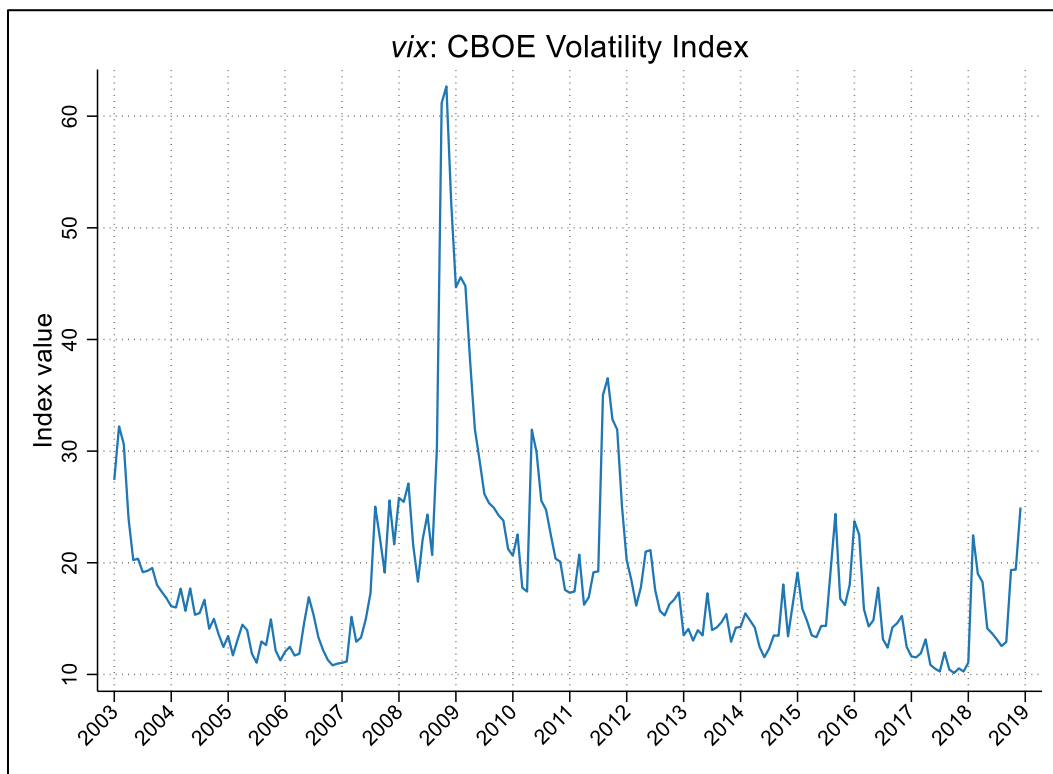


Figure 15 - *vix*: Chicago Board Option Exchange Volatility Index

i. PSPP & T00: Eurosystem's PSPP purchases & total Eurosystem assets

In order to measure for the flow effects of the PSPP, i.e. the effects of the Eurosystem's actual government bond purchases, I use monthly data published by the ECB ex post. Data is available for every month after and including March 2015, when the PSPP purchases were started, until the end of the sample period. Figure 16 visualizes variable *PSPP_p_i*, which denotes monthly net purchases in billions of Euros split by countries (issuers of purchases bonds). The vertical axis shows stacked purchases. Variable *PSPP_p*, also depicted in the graph, denotes total monthly net PSPP purchases across all eligible entities while the Eurosystem's total monthly net purchase limit for its Asset Purchases Programmes (APP) is shown as a dashed line (PSPP is a subprogram of APP; limit varying over time). The data shows (as discussed in section 1c) that PSPP purchases made up the bulk of the APP envelope purchases and that countries (bond issuers) included in the sample made up the bulk of targeted countries. Supranational entities account for the bulk of the difference between total PSPP purchases and PSPP purchases of sample countries (see figure 7) and "bigger" countries account for the bulk of purchases as per PSPP's framework setup.

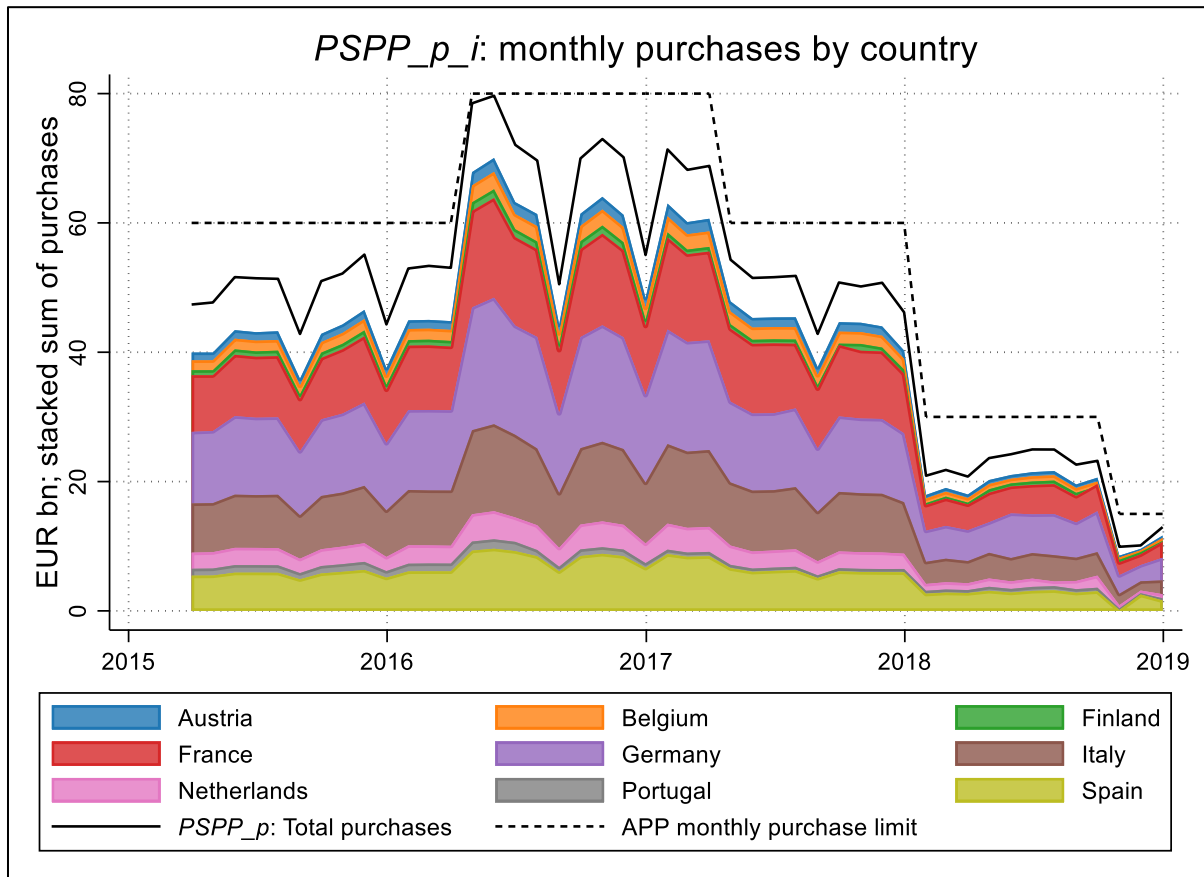


Figure 16 - *PSPP_p_i* and *PSPP_p*: Eurosystem's monthly PSPP purchases

For a brief discussion how the PSPP data relates to other Eurosystem balance sheet items, please refer to appendix 6g. There, also a graph of variable *T00* (total Eurosystem balance sheet assets, Eur bn) can be found. Following Kinatader and Wagner (2017), I also use this variable as a very loose overall proxy for the ECB's unconventional monetary policy.

Monthly PSPP data published by the ECB is not split by bond maturity, i.e. it includes the total of net bond purchases for a given country across all included bond maturities. For a brief discussion of the weighted average maturity of the Eurosystem's PSPP holdings, please refer to appendix 6i.

j. Sample statistics and graphs

Variable		Mean	Std. dev.	Min	Max	Observations	
yield	overall	3.083513	1.738642	-.15	13.85	N =	1728
	between		.7708164	2.345417	4.721719	n =	9
	within		1.579377	.021794	12.21179	T =	192
bas	overall	.0269288	.0906418	-.0546429	1.307318	N =	1724
	between		.0351427	.0090786	.1195796	n =	9
	within		.0843508	-.100746	1.214667	T-bar =	191.556
debt	overall	80.72177	27.00723	28.693	138.834	N =	1728
	between		23.38023	47.64222	120.7321	n =	9
	within		15.59524	40.64224	116.5402	T =	192
rating	overall	18.82426	3.10447	9.25	21	N =	1728
	between		2.566152	14.27148	20.99509	n =	9
	within		1.944451	13.03451	23.55279	T =	192
esi	overall	99.34178	9.450484	66.4	119.3	N =	1728
	between		.7503428	98.40208	100.5349	n =	9
	within		9.423953	65.82616	118.8027	T =	192
hicp	overall	1.707002	1.164463	-1.8	5.9	N =	1728
	between		.1957755	1.494792	1.970313	n =	9
	within		1.149732	-1.96331	5.63669	T =	192
eurib~3m	overall	1.382189	1.565869	-.3304	5.1131	N =	1728
	between		0	1.382189	1.382189	n =	9
	within		1.565869	-.3304	5.1131	T =	192
vix	overall	18.63607	8.347701	10.12545	62.66895	N =	1728
	between		0	18.63607	18.63607	n =	9
	within		8.347701	10.12545	62.66895	T =	192
PSPP_p_i	overall	4.545186	4.521262	-.355	19.573	N =	414
	between		4.053161	.7202826	11.27298	n =	9
	within		2.409084	-4.223792	12.84521	T =	46
T00	overall	2147.376	1112.015	772.812	4667.064	N =	1728
	between		0	2147.376	2147.376	n =	9
	within		1112.015	772.812	4667.064	T =	192

Table 3 - Summary statistics

Table 3 gives summary statistics (sample mean, standard deviation, minimum and maximum values) of the panel data consisting of 9 countries (n) and 192 months from January 2003 until December 2018 (T). All time series are complete except for bid-ask spreads as discussed above. PSPP purchases are available only from March 2015 onwards after the start of the programme. The table includes **overall** statistics ($\bar{x}$, e.g. yield: mean yield calculated across all countries and months) and a decomposition into **between** statistics ($\bar{x}_i$, e.g. yield: mean yield per country across all months, resulting 9 mean values used to calculate statistics) as well as **within** statistics ($x_{it} - \bar{x}_i + \bar{x}$, e.g. yield: individual value – mean yield of given country + mean of all yields, resulting 1728 values used to calculate statistics).

Table 4 summarizes **correlations** between the main variables in level over the entire sample period. *rating* exhibits a highly negative correlation with yields, which seems logical (lower values of *rating*, i.e. bad ratings correlate with higher yields, reflecting higher sovereign risks). The high correlation between *yield* and *bas* (bid-ask-spread) as well as *debt* (Public debt/GDP) can be interpreted in the context of episodes of simultaneous bond market illiquidity and high yields as well as higher yields for highly indebted countries during the European sovereign debt crisis. Short-term money market interest rates (*euribor_3m*) have a somewhat negative correlation with economic sentiment and inflation. Interestingly, higher PSPP purchases seem to correlate with episodes of rather low bid-ask-spreads, which could mean that the Eurosystem's secondary market purchases increased market liquidity.

	yield	bas	debt	rating	esi	hicp	euribor_3m	vix	PSPP_p_i	T00
yield	1.0000									
bas	0.5628	1.0000								
debt	0.7738	0.3481	1.0000							
rating	-0.8888	-0.4493	-0.8848	1.0000						
esi	0.2056	0.1969	0.0548	-0.1570	1.0000					
hicp	-0.0634	0.1630	-0.0405	0.1912	0.4138	1.0000				
euribor_3m	0.0398	-0.0712	0.0502	-0.0505	-0.5167	-0.5321	1.0000			
vix	0.0331	-0.0590	-0.0051	-0.0031	-0.2384	-0.2555	0.3118	1.0000		
PSPP_p_i	-0.1505	-0.4277	0.0769	0.0009	-0.0617	-0.2611	0.0439	-0.1242	1.0000	
T00	0.0063	0.0713	-0.0688	0.0627	0.6293	0.6798	-0.8677	-0.2233	-0.2103	1.0000

Table 4 - Correlation matrix

4. Methodology and estimation

a. Univariate analysis

Methodology

I follow the univariate time series model-building cycle as outlined by Verbeek (2017; chapter 8). Verbeek argues that a pure time series approach can be a powerful modelling building block. Beyond that, I find it worthwhile to analyze the dependence of the time series used in my panel on themselves to potentially gain further insights into their joint multivariate/panel dynamics. The aim of this section is to model as many individual variables in my dataset as possible as country-by-country univariate ARIMA processes. In the following, I use Verbeek's notation, definitions and conclusions on univariate time series to briefly explain the underlying methodology and sequence of analytical steps to follow.

A **general ARMA(p,q) process** for the univariate series Y_t ($Y_t = y_t + \mu$) can be written as $Y_t = \mu + \theta_1 y_{t-1} + \theta_2 y_{t-2} + \dots + \theta_p y_{t-p} + \varepsilon_t + \alpha_1 \varepsilon_{t-1} + \alpha_2 \varepsilon_{t-2} + \dots + \alpha_q \varepsilon_{t-q}$ where μ is the mean, ε_t is a white noise process and the θ & α coefficients respectively describe the autoregressive & moving average specifications with lag orders p and q . Equivalently, ARMA(p,q) can be written as $\theta(L)y_t = \alpha(L)\varepsilon_t$ where $\theta(L)$ and $\alpha(L)$ are lag polynomials of order p and q respectively.

A process $\{Y_t\}$ is defined as **weakly stationary** if the following three conditions are fulfilled for all t : (1) $E\{Y_t\} = \mu < \infty$ and (2) $V\{Y_t\} = E\{(Y_t - \mu)^2\} = \gamma_0 < \infty$ and (3) $cov\{Y_t, Y_{t-k}\} = E\{(Y_t - \mu)(Y_{t-k} - \mu)\} = \gamma_k, k = 1, 2, 3 \dots$. The conditions imply that the process has a finite & time-invariant mean and variance and that the covariance function γ_k is only dependent on the time lag between two observations. An ARMA(p,q) process is weakly stationary if the characteristic AR polynomial $\theta(L)$ is invertible, or equivalently the roots $z_1 \dots z_p$ of the characteristic equation ($\theta(z) = 1 - \theta_1 z - \theta_2 z^2 - \dots - \theta_p z^p = 0$) are larger than 1 in modulus. One situation where this is not the case is if one root z_i is exactly equal to 1, i.e. the presence of a **single unit root**. However, in this special case the non-stationarity can be removed by first-differencing, after which the process Δy_t can be described by a stationary ARMA model and y_t is said to be **integrated** of order 1 or $I(1)$.

In order to understand whether a time series can be appropriately described by an ARMA(p,q) model, testing for the presence of a unit root is an indispensable first step. In case there is a unit root, the series becomes stationary after taking first differences only. I use the (augmented) **Dickey-Fuller test**, which allows to test the H_0 of non-stationarity due to a single unit root. The Dickey-Fuller procedure is based on rewriting the AR(p) process for Y_t as $\Delta Y_t = \delta + \pi Y_{t-1} + \gamma t + c_1 \Delta Y_{t-1} + \dots + c_{p-1} \Delta Y_{t-p+1} + \varepsilon_t$. The drift term δ and deterministic trend term γt are included when appropriate (H_0 : process follows random walk with drift; visual inspection of the data indicates a positive or negative trend over time). Here, $\pi = \theta_1 + \dots + \theta_p - 1$ where θ_i are the autoregressive coefficients from the AR(p) model. When $\pi=0$, it is implied that $z=1$ is a root of the characteristic equation and that there is a single unit root. Hence, the $H_0 \pi=0$ (presence of a single unit root) must be rejected to allow for the hypothesis

of stationarity ($H_1: \pi < 0$). When the H_0 of presence of a single unit root cannot be rejected at a given significance level, a reasonable next step is first-difference the time series and check once more via Dickey-Fuller test whether this has removed the non-stationarity. The Dickey-Fuller test framework relies on the assumption that the lag order p is known. As this is not the case a priori in an empirical context, Verbeek advises to start from a high lag order (using fewer lags than $AR(p)$ risks inconsistency) or use information criteria to select the lag order specification. I run Stata code to include up to 12 lag terms in the Dickey Fuller regression defined above (all months within 1 year prior to current data point are included).

After performing unit root tests, my **primary model selection strategy** for choosing $ARMA(p,q)$ is to minimize **Schwarz's Bayesian information criterion**. I run Stata code to estimate $ARMA(p,q)$ models for a given p and q up to a certain threshold and compute Akaike's and Schwarz's Bayesian information criteria (AIC & BIC), which both offer a tradeoff between parsimony (using small values for p and q) and model fit as measured by the log-likelihood value. I choose the model specification with the lowest BIC since Verbeek argues that BIC "*can be preferred*" to AIC as it tends to select the correct model when $T \rightarrow \infty$ and following the overall argument in the univariate context in favor of parsimonious models. I perform **additional checks** for model adequacy by "overfitting" and residual analysis. Overfitting: I add +1 to either p or q and check the significance of the estimated additional parameters to see whether their inclusion is warranted. The error terms of $ARMA(p,q)$ should be close to white noise, i.e. in a good model they should not contain any dynamics. I run a Ljung-Box portmanteau test on the residuals of my model specification. The test is based on the statistic $Q_k = T(T+2) \sum_{k=1}^K \frac{1}{T-k} r_k^2$ where r_k refers to the estimated autocorrelation coefficients of the residuals and K is the Stata default value $K = \min(n/2-2, 40)$. Under the H_0 of a correct model, $Q_k \sim \chi_{K-p-q}^2$. If H_0 is not rejected, I conclude the analysis, while in the (rare) cases where the Ljung-Box test rejects, I opt to additionally estimate the model specification suggested by the minimal AIC and repeat the additional checks (overfitting, Ljung-Box test).

Results

Government bond yields (*yield*)

Visual inspection of *yield* (see figure 8) indicates that the hypothesis of stationarity does not make much sense due to the possible presence of a (potentially complex) time trend. I use the Dickey Fuller unit root test, allowing for a random walk with drift and/or trend under the H_0 by including both terms in the Dickey-Fuller regressions. Using a general-to-specific approach, I start from a high number of augmenting differenced terms (lags) to check whether the H_0 of non-stationarity can be rejected. For all countries in the sample, the H_0 that there is a unit root in *yield* in levels cannot be rejected with reasonable assumptions on the process under the H_0 . This is in line with Verbeek, who notes that H_0 is often hard to reject for interest rates due to their high persistence. Here, I am not trying to make any statements on whether the existence of a unit root process in interest rates is a realistic conclusion, I am merely interested in modelling *yield* as good as possible using ARIMA models.

Hence, I proceed with country-wise first-differenced series (*D.yield*) shown by Figure 23 in appendix 6j. Again, I use the Dickey-Fuller test, checking both for a H_0 with or without a drift term; however, here I disregard the case of existence of a deterministic time trend under the H_0 as visual inspection seems to rule out this possibility. I find that for almost all countries the H_0 can be rejected at the 1% significance level both irrespective of inclusion of a drift term; for Italy, Portugal and Spain, there are a small number of cases where the H_0 is not rejected at the 1% significance level when including 9+ lag terms, however, even in those cases, for lower numbers of lag terms and larger significance levels (5% or 10%), the H_0 is rejected. Hence, I proceed with the conclusion that country-by-country *yield* can be considered $I(1)$.

I use Stata code to estimate ARMA models for *D.yield* for all combinations of p and q up to 5 and tabulate AIC and BIC, selecting for the lowest value among the criteria respectively. While AIC tends to recommend “larger” models with up to 5 p and 4 q , BIC recommends ARMA(0,1) models for all countries but Portugal (ARMA(3,0)). For Italy and the Netherlands, both criteria favor ARMA(0,1). I follow the BIC model specification and estimate ARMA(0,1) for *D.yield* (ARIMA(0,1,1) for *yield*) for all countries but Portugal & Spain. Lastly, I “overfit” the model by adding one p or q parameter respectively and check the significance of the additional parameters; while in case of Belgium and Portugal, the case for an additional moving average term can be made, in all other cases the additional parameters are not significant. I test whether the residuals are similar to white noise using the Ljung-Box test; the H_0 is not rejected in all cases except Spain, for which case the Ljung-Box test finds not residual autocorrelation after estimating the AIC-selected specification. Appendix section 6j “univariate analysis” includes a summary table of these results and the estimated coefficients.

I use Stata to compute the “model fit” for the original series *yield* (Stata command for one-step-ahead forecasts after ARIMA estimation: *predict fit_yield, y*). Overall, my approach seems to do a reasonable job at approximating the value of yield in a given month (see figure 17); however, the models have difficulty in fitting periods with bigger volatility, such as in Italy and especially Portugal & Spain around 2011-2013. Furthermore, in a univariate context, I cannot make any statements about the impact of other macroeconomic variables on the level government bond yields. An extension of the univariate analysis could investigate more complex reasons of non-stationarity other than single unit roots or (G)ARCH models to capture the volatility around 2011-13. However, since I am more interested in the interaction of the various variables in my dataset, after briefly describing the results for my univariate analysis of other variables in the dataset (obtained in analogy to the process described above), I move on to multivariate panel analysis.

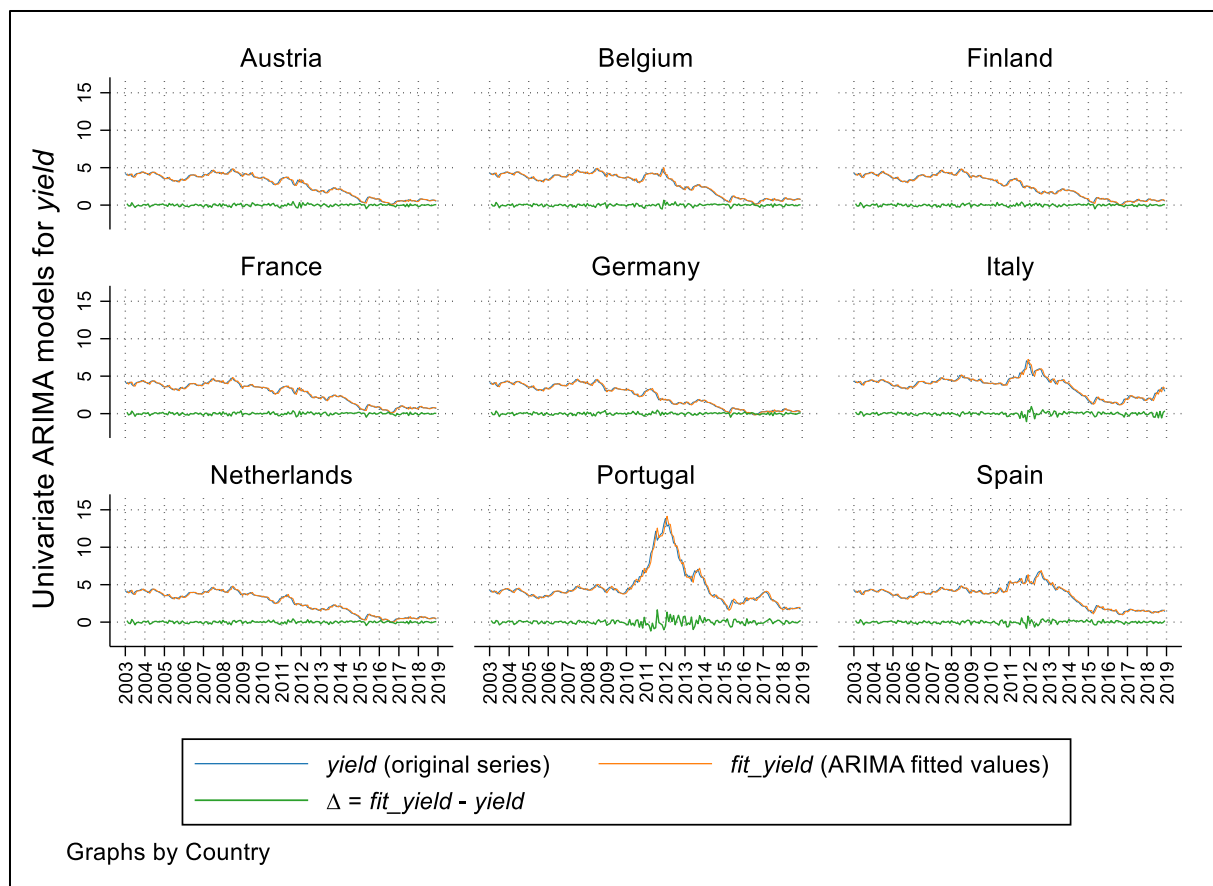


Figure 17 - Univariate ARMA models for yield

Bid-ask spread (*bas*)

For a low number of lag terms (as indicated by the PACF) and with/without drift under H_0 , the Dickey Fuller tests reject the H_0 . Estimating ARMA models for the first-differenced series sometimes reveals “over-differencing” (sum of MA coefficients close to 1). Hence, I proceed with the undifferenced series. Model specification selection through information criteria reveals a more heterogeneous picture across countries than for *D.yield*, with the preferred models including up to 7 AR/MA terms. I opt for the more parsimonious BIC-selected models except for Finland, France and Spain where the Ljung-Box test reveals remaining dynamics in the residuals of these models. I additionally estimate the models suggested by AIC with high lag orders, which results in residuals close to white noise. Overspecification also confirms the chosen specifications (see summary table in the appendix). Again, the model fit seems reasonable for most periods/countries except for episodes with high volatility (e.g. Portugal 2011/12).

CBOE volatility index (*vix*)

Visual inspection of *vix* (monthly average over daily index values; common for all countries) shows a non-zero mean but no deterministic time trend; Dickey Fuller tests with or without drift term under the H_0 reject the H_0 for low numbers of lags. Hence, while a stationary daily volatility index might seem counterintuitive, the time aggregation to the monthly level might

make the ARMA modelling approach in levels more viable. I proceed to analyze the undifferenced *vix* series. Both AIC and BIC favor ARMA(1,1). For this specification, all estimated coefficients are highly significant and the overfitted additional coefficients are not significant; the residuals seem similar to white noise (ACF exhibits no autocorrelation and Ljung-Box test does not reject H_0). I conclude that *vix* can be reasonably well modeled by ARMA(1,1); see appendix for a figure of model fit/residuals.

Harmonized Consumer Price Inflation (*hicp*)

ACF/PACF indicate a high persistence and visual inspection of the time series shows a non-zero mean but no clear deterministic time trend. The Dickey Fuller test reject the H_0 (unit root) when including drift under the H_0 at the 1% or 5% significance level for all countries; when excluding a drift term under the H_0 , the test fails to reject when a low number of lags is included, but becomes more likely to reject at the 10% significance level as the number of (“augmenting”) lag terms is increased. Hence, an argument for either $I(1)$ or $I(0)$ with large persistence can be made. I proceed to analyse the *hicp* series as $I(0)$. After estimating the more parsimonious model specification determined by BIC, for some countries the Ljung-Box test rejects the H_0 . In those cases, I also estimate the specification determined by AIC, which tend to include many AR terms (up to 12). Even for those models, overfitting sometimes suggests adding autoregressive and/or moving average terms (estimated additional coefficients are significantly different from 0). I conclude that inflation in the Euro Area seems to be highly persistent, as also indicated by the estimated sum of autoregressive coefficients, which is very high for some countries (up to 0.98).

Economic sentiment indicator (*esi*)

Similarly to inflation, sentiment indicator *esi* shows relatively high persistence. A possible interpretation is that economic sentiment exhibits bouts of above- or below-average levels which tend to last; at the same time, the index is designed to eventually return to its mean of 100 (“average sentiment”). The Dickey Fuller test rejects the H_0 (unit root) with drift under the H_0 at the 1% or 5% significance level for all countries; when excluding the drift term under the H_0 , the test fails to reject when a low number of lags is included, but becomes more likely to reject at the 10% significance level as the number of (“augmenting”) lag terms is increased. Again, an argument for either $I(1)$ or $I(0)$ with large persistence can be made. I proceed to analyse *esi* as $I(0)$. Model selection via the “more parsimonious” BIC yields models with at most 4 AR and/or 3 MA terms. Overspecification as well as residual analysis via Ljung-Box supports the chosen models. While high persistence is demonstrated by the sum of estimated autoregressive coefficients, which ranges from 0.96 (Netherlands) to 0.99 (Spain), all estimated AR parameters satisfy the stability condition (characteristic roots larger than one in modulus). Hence, I conclude that ARMA models for the undifferenced series can be applied (for details see summary table in the appendix).

Euribor (*euribor_3m*)

Similarly to *yield*, visual inspection of the series, ACF and PACF indicates that the hypothesis of stationarity does not make much sense due to the presence of a time trend. The Dickey Fuller tests fail to reject unit root hypothesis regardless of assumptions on drift/trend. Hence, I proceed to analyse first-differenced *D.euribor_3m*. Now, the Dickey Fuller tests reject the H_0 regardless of assumptions on drift/trend (at least 5% significance level for up to 10 lags). Both AIC and BIC support ARMA(5,4). The estimated AR and MA parameters satisfy the stability and invertibility conditions respectively and are highly significant. The residuals seem similar to white noise (ACF exhibits no autocorrelation and Ljung-Box test does not reject H_0). Hence, I conclude that *D.euribor_3m* can be reasonably well modeled by ARMA(5,4); see appendix for a figure of model fit/residuals.

Other variables: *rating*, *debt*, *PSPP_p_i* and *T0*

There are variables in my dataset for which a univariate ARIMA framework does not make much sense. Sovereign ratings (*rating*) stay constant for many countries for years, and in most cases exhibit no month-month fluctuations which would justify assuming a white noise error term. Assuming that the data-generating process can be modelled by an ARIMA framework would not align well with reality where the actual sovereign process is based on models and analysts' judgement. Also, sovereign debt as a % of GDP (*debt*) available on a quarterly basis and used in this dataset as monthly values (values updated once new data is available, constant in intermittent months) is ill-suited for ARIMA analysis due to a lack of month-on-month fluctuations. Lastly, PSPP monthly purchases on a country-by-country basis (*PSPP_p_i*), which are available for the timeframe 2015-2018, were determined directly by policy decisions within a framework of rules and restrictions as outlined above, and hence modelling this variable using ARIMA would not align well with the actual data-generating process.

As a last step within the univariate framework, I analyse as a very loose proxy for the ECB's unconventional monetary policy the Eurosystem's total assets (*T0* - see figure 18 in the appendix). Clearly, this variable is not stationary and a time trend seems to be present. Dickey Fuller tests reject the H_0 regardless of assumptions on drift/trend. After first-differencing, the series becomes stationary. Using BIC for model selection, ARMA(1,1) for *D.T0* is chosen as the model which best fits the data. The estimated coefficients are highly significant; overspecification and residual analysis confirms the model choice. However, unsurprisingly, this framework fails to fully capture large swings in Eurosystem assets, in particular during 2008/9 (see appendix for a figure of model fit/residuals).

Conclusions

Univariate analysis has helped to find useful conclusions about the stationarity of the variables in my dataset. While stationarity in levels can be defended for Bid-Ask spreads (*bas*), inflation (*hicp*), volatility index *vix* and the Economic Sentiment Indicator (*esi*), the series of government bond yields (*yield*), short-term interbank rate Euribor (*euribor_3m*) and total assets of the Eurosystem (*T0*) must be first-differenced to obtain stationary time series. Furthermore,

univariate analysis has affirmed the intuition that the dataset contains many macroeconomic variables which exhibit a rather high persistence, such as *yield* and *euribor_3m*, but also *hicp* and *esi*. For *hicp* and *esi*, the sum of estimated autoregressive coefficients can be high, pointing to borderline cases where first-differencing is also a viable path. There are also variables in my dataset for which the data-generating process follows conditions not particularly well suited for analysis within the ARIMA framework, such as *rating*, *debt* and *PSPP_p_i*. Further drawbacks of the univariate ARIMA approach are the inherent inability to model episodes of higher volatility around 2011-13 (even after first-differencing) as well as the lack of insights on the interdependencies and relations between various variables in my data set. Hence, I move on to multivariate panel analysis.

b. Static panel modelling

Methodology

In the following I use the notation and conclusions from Baltagi (2013; chapters 2-5) and Verbeek (2017, chapter 10) to briefly outline the methodology used in the static panel modelling discussed below. In a first step, I use **pooled OLS** (POLS) ignoring the panel structure of the data. This simplified approach, which imposes identical intercept and slope coefficients for all individuals and time periods, serves as a baseline for comparison. The model can be written as $y_{it} = \alpha + \beta'X_{it} + u_{it}$ ($i = 1, \dots, N$; $t = 1, \dots, T$) where α is a scalar, β is the $k \times 1$ matrix of coefficients and X_{it} is the it -th observation on k explanatory variables. In pooled OLS, assumptions on the error term include that u is independent across i and t , $E(u) = 0$ and $var(u) = \sigma^2$. Since these assumptions are unrealistic in a panel context where error terms from different periods from the same country are likely to be correlated, I proceed to estimate one-way error component models.

In the **fixed effects (FE)** model $u_{it} = \mu_i + v_{it}$ where μ_i are the country-specific parameters with the identifying condition $\sum_{i=1}^N \mu_i = 0$, the remainder disturbances v_{it} are assumed to be $IID(0, \sigma^2)$ and the X_{it} are assumed to be independent of the v_{it} for all i and t . The least-squares dummy variable (LSDV) regression can be written in matrix form as $y = \alpha\iota_{NT} + X\beta + Z_\mu\mu + v$ where y is the NT vector of dependent variables, $\alpha\iota_{NT}$ is a NT vector of constants, X is the $NT \times K$ matrix of regressor observations, β is the K vector of slope coefficients, Z_μ is the $NT \times N$ matrix multiplied by the N vector μ of unobserved country effects and v is the NT vector of disturbances. Equivalently, β can be estimated by application of the Frisch-Waugh theorem by pre-multiplying with the purging matrix Q , which adjusts all observations for the country-specific time averages. OLS in the FE model may be biased, inconsistent and/or inefficient as the Gauss-Markov conditions are not met since $E(u_{it}) = \mu_i \neq 0$.

In the **random effects (RE)** model, $u_{it} = \mu_i + v_{it}$, and assumptions include: μ_i is random with $\mu_i \sim IID(0, \sigma_\mu^2)$ and $v_{it} \sim IID(0, \sigma_v^2)$ for $i=1, \dots, N$ and $t=1, \dots, T$ μ and v mutually independent, and X_{it} is independent of the μ_i and v_{it} for all i and t . The latter is a key difference between FE and RE models where the RE model imposes more restrictive assumptions. GLS estimation of the RE model requires an estimation of the weight parameter $\theta = 1 - \sigma_v / \sqrt{T\sigma_\mu^2 + \sigma_v^2}$, which in

turn determines the RE estimate as it is part of the GLS estimator $\beta_{\#,RE} = \left\{ (\tilde{P}_1 X_{\#})' (\tilde{P}_1 X_{\#}) \right\}^{-1} (\tilde{P}_1 X_{\#})' \tilde{P}_1 y$ where $\tilde{P}_1 = (1 - \theta)P + Q$. If $\theta=0$, RE becomes OLS ($\tilde{P}_1 = I$) and if $\theta=1$, RE becomes FE ($\tilde{P}_1 =$ within “purging matrix” Q). θ in turn depends on estimating $\sigma_1^2 = T\sigma_{\mu}^2 + \sigma_v^2$ and σ_v^2 , which can be done via OLS or LSDV residuals. In empirical application, implied $\hat{\sigma}_{\mu}^2$ can be negative; Baltagi notes that “one solution is to replace these negative estimates by zero” in which case RE becomes OLS as $\theta=0$. Both OLS and LSDV are consistent and unbiased in the RE model, but not necessarily efficient.

I estimate the three models pooled OLS, one-way individual FE & RE with left-hand variable $y = \text{yield}$ and $X = \text{explanatory variables}$ in my dataset and **compare** the results with each other. A priori, the fixed effects model seems to be the most appropriate specification for my data. I have a time-series panel with $T \gg N$ where N are 9 Eurozone countries with substantial heterogeneity. In this context, pooled OLS falls short by ignoring the panel structure of the data and the random effects framework seems inappropriate as it relies on the plausibility of viewing the N individuals as randomly drawn from a large population. I run several **tests** to check this idea and select the most appropriate **model specification**.

A test for the joint significance of one-way individual FE is provided by the **F-test** (Stata: automatically performed when running FE regression). The test statistic is computed based on the restricted residual sum of squares (RRSS) of the pooled OLS model and the unrestricted residual sum of squares (URSS) from the LSDV regression. Under $H_0: \mu_i = 0, i = 1, \dots, N$, test statistic $F_{1-way} = \frac{(RRSS-URSS)/(N-1)}{URSS/(NT-N-K)} \sim F_{N-1, N(T-1)-K}$. If H_0 is rejected, it means that FE is to be preferred over pooled OLS as the individual fixed effects are jointly significant.

A test for one-way random effects is the **Breusch-Pagan Lagrange Multiplier test** (BPLM) (Stata: `xttest0` after RE regression). The test statistic is computed using the pooled OLS residuals $\tilde{u}$. Under $H_0: \sigma_{\mu}^2 = 0$, test statistic $LM_1 = \frac{NT}{2(T-1)} \left[1 - \frac{\tilde{u}'(I_N \otimes J_T)\tilde{u}}{\tilde{u}'\tilde{u}} \right]^2 \sim \chi_1^2$. If H_0 is rejected, it means that RE is to be preferred over pooled OLS (variance of random effects $\neq 0$).

A comparison between FE and RE is performed by the **Hausman test** (Stata: `hausman FE RE` where `FE` and `RE` are the stored estimates of the respective models). The test relies on the fact that FE is consistent in both FE and RE model while RE is efficient in the RE model but inconsistent in the FE model. The test statistic is based on $\hat{\beta}_{FE}$ and $\hat{\beta}_{RE}$. Under $H_0: E(u_{it}|X_{it}) = 0$ (i.e. in the restricted RE model), test statistic $m = q' (var \hat{\beta}_{FE} - var \hat{\beta}_{RE})^{-1} q \sim \chi_k^2$ where $q = \hat{\beta}_{FE} - \hat{\beta}_{RE}$ and k is the dimension of slope vector β . If H_0 is rejected, FE is to be preferred over RE.

After performing these tests, I analyse the **estimated coefficients** of the preferred (one-way FE) model specification. Another extension is the two-way framework including both individual and time effects, for which appropriate tests can be performed in analogy to the one-way analysis above. However, as Baltagi notes, the 2-way FE model “cannot estimate the

effect of time-invariant and individual-invariant variables, because the Q transformation wipes out these variables". In my sample, this means that I cannot estimate two-way FE coefficients for *euribor_3m*, *vix* and *T00* and also cannot use the Hausman test to directly compare the two-way RE model (which can estimate coefficients for these variables) to the two-way FE model. Since these variables turn out to be relevant in the one-way context in my sample and the FE model is preferred over the RE model, I focus on the one-way framework.

Results

I estimate the 3 models with 3 main sets of regressors: including all variables in my dataset i) without any proxy variable for the ECB's unconventional monetary policy, ii) with the variable *PSPP_p_i* (monthly net PSPP purchases during active phase of program from 2015-2018, EUR bn; 0 otherwise) and iii) with *T00* (Eurosystem's total assets, EUR bn, available for full sample). Furthermore, I optionally include a squared term for government debt/GDP (*debt* and *debt2*) following De Santis (2020) who finds a quadratic and convex relationship between debt and sovereign yields. I also check various time periods, starting with my full sample (2003-2018) and reducing it by one year up to 2010-2018. I run F-tests, BPLM tests and Hausman tests for all these specifications – a table in appendix 6k displays the results of these tests.

For the specifications excluding *PSPP_p_i* and *debt2*, the BPLM tests reject H_0 except for 2003-2018 where RE becomes POLS. When including *PSPP_p_i*, BPLM tests do not reject when including the early sample years up to 2007 but do reject when starting the sample in 2007 and more recent years. When including *debt2*, the RE models with *PSPP_p_i* become POLS but the BPLM tests reject H_0 for the RE models with or without *T00* when starting the sample in 2007 and more recent years. A reason for this could be that there is little variation between countries in the first couple of years of the sample before the global financial and the Eurozone sovereign debt crises in later years. The F-tests reject the H_0 for all combinations of regressors and sample periods. Also, the Hausman tests reject H_0 for all specifications, showing a clear preference for the FE framework. Hence, I conclude that the **FE model is the most appropriate model** for my data. I also conduct a modified Wald test for groupwise heteroskedasticity in the residuals of the FE model (*xttest3* in Stata), which rejects the H_0 of homoskedasticity; hence, I use robust standard errors (Stata: , *vce(robust)*).

Table 5 shows the estimated coefficients of the pooled OLS and FE models with robust standard errors and *debt2* for the full sample. A table with the default standard errors used for the tests above can be found in appendix 6k. Comparing the FE estimates to FE with PSPP, the coefficients of the variables other than PSPP does not change dramatically, hence I argue that the specification including PSPP can be used (although PSPP is 0 for most years and subtracting the country-specific time average leads to negative values for transformed *PSPP_p_i* for those years when PSPP was not active). The within and overall R^2 for the FE models provide further support for the use of fixed effects: for all sets of regressors, the within R^2 (maximized by LSDV) differs from the overall R^2 (shown in table 5, maximized by POLS).

Within vs overall R^2 : FE: 0.753 vs 0.728, FE_w_PSPP: 0.792 vs 0.729, FE_w_T00: 0.814 vs 0.805. Both AIC and BIC seem to favor the FE_w_T00 model.

Pooled OLS and static individual FE panel models: 01/2003 - 12/2018				
	POLS	FE	FE_w_PSPP	FE_w_T00
bas	8.221 *** (0.426)	8.590 *** (0.796)	8.591 *** (0.668)	8.161 *** (0.590)
debt	-0.014 *** (0.005)	-0.030 (0.030)	-0.042 (0.029)	-0.028 (0.017)
debt2	0.000 ** (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)
rating	-0.173 *** (0.015)	-0.177 (0.168)	-0.229 (0.139)	-0.300 ** (0.099)
hicp	0.072 *** (0.023)	0.023 (0.062)	-0.005 (0.056)	0.206 *** (0.058)
esi	-0.049 *** (0.002)	-0.045 *** (0.009)	-0.032 *** (0.008)	-0.021 * (0.010)
euribor_3m	0.764 *** (0.014)	0.665 *** (0.080)	0.561 *** (0.095)	0.286 *** (0.048)
vix	-0.018 *** (0.003)	-0.017 *** (0.004)	-0.012 ** (0.004)	0.009 ** (0.004)
PSPP_p_i			-0.130 *** (0.023)	
T00				-0.001 *** (0.000)
_cons	10.720 *** (0.440)	11.975 ** (3.667)	12.564 *** (3.013)	13.016 *** (2.386)
Number of observations	1724	1724	1724	1724
Adjusted R-squared	0.78	0.75	0.79	0.81
AIC	4186.61	4070.19	3775.16	3579.48
BIC	4235.68	4113.81	3818.77	3623.10

*** p<.01, ** p<.05, * p<.1

Table 5 - Pooled OLS and static individual FE panel models

The *ceteris paribus* **economic interpretation** of the estimated FE coefficients is as follows:

- An increase in **bid-ask spreads** by 1 basis point implies an increase in 10-year government bond yields by 8.16 basis points (FE_w_T00). The estimated effect is a bit higher for the other two FE specifications and the coefficient is significant at the 1% level in all 3 models. Thus, lower liquidity in a given bond market as proxied by higher *bas* has a yield-increasing effect, which seems logical.
- I do not find evidence for a significant impact of **public debt/GDP** (*debt* & *debt2*).
- An increase in the numerical value of **S&P ratings** by 1 implies a decrease in bond yields by 0.3 percentage points (FE_w_T00, significant at 5% level; no significant effect in other FE models). Thus, better sovereign credit ratings imply lower yields, an intuitively sensible finding.

- Higher nominal **inflation** (*hicp*) leads to higher nominal bond yields: an increase in *hicp* by 1 pp leads to an increase in *yield* by 0.2 pp (*FE_w_T00*, significant at 1% level, no significant effect found in other FE models). This is also economically plausible as investors are interested in real yields.
- Better **economic sentiment** implies lower yields: an increase in the *esi* (Index) by 1 point leads to a decrease in *yield* by 2 basis points (*FE_w_T00*, significant at 10% level; estimated effect is larger and more significant in other FE models). This effect seems also credible as a better overall economic picture reduces sovereign risk.
- Higher **short-run interbank rates** (*euribor_3m*) have a yield-increasing impact: an increase in 3-month-Euribor by 1 pp leads to an increase in bond yields by 0.28 pp (*FE_w_T00*, significant at 1% level, impact found to be higher in other FE models, where estimated coefficients are also highly significant). This finding makes sense: when banks' funding conditions deteriorate, financial intermediaries will also demand higher yields in other financial markets, including the government bond markets.
- I find no clear-cut effect for **economic uncertainty** as proxied by *vix*: while the effect is found to be yield-decreasing for FE and *FE_PSPP*, the estimated coefficient is small, but positive for *FE_w_T00* and significant at least at the 5% level across the 3 models.
- Finally, I do find evidence for a yield-decreasing impact of **unconventional monetary policy**. In the specification *FE_w_PSPP*, I find a highly significant effect at the 1% significance level for PSPP purchases. An increase in monthly net purchases by EUR 1bn decreases bond yields by 13 basis points. Also, in the specification *FE_w_T00*, I find a highly significant effect at the 1% significance level for an increase in total Eurosystem assets. An increase in *T00* by EUR 1bn leads to a decrease in yields by 0.1 basis points.

Estimates for the shortened sample 2008-2018 can be found in appendix 6k. There are some slight changes relative to table 5, e.g. the effect of *rating* and *euribor_3m* is not significant in *FE_w_T00* anymore, the effect of *hicp* is found to be significant and yield-increasing for all 3 FE models, while the effect of *vix* is not significant anymore in any FE model. However, the estimated coefficients for unconventional monetary policy remain highly significant and yield-decreasing in both *FE_w_PSPP* and *FE_w_T00* respectively.

Furthermore, estimates for the two-way FE and RE models produced with R can also be found in appendix 6k. As the 2-way FE model cannot estimate coefficients for *euribor_3m*, *vix* or *T00*, I do not directly compare the estimated coefficients to the 1-way FE or 2-way RE models with the full sets of regressors and focus on the one-way framework. However, it is worth noting that while the size of the estimated coefficients changes slightly in the 2-way FE model, *bas* is still highly significant and yield-decreasing, as is *rating* and *esi*. *hicp* loses significance while I now find a significant quadratic and convex impact of government debt/GDP on *yield*. Furthermore, I do find evidence for a significant impact of unconventional monetary policy: in the 2-way FE model (reduced set of regressors) I still find a highly significant yield-decreasing

impact of $PSPP_p_i$, albeit lower than in the one-way model and in a 2-way RE model, I also find a highly significant yield-decreasing impact of $T00$.

Conclusions

Static panel analysis has yielded preliminary answers to my research questions (main factors driving yields and flow effects of Eurosystem unconventional monetary policy). I have found that in the one-way FE model, which is more suited to my dataset than pooled OLS or RE based on the results of BPLM tests, F-test and Hausman tests, *ceteris paribus*: higher bond **bid-ask spreads** and short-term **interbank rates** have a highly significant yield-increasing effect, while better **economic sentiment** has a highly significant yield-decreasing effect. I also find a highly significant yield-decreasing effect for 2 different proxies of **unconventional monetary policy**, namely increases in monthly net PSPP purchases or increases in total Eurosystem assets. These results are robust to changes in the used sample period as well as extensions to the 2-way framework. In some models I also find evidence for a yield-increasing effect of higher inflation, a yield-decreasing effect for better ratings and/or a quadratic and convex impact of higher public debt/GDP on yields. These findings all seem economically plausible, and are broadly in line with findings of De Santis (2020; slightly different set of regressors over period 2004-2015). Lastly, evidence for the impact of economic uncertainty as proxied by *vix* is ambiguous. A key drawback of static panel models is their bias in estimating regressions where the dependent variable also depends on its own lag(s), which is a potentially important extension in my context. Hence, I move on to dynamic panel modelling.

c. Dynamic panel modelling

Methodology

In the following I use the notation and conclusions from Baltagi (2013, chapter 8) and Verbeek (2017, chapter 10) to briefly outline the methodology used in the dynamic panel modelling discussed below. As pointed out in the univariate analysis section, some variables in my dataset, including *yield*, exhibit a substantial degree of persistence, i.e. dependence on their own past. One way to account for this is by including the lagged dependent variable as a regressor. Consider the equation $y_{it} = \delta y_{i,t-1} + x'_{it}\beta + u_{it}$ ($i=1\dots N$; $t=1\dots T$) where δ is a scalar, x'_{it} and β are the usual vectors of regressors and coefficients with fitting dimensions and $u_{it} = \mu_i + v_{it}$ with $\mu_i \sim IID(0, \sigma_\mu^2)$ and $v_{it} \sim IID(0, \sigma_v^2)$ where μ and v are independent of each other and among themselves. $y_{i,t-1}$ is a function of μ_i and hence correlated with the error term. In the dynamic panel model context, OLS and FE suffer from the **Nickell bias**. OLS is biased and inconsistent. While the within transformation of the FE estimator wipes out the μ_i , the resulting terms $(y_{i,t-1} - \bar{y}_{i,-1})$ and $(v_{it} - \bar{v}_i)$ are still correlated and FE is biased and inconsistent. While the bias disappears for $T \rightarrow \infty$, in empirical application this is not the case and hence, dynamic panel methods using valid instruments are needed. Out of the many potential avenues, I use two one-step GMM methods to estimate my model: difference and system GMM.

In the **difference GMM** framework (Arellano-Bond estimator) the model is estimated in first differences, i.e. $(y_{it} - y_{i,t-1}) = \delta(y_{i,t-1} - y_{i,t-2}) + (x'_{it} - x'_{i,t-1})\beta + (v_{it} - v_{i,t-1})$ or $\Delta y_{it} = \delta \Delta y_{i,t-1} + \Delta x'_{it}\beta + \Delta v_{it}$. In this case, Δv_{it} is correlated with $\Delta y_{i,t-1}$ and autocorrelated. However, $y_{i,t-s}$ for $s > 1$ can be used as a valid instrument as it is highly correlated with $\Delta y_{i,t-1}$ but not correlated with Δv_{it} . The set of valid instruments grows with each period and is longest for period $t=T$. The differenced equation is pre-multiplied by an appropriately defined matrix of instruments W' ; GLS on this new equation yields the consistent one-step Arellano-Bond estimator. The matrix of instruments includes the list of valid instruments for the lagged dependent variable and also - if the variables x are strictly exogenous - the first-differenced x variables, which in this case serve as their own instruments.

In the **system GMM** framework (Arellano-Bover/Blundell-Bond estimator), in addition to the equation in differences with lagged level terms as instruments, also the equation in levels with lagged difference terms as instruments is considered. Baltagi provides a step-by-step derivation of the system GMM estimator based on $y_{it} = \delta y_{i,t-1} + x'_{it}\beta + Z'_i\gamma + u_{it}$, where the regressors are split into time-invariant variables Z'_i and variables that vary across individuals and time x'_{it} . Defining $\eta' = (\delta, \beta', \gamma')$ and $W_i = [y_{(-1)}, X_i, Z_i]$, the equation can be rewritten in vector form as $y_i = W_i\eta + u_i$ with $u_i = \mu_i + v_i$. Then, stacking y_i , W_i , and u_i respectively across individuals and pre-multiplying the equation by appropriately defined $M'\bar{H}$ (M is the matrix of instruments; $\bar{H} = I_N \otimes H$ where the first $T-1$ rows of H are defined such that they transform the effect vectors to 0) one receives $M'\bar{H}y = M'\bar{H}W\eta + M'\bar{H}u$. GLS yields the system GMM estimator $\hat{\eta} = [W'\bar{H}'M(M'\bar{H}\bar{\Omega}\bar{H}'M)^{-1}M'\bar{H}W]^{-1}W'\bar{H}'M(M'\bar{H}\bar{\Omega}\bar{H}'M)^{-1}M'\bar{H}y$ where the covariance matrix $\bar{\Omega}$ contains consistent estimates from a preliminary step. System GMM can be more efficient than difference GMM under certain conditions. (Roodman 2009 - assumption about initial conditions: *"controlling for the covariates, faster-growing individuals are not systematically closer or farther from their steady states than slower-growing ones"*).

Both difference and system GMM estimators have been implemented in the Stata command `xtabond2` which I use for my purposes. After estimation, `xtabond2` automatically provides the results of several **tests**. For a well-specified model the disturbances will exhibit first-order **serial correlation**, but will not exhibit second-order serial correlation. In the `xtabond2` command, the **Arellano-Bond test** is applied after estimation to the residuals in differences (H_0 : no serial correlation of given order) to test for AR(1) and AR(2) (Roodman 2009). As also outlined by Baltagi, the test should reject for AR(1) and not reject for AR(2). Furthermore, `xtabond2` reports tests for **overidentifying restrictions**. When there are more moment conditions than needed to for exact identification, a specification test (H_0 : the imposed restrictions are jointly valid) can be calculated- the **Sargan/Hansen test**. The desired result is that H_0 is not rejected, hence, the restrictions are jointly valid. As Roodman (2009) discusses, for robust one-step GMM, the Sargan statistic is inconsistent and, in this case, `xtabond2` additionally provides a consistent Hansen statistic for the same H_0 . However, Roodman also warns that this test grows weaker the more moment conditions there are.

Results

For the application of *xtabond2*, I rely on examples provided by Baltagi (2013) and documentation by Roodman (2009). I estimate difference GMM (*DIF_GMM*) and system GMM (*SYS_GMM*) with the full set of regressors used also in the static context, i.e. including all variables in my dataset i) without any proxy variable for the ECB's unconventional monetary policy, ii) with the variable *PSPP_p_i* (monthly net PSPP purchases during active phase of program from 2015-2018, EUR bn; 0 otherwise) and iii) with *T00* (Eurosystem's total assets, EUR bn, available for full sample). On top, I add the lagged dependent variable as a regressor (*L.yield*). Following the recommendations of Roodman, I also add time dummies for each month and include all but one of them as additional regressors. This is done to account for shocks that affect all countries at the same time, thus controlling for contemporaneous correlation between countries. *L.yield* is used as a GMM-style instrument, i.e. the list of instruments grows in *t*. Since my data set has a large time dimension, this results in a large number of instruments. Therefore, I also use the collapse option in *xtabond2*, which is a viable way to “*limit instrument proliferation*” without losing too much information (Roodman). All other regressors serve as their own instruments (*ivstyle* option in *xtabond2*). Finally, I follow Baltagi in using the *robust* option, which produces “*standard errors that are robust to heteroskedasticity and arbitrary patterns of autocorrelation within individuals*” for one-step GMM routines (Roodman).

Table 6 shows the results of the GMM estimation for the full sample, with the columns indicating the different sets of regressors and GMM models. For brevity, the coefficients for time dummies and the constant are omitted from the tables of results. First, the estimated coefficients along with their significance levels denoted by asterisks are shown. Then, the test results of the aforementioned tests are shown below the horizontal line in the table. First, the p-values of the Sargan and Hansen test statistics for overidentifying restrictions are given by *sarganp* and *hansenp*. Here, the picture across the different GMM specifications is the same: The Sargan test rejects the H_0 of joint validity of the restrictions, while the Hansen test (which according to Roodman is consistent in the case of robust one-step GMM estimation but grows weaker as the number of instruments increases) does not reject the H_0 . Ideally, these tests would give the result that the restrictions are jointly valid; the high number of instruments relative to the number of countries might be problematic for using GMM in my dataset. Next, the p-values of Arellano-Bond test statistics for serial correlation AR(1) and AR(2) are given by *ar1p* and *ar2p*. Here, the results are more in line with the expected outcome for a well-specified GMM model: the AR(1) test rejects the H_0 of no serial correlation of order 1 in the disturbance terms (albeit only at a significance level slightly above 10% for the difference GMM specifications *DIF_GMM* and *DIF_GMM_T00*) while the AR(2) test does not reject the H_0 of no serial correlation of order 2 in the disturbance terms, which indicates that the models are not mis-specified. Lastly, the number of instruments is given by *j0*.

Dynamic one-step panel models: 01/2003 - 12/2018, robust standard errors						
Variable	DIF_GMM	DIF_GMM_PSPP	DIF_GMM_T00	SYS_GMM	SYS_GMM_PSPP	SYS_GMM_T00
yield L1.	.82100912***	.81404861***	.82100911***	.86173644***	.86299429***	.8617364***
bas	1.9526185***	2.0153108***	1.9526185***	1.252705***	1.2445937***	1.2527052***
debt	-.01033346	-.01374263	-.01033347	-.00087832	-.00086349	-.00087831
debt2	.00005033	.00005641	.00005033	3.444e-06	3.420e-06	3.444e-06
rating	-.00298478	-.01095734	-.00298477	-.03017193***	-.02979152***	-.03017194***
hicp	.0171982**	.00075472	.01719819**	.00962543*	.00939588*	.00962543*
esi	.0015941	.00133833	.0015941	-.00363457***	-.0036066***	-.00363457***
euribor_3m	.13464056***	.13419509***	.11712411***	.04977762	.01873442	.06482501*
vix	-.00421032***	-.00468641***	-.00496466***	.00013929	-.000385	.00151799**
PSPP_p_i T00		-.01669369	-.0000441**		-.00081092	.00005123***
sarganp	7.925e-24	1.428e-22	7.925e-24	8.999e-30	1.077e-29	8.999e-30
hansenp	1	1	1	1	1	1
ar1p	.10107141	.09871016	.10107141	.08757772	.08825164	.08757773
ar2p	.44712216	.44851015	.44712219	.42006146	.41955587	.42006146
j0	389	390	390	391	392	392

Legend: * p<.1; ** p<.05; *** p<.01

Table 6 - One-step GMM models: 2003-2018

The *ceteris paribus* **interpretation** of the estimated coefficients is as follows:

- The estimated coefficient for **lagged yield** is relatively high, ranging from 0.81 for *DIF_GMM_PSPP* to 0.86 for *SYS_GMM_PSPP* and is highly significant at the 1% level for all specifications. E.g. an increase in yields by 1 percentage point in the previous month will lead to an increase in the current month's yield by 0.86 percentage points for *SYS_GMM_PSPP*. Dynamic panel analysis thus further underlines the high persistence of yields, evidence for which was also found in the univariate analysis. The estimated autoregressive coefficient is higher for the system GMM specifications compared to the difference GMM specifications.
- An increase in **bid-ask spreads** by 1 basis point implies an increase in 10-year government bond yields by 1.95 basis points (*DIF_GMM_T00*). The estimated effect is a bit higher for difference GMM versus system GMM, but the coefficient is significant at the 1% level in all models. Thus, lower liquidity in a given bond market as proxied by higher *bas* has a yield-increasing effect, which seems logical.
- I do not find evidence for a significant impact of **public debt/GDP** (*debt* & *debt2*).
- I find partial evidence that an increase in the numerical value of **S&P ratings** by 1 implies a decrease in bond yields by around 0.03 percentage points (system GMM, significant at 1% level; no significant effect in difference GMM). Thus, better sovereign credit ratings imply lower yields, an intuitively sensible finding.
- I find evidence that higher nominal **inflation** (*hicp*) leads to higher nominal bond yields: an increase in *hicp* by 1 pp leads to an increase in *yield* by 0.017 pp (*DIF_GMM_T00*, significant at 5% level; the estimated effect is a bit lower and less significant for system GMM). This is also economically plausible as investors are interested in real yields.

- I find partial evidence that better **economic sentiment** implies lower yields: an increase in the *esi* (Index) by 1 point leads to a decrease in *yield* by around 0.3 basis points (*system GMM*, significant at 1% level; no significant effect in difference GMM). This effect also seems credible as a better overall economic picture reduces sovereign risk.
- I find partial evidence that higher **short-run interbank rates** (*euribor_3m*) have a yield-increasing impact: an increase in 3-month-Euribor by 1 pp leads to an increase in bond yields by 0.11-0.13 pp (difference GMM, significant at 1% level, smaller and less significant impact for SYS_GMM_T00, no effect for other system GMM specifications). This finding makes sense: when banks' funding conditions deteriorate, financial intermediaries will also demand higher yields in other financial markets, including the government bond markets.
- I find no clear-cut effect for **economic uncertainty** as proxied by *vix*: while the effect is found to be yield-decreasing for difference GMM, the estimated coefficient is small, but positive for SYS_GMM_T00.
- Also, I do not find clear-cut evidence for the impact of **unconventional monetary policy**. In the difference and system GMM specifications where PSPP purchases are included, the estimated coefficient is not significant. In the difference GMM specifications where T00 (total Eurosystem assets) is included, the estimated coefficient is found to be significant at the 5% level and yield-decreasing. However, in the system GMM specification with T00, I find a highly significant effect at the 1% significance level for a yield-increasing effect for an increase in T00.

These results do not change substantially when leaving out the regressors *debt* and *debt2* (main differences: *hicp* less significant in system GMM case, estimated coefficients for T00 are bigger but significance and direction unchanged). Similarly, when estimating the models with the full set of regressors but only for 2008-2018, there are no major changes (main differences: *rating* is less significant in system GMM, *euribor_3m* has a highly significant yield-increasing effect also in system GMM, *vix* is significant and unambiguously yield-decreasing effect in both system and difference GMM; however, significance and direction for T00 are unchanged). These results can be found in appendix I.

Conclusions

Dynamic panel analysis has corroborated some findings on my research questions (main factors driving yields and flow effects of Eurosystem unconventional monetary policy) from the static panel analysis and univariate analysis sections. I have confirmed that yields exhibit a substantial degree of **persistence** (highly significant estimated coefficient for lagged yield). Furthermore, *ceteris paribus*, I find that in both difference and system GMM frameworks, there is clear and highly significant evidence for a yield-increasing impact of higher bond **bid-ask spreads**. I find weaker evidence (coefficients not highly significant across all specifications) for a yield-increasing impact of higher **short-term interest rates** and **inflation**. I find partial evidence (coefficients only significant in system GMM) for a yield-decreasing impact of better

ratings and economic sentiment. These results are broadly robust to changes in the sample period and exclusion of insignificant regressors. As in the static panel analysis section, evidence for the impact of economic uncertainty is ambiguous. However, I could not corroborate a clear-cut effect of the Eurosystem's unconventional monetary policy in the dynamic panel analysis. The estimated coefficients for PSPP purchases were insignificant, while the estimated coefficients for T00 (total Eurosystem assets) were highly significant, but pointed into different directions for difference and system GMM.

Overall, the estimated coefficients in GMM were smaller than in static panel models – this seems logical as the static models do not capture the high persistence of *yield*. While both difference and system GMM estimate a high coefficient for *L.yield*, the coefficient is slightly higher for system GMM (around 0.86 vs 0.82). Verbeek (2017) discusses that in the case when $\delta \rightarrow 1$, system GMM may perform better than difference GMM, but at the same time, in the system GMM framework the *“validity of the additional instruments depends upon the assumption that changes in $y(it)$ are uncorrelated with the fixed effects”*, which Verbeek cautions is less likely to be fulfilled precisely when $\delta \rightarrow 1$. In any case, my dataset is not ideally suited for analysis in a GMM framework: as Roodman (2009) notes, GMM estimators should be applied to small T, large N datasets; in my dataset, the number of instruments (even using the collapse option) is clearly very high relative to the low number of countries (9).

5. Conclusions

This master thesis was motivated by a desire to gain a deeper understanding of the complicated interplay between central bank policy and government bond yields in the Euro Area against the backdrop of other key macroeconomic processes. Sovereign bonds are a crucial benchmark for financial markets and of central importance to public deficit financing. Thus, modelling the dynamics of sovereign bond yields remains highly relevant, especially since the central bank/bond market nexus has deepened drastically.

I start out by discussing which unconventional tools the ECB and the Eurosystem deployed in the wake of the global financial crisis and the Eurozone sovereign debt crisis. Among these tools, I focus on the Eurosystem's asset purchases program (APP) and describe in detail the chronology and technicalities of the Public Sector Purchase Program (PSPP), which accounted for the lion share of APP and contributed greatly to the unprecedented expansion of the Eurosystem's balance sheet from 2015 onwards. I follow a strain of economic literature that attempts to estimate the “flow effects” of these purchases (as opposed to the “stock effect” of the announcement of a given program) over a longer sample period in a dataset including both proxies for unconventional monetary policy as well as other key macro-financial variables identified in the literature as important drivers of government bond rates.

In univariate analysis, I find that several variables in my dataset exhibit substantial persistence, including sovereign bond yields. In static panel analysis, I estimate various models for bond yields and a set of regressors across my panel of 9 countries and 192 months from 2003–2018. In a one-way fixed effects model, I find inter alia that bid-ask spreads and short-term interbank

rates have a highly significant yield-increasing effect, while better economic sentiment has a highly significant yield-decreasing effect; these findings are broadly in line with De Santis (2020). I also find a highly significant yield-decreasing effect for increases in PSPP purchases and total Eurosystem assets. Considering yield persistence, I move to a dynamic panel setting by adding lagged yields to the set of regressors. Difference and system GMM estimation confirms the high persistence of yields and corroborates the importance of bid-ask spreads while supporting (a bit less clearly) the evidence on short-term interbank rates and economic sentiment. However, the GMM estimation does not corroborate the finding of a clear-cut effect of the Eurosystem's unconventional monetary policy. Contrary to my expectations based on economic intuition, I do not find evidence for a clear negative effect of higher government debt and only relatively weak evidence for an effect of sovereign ratings or inflation on nominal bond yields.

Inevitably, the methodological approach and dataset chosen in this thesis has various shortcomings. The monthly frequency of data might obscure more granular effects on fast-moving financial markets. Yield data are taken from an ECB-compiled dataset which includes only bonds with a remaining maturity of 10 years – the chosen regressors' effects on other segments of the bond market might differ. The dataset lacks good proxies to encompass highly relevant political uncertainty or shifts in announced government fiscal policy. While GMM models account for the dynamic dependence of yields, the dataset with low N and high T dimensions is not ideally suited for these methods. More generally, the *ceteris paribus* attribution of effects is challenging in a complex interdependent financial system. Further avenues of research could look into more granular and broader datasets or investigate panel unit roots, panel cointegration or more general panel VAR models.

6. Appendix

a. Data description and sources

Variable name	Description	Unit	Frequency in source	Frequency transformation	Source	Link to source
<i>yield</i>	"Long-term interest rate for convergence purposes - 10 years maturity, denominated in Euro; percentages p.a.; period averages; secondary market yields of government bonds with maturities of close to ten years"	% p.a.	monthly	none	ECB	https://sdw.ecb.europa.eu/browse.do?node=9691124
<i>bas</i>	Bid-Ask spread as calculated by author from Refinitiv data; Bid-Ask-spread (% p.a.) := Bid yield (% p.a.) - Ask yield (% p.a.). Bid and Ask yield-to-maturity for 10-year maturity benchmarks	% p.a.	daily	monthly data calculated as average over daily values in given calendar month	Refinitiv Data-stream	Access via https://vdc.univie.ac.at/
<i>rating</i>	S&P Long-Term Sovereign Foreign Currency Rating Outlooks Or CreditWatch - numerical transformation as calculated by author from S&P source data	value from 0 to 21; 21 = best	daily	monthly data calculated as average over daily values in given calendar month	Standard and Poor's	https://www.spglobal.com/ratings/en/research/articles/210726-sovereign-ratings-history-12029439
<i>debt</i>	Consolidated government debt as % of GDP; debt at face value as ratio to annual moving sum of GDP at end of period, neither seasonally nor calendar adjusted; ESA 2010 definitions	% of GDP	quarterly	monthly values: end of Q1 data used for April, May, June, end of Q2 data used for July, August, September etc.	ECB / Eurostat	https://sdw.ecb.europa.eu/browse.do?node=9693772
<i>hicip, hicip_ea</i>	All-items harmonised index of consumer prices (HICP); Annual rates of change (m/m-12) (table: prc_hicip_manr)	% change y/y	monthly	none	Eurostat	https://ec.europa.eu/eurostat/databrowser/view/PRC_HICP_MANR/default/table?lang=en

Table 7 - Data description and sources; table continued on next page

Variable name	Description	Unit	Frequency in source	Frequency transformation	Source	Link to source
<i>esi</i>	EC Economic Sentiment Indicator; based on monthly survey data; purpose: to "track GDP growth"; scaled to have long-term mean of 100; values > 100 or < 100 indicate above or below-average economic sentiment respectively	Index	monthly	none	EC	https://economy-finance.ec.europa.eu/economic-forecast-and-surveys/business-and-consumer-surveys_en
<i>vix</i>	CBOE Volatility Index value; designed to measure "30-day expected volatility of the S&P 500 Index"; higher VIX is interpreted as a proxy for larger expected volatility	Index	daily	monthly data calculated as average over daily close values in given calendar month	CBOE	https://www.cboe.com/tradable_products/vix/vix_historical_data/
<i>euribor_3m</i>	Euro area money market - Euro Interbank Offered Rate 3-month - average of daily observations through period	% p.a.	monthly	none	ECB / Refinitiv / EMMI	https://sdw.ecb.europa.eu/quickview.do?SERIES_KEY=143.FM.M.U2.EUR.RT.MM.EURIBOR3MD.HSTA
<i>PSPP_p_i</i> , <i>PSPP_p</i>	Eurosystem's monthly net PSPP purchases per country in sample (<i>PSPP_p_i</i>) and in total (<i>PSPP_p</i> : including all eligible Euro Area countries + supranationals)	EUR bn	monthly	none	ECB	https://www.ecb.europa.eu/mopo/implementation/app/html/index.en.html
<i>PSPP_h_i</i> , <i>PSPP_h</i>	Eurosystem's PSPP holdings per country in sample (<i>PSPP_h_i</i>) and in total (<i>PSPP_h</i> : including all eligible Euro Area countries + supranationals)	EUR bn	monthly	none	ECB	https://www.ecb.europa.eu/mopo/implementation/app/html/index.en.html
<i>A1 - A9</i> ; <i>T0</i>	Consolidated Eurosystem balance sheet assets reported for end of each week. Total assets <i>T0</i> consist of subitems <i>A1 - A9</i> as per ECB definitions	EUR bn	weekly	monthly data calculated as average over weekly values in given calendar month	ECB	https://sdw.ecb.europa.eu/browse.do?node=9691294

Table 8 - Data description and sources

b. ISO 3166-1 alpha-3 country codes

As is common praxis, throughout this thesis country codes as defined by the International Organization for Standardization (ISO) are used to represent country names of EA members. Below is a list of the ISO 3166-1 alpha-3 codes (ISO, 2019) used in this thesis in alphabetical order.

AUT Austria
BEL Belgium
DEU Germany
ESP Spain
FIN Finland
FRA..... France
ITA..... Italy
NLD Netherlands (the)
PRT..... Portugal

c. ECB definition of yield

ECB (2022) defines *yield* as follows: “For a fully paid fixed coupon bond paying h times interest per year with an assumed single redemption date, the redemption yield formula can be written as follows:”

Equation 1 – ECB redemption yield formula

$$P = v^{f_1} * \left(k + \sum_{i=1}^{n-1} \frac{g}{h} * v^i \right) + \left(C + \frac{g}{h} * f_2 \right) * v^{n+f_1+f_2-1}$$

P = gross price of the bond at the close of the previous business day

g = annual coupon interest rate as a percentage

k = first/next coupon payment as a percentage

h= number of coupon payments a year

n = number of coupon payments to redemption

f1 = fraction of the number of calendar days from value date to the first/next interest payment

f2 = fraction of the number of calendar days from the last normal coupon date to redemption

C = redemption value

v = discount factor. $v = 1/(1+y/h)$, where y = required redemption yield compounded h times per annum.

Economic interpretation: The first right-hand-side summand represents the present value of all upcoming coupon payments. The second right-hand-side summand represents the present value of the redemption value plus the present value of the fraction of interest accumulated from the last coupon payment until the redemption date.

d. S&P rating mapping table

I follow Kinatder & Wagner (2017) in assigning numerical values to S&P sovereign credit ratings and outlooks. Numerical values range from 0 to 21, where 21 represents the best attainable rating level while 0 is the worst attainable level. Table 9 shows the mapping between S&P's long-term foreign currency issuer rating levels & outlooks and corresponding numerical values as well as an example.

According to S&P's methodology as disclosed on their website (S&P 2022b): *"S&P Global Ratings uses CreditWatch and rating outlooks to indicate its view regarding the degree of likelihood of a rating change and, in most cases, the probable direction of that change. Although many rating changes are preceded by a CreditWatch placement or a non-stable outlook, changes can and should occur even when the outlook is stable or the rating is not on CreditWatch but when a sudden change in the creditworthiness can be assessed immediately. Also, a positive or negative outlook is not necessarily a precursor of a future rating action or CreditWatch listing."*

The order of the numerical values assigned to "credit outlook" are based on the understanding that the assignment of an outlook or CreditWatch other than "stable" is likely to be understood as an indication of possible future movements of the rating level. Therefore, the decision to include both credit rating levels as well as outlook is taken as defined in table 9:

Credit rating level	Numerical value A		Credit outlook	Numerical value B
AAA	21		Watch Positive	-0.75
AA+	20		Positive	0.50
AA	19		Stable	0.00
AA-	18		Negative	-0.50
A+	17		Watch Negative	-0.75
A	16			
A-	15			
BBB+	14	Credit rating =	combination of:	
BBB	13		Credit rating level	
BBB-	12		Credit outlook	
BB+	11			
BB	10			
BB-	9	variable rating =	sum of:	
B+	8		Numerical value A	
B	7		Numerical value B	
B-	6			
CCC+	5			
CCC	4	Example rating:	AA+, Positive outlook	
CCC-	3	Credit rating level:	AA+; value A = 20	
CC	2	Credit outlook:	Positive; value B = 0.5	
SD	1	variable rating =	20 + 0.5 = 20.5	
D	0			

Table 9 - Numerical mapping of S&P ratings

e. EC ESI definition

EC (2022) gives a detailed description of the calculation of the Economic Sentiment Indicator (ESI) for each month t . The following is a brief summary of the procedure. The ESI is based on monthly surveys conducted in all EU countries across firms in manufacturing, construction, retail trade and services as well as across consumers. Around 134,000 firms and 32,000 consumers are surveyed, with country sample sizes varying according to size and structure of the respective economies. Out of these monthly surveys, 15 questions or “components” are used to calculate ESI separately for each country. For each component, the *balance* is calculated as “the difference between the percentages of respondents giving positive and negative replies” in the following way for questions with 3 and 6 possible answers:

- “if P , E and M (with $P+E+M=100$) denote percentages of respondents having chosen respectively options positive, neutral, negative, the balance is calculated as $B = P - M$ ”
- “If P , E and M have the same meaning as in the previous paragraph, while PP denotes the percentage of respondents having chosen the option “very positive”, MM the percentage of respondents having chosen the option “very negative” and N is the percentage of respondents without any opinion (so that $PP+P+E+M+MM+N=100$), balances are calculated as $B = (PP + \frac{1}{2}P) - (\frac{1}{2}M + MM)$ ”

Furthermore, to calculate ESI for each country, 3 steps are required:

1st step: For each component $j = 1, \dots, 15$ the balance X is standardized using its sample mean and standard deviation where the sample period is given by $t=1$ is January 2000 and T is frozen at December of the previous year.

$$Y_{j,t} = \frac{X_{j,t} - \bar{X}_j}{S_j} \text{ where } \bar{X}_j = \frac{1}{T} \sum_{t=1}^T X_{j,t} \text{ and } S_j = \sqrt{\frac{1}{T-1} \sum_{t=1}^T (X_{j,t} - \bar{X}_j)^2}$$

2nd step: All standardized component balances are weighted and divided by the sum of available series' weights (division: to account for potential missing components in given month):

$$Z_t = \frac{\sum_j w_j Y_{j,t}}{(\sum_j w_j)_t} \text{ where } \left(\sum_j w_j \right)_t \text{ is sum of weights of the available series at time } t$$

Weights w_j are determined based on a) sector weights (based on the “representativeness” of the sector in question and tracking performance vis-à-vis the reference variable [...] GDP growth” and b) number of components included per sector. Weights:

Sector	Sector weight	# of components per sector	Weight of component = w(j) within sector
Industry	40%	3	13.33%
Services	30%	3	10.00%
Consumers	20%	4	5.00%
Construction	5%	2	2.50%
Retail trade	5%	3	1.67%
Total	100%	15	100.00%

Table 10 - ESI component weights

3rd step: ESI is “scaled to have a long-term mean of 100 and a standard deviation of 10”. Sample period to calculate the sample moments is the same as in step 1. “Values greater than 100 indicate an above-average economic sentiment, whereas values below 100 indicate a below-average position. Assuming approximate normality, the imposed standard deviation of 10 implies that in about 68% of the cases the ESI will be within the range of 90 to 110.”

$$ESI_t = \left(\frac{Z_t - \bar{Z}}{S_Z} \right) * 10 + 100 \text{ where } \bar{Z} = \frac{1}{T} \sum_{t=1}^T Z_t \text{ and } S_Z = \sqrt{\frac{1}{T-1} \sum_{t=1}^T (Z_t - \bar{Z})^2}$$

Equation 2 - ESI formula

f. CBOE VIX definition

The Chicago Board Options Exchange (CBOE) writes about its volatility index in its white paper (2019): “The VIX Index is a volatility index comprised of options rather than stocks, with the price of each option reflecting the market’s expectation of future volatility. Like conventional indexes, the VIX Index calculation employs rules for selecting component options and a formula

to calculate index values. [...] The VIX Index measures 30-day expected volatility of the S&P 500 Index". The formula used to calculate VIX is given by (CBOE 2019):

Equation 3 - CBOE VIX generalized formula

$$\sigma^2 = \frac{2}{T} \sum_i \frac{\Delta K_i}{K_i^2} e^{RT} Q(K_i) - \frac{1}{T} \left[\frac{F}{K_0} - 1 \right]^2$$

where:

VIX = σ * 100

and

T = time to expiration

F = Forward index level derived from index option prices

K_0 = First option strike price below forward index level F

$\Delta K_i = \frac{K_{i+1} - K_{i-1}}{2}$ = Interval between strike prices; half of difference between strikes on either side of K_i

K_i = Strike price of i-th out-of-the-money option; a call option price if $K_i > K_0$ and a put option price if $K_i < K_0$; both put and call if $K_i = K_0$

R = Risk-free interest rate to expiration; R assumed to be yields derived from U.S. Treasury yield curve

$Q(K_i)$ = midpoint of the bid-ask spread for each option with strike K_i

Economic interpretation: The CBOE white paper (CBOE 2019) provides a detailed example of the procedure used to select options which are then used for the VIX calculation. In short, VIX is an index which uses observable information on options on the US "S&P 500" stock market index to calculate a proxy measure of expected volatility of the index over the next 30 days. VIX is always a positive value by definition. Higher VIX is interpreted as a proxy for larger expected stock index volatility. The size of VIX inter alia depends on

- i) Level of forward index level derived from index option prices (F)
- ii) Prices of out-of-the money call and put option contracts with non-zero bid prices to either side of F
- iii) Size of ΔK_i : size of steps between included options' strike prices

VIX uses "out-of-the-money SPX calls and out-of-the-money SPX puts centered around an at-the-money strike price, K_0 . Only SPX options quoted with non-zero bid prices are used in the VIX Index calculation". Interpretation: Generally, the more options are being bid for on either side of the forward index level derived from index option prices, the more market participants assume that the future index level will differ from F. The larger the interval between out-of-the-money option strike prices with non-zero bids, the larger VIX.

g. PSPP and Eurosystem balance sheet variables

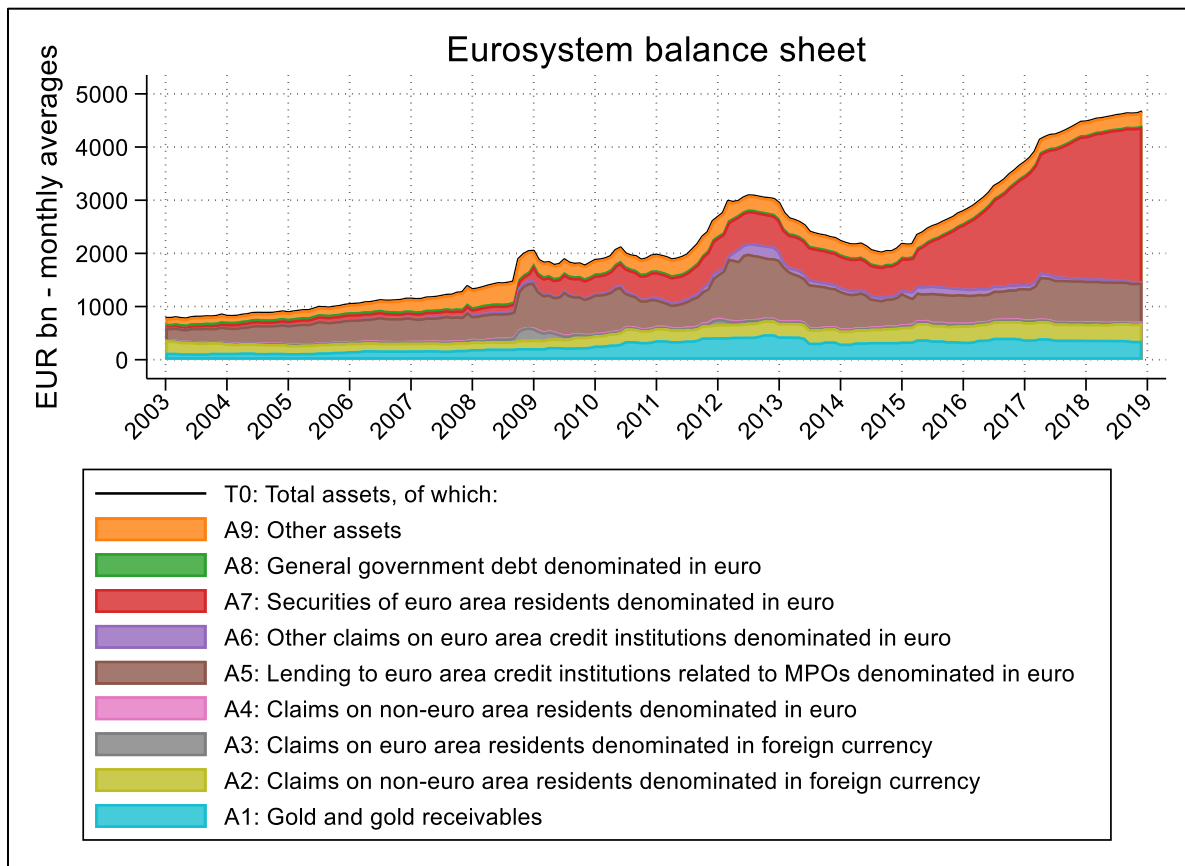


Figure 18 - Eurosystem balance sheet

The ECB publishes data (ECB, 2022c) on the consolidated balance sheet of the Eurosystem on a weekly basis (end of period). I have used the published asset-side data to calculate monthly averages to add to my dataset. Figure 18 shows variables *T0* (total assets), which is the sum of variables *A1* through *A9* as defined by the ECB. As the figure indicates, major shifts in the volume of the Eurosystem's assets over the sample period were largely driven by changes in balance sheet item *A7* (Securities of euro area residents denominated in euro) and *A5* (Lending to euro area credit institutions related to monetary policy operations (MPOs) denominated in euro). In this section, I briefly discuss *A7* followed by a brief discussion of *A5* in the subsequent section 6j.

As per ECB (2022c), *A7* is mainly made up of “*securities held for monetary policy purposes*”, which “*reflects the Eurosystem's holdings of securities purchased within the scope of its monetary policy-related asset purchase programmes, broken down into different portfolios*”. Figure 19 shows that the major increase in *A7* after late 2014 were to a large extent driven by the ramp-up of the Eurosystem's massive APP programs, the bulk of which was made up of the PSPP. *A7* is the same as in figure 18 while data on APP subprograms' holdings is end-of-month data from a different source (ECB 2019a). Both sources report book value of securities at amortised cost. Figure 19 once more motivates the focus of this thesis on PSPP, which was by far the largest APP subprogramme in the sample period.

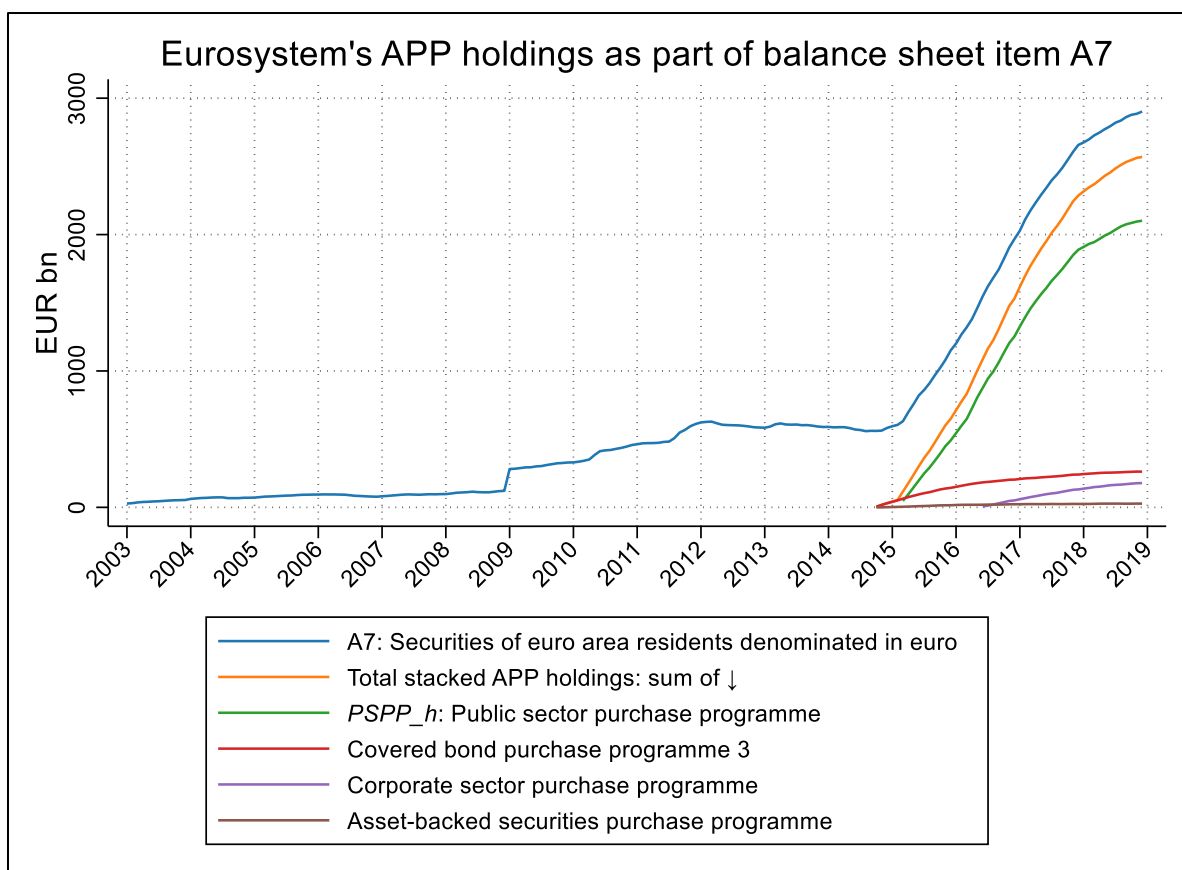


Figure 19 - Eurosystème's APP holdings as part of balance sheet item A7

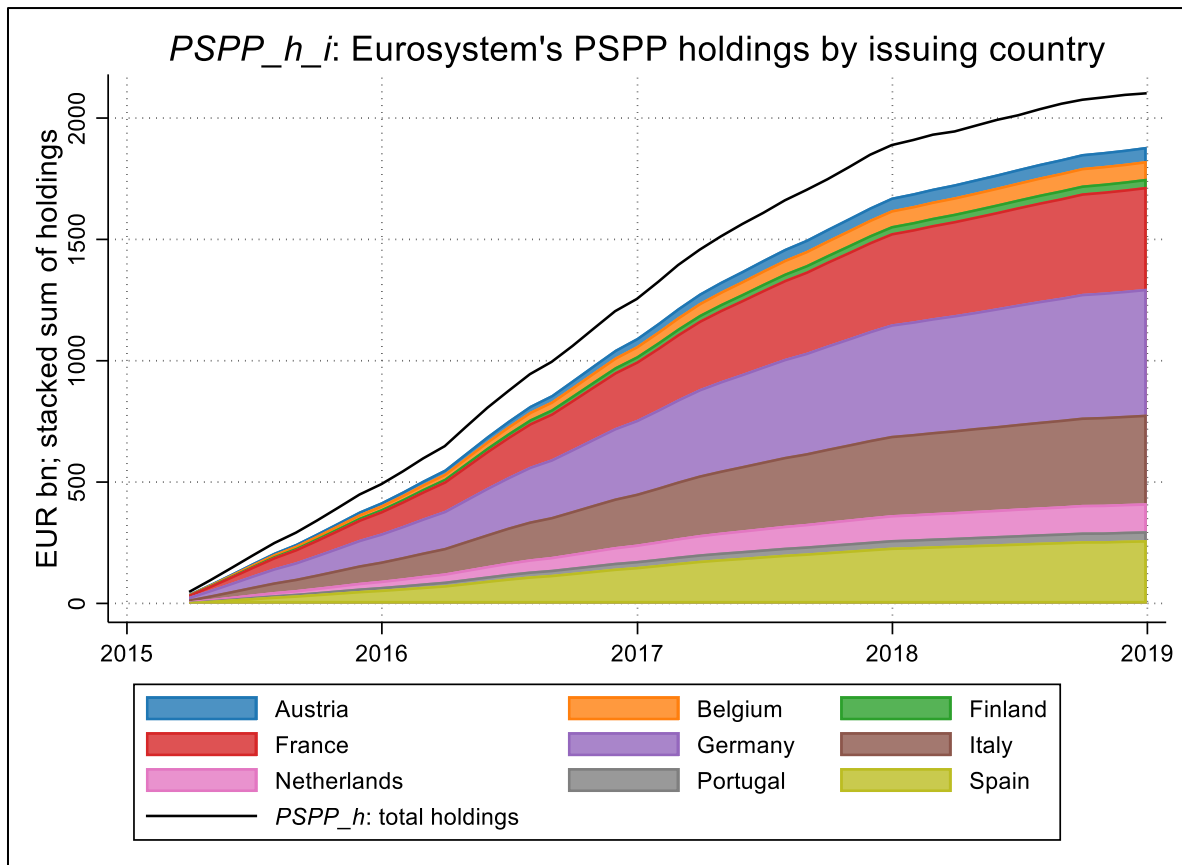


Figure 20 - PSPP_h_i: Eurosystème's PSPP holdings by issuing country

Figure 20 further zooms further in and shows the envelope of variable $PSPP_h$ (Eurosystem's end-of-month PSPP holdings across all jurisdictions as reported by ECB 2019a) versus holdings of individual countries' securities as calculated by the author (variable $PSPP_h_i$). $PSPP_h_i$ is calculated from data on monthly net purchases (as reported by ECB 2019a; see Figure 16 depicting variable $PSPP_p_i$) by adding up net purchases of government bonds over time. Figure 20 again motivates the selection of sample countries, which accounted for the bulk of PSPP purchases over the sample period. As discussed in section 3, the bulk of the remainder (difference between stacked $PSPP_h_i$ and total $PSPP_h$) is made up of holdings of supnationals' securities purchased by the Eurosystem.

h. Eurosystem's (T)LTRO

As discussed in section 6i, **balance sheet A5 (Lending to euro area credit institutions related to monetary policy operations (MPOs) denominated in euro)** was a major driver of changes of the Eurosystem's overall balance sheet size over the sample period. This section aims to depict the components of A5 in some further detail. As per ECB (2022c), the Eurosystem's net lending to credit institutions can be calculated by subtracting certain liability-side balance sheet elements from A5, but I will focus on the asset side perspective for the sake of visualizing the easing of monetary policy conditions vis-a-vis the interbank market.

The **left-hand-side chart** in figure 21 shows visualizes the ECB's readily available weekly data (ECB 2022c) on the subitems of A5. As per ECB (2022c), the balance sheet item and its subcomponents reflect *"the liquidity-providing monetary policy credit operations used by the Eurosystem"*. Main refinancing operations (MROs) and longer-term refinancing operations (LTROs) account for almost the entirety of A5. Key trends include the shift towards longer-term refinancing sharp expansion of overall liquidity provision in 2008 as well as a shift towards more longer-term refinancing operations; another sharp spike of longer-term liquidity provisioning can be observed in 2012 in the broader context of the "European sovereign debt crisis" as discussed in section 1c. Furthermore, the longer-term trend of waning use of the ECB's main refinancing operations towards the end of the sample period is evident.

Further deeper insights into Eurosystem's liquidity provisioning are provided by data compiled and published by Schnabl and Sonnenberg (2020) depicted on the **right-hand-side chart** of figure 21. The authors use weekly ECB data on (T)LTRO allotments and repayments to compile a daily dataset on the Eurosystem's outstanding refinancing operations. Their dataset gives a more detailed picture, showing the extension of LTRO financing to include other tenors than the "traditional" 3-month LTROs from 2008 onwards, including 1-year LTROs in 2009/10 as well as a further maturity expansion towards 3-year LTROs in 2011/12. The chart also depicts the introduction of Targeted LTROs (TLTROs) from 2014 onwards. As per Schnabl and Sonnenberg (2020), these programs *"offer longer-term funding at accommodative terms, if banks subsequently hand out credits to households and enterprises"*, with borrowing rates having been repeatedly reduced throughout the sample period after the initial introduction of this instrument.

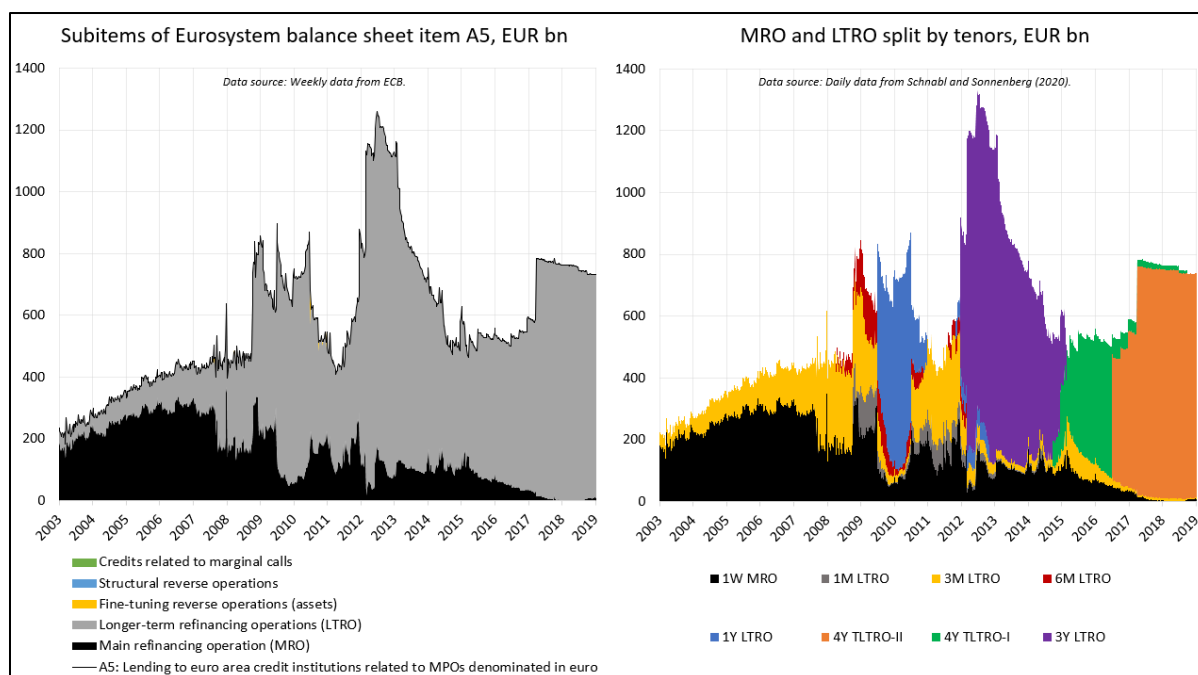


Figure 21 - Eurosystem's balance sheet item A5

Evidently, the Eurosystem's lending to credit institutions has been expanded to longer tenors and has been significantly ramped up in times of crises such as the global financial crisis or the European sovereign debt crisis (in parallel with key rate cuts). This opens the question to which extent the Eurosystem's monetary policy actions towards the interbank market indirectly might have influenced sovereign bond yields over the sample period. However, instead of including the Eurosystem's balance sheet item A5 or any of its subcomponents, I am following economic literature as discussed in section 3g in using short-term market interest rates (3-month Euribor) to control for the effects of monetary policy and banks' funding conditions on sovereign long-term yields.

i. PSPP and WAM

Data on the weighted average maturity (WAM) of securities held by the Eurosystem under the PSPP envelope over the timeframe of my analysis suggests that the WAM of the portfolio moved between 6.3 and 11.65 years across the nine countries in question, with the average

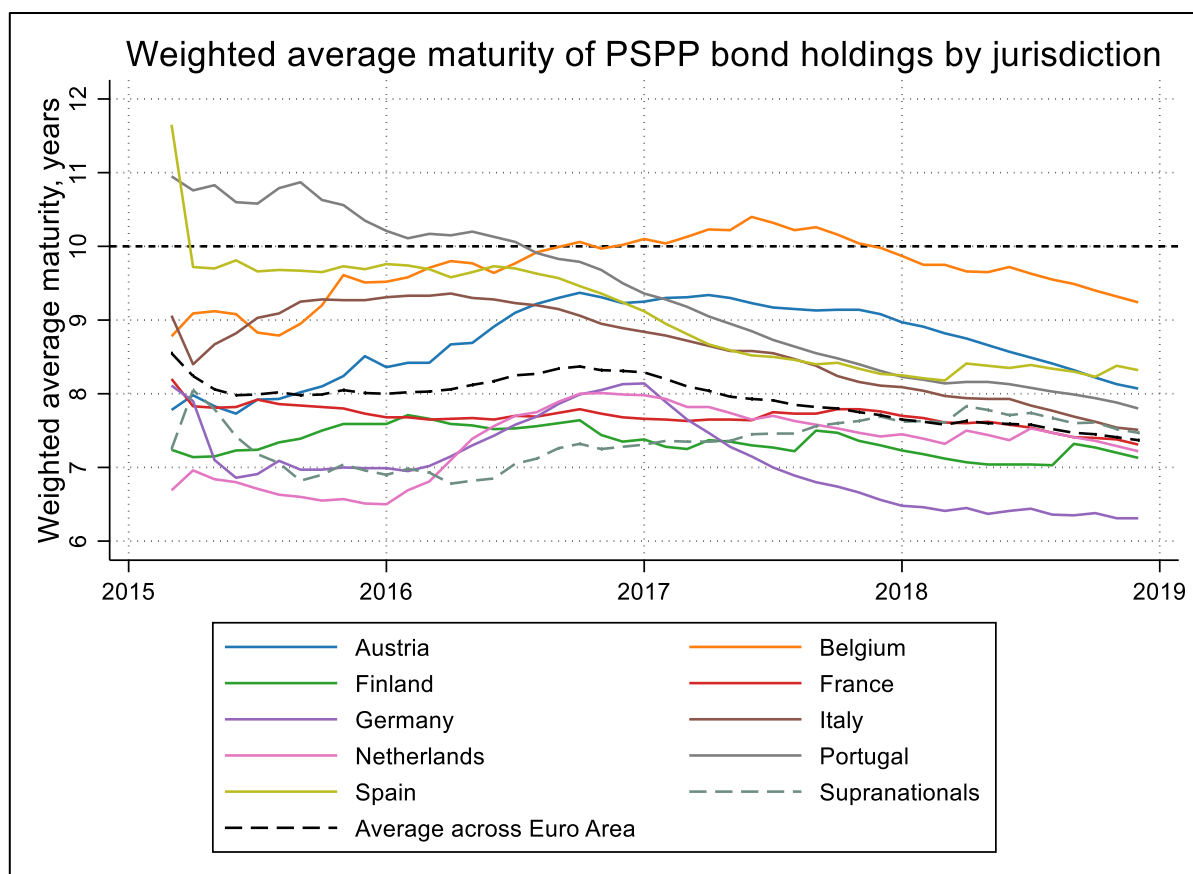


Figure 22 - WAM of Eurosystem PSPP holdings

of the total portfolio across all Eurozone countries close to around 8 years across the period under consideration (see Figure 22; monthly, end of period data, source: ECB 2019a). This suggests that analysing dynamics of bonds with a 10-year maturity is a reasonable choice as a proxy covering a large and relevant segment of sovereign bonds affected by the PSPP. However, it is important to underline that PSPP purchases over the sample period targeted bonds with varying maturities, the weighted average maturity of which was shorter than 10 years for most countries/months. Data on the detailed split of monthly PSPP purchases by country/maturity of purchased bonds is not available publicly. Due to these data limitations, the inclusion of nominal € sums of PSPP purchases or holdings encompassing bonds of all maturities is a simplifying step and implicitly assumes that PSPP purchases of bonds across various maturities might affect yields across the maturity range, including specifically yields of 10-year bonds. More granular analysis using data covering exclusively PSPP purchases of bonds with a 10-year maturity could yield a more clear-cut picture; however, I do not have

access to such data. Keeping this caveat in mind, the results obtained in this thesis should not be generalized for sovereign bonds with shorter or longer maturities.

j. Univariate analysis

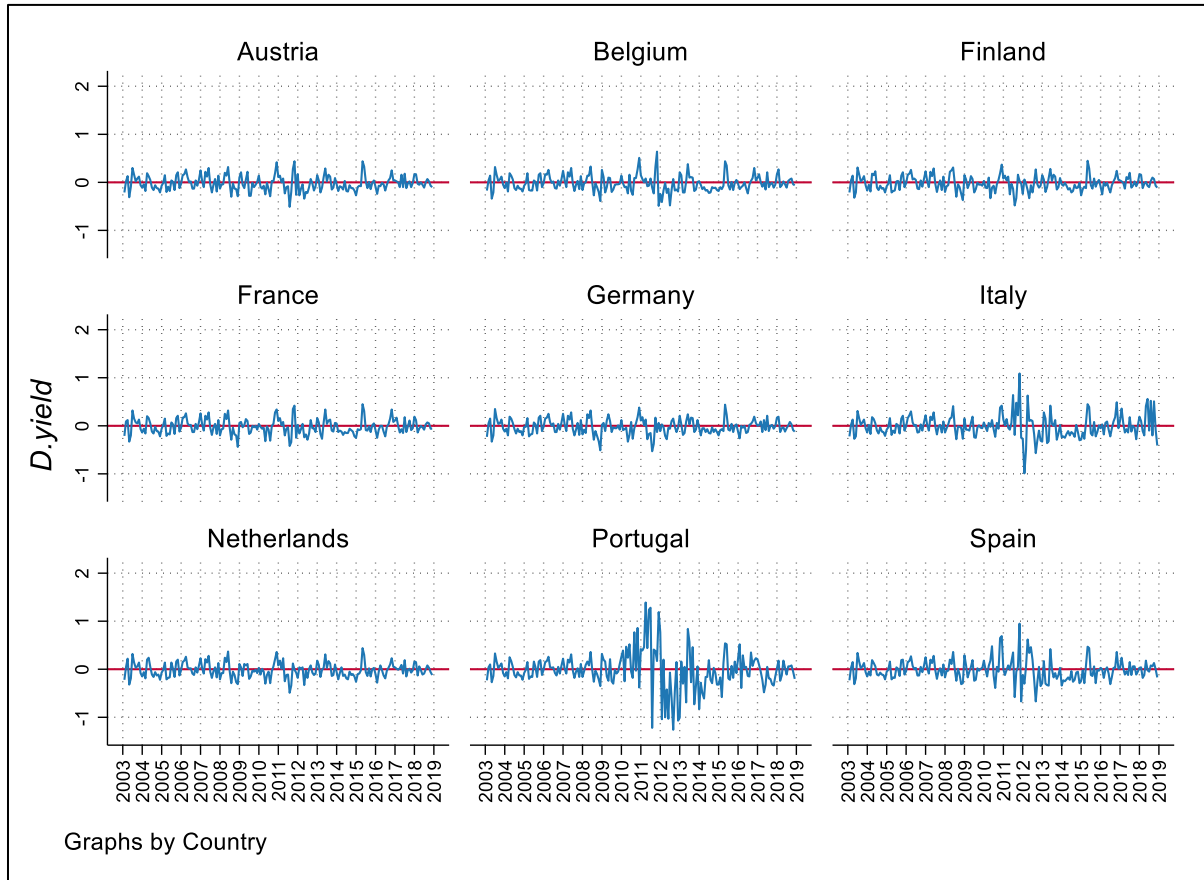


Figure 23 - *D.yield*

Table 11 shows coefficient estimates for the univariate ARMA models for *D.yield*. Significance levels: no star: not significant, *10% significance, ** 5 % significance, *** 1% significance.

<i>D.yield</i>	Austria	Belgium	Finland	France	Germany	Italy	Netherlands	Portugal	Spain	
ARMA(p,q)	(0,1)	(0,1)	(0,1)	(0,1)	(0,1)	(0,1)	(0,1)	(3,0)	(4,2)	
constant	-0.020	-0.019	-0.020	-0.019	-0.021	-0.008	-0.020	-0.015	-0.015	
ar L1.								0.108	-1.212	***
ar L2.								0.147	-0.718	***
ar L3.								0.242	-0.046	***
ar L4.									-0.140	*
ma L1.	0.286	0.351	0.349	0.389	0.301	0.185	0.305		1.455	***
ma L2.									0.962	***

Table 11 - ARMA coefficient estimates for *D.yield*

Tables 12-18 summarize the results of the univariate analysis. Left to right: first 2 columns after “country” indicate when empirical autocorrelation and partial autocorrelation functions cut off respectively; next 2 columns show for which p,q specification information criteria were minimal. Based on BIC, an ARMA(p,q) specification is estimated. Additional checks are performed: overfitting (next 2 columns: significance of estimated added coefficient) and Ljung-Box residual test (rightmost column). In the (rare) case the Ljung-Box test rejects H_0 , the respective entries for overfitting and Ljung-Box refer to the estimated AIC specification.

<i>D.yield</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
Country	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
Austria	1	1	1 ARMA(4,2)	ARMA(0,1)	ARMA(0,1)	0.541	0.528	0.461
Belgium	1	2	2 ARMA(0,2)	ARMA(0,1)	ARMA(0,1)	0.044	0.053	0.910
Finland	1	2	2 ARMA(4,3)	ARMA(0,1)	ARMA(0,1)	0.485	0.602	0.574
France	1	2	2 ARMA(5,2)	ARMA(0,1)	ARMA(0,1)	0.107	0.180	0.955
Germany	1	1	1 ARMA(4,3)	ARMA(0,1)	ARMA(0,1)	0.893	0.893	0.505
Italy	1	1	1 ARMA(0,1)	ARMA(0,1)	ARMA(0,1)	0.589	0.785	0.980
Netherlands	1	1	1 ARMA(0,1)	ARMA(0,1)	ARMA(0,1)	0.682	0.737	0.887
Portugal	3	3	3 ARMA(4,4)	ARMA(3,0)	ARMA(3,0)	0.006	0.328	0.641
Spain	1	1	1 ARMA(4,2)	ARMA(0,1)	ARMA(4,2)	0.962	0.084	0.420

Table 12 - Univariate modelling for *D.yield*

<i>bas</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
Country	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
Austria	tails off		1 ARMA(6,5)	ARMA(1,0)	ARMA(1,0)	0.037	0.166	0.569
Belgium	tails off		1 ARMA(4,3)	ARMA(1,0)	ARMA(1,0)	0.090	0.104	0.477
Finland	tails off		2 ARMA(4,4)	ARMA(1,1)	ARMA(4,4)	non-convergence	0.766	0.337
France	tails off		1 ARMA(7,5)	ARMA(1,3)	ARMA(7,5)	0.999	0.822	0.922
Germany	tails off		1 ARMA(6,2)	ARMA(6,2)	ARMA(6,2)	0.996	0.003	0.794
Italy	tails off		1 ARMA(5,3)	ARMA(3,3)	ARMA(3,3)	0.074	0.139	0.589
Netherlands	tails off		2 ARMA(2,5)	ARMA(0,3)	ARMA(0,3)	0.683	0.664	0.940
Portugal	tails off		2 ARMA(5,6)	ARMA(5,6)	ARMA(5,6)	0.992	0.937	0.407
Spain	tails off		1 ARMA(6,3)	ARMA(1,0)	ARMA(6,3)	0.399	0.003	0.950

Table 13 - Univariate modelling for *bas*

<i>vix</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
All countries	tails off		3 ARMA(1,1)	ARMA(1,1)	ARMA(1,1)	0.321	0.430	0.780

Table 14 - Univariate modelling for *vix*

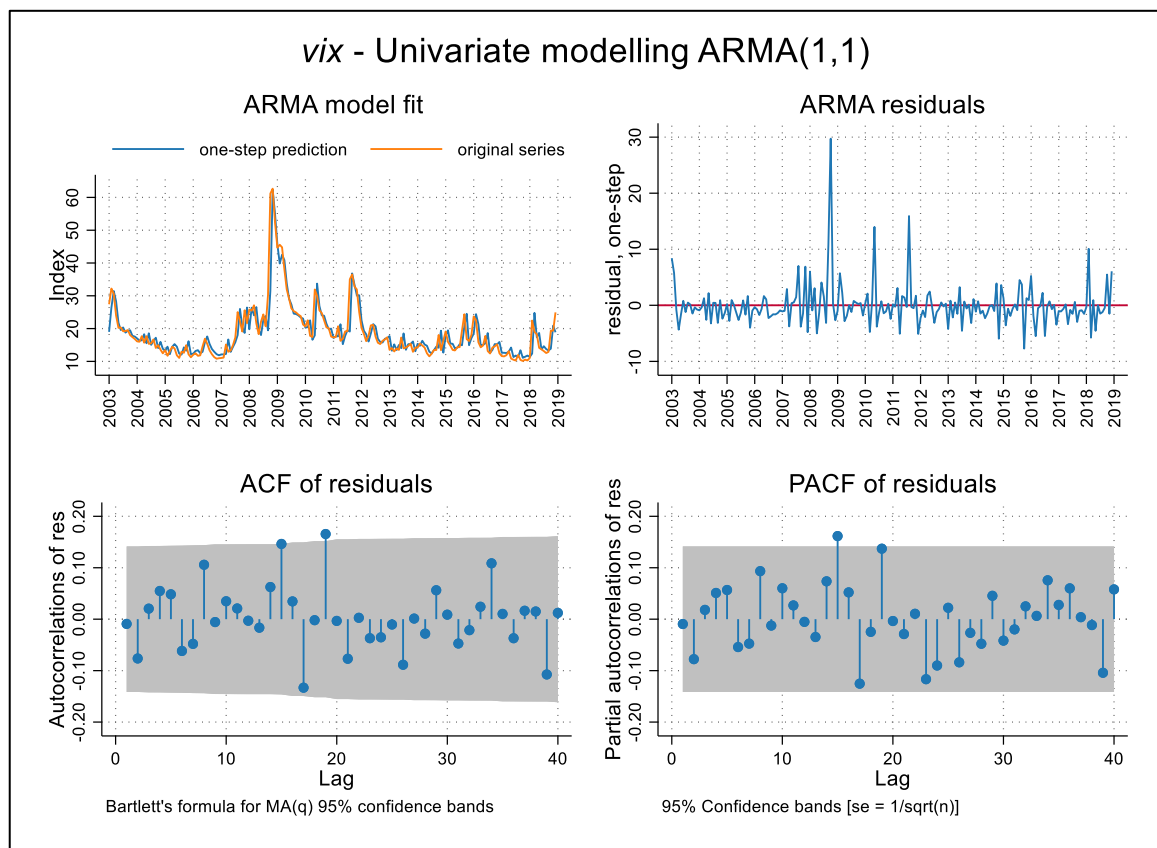


Figure 24 - Univariate modelling for *vix*: model fit and residuals

<i>hicp</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
Country	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
Austria	tails off		2 ARMA(12,1)	ARMA(2,1)	ARMA(12,1)	0.011	0.000	0.409
Belgium	tails off		2 ARMA(12,2)	ARMA(4,2)	ARMA(12,2)	n.a.	0.000	0.296
Finland	tails off		1 ARMA(12,3)	ARMA(1,0)	ARMA(12,3)	0.000	0.000	0.890
France	tails off		2 ARMA(4,2)	ARMA(4,2)	ARMA(4,2)	0.000	0.350	0.637
Germany	tails off		3 ARMA(12,4)	ARMA(4,4)	ARMA(12,4)	0.814	0.703	0.964
Italy	tails off		1 ARMA(8,5)	ARMA(3,2)	ARMA(3,2)	0.975	0.496	0.587
Netherlands	tails off		1 ARMA(12,4)	ARMA(1,0)	ARMA(1,0)	0.761	0.753	0.177
Portugal	tails off		1 ARMA(10,5)	ARMA(1,0)	ARMA(10,5)	0.997	0.317	0.525
Spain	tails off		2 ARMA(5,4)	ARMA(4,2)	ARMA(4,2)	0.003	0.377	0.246
<i>Euro Area</i>	tails off		3 ARMA(12,2)	ARMA(3,0)	ARMA(12,2)	0.826	0.000	0.795

Table 15 - Univariate modelling for *hicp*

<i>esi</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
Country	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
Austria	tails off		1 ARMA(2,2)	ARMA(2,1)	ARMA(2,1)	0.069	0.087	0.836
Belgium	tails off		3 ARMA(2,1)	ARMA(2,1)	ARMA(2,1)	0.469	0.441	0.792
Finland	tails off		1 ARMA(9,4)	ARMA(3,3)	ARMA(3,3)	0.510	0.672	0.913
France	tails off		5 ARMA(7,2)	ARMA(2,1)	ARMA(2,1)	0.181	0.323	0.561
Germany	tails off		4 ARMA(6,2)	ARMA(4,0)	ARMA(4,0)	0.585	0.408	0.887
Italy	tails off		1 ARMA(4,3)	ARMA(1,0)	ARMA(1,0)	0.217	0.172	0.142
Netherlands	tails off		4 ARMA(7,2)	ARMA(4,0)	ARMA(4,0)	0.790	0.840	0.998
Portugal	tails off		3 ARMA(3,1)	ARMA(3,0)	ARMA(3,0)	0.001	0.359	0.743
Spain	tails off		1 ARMA(5,0)	ARMA(2,2)	ARMA(2,2)	0.028	0.538	0.780

Table 16 - Univariate modelling for *esi*

<i>D.euribor</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
Country	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
<i>Euro Area</i>		4	1 ARMA(5,4)	ARMA(5,4)	ARMA(5,4)	non-conv.	0.560	0.994

Table 17 - Univariate modelling for *D.euribor_3m*

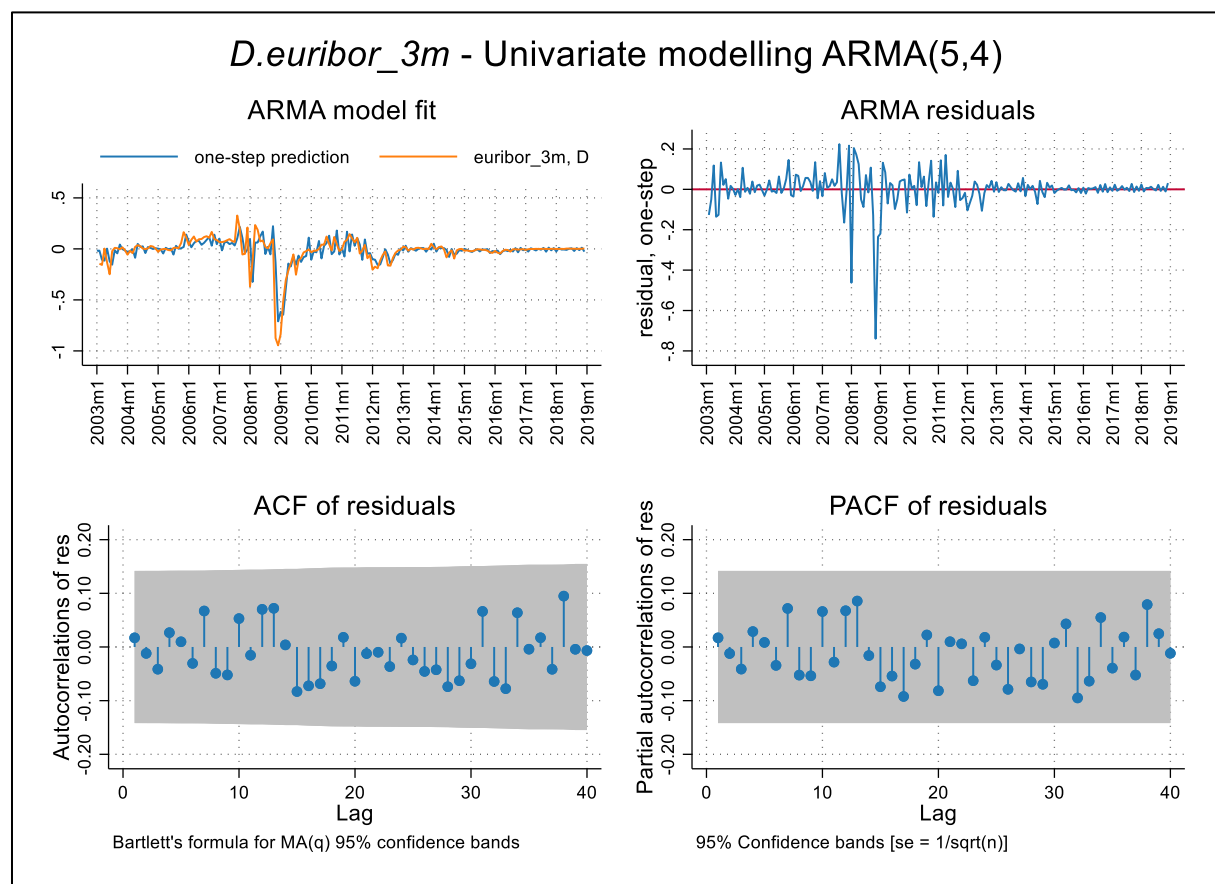


Figure 25 - Univariate modelling for *D.euribor_3m*: model fit and residuals

<i>D.T0</i>	Visual: function cuts off after...		Information criteria suggest...		Chosen specification	Overfitting vs ARMA(p,q)		Ljung-Box residual test
Country	ACF (q)	PACF (p)	AIC	BIC	ARMA(p,q)	q+1: P> z	p+1: P> z	Prob > chi2
Euro Area	3	3	ARMA(7,1)	ARMA(1,1)	ARMA(1,1)	0.420	0.483	0.825

Table 18 - Univariate modelling for *D.T0*

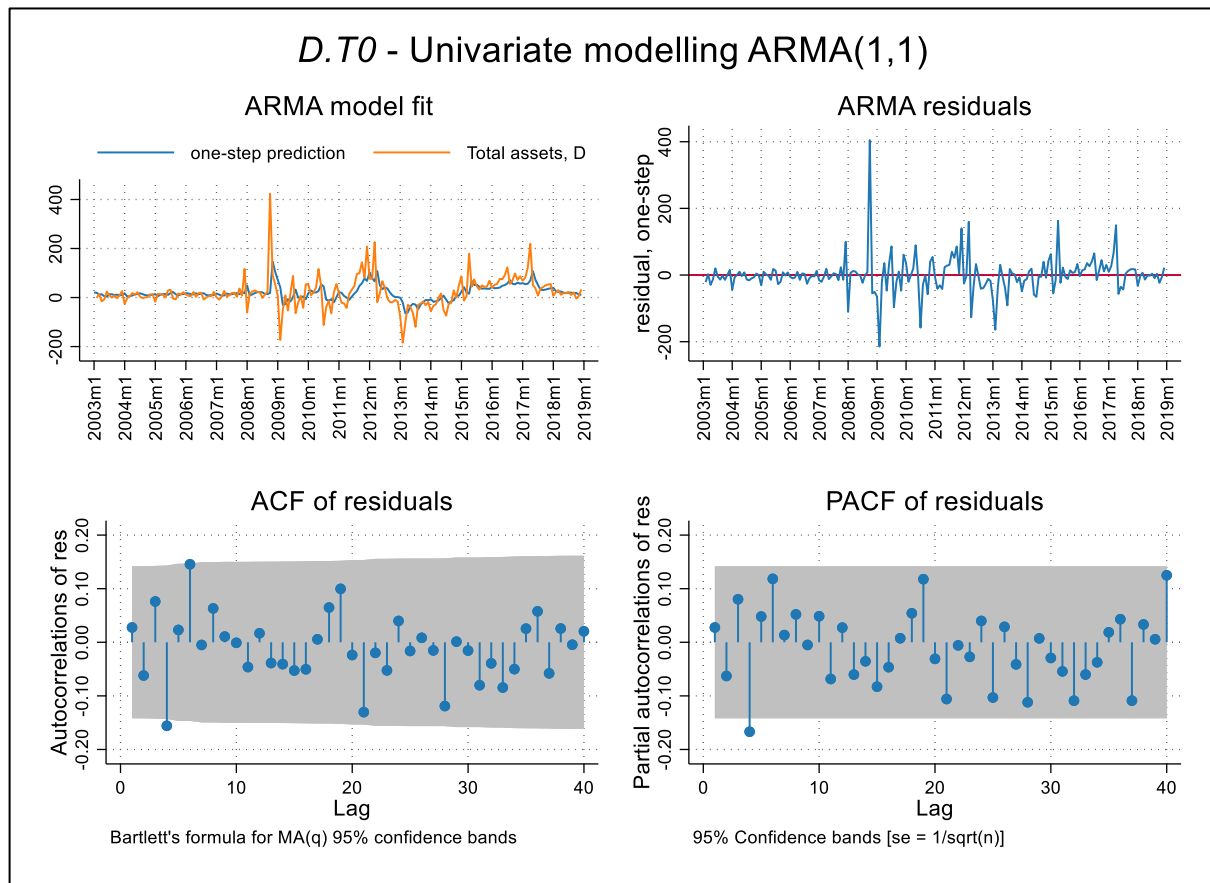


Figure 26 - Univariate modelling for *D.T0* - model fit and residuals

k. Static panel modelling

Table 19 shows the Stata results of BPLM tests (individual random effects vs pooled OLS), F-tests (individual FE vs pooled OLS) and Hausman tests (FE vs RE) for different combinations of periods and regressors. The upper part of the table shows the results for regressions excluding squared government debt/GDP (*debt2*), the lower part shows the results including *debt2*. The columns refer to regressions excluding both *PSPP_i* and *T00*, including only *PSPP_i* and including only *T00* respectively. In the Hausman tests, the FE and RE models with the same list of regressors are compared with each other.

regression w/o debt2	BPLM test: ttest0, Prob > chibar2 =			Test for FE: F-test in Stata output, Prob > F =			Hausman test: hausman FE RE, Prob > chi2 =		
Sample	RE	RE_w_PSPP	RE_w_T00	FE Prob > F	FE_w_PSPP	FE_w_T00	FE	FE_w_PSPP	FE_w_T00
2003-2018	1	1	1	0	0	0	0	0	0
2004-2018	0.0000	1	0.0165	0	0	0	0	0	0
2005-2018	0.0014	1	0.0110	0	0	0	0	0	0
2006-2018	0.0008	1	0.0003	0	0	0	0	0	0
2007-2018	0.0001	0	0	0	0	0	0	0.001	0
2008-2018	0	0	0	0	0	0	0	0	0
2009-2018	0	0	0	0	0	0	0	0	0
2010-2018	0	0	0	0	0	0	0	0	0

regression w/debt2	BPLM test: ttest0, Prob > chibar2 =			Test for FE: F-test in Stata output, Prob > F =			Hausman test: hausman FE RE, Prob > chi2 =		
Sample	RE	RE_w_PSPP	RE_w_T00	FE Prob > F	FE_w_PSPP	FE_w_T00	FE	FE_w_PSPP	FE_w_T00
2003-2018	1	1	1	0	0	0	0	0	0
2004-2018	1	1	1	0	0	0	0	0	0
2005-2018	1	1	1	0	0	0	0	0	0.001
2006-2018	1	1	1	0	0	0	0	0	0
2007-2018	0.0001	1	0	0	0	0	0	0	0
2008-2018	0	1	0	0	0	0	0	0	0
2009-2018	0	1	0	0	0	0	0	0	0
2010-2018	0.0001	1	0	0	0	0	0	0	0

Table 19 - BPLM tests, F-tests, Hausman tests

Tables 20 and 21 show the estimated coefficients for different combinations of sample periods, regressors and standard errors. The test results shown in table 19 are obtained using Stata default standard errors.

Pooled OLS and static one-way (individual effects) panel models: 01/2003 - 12/2018, default standard errors									
	OLS	OLS_w_PSPP	OLS_w_T00	RE	RE_w_PSPP	RE_w_T00	FE	FE_w_PSPP	FE_w_T00
bas	8.221 *** (0.250)	8.124 *** (0.234)	8.053 *** (0.214)	8.221 *** (0.250)	8.124 *** (0.234)	8.053 *** (0.214)	8.590 *** (0.256)	8.591 *** (0.235)	8.161 *** (0.223)
debt	-0.014 *** (0.005)	-0.015 *** (0.005)	-0.012 *** (0.004)	-0.014 *** (0.005)	-0.015 *** (0.005)	-0.012 *** (0.004)	-0.030 *** (0.006)	-0.042 *** (0.006)	-0.028 *** (0.005)
debt2	0.000 ** (0.000)	0.000 *** (0.000)	0.000 ** (0.000)	0.000 ** (0.000)	0.000 *** (0.000)	0.000 ** (0.000)	0.000 * (0.000)	0.000 *** (0.000)	0.000 (0.000)
rating	-0.173 *** (0.013)	-0.171 *** (0.012)	-0.214 *** (0.011)	-0.173 *** (0.013)	-0.171 *** (0.012)	-0.214 *** (0.011)	-0.177 *** (0.026)	-0.229 *** (0.024)	-0.300 *** (0.024)
hicp	0.072 *** (0.020)	0.044 ** (0.019)	0.229 *** (0.018)	0.072 *** (0.020)	0.044 ** (0.019)	0.229 *** (0.018)	0.023 (0.020)	-0.005 (0.018)	0.206 *** (0.019)
esi	-0.049 *** (0.003)	-0.039 *** (0.002)	-0.024 *** (0.002)	-0.049 *** (0.003)	-0.039 *** (0.002)	-0.024 *** (0.002)	-0.045 *** (0.003)	-0.032 *** (0.002)	-0.021 *** (0.002)
euribor_3m	0.764 *** (0.016)	0.693 *** (0.016)	0.346 *** (0.022)	0.764 *** (0.016)	0.693 *** (0.016)	0.346 *** (0.022)	0.665 *** (0.019)	0.561 *** (0.019)	0.286 *** (0.023)
vix	-0.018 *** (0.003)	-0.014 *** (0.003)	0.008 *** (0.003)	-0.018 *** (0.003)	-0.014 *** (0.003)	0.008 *** (0.003)	-0.017 *** (0.003)	-0.012 *** (0.003)	0.009 *** (0.003)
PSPP_p_i		-0.109 *** (0.007)			-0.109 *** (0.007)			-0.130 *** (0.007)	
T00			-0.001 *** (0.000)			-0.001 *** (0.000)			-0.001 *** (0.000)
_cons	10.720 *** (0.390)	9.920 *** (0.369)	10.275 *** (0.333)	10.720 *** (0.390)	9.920 *** (0.369)	10.275 *** (0.333)	11.975 *** (0.708)	12.564 *** (0.651)	13.016 *** (0.616)
Number of observations	1724	1724	1724	1724	1724	1724	1724	1724	1724
R-squared	0.78	0.81	0.84				0.75	0.79	0.81
Adjusted R-squared	0.78	0.81	0.84				0.75	0.79	0.81
AIC	4186.61	3964.49	3646.07				4072.19	3779.16	3583.48
BIC	4235.68	4019.02	3700.59				4121.26	3833.68	3638.00

Table 20 - Pooled OLS and country-specific effects panel models: 2003-2018

Pooled OLS and static one-way (individual effects) panel models: 01/2008 - 12/2018, robust standard errors									
	OLS	OLS_w_PSPP	OLS_w_T00	RE	RE_w_PSPP	RE_w_T00	FE	FE_w_PSPP	FE_w_T00
bas	7.746 *** (0.423)	7.742 *** (0.397)	7.353 *** (0.368)	7.751 *** (0.861)	7.742 *** (0.687)	7.356 *** (0.713)	7.712 *** (0.969)	7.917 *** (0.831)	7.213 *** (0.659)
debt	0.004 (0.006)	0.002 (0.006)	-0.000 (0.006)	0.004 (0.012)	0.002 (0.011)	-0.003 (0.010)	-0.000 (0.035)	-0.019 (0.034)	-0.016 (0.025)
debt2	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)
rating	-0.219 *** (0.016)	-0.212 *** (0.015)	-0.253 *** (0.015)	-0.208 *** (0.056)	-0.212 *** (0.043)	-0.242 *** (0.043)	-0.103 (0.211)	-0.174 (0.168)	-0.222 (0.134)
hicp	0.167 *** (0.028)	0.129 *** (0.027)	0.357 *** (0.030)	0.169 ** (0.076)	0.129 * (0.073)	0.363 *** (0.087)	0.175 ** (0.072)	0.132 * (0.063)	0.368 *** (0.072)
esi	-0.055 *** (0.003)	-0.045 *** (0.003)	-0.033 *** (0.003)	-0.055 *** (0.012)	-0.045 *** (0.011)	-0.033 *** (0.012)	-0.055 *** (0.011)	-0.040 *** (0.011)	-0.030 * (0.013)
euribor_3m	0.589 *** (0.022)	0.554 *** (0.021)	0.220 *** (0.028)	0.581 *** (0.065)	0.554 *** (0.070)	0.202 ** (0.083)	0.475 *** (0.103)	0.405 *** (0.102)	0.081 (0.079)
vix	0.001 (0.004)	0.002 (0.004)	0.017 *** (0.003)	0.001 (0.005)	0.002 (0.006)	0.016 *** (0.005)	-0.007 (0.006)	-0.004 (0.006)	0.010 (0.006)
PSPP_p_i		-0.092 *** (0.006)			-0.092 *** (0.019)			-0.114 *** (0.020)	
T00			-0.001 *** (0.000)			-0.001 *** (0.000)			-0.001 *** (0.000)
_cons	10.932 *** (0.517)	10.119 *** (0.498)	11.153 *** (0.447)	10.824 *** (1.968)	10.119 *** (1.793)	11.151 *** (1.859)	9.826 * (4.985)	11.063 ** (3.902)	12.206 *** (2.894)
Number of observations	1188	1188	1188	1188	1188	1188	1188	1188	1188
R-squared	0.81	0.83	0.85				0.77	0.80	0.82
Adjusted R-squared	0.81	0.83	0.85				0.77	0.80	0.82
AIC	3002.53	2865.89	2733.74				2845.83	2651.99	2543.02
BIC	3048.25	2916.69	2784.54				2886.47	2692.63	2583.66

*** p<.01, ** p<.05, * p<.1

Table 21 - Pooled OLS and country-specific effects panel models: 2008-2018

Table 22 shows two-way FE & RE estimates produced with R for the full sample 2003-2018:

	FE2	FE2_w_PSPP	FE2_w_T00	RE2	RE2_w_PSPP	RE2_w_T00
bas	6.843 *** (0.128)	6.879 ** (0.127)	6.843 ** (0.128)	6.965 *** (0.350)	7.025 *** (0.350)	7.054 *** (0.345)
debt	-0.027 *** (0.003)	-0.029 *** (0.003)	-0.027 *** (0.003)	-0.023 ** (0.008)	-0.027 *** (0.008)	-0.019 ** (0.007)
debt2	0.000 *** (0.000)	0.000 *** (0.000)	0.000 *** (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)
rating	-0.356 *** (0.013)	-0.364 *** (0.013)	-0.356 *** (0.013)	-0.315 *** (0.032)	-0.331 *** (0.034)	-0.292 *** (0.028)
hicp	0.010 (0.016)	0.004 (0.016)	0.010 (0.016)	0.011 (0.041)	0.003 (0.041)	0.040 (0.040)
esi	-0.024 *** (0.002)	-0.023 *** (0.002)	-0.024 *** (0.002)	-0.027 *** (0.005)	-0.026 *** (0.005)	-0.025 *** (0.005)
PSPP_p_i		-0.026 *** (0.005)			-0.036 ** (0.013)	
(Intercept)				11.812 *** (0.999)	12.409 *** (1.025)	12.355 *** (0.925)
euribor_3m				0.784 *** (0.083)	0.748 *** (0.072)	0.462 *** (0.100)
vix				-0.002 (0.015)	-0.002 (0.013)	0.007 (0.013)
T00						-0.001 *** (0.000)
Observations	1724	1724	1724	1724	1724	1724
R ² / R ² adjusted	0.845 / 0.824	0.848 / 0.827	0.845 / 0.824	0.748 / 0.747	0.760 / 0.759	0.818 / 0.817

* p<0.05 ** p<0.01 *** p<0.001

Table 22 - Two-way panel models: 2003-2018

I. Dynamic panel modelling

Table 23 shows the results of estimated GMM models excluding *debt* & *debt2*, which were not significant when included in the full set of regressors. Table 24 shows the results of estimated GMM models for the full set of regressors for a shorter sample (2008-2018). *sarganp* & *hansenp*: p-values of Sargan and Hansen test statistics. *ar1p* and *ar2p*: p-values of Arellano-Bond test statistics for serial correlation AR(1) and AR(2). *j0*: number of instruments.

Dynamic one-step panel models: 01/2003 - 12/2018, robust standard errors, no debt / debt2						
Variable	DIF_GMM	DIF_GMM_PSPP	DIF_GMM_T00	SYS_GMM	SYS_GMM_PSPP	SYS_GMM_T00
yield L1.	.81795537***	.81340886***	.81795536***	.85884926***	.86035677***	.85884928***
bas	1.9406985***	2.0548446***	1.9406986***	1.2786583***	1.2685523***	1.2786582***
rating	-.01180865	-.01545728	-.01180866	-.02936208***	-.0289856***	-.02936208***
hicp	.01684143**	.00214214	.01684143**	.00898943	.00873051*	.00898943
esi	.00204582	.00189388	.00204582	-.00375216***	-.00371667***	-.00375216***
euribor_3m	.16116782***	.16768996***	.0230153***	.05315818	.02165854	.06887173*
vix	-.00554459***	-.00573727***	-.0010005	.00010843	-.00041994	.00143847**
PSPP_p_i T00		-.01467367	-.00021083***		-.00099552	.0000524***
sarganp	7.628e-24	9.593e-23	7.628e-24	1.148e-29	1.412e-29	1.148e-29
hansenp	1	1	1	1	1	1
ar1p	.09574435	.09414564	.09574435	.08751314	.0881977	.08751314
ar2p	.43813158	.44474788	.43813161	.41927032	.41863912	.41927032
j0	387	388	388	389	390	390

Legend: * p<.1; ** p<.05; *** p<.01

Table 23 - One-step GMM models: 2003-2018 without debt, debt2

Dynamic one-step panel models: 01/2008 - 12/2018, robust standard errors						
Variable	DIF_GMM	DIF_GMM_PSPP	DIF_GMM_T00	SYS_GMM	SYS_GMM_PSPP	SYS_GMM_T00
yield L1.	.81715751***	.81542655***	.81715751***	.86213628***	.86307952***	.86213628***
bas	2.1148618***	2.1299209***	2.1148618***	1.1894702***	1.1837523***	1.1894702***
debt	-.03119572	-.03055525	-.03119572	.00121134	.00121234	.00121134
debt2	.00015326	.0001443	.00015326	-8.829e-06	-8.799e-06	-8.829e-06
rating	.01603559	.00819717	.01603559	-.03385823**	-.03356936**	-.03385823**
hicp	.00987256	.0026417	.00987256	.02278405*	.02263977*	.02278405*
esi	.00191251	.00205778	.00191251	-.00526063***	-.00522844***	-.00526063***
euribor_3m	.11372247***	.11549798***	-.01964427	.12981218***	.1005426***	.15385238***
vix	-.00432667***	-.00446035***	-.00105882	-.00788266***	-.00828216***	-.00834617***
PSPP_p_i T00		-.00995202	-.00021814***		-.00025364	.00006215***
sarganp	4.733e-12	6.354e-12	4.733e-12	2.368e-16	2.539e-16	2.368e-16
hansenp	1	1	1	1	1	1
ar1p	.10504313	.10271627	.10504313	.08819461	.08823003	.08819461
ar2p	.4538318	.44542849	.4538318	.42842443	.42834279	.42842443
j0	269	270	270	271	272	272

Legend: * p<.1; ** p<.05; *** p<.01

Table 24 - One-step GMM models: 2008-2018

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