European economic integration – the role of monetary policy

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vorgelegt von
Isabella Moder, B.A. M.A

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Chapter 1

Introduction and motivation

International spillovers from the monetary policy of one country to other economies are a corollary of globalisation. This entails that we, as policymakers, have to rise to the challenge of conducting monetary policy in the presence of these unintended side-effects. [...] We should not underestimate the challenges of living in the ever more closely interconnected global economy.

– Vítor Constâncio, former Vice President of the European Central Bank
Joint FED/ECB/FED Dallas/HKMA Conference, 15 October 2015, Hong Kong[1]

In a world that is more than ever economically interconnected, the question arises to what extent national policy making is still able to control domestic financial and economic conditions. This applies especially to small open economies in Europe that are in the vicinity but not yet part of the euro area, as the European integration project may be considered globalisation on a smaller scale (Ther, 2019). The overarching research focus of this thesis is thus to what extent financial and economic conditions in central, eastern and south-eastern Europe (CESEE) are dominated by euro area monetary policy, whether this leaves room to manoeuvre for domestic monetary policy (especially in the presence of unofficial financial euroisation) and how costly the potential loss in independent policy making ultimately is.

Roughly three decades after Francis Fukuyama's famously proclaimed 'the end of history' that marked the beginning of the European transformation process, the countries located in CESEE are integrated to a different extent with the European institutions. While some CESEE countries have already adopted the euro[2], other countries are EU members but not yet part of the euro area[3], or even not yet part of the EU[4]. Economically, the CESEE countries are already very closely interlinked with the euro area. High trade integration and sizeable remittance

[1] see `https://www.ecb.europa.eu/press/key/date/2015/html/sp151015.en.html`

[2] Those countries consist of Estonia, Latvia, Lithuania, Slovenia and Slovakia and are not covered in this thesis.

[3] Those include Bulgaria, Croatia, Czech Republic, Hungary, Poland and Romania

[4] Those include Albania, Bosnia and Herzegovina, Kosovo, Montenegro, North Macedonia and Serbia.

flows (especially to south-east European countries) constitute real economic links, while on the financial side the euro area and CESEE are connected through the presence of a number of euro area headquartered bank subsidiaries and a high degree of unofficial financial euroisation.[5]

The convergence and monetary policy history of the CESEE countries can broadly be grouped into three decades (Backé et al., 2019; Grieveson et al., 2019): In the first decade that marked the early transition years characterised by rapid transformation, monetary policy focused mainly on disinflation while monetary policy frameworks, targets and instruments were established. During the second decade of transition, the 'boom years', CESEE countries experienced rapid economic growth and buoyant credit growth, which was very often accompanied by the build-up of considerable foreign currency debt. In the third decade since the global financial crisis, CESEE countries continued to catch-up, although to a much lesser extent than before the crisis, prompting some discussion among academics and policy makers whether the convergence process has stalled and what prospects remain for convergence going forward (see e.g. Papi et al., 2018).

In the early years of transition, the currencies of the CESEE countries were often pegged to a key currency (the Deutschmark) or to a currency basket (Backé et al., 2019). Over time, some countries adopted inflation targeting regimes and made exchange rates more flexible. Therefore, despite the many common characteristics CESEE countries share otherwise, today's landscape of monetary and exchange rate policies in CESEE is very heterogeneous (table 1.1). For the group of CESEE EU members, a majority of countries operate under inflation-targeting frameworks (Czech Republic, Hungary, Poland and Romania) with the remainder using exchange rate anchors (Bulgaria and Croatia) with the euro as base currency. In countries with flexible exchange rate regimes, exchange rates are allowed to vary to a different degree, from a stabilised arrangement in the case of Romania to Poland where the exchange rate is allowed to float freely. In contrast, non-EU CESEE countries seem to prefer arrangements with more exchange rate stability, as inflation targeters (Albania and Serbia) are outnumbered by other monetary policy regimes and only one country (Albania) is classified as letting its exchange rate floating at all. Moreover, Kosovo and Montenegro are special cases as they do not have a separate legal tender, but use the euro as their legal tender instead, without being members of the euro area.

During the transition of CESEE countries from central planning economies to open market economies, the project of a common European currency took concrete form with the euro being introduced as an electronic currency in January 1999 and physically in January 2002. While the first decade of the euro was characterised by decent growth in real incomes and an upside distribution of shocks to inflation, the second decade of the euro was marked by instability and crisis with the euro area being hit by disinflationary forces (Rostagno et al., 2019). In response to the new challenges, the European Central Bank (ECB) introduced a number of non-standard monetary policy measures, which were unprecedented in nature, scope and magnitude, ranging from significant changes in its operational framework to large asset purchasing programmes. At

[5] Unofficial financial euroisation describes a situation where economic agents voluntarily choose to hold a share of their deposits in euro, or take out loans in euro instead of the national legal tender.

Table 1.1: Monetary policy and exchange rate regimes in CESEE

Country	*Monetary policy regime*	*Exchange rate regime*
EU members		
Bulgaria	Exchange rate anchor	Currency board
Croatia	Exchange rate anchor	Stabilised arrangement
Czech Republic	Inflation-targeting framework	Floating exchange rate
Hungary	Inflation-targeting framework	Floating exchange rate
Poland	Inflation-targeting framework	Free floating exchange rate
Romania	Inflation-targeting framework	Stabilised arrangement
Prospective EU members		
Albania	Inflation-targeting framework	Floating exchange rate
Bosnia and Herzegovina	Exchange rate anchor	Currency board
Kosovo	Exchange rate anchor	No separate legal tender
Montenegro	Exchange rate anchor	No separate legal tender
North Macedonia	Exchange rate anchor	Stabilised arrangement
Serbia	Inflation-targeting framework	Stabilised arrangement

Source: IMF AREAER 2018 database. Note: The table does not include CESEE countries that have already adopted the euro as they are not covered in this thesis.

the same time, during the 21 years of its existence, the euro area extended from its initial eleven members to a total of 19 countries, including some of the former CESEE transition economies.

Overall, the relationship between the euro area and CESEE countries provide an interesting case for answering the overarching questions of this thesis for several reasons: Because of the very close economic ties between the euro area and CESEE, but at the same time a considerable asymmetry in their economic size, it can be investigated how financial and economic conditions in small open economies are dominated by monetary policy of a large advanced economy (or currency union in the case of the euro area), both in conventional and unconventional times. Moreover, their heterogeneous monetary policy and exchange rate regimes allow a comparison between CESEE countries to assess whether domestic monetary policy can counteract monetary policy spillovers, especially in the presence of unofficial financial euroisation. Lastly, in order to assess the cost related to the loss in monetary policy autonomy, the strong economic ties between the euro area and CESEE allow for an assessment in the drivers of business cycle synchronization.

This thesis consists of three self-contained research papers. The first and the second paper both focus on the implications of euro area monetary policy for CESEE countries, but under special circumstances: First, spillovers of the ECB's non-standard monetary policy measures on a subset of CESEE countries are assessed and compared to the literature on spillovers from conventional monetary policy. In a similar vein, the second paper investigates the interest rate pass-through of euro area monetary policy to interest rates of euro deposits and euro loans in CESEE countries, shedding light on monetary policy spillovers in the presence of unofficial financial euroisation and the consequences for domestic monetary policy. The third paper complements the first two papers by addressing the consequences of the relative loss in domestic monetary policy found in the first two papers through examining the cost in terms of business cycle stabilisation and the drivers of business cycle synchronization.

Based on the literature on the international effects of conventional monetary policy, the **first paper** looks at the consequences of the ECB's non-standard monetary policy measures for south-eastern Europe (SEE) that can be regarded as transition countries with respect to their economic development stage and are characterised by relatively shallow capital markets. Three questions are addressed in this paper: first, how have the ECB's non-standard monetary policy measures been affecting the SEE countries? Second, which channels are transmitting these shocks to SEE? Third, do different exchange rate regimes play a role in the SEE countries' responses to the shock? I overcome the challenges of empirically assessing non-standard monetary policy measures by using Eurosystem balance sheet assets as the main proxy for non-standard monetary policy measures, and through an identification strategy that allows to distinguish between endogenous (demand-driven) and exogenous monetary expansions (following the approach developed by Boeckx et al., 2017; Burriel and Galesi, 2018). To model spillovers from the euro area's non-standard monetary policy to SEE countries, I employ structural Bayesian models for each SEE country, through which I am able to assess the response of output and prices for each country as well as to shed more light on potential channels of international shock transmission.

The results suggest that the price level of all countries is positively affected by an expansionary non-standard monetary policy shock originating in the euro area, in line with the importance of euro-area imports in total imports and the expansion of domestic activity. With regard to the output response, the shock has an expansionary effect in approximately half of the countries. These results are confirmed by robustness checks. With respect to possible transmission channels, I find that spillovers seem to be mostly transmitted via the export channel. On the contrary, the interest rate channel exhibits a pronounced response only in a few countries, which might be driven by the relative illiquid interbank market in the SEE countries. Nevertheless, financial flows in the form of foreign direct or portfolio investments, which are not captured in the model, still might play a role. Furthermore, the results suggest that the exchange rate regime does not play a role in determining the sign and magnitude of price level and output responses. This is in line with the absence of distinct exchange rate responses in the model output, suggesting that exchange rates did not act as buffers for spillovers of euro area non-standard monetary policy measures on south-eastern Europe during the sample period.

In the **second paper**, I undertake a comprehensive analysis of the interest rate pass-through of euro area monetary policy to euro retail rates outside the euro area, contributing to the literature on transmission channels of monetary policy spillovers and on the consequences of unofficial financial euroisation. Unofficial financial euroisation is a common phenomenon in

CESEE countries, as over 40% of total outstanding loans and more than one third of total deposits are denominated in euro instead of the national currency. Two main questions are addressed: First, does a long-run relationship between euro area monetary policy and euro retail rates in unofficially euroised economies exist? Second, if such a relationship exists, to what extent can domestic monetary policy influence the 'euro part' of the interest rate channel? In order to answer the questions, I use two different methods. First and foremost, I estimate the interest rate pass-through of the EONIA to a comprehensive dataset of 200 time series of euro retail rates in eight CESEE countries, using an ARDL model in the spirit of Pesaran and Shin (1998) and Pesaran et al. (2001). To be able to compare the pass-through of euro area monetary policy with the pass-through of domestic monetary policy, the same exercise is thereafter undertaken with domestic monetary policy rates (where applicable). Lastly, to complement the evidence found in the previous steps, I investigate the dynamic interdependence between the respective interest rates through employing structural VAR models for a subset of three countries.

The results suggest that in the long run more than one third of all euro retail rates in euroised CESEE countries are linked to euro area monetary policy. Moreover, euro retail rates in CESEE adjust excessively to a change in the EONIA, and a deviation from the long-run relationship is corrected by 20% already within the next month. Compared to euro area monetary policy, domestic monetary policy has less of an influence on euro retail rates. This suggests that domestic central banks in countries with independent monetary policy can only partially control the 'euro part' of the interest rate channel, raising the question to which extent monetary policy in those CESEE countries can lean against the 'euro area wind'. For countries with a fixed exchange rate regime linked to the euro, the consequences seem to be less dramatic, with the direct interest rate pass-through to euro retail rates likely to constitute an additional channel of monetary policy transmission.

Finally, the aim of the **third paper** is to assess the cost of the partial loss of independent monetary policy making in terms of business cycle stabilisation for a subset of Western Balkan countries. More specifically, it investigates the degree of business cycle synchronization (BCS) between the Western Balkans and the European Union, and empirically tests which factors are driving business cycle synchronization. To that end, we make use of some recent advances in the business cycle literature by using time-varying correlation index developed by Cerqueira and Martins (2009) and Cerqueira (2013) to analyze the convergence process on a yearly basis and to obtain a panel data set for the regressions. To identify the determinants of BCS, we start by looking at well-known factors identified in the literature (trade, specialization, fiscal policy) but extend the choice of variables by adding other potential channels such as common monetary policy, financial flows and remittances, which might be relevant especially for the Western Balkans.

The results indicate that while the degree of synchronization had been low or even negative before 2000, business cycles have clearly converged ever since. Since 2010, in particular, BCS with the EU has been high for all Western Balkan countries. With respect to the drivers of

BCS, we find that foreign trade is the most important positive contributor to business cycle convergence. Although fiscal differences are associated with negative BCS in the same year, our results suggest that they have a positive influence in subsequent periods. In contrast, we find that financial flows lead to business cycle decoupling because of their procyclical behavior in the respective domestic economy. The same relationship applies, according to our analysis, to remittances, whose impact on BCS is an underresearched topic in the empirical literature. We find that remittances from the EU to the Western Balkans behave similarly to financial flows, which supports the hypothesis that remittances are sent home to take advantage of favorable economic conditions.

Subsequent to the third paper, I provide a final conclusion and discussion of the consolidated results of the individual papers and some reflections on the road ahead.

Chapter 2

Spillovers from the ECB's non-standard monetary policy measures on south-eastern Europe

Author: Isabella Moder, B.A. M.A.

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Abstract

This paper is the first to comprehensively assess the impact of the euro area's non-standard monetary policy measures on south-eastern Europe. The outcomes of bilateral BVAR models suggest that the ECB's non-standard monetary policy measures have had pronounced price effects on all, and output effects on approximately half, of the countries in south-eastern Europe. While I find evidence that exports have posed as relevant transmission channels in most cases, the role of the interbank market rate as a channel of shock transmission is less clear. Furthermore, the results suggest that exchange rates' responses have been relatively muted.

2.1 Introduction

Since October 2008 the European Central Bank (ECB) has introduced a number of non-standard monetary policy measures, which are unprecedented in nature, scope and magnitude; and have ranged from significant changes in the operational framework to large bond purchasing programmes. Assessing potential spillovers from monetary policy measures of advanced economies has become important in a globalised world; and it does not only incorporate potential spillovers via real channels like trade links and remittance flows, but more and more also the impact of financial spillovers, as monetary policy measures often generate sizable changes in capital flows and exchange rate dynamics. This mechanism could be very well observed in the so-called 'taper tantrum episode' in mid-2013, when the Fed announced to gradually turn off its bond-buying programme, which provoked a pronounced shift in market sentiment vis-à-vis emerging markets (see Sahay et al., 2014). Quantifying the direction and magnitude of international spillovers caused by advanced economies' monetary policy measures – and identifying the main transmission channels – is thus of utmost importance for policy makers in order to design optimum policy responses, both to spillovers from the introduction of such measures as well as to spillovers from their potential reversal.

The focus of interest for this paper lies in potential spillovers of ECB monetary policy measures to European countries that are not yet part of the euro area, or are in the process of EU accession. More specifically, this paper deals with the countries of south-eastern Europe (SEE)[1] that can be regarded as transition countries with respect to their economic development stage.[2] SEE countries are interlinked with the euro area through various channels. High trade integration and sizable remittance flows constitute potential real transmission channels, while the presence of a number of euro area headquartered bank subsidiaries and (correspondingly) a high degree of euroisation[3] represent financial links.

Additionally, the heterogeneous monetary policy regimes of SEE countries provide an interesting case for cross-country comparisons with regard to the role of exchange rate regimes in shaping spillovers: exchange rate regimes in SEE range from inflation targeters with (managed) floating exchange rates (Albania, Romania and Serbia), to stabilised arrangements with the euro as a reference currency (Croatia and North Macedonia), to euro-based currency boards (Bosnia and Herzegovina as well as Bulgaria), to the unilateral adoption of the euro as the sole legal tender (Montenegro).

The aim of this paper is thus to answer three questions: First, in what direction and to which magnitude have the ECB's non-standard monetary policy measures been affecting the SEE countries? Second, through which channels are these shocks transmitted to SEE? Third,

[1] Albania, Bosnia and Herzegovina, Bulgaria, Croatia, Montenegro, North Macedonia, Romania and Serbia.

[2] In contrast to Baltic and central European countries, where convergence towards the 'old' EU member states progressed further than in SEE countries.

[3] Either through high unofficial asset and liability euroisation of the banking systems or, in the case of Montenegro, through the use of the euro as the legal tender, see European Central Bank (2019) for more information.

do different exchange rate regimes play a role in shaping the SEE countries' responses to a non-standard monetary policy shock?

The main contribution of this paper to the literature is the systematic examination of spillovers from the ECB's non-standard monetary policy measures to the whole SEE region. While its three EU members (Bulgaria, Croatia and Romania) have already been covered to a certain extent in the spillover literature, no research has been undertaken yet for the remaining countries, which are five candidate and potential candidate countries to the EU.[4] By employing impulse response functions in a structural BVAR setting, the effect of non-standard monetary policy shocks on each country's output, price level, exports, short-term interest rate and (if applicable) exchange rate is estimated. In identifying the non-standard monetary policy shocks, I follow the approach developed by Boeckx et al. (2017) and Burriel and Galesi (2018) to examine the domestic transmission of monetary policy, and extend it to investigate the international effects of euro area monetary policy.

The results show that the price level of all countries is positively affected by an expansionary non-standard monetary policy shock originating in the euro area, and for approximately half of the countries output also responds in a positive and pronounced way. Furthermore, spillovers seem to be mostly transmitted via exports (which impact SEE countries' output positively in the absence of exchange rate responses), while only in a few cases the interbank interest rate (used as a proxy for the financial channel) exhibits a clear response to the shock. Additionally, the results suggest that in countries operating under a flexible exchange rate regime, exchange rates do not react strongly to the non-standard monetary shock, which is in line with the relatively stable exchange rates during the sample period.

The structure of the paper is as follows: Section 2.2 provides an overview of the literature, while section 2.3 introduces the methodological approach. The corresponding results and potential transmission channels are discussed both from a cross-country perspective and for each country individually in section 2.4. Section 2.5 documents robustness tests undertaken in order to cement the results. Section 2.6 concludes.

2.2 Related literature

2.2.1 Theoretical foundations

This subsection provides a brief overview of possible spillovers of a foreign monetary policy shock in the Mundell-Fleming and AD-AS models, as well as in the Corsetti-Pesenti model, based on Gärtner (2013) and Harms (2016).

The classic Mundell-Fleming model (see Mundell, 1962; Fleming, 1962) represents a small

[4] Due to data limitations, the remaining prospective EU member Kosovo cannot be included in the empirical analysis.

open economy under the assumption of perfect capital mobility, and distinguishes between flexible and fixed exchange rate regimes. In the dynamics of the model, an expansionary foreign monetary policy shock decreases the world interest rate, which is treated as an exogenous variable for the small open economy. Because the domestic interest rate is higher than the foreign interest rate, the interest rate differential attracts financial inflows, leading to appreciation pressures on the domestic exchange rate.

How the domestic exchange rate reacts depends on whether the country operates under a fixed or a flexible exchange rate regime. In the case of a fixed exchange rate regime, the central bank cannot let the exchange rate appreciate and thus has to intervene in the foreign exchange market by buying foreign currency and selling local currency in order to keep the exchange rate at its pre-determined level. As a consequence, output increases to a new equilibrium level in line with the lower world and domestic interest rate. In contrast, if the small open economy operates under a freely floating exchange rate regime, both its nominal and real (in the absence of price adjustments) exchange rates appreciate as the interest rate differential between the lower world interest rate and the higher domestic interest rate attracts financial inflows. The exchange rate appreciation affects output negatively as long as the domestic interest rate remains unchanged. In case the central bank responds by lowering the domestic interest rate as well, the interest rate differential with the world interest rate narrows, which dampens capital inflows and the exchange rate appreciation, and stimulates domestic output.

Besides the financial channel, a foreign monetary policy shock also affects the small open economy's net exports, assuming that the foreign country and the domestic country trade with each other. If a foreign expansionary monetary policy shock causes foreign output to rise, the domestic economy faces a positive export demand shock. Increased domestic exports again exert appreciation pressures on the domestic currency. In the case of a fixed exchange rate regime, the central bank has to intervene on the foreign exchange market to prevent the currency from appreciating. Under a floating exchange rate regime, the currency appreciates, offsetting the positive effect on exports and keeping output unchanged. Therefore, in the Mundell-Fleming model, spillovers of an expansionary monetary policy shock that are transmitted via foreign exports only affects output only in fixed exchange rate regimes, while in floating regimes the currency appreciation offsets the positive export demand shock.

The predictions of the Mundell-Fleming model have fed into the framework of the impossible trinity (or 'trilemma'). The impossible trinity states that – in the absence of capital controls – countries can counteract unwanted effects stemming from foreign monetary policy shocks with their own monetary policy measures only if they let their exchange rates float, as the policy choice of fixed exchange rates cannot be pursued together with the policy choice of independent monetary policy (see e.g. Obstfeld et al., 2004, for an empirical validation of the policy choices stipulated by the impossible trinity). The paradigm of the impossible trinity has however been disputed by Rey (2013), who argues that under free capital mobility, the global financial cycle constrains national monetary policy options regardless of the exchange rate regime, transforming the 'trilemma' into a 'dilemma' (see e.g. Passari and Rey, 2015, who show empirically that the global financial cycle affects economies with different exchange rate regimes in a similar way).

The analysis in the standard Mundell-Fleming framework has some shortcomings, especially with respect to the assumption of constant prices.[5] This is overcome by the extension of the Mundell-Fleming framework into the AD-AS model, which links the demand side of the economy (aggregate demand – AD) to aggregate supply (AS) and includes flexible prices. In the AD-AS model, assuming that the economy is operating at its potential, the expansion in domestic output due to a foreign monetary policy shock exerts price pressures, causing output to return to its initial level over time. Thus, while a temporary foreign monetary policy shock might lead to an increase of domestic output in the short run (depending on the reaction of the exchange rate as discussed above), in the long-run equilibrium domestic output is determined only by domestic supply.

Further limitations of the Mundell-Fleming model are overcome by the framework of new open economy macroeconomics (NOEM) models, for example the standard (i.e. operating under the assumption of perfect international capital mobility) version of the Corsetti-Pesenti (Corsetti and Pesenti, 2001) model. For an economy operating under a flexible exchange rate regime, an expansionary foreign monetary policy shock leads to an improvement in its terms of trade, which increases consumption in the short-run while the temporary effect on domestic production is ambiguous and depends on the intertemporal elasticity of substitution. In the long-run, all prices adjust to the increase in money supply and monetary neutrality, causing the real variables to return to their previous equilibrium levels. On the contrary, for a country operating under a fixed exchange rate regime, the domestic central bank must instantly react to the change in foreign money supply to keep the exchange rate fixed. Therefore, neither the relative money supply nor the terms of trade change for the domestic economy, and the foreign shock has no effect on consumption and production.

In terms of price effects of the foreign monetary policy shock, in the Corsetti-Pesenti model with flexible exchange rates, the domestic exchange rate appreciates in the short-run, dampening import prices if prices are set in the producer's currency.[6] In the long-run, the price level adjusts to the permanent increase in money supply and the terms of trade return to equilibrium, implicating either a drop in domestic prices or an increase in foreign prices to counterbalance the exchange rate appreciation. Therefore, the long-run effect on the domestic price level is ambiguous. While in fixed exchange rate regimes relative money supply remains unchanged, the nominal money increase in the domestic economy causes the price level to rise in the long-run.

[5]Additional shortcomings are that it provides only a static analysis, assumes that the economy operates below potential output, lacks a budget constraint and does not include welfare effects.

[6]It should be noted that a recent literature strand has disputed the traditional assumption of producer-currency pricing found in open-economy models, see e.g. Gopinath (2015), with far-reaching implications for the models' predictions.

2.2.2 Empirical evidence

A vast amount of literature on cross-border monetary policy spillovers has emerged in the past decades, which in the beginning focused mostly on spillovers between advanced economies. Canova (2005) was among the first ones to investigate monetary policy spillovers from an advanced economy (the US) to emerging economies (eight countries in Latin America), followed by other papers modelling spillovers from US monetary policy to Latin America, Canada and Asian economies (see e.g. Maćkowiak, 2007), and from euro area monetary policy to other European countries (both emerging and advanced, see e.g. Maćkowiak, 2006; Jarociński, 2010; Benkovskis et al., 2011). On SEE countries, the literature on conventional monetary policy spillovers is less abundant, which is mainly related to the short time series available. Nevertheless, available results are very heterogeneous and seem to depend on the model and specifications used (see Jiménez-Rodriguez et al., 2010; Minea and Rault, 2011; Feldkircher, 2015; Hájek and Horváth, 2016; Petrevski et al., 2015; Potjagailo, 2017, using near-VAR, VAR, GVAR, SVAR and FAVAR models). Moreover, to the best of my knowledge, four SEE countries have not yet been covered in the spillover literature at all.

The introduction of non-standard measures in October 2008 and the subsequent expansion of several different unconventional instruments brought a new angle into the academic and policy discussion of euro area monetary policy spillovers to countries outside the euro area.[7] However, given that the global experience with non-standard monetary policy is restricted (with a few exceptions) to the aftermath of the global financial crisis, the literature on spillovers from advanced economies' non-standard or unconventional monetary policy measures to emerging markets is relatively scarce. It can be divided into two categories: One strand investigates the impact of monetary policy announcements (and in some cases also implementations) on high-frequency financial indicators (e.g. sovereign bond yields, stock market indices, CDS spreads or exchange rates); see for example Fratzscher et al. (2018) for spillovers of US, and Georgiadis and Gräb Georgiadis and Gräb (2015) as well as Falagiarda et al. (2015) for spillovers of euro area non-standard monetary policy. The latter examine the effects of more than seventy announcement-related events on financial assets of four non euro area EU countries. For Romania, which is the only SEE country covered, they find a significant effect on the short-term money market rate and an especially pronounced effect on long-term government bond yields, while the exchange rate seems not to respond immediately to an ECB announcement. Ciarlone and Colabella (2016) test the effect of the ECB's asset purchase programmes on financial conditions of a panel of eleven countries in central, eastern and south-eastern Europe, including all countries covered in this paper. They find significant short-term spillover effects on financial markets as well as long-term spillovers on portfolio and cross-border banking flows, pointing to the existence of a portfolio rebalancing and a banking liquidity channel.

[7] For the purpose of this paper only spillovers outside the euro area are discussed. For an assessment of the effects *within* the euro area see e.g. Peersman (2011), Boeckx et al. (2017) and Burriel and Galesi (2018).

Focusing on longer-lasting macroeconomic effects instead, the second strand of literature has been following the methods of the literature on 'conventional' monetary policy spillovers by using some kind of VAR model to assess spillovers on macroeconomic variables. The literature on macroeconomic effects of non-standard monetary policy spillovers from the euro area to central and eastern Europe (CEE) is scarce, whereas it is non-existent for most countries in SEE. Babecká Kucharčuková et al. (2016) investigate spillovers on six EU non-euro area countries, among them three in CEE (Czech Republic, Hungary, Poland). They conclude that the spillovers of unconventional shocks are transmitted differently compared to conventional shocks, and while exchange rates respond quickly, the effect on inflation is ambiguous. Bluwstein and Canova (2016) use a Bayesian mixed-frequency VAR to incorporate both high-frequency financial as well as low-frequency macroeconomic data. They find that output effects of unconventional monetary policy measures were insignificant for CEE (Czech Republic, Hungary, Poland), and slightly negative for SEE (Bulgaria, Romania), and that the impact on inflation was slightly positive for both groups. With regard to the exchange rate channel, they conclude that it does not seem to shape the response of macroeconomic variables in the case of unconventional monetary policy shocks, as opposed to the case for conventional monetary policy. Halova and Horváth (2015) employ a PVAR model for eleven CEE and SEE countries (among them Bulgaria, Croatia and Romania). On the contrary to Bluwstein and Canova (2016) and Babecká Kucharčuková et al. (2016), their results suggest sizable spillovers and that a significant amount of output fluctuations in the CEE and SEE countries can be explained by the euro area's non-standard monetary policy measures.

Another open issue that has been discussed in the spillover-literature is what shapes the response of an economy to a foreign monetary policy shock (both conventional and non-standard). The role of the exchange rate regime has featured prominently in the spillover discussion, following the argument that flexible exchange rates are better suited to buffer real external shocks (based on Meade, 1951; Friedman, 1953). Moreover, with the policy options stipulated by the impossible trinity (see section 2.2.1), only flexible exchange rates allow for the possibility to react to foreign monetary policy shocks by means of domestic monetary policy. However, in a world where the 'trilemma' is replaced by a 'dilemma' (Rey, 2013), the exchange rate regime plays less of a role in shaping the size and magnitude of spillovers stemming from foreign monetary policy shocks, as leaning against the wind is not possible even for countries that pursue independent monetary policy. Besides exchange rate regimes, other potential determinants of spillovers identified by the literature are the degrees of trade and financial openness (see e.g. Miniane and Rogers, 2007). Georgiadis (2016) systematically examines U.S. monetary policy spillovers and finds that the role of the exchange rate regime is non-linear, and that non-advanced economies operating under an inflexible exchange rate regime experience larger spillovers the more strongly they are integrated in global trade. Furthermore, the results suggest that trade integration amplifies spillovers to non-advanced economies if the share of manufactured goods in aggregate output is large and the country participates in global value chains. Crespo-Cuaresma et al.

(2016) additionally find that macroeconomic vulnerabilities such as a high external imbalances tend to amplify spillovers from US monetary policy shocks. For spillovers of euro area monetary policy, Potjagailo (2017) presents some evidence that spillovers on other EU countries' output are larger if the exchange rate regime is fixed.

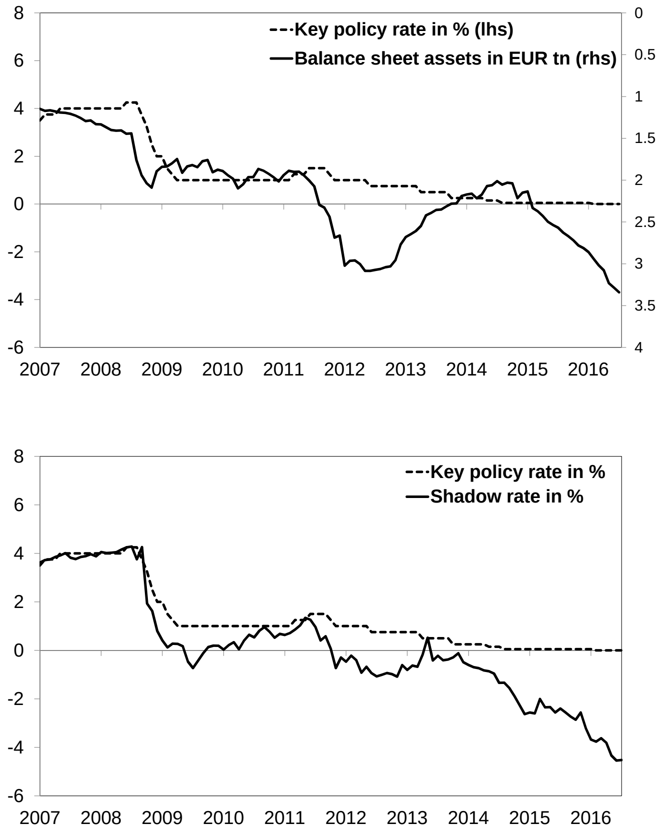
2.3 Methodology

The methodology used in this paper follows the strand of literature that investigates the effects of non-standard monetary policy measures on the real economy by employing some specification of a VAR model. There are two reasons to choose this approach: First, although event studies could identify significant financial market spillovers for some European countries outside the euro area (e.g. Ciarlone and Colabella, 2016), this does not necessarily imply that the real economy is equally affected, since financial variables often exhibit overshooting behaviour that does not necessarily transmit into output and inflation. Second, the event study approach requires developed financial markets to investigate the behaviour of high-frequency indicators. This is a major drawback in the case of emerging markets in general and SEE in particular, as these countries have very shallow financial markets in line with their small economic size and comparatively low GDP per capita levels.

2.3.1 Issues in dealing with non-standard monetary policy

Empirically assessing non-standard monetary policy measures brings a number of additional challenges compared to conventional monetary policy. First, as the key monetary policy rate does not incorporate non-standard measures such as longer-term refinancing operations or purchases of private and public sector securities, alternative indicators for the stance of non-standard monetary policy have to be found. Those used in the literature so far have been the term spread between government bonds of different maturities (e.g. Chen et al., 2012), central bank balance sheet assets (e.g. Gambacorta et al., 2014), or shadow rates that are constructed in order to be directly comparable to key policy rates (Lombardi and Zhu, 2018; Krippner, 2015; Wu and Xia, 2016). In this paper I use Eurosystem balance sheet assets as the main measure of non-standard monetary policy. In the upper chart of figure 2.1, inverted Eurosystem balance sheet assets are plotted together with the key policy rate. It can be seen that the key policy rate was decreased in various steps to 1 percent in May 2009, from where it slowly and gradually moved towards the zero lower bound, which was reached in March 2016. In contrast, Eurosystem balance sheet assets started to increase already with the switch of liquidity operations to a fixed rate tender with full allotment in October 2008, and thereafter fluctuated with the introduction and phase-out of the respective non-standard programmes. Besides Eurosystem balance sheet assets, I additionally use the shadow rate developed by Wu and Xia (2016) for robustness testing. It is calculated by assessing bond prices in a framework of a multifactor term structure model; and is directly comparable to the key policy rate as both interest rates are equal in conventional times (see the lower chart of figure 2.1). In contrast to balance sheet assets, this indicator also includes announcement effects of non-standard monetary policy measures whenever they affect bond yields.

Figure 2.1: Indicators of non-standard monetary policy measures



Sources: ECB, Wu and Xia (2016)

Second, the way some of the ECB's non-standard monetary policy measures have been designed makes it necessary to find an empirical strategy that disentangles exogenous monetary policy shocks from endogenous or demand-driven monetary expansions. Since the change in the operational framework from standard tender-based allotment to fixed-rate full allotment in October 2008, monetary policy operations have been essentially endogenous or demand-driven, as banks have unlimited access to liquidity at the interest rate on the main refinancing operations (MRO) under the condition that they can provide enough collateral (Boeckx et al., 2017). Moreover, the (targeted) longer-term refinancing operations, which were increased in duration and size in October 2008, are also endogenous to a certain extent since the ECB only fixes the upper ceiling of these operations, whereas banks decide how much to draw upon that limit. This paper deals with endogeneity issues in several ways. First, I follow the approach proposed by Boeckx et al. (2017), who complement the Eurosystem's balance sheet assets as the main measure for non-standard monetary policy measures with certain assumptions on shock identification (see subsection 2.3.3). Moreover, I perform robustness checks by using only the position 'Securities held for monetary policy purposes' (A070100) of the Eurosystem's balance sheet, which incorporates all securities purchased under the various purchasing programmes. Compared to other positions of the Eurosystem balance sheet, this is the most exogenous part since the size and frequency of bond purchases are ex-ante determined by the ECB and not shaped by commercial banks' behaviour.

2.3.2 Model

To model spillovers from the euro area's non-standard monetary policy to SEE countries I employ structural Bayesian (BVAR) models for each SEE country. Compared to frequentist (i.e. non-Bayesian) models, the principle of Bayesian inference is that the estimation outcome (the 'posterior') does not only depend on the information contained in the data, but also on prior information on the distribution of the parameters (the 'prior'). The ability of including prior information in the estimation bears several advantages, of which two are especially useful for the purpose of this paper: First, the additional relevant information on the distribution of coefficients that is included in the prior specification is especially suited for shorter data samples.[8] Second, to account for autoregressive behaviour of a variable, a prior can be chosen that postulates a unit root on the respective variable. Therefore, in Bayesian estimation non-stationary variables do not need to be transformed into first differences, which brings the advantage of avoiding a transformation bias that occurs when estimation is carried out with data transformed into first differences.

The following structural BVAR model is employed for each SEE country at monthly frequency:

$$\sum_{s=0}^{p} \begin{bmatrix} A_{11}(s) & A_{12}(s) \\ A_{21}(s) & A_{22}(s) \end{bmatrix} \begin{bmatrix} y_1(t-s) \\ y_2(t-s) \end{bmatrix} + \begin{bmatrix} c_{11} \\ c_{21} \end{bmatrix} = \begin{bmatrix} \varepsilon_1(t) \\ \varepsilon_2(t) \end{bmatrix} \tag{2.1}$$

where $y_1(t)$ represents a vector of macroeconomic variables of the SEE country, $y_2(t)$ a vector of macroeconomic variables of the euro area, and the vectors c_{11}, c_{21} are constants. The vectors $\varepsilon_1(t) \sim N(0, \Sigma_1)$ and $\varepsilon_2(t) \sim N(0, \Sigma_2)$ denote structural shocks of domestic and euro-area origin, respectively.

For each s, $A_{21}(s) = 0$, implying that the variables of the SEE country are set to be exogenous to the variables of the euro area under the assumption that neither current nor past economic

[8]In turn, for larger samples the results obtained through Bayesian estimation become more similar to the results produced by frequentist methods.

developments in the SEE countries influence developments in the euro area. This so-called block exogeneity feature introduced by Cushman and Zha (1997) has been used frequently in the literature (see e.g. Canova, 2005; Maćkowiak, 2007; Benkovskis et al., 2011) and is well suited for modelling spillovers from large to small economies, as it helps to identify spillovers from the viewpoint of the small open economy and reduces the number of parameters to be estimated (Cushman and Zha, 1997).

The vector y_1 consists of the following variables:

$$y_1 = \begin{pmatrix} y_t^{SEE} & p_t^{SEE} \end{pmatrix}' \tag{2.2}$$

where y_t^{SEE} denotes real GDP in local currency and p_t^{SEE} denotes prices of the respective SEE country. At a second stage, in order to investigate potential transmission channels, the vector y_1 includes either exports[9], x_t^{SEE}, the interbank market rate of the respective SEE country, i_t^{SEE}, or the exchange rate of the local currency vis-à-vis the euro, e_t, for countries that are operating under a flexible exchange rate regime.

The vector y_2 represents the euro area and includes six variables:

$$y_2 = \begin{pmatrix} y_t^{EA} & p_t^{EA} & assets_t^{EA} & CISS_t & spread_t^{EA} & MRO_t \end{pmatrix}' \tag{2.3}$$

where y_t^{EA} and p_t^{EA} again denote output and prices, respectively, but this time for the euro area and $assets_t^{EA}$ represents Eurosystem balance sheet assets as the main measure for non-standard monetary policy (as discussed in subsection 2.3.1). Moreover, following Gambacorta et al. (2014), Boeckx et al. (2017) and Burriel and Galesi (2018), I include the CISS indicator (composite indicator of systemic stress) developed by Holló et al. (2012) ($CISS_t$), which serves two purposes: First, it controls for the impact of euro-area financial stress and economic risk, which is important to capture in the model as it has had pronounced effects on euro-area macroeconomic developments. Second, the inclusion of the CISS indicator helps to disentangle exogenous balance sheet movements from endogenous ones, and thus enables a proper identification of monetary policy shocks (see subsection 2.3.3). For the same purpose, I also include the spread between EONIA and the MRO rate (denoted $spread_t^{EA}$). Moreover, in order to disentangle conventional from non-standard monetary policy shocks, the model incorporates the MRO rate (MRO_t).

For the prior specification I use an independent normal-Wishart prior and obtain the scale matrix S_0 from individual AR regressions. The autoregressive coefficient of the prior is set to 1, since the variables enter the model in levels as discussed above. The remaining hyperparameters that specify the prior are chosen following Dieppe et al. (2016), which are standard values found in the literature.[10]

[9] In order to account for indirect spillovers, exports to the world (instead of only to the euro area) are used.

[10] Those remaining hyperparameters specify the overall tightness of the prior, i.e. how large the prior variance

The posterior is derived by Gibbs sampling with a total number of 5,000 iterations and a burn-in sample of 1,000 iterations. The Bayesian information criterion (BIC) suggests a lag length of 1; however, testing for autoregressive behaviour of the residuals suggests that a model specification of 4 lags is best to avoid residual autocorrelation. Therefore I define $p = 4$. Estimations are carried out by employing the BEAR (Bayesian Estimation, Analysis and Regression) toolbox developed by Dieppe et al. (2016).

2.3.3 Identification

In order to generate impulse response functions, the shocks are identified via sign and zero restrictions, following the method proposed by Arias et al. (2014) (see Dieppe et al., 2016). The non-standard monetary policy shock is the only identified shock in the model and specified according to table 2.1. The first six variables define the non-standard monetary policy shock and its effects on the euro area, while the remaining variables apply to the respective SEE country's output and price level. An expansionary non-standard monetary policy shock increases the Eurosystem balance sheet assets on impact and in the first month following the shock, while both the CISS-indicator as well as the spread between the EONIA and the MRO decrease immediately (on impact) and in the first month after the shock. These identifying assumptions are taken to distinguish demand-driven from exogenous balance sheet shocks, following Boeckx et al. (2017) and Burriel and Galesi (2018). More specifically, in periods of financial stress or other shocks, increased demand for liquidity expands the balance sheet, implying that the CISS indicator as well as EONIA increase (see Boeckx et al., 2017). Vice versa, a balance sheet expansion that is caused by an ECB monetary policy measure should not increase but *decrease* both financial stress and the demand for liquidity, which is reflected in the sign restrictions for the shock identification. Finally, the zero restriction of the MRO rate ensures that the balance sheet increase is orthogonal to a conventional monetary policy shock. For the response of output and prices, I follow the standard approach of defining conventional monetary policy shocks by imposing zero restrictions to disentangle it from other shocks. Similarly, zero restrictions are placed on output and price responses of the SEE country, in order to disentangle the potential spillover from domestic real economy disturbances.[11] The acceptance rates (i.e. the percentage of draws from the Gibbs sampling algorithm satisfying the sign and zero restrictions) of the baseline models are depicted in table 2.2, suggesting that the chosen shock identification is reasonable.

is and thus how much the prior is forced upon the posterior, the speed at which coefficients for lags greater than 1 converge to 0, and the coefficients relating variables to past values of other variables. For further info see Dieppe et al. (2016).

[11]For the same reason, a zero restriction is put on exports in the subsequent estimations on potential transmission channels.

Table 2.1: Sign and zero restrictions for the shock identification of the baseline model

$assets_t^{EA}$	$CISS_t$	$spread_t^{EA}$	MRO_t	y_t^{EA}	p_t^{EA}	y_t^{SEE}	p_t^{SEE}
+	−	−	0	0	0	0	0
0-1	0-1	0-1					

Note: 0 indicates that the immediate response is restricted, while + (−) indicates that only a positive (negative) reaction is permitted in the respective period.

Table 2.2: Acceptance rates of structural matrices (in %)

Albania	13.93
Bosnia and Herzegovina	12.21
Bulgaria	13.71
Croatia	11.95
Montenegro	12.24
North Macedonia	13.40
Romania	11.96
Serbia	13.12

2.3.4 Data

The time span of the baseline model covers the period between January 2008 and December 2017 and all variables enter the model in levels. As a measure of output I use real GDP in local currency, interpolated by the Chow-Lin method with industrial production to obtain data with monthly frequency.[12] The price level enters the model as the consumer price (harmonised consumer price inflation in the case of EU countries) index. As indicator for a potential financial channel I use short-term interest rates, since time series on asset prices or longer-term interest rates are not available for all countries. More specifically, I include monthly values of three-month interbank market rates, with the exception of Bosnia and Herzegovina as well as Montenegro, which do not publish interbank market rates. In the case of Montenegro, an unweighted average of three-month and six-month government T-bill rates is used as a proxy for interbank market rates. For Bosnia and Herzegovina, no such short-term interest rate exists. Therefore, following the approach of Cerutti et al. (2010), I create a composite series that consists to two-thirds of retail deposit rates and to one-third of retail lending rates, both in the corporate sector. All exchange rates are expressed in average local currency vis-à-vis the euro, so an increase in the exchange rate depicts depreciation and vice versa. Data on exports, which is derived from the IMF's Direction Of Trade Statistics, is limited to merchandise exports, which means that service exports are not captured in the model. All variables are seasonally adjusted by the U.S. Census

[12] With the exception of Montenegro, where GDP is not available for the whole time span and therefore industrial production is used as a measure for output.

Bureau's X-13 seasonal adjustment procedure and are transformed into their natural logs (with the exception of financial variables). The data sources are national central banks, national statistical offices, Eurostat, the ECB, the IMF's Direction Of Trade Statistics and Bloomberg.

2.4 Results

2.4.1 The effect of a non-standard monetary policy shock on the euro area economy

I start by looking at the transmission of an expansionary Eurosystem balance sheet shock within the euro area. The impulse response functions of the euro area are displayed in figure 2.2, where the continuous line depicts the median posterior response and the shaded area represents 68 percent of the credibility interval. Because the variables enter the model in natural logs, the y-axis reports percentage changes, except for financial variables which are depicted in percentage point changes. It can be observed that the one standard deviation (1.1 percent increase) balance sheet shock is very persistent, as it remains almost unchanged for two years, driven by the sample period of continuous expansionary monetary policy. Both the accompanying decline of the CISS indicator and the decrease of the EONIA-MRO spread are less persistent, but do not fade out completely until the end of the two-year horizon. The response of the MRO (key policy rate) to the non-standard monetary policy shock is ambiguous. Turning to the macroeconomic effects, the impulse responses suggest that output rises gradually with a peak increase of 0.03 percent after eleven months. The price level increase reaches 0.04 percent after two years and its response seems to be relatively persistent.

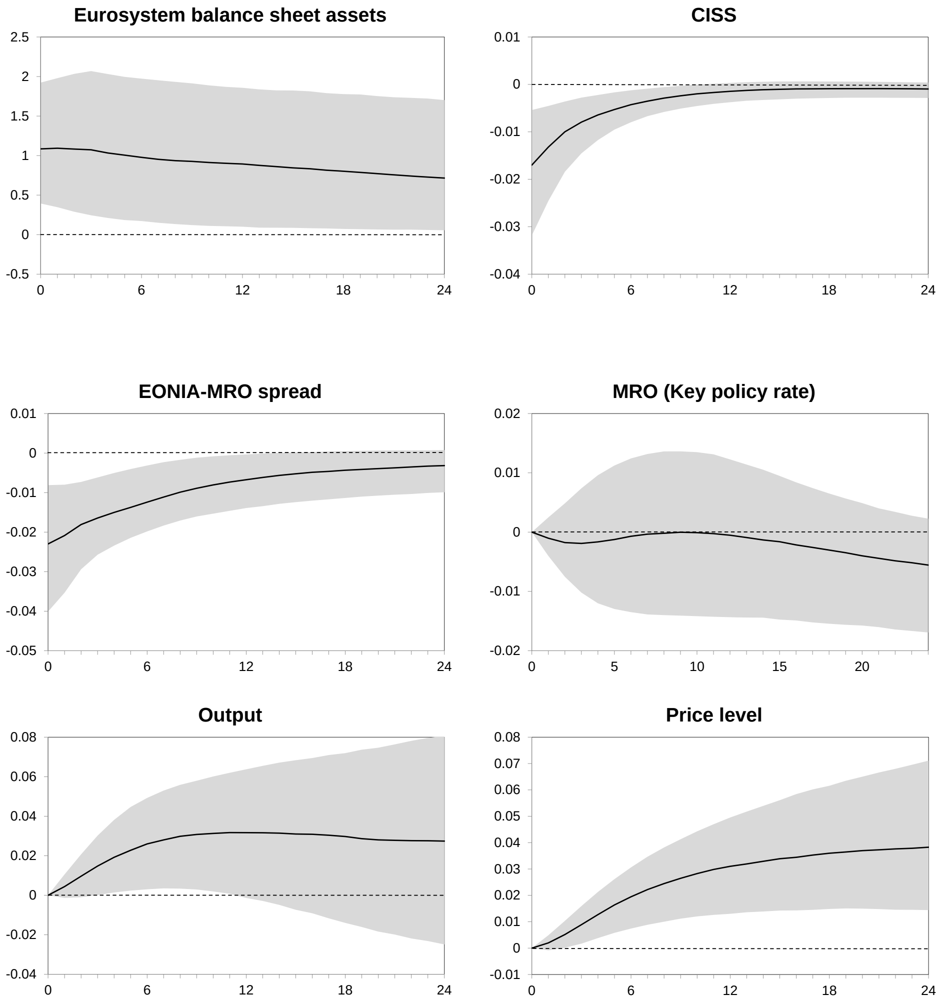
2.4.2 Spillovers to SEE countries and their transmission channels

Turning to the SEE countries, the results suggest that spillovers from a euro area expansionary non-standard monetary policy shock are positive (see figure 2.3, upper-left panel). In the three countries that exhibit the highest magnitude in their shock response (Montenegro, Albania and Serbia), it is credible within the 68 percent interval. Moreover, in half of the countries the response of output seems to be stronger than in the euro area.

Price level responses are positive in all countries and lie (with one exception) within the 68 percent credibility interval (see figure 2.3, upper-right panel). It can be observed that the peak response is by far the strongest for Serbia (at 0.13 percent), followed by Romania and Bulgaria. The relatively strong price responses are in line with the high share of imports from the euro area that range from around one third to over 50 percent of all imports in SEE countries. In most countries the price response is even more pronounced than in the euro area.

In order to shed light on potential transmission channels, I estimate spillovers on exports and interbank interest rates in separate models, where the vector $y_2(t)$ for the identification of the

Figure 2.2: Euro area: Response to a balance sheet shock

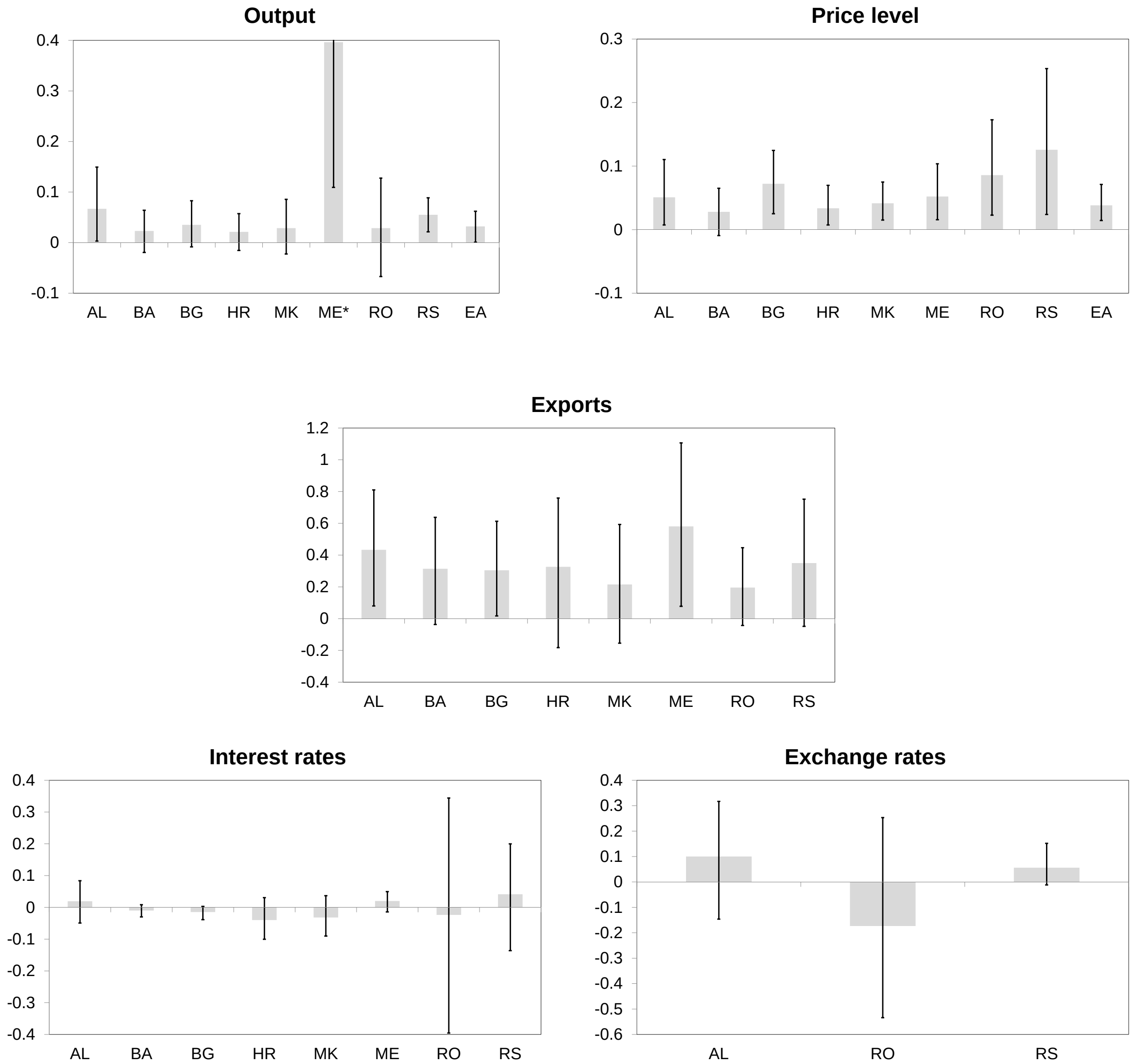


Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for financial variables, where percentage point changes are depicted.

euro area shock remains unchanged, but the vector $y_1(t)$ contains either exports or short-term interest rates. For exports, the peak response is depicted in the centre-left panel of figure 2.3. The impact of a non-standard monetary policy shock on SEE exports is positive in all countries, suggesting that exports are indeed an important channel of shock transmission, and it is credible within a 68 percent interval in three countries. The largest magnitude of the shock response can be observed for Montenegro, Albania and Serbia, which corresponds to the relative magnitude of their output reaction and thus suggests that exports are indeed a relevant transmission channel.

Turning to short-term interest rates, the peak responses across countries are heterogeneous in sign and magnitude, and surrounded by large uncertainty bands. One reason for the weak model output might be the relatively illiquid interbank money markets in SEE countries. Thus this result should not necessarily be taken as proof that financial channels do not transmit

Figure 2.3: Comparison of responses to a non-standard monetary policy shock



Note: Peak response to expansionary balance sheet shock within the first 24 months in percent. AL refers to Albania, BA to Bosnia and Herzegovina, BG to Bulgaria, HR to Croatia, MK to North Macedonia, ME to Montenegro, RO to Romania, RS to Serbia and EA to the euro area. *For Montenegro, output refers to industrial production instead of GDP. The results are derived from the baseline model (output and price level) as well as separate models including exports, interest rates and exchange rates, respectively. They depict growth rates in percent for all variables except for interest rates, where percentage point changes are depicted.

non-standard monetary policy shocks from the euro area to SEE countries, as changes in the interbank market rate do not capture foreign direct investment or portfolio inflows.

Comparing the peak responses across countries, the results suggest that the inflation targeting countries (Albania, Romania and Serbia) that operate under a managed or flexible exchange regime were equally affected by spillovers as countries that have pegged their currency to the euro. This result is not surprising when the model output on exchange rate responses is taken into account (figure 2.3, lower-left panel), where the peak responses in Albania and Serbia suggest an exchange rate depreciation (rather than an appreciation that would be consistent with the response of exports), and in Romania the peak appreciation is surrounded by a large uncertainty band. This is in line with the very stable exchange rates observed for Albania and Romania in the sample period. Moreover, between January 2008 and December 2017 the central banks of Albania, Romania and Serbia eased their monetary policy stance by decreasing interest rates by a total of 500, 625 and 650 basis points, respectively, which might have counteracted appreciation pressures on the exchange rate. The finding that in SEE flexible exchange rates did not respond strongly to non-standard monetary policy shocks during the period under review is in line with Bluwstein and Canova (2016), who argue that the exchange rate channel is not important for unconventional monetary policy transmission (which is different from the conventional case).

2.4.3 The importance of euro area non-standard monetary policy shocks for SEE countries

Besides analysing the peak effect from non-standard monetary policy shocks on certain macroeconomic variables in SEE, it is also important to assess how much of the variance of output and prices can be explained by non-standard monetary policy shocks. Table 2.3 presents the percentage of the forecast error variance of output and price explained by a non-standard monetary policy shock. For the euro area itself, non-standard monetary policy shocks account for 1.24 percent of the fluctuations of GDP and 4.94 percent of the fluctuations of the price level after two years. In a similar vein, for most SEE countries the degree of variance explained by euro-area non-standard monetary policy shocks is higher for prices than for output. Comparing the forecast error variance of output among SEE countries, Albania, Serbia and Bulgaria are the countries most exposed to euro-area non-standard monetary policy shocks, and in half of the countries the movements in output are larger than the output variability in the euro area, which is in line with the analysis from impulse response functions. With regard to the price level, the highest variability among SEE countries can be observed in Romania, Bulgaria and Montenegro, which is again mostly in line with the previous analysis. However, the explanatory power of the non-standard monetary policy shock in price levels is lower in SEE countries than in the euro area.

2.4.4 Results for individual SEE countries

The response of the Albanian economy to a non-standard monetary policy shock is shown in figure 2.4. An expansionary euro area balance sheet shock raises output by 0.07 percent and prices by 0.05 percent after two years. Albania's exports rise as a response to the balance sheet shock, peaking at 0.43 percent after five months, suggesting that exports are an important transmission channel to explain the relatively strong spillover to the Albanian economy. The response of the interbank interest rate on the other hand seems to be muted. Also, the exchange

Table 2.3: Forecast error variance

	Output	**Price level**
Euro area	1.24	4.94
Albania	3.98	3.10
Bosnia and Herzegovina	0.67	1.21
Bulgaria	1.45	4.23
Croatia	0.73	1.66
Montenegro	1.29	3.78
North Macedonia	1.04	2.68
Romania	0.83	4.27
Serbia	2.45	3.04

Note: Percentage of the variance of the respective variables explained by a non-standard monetary policy shock after twenty-four months.

rate of the lek vis-à-vis the euro does not exhibit a distinct response to the shock, which is in line with the fact that it has fluctuated only slightly against the euro in the sample period.[13]

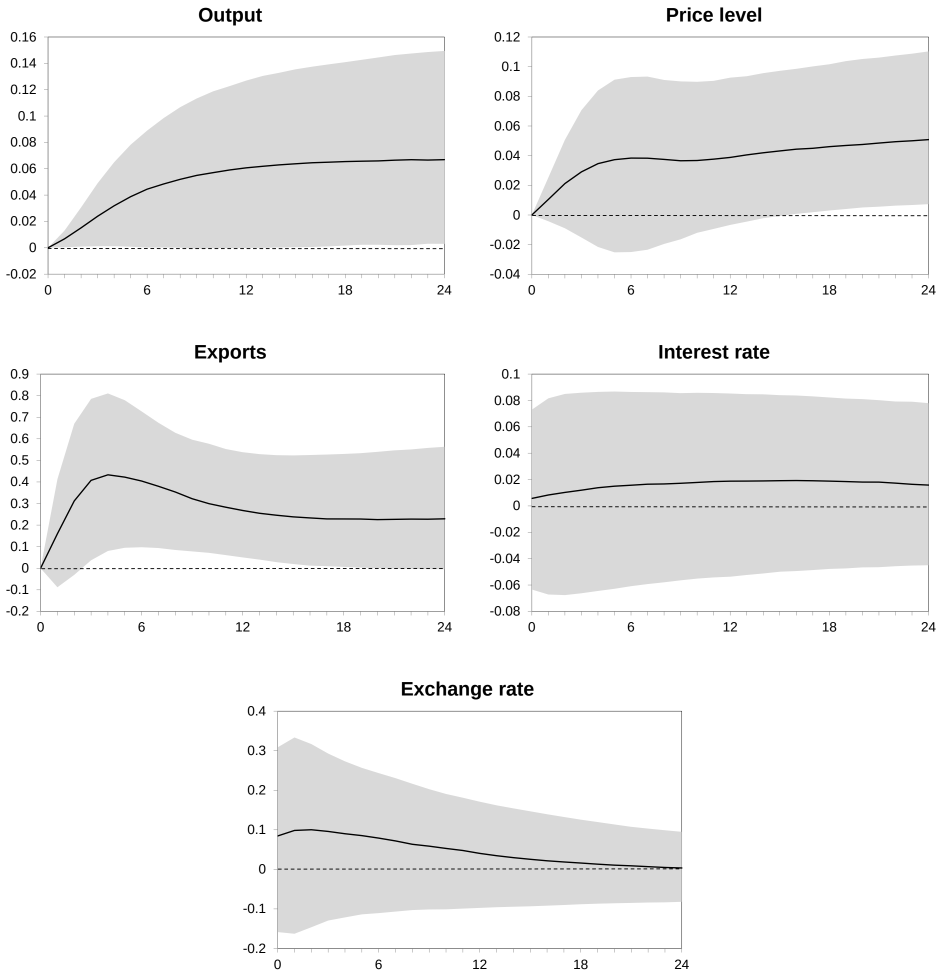
In the case of Bosnia and Herzegovina, the economic response to a non-standard euro area monetary policy shock is depicted in figure 2.5. Output shows initially a slightly negative response which turns positive after seven months, but the uncertainty band surrounding the response is relatively high. Notwithstanding the mixed output response, the price level increases gradually with a peak increase of 0.03 percent after twelve months. Exports react pronounced with a peak response of 0.32 percent after nine months. The uncertainty band of the interest rate response (which in Bosnia and Herzegovina's case is a composite retail rate) is relatively wide in the short-term, but becomes more significant in the medium-term with a decrease of 0.01 percentage points.

For Bulgaria (see figure 2.6), the output response is positive with an increase of 0.04 percent after two years. The price level exhibits a pronounced increase, reaching the peak of 0.07 percent after 19 months. The export channel seems to be relevant also in the case of Bulgaria, as exports rise with a peak of 0.30 percent after eight months. The interbank interest rate does not react in the short-run, but in the medium term is exhibits a decrease of 0.01 percentage points at the end of the two-year horizon. The marked reaction of output and prices in Bulgaria are in line with the results of Hájek and Horváth (2016) for spillovers of short-term interest rate shocks.

In the case of Croatia (see figure 2.7), the response of output is positive and peaks at 0.02 percent after twelve months, although it is surrounded by some uncertainty. The response of the price level is more significant and exhibits an increase of 0.03 percent at the end of the two-year horizon. Both the output as well as the price response are in line with what Hájek and Horváth

[13]Between 2008 and 2017 the average monthly fluctuation against the euro amounted to 0.5 percent.

Figure 2.4: Albania: Response to a balance sheet shock



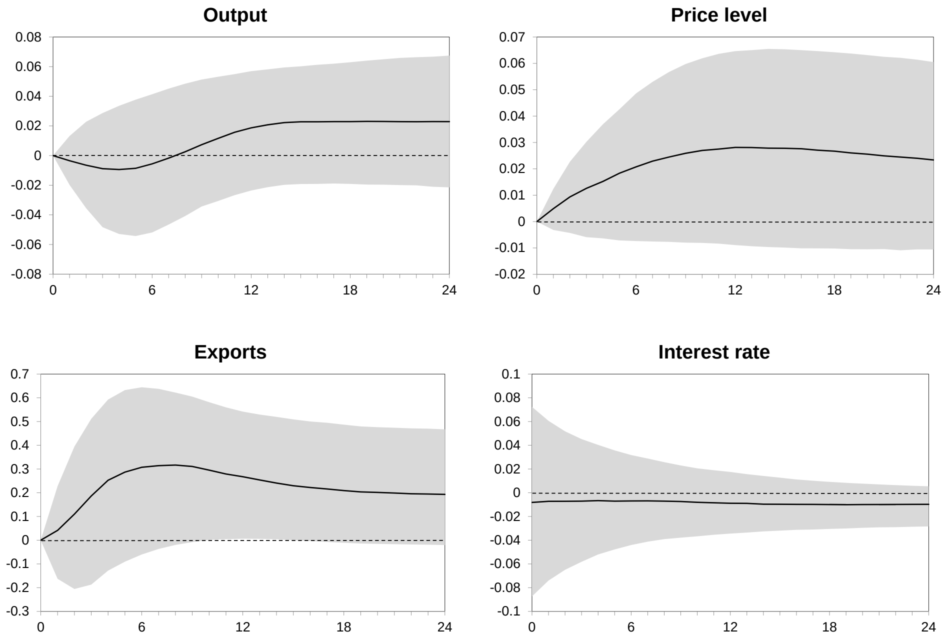
Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports, the interest rate and exchange rate, respectively.

(2016) find for policy spillovers of short-term interest rate shocks.[14] Furthermore, the export response peaks at 0.33 percent after six months, while the interbank market rate decreases with a trough of 0.04 percentage points after eight months. The results suggest that both exports as well as financial channels might be relevant in transmitting shocks from non-standard monetary policy measures in the euro area to Croatia.

For Montenegro (see figure 2.8), an exogenous expansion of the Eurosystem's balance sheet translates into a pronounced rise of industrial production by 0.40 percent after three months. The price level also increases by 0.05 percent at the end of the two-year. Montenegro's exports seem to rise by 0.58 percent after six months, suggesting that an increase in exports might explain

[14]Conversely, Petrevski et al. (2015) find that a positive interest rate shock *increases* Croatia's price level.

Figure 2.5: Bosnia and Herzegovina: Response to a balance sheet shock



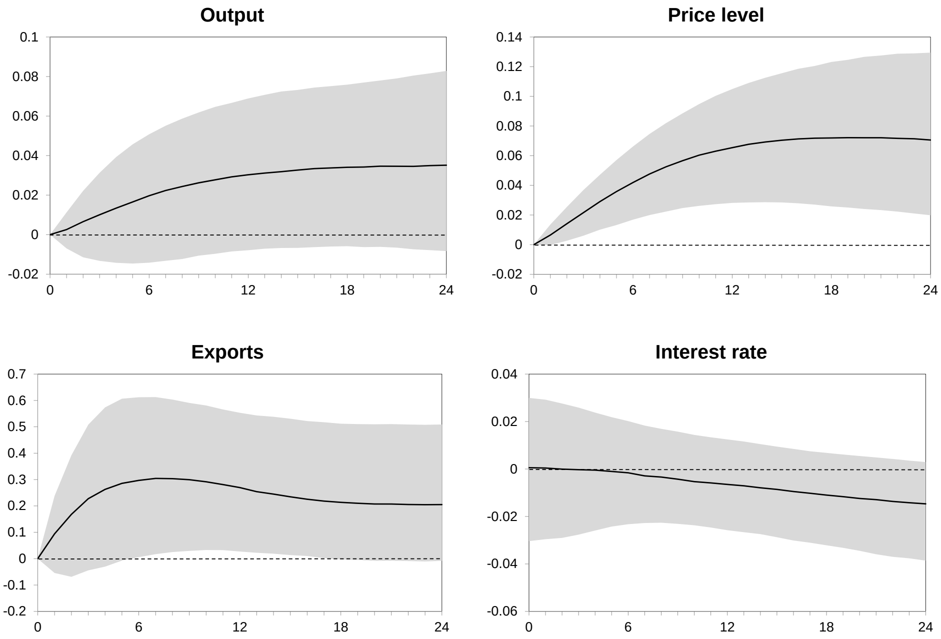
Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports and the interest rate, respectively.

the rise in industrial production and prices. The uncertainty of the interest rate response, which is in the case of Montenegro a composite of three- and six-month T-bill rates, is very high in the short-term, while in the medium-term the increase in interest rates becomes more significant.

The response of the economy of North Macedonia is depicted in figure 2.9. The results suggest that the euro-area non-standard monetary policy shock does not trigger a significant output response, while the price level response is positive with a peak of 0.04 percent after 23 months. Exports peak — although not strictly significantly — at 0.22 percent after twelve months. The interbank market rate decreases by a maximum of 0.03 percentage points after eight months with the response being relatively persistent and becoming more significant towards the end of the horizon, suggesting that financial spillovers could play a role at least in the medium term. The results for the interest rate response are in line with Petrevski et al. (2015) for a positive short-term interest rate shock; however, they find a different response of the price level.

Spillovers to Romanian output from the non-standard monetary policy shock are not very pronounced (compare figure 2.10). This result is different from the findings of Hájek and Horváth (2016) and Bluwstein and Canova (2016), who conclude that a contractionary euro-area short-term interest rate shock initially increases Romanian output (or that an expansionary non-standard monetary policy output decreases Romanian output, respectively). The response of

Figure 2.6: Bulgaria: Response to a balance sheet shock



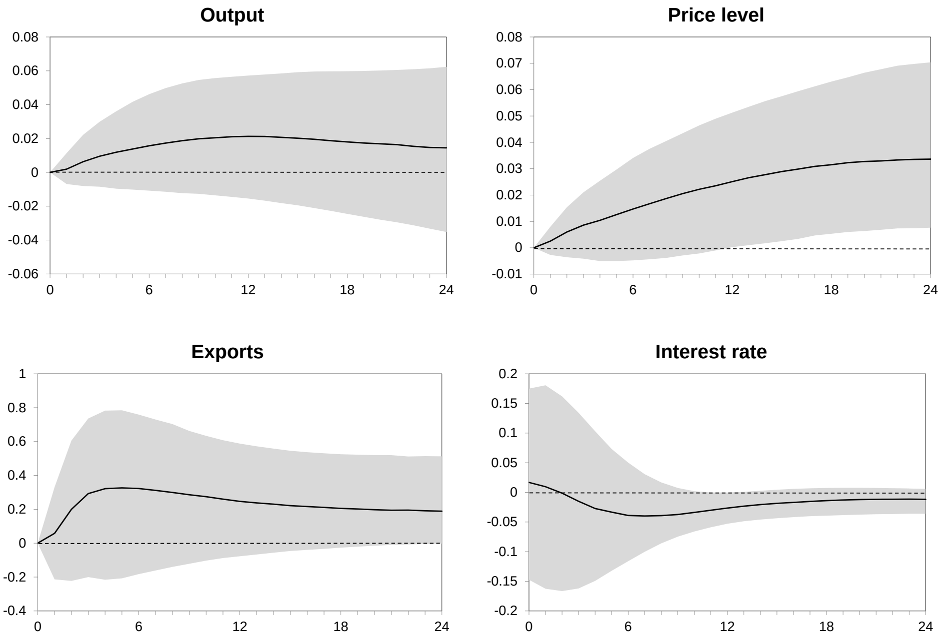
Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports and the interest rate, respectively.

prices, on the other hand, is positive and relatively strong with an increase of 0.09 percent after two years. Exports are peaking at 0.20 percent after a period of seven months. On the contrary, the response of the short term interest rate is muted, suggesting that the euro area shock does not affect the interbank market rate in Romania. Initially the exchange rate seems to appreciate, which is however subject to high uncertainty, but turns into a more pronounced and persistent depreciation after five months. The blurred response is in line with the relative exchange rate stability of the lei vis-à-vis the euro since mid-2012.[15] The muted exchange rate response confirms the outcome of the event study by Falagiarda et al. (2015), while they also find a pronounced reaction of the short-term money market rate which is different to the results obtained here.

Serbia's output (see figure 2.11) increases with a peak of 0.06 percent after eleven months, which is one of the strongest output responses compared with the other countries in the region. Also, the price response is very pronounced with a peak increase of 0.13 percent at the end of the two-year horizon, which inter alia can be explained by the strong contribution of euro-area import prices to inflation pressures in the past (see e.g. International Monetary Fund, 2011). Serbia's exports seem to react positively to the shock with a peak response of 0.35 percent after

[15]From mid-2012 to end-2017 the average monthly fluctuation against the euro amounted to 0.5 percent, as compared to 1.0 percent from 2008 up to mid-2012.

Figure 2.7: Croatia: Response to a balance sheet shock

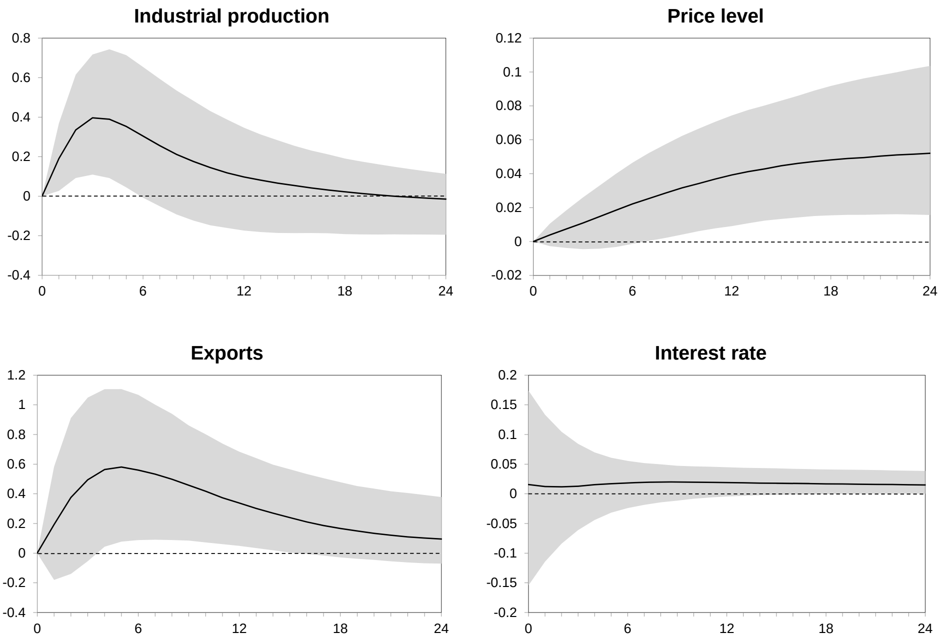


Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports and the interest rate, respectively.

seven months. On the contrary, neither the interbank interest rate nor the exchange rate seem to be affected significantly by the shock.[16]

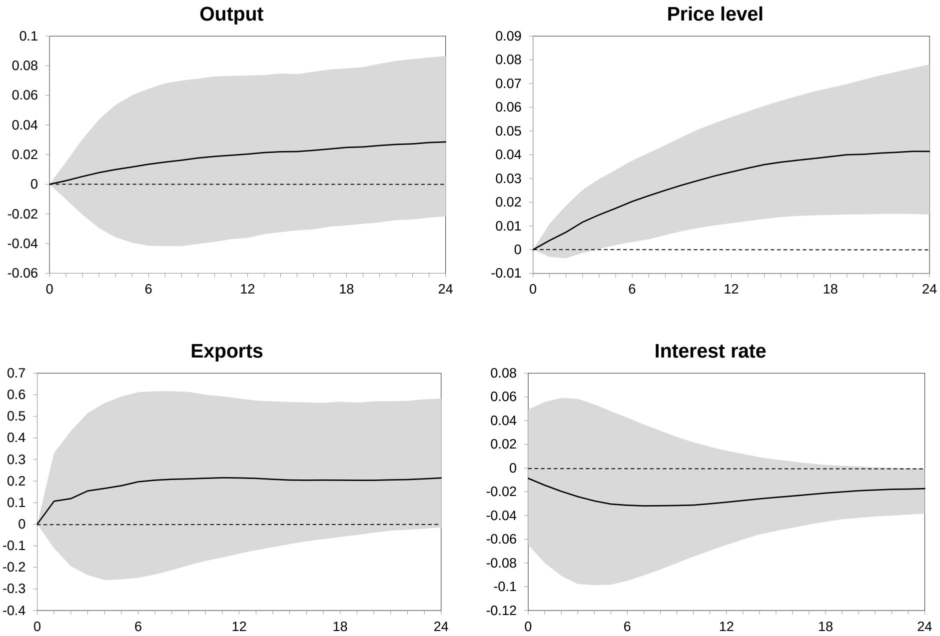
[16]This is despite the fact that, compared with Albania and Romania, Serbia's exchange rate fluctuated relatively strong vis-à-vis the euro in the sample period, with an average monthly fluctuation of 0.9 percent.

Figure 2.8: Montenegro: Response to a balance sheet shock



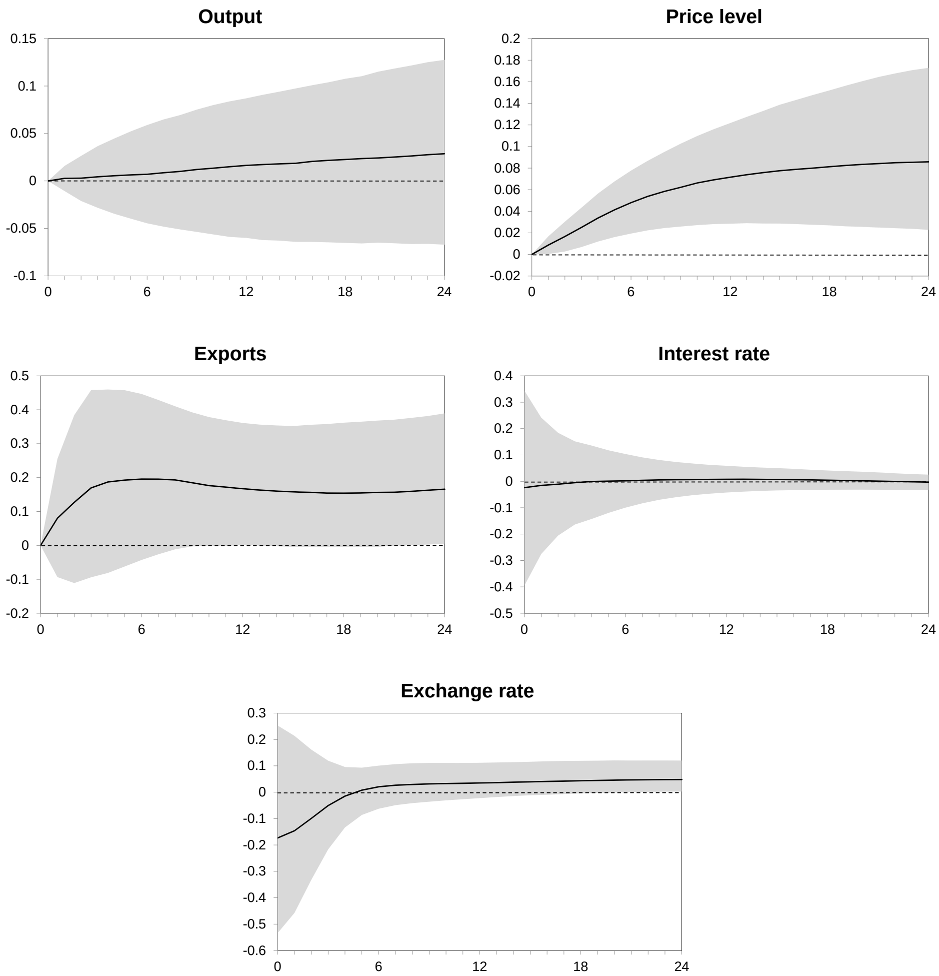
Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports and the interest rate, respectively.

Figure 2.9: North Macedonia: Response to a balance sheet shock



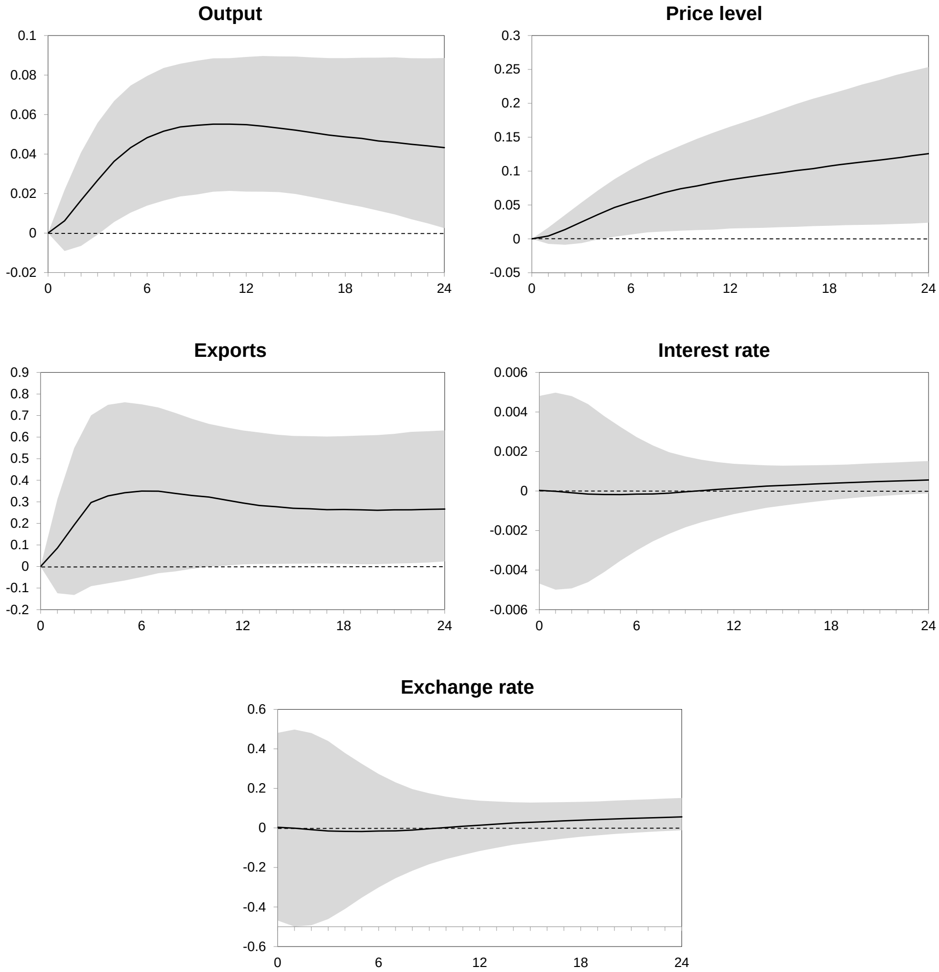
Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports and the interest rate, respectively.

Figure 2.10: Romania: Response to a balance sheet shock



Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports, the interest rate and exchange rate, respectively.

Figure 2.11: Serbia: Response to a balance sheet shock



Note: Response of variables to an expansionary one standard deviation Eurosystem balance sheet shock. The shaded regions report pointwise 68 percent credibility intervals. The x-axis reports months, the y-axis monthly growth rates in percent for all variables except for interest rates, which depict changes in percentage points. The results are derived from the baseline model (output and price level) as well as separate models including exports, the interest rate and exchange rate, respectively.

2.5 Robustness checks

To test whether the results hold for different model specifications or variable choices, I perform the following robustness checks: As discussed in subsection 2.3.1, besides balance sheet assets, shadow rates can also be used as an indicator for non-standard monetary policy measures. To see whether the results are robust to an expansionary shadow rate shock (where I use the shadow rate developed by Wu and Xia, 2016), I keep all other variables and the shock identification unchanged. The only difference is that an expansionary shock implies that the shadow rate decreases, which means that the sign restriction for the shadow rate is turned into negative (see table 2.4).

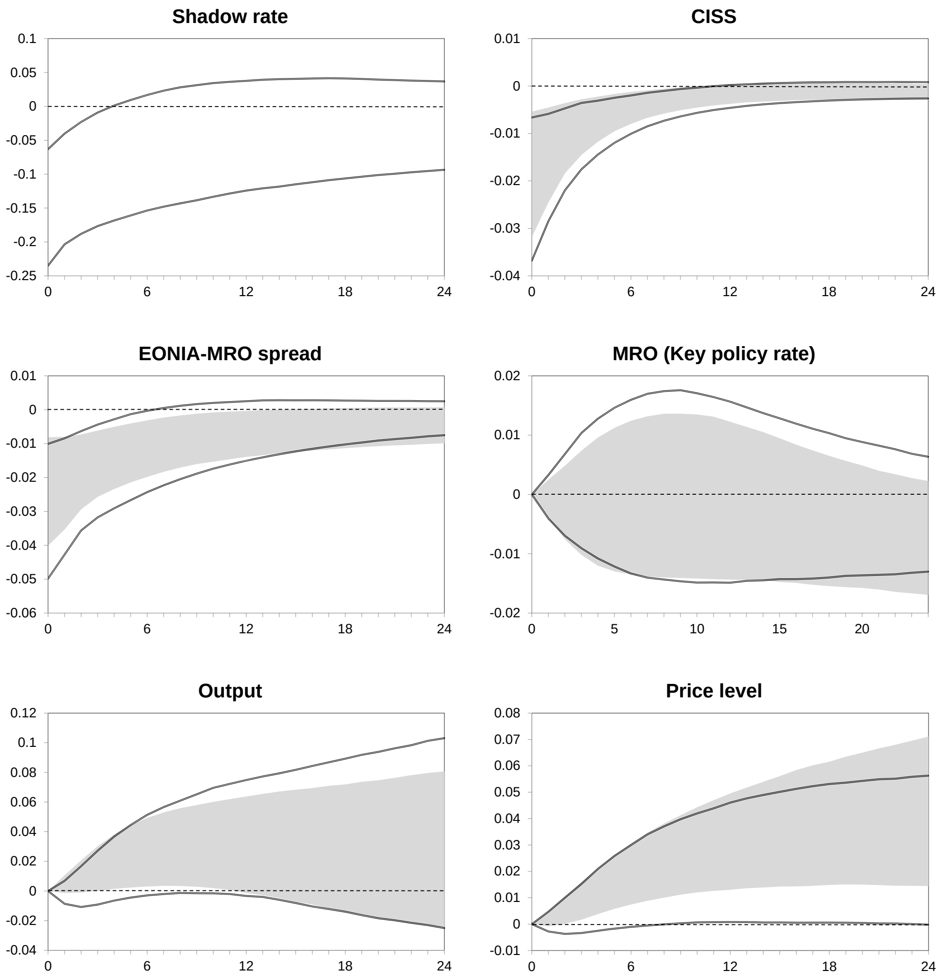
Table 2.4: Sign and zero restrictions for the shock identification of the shadow rate model

$shadow_t^{EA}$	$CISS_t$	$spread_t^{EA}$	MRO_t	y_t^{EA}	p_t^{EA}	y_t^{SEE}	p_t^{SEE}
$-$	$-$	$-$	0	0	0	0	0
0-1	0-1	0-1					

Note: 0 indicates that the immediate response is restricted, while + ($-$) indicates that only a positive (negative) reaction is permitted in the respective period.

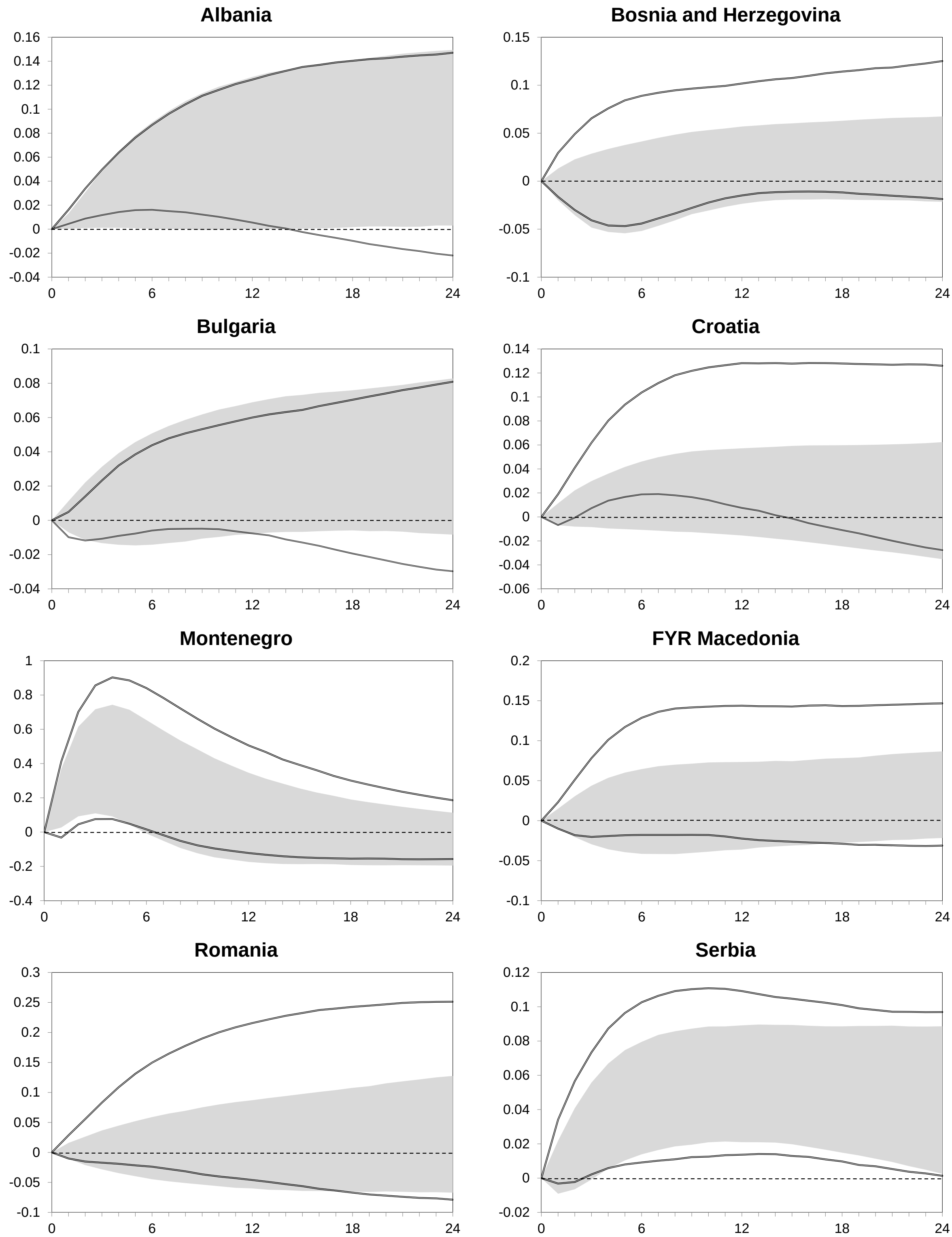
The results of a one standard deviation shadow rate shock for the euro area are depicted in figure 2.12. The double lines represent the credibility interval of the shadow rate shock, while the shaded area indicates the credibility interval of the baseline model that uses balance sheet assets to capture non-standard monetary policy. Compared to a balance sheet shock (see figure 2.2), the shadow rate shock is less persistent. Regardless of the different time horizon, the responses of the financial variables (CISS, EONIA-MRO spread and MRO) are very similar. The effects on output and prices are also in line with the baseline model, although the median price level response is lower and the credibility intervals are larger when using the shadow rate as an indicator for non-standard monetary policy measures. The spillovers to output and price levels of the SEE countries are depicted in figures 2.13 and 2.14. The outcome is qualitatively in line with the baseline model. In some cases, the price level response is less pronounced compared to the baseline model, which can be explained by the more muted response of euro area prices (see figure 2.12). As an additional robustness check, I re-estimate all models by using only the position 'Securities held for monetary policy purposes' (A070100) of the Eurosystem's balance sheet. As discussed in subsection 2.3.1, this balance sheet position reflects the most exogenous non-standard monetary policy measure and therefore serves as another robustness test for exogenous monetary policy shocks. Furthermore, I can also infer from the results whether spillovers of securities purchases are different to composite spillovers of all programmes. Again the results are qualitatively robust.

Figure 2.12: Euro area: Response to a shadow rate shock



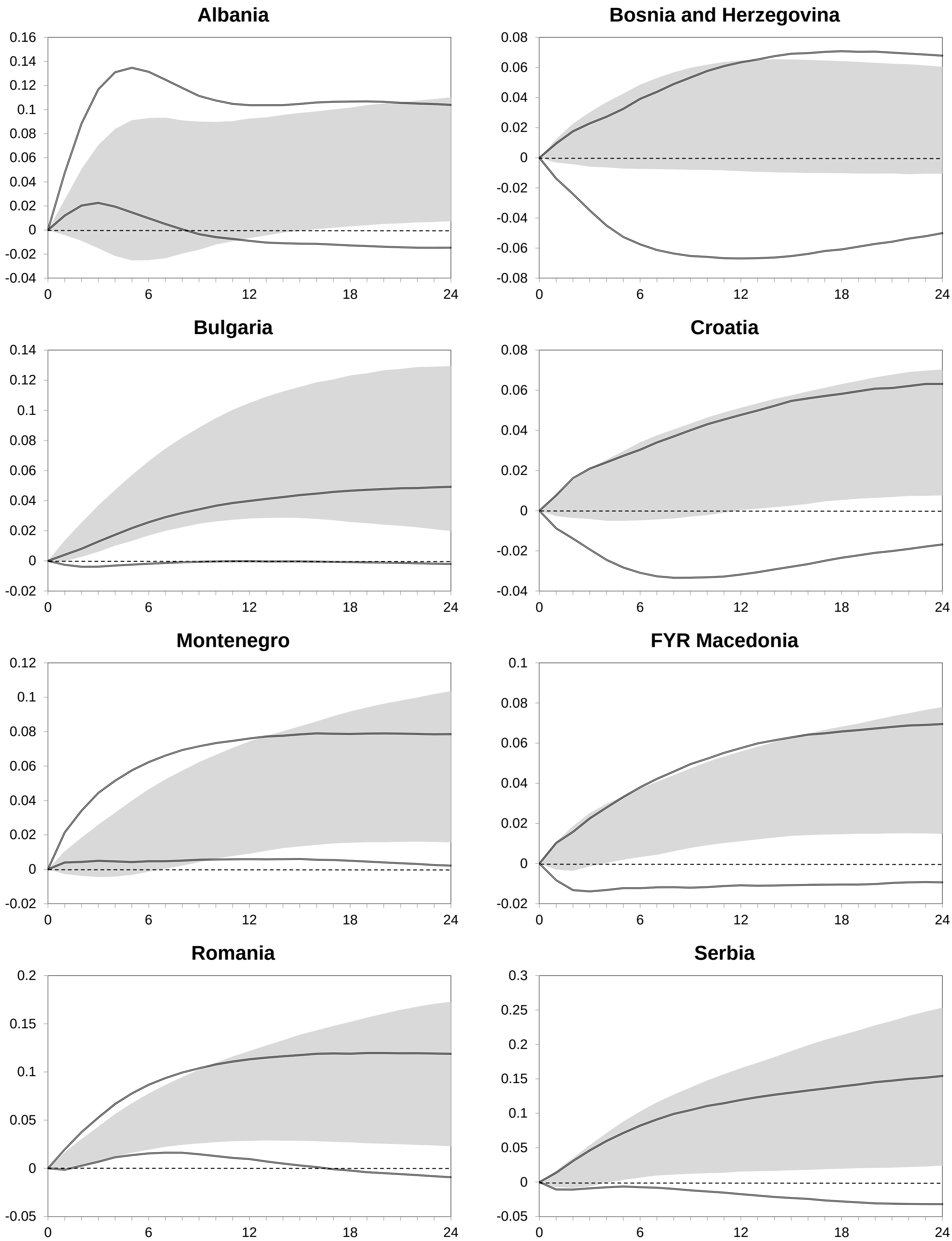
Note: The shaded area represents the 68 percent credibility interval of the baseline model, while the double lines indicate the credibility interval of the response to an expansionary one standard deviation shadow rate shock.

Figure 2.13: Response to a shadow rate shock: effect on output



Note: The shaded area represents the 68 percent credibility interval of the benchmark model, while the double lines indicate the credibility interval of the response to an expansionary one-standard-deviation shadow rate shock. For Montenegro industrial production is taken as a measure for output.

Figure 2.14: Response to a shadow rate shock: effect on price level



Note: The shaded area represents the 68 percent credibility interval of the benchmark model, while the double lines indicate the credibility interval of the response to an expansionary one-standard-deviation shadow rate shock.

2.6 Conclusion

This paper is the first one to comprehensively assess the economic impact of the euro area's non-standard monetary policy measures on the countries of south-eastern Europe (SEE). By employing bilateral structural BVAR models, I am able to identify macroeconomic spillovers as well as potential transmission channels for each country individually. Three questions are addressed in this paper: first, how have the ECB's non-standard monetary policy measures been affecting the SEE countries? Second, which channels are transmitting these shocks to SEE? Third, do different exchange rate regimes play a role in the SEE countries' responses to the shock?

The results show that the price level of all countries is positively affected by an expansionary non-standard monetary policy shock originating in the euro area, in line with the importance of euro-area imports in total imports and the expansion of domestic activity. Compared to the euro-area response of the price level, the inflationary effect on SEE is larger in most SEE countries. With regard to the output response, the shock has an expansionary effect in approximately half of the countries, which is in some cases also more pronounced than the euro area output response. These results are confirmed by robustness checks.

Regarding possible transmission channels, I find that spillovers seem to be mostly transmitted via the export channel. On the contrary, the interest rate channel exhibits a pronounced response only in a few countries, which might be driven by the relative illiquid interbank market in the SEE countries. Nevertheless, financial flows in the form of foreign direct or portfolio investments, which are not captured in the model, still might play a role.

With respect to the exchange rate regime, I find no evidence that it influences the price level or output responses. This is in line with the absence of a distinct exchange rate response in the model output for the countries under a flexible regime, driven by the very stable exchange rate the respective currencies have exhibited vis-à-vis the euro in the past years.

The current work could be extended into various directions. A comparison between the spillovers of euro-area non-standard and euro-area conventional monetary policy measures could indicate whether these measures have different international effects. Moreover, future research might include additional variables to shed more light on potential transmission channels; especially the role of financial transmission could be further explored and other channels not covered here could be added (e.g. confidence channel). Finally, a comparison with spillovers from non-standard monetary policy measures undertaken by the central banks of other large advanced economies (notably the United States) would shed light on the relative importance of euro area non-standard monetary policy, and be helpful for policymakers to design optimal policy responses to advanced economies' monetary policy measures and their (potential) reversal.

Chapter 3

The transmission of euro area monetary policy to financially euroised countries

Author: Isabella Moder, B.A. M.A.

Abstract

This paper provides a comprehensive analysis of the interest rate pass-through of euro area monetary policy to retail rates outside the euro area, contributing to the literature on transmission channels of monetary policy spillovers and on the consequences of unofficial financial euroisation. The results suggest that in the long run, more than one third of all euro retail rates in euroised countries of central, eastern and south-eastern (CESEE) Europe are linked to euro area monetary policy. Moreover, euro retail rates in CESEE adjust excessively to a change in the EONIA, and a deviation from the long-run relationship is corrected by one fifth already within the next month. Compared to euro area monetary policy, domestic monetary policy has less influence on euro retail rates, suggesting that domestic central banks in euroised countries with independent monetary policy can only partially control the 'euro part' of the interest rate channel.

3.1 Introduction

This paper provides a comprehensive analysis of the interest rate pass-through of euro area monetary policy to retail rates of euro loans and euro deposits *outside the euro area*. To the best of my knowledge, the cross-border pass-through of euro area monetary policy to financially euroised countries has not been systematically examined yet.

Since the seminal works on the interest rate pass-through undertaken by Cottarelli and Kourelis (1994) as well as Borio and Fritz (1995), a vast amount of research has emerged in order to estimate the interest rate pass-through in the euro area, linking changes in monetary policy and/or money market rates to changes in commercial banks' retail interest rates.

However, the use of the euro is not restricted to the euro area. Instead, in a number of European countries outside the euro area, mostly in the economies of central, eastern and south-eastern Europe (CESEE), economic agents voluntarily choose to hold a share of their deposits in euro, or take out loans in euro – a phenomenon which is called unofficial financial euroisation.[1] Unofficial financial euroisation is a common phenomenon in CESEE countries, as over 40% of total outstanding loans and more than one third of total deposits are denominated in euro instead of the national currency (see Table 3.1).[2]

Two main questions thus arise: First, does a long-run relationship between euro area monetary policy and euro retail rates in unofficially euroised economies exist? Second, if such a relationship exists, to what extent can domestic monetary policy influence the euro part of the interest rate channel?

This paper is related to two strands in the literature: First, it contributes to the discussion on international spillovers of advanced economies' monetary policy, and the respective transmission channels. A significant interest rate pass-through of euro area monetary policy to CESEE countries' euro retail rates would indicate that an additional transmission channel of monetary policy spillovers for euroised countries exists. The presence of such a 'forex interest rate channel' would imply that euro area monetary policy directly affects the euroised share of domestic financial conditions irrespective of the domestic monetary policy stance.

Second, and closely related, this paper contributes to the literature on the consequences of unofficial financial euroisation. While the discussion has focused mostly on its impact on financial stability and inflation, the implications for monetary policy effectiveness have not received much attention yet. A stronger interest rate pass-through to CESEE countries' retail rates for euro loans and deposits of euro area monetary policy compared to domestic monetary policy would indicate that domestic central banks can only partially control the euro part of the interest rate

[1] Other types of unofficial euroisation include currency substitution, where the euro is used for financial transaction purposes, as well as real euroisation when prices and wages are indexed to euro.

[2] Unweighted average of CESEE countries excluding the Czech Republic and Poland where unofficial euroisation is less prevalent.

Table 3.1: Unofficial financial euroisation in CESEE

Country	*Share of outstanding euro deposits in total deposits*	*Share of oustanding euro loans in total loans*
Albania	44.3%	46.4%
Bosnia and Herzegovina	33.2%	54.7%
Bulgaria	29.5%	33.7%
Croatia	51.0%	53.4%
Czech Republic	6.3%	13.2%
Hungary	16.1%	22.8%
North Macedonia	35.8%	40.4%
Poland	7.7%	10.2%
Romania	28.6%	31.5%
Serbia	61.1%	63.8%
Unweighted average	31.4%	37.0%
Unweighted average excl. Czech Republic and Poland	37.5%	43.3%

Source: Statistical annex of European Central Bank (2019). Note: Where applicable, euro-indexed deposits and loans are included. The table excludes Kosovo and Montenegro as they are unilaterally euroised without using a separate legal tender. Data as of December 2018.

channel. This has adverse implications for monetary policy effectiveness of CESEE countries that are pursuing independent monetary policy.

In order to investigate whether a long-run relationship between euro area monetary policy and euro retail rates in euroised CESEE countries exists, I use two different methods. First and foremost, I estimate the interest rate pass-through of the EONIA to a comprehensive dataset of 200 time series of euro retail rates in eight CESEE countries using an ARDL model in the spirit of Pesaran and Shin (1998) and Pesaran et al. (2001). To be able to compare the pass-through of euro area monetary policy with the pass-through of domestic monetary policy, the same exercise is thereafter undertaken with domestic monetary policy rates (where applicable). Lastly, to complement the evidence found in the previous steps, I investigate the dynamic interdependence between the respective interest rates through employing structural VAR models for a subset of three countries.

The results suggest that in the long run more than one third of all euro retail rates in euroised CESEE countries are linked to euro area monetary policy. Moreover, euro retail rates in CESEE adjust excessively to a change in the EONIA, and a deviation from the long-run relationship is corrected by 20% already within the next month. Compared to euro area monetary policy, domestic monetary policy has less of an influence on euro retail rates. This suggests that domestic central banks in countries with independent monetary policy can only partially control the euro part of the interest rate channel. Therefore, the question arises to which extent monetary policy in those CESEE countries can lean against the 'euro area wind'. For countries with a fixed exchange rate regime linked to the euro, the consequences seem to be less dramatic, with the direct interest rate pass-through to euro retail rates likely to constitute an additional channel of monetary policy transmission.

Section 3.2 of this paper provides an overview of related literature, followed by section 3.3 to introduce the empirical set-up and the underlying data. Results are presented and discussed in section 3.4. Section 3.5 provides robustness checks and section 3.6 concludes.

3.2 Related literature

The traditional literature on the interest rate pass-through examines how a change in a domestic monetary policy reference rate (typically either the key policy rate or a short-term money market rate) affects domestic banks' lending or deposit rates. The outcome yields important information about the interest rate channel of monetary policy transmission, which is one of the main channels how monetary policy decisions transmit into the real economy. The 'completeness' in the interest rate pass-through (i.e. how well the transmission of monetary policy to lending or deposit rates works) is an important metric in evaluating monetary policy effectiveness and thus the ability of a central bank to stabilise short-run economic fluctuations.

Most studies on the interest rate pass-through use small-scale error correction models following the cost pricing model developed by Rousseas (1985), Winker (1999) and De Bondt (2005), where bank retail (i.e. lending and/or deposit) rates are explained by a constant markup on the pass-through of banks' marginal cost of funds, approximated by a market interest rate. Based on this model two distinct approaches have evolved in the literature: The *monetary policy approach* tests for the transmission of the key policy rate (or of an overnight or short-term money market rate as a more volatile proxy for monetary policy) to bank retail rates. On the contrary, the *cost of funds approach* matches retail rates with money market rates of a comparable maturity to estimate the pass-through of banks' funding costs (Andries and Billon, 2016). The *monetary policy approach* thus covers the whole pass-through process from the central bank policy rate along the yield curve to longer-term money market rates and eventually to bank retail rates, while the *cost of funds approach* covers only the second stage of the monetary policy transmission process.

A large amount of literature has been dedicated to study the interest rate pass-through in Europe, see e.g. Sander and Kleimeier (2004) as well as Sørensen and Werner (2006) for the euro area on a macro level, and Gambacorta (2008) on Italian banks as an example for a micro-level analysis. According to the meta study of Andries and Billon (2016), the magnitude of the long-term pass-through has been found to be stronger for lending rates compared to deposit rates. For corporate loans, the long-run pass-through of lending rates has been estimated to be close to unity by a number of studies and thus of a higher magnitude compared to household loans.

In the short-run, the empirical evidence suggests that the pass-through is incomplete, especially for household loans and deposits with short maturities. Since the onset of the financial crisis in 2008 it seems that the magnitude of the long-term pass-through has in general decreased, while cross-country heterogeneity has increased further in the euro area (see e.g. Darracq Paries et al., 2014; Hristov et al., 2014).

A number of studies have also investigated the interest rate pass-through of domestic monetary policy in CESEE, as summarized in the meta study by Égert and MacDonald (2009). The empirical literature has found, similar to the euro area, that the most complete pass-through is present for corporate lending rates and that the most incomplete pass-through can be observed for consumer loans. For deposits, the pass-through seems to be less complete, but increases for products with higher maturities. Within the region, there exists considerable cross-country heterogeneity on the estimated pass-through coefficients. Closest to the research questions of this paper comes the work by Petrevski and Bogoev (2012), who estimate the interest rate pass-through for various lending rates of three countries in south-eastern Europe. For North Macedonia they investigate the pass-through from the EURIBOR (following the *cost of funds approach*) to lending rates in domestic currency only, but for Bulgaria and Croatia the pass-through from the EURIBOR to lending rates denominated in euro is estimated as well. The results suggest that for euro-denominated loans in Bulgaria the adjustment for long-term corporate loans is lower than for short-term corporate loans. On the other hand, no cointegrating relationship between the EURIBOR and euro lending rates is found for Croatia.

This paper is relevant in the context of the literature on international monetary policy spillovers and, more specifically, the transmission channels of monetary policy spillovers. The literature typically distinguishes between trade and financial transmission channels, see e.g. Kim (2001) for an early contribution on transmission channels of US conventional monetary policy spillovers, and Moder (2019) for recent work on spillover transmission channels of euro area non-standard monetary policy measures. With respect to financial transmission channels, overwhelming evidence has been found in the literature that US monetary policy influences global financial conditions and thus money market rates.[3] For the euro area, evidence on spillovers of ECB monetary policy shocks to other countries' money market rates is mixed (see e.g. Jiménez-Rodriguez et al., 2010; Moder, 2019). To the best of my knowledge however, the effects of US (euro) monetary policy shocks on dollar (euro) retail rates in dollarised (euroised) economies have not been investigated yet.

Additionally, this paper is related to the literature on financial dollarisation/euroisation, and more specifically on its macroeconomic consequences.[4] One strand of the literature has emphasised potentially adverse implications of dollarisation/euroisation for financial stability

[3]Related to the financial transmission channel of monetary policy shocks is the discussion on the global financial cycle and monetary policy independence kicked off by Rey (2013).

[4]On the drivers of dollarisation/euroisation, see e.g. Ize and Levy Yeyati (2003), Basso et al. (2011), Fidrmuc et al. (2013) and Brown and Stix (2015). For a discussion on regulatory policies aimed at reducing dollarisation/euroisation, see e.g. Rosenberg and Tirpák (2009) and Windischbauer (2016).

through adverse balance sheet effects (see e.g. Levy Yeyati, 2006; Bruno and Shin, 2015). In the presence of large foreign-currency exposures, domestic monetary policy might therefore on purpose mimic the monetary policy of the base country for financial stability considerations and thus exhibit 'fear of floating', even if it is at another stage of the business cycle (Reinhart et al., 2003; Georgiadis and Zhu, 2019). The literature has also investigated the consequences of dollarisation/euroisation for inflation, based on the observation that dollarised economies exhibit higher inflation rates due to a higher price elasticity to monetary shocks (Levy Yeyati, 2006). However, the impact of dollarisation/euroisation on the effectiveness of monetary policy, i.e. the transmission power of domestic monetary policy, has received very little attention so far. Brzoza-Brzezina et al. (2010) focus on the credit channel of monetary policy transmission in four central European countries and conclude that a rise in domestic interest rates prompts consumers to turn to foreign credit loans, as their results indicate that the volume of foreign currency loans increases at the expense of the volume of domestic currency loans following a domestic monetary tightening. With regard to the interest rate channel of monetary policy transmission, despite the fact that the loss of control has been mentioned in the literature (see e.g. Ize and Levy Yeyati, 2005), to the best of my knowledge no systematic attempt has been made so far to quantify those effectiveness losses.

This paper therefore aims to fill two gaps in the literature: First, it provides an assessment on the importance of the forex interest rate channel in transmitting spillovers of foreign monetary policy shocks. Second, it examines whether unofficial euroisation constrains the effectiveness of domestic monetary policy.

3.3 Methodology

3.3.1 Empirical set-up

In the first step, I estimate the interest rate pass-through of the euro area monetary policy rate to a large dataset of 200 time series of euro retail rates in eight CESEE countries[5], for which the proportion of currency replacement by the euro in both deposits as well as loans is substantial.

More specifically, following Pesaran and Shin (1998) and Pesaran et al. (2001) I estimate a general Autoregressive Distributed Lag (ARDL) model for each interest rate series:

$$\Delta i_t^R = \alpha + \beta_1 i_{t-1}^R + \beta_2 i_{t-1}^M + \sum_{p=1}^{P-1} \gamma_{1p} \Delta i_{t-p}^R + \sum_{q=0}^{Q-1} \gamma_{2q} \Delta i_{t-q}^M + \varepsilon_t \tag{3.1}$$

which is another representation of the following error correction model:

[5]Those countries are Albania, Bosnia and Herzegovina, Bulgaria, Croatia, Hungary, North Macedonia, Romania and Serbia.

$$(3.2) \qquad \Delta i_t^R = \sum_{p=1}^{P-1} \gamma_{1p} \Delta i_{t-p}^R + \sum_{q=0}^{Q-1} \gamma_{2q} \Delta i_{t-q}^M + \beta_1 (i_{t-1}^R + \frac{\beta_2}{\beta_1} i_{t-1}^M + \frac{\alpha}{\beta_1}) + \varepsilon_t$$

and the error correction term is thus defined as follows:

$$(3.3) \qquad EC_t = i_{t-1}^R + \frac{\beta_2}{\beta_1} i_{t-1}^M + \frac{\alpha}{\beta_1}$$

In the equations above i_t^R and i_t^M represent the retail rate and the euro area monetary policy rate, respectively, at time t, while Δ is the first difference operator. α denotes the constant markup, representing the risk premium, maturity premium and the banks' profit margin. The associated long-run pass-through coefficient of the euro area monetary policy rate to the retail rate is given by β_2/β_1 and the speed of adjustment is indicated by the negative error correction coefficient β_1. The coefficients γ_{1p} and γ_{2q} denote the short-run adjustment dynamics towards the long-run equilibrium. The number of lags (P, Q) is determined by the Bayesian Information Criterion (BIC). Out of the five alternative representations offered by Pesaran and Shin (1998) and Pesaran et al. (2001) for ARDL models, case 2 is chosen that integrates the constant α into the error correction term and thus into the long-run relationship to represent the markup α/β_1.

ARDL models are superior to other approaches when testing for the existence of a relationship between variables in levels, since they are robust to misspecification of integration orders as the regressors do not necessarily all have to be integrated of order one (Pesaran et al., 2001). Instead, the underlying regressions can be I(0), I(1) or mutually cointegrated, which reduces the necessary degree of pre-testing and the associated uncertainty. Furthermore, estimating both short- and long-run parameters together in a single equation as is the case in the ARDL approach might help to avoid small-sample bias that occurs when estimating the long-run parameters separately in the first estimation step (Romilly et al., 2001).

Before the model ist estimated, I test each series for the presence of I(2) and, if a cointegrating order of two is found, exclude it from the analysis. Subsequently the model ist estimated with a maximum of 8 lags for both variables, with insignificant lags being dropped according to the BIC. Cointegration is then tested by using the bounds test for cointegration from Pesaran et al. (2001), which is based on a standard F-test, with the critical values provided by Narayan (2004) for smaller sample sizes. The following null hypothesis of no relationship in levels is tested:

$$(3.4) \qquad H_0 : \beta_1 = \beta_2 = 0$$

If cointegration is found between the respective retail rate i^r and the euro area monetary policy rate i^m, the analysis continues with estimating the size of the long-run pass-through and the speed of adjustment. The estimations are tested for serial correlation and heteroskedasticity, and the model is adjusted if necessary by including more lags or by using a HAC covariance matrix adjustment to correct the value of the test statistics used for heteroskedasticity.

As discussed in section 3.2, the explanatory variable in the ARDL model can either represent the policy rate to test for the pass-through of monetary policy, or a money market rate with similar maturity to estimate the pass-through of funding costs to retail rates (see Sander and Kleimeier, 2004). This paper follows the first approach, since the interest lies in the international transmission of euro area *monetary policy*. Moreover, after the global financial crisis, commercial banks in CESEE countries have shifted their funding away from money markets and/or parent bank funding towards domestic (euro) deposits (reflected in the decline in loan-to-deposit ratios, see e.g. Vienna Initiative, 2019). Thus a money market rate with a similar maturity might not even represent the real funding costs banks are facing.

While I estimate the same models for countries with fixed and floating exchange rate regimes in the first part, I focus on the interest rate pass-through of domestic monetary policy to euro retail rates in the second step. Out of the countries included in this paper, Albania, Hungary, Romania and Serbia follow inflation targeting regimes with managed or freely floating exchange rates, and are thus the focus of the analysis of the domestic monetary policy pass-through.[6] To this end, I follow the procedure outlined above, except that i_t^M now represents the domestic monetary policy rate i_t^{Mdom} instead of the euro area monetary policy rate, and that the analysis is – naturally – confined to countries with independent monetary policy, which excludes four out of the eight countries in the sample.[7]

Lastly, in order to obtain additional evidence on the interest rate pass-through of euro area monetary policy, I undertake a similar analysis by employing VAR models, which are better suited to capture the interdependency between the respective interest rates than a single-equation ARDL model.[8] In particular for countries with independent monetary policy, the domestic policy rate can be included in the VAR analysis, providing valuable information on the interdependence of euro area and domestic monetary policy and the effect on the euro retail rates. Moreover, VAR models provide valuable information on the dynamic transmission, i.e. after which period the long-run pass-through of monetary policy has fully materialised. In order to reduce the dimensionality of the data, I first compute the first principal component of the estimated covariance matrix of all lending and deposit series for each country, respectively. Since the scales are the same (interest rates) and the variances of the respective series are similar, the covariance matrix is used to compute the first principal component. Next, I use the first principal component series as input into the following country-specific vector autoregression (VAR) model

[6] The remainder of the countries peg their currencies to the euro and does do not have independent monetary policy tools at their disposal.

[7] These countries are Bosnia and Herzegovina, Bulgaria, Croatia and North Macedonia.

[8] Unfortunately, due to data limitations (see 3.3.2), this last part of the analysis can only be undertaken for a subset of three countries.

at monthly frequency:

$$\sum_{s=0}^{p} \mathbf{A}(s)\mathbf{y}(t-s) + \mathbf{c} = \boldsymbol{\varepsilon}(t) \tag{3.5}$$

$\mathbf{A}(s)$ denotes the matrix of the estimated coefficients, $\mathbf{y}(t)$ represents a vector of macroeconomic variables and $\mathbf{c}$ a vector of constants. The vector $\boldsymbol{\varepsilon}(t) \sim N(0, \Sigma)$ denotes structural shocks. Each country's VAR model includes the euro area monetary policy rate, i_t^M, the domestic monetary policy rate (if applicable), i_t^{Mdom} and the estimated first principal component series of both the lending and deposit rates, $PC1_t^{Lending}$ and $PC1_t^{Deposit}$. The domestic monetary policy rate, as well as the principal component series of the lending and deposit rates are set block exogenous to the euro area monetary policy rate, constraining the model so that neither current nor past economic developments in the CESEE countries affect euro area monetary policy. This feature introduced by Cushman and Zha (1997) has been used frequently in the literature and is well suited for modelling spillovers from large to small economies, as it helps to identify spillovers from the viewpoint of the small open economy and reduces the number of parameters to be estimated (Cushman and Zha, 1997).

For the shock identification necessary to generate impulse response functions, a standard Cholesky decomposition is used. The variables enter the model in the following decreasing order of exogeneity: (1) Euro area monetary policy rate i_t^M, (2) domestic monetary policy rate (if applicable) i_t^{Mdom}, (3) first principal component series of deposit rates $PC1_t^{Deposit}$ and (4) first principal component series of lending rates $PC_t^{Lending}$. While the monetary policy rates are clearly more exogenous than the retail rates and thus put first in the model, the ordering of the latter is not straightforward. The results of the VAR models are however robust to a reverse ordering of (3) and (4) in the Cholesky decomposition (see section 3.5). As the data enters the model in monthly frequency, I use 12 lags for the VAR analysis, thus $p = 12$. All estimations are carried out in EViews10.

3.3.2 Data

In line with the literature following the *monetary policy approach* (see e.g. Hristov et al., 2014; Gambacorta et al., 2015; von Borstel et al., 2016), the EONIA is used as a proxy for euro area monetary policy, because it exhibits more volatility and is not restricted by the zero lower bound in times of unconventional monetary policy. However, the EONIA has been varying less when hit its effective lower bound constrained by the deposit rate in early-2016. This raises the question whether another measure of ECB monetary policy, e.g. the Eurosystem balance sheet or a shadow rate would be better suited for assessing the pass-through of monetary policy. Plotting the EONIA together with the (medians of) the respective lending and deposit rates for each country suggests that the EONIA and the respective retail rates (see figure 3.1) generally move into the same direction. Interest rates of both euro loans and euro deposits in CESEE appear to be restricted as well by some effective lower bound, namely the deposit rates close to 0% and the lending rates at some markup above the deposit rates. This indicates that the EONIA is well suited as a measure for the pass-through of euro area monetary policy. Moreover, it should be noted that the EONIA does incorporate non-standard monetary policy measures to a certain extent, which is evident in the decoupling of the EONIA from the key policy (MRO) rate since the onset of the global financial crisis that marked the introduction of non-standard measures. For the estimations of the pass-through of domestic monetary policy rates to euro retail rates, I use the respective domestic key monetary policy rates since overnight or short-term money market rates are not available over the time span for all countries.

Excluding two countries with relatively low loan and deposit euroisation (Czech Republic and Poland, where euro loans and deposits account for less than 15% and 10%, respectively, of the total volumes) leaves eight countries in CESEE at my disposal[9]. For each country, all publicly available statistics of euro lending or deposit rates for new business[10] that do not exhibit data gaps are used as individual input into each ARDL model. Including all available time series in the ARDL models instead of focusing on a few main statistics makes the display and discussion of results challenging. Nevertheless, constraining the analysis to a few statistics instead of all available ones would come at a considerable cost as it would greatly reduce the richness of the results, given the very heterogeneous composition of available statistics across countries that are not harmonised.

Table 3.2 shows the decomposition of statistics for each country into lending and deposit rates. In total 200 retail rates for eight countries are analysed in this paper. The series cover interest rates of both loans and deposits to/from households and non-financial corporates, respectively, and differ in terms of sectors, currencies[11], maturity, purpose, interest rate fixation and loan amount. By far the largest number of interest rate statistics for euro loans and deposits are available for Bulgaria and Croatia, followed by Albania and Bosnia and Herzegovina. The composition of deposit versus lending rates is – with two exceptions – tilted towards lending rates, for which generally more time series are available. A list of all statistics included can be found in the appendix.

The statistics are also heterogeneous in their starting points. While data is available from January 2005 (Hungary) and January 2007 (Bulgaria and Romania), it starts much later for other countries, especially for North Macedonia (January 2015) and Albania (December 2015).[12]

[9] The respective countries are Albania, Bulgaria, Bosnia and Herzegovina, Croatia, North Macedonia, Hungary, Romania and Serbia.

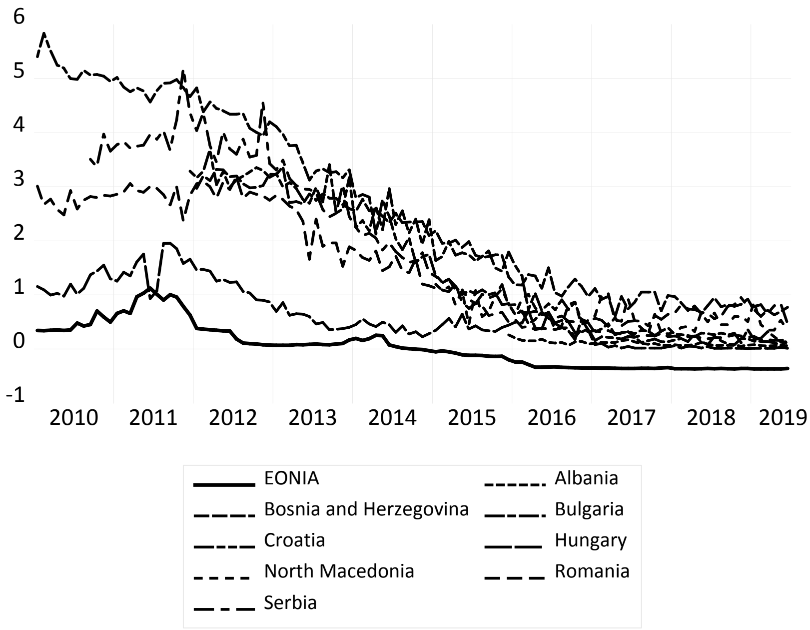
[10] Series for new business as opposed to outstanding loans are used, as the former provide a much better indicator of current interest rates while the latter react with a considerable time lag given their composition.

[11] Some series cover interest rates of so-called euro-indexed deposits and loans. Those products are paid and technically repaid in domestic currency, but cash flows are indexed to changes in the exchange rate between the euro and the domestic currency.

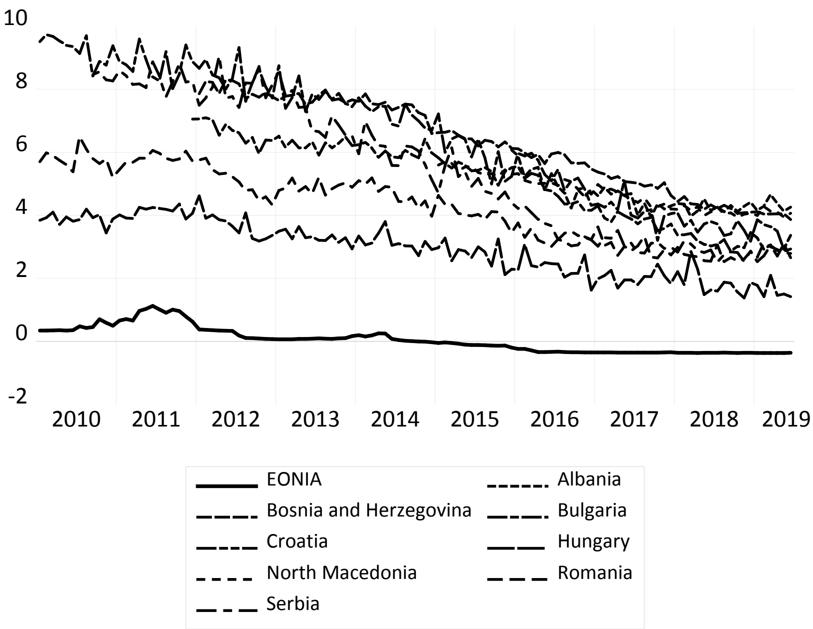
[12] For Albania, all series are available from 2015, for Bosnia and Herzegovina from 2012, for Bulgaria some from 2004 and some from 2007, for Croatia from 2011, for Hungary most from 2005 and a few from 2012, for North

Figure 3.1: Evolution of retail rates over time

(a) Deposit rates



(b) Lending rates



Source: Author's calculations and national central banks. Note: Lines depict the median of the respective time series listed in the Appendix.

In order to make the series somewhat comparable without loosing too many observations, and to account for structural breaks in the time series during the global financial crisis (see e.g. Gambacorta et al., 2015), I start the estimation sample in January 2010 or later whenever the time series become available (table 3.2). With a monthly frequency, at most 114 observations are available for each estimation. All statistics are derived from the webpages of the respective national central banks.

Table 3.2: Summary of statistics

Country	*#of series*	*of which: deposit rates*	*of which: lending rates*	*Sample*
Albania	28	11	17	12/15-06/19
Bosnia and Herzegovina	26	10	16	01/12-06/19
Bulgaria	52	25	27	01/10-06/19
Croatia	41	12	29	12/11-06/19
Hungary	14	5	9	01/10-06/19
North Macedonia	11	5	6	01/15-06/19
Romania	16	9	7	01/10-06/19
Serbia	12	9	3	09/10-06/19
Sum	200	86	114	

For the VAR models, the available statistics are separated between loan and deposit interest rates and by country to derive the series of first principal components. As the VARs are richer models compared to the ARDL models, more observations are needed. Therefore, unfortunately, the VAR analysis can only be performed for a subset of three countries (Bulgaria, Hungary and Romania) where the availability of the interest rate series from January 2007 (Bulgaria and Romania) and January 2005 (Hungary)[13] provides enough observations to obtain meaningful results.

3.4 Results

In the following two subsections, interest rate statistics are reported to be cointegrated with the EONIA (or the domestic monetary policy rate) if the bounds test by Pesaran et al. (2001) yields a conclusive result at least at the 5% significance level, and if the coefficient of the long-run pass-through of the EONIA (domestic monetary policy rate) is found to be significant at least

Macedonia from 2015, for Romania from 2007 and for Serbia from 2010.

[13]While the bulk of time series are available as of January 2005 for Hungary, five time series have to be excluded from the principal components analysis as their availability starts much later.

at the 5% level.[14] Only if cointegration is found, the long-run pass-through and the speed of adjustment are estimated (see section 3.3.1).

Given the high number of coefficients estimated in the 200 ARDL models and the large heterogeneity in terms of the available statistics, the presentation of the results focuses on two metrics: the differences between lending and deposit rates on the one hand, and between countries on the other hand. For the presentation of the estimation results, I focus on the long-run pass-through and on the speed of adjustment. The short-run pass-through is in many cases insignificant and thus not further discussed here.[15]

3.4.1 Pass-through of euro area monetary policy

In total, out of 200 euro interest rate statistics tested, 74 have been found to be cointegrated with the EONIA. In other words more than one third or 37% of all published euro interest rate series exhibit a long-run relationship with euro area monetary policy (table 3.3). Distinguishing by products, the percentage of deposit rates that are cointegrated with the EONIA is slightly higher than the cointegration of lending rates. Turning to the estimated coefficients of the cointegrated series, the long-run pass-through is clearly above unity for both deposit as well as lending rates in most cases (see boxplot in figure 3.2), implying that deposit and lending rates adjust excessively to a change in the EONIA. This result is puzzling, as in the literature such an over-adjustment has neither been found for the interest rate pass-through within the euro area (which even points to a decreasing magnitude of the long-run pass-through after the global financial crisis), nor for the pass-through of domestic monetary policy in CESEE countries (see section 3.2). Still, the result of this paper that the magnitude of the long-run pass-through is stronger for lending rates compared to deposit rates is in line with the findings of the literature on the euro area (see the metastudy by Andries and Billon, 2016).

One potential explanation for the unusually large coefficient of the EONIA in the estimation could be related to decreasing markups between the respective retail rates and the EONIA (see figure 3.1).[16] As the estimated ARDL models assume a constant markup in the data generating process, and the decreasing markups coincide with the drop of the EONIA towards its effective lower bound during the sample period, the resulting relationships between the EONIA and the respective retail rates might be picked up as being multiplicative instead of additive by the model. Another factor that might potentially be related to the results could be a non-linear or

[14]It should be noted that the test statistic used implies that the interest rate statistics are independent from each other. In subsection 3.5 I report the results of robustness checks of adjusting the significance level for multiple testing.

[15]The insignificant short-run pass-through is confirmed in the results of the VAR models in section 3.4.3.

[16]The decreasing markups in turn can be explained by the decreasing competition of banks' for deposits, in line with the steadily growing deposit base in many countries (see Vienna Initiative, 2019). On the credit side, the decrease in the markup could be related to a number of factors, e.g. stronger competition between banks for lending, or a shift to less risky credit classes and thus lower aggregate credit risk.

asymmetric interest rate pass-through of the EONIA to euro retail rates, where retail rates are more responsive to a fall in the EONIA compared to a situation where the EONIA rises.

Table 3.3: Interest rate pass-through of the EONIA by country

Country	*Cointegrated all rates*	*Cointegrated deposit rates*	*Cointegrated lending rates*
Albania	35.7%	36.4%	35.3%
Bosnia and Herzegovina	19.2%	30.0%	12.5%
Bulgaria	28.8%	56.0%	3.7%
Croatia	56.1%	50.0%	58.6%
Hungary	50%	60%	44.4%
North Macedonia	45.5%	80%	16.7%
Romania	37.5%	11.1%	71.4%
Serbia	25.0%	11.1%	66.7%
Sum	37.0%	41.9%	33.3%

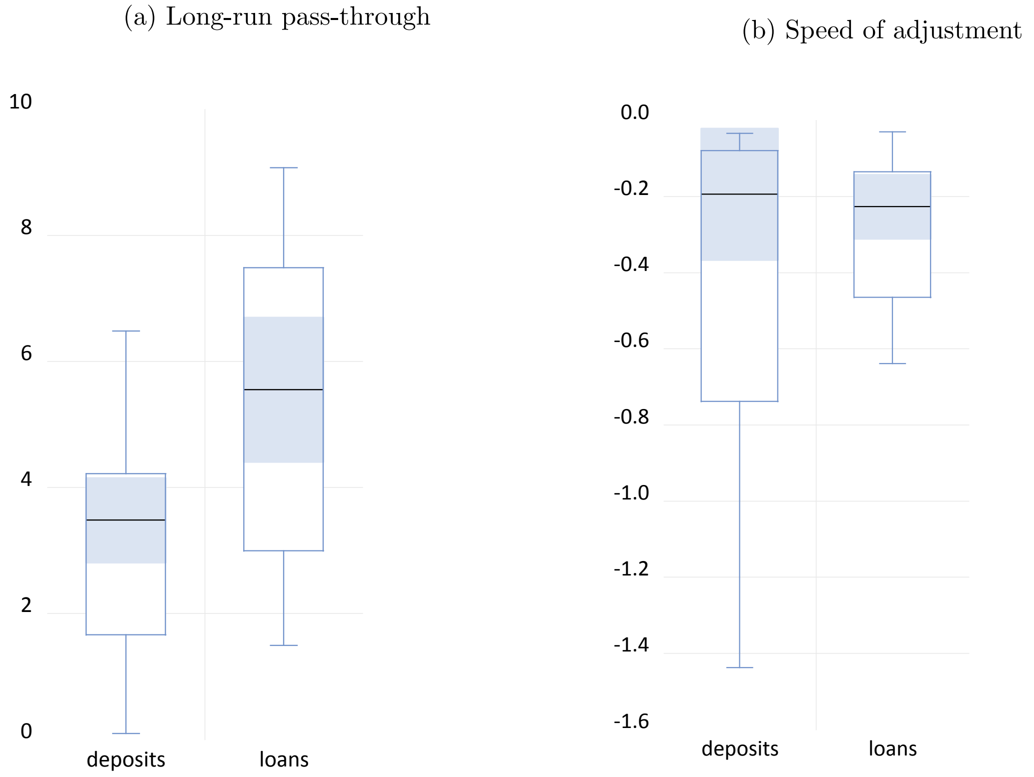
Note: A series is considered to be cointegrated with the EONIA if the test by Pesaran et al. (2001) with the critical values provided by Narayan (2004) yields a conclusive result at least at the 5% significance level, and if the long-run coefficient of the EONIA is significant at least at the 5% significance level.

The speed of adjustment is by definition negative and measures to what extent movements into disequilibrium are corrected within one period. For both deposit as well as for lending rates the results suggest that about 20% of a deviation from the estimated long-run relationship is corrected already in the next month (figure 3.2). The variance of the distribution of estimated coefficients is higher for deposit than for lending rates, with the speed of adjustment for deposit rates ranging from almost zero to values below -1 that indicate overshooting (a deviation from the equilibrium is more than corrected in the next period).

Grouping the results by countries (table 3.3) suggests that the highest share of cointegrated retail rate series is observed in Croatia, followed by Hungary. In both countries half or even more than half of all euro retail rate statistics available are cointegrated with the EONIA. On the contrary, cointegration is the lowest for euro retail rates in Bosnia and Herzegovina and Serbia. Table 3.3 also reveals that no clear pattern can be detected regarding the exchange rate regimes of countries.[17] In terms of EU membership however, the results suggest that EU members might have a somewhat higher share of cointegrated interest rate statistics. In general, keeping in mind the heterogeneity of the available statistics across countries and the estimated confidence bands, the results and cross-country comparisons should be treated with caution. In five out of

[17]For example, the two countries with the highest degree of cointegrated series are Croatia and Hungary, which operate under a fixed and a flexible exchange rate regime, respectively. The same holds for the countries with the lowest percentage of cointegrated series, Serbia and Bosnia and Herzegovina, with the former letting its currency float (although with substantial interventions) and the latter's currency being tied to the euro through a currency board.

Figure 3.2: Estimated coefficients of the interest rate pass-through of the EONIA: by product

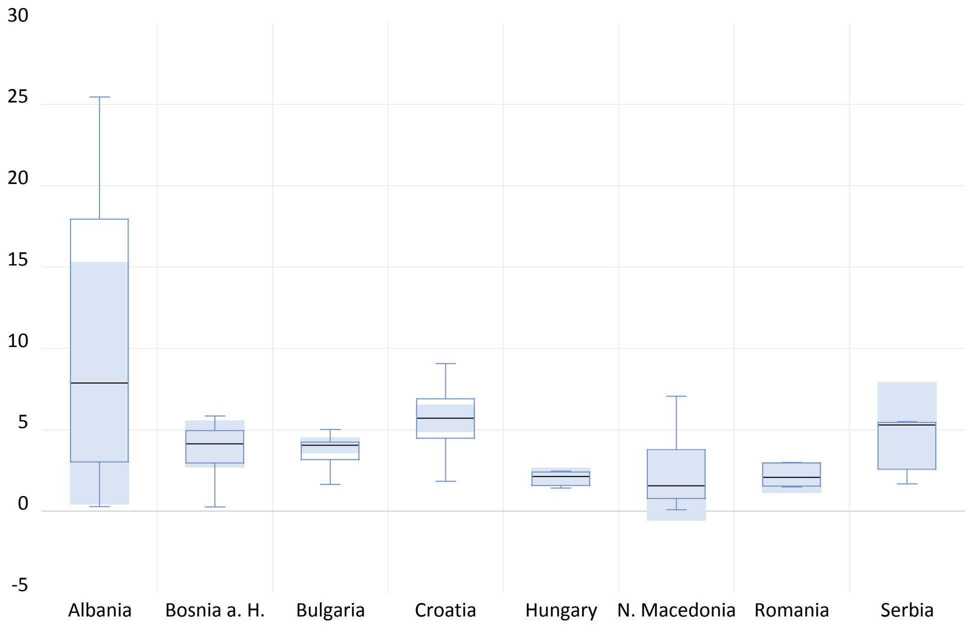


Note: Distribution of estimated coefficients in ARDL models. Distribution consists of 36 and 38 coefficients for deposit and lending rates, respectively. The shaded areas represent the 95% confidence intervals for the medians, which are depicted as the black horizontal lines.

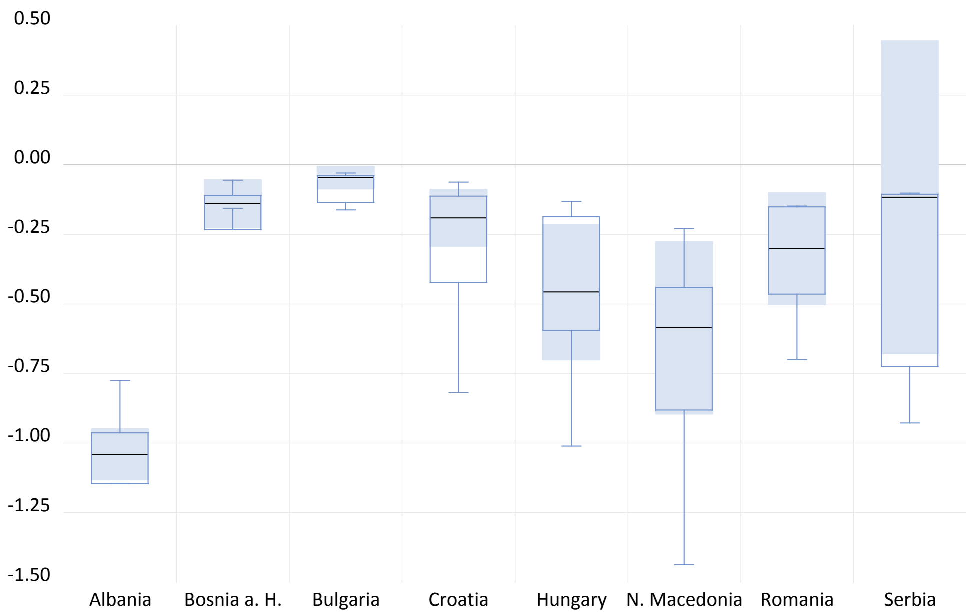
eight countries the share of cointegrated deposit rates is higher compared to lending rates. The heterogeneity across countries is also reflected in the distribution of the estimated coefficients of the cointegrated series (figure 3.3). The median of the estimated long-run pass-through is the lowest in North Macedonia, followed by Romania and Hungary, and the highest in Albania (although with considerable uncertainty). The adjustment to disequilibria is also heterogeneous, with the most sluggish adjustment suggested for Bulgaria, Bosnia and Herzegovina and Serbia (the countries where the percent of cointegrated series in total series is the lowest), and an overshooting tendency for many interest rate statistics in Albania.

Figure 3.3: Estimated coefficients of the interest rate pass-through of the EONIA: by country

(a) Long-run pass-through



(b) Speed of adjustment



Note: Distribution of estimated coefficients in ARDL models. Distribution consists of 10, 5, 15, 23, 7, 5, 6 and 3 coefficients for each of the countries, respectively. The shaded areas represent the 95% confidence intervals for the medians, which are depicted as the black horizontal lines.

3.4.2 Pass-through of domestic monetary policy

The second part of the analysis deals with the pass-through of domestic monetary policy to euro deposit and lending rates in CESEE countries with independent monetary policy. Table 3.4 displays the percentage of cointegrated series for Albania, Hungary, Romania and Serbia. In sum, the percentage of interest rate statistics cointegrated with the domestic monetary policy rate is *lower* than the percentage of interest rate statistics cointegrated with the EONIA, suggesting that domestic central banks can only partially control the 'euro' part of the interest rate channel. For the countries with independent monetary policy that are CESEE EU members (Hungary and Romania), the share of euro retail rates cointegrated with the EONIA is higher than the share of retail rates cointegrated with the domestic monetary policy, while the opposite holds true for CESEE non-EU members (Albania and Serbia). As CESEE EU members tend to be more financially integrated with the euro area compared to CESEE non-EU members, this result provides evidence for the hypothesis of Passari and Rey (2015), who show that in determining a country's exposure to the global financial cycle, capital mobility matters more than exchange rate policy, as the global financial cycle affects countries with different exchange rate regimes in a similar way. Looking at the specific statistics, it is not necessarily the same series that are cointegrated with both the EONIA and the domestic monetary policy rate.[18]

Table 3.4: Pass-through of domestic monetary policy

Country	*Cointegrated all rates*	*Cointegrated deposit rates*	*Cointegrated lending rates*
Albania	39.3%	63.6%	23.5%
Hungary	28.6%	40.0%	22.2%
Romania	25.0%	22.2%	28.6%
Serbia	33.3%	11.1%	100.0%
Sum	32.9%	35.3%	30.6%

Note: A series is considered to be cointegrated with the domestic monetary policy rate if the test by Pesaran et al. (2001) with the critical values provided by Narayan (2004) yields a conclusive result at least at the 5% significance level, and if the long-run coefficient of the domestic monetary policy rate is significant at least at the 5% significance level.

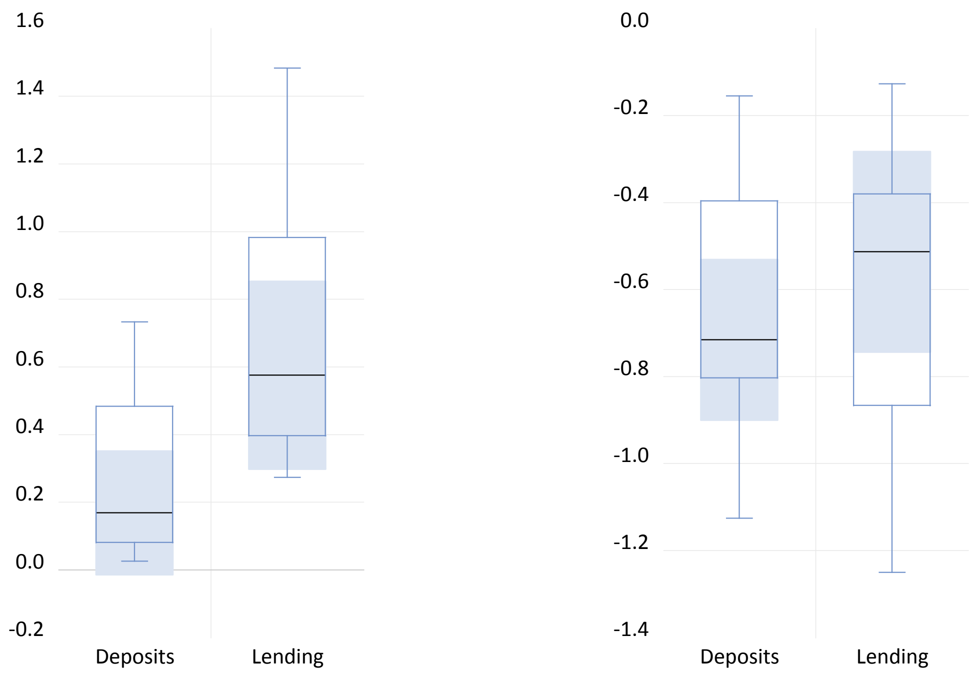
The long-run pass-through is below unity for both deposit and lending rates, implying an incomplete long-run pass-through (figure 3.4). This suggests that in the long-run relationship, the euro retail rates react less than one-to-one to a change in the domestic monetary policy rate, and that the puzzling result of the sizeable overadjustment of retail rates to the EONIA does not extend to domestic monetary policy. On the other hand, the speed of adjustment seems to be faster to domestic monetary policy for both deposit and lending rates as compared to the EONIA, but the estimated coefficients are subject to a higher variance. Like the results in the

[18]See a list of statistics in the appendix, where series are marked if they are cointegrated with the **EONIA** and/or *domestic monetary policy rate.*

previous section the results of the pass-through of domestic monetary policy to interest rates in domestic currency in CESEE found in the literature, the long-run pass-through is less complete for deposits compared to lending rates, whereas the speed of adjustment is higher for deposit rates.

Figure 3.4: Estimated coefficients of domestic monetary policy pass-through: by product

(a) Long-run pass-through (b) Speed of adjustment

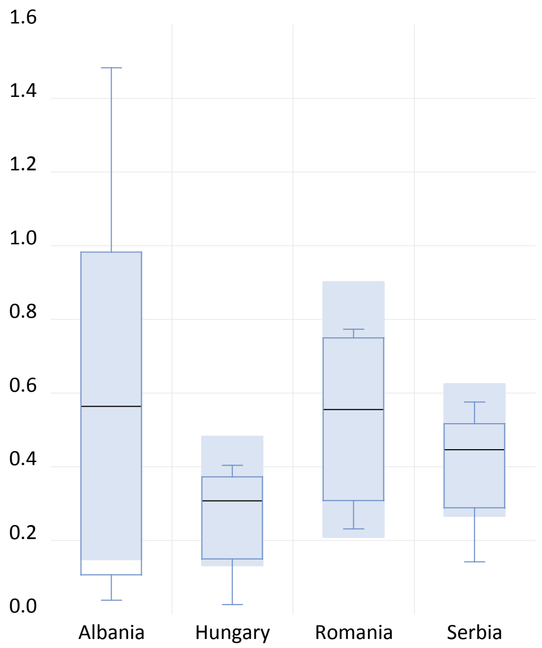


Note: Distribution of estimated coefficients in ARDL models. Distribution consists of 12 and 11 coefficients for deposit and lending rates, respectively. The shaded areas represent the 95% confidence intervals for the medians, which are depicted as the black horizontal lines.

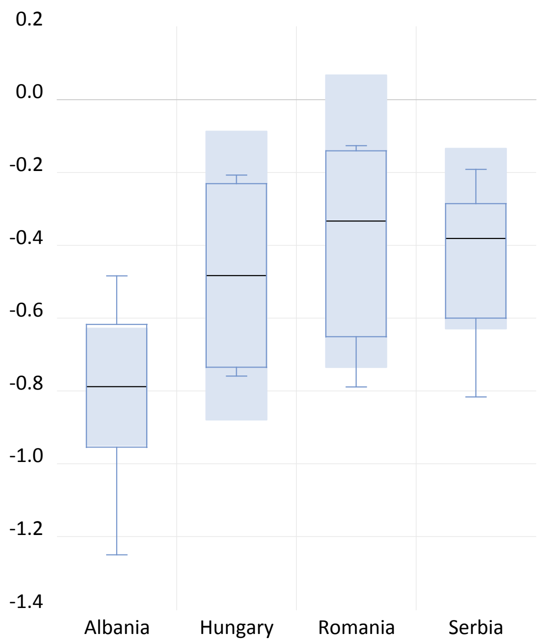
The estimated coefficients vary again considerably across countries, but are this time subject to high uncertainty given the large confidence intervals (figure 3.5). It is noteworthy to point out that the fast speed of adjustment seems to be in part driven by Albania, where euro retail rates react faster to domestic monetary policy changes compared to the other countries.

Figure 3.5: Estimated coefficients of domestic monetary policy pass-through: by country

(a) Long-run pass-through



(b) Speed of adjustment



Note: Distribution of estimated coefficients in ARDL models. Distribution consists of 11, 4, 4, and 4 coefficients for each of the countries, respectively. The shaded areas represent the 95% confidence intervals for the medians, which are depicted as the black horizontal lines.

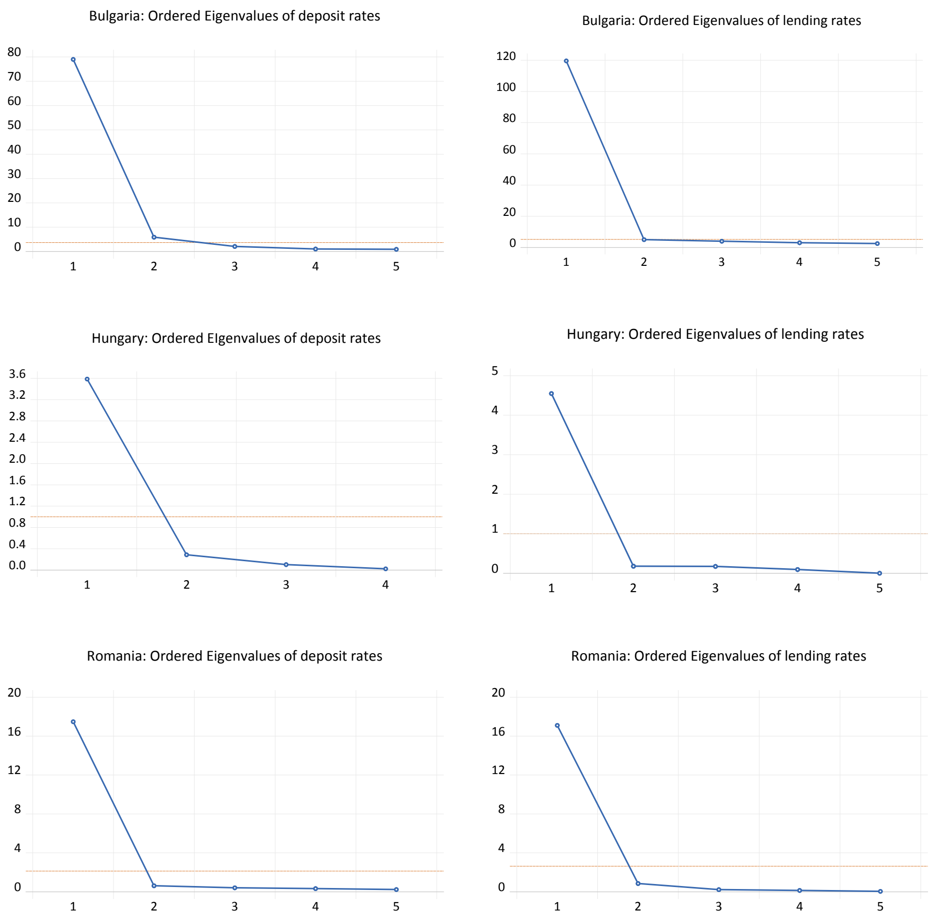
3.4.3 Evidence on dynamic transmission of euro area monetary policy

In this subsection the previous findings based on ARDL models are complemented with additional analysis obtained from employing structural VAR models. More specifically, I am interested in the dynamic transmission of euro area monetary policy to euro retail rates in CESEE. As the short-term pass-through is in most cases insignificant, a VAR analysis can indicate how long it takes for an EONIA shock to be transmitted to euro retail rates in CESEE. Additionally, the analysis can shed light on the dynamic interplay between euro area monetary policy and domestic monetary policy. However, as discussed in section 3.3.1, unfortunately only the time series of three countries (Bulgaria, Hungary and Romania) are long enough to provide enough degrees of freedom for the VAR analysis.

The outcome of the principal component analysis shows that the first principal direction already captures most of the common movement of the respective variables (figure 3.6), as the second principal direction is close to or below the average of the eigenvalues. Thus, using the first principal direction in the VAR analysis should be sufficient.

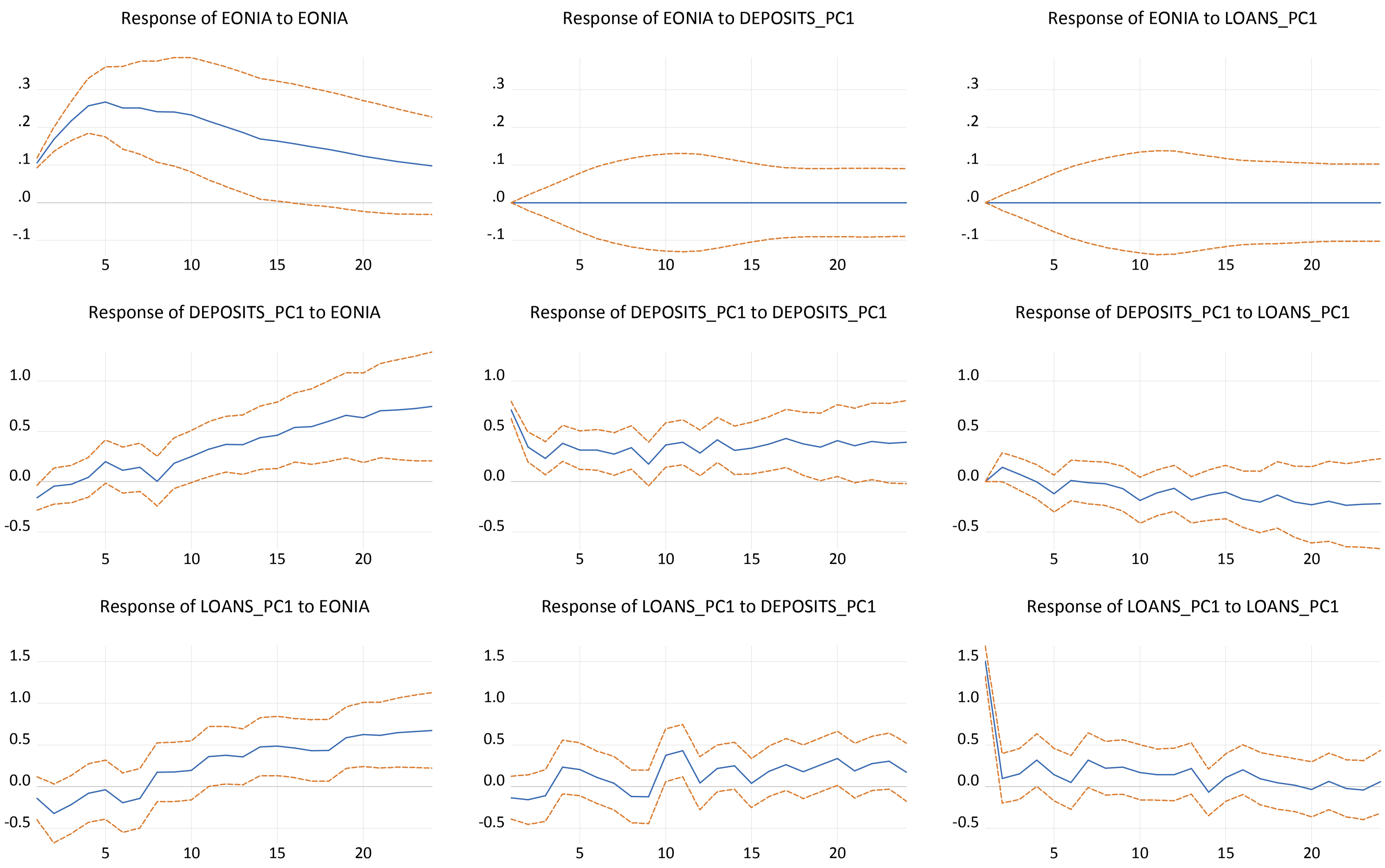
The impulse response functions for Bulgaria are shown in figure 3.7, where the three figures in the first column are the main focus of interest. A shock of the EONIA rate clearly affects the principal component of the deposit and the lending rates. The effect becomes significant only in the medium-term after approximately 11 months, and reaches its peak response after approximately 20 months. This confirms the result of the ARDL models that the short-run pass-through of the EONIA to the respective euro retail rates is mostly insignificant. It should also be noted that the response of both deposit and lending rates to the EONIA shock appears to be very persistent. The magnitude of the response of deposit and lending rates to an EONIA shock is also in line with the ARDL model output. The magnitude of the peak response of the deposit and lending rates (0.66 and 0.62 percentage points, respectively) is more than twice the size of the peak EONIA shock (0.27 percentage points), which points to a long-run pass-through of more than 200%.

Figure 3.6: Scree plots



Note: Eigenvalues arranged in order of principality with the horizontal line equal to the average of the eigenvalues.

Figure 3.7: Bulgaria: Impulse Response Functions



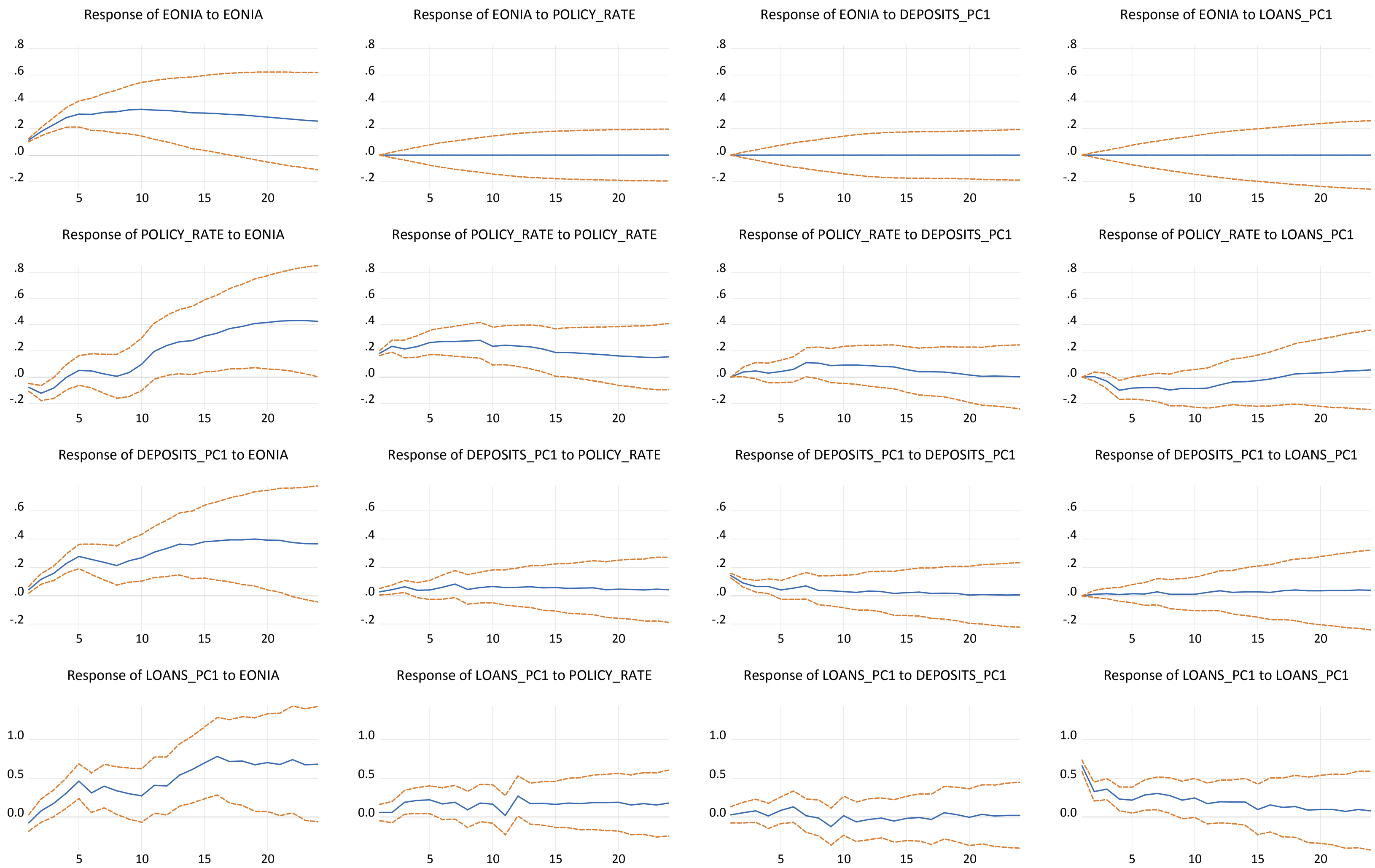
Note: PC1 refers to the first principal component of the respective series. Shocks are identified via a Cholesky decomposition. The dashed lines report the 95% confidence intervals. The x-axis reports months, the y-axis percentage point changes.

For Hungary the VAR model includes additionally the domestic monetary policy rate in order to assess the interaction of a euro area monetary policy shock with domestic monetary policy (figure 3.8). A shock in the EONIA triggers an immediate but short-lived negative response of the domestic monetary policy rate, which turns positive after 11 months. Despite the immediate counteracting response of domestic monetary policy, the response of the principal component series of deposits is positive and significant from the start, which is different from the case of Bulgaria discussed above.[19] Similarly, the response of the principal component of the lending rates also turns significant after 3 months, much earlier than in the case of Bulgaria. Still, both responses reach their peak again only in the medium term after approximately 16 and 19 months. Comparing the magnitude of the response of the principal component of the deposit and lending rates (0.40 and 0.83, respectively) to the EONIA shock (0.33 percentage points) suggests that in Romania the long-run pass-through of the EONIA to lending rates is again more than 200%, but much stronger compared to deposit rates.

By means of including the domestic monetary policy rate, the response of the principal components of deposit and lending rates to an EONIA shock can be compared to the response of the same variables to a shock of the domestic policy rate. A domestic monetary policy shock (figure 3.8, second column), which is broadly similar in magnitude compared to the EONIA shock, triggers only a relatively muted and barely significant response of the euro deposit and lending rates.

[19] While the model for Hungary includes the domestic monetary policy rate and are thus not directly comparable to the model of Bulgaria, the response of the principal components of the deposit and lending rates to a shock in the EONIA are robust to excluding the domestic policy rate.

Figure 3.8: Hungary: Impulse Response Functions

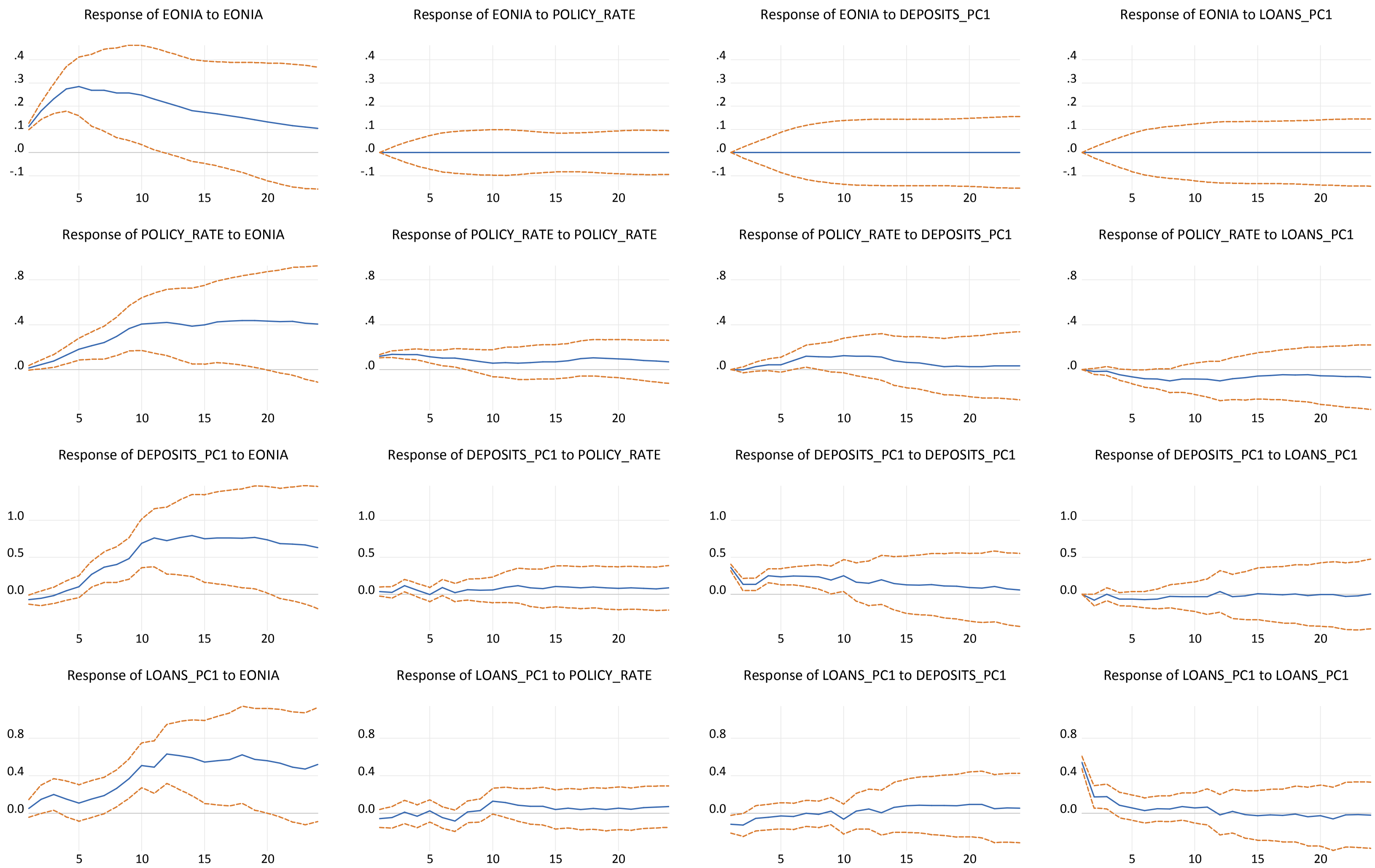


Note: PC1 refers to the first principal component of the respective series and POLICY_RATE refers to the key policy rate of the Central Bank of Hungary. Shocks are identified via a Cholesky decomposition. The dashed lines report the 95% confidence intervals. The x-axis reports months, the y-axis percentage point changes.

In the case of Romania the VAR model again includes the domestic monetary policy rate (figure 3.9). According to the impulse response functions, a shock in the EONIA rate triggers a positive and sizable interest rate response of domestic monetary policy, which reaches its maximum after approximately 10 months. This result indicates that – on the contrary to Hungary – Romanian monetary policy has closely followed euro area monetary policy in the period under investigation, either because the country was at a same stage of its business cycle as the euro area, or because Romanian monetary policy can be characterised as 'fear of floating' (i.e. despite claiming that the exchange rate is allowed to float, in practice it is stabilised through the active use of interest rate policy; see Calvo and Reinhart, 2002). The responses of the principal component series of deposit and lending rates becomes significant earlier than in the case of Bulgaria but later compared to Hungary, after approximately 6 and 8 months. The magnitude of the peak response amounts to 0.79 and 0.63 percentage points and is reached after 14 and 12 months, respectively. Excluding the domestic monetary policy rate from the VAR estimation yields very similar but slightly less pronounced responses of the deposit and lending rates.

The VAR model for Romania also indicates the magnitude of pass-through of a domestic monetary policy shock to euro deposit and lending rates (figure 3.9, second column). A domestic monetary policy shock does not seem to transmit to a large extent into domestic euro retail rates, as the impulse responses are again small and mostly insignificant.

Figure 3.9: Romania: Impulse Response Functions



Note: PC1 refers to the first principal component of the respective series and POLICY_RATE refers to the key policy rate of the National Bank of Romania. Shocks are identified via a Cholesky decomposition. The dashed lines report the 95% confidence intervals. The x-axis reports months, the y-axis percentage point changes.

The forecast error variance decompositions (tables 3.5 and 3.6) confirm the results obtained through the impulse response analysis. After 24 months, more than 50% of the degree of variance of Bulgarian euro retail rates, more than 70% of the degree of variance of Hungarian euro retail rates, and almost 90% of the degree of variance of Romanian euro retail rates can be explained by EONIA shocks. On the contrary, the degree of variance of Hungarian and Romanian euro retail rates that can be explained by domestic monetary policy shocks is almost negligible, especially for Romania.

Table 3.5: Forecast error variance decomposition due to EONIA shock

Country	*PC1 deposit rates*	*PC1 lending rates*
Bulgaria	55.7%	51.2%
Hungary	94.5%	76.0%
Romania	89.9%	89.4%

Note: Percentage of the variance of the respective variables explained by an EONIA shock after twenty-four months.

Table 3.6: Forecast error variance decomposition due to domestic monetary policy shock

Country	*PC1 deposit rates*	*PC1 lending rates*
Hungary	5.6%	12.2%
	(7 months)	(12 months)
Romania	2.1%	3.5%
	(12 months)	(11 months)

Note: Percentage of the variance of the respective variables explained by domestic monetary policy shock at peak.

3.5 Robustness checks

In order to test the validity of the results, a number of robustness checks are undertaken. One potential issue could arise from the test statistic used for the analysis of the results described in subsections 3.4.1 and 3.4.2, which implies 200 independent time series. However, the treatment of those 200 interest rate statistics as independent from each other can be questioned on the grounds that they might be correlated at least at country level, and potentially even across countries. This creates a multiple testing problem, where the confidence interval and the significance thresholds need to be adjusted accordingly in order to avoid a type I error (i.e. a false positive result).

In the inference of cointegration of an individual interest rate statistic in subsections 3.4.1 and 3.4.2 two conditions needed to be fulfilled: First, the series passes the bounds test for cointegration developed by Pesaran et al. (2001) at least at the 5% significance level, using the critical values provided by Narayan (2004) for small data samples. Second, the coefficient of the long-run pass-through of the EONIA (or domestic monetary policy rate) to the respective interest rate statistics is significant at least at the 5% level.

Adjusting the critical values of the bounds test for cointegration developed by Pesaran et al. (2001) (the first condition) for multiple statistics is unfortunately not straightforward and subject to future research. However, I am able to adjust the critical values of the coefficient of the EONIA in the long-run equation (the second condition). More specifically, I adjust the confidence interval through the Bonferroni correction, which is one of the more conservative approaches to correct for multiple testing (for details see e.g. Sokal and Rohlf, 1995). With the Bonferroni correction the confidence interval, CI, is adjusted through the following formula:

$$CI = 1 - \frac{\alpha}{m} \tag{3.6}$$

where the initial significance level, α, is divided by the number of multiple tests undertaken, m. Since 200 time series are tested, the long-run coefficient of the EONIA therefore needs to be significant at the 0.025% level, leading to a Bonferroni-adjusted confidence interval of 99.975%.[20].

While testing 200 time series with unadjusted critical values suggested that 74 time series, or 37% were cointegrated with the EONIA, the result does not change significantly when adjusting the critical values for multiple testing through the Bonferroni correction. In fact, 60 out of 200, or 30% of the euro retail rates in CESEE countries still appear to be cointegrated with the EONIA, even when applying a conservative significance level of 0.025% (see table 3.7). The medians of the estimated coefficients are also broadly unchanged when adjusting the sample for multiple testing. Overall, this confirms the robustness of the results obtained earlier.

Adjusting the critical values of the interest rate pass-through of domestic monetary policy to euro retail rates by the same method with a significance level of 0.07%[21] suggests that 22.9% of all interest rate statistics are cointegrated with domestic monetary policy rates, compared to the result of 32.9% when using unadjusted critical values (see table 3.4). Therefore, it can be concluded that after adjusting for multiple comparisons, the pass-through of euro area monetary policy still appears stronger than the pass-through of domestic monetary policy.

As a robustness check of the results of the VAR analysis, the order of the variables in the Cholesky decomposition is switched, so the first principal direction of lending rates enters the model before the first principal direction of deposit rates, i.e. lending rates are regarded as more exogenous compared to deposit rates. The results of the robustness check are practically

[20] The Dunn–Sidak correction, another method of correcting for multiple testing through the formula $1-(1-\alpha)^{\frac{1}{m}}$, yields a very similar adjusted confidence interval of 99.9744%.

[21] This is derived by $\frac{5\%}{70}$, as the total number of interest rate statistics in countries with independent monetary policy amounts to 70.

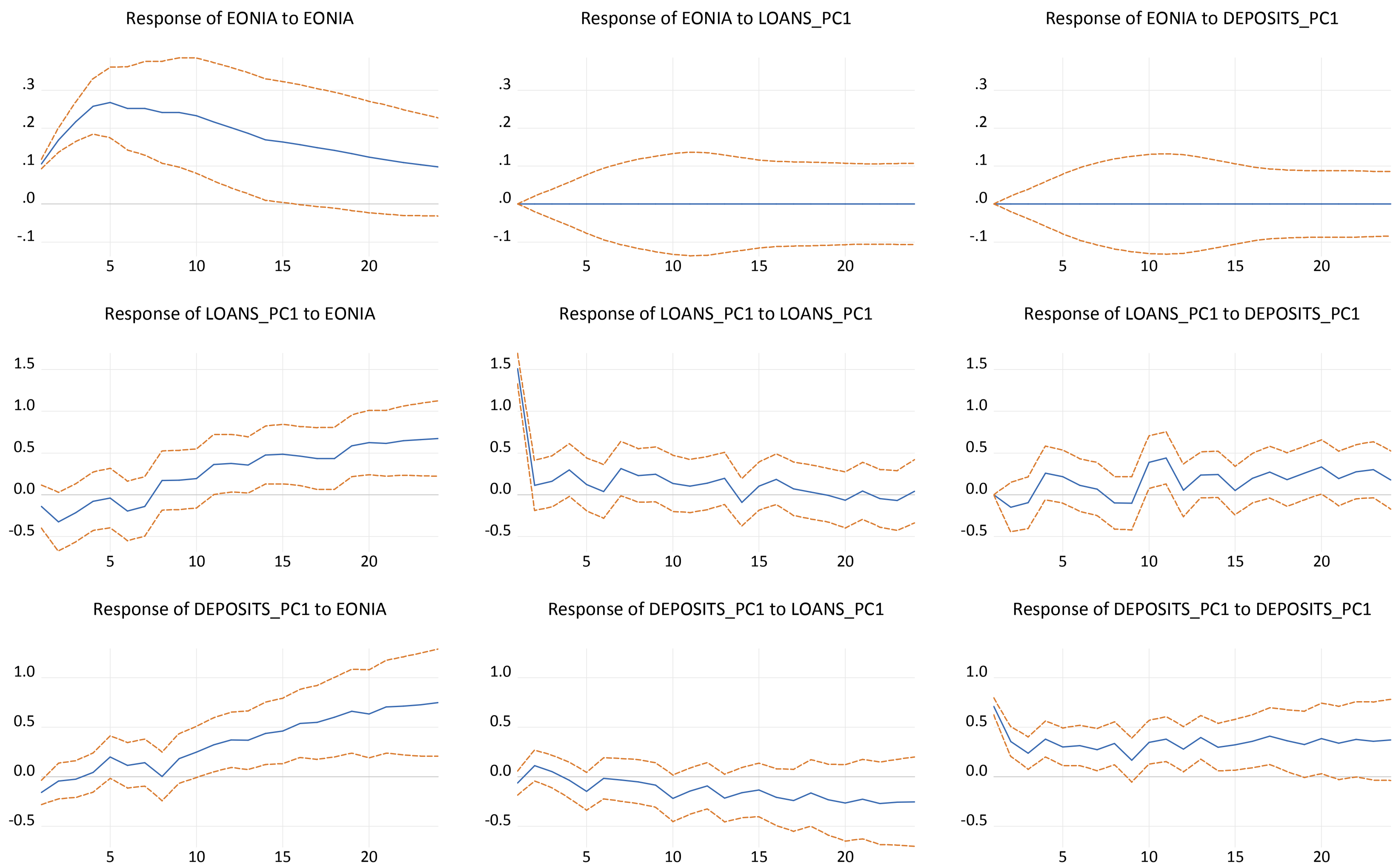
Table 3.7: Interest rate pass-through by country adjusted for multiple testing

Country	*Cointegrated with unadjusted critical values*	*Cointegrated with adjusted critical values*
Albania	35.7%	10.7%
Bosnia and Herzegovina	19.2%	11.5%
Bulgaria	28.8%	26.9%
Croatia	56.1%	53.7%
Hungary	50%	50%
North Macedonia	45.5%	18.2%
Romania	37.5%	37.5%
Serbia	25.0%	25.0%
Sum	37.0%	30%

Note: A series is considered to be cointegrated with the EONIA if the test by Pesaran et al. (2001) with the critical values provided by Narayan (2004) yields a conclusive result at least at the 5% significance level, and if the long-run coefficient of the EONIA is significant at least at a significance level of 5% (second column) or 0.025% (third column).

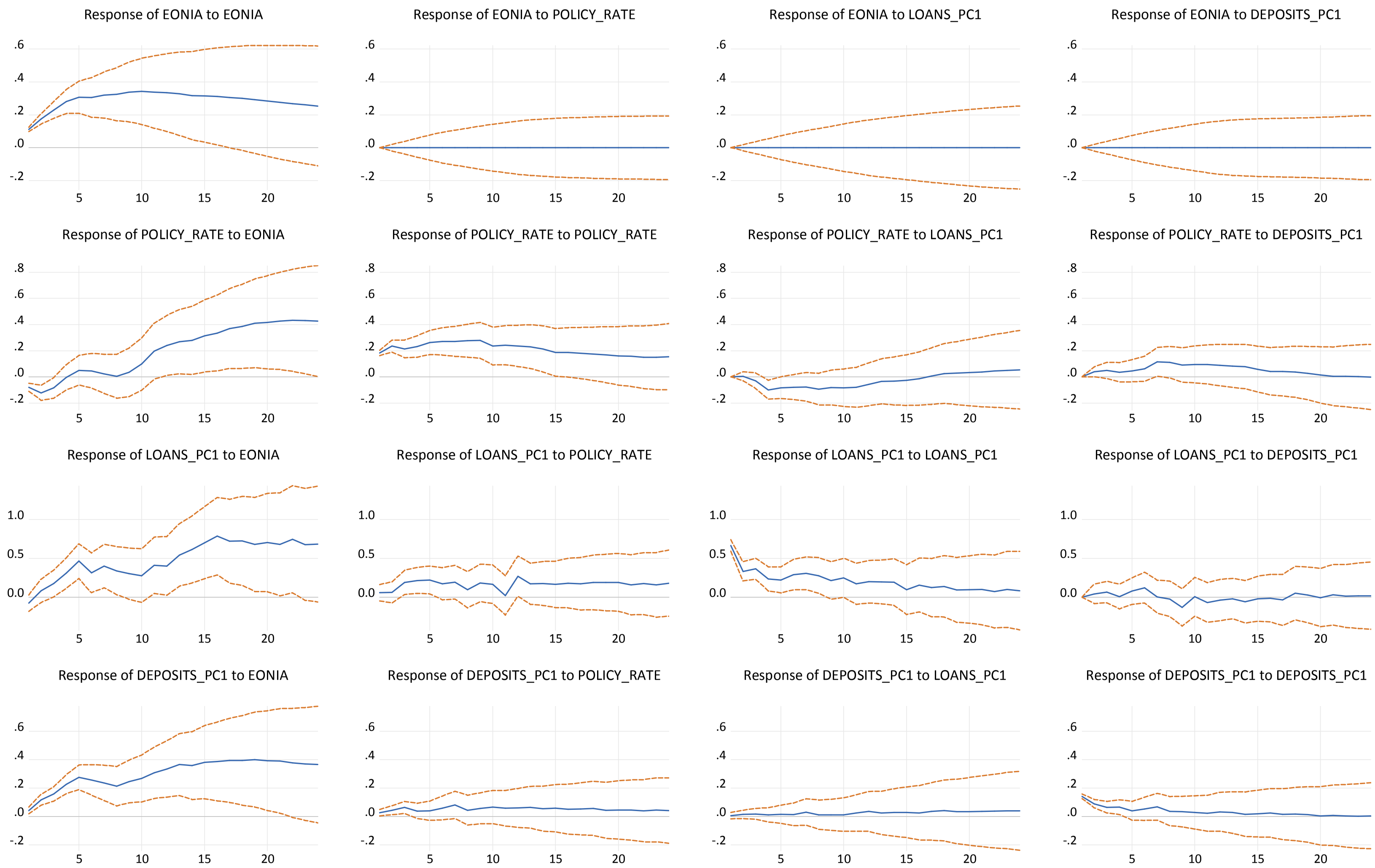
unchanged compared to the results obtained from the previous ordering (see figures 3.10, 3.11 and 3.12).

Figure 3.10: Bulgaria: Impulse Response Functions – robustness check



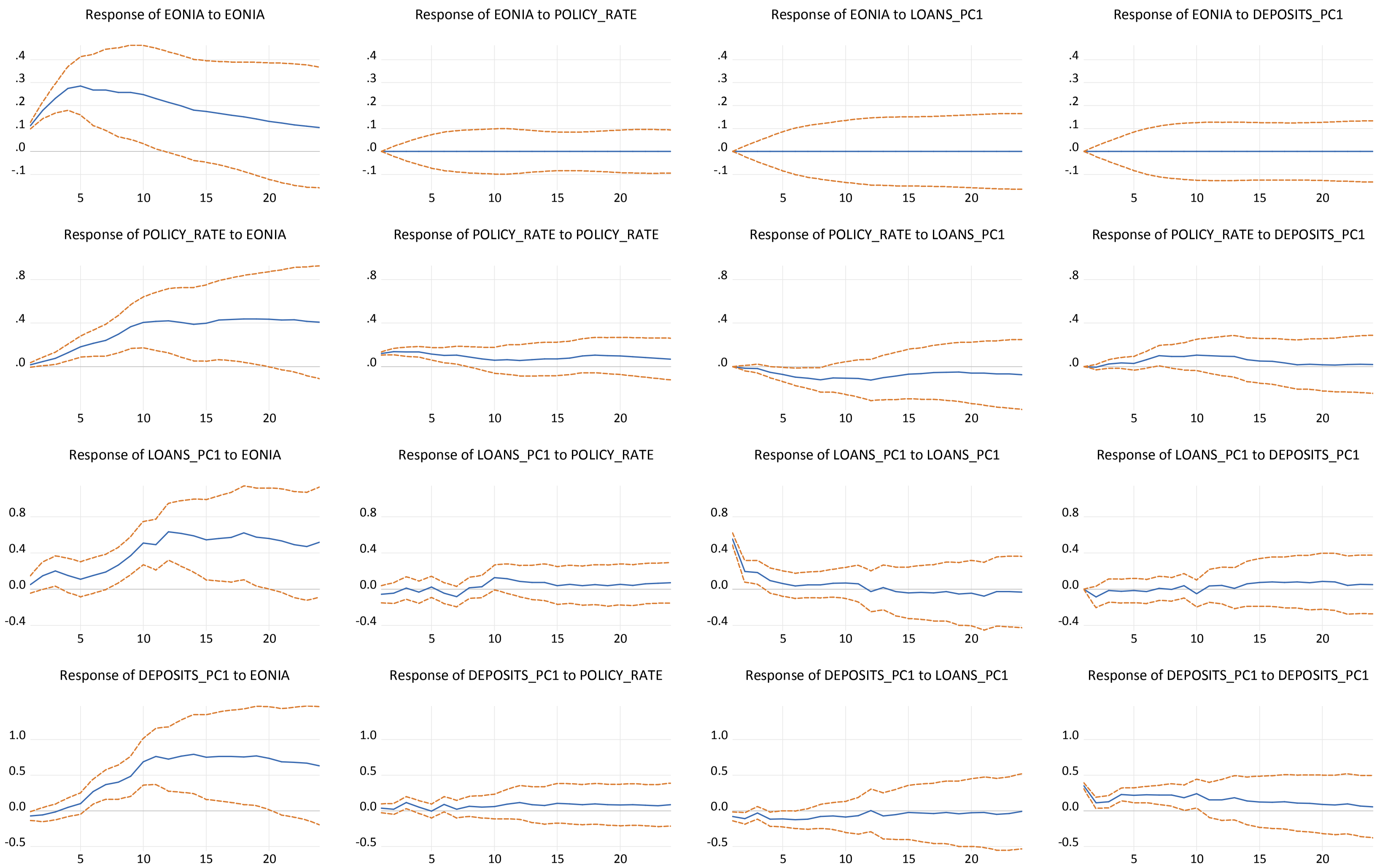
Note: PC1 refers to the first principal component of the respective series. Shocks are identified via a Cholesky decomposition. The dashed lines report the 95% confidence intervals. The x-axis reports months, the y-axis percentage point changes.

Figure 3.11: Hungary: Impulse Response Functions – robustness check



Note: PC1 refers to the first principal component of the respective series and POLICY_RATE refers to the key policy rate of the Central Bank of Hungary. Shocks are identified via a Cholesky decomposition. The dashed lines report the 95% confidence intervals. The x-axis reports months, the y-axis percentage point changes.

Figure 3.12: Romania: Impulse Response Functions – robustness check



Note: PC1 refers to the first principal component of the respective series and POLICY_RATE refers to the key policy rate of the National Bank of Romania. Shocks are identified via a Cholesky decomposition. The dashed lines report the 95% confidence intervals. The x-axis reports months, the y-axis percentage point changes.

3.6 Conclusion

This paper provides a comprehensive analysis of the interest rate pass-through of euro area monetary policy to euro retail rates outside the euro area, and compares it with the pass-through of domestic monetary policy to euro retail rates. More specifically, two questions are addressed: First, does a long-run relationship between euro area monetary policy and euro retail rates in unofficially euroised economies exist? Second, if such a relationship exists, to what extent can domestic monetary policy influence the euro part of the interest rate channel?

The results suggest that in the long run, more than one third of all euro retail rates in euroised CESEE countries are linked to euro area monetary policy, and that a higher share of deposit rates compared to lending rates is cointegrated with the EONIA . Across countries, some evidence can be found that the percentage of cointegrated retail rates is higher for EU members compared to non-EU members.

The estimated coefficients of the long-run pass-through suggest that euro retail rates in CESEE adjust excessively to a change in the EONIA, which is a puzzling result and contrary to both the interest rate pass-through within the euro area as well as the pass-through of domestic monetary policy in CESEE. With regard to the speed of adjustment, around 20% of a deviation of the long-run relationship between the respective retail rate and the EONIA is corrected within the next month, with considerable heterogeneity across countries.

Investigating the interdependence with domestic monetary policy suggests that euro retail rates are more likely to be cointegrated with the EONIA than with the domestic monetary policy rate. Also, if a long-run relationship between the euro retail rate and the domestic monetary policy rate exists, the pass-through of a domestic monetary policy change to the respective retail rate is incomplete. This suggests that domestic central banks in countries with independent monetary policy can only partially control the 'euro part' of the interest rate channel, which seems to be especially the case for CESEE EU members.

Additional evidence from a VAR analysis covering a subset of three countries (Bulgaria, Hungary and Romania) confirms the significant long-run relationship between the EONIA and euro retail rates, with the transmission of an EONIA shock materialising fully after between 12 and 20 months. Furthermore, the results suggest that the effect of a domestic monetary policy shock on euro retail rates is negligible and mostly insignificant.

The results of this paper are particularly relevant for CESEE countries with independent monetary policy, as they suggest a loss in monetary policy effectiveness for financially euroised countries. The question arises to what extent monetary policy in those CESEE countries can lean against the 'euro area wind', when it can only partially control parts of the domestic interest rate channel. For countries with a fixed exchange rate regime linked to the euro, the consequences seem to be less dramatic given that domestic financial conditions are anyway linked to euro area monetary policy. The direct interest rate pass-through to euro retail rates might however constitute an additional channel of the transmission of euro area monetary policy shocks.

While this paper provides evidence of a clear link between euro area monetary policy and euro retail rates in CESEE and finds that domestic central banks can only partially control the 'euro part' of the interest rate channel, a number of questions remain for future research: First, what determines the interest rate pass-through and the estimated coefficients, especially with respect to the excessive adjustment of euro retail rates to the EONIA? Second and related to the previous question, as the analysis focuses on the post-crisis low interest rate-period, do the findings of this paper apply to 'normal' times as well? Third, what can domestic (monetary) policy do to gain more control over the 'euroised' interest rate channel? Last but not least, are the findings of this paper confined to the euro area and the closely linked CESEE region, or do they also apply to US monetary policy and dollarised countries across the world?

3.7 Annex – list of euro retail rates

1 Albania	
1.1 Deposit rates	1.2 Lending rates
1.1.1 All sectors	1.2.1 Households
Current account	Total
Demand deposits	Overdraft
1 Month deposits	**Consuming of non durable goods**
3 Months deposits	***Consuming of durable goods***
6 Months deposits	Loans for house purchase
12 Months deposits	Other lending (loans for other purposes)
24 Months deposits	1.2.2 NFCs
36 Months deposits	Total
60 Months deposits	Overdraft
	Working capital
	Equipment loans
	Real estate loans
	1.2.3 all sectors
	Up to 6 months
	6 months-1 year
	1-3 years
	3-5 years
	1.2.3 All sectors
	Average interest rate (for all maturities)

2 Bosnia and Herzegovina	
2.1 Deposit rates	2.2 Lending rates
2.1.1 Households	2.2.1 Households
Up to 1 year maturity (indexed)	**House purchases: Floating rate and up to 1 year IRF (indexed)**
> 1 and up to 2 years maturity (indexed)	House purchases: Floating rate and up to 1 year IRF; maturity > 1 year (indexed)
> 2 years maturity (indexed)	House purchases: > 1 and up to 5 years IRF (indexed)
Up to 1 year maturity	House purchases: > 5 years IRF (indexed)
> 1 and up to 2 years maturity	House purchases: > 10 years IRF (indexed)
> 2 years maturity	Other purposes: Floating rate and up to 1 year IRF (indexed)
2.1.2 NFCs	Other purposes: Floating rate and up to 1 year IRF and maturity > 1 year (indexed)
Overnight	Other purposes: > 1 and up to 5 years IRF (indexed)
Up to 1 year maturity (indexed)	Other purposes: > 5 years IRF (indexed)
> 1 and up to 2 years maturity (indexed)	2.2.2 NFCs
Overnight (indexed)	Up to EUR 0.25mn: Floating rate and up to 1 year IRF (indexed)
	Up to EUR 0.25mn: Floating and up to 1 year IRF; maturity > 1 year (indexed)
	Up to EUR 0.25mn: > 1 and up to 5 years IRF (indexed)
	Up to EUR 0.25mn: > 5 years IRF (indexed)
	From EUR 0.25mn to EUR 1mn: Floating rate and up to 1 year IRF (indexed)
	EUR 0.25mn to EUR 1mn: Floating + up to 1 year IRF; maturity > 1 year (indexed)
	> EUR 1mn: Floating rate and up to 1 year IRF (indexed)

Note: Interest rates depicted in **bold**/*italics* have been found to be cointegrated at least at the 5% significance level with the **EONIA**/*domestic monetary policy rate.*

3 Bulgaria	
3.1 Deposit rates	3.2 Lending rates
3.1.1 Households	3.2.1 Households
Agreed maturity: short-term	Consumer loans IRF
Agreed maturity: > 1 day up to 1 month	Consumer loans: up to 1 year IRF
Agreed maturity: > 1 up to 3 months	Consumer loans: > 1 and up to 5 years IRF
Agreed maturity: > 3 up to 6 months	House purchases IRF
Agreed maturity: > 6 up to 12 months	**House purchases: up to 1 year IRF**
Agreed maturity: > 1 up to 2 years	Other loans IRF
Agreed maturity: > 2 years	Other loans: up to 1 year IRF
Time deposits: short-term	Consumer loans: all maturities
Time deposits: > 1 day up to 1 month	Consumer loans: up to 1 year maturity
Time deposits: > 1 up to 3 months	Consumer loans: > 1 and up to 5 years maturity
Time deposits: > 3 up to 6 months	Consumer loans: > 5 years maturity
Time deposits: > 6 up to 12 months	House purchases: all maturities
Time deposits: > 1 up to 2 years	House purchases: > 1 and up to 5 years maturity
Time deposits: > 2 years	House purchases: > 5 and up to 10 years maturity
3.1.2 NFCs	House purchases: > 10 years maturity
Agreed maturity: short-term	Other loans: all maturities
Agreed maturity: > 1 day up to 1 month	Other loans: > 1 and up to 5 years maturity
Agreed maturity: > 1 up to 3 months	Other loans: > 5 years maturity
Agreed maturity: > 3 up to 6 months	3.2.2 NFCs
Agreed maturity: > 6 up to 12 months	Overall IRF
Agreed maturity: > 2 years	Up to 1mn EUR IRF
Time deposits: short-term	Up to 1mn EUR IRF
Time deposits: > 1 day up to 1 month	> 1mn EUR IRF
Time deposits: > 1 up to 3 months	> 1mn EUR IRF
Time deposits: > 3 up to 6 months	Total: all maturities
Time deposits: > 6 up to 12 months	Up to 1 year maturity
	> 1 and up to 5 years maturity
	> 5 years maturity

4 Croatia	
4.1 Deposit rates	4.2 Lending rates
4.1.1 Households	4.2.1 Households
Time deposits	**Consumer and other loans: short-term**
Time deposits: short-term	Consumer loans (indexed)
Time deposits: long-term	**Consumer and other loans: short-term (indexed)**
Time deposits (indexed)	**Consumer and other loans: long-term (indexed)**
Time deposits: short-term (indexed)	House purchases: short-term
Time deposits: long-term (indexed)	House purchases (indexed)
4.1.1 NFCs	**Loans for other purposes (indexed)**
Time deposits	4.2.2 NFCs
Time deposits: short-term	**Total short-term (indexed)**
Time deposits: long-term	**Total long-term (indexed)**
Time deposits (indexed)	**Up to HRK 2mn (indexed)**
Time deposits: short-term (indexed)	**Up to HRK 2mn: short-term (indexed)**
Time deposits: long-term (indexed)	**Up to HRK 2mn: long-term (indexed)**
	From HRK 2mn to HRK 7.5mn (indexed)
	From HRK 2mn to HRK 7.5mn: short-term (indexed)
	From HRK 2mn to HRK 7.5mn: long-term (indexed)
	Loans > HRK 7.5mn (indexed)
	Loans > HRK 7.5mn: short-term (indexed)
	Loans > HRK 7.5mn: long-term (indexed)
	Total short-term
	Total long-term
	Up to HRK 2mn
	Up to HRK 2mn: short-term
	Up to HRK 2mn: long-term
	From HRK 2mn to HRK 7.5mn
	From HRK 2mn to HRK 7.5mn: short-term
	From HRK 2mn to HRK 7.5mn: long-term
	Loans > HRK 7.5mn
	Loans > HRK 7.5mn: short-term
	Loans > HRK 7.5mn: long-term

Note: Interest rates depicted in **bold**/*italics* have been found to be cointegrated at least at the 5% significance level with the **EONIA**/*domestic monetary policy rate*.

5 Hungary	
5.1 Deposit rates	5.2 Lending rates
5.1.1 Households	5.2.1 Households
Overnight	Bank overdrafts
Up to 1 year	5.2.2 NFCs
> 1 year and up to two years	Bank overdrafts
5.1.2 NFCs	**Floating rate and up to 1 year IRF**
Overnight	Floating rate and up to 1 year IRF; secured
Up to 1 year	***Floating rate and up to 1 year IRF; > 1mn euro***
	Floating rate and up to 1 year IRF; up to 1mn euro; secured
	Floating rate and up to 1 year IRF; up to 0.25mn EUR
	Floating rate and up to 1 year IRF; > 0.25mn EUR up to 1mn EUR
	Original maturity > 1 year

6 North Macedonia	
6.1 Deposit rates	6.2 Lending rates
6.1.1 Households	6.2.1 Households
All deposits	All loans
Overnight deposits	Overdraft and credit card loans
6.1.2 NFCs	All loans (indexed)
All deposits	6.2.2 NFCs
Overnight deposits	All loans
All deposits (indexed)	Overdraft and credit card loans
	All loans (indexed)

7 Romania		
7.1 Deposit rates		7.2 Lending rates
7.1.1 Households	7.1.3 All sectors	7.2.1 Households
Time deposits	Time deposits	**All loans**
Original maturity up to 1 year		***Consumer loans: floating rate and up to 1 year IRF***
Original maturity > 1 and up to 2 years		**Consumer loans: > 1 and up to 5 years IRF**
Original maturity > 2 years		7.2.2 NFCs
7.1.2 NFCs		**All loans**
Time deposits		*Other loans up to EUR 1mn; floating rate and up to 1 year IRF*
Original maturity up to 1 year		Other loans > EUR 1mn; floating rate and up to 1 year IRF
Original maturity > 1 and up to 2 years		7.2.3 All sectors
Original maturity > 2 years		**All loans**

8 Serbia		
8.1 Deposit rates	8.1.2 NFCs	8.2 Lending rates
8.1.1 Households	Up to 1 year (euro and indexed)	8.2.1 Households
Up to 1 year (euro and indexed)	> 1 and to 2 years (euro and indexed)	*Total loans (euro and indexed)*
> 1 and up to 2 years (euro and indexed)	***> 2 years (euro and indexed)***	8.2.2 NFCs
> 2 years (euro and indexed)	Total (euro and indexed)	***Total loans (euro and indexed)***
Total (euro and indexed)	8.1.3 All sectors	8.2.3 All sectors
	Total deposits (euro and indexed)	***Total loans (euro and indexed)***

Note: Interest rates depicted in **bold**/*italics* have been found to be cointegrated at least at the 5% significance level with the **EONIA**/*domestic monetary policy rate.*

Chapter 4

Business cycle synchronization between the Western Balkans and the European Union

Authors: Isabella Moder, B.A. M.A. and Dr. Antje Hildebrandt

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Abstract

This paper investigates the degree of business cycle synchronization (BCS) between the Western Balkans and the European Union. While the degree of synchronization had been low or even negative before 2000, business cycles have clearly converged ever since. Since 2010, in particular, BCS with the EU has been high for all Western Balkan countries. In addition, this paper empirically tests which factors were responsible for the convergence process. Using a period-by-period correlation index for the time span from 1994 to 2013, we find that foreign trade is the most important positive contributor to business cycle convergence. Although fiscal differences are associated with negative BCS in the same year, our results suggest that they have a positive influence in subsequent periods. In contrast, we find that financial flows lead to business cycle decoupling. The same relationship applies to remittances, where we find that remittances from the EU to the Western Balkans behave similarly to financial flows.

4.1 Introduction

Four of the seven Western Balkans economies[1] follow a monetary policy that is very closely linked to that of the euro area, either through the use of the euro as their official currency (Montenegro and Kosovo) or through fixed exchange rate regimes (euro-based currency board in Bosnia and Herzegovina, euro peg in North Macedonia). Although Albania, Croatia and Serbia operate under a managed, or rather free-floating, exchange rate regime, respectively, in practice they also face substantial monetary policy constraints, given the high asset and liability euroisation in their banking systems.

According to the optimum currency area (OCA) theory developed by Mundell (1961), McKinnon (1963) and Kenen (1969), one prerequisite for the efficient use of a common currency or for following the monetary policy of another country or currency area is that the business cycles of the countries involved are sufficiently synchronized. Otherwise – if output patterns are divergent – monetary policy cannot bring about optimal reactions for each country at the same time. When evaluating the economic costs of a country's lack of, or tight constraint on, independent monetary policy, it is therefore important to know the degree of BCS of the respective country with the euro area (and, in a broader, forward-looking sense, with prospective euro area countries).

This paper is structured as follows. Section 4.2 provides an overview of the relevant literature. In section 4.3, we investigate the degree of BCS between the Western Balkan countries and the EU-25 aggregate to find out which countries exhibit higher or lower BCS and whether a convergence process can be identified. Section 4.4 discusses the main drivers and transmission channels of BCS and the choice of explanatory variables. Subsequently, we employ a regression model for analyzing what factors drive the development of BCS over time. Finally, in section 4.5, we draw conclusions from our analysis.

4.2 Related literature

The degree of BCS between countries is driven by two main factors: on the one hand, it depends on the presence and dominance of common transnational shocks over idiosyncratic (i.e. country-specific) shocks. On the other hand, idiosyncratic shocks by themselves can be spread across countries through certain transmission channels. Drivers of BCS are a well-researched topic in the empirical literature. Focusing mainly on industrial countries (e.g. Frankel and Rose, 1998; Darvas et al., 2005; Artis et al., 2008; Inklaar et al., 2008), the literature finds that trade is one of the main transmission channels of BCS. The evidence for other factors like industrial specialization or fiscal and monetary policy, which have been proposed by theoretical literature

[1] In the regional definition of the Western Balkans, we follow the IMF (see `https://www.imf.org/external/pubs/ft/reo/2015/eur/eng/pdf/erei_sr_030915.pdf`) and include Albania, Bosnia and Herzegovina, Croatia, North Macedonia, Kosovo, Montenegro and Serbia.

and investigated empirically, is mixed (de Haan et al., 2008). Prompted by the accession of ten Central, Eastern and Southeastern European (CESEE) countries to the EU in the 2010s, several studies also investigated the patterns and drivers of synchronization between the industrialized and transition countries in Europe (e.g. Artis et al., 2008; Babetskii, 2005; Crespo Cuaresma et al., 2011). Their results suggest that the determinants of BCS between emerging markets and industrialized countries seem to be similar to those that dominate BCS between industrial countries. Distinguishing between industrial and developing countries, Calderon et al. (2007) find, however, that the impact of trade integration is higher for BCS between industrial countries than for BCS between developing countries or between 'mixed' pairs. Another strand of business cycle literature is concerned with the endogeneity of OCA criteria, as pointed out first by Frankel and Rose (1998). This endogeneity implies that a country is more likely to fulfill the OCA criteria after having joined a currency union than before accession (see Gächter and Riedl, 2014, for a discussion).

Only a few studies so far have covered the degree of BCS of the Western Balkan economies with the euro area or the European Union (EU). Velickovski (2013) investigates shock synchronization between selected Western Balkan countries (Albania, Croatia, North Macedonia and Serbia) and the euro area by two approaches. First, he calculates correlation coefficients between the four Western Balkan countries and the euro area and finds that output correlation is highest between Croatia and the euro area and between North Macedonia and the euro area whereas correlation between Albania and Serbia vis-à-vis the euro area is comparably low. Second, the author estimates time-varying coefficients of shock symmetry (both supply and demand shocks) between the Western Balkans and the euro area in a vector autoregressive (VAR) framework. Gouveia (2014) uses a data set of eight Balkan economies (including Croatia, North Macedonia and Serbia from the Western Balkans) and compares various measures of trade intensity and BCS between these countries and the euro area average. For Croatia and Serbia, the degree of BCS is found to be moderate and well below the average of intra-euro area correlation, while North Macedonia exhibits higher output synchronization with the euro area. With respect to output volatility, the author concludes that the volatility of business cycles is substantially higher in the Balkan countries than in the euro area.

To the best of our knowledge, there is no study on BCS that covers all Western Balkan countries nor one that investigates the determinants of business cycle convergence between the Western Balkans and the euro area or the EU. This paper aims to fill this gap in the empirical literature, i.e. to identify the degree of synchronization between the Western Balkans and the EU, to find how BCS has changed since the beginning of transition and to identify the main drivers of business cycle convergence.

More precisely, our analysis covers the EU-25, i.e. all EU Member States with the exception of Croatia, which joined the EU only in mid-2013 (and is covered here in the group of Western Balkan countries), and Denmark and the United Kingdom, which have been granted an opt-out clause and are thus not required to participate in Stage Three of Economic and Monetary Union

(EMU). All other EU Member States are required to join the euro area once they fulfill the convergence criteria, as 19 of them have done so far. We investigate BCS between the Western Balkans and the current euro area as well as all EU Member States (except Croatia) that are obliged to adopt the euro and thus are the relevant counterpart for the Western Balkans' future business cycle convergence and economic integration. A further advantage of this approach is that, by doing so, we obtain a larger data set, which is particularly helpful in underpinning the robustness of our results regarding the determinants of BCS.

In our analysis, we make use of some recent advances in the business cycle literature. First, we use a new time-varying correlation index developed by Cerqueira and Martins (2009) and Cerqueira (2013) to analyze the convergence process on a yearly basis. As can be seen in the results in the subsequent sections, using a time-varying correlation index considerably improves the measurement of BCS, as synchronization varies greatly over time and this variation could not be captured by a single correlation coefficient spanning the whole observation period. Additionally, we obtain a panel data set instead of a cross-sectional sample for the regressions, which allows us to explore the full-time variability of the data. Most studies that identify determinants of BCS use a two-step instrumental variables approach; however, we employ the system Generalized Method of Moments (GMM) estimator developed by Blundell and Bond (1998) as it offers several advantages (see below). To identify the determinants of BCS, we start by looking at well-known factors identified in the literature (trade, specialization, fiscal policy) but extend the choice of variables by adding other potential channels that might be relevant especially for the Western Balkans. We incorporate common monetary policy and financial flows in the regressions. In addition, we include remittances as they constitute a large part of income in the Western Balkan countries, are less volatile than other financial flows to the region (Petreski and Jovanovic, 2013) and, as already argued by Barajas et al. (2012), constitute an important BCS channel. According to the World Bank, remittances amounted to more than 16% of GDP in Kosovo and to more than 10% of GDP in Bosnia and Herzegovina in 2013. The inflow of remittances was quite substantial also in Albania, Montenegro and Serbia (between about 6% to 8% of GDP). Only Croatia and North Macedonia registered inflows of below 4% in 2013.

4.3 Business cycle synchronization

4.3.1 Measurement issues

In the business cycle literature, a variety of indicators of economic activity – like GDP (in levels or first differences), industrial production or consumption measures – have been used to identify the cyclical component of economy activity (see Darvas and Szapáry, 2008, for a comparison and discussion). Since the Western Balkans underwent a process of de-industrialization particularly at the beginning of the transition period in the 1990s and since industrial production does not include all sectors of the economy, we do not consider industrial production a good indicator of economic activity. We therefore make use of annual GDP data (year on year), as quarterly GDP data do not provide long time series and are not even available for some of the countries covered. Instead of using growth rates, we use the logs of real GDP data in levels as these are better suited for heterogeneous samples of countries at different stages of economic development (see Gächter and Riedl, 2014).

To separate the cyclical component (i.e. the output gap) from the trend component (i.e. potential output), we use the Hodrick-Prescott (HP) filter as it is the standard method and easy to implement. Moreover, the resulting cyclical components are similar to those of the band-pass filter (Belke and Heine, 2007; de Haan et al., 2008). Following the Ravn-Uhlig rule, the smoothing parameter is set to 6.25 as proposed for yearly data. One drawback of the HP filter is that it delivers suboptimal results at the end of the sample (see e.g. Mise et al., 2005). To overcome this problem, we complement the time series for each country by forecasts from the IMF World Economic Outlook until 2019. The filtered cyclical components are tested for stationarity with the augmented Dickey-Fuller unit root test including a constant. With the exception of the cyclical component for Greece, all cyclical components in the sample are stationary at a 5% confidence level.

Most studies on this issue are affected by the nonavailability of a year-by-year index for BCS, which is why they had to investigate the topic on a cross-sectional basis. To account for at least some time variability, most studies use moving averages, sample period splits and other methods. The lack of year-by-year BCS indices can be overcome by using the period correlation index developed by Cerqueira and Martins (2009):

$$\rho_{ij,t} = 1 - \frac{1}{2}\left(\frac{d_{j,t} - \bar{d}_j}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}(d_{j,t} - \bar{d}_j)^2}} - \frac{d_{i,t} - \bar{d}_i}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}(d_{i,t} - \bar{d}_i)^2}}\right)^2 \tag{4.1}$$

Here, d_i and d_j denote any two time series and $\bar{d}_j$, $\bar{d}_i$ denote the respective averages over time. The index thus measures the correlation between d_i and d_j at each point in time ($t = 1, \ldots, T$). Taking the average of $\rho_{ij,t}$ equals the linear correlation index ρ_{ij} conventionally used in cross-sectional studies. The index is of an asymmetric nature with $max(\rho_{ij,t}) = 1$ and $min(\rho_{ij,t}) =$ 3–2T (see Cerqueira, 2013).

4.3.2 Data sample and descriptive results

GDP data (in euro) at constant prices are extracted from the IMF World Economic Outlook Database. Our country sample comprises 25 EU Member States (EU-25, i.e. all current members excluding Croatia and Denmark) and the Western Balkan countries (Albania, Bosnia and Herzegovina, Croatia, Kosovo[2], North Macedonia, Montenegro and Serbia). The business cycle

[2]Due to data limitations, Kosovo is included in the descriptive analysis of BCS but not in the empirical part.

measures are calculated for the maximum period from 1989 to 2013.[3]

Table 4.1 summarizes the yearly measures of correlation between the Western Balkans and the EU-25 aggregate for different subperiods.[4] When comparing the development of this synchronization index over the four subperiods (transition period, precrisis period, crisis peak and crisis aftermath), we can indeed identify a convergence process. Except in the crisis peak subperiod, the correlation index has clearly increased since the transition period. The different subperiods are characterized by specific global and country-specific shocks. During the transformation period from 1989 to 1995, the Western Balkans were marked by the break-up of former Yugoslavia, trade interruptions and the Balkan Wars. The latter also led to EU financial and economic sanctions against Montenegro and Serbia (World Bank, 2004). Also in this period, Albania started to ease its isolation policy. These country-specific shocks of the 1990s make the analysis of BCS of the Western Balkans a difficult task. After 1995, the region saw some economic and political stability but experienced transformational recession. In addition, the limited availability of GDP data for this period also puts limits on the interpretation of business cycle correlations.

From 2001 to 2008, BCS with the EU-25 increased in particular in Bosnia and Herzegovina, Croatia, North Macedonia and Montenegro. In this period, the EU integration process started in all Western Balkan countries, but it progressed at different speeds. Supposedly, closer ties with the EU and increasing trade relations have had a positive impact on business cycle comovements. At the Thessaloniki European Council summit in 2003, the EU confirmed that the future of the Western Balkans lies within the EU and granted potential EU candidate status to the Western Balkan countries (European Council, 2003). For North Macedonia, however, an agreement on trade and trade-related matters had already entered into force in 2001, and the Stabilisation and Association Agreement (SAA) became effective in 2004. In the remaining countries of our sample, the SAA became effective at a later stage (European Commission, 2015).

For Albania, Kosovo and Serbia, synchronization with the EU-25 was much weaker and more volatile. Kosovo strongly decoupled from the EU-25 aggregate in 2006, and Albania in 2007. Kosovo showed pronounced cyclical movements over this period, posting very high GDP growth (more than 8%) in 2007, while growth was much lower in the years before 2007 and afterward. This helps explain the large fluctuations of Kosovo's BCS. Albania experienced no strong cycle movement over time. However, in 2007, growth was relatively weak in Albania compared with a booming EU-25, which is mirrored in a low co-movement of the respective business cycles.

In 2009, the economic and financial crisis hit the EU countries hard, as reflected in the slump of the cyclical component of the EU-25 aggregate and eventually in the decoupling of the business

[3]For most Western Balkan countries, GDP data series start later than 1989.

[4]To be consistent with the empirical part of our paper, table 4.1 displays the aggregate of the EU-25 and not that of the euro area. However, synchronization between the Western Balkan countries and the euro area aggregate is very similar.

Table 4.1: Synchronization of the Western Balkans with the EU-25 aggregate in different subperiods

Country	*Transition period* (until 2000)	*Precrisis period* (2001 – 08)	*Crisis peak* (2009)	*Crisis aftermath* (2010 – 13)	*Overall*
Montenegro	*n/a*	0.860	–0.668	0.871	0.760
Croatia	0.401	0.765	–0.778	0.909	0.572
North Macedonia	0.473	0.669	–1.383	0.871	0.532
Bosnia and Herzegovina	0.313	0.720	–1.615	0.832	0.526
Serbia	–1.381	0.568	–0.869	0.839	0.181
Kosovo	*n/a*	0.195	–3.072	0.908	0.175
Albania	–0.073	0.287	–2.424	0.850	0.096

Source: Authors' calculations. Note: Countries are sorted by overall synchronization level. Data are available from 1989 for Albania, from 1992 for Croatia and North Macedonia, from 1998 for Bosnia and Herzegovina as well as Serbia, and from 2000 for Kosovo and Montenegro.

cycles of all Western Balkan countries from the EU-25. The Western Balkans also experienced some economic downturn, but it occurred later and was less pronounced than in the EU-25. According to Bonetto et al. (2009), the Western Balkan countries were partially protected from the economic and financial crisis of 2009 as they had a low exposure to international financial markets and the foreign banks active in the region were strongly capitalized. Albania and Kosovo even overcame the crisis years without dipping into a recession.

From 2010 to 2013, convergence of the business cycles of the EU-25 and the Western Balkans increased. Even the group of Western Balkan countries recording less synchronized cycles exhibited a high degree of synchronization in that period.

As discussed above, studies which also analyze the degree of BCS between the EU and the Western Balkans generally come to similar results: North Macedonia shows relative strong comovements with the EU whereas EU correlations with Albania and Serbia are comparatively low. Results for Croatia are ambiguous. Furthermore, Gouveia (2014) also shows that business cycle correlation has in general increased over time between the Western Balkan countries and the EU.

4.4 Determinants of business cycle synchronization

The result of our descriptive analysis, namely that the business cycles of the Western Balkans have clearly converged with the EU business cycle over the past 15 years, leads to the question which factors drove this convergence process. We therefore proceed to empirically test the determinants of BCS.

4.4.1 Empirical model

In order to investigate determinants of BCS between the EU and the Western Balkans, we use the following model:

$$Correl_{ij,t} = \alpha + \beta Correl_{ij,t-1} + Z'_{ij,t}\gamma + \mu_{ij} + \lambda_t + \nu_{ij,t} \tag{4.2}$$

where $Correl_{ij,t}$ denotes the bilateral correlation index of BCS between countries i and j in year t. As pointed out by Cerqueira (2013), one shortcoming of the index is that it is asymmetric, which could lead to biased results when used in regressions. Thus, for the purpose of the regressions, we use the following transformation developed by Cerqueira (2013):

$$\rho_{ij,t} = \frac{1}{2}\ln\left(\frac{1+\frac{\rho_{ij,t}}{2T-3}}{1-\rho_{ij,t}}\right) \tag{4.3}$$

that yields a nonbounded index with a symmetric support around 0 and a symmetric range between $-\infty$ and $+\infty$.

$Z'_{ij,t}$ is a matrix consisting of the potential determinants of BCS such as bilateral trade, asymmetry of production, fiscal differences, common monetary policy, FDI, bank flows and remittances. Data are available yearly in an unbalanced panel for 162 country pairs with a maximum time span from 1989 to 2013. To control for autocorrelation, the one-period lagged BCS $Correl_{ij,t-1}$ is included in the model. Additionally, we control for country-pair fixed effects μ_{ij} to account for specific unobservable country-pair factors, and time effects λ_t to account for common global shocks. Our main interest lies in the signs and magnitudes of vector γ, which indicates what drives BCS between the EU and the Western Balkans. While some authors (Imbs, 2006; Dées and Zorell, 2012) investigate the determinants of BCS with a system of equations to account for indirect effects as well, we focus on a single equation approach. However, future research on this topic might expand our approach into a multi-equation system.

Frankel and Rose (1998) pointed out that endogeneity plays an important role in the relationship between trade integration and BCS. They argue that countries with stronger trade integration and countries with similar output patterns are more likely to join a currency union, and joining a currency union in turn increases trade integration and business cycle correlation. Following this argument, several endogeneity issues with respect to BCS have been discussed in the literature, not only with respect to trade integration but also regarding financial integration (Grauwe and Mongelli, 2005) or remittance flows (Frankel, 2011). Most studies tackle the endogeneity issue by using two-step instrumental variable approaches. However, we follow Cerqueira and Martins (2009) and make use of a two-step system GMM estimator developed by Blundell and Bond (1998), which offers several advantages. First, it allows us to draw a large number of instruments from within the data set by instrumenting endogeneous variables with their own lagged values. Second, additional time-invariant instruments can be included in the regression (in contrast to the difference GMM estimator used by Arellano and Bond, 1991). Additionally, the present data set is a small T, large N panel data set, for which the estimator is well suited as it controls for the dynamic panel bias (Roodman, 2009).

4.4.2 Choice of explanatory variables

In this section we briefly describe the set of variables identified in the literature as important determinants of business cycle convergence. There is a broad consensus that trade integration is an important driver of BCS. It is argued that the elimination of trade barriers results in a stronger transmission of demand shocks and eventually higher business cycle co-movement (Frankel and Rose, 1998). For testing the impact of trade on BCS between the EU-25 and the Western Balkans, we use the IMF's database 'Direction of Trade Statistics' (DOTS). DOTS provides data for the calculation of bilateral trade intensities between individual countries of the EU-25 and the Western Balkans. We calculate bilateral trade intensity as follows:

$$BTI_{ij,t} = \left(\frac{Exports_{ij,t} + Imports_{ij,t}}{Trade_{i,t} + Trade_{j,t}} \right) \tag{4.4}$$

$BTI_{ij,t}$ is obtained by dividing the sum of bilateral trade flows between i and j by the sum of total trade of country i, $Trade_{i,t}$, $Trade_{j,t}$ and j, respectively, at time t. In the literature, it is common to use either this approach or to scale the sum of total trade by the sum of both countries' GDP. However, as argued in Romer and Frankel (1999), countries with higher income are possibly more active in trading. Therefore, using bilateral trade as a percentage of GDP as a determinant of BCS could lead to biased results as our country sample is very heterogeneous with respect to income levels.

As discussed above, the endogeneity of the relationship between trade integration and BCS has been widely acknowledged in the literature. To tackle this issue, we follow e.g. Calderon et al. (2007), Frankel and Rose (1997) or Romer and Frankel (1999) by instrumenting bilateral trade intensity with a gravity model variable. As an instrument, we use the log of the distance between the capital cities of each country pair, which is taken from the CEPII's database on geographical variables.[5]

Another commonly used variable is the degree of economic specialization, which could affect the impact of trade on BCS. In this case, we follow Calderon et al. (2007) and Barajas et al. (2012) and calculate a simple asymmetry of production index

$$ASP_{ij,t} = \frac{1}{k}\sum_{k=1}^{3} |\nu_{ki,t} - \nu_{kj,t}| \tag{4.5}$$

where $ASP_{ij,t}$ denominates the mean[6] of absolute differences of the value added share in the total production of each country ν of sector k for each country pair in each year. Like Barajas et al. (2012), we focus on three sectors, namely agriculture, industry and services.

Fiscal policy is another possible determinant of BCS. On the one hand, fiscal policy can be used as a stabilizer at the national level to help smoothing the business cycle but on the other hand, fiscal policy can by itself be the source of idiosyncratic shocks. To account for the role of fiscal policy, usually the difference between a country pair's budget balances is used for measuring fiscal differences between country pairs; but as the budget balance itself is affected by the business cycle, the issue of endogeneity has to be tackled to avoid reverse causality issues. In this paper, we follow a popular approach by using the difference of the cyclically adjusted budget balances between each country pair i and j in each year t.

$$FD_{ij,t} = |CAB_{i,t} - CAB_{j,t}| \tag{4.6}$$

For the EU countries, the cyclically adjusted budget balance is available from the European Commission's annual macro-economic (AMECO) database. Estimates of the cyclically adjusted budget balance do not exist for the Western Balkan countries, however.[7] Therefore, we use the budget balance reported for each country in the Vienna Institute for International Economic Studies (wiiw) database and adjust it using the method employed by the European Commission (see Mourre et al., 2014). Following the European Commission, we calculate the cyclically adjusted budget balance by

[5] The CEPII database does not provide individual data for Serbia and Montenegro. Therefore, we use the same measure for both countries.

[6] Instead of using the sum of absolute differences as in Barajas et al. (2012), we use the mean given that data points are missing in some sectors.

[7] Croatia's cyclically adjusted budget balance has been available on AMECO only since 2001; in our paper, we therefore treat Croatia like the other Western Balkan economies.

$$CAB_{i,t} = \frac{B_{i,t}}{Y_{i,t}} - \varepsilon * \frac{(Y_{i,t} - Y^p_{i,t})}{Y^p_{i,t}} \tag{4.7}$$

where $\frac{B_{i,t}}{Y_{i,t}}$ denotes the nominal budget balance in terms of country i's GDP at year t, ε stands for budgetary semielasticity, which is a measure of the budget balance's reaction to the level of the output gap $\frac{(Y_{i,t}-Y^p_{i,t})}{Y^p_{i,t}}$. To calculate the output gap, we take the potential output $Y^p_{i,t}$ obtained from the HP filter described in section 4.3.1; for ε we assume a semielasticity of 0.42, which is the unweighted average of the budgetary semielasticities of the individual countries that joined the EU after 2004.

The role of common monetary policy in BCS has entered the discussion mainly with respect to the question whether the business cycles of members of currency areas tend to synchronize – either because currency areas boost trade, which in turn increases output synchronization, or because there is a currency union effect per se (see Gächter and Riedl, 2014). Belke and Heine (2007) argue that the choice of exchange rate regime plays an important role in making the Western Balkan countries' transition process, and eventually their further EU integration, a success.[8] In the regression, common monetary policy is accounted for by using a bilateral dummy variable. It takes the value of 0 at all points in time for Albania, Croatia and Serbia, since these countries – at least de jure – do not fix their exchange rates. For the remaining Western Balkan countries, the dummy variable is set to 1 vis-à-vis Germany, starting from the year in which their currencies were pegged to the Deutsche mark (1995 for North Macedonia, 1998 for Bosnia and Herzegovina, 1999 for Montenegro). Following the euro cash changeover in 2002, the dummy variable takes the value of 1 vis-à-vis the 12 original euro area countries. Subsequently, the dummy variable for Bosnia and Herzegovina, North Macedonia and Montenegro changes to the value of 1 vis-à-vis Slovenia (2007), Cyprus and Malta (2008), Slovakia (2009) and Estonia (2011).

To test for the impact of financial flows, we follow Artis et al. (2008) and use FDI flows scaled by both countries' GDP, compiled from the United Nations Conference on Trade and Development (UNCTAD) FDI database and the wiiw database:

$$FDI_{ij,t} = \left(\frac{FDI^I_{ij,t}}{Y_{i,t} + Y_{j,t}} \right) \tag{4.8}$$

Because only data on inward FDI are available for some Western Balkan countries, we use bilateral net inward FDI flows between the countries i (EU country, sender) and j (Western

[8] The authors would like to point out that other factors such as legal and institutional reforms are relevant as well. However, these factors are largely time invariant and have already been incorporated in the regression by country-pair dummies.

Balkan country, recipient), scaled by the sum of both countries' GDP. We do not expect that excluding outward FDI flows will bias our results as the available data show that outward FDI flows from the Western Balkans to the EU are negligible.

Another proxy for financial flows is taken from the Bank for International Settlement (BIS) locational[9] banking statistics database, which provides data on international financial claims and liabilities of bank offices residing in the BIS reporting countries vis-à-vis the country of residence of the bank's respective counterparty. However, reporting countries neither include the CESEE EU Member States nor Malta, which substantially constraints the observations. In a manner analogous to the construction of the trade and FDI variable, we define

$$BF_{ij,t} = \left(\frac{A_{ij,t} + L_{ij,t}}{Y_{i,t} + Y_{j,t}}\right) \tag{4.9}$$

where bank flows between two countries i and j, $BF_{ij,t}$, are measured as the sum of exchange rate-adjusted flows in assets A_{ij} and liabilities L_{ij} divided by the sum of both countries' GDP.

Remittances are an important source of income in most Western Balkan countries. To take account of the importance of these flows to the Western Balkans, we investigate the effect remittances on BCS between the sending and the recipient country. No data are available on bilateral remittances that fully cover our country sample and time period. However, the World Bank provides data on the aggregate remittance flows the Western Balkan countries received from the EU in a specific year. We use these data for calculating proxies for bilateral flows of remittances.[10] Bilateral migration data are provided by UNCTAD:[11]

$$R_{ij,t} = \frac{M_{ij,t}}{M_{j,t}^{EU}} * R_{i,t}^{EU} \tag{4.10}$$

Remittances R sent from country i to country j at time t are calculated by dividing the number of migrants M of home country j living in host country i at time t by the total number of migrants of the home country (migration share), which is multiplied by total remittances originating from the EU. It has to be noted that, by construction, business cycles of the respective EU countries are not reflected in the constructed variable. However, there is empirical evidence that cyclical output developments of the sending country are more or less irrelevant for the

[9] We use the locational instead of the consolidated banking statistics because they include lending to subsidiaries and affiliates, which would be netted out otherwise.

[10] As inflows of remittances to the home country are significantly correlated with the degree of migrants living in the migrants' host country (see e.g. IMF, 2005; Lueth and Ruiz-Arranz, 2008), we assume that the amount remitted to a Western Balkan country from a certain EU country depends on the number of migrants living in the EU host country as a share of the total number of migrants living in the EU.

[11] Bilateral migration flows are only available for the years 1990, 2000, 2010, 2012 and 2013. For missing years, we approximate the values by calculating moving averages.

propensity to remit (Akkoyunlu and Kholodilin, 2008; Sayan and Tekin-Koru, 2012; Vargas-Silva, 2008).

For estimation[12] we use a system GMM estimator developed by Blundell and Bond (1998), which allows us to instrument endogenous variables with their own lags. Additionally, we employ one external instrument (distance between capital cities) for all regressions. Asymmetry of production, fiscal differences, quasi-common monetary policy and the distance between capital cities are considered exogeneous control variables or instruments, while the other variables are treated as endogenous variables because they are correlated with past and possibly current realizations of the error term. Because the two-step estimation typically yields standard errors that are downward biased, we choose the more efficient Windmeijer's finite sample correction for the two-step covariance matrix.

4.4.3 Regression results

Table 4.2 shows the estimation results of the regressions performed. In addition to the obtained coefficients, we show the results of the Arellano-Bond tests for autocorrelation as, by construction, in first differences an autoregressive [AR](1) process is expected but autocorrelation of AR(2) should not be present because that would indicate that the second lags of endogeneous variables are poor instruments. We also include the p-values of the Hansen test for overidentifying restrictions. Column (1) reports the results from the 'baseline' regression that includes bilateral trade, asymmetry of production and fiscal differences. Columns (2) to (5) add several variables, one at a time, to the baseline model, namely common monetary policy (3), bilateral FDI flows (4), bilateral bank flows (5) and bilateral remittances (6). The results of the Arellano-Bond tests confirm that the lags used as instruments are valid and the Hansen p-values indicate that the results are robust to overfitting.

The results of the baseline regression are in line with the findings of empirical literature. The coefficient of bilateral trade is positive as expected. As pointed out above, some authors have argued that the trade effect for uneven country pairs is usually negligible. The magnitude of the trade coefficient in our baseline regression does not support this hypothesis; however, it can be noted that the coefficient obtained is only one-third of the size of the trade coefficient identified in Gächter and Riedl (2014), who use a similar econometric model but estimate the determinants of BCS between countries within the EU. The asymmetry of the production index yields no significant result; thus we conclude that economic specialization did not play a role in the business cycle convergence process between the Western Balkans and the EU in the observation period. In the baseline regression we also test the effect of fiscal differences; the resulting negative coefficient is in line with the majority of earlier studies (see e.g. Darvas et al., 2005; Crespo Cuaresma et al., 2011), which argued that fiscal policy was the source of idiosyncratic shocks and thus led to greater business cycle divergence.

[12] Our estimation was carried out in STATA, using the xtabond2 environment developed by Roodman (2009).

Table 4.2: Regression results

Variable/model	(1)	(2)	(3)	(4)	(5)
Lagged bilateral BCS	0.080***	0.085***	0.071**	–0.013	0.038
	(0.028)	(0.028)	(0.030)	(0.046)	(0.030)
Bilateral trade	0.046*	0.038*	0.075***	–0.020	0.044**
	(0.024)	(0.021)	(0.028)	(0.035)	(0.020)
Asymmetry of production	0.003	0.005	–0.003	–0.016	0.012
	(0.007)	(0.007)	(0.009)	(0.012)	(0.007)
Fiscal differences	–0.018**	–0.018**	–0.028***	0.006	–0.020**
	(0.009)	(0.008)	(0.009)	(0.012)	(0.008)
Common monetary policy		0.065			
		(0.071)			
Bilateral FDI			–0.009**		
			(0.004)		
Bilateral bank flows				–0.011**	
				(0.004)	
Bilateral remittances					–0.015*
					(0.008)
Constant	1.006**	1.102***	0.598***	0.365	0.759***
	(0.514)	(0.214)	(0.177)	(0.670)	(0.227)
Arellano-Bond test AR(1)	–8.16***	–8.20***	–7.40***	–4.33***	–7.58***
Arellano-Bond test AR(2)	1.24	1.32	0.89	–0.31	0.75
Hansen p-value	0.989	0.991	0.956	0.997	0.966
Number of instruments	205	206	192	135	192
Observations	2,245	2,245	1,994	994	1,905

Source: Authors' calculations. Note: Dependent variable: bilateral BCS. Standard errors are reported in parentheses. $*$, $**$, $***$ indicate a significance level of 10%, 5% and 1%, respectively. Out-of-sample instrument included: logdistcap. In-sample instruments: up to 4 lags. Time dummies are included but not reported. Maximum time span: 1994–2013.

In the next step (2), we add a dummy variable for common monetary policy as described above. However, using the dummy variable does not yield any statistically significant result, meaning that ceteris paribus a Western Balkan country that uses the euro as its currency or nominal anchor does not exhibit higher business cycle correlation with the euro area countries than other Western Balkan countries. In column (3) we add FDI to test for the effect of financial flows on BCS. The result confirms the negative impact of financial flows found by Kalemli-Ozcan et al. (2009) and García-Herrero and Ruiz (2008). Our result points to the argument that FDI flows are procyclical and thus reduce BCS between the Western Balkans and the EU. Although the coefficient seems small, it is rather large when we look at standardized coefficients, which is useful when the regressors are scaled differently (Wooldridge, 2009), as is the case for trade and FDI in our sample. The standardized coefficient[13] is 0.130 for trade and –0.182 for FDI; thus, an increase in the standard deviation of FDI has a stronger effect on BCS than an increase in the standard deviation of trade. A comparison of the coefficient of trade in model (3) with the coefficient obtained in the baseline model (1) supports the argument brought forward by Dées and Zorell (2012) that FDI has an indirect positive effect on BCS via trade.

In model (4) we attempt to test whether the negative correlation coefficient of FDI obtained in model (3) holds when another measure of financial flows is used, namely the flows of bank assets and liabilities between two countries. Unfortunately, because data are not available for all CESEE EU countries and for Malta, about half of the observations have to be dropped in the regression, which causes almost all indicators to become insignificant. However, the coefficient of bank flows is significant, negative and about the same size as the coefficient of FDI in model (3), which tends to confirm the negative impact of financial flows on BCS.

Column (5) reports the regression results obtained when taking into account bilateral remittance flows. In contrast to the results of Barajas et al. (2012), the obtained coefficient is negative, indicating that remittance flows from the EU to the Western Balkans decrease BCS. To the best of our knowledge, no theoretical or empirical research apart from Barajas et al. (2012) has so far included the role of remittances in BCS. However, there is literature that investigates the relationship between remittances and the business cycle of the recipient country.[14] There are two hypotheses on the motives behind remittances (Chami et al., 2008). On the one hand, when sent in order to take advantage of high returns or favorable economic conditions, remittances could exhibit procyclical properties (see Sayan and Tekin-Koru, 2012; Isakovic and Ilgun, 2015; Lueth and Ruiz-Arranz, 2008, for Turkey, Bosnia and Herzegovina and for a global data set, respectively). If this was the case, remittance flows would have a negative effect on BCS by enhancing the size of the business cycle of the recipient country, similar to financial flows. On the other hand, remittances could be used to compensate recipients for unfavorable economic conditions and thus help smooth their consumption patterns. This hypothesis would imply anticyclical behavior with respect to the business cycle (see e.g. Frankel, 2011; Sayan, 2006) and a positive impact of remittances on BCS between the sending and the receiving country.[15] Our result suggests that remittances from the EU to the Western Balkans exhibit procyclical behavior with respect to the business cycle in the Western Balkan economies. Standardizing the coefficient of remittances yields a value of 0.036, while the standardized coefficient of trade in

[13]The standardized coefficient is calculated by multiplying the derived coefficient by the standard deviation of the respective independent variable and dividing it by the standard deviation of the dependent variable (i.e. BCS). The transformation of trade is thus obtained by $0.075 * \frac{2.028}{1.174} \approx 0.130$; the transformation of FDI by $-0.009 * \frac{23.689}{1.174} \approx -0.182$.

[14]As discussed above, empirical evidence points to the fact that cyclical output developments of the sending country are more or less irrelevant for the propensity to remit.

[15]Motives behind remittances do not necessarily have to be congruent with their final use; microeconomic evidence suggests that the majority of remittances is often used for consumption, but that a significant part of remittances also goes into saving or investment (Chami et al., 2008).

this regression is 0.076.[16] Thus, the effect of an increase in a standard deviation of remittances on BCS is about half the size of the effect of a decrease in a standard deviation of trade.

According to the literature, it may be possible that idiosyncratic shocks are transmitted with lags; if this was the case, explanatory variables in one period would affect BCS mainly in the subsequent period. To test this assumption and to put additional restrictions on reverse causality issues at the same time, the most important regressions provided in table 4.2 are re-estimated by lagging all explanatory variables by one period (the lagged bilateral dependent variable is kept at the lag of one period).

Table 4.3: Regression results when lagging explanatory variables

Variable/model	(1)	(3)	(5)
Lagged bilateral BCS	0.070**	0.046	0.041
	(0.028)	(0.028)	(0.030)
Lagged bilateral trade	0.049**	0.076***	0.059***
	(0.020)	(0.024)	(0.022)
Lagged asymmetry of production	–0.001	–0.013	0.003
	(0.007)	(0.009)	(0.007)
Lagged fiscal differences	0.042***	0.046***	0.048***
	(0.013)	(0.015)	(0.011)
Lagged bilateral FDI		–0.009***	
		(0.003)	
Lagged bilateral remittances			–0.024*
			(0.013)
Constant	0.989	0.367**	1.647**
	(0.762)	(0.182)	(0.651)
Arellano-Bond test AR(1)	–8.11***	–7.33***	–7.35***
Arellano-Bond test AR(2)	0.82	0.05	0.20
Hansen p-value	0.957	0.934	0.944
Number of instruments	196	182	182
Observations	2,127	1,878	1,760

Source: Authors' calculations. Note: Dependent variable: bilateral BCS. Standard errors are reported in parentheses. *, **, *** indicate a significance level of 10%, 5% and 1%, respectively. Out-of-sample instrument included: logdistcap. In-sample instruments: up to 4 lags. Time dummies are included but not reported. Maximum time span: 1995–2013.

Performing a robustness test by lagging the explanatory variables by one year yields rather interesting results (see table 4.3). The signs of the coefficients of trade, FDI and remittances do not change; instead, the magnitude of the coefficient of trade and of the coefficient of remittances

[16] The transformation of trade is obtained by $0.044 * \frac{2.028}{1.174} \approx 0.076$; the transformation of remittances by $-0.015 * \frac{2.790}{1.174} \approx -0.036$.

even grows. This leads to the conclusion that idiosyncratic shocks are indeed transmitted with lags. In contrast to the other explanatory variables, the coefficient of fiscal differences becomes positive, larger in magnitude and more significant. The regression results presented in table 4.2 show that fiscal differences are negatively correlated with BCS in the same year. Thus, we argued that fiscal differences are the result of idiosyncratic shocks. However, the outcome of lagging the explanatory variables suggests that fiscal differences actually lead to higher BCS in the following year. The positive sign of the coefficient even holds when the fiscal difference variable is lagged by two years.[17] As argued above, while we cannot instrument fiscal differences, we use the cyclically adjusted budget balance to correct the cyclical component of fiscal spending. However, regressing the fiscal differences on the BCS of the same year does not completely solve the issue of reverse causality. If policymakers anticipated a recession (boom) during the current year or even before the year begins, they could increase (decrease) public consumption or investment to smooth the business cycle. Under the assumption that anticyclical policy needs some time to become effective, such a policy reaction would suggest a negative relationship between BCS fiscal differences in the same year, but as soon as the anticyclical measures unfold, fiscal differences would lead to higher synchronization in subsequent years. To conclude, the changing sign of fiscal differences suggests that fiscal policy is used as an economic stabilizer to help smoothing the business cycle rather than being the source of idiosyncratic shocks.

Moreover, to rule out that the results obtained are driven by the dynamics of one country we run the regressions of 4.2 again, excluding one Western Balkan country at a time. The estimated coefficients prove to be robust in the sense that the signs do not change.[18] However, some of the coefficients become insignificant, which can be traced back to the loss of observations (25 country pairs multiplied by the available time span) in each regression. Similarly, we also estimate each of the models in table 4.2 with a smaller data set that comprises only the current 19 euro area countries instead of the EU-25. Again, our results do not change qualitatively.

4.5 Conclusions

This paper fills several gaps in the literature. It is the first that investigates the business cycle synchronization (BCS) of all Western Balkan economies with the EU-25, i.e. the EU excluding Denmark, Croatia and the U.K., and the first that empirically identifies the determinants of BCS between the two regions. For this purpose, we use a period-by-period correlation index to analyze the convergence process on a yearly basis. Because BCS estimations are prone to endogeneity problems, we employ a system GMM estimator that instruments potentially endogenous variables with their own lagged values.

We clearly identify a process of business cycle convergence between the Western Balkan economies and the EU-25 aggregate from the early transition phase in the 1990s up to the

[17] For brevity reasons, results are not reported here but are available from the authors upon request.
[18] See previous footnote.

year 2013. While prior to 2009, convergence was higher for Bosnia and Herzegovina, Croatia, Montenegro and North Macedonia than for the other Western Balkan countries, after the 2009 crisis year BCS has been high for all Western Balkan countries. Thus, in recent years, the lack or narrow limits of independent monetary policy in the Western Balkans seem not to have been very costly from the perspective of business cycle developments. However, it remains to be seen whether the high degree of business cycle convergence will continue in the longer term. Moreover, it should be recalled that in this study we only examine one of the multiple OCA criteria.

With respect of the determinants of BCS between the Western Balkans and the EU, we find that foreign trade is the most important positive factor. This result is in line with earlier studies that used different regional or global samples. Another variable that is usually tested for in the BCS literature are the fiscal differences between two countries. Here, our results differ somewhat from the findings of other papers. While we also find a negative coefficient of fiscal differences for BCS in the same year, the sign of the coefficient becomes positive in the subsequent years. It therefore seems that fiscal policy is used as an economic stabilizer to help smooth the business cycle rather than being a source of idiosyncratic shocks as has been argued in earlier studies.

We also include two possible determinants of BCS that are less frequently used in the literature but which we assume to be very important in our specific country sample. One is FDI inflows, which poured into the Western Balkans especially in the years preceding the crisis. While empirical studies have not delivered a definite answer on the effect of financial flows on BCS so far, our results show that in the case of the Western Balkans, FDI has led to business cycle divergence. We explain this outcome by the procyclical nature of FDI.

There is hardly any literature on the impact of remittance flows on BCS. One study discovered a positive impact of remittances on BCS. Our results are to the contrary, as we find that remittances sent from EU countries to the Western Balkans actually lead to business cycle decoupling. This supports the hypothesis that remittances exhibit procyclical properties with respect to the receiving economy, similarly to FDI flows.

4.6 Annex

Data sources

Variable	*Source*
Real GDP	*WEO*
Trade	*IMF DOTS*
Gravity variable (distcap)	*CEPII*
Asymmetry of production	*WDI*
Cyclically adjusted budget balance	*Ameco, wiiw*
FDI	*UNCTAD FDI, wiiw*
Bank flows (BF)	*BIS statistics*
Remittances	*World Bank*
Migration data	*UNCTAT*

Source: Authors' compilation.

Descriptive statistics

Variable	*Number of observations*	*Mean*	*Standard deviation*	*Minimum*	*Maximum*
BCS	2,801	1.037	1.174	–1.289	7.991
Log(trade)	2,402	–8.101	2.028	–16.888	–2.694
Asymmetry of production	2,986	9.122	6.157	0.047	38.139
Fiscal differences	2,554	3.579	3.473	0	51.149
Common monetary policy	3,750	0.141	0.348	0	1
Log(FDI)	2,146	–34.846	23.689	–57.565	–4.501
Log(bank flows)	1,244	–35.229	23.829	–57.565	–3.766
Remittances	2,275	0.842	3.007	0	27.745

Source: Authors' calculations.

Chapter 5

Conclusion and discussion

This thesis investigates the implications of euro area monetary policy for non-euro area countries in central, eastern and south-eastern Europe (CESEE), which have either already joined the European Union or are prospective EU members. More specifically, it assesses the extent through which financial and economic conditions in CESEE are dominated by euro area monetary policy, whether this leaves room to manoeuvre for domestic monetary policy (especially in the presence of unofficial financial euroisation) and how costly the potential loss in independent policy making ultimately is.

The first part of the thesis deals with the ECB's non-standard monetary policy measures that were introduced since October 2008, with the results suggesting that spillovers of non-standard monetary policy measures are similar to the spillovers of conventional monetary policy measures identified in the literature. More specifically, the results indicate that the price level of all eight south-east European (SEE) countries included in the study is positively affected by an expansionary non-standard monetary policy shock originating in the euro area, in line with the importance of euro-area imports in total imports and the expansion of domestic activity. Moreover, an expansionary non-standard monetary policy shock seems to have a positive effect on output in approximately half of the countries. The analysis also sheds light on potential transmission channels, as euro area monetary policy spillovers to SEE countries seem to be mostly transmitted via the export channel, and less via short-term interest rates. Nevertheless, financial flows in the form of foreign direct or portfolio investments, which are not captured in the model, still might play a role. Furthermore, the results suggest that the exchange rate regime does not play a role in determining the sign and magnitude of price level and output responses. This is in line with the absence of distinct exchange rate responses in the model output, suggesting that exchange rates did not act as buffers for spillovers of euro area non-standard monetary policy measures on south-eastern Europe during the sample period.

The second part of this thesis also investigates the impact of euro area monetary policy on CESEE countries, but focuses on a specific phenomenon called unofficial financial euroisation.

The results suggest that in the long run more than one third of all euro retail rates in euroised CESEE countries are linked to euro area monetary policy. Moreover, euro retail rates in CESEE adjust excessively to a change in the EONIA, and a deviation from the long-run relationship is corrected by 20% already within the next month. Compared to euro area monetary policy, domestic monetary policy has less of an influence on euro retail rates. This suggests that domestic central banks in countries with independent monetary policy can only partially control the 'euro part' of the interest rate channel, raising the question to which extent monetary policy in those CESEE countries can lean against the 'euro area wind'. For countries with a fixed exchange rate regime linked to the euro, the consequences seem to be less dramatic, with the direct interest rate pass-through to euro retail rates likely to constitute an additional channel of monetary policy transmission.

Finally, complementing the first two parts, the third part of this thesis assesses the cost of the partial loss of independent monetary policy making in terms of business cycle stabilisation for a subset of Western Balkan countries. More specifically, it investigates the degree of business cycle synchronization (BCS) between the Western Balkans and the European Union, and empirically tests which factors are driving business cycle synchronization. The results indicate that while the degree of synchronization had been low or even negative before 2000, business cycles have clearly converged ever since. Since 2010, in particular, BCS with the EU has been high for all Western Balkan countries. With respect to the drivers of BCS, we find that foreign trade is the most important positive contributor to business cycle convergence. Although fiscal differences are associated with negative BCS in the same year, our results suggest that they have a positive influence in subsequent periods. In contrast, we find that financial flows lead to business cycle decoupling because of their procyclical behavior in the respective domestic economy. The same relationship applies, according to our analysis, to remittances, whose impact on BCS is an underresearched topic in the empirical literature. We find that remittances from the EU to the Western Balkans behave similarly to financial flows, which supports the hypothesis that remittances are sent home to take advantage of favorable economic conditions.

Overall, the findings of this thesis confirm that financial and economic conditions in CESEE countries are significantly affected by euro area monetary policy. Moreover, monetary policy independence in CESEE countries seems to be somewhat limited, especially in the presence of unofficial euroisation. This is supported by the fact that no distinct differences in the outcomes of the respective countries with heterogeneous monetary and exchange rate regimes but otherwise very similar characteristics can be found. This raises the question whether the benefits of independent monetary policy for small open economies in the vicinity of the euro area should be reassessed, also in view of a potential euro adoption many countries are aiming at. This is even more the case given the high business cycle synchronization between CESEE countries and the euro area since 2010, which suggests that the absence or limitation of independent monetary policy is not very costly. In fact, business cycle convergence appears to be driven by economic integration through trade, while two other facets of economic integration, namely financial and remittance flows, seem to have a negative effect on business cycle synchronization. At the same time, a common monetary policy (through using the euro as exchange rate anchor in the monetary policy framework) does not appear to make a difference in the degree of business cycle synchronization.

Recently, on 10 July 2020, at the request of the Bulgarian and Croatian authorities, the European authorities have decided to include the Bulgarian lev and the Croatian kuna in the EU's Exchange Rate Mechanism II, a pre-stage of euro adoption.[1] As a consequence, two years ahead of now (or later, depending on when the convergence criteria will be fulfilled), Bulgaria and Croatia will become members of the euro area. Thus, the pool of countries outside of the euro area will become smaller going forward, while the euro area will continue to grow. At the same time, for many other CESEE countries covered in this thesis, euro adoption remains further away, either because those countries are not yet part of the EU, or because euro adoption is currently not a policy priority.

While the ECB's net purchases under the asset purchase programme were phased out temporarily at the end of 2018, the subsequent economic downturn in the euro area forced the ECB to resume its net purchases of assets from November 2019 on. Furthermore, as a response to the unprecedented economic shock posed by the COVID-19 crisis in the first half of 2020, the ECB launched a new temporary asset purchase programme and decided to conduct a new series of seven additional longer-term refinancing operations. Threfore, the issue of spillovers of non-standard monetary policy measures of euro area on CESEE countries will remain topical also in the near future. Finally, the medium-term impact of COVID-19 on business cycle synchronization remains to be seen. While the sudden stop in activity as a response to the first wave affected both the euro area as well as CESEE countries, the recovery might be uneven across countries and thus potentially contribute to a divergence of business cycles going forward.

[1] For further info see `https://www.consilium.europa.eu/en/press/press-releases/2020/07/10/communiques-by-the-erm-ii-parties-on-bulgaria-and-croatia/`.

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