Macroeconomic Policy under Uncertainty and Inequality

DISSERTATION

zur Erlangung des akademischen Grades
doctor rerum politicarum
(Doktor der Wirtschaftswissenschaft)

eingereicht an der

Wirtschaftswissenschaftlichen Fakultät
der Humboldt-Universität zu Berlin

von

M.Sc. Jan Philipp Fritsche

Präsidentin der Humboldt-Universität zu Berlin:
Prof. Dr.-Ing. Dr. Sabine Kunst
Dekan der Wirtschaftswissenschaftlichen Fakultät:
Prof. Dr. Daniel Klapper
Gutachter: 1. Prof. Marcel Fratzscher, Ph.D.
2. Prof. Dr. Lukas Menkhoff
Tag des Kolloquiums: 31.01.2022

Acknowledgements

Throughout the writing of this dissertation, I have received a great deal of support and assistance.

I would first like to thank my supervisor, Professor Marcel Fratzscher, whose expertise was invaluable in formulating the research questions and streamlining my papers. Your insightful feedback pushed me to sharpen my thinking and brought my work to a higher level. Many thanks for guiding me through the research process, while giving me a lot of freedom to explore.

I also thank the Chair of Econometrics and the Chair of Macroeconomics at Free University of Berlin, Helmut Lütkepohl, Dieter Nautz, Matthias Trabandt and Lars Winkelmann for their extensive feedback at seminars. I also thank my second supervisor Lukas Menkhoff for the feedback on the home strait of my thesis. You provided me with the tools that I needed to choose the right direction and successfully complete my dissertation.

I would like to thank my colleagues from DIW Berlin for their advice, support, and all the feedback during research seminars and personal talks. A special thanks goes to my supervisors Kerstin Bernoth and Alexander Kriwoluzky, who gave me the chance to showcase my work in policy papers and pushed me to make my research concise and relevant to the general public as well as to my colleagues who became my co-authors, Malte Rieth and Mathias Klein, for the amazing cooperation, honesty and hours of discussions. Jakob Miethe, Chi Hyun Kim, Claus Michelsen, Sandra Pasch, Nils Aka, Stefan Gebauer, Thore Schlaak, Stephanie Ettmeier, Daniel Bierbaumer, Tobias König, and Alexander Haas many thanks for your feedback and support during my time at DIW.

I would like to acknowledge my colleagues from the European Central Bank for their wonderful collaboration. I would particularly like to single out my supervisors, Fatima Pires, Karen Braun-Munzinger, Michael Grill and Claudia Lambert, I want to thank you for your patient support and for all the opportunities to learn about financial stability, the plumbing of central banking as well as the details of market-based finance and money markets.

A particular thanks goes also to my co-authors Patrick Harms and Lea Steininger, who went through their own PhD journey and, while working together, also became inspiring friends. You made my PhD fun and engaging. I learned a lot from our cooperation and hope to continue it in the future.

In addition, I would like to thank my parents for their sympathetic ear and happy distractions. Finally, I could not have completed this dissertation without the support of my friends and allies. Many thanks for good vibes, creativity, clandestine culture, spontaneous debauchery and plans with perfection.

Abstract

This thesis includes three chapters that inform the debate about macroeconomic policy under uncertainty and inequality. The first chapter shows how low levels of uncertainty are associated with more effective fiscal policy. The second chapter provides evidence for the rule-based policy approach of the ECB reducing monetary policy stress – i.e., *identified* monetary policy uncertainty – in the euro area. This has contributed to the euro becoming a globally dominant currency. Giving up the euro would, for any country, likely result in a situation where the domestic economy would face the adverse effects of globally dominating currencies. The third chapter investigates how monetary policy affects the redistribution of income between employees and owners of companies, and how heterogeneity across firms affects monetary policy transmission at country level. Overall, the results highlight the important role of uncertainty in shaping the fiscal policy transmission mechanism, they contribute to the public debate of euro-skepticism, and they show that between-country inequality can be a result of firm-heterogeneity. Based on the findings of this thesis, a deeper integration of the euro area – i.e. a European fiscal union, joint debt issuance, unified labor rights as well as a completed capital markets and banking union - seem advisable.

Abstract (deutsch)

Diese Arbeit umfasst drei Kapitel zur Debatte über makroökonomische Politik unter Unsicherheit und Ungleichheit. Das erste Kapitel zeigt auf, dass ein geringes Maß an Unsicherheit mit einer effektiveren Ausgabenpolitik einhergeht, und dass fiskalpolitische Ausgaben grundsätzlich ein wirksames Instrument zur Stabilisierung von Konjunkturzyklen sind. Das zweite Kapitel liefert Belege dafür, dass der regelbasierte Politikansatz der EZB den geldpolitischen Stress - d.h. identifizierte geldpolitische Unsicherheit - im Euroraum verringert hat. Dies hat dazu beigetragen, dass sich der Euro zu einer globalen Leitwahrung entwickelt hat. Ein Verzicht auf den Euro würde für jedes der Euroländer wahrscheinlich zu einer Situation führen, in der die heimische Wirtschaft mit den nachteiligen Auswirkungen anderer weltweit dominierender Wahrungen konfrontiert ware. Das dritte Kapitel untersucht, wie sich Geldpolitik auf die Verteilung zwischen Einkommen aus Arbeit und Kapital auswirkt, und wie Heterogenitat zwischen Unternehmen die geldpolitische Transmission beeinflussen kann. Insgesamt unterstreichen die Ergebnisse die entscheidende Rolle von Unsicherheit im Transmissionsmechanismus von Fiskalpolitik, sie leisten einen Beitrag zur offentlichen Debatte zur Euroskepsis und zeigen, dass Ungleichheit zwischen Mitgliedsstaaten eine Folge von Unternehmensheterogenität sein kann. Auf der Grundlage der Ergebnisse dieser Arbeit erscheint eine tiefere Integration der Eurozone - d.h. ein europaisches Finanzministerium mit dem Recht eigene Schulden aufzunehmen, eine Vereinheitlichung des europaischen Arbeitsrechts sowie die Vervollstandigung der Banken und Kapitalmarktunion – als ratsam.

Contents

List of Figures

List of Tables

List of Abbreviations

AT	Austria
AU	Australia
BE	Belgium
BvD	Bureau van Dijk
CA	Canada
CH	Switzerland
D	Doppelganger
DE	Germany
DK	Denmark
e.g.	exempli gratia
EA	Euro Area
ECB	European Central Bank
EM	Expectation–Maximization
EMS	European Monetary System
EMU	European Monetary Union
ES	Spain
et al.	et alia
EU	European Union
(EU) KLEMS	(EU level analysis of) Capital, Labour, Energy, Materials and Service inputs
FI	Finland
FR	France
GARCH	Generalized Autoregressive Conditional Heteroskedasticity
GDP	Gross Domestic Product
GR	Greece
i.e.	id est
IE	Ireland
IH	Identification using Hheteroskedasticity
IT	Italy
JP	Japan
LR	Leverage Ratio
LR	Leverage Ratio
LU	Luxembourg
ML	Maximum Likelihood

MP	Monetary Policy
MS	Markov Switching
MS-SVAR	Markov Switching in Heteroskedasticity Structural Vector Autoregression
MX	Mexico
NACE (Rev.2)	Statistical Classification of Economic Activities in the European Community (Revision 2)
NBER	National Bureau of Economic Research
NK	New Keynesian
NL	Netherlands
NO	Norway
NUTS	Nomenclature of Territorial Units for Statistics
NZ	New Zealand
OCA	Optimum Currency Area
OPEC	Organization of the Petroleum Exporting Countries
PT	Portugal
SCM	Synthetic Control Method
SE	Sweden
SR	Sign Restrictions
SVAR	Structural Vector Autoregression
T-Bill	Treasury bill
TA	Total Assets
UK	United Kingdom
US(A)	United States (of America)
VAR	Vector autoregression
WWII	World War II

General Introduction

The last decade of the euro area has been characterized by macroeconomic uncertainty and inequality. With the emergence of better datasets and improved computational capacities, studying these aspects of macroeconomics has become feasible. This thesis includes three chapters that inform the debate about macroeconomic policy under uncertainty and inequality. The first chapter shows how low levels of uncertainty are associated with more effective fiscal policy. The second chapter provides evidence for the rule-based policy approach of the ECB reducing monetary policy stress – i.e., *identified* monetary policy uncertainty – in the euro area. This has contributed to the euro becoming a globally dominant currency. The third chapter investigates how monetary policy affects the redistribution of income between employees and owners of companies, and how heterogeneity across firms affects monetary policy transmission at country-level.

The aftermath of the global financial crisis triggered the European sovereign debt crisis. Some saw widening yield spreads and macroeconomic imbalances as presages of a collapse of the euro area and suggested that the countries which were hit hardest should abandon the euro. Others urged for the deeper integration of European institutions. Debt levels were declared unbearable in the media, and unemployment was peaking in several European member states when austerity programs were launched with the hope of calming down the markets. The lack of a euro-denominated save asset aggravated the situation and complicated the stabilization of monetary policy in many countries.

During the COVID pandemic - only one decade after the global financial crisis - uncertainty indicators peaked again (Altig et al., 2020), and worldwide fiscal spending programs reached unprecedented levels (IMF, 2020). In the euro area, sovereign debt to GDP levels reached an all-time high and sovereign bond yields reached an all-time low (Corradin et al., 2021). Since 2007, between-country inequality in the euro area has been on a slow but steady rise (Fischer and Filauro, 2021). The European Commission expects within-country inequality to increase further, as the pandemic has hit lower income groups especially hard (European Commission, 2020a,b).

Facing high uncertainty and the risk of rising inequality, there is a need to design macro policies that take these factors into account. The subsequent chapters evaluate macroeconomic policy empirically, and try to derive conclusions for future policy. The two main tools of macroeconomics for stabilizing prices and businesses are fiscal policy and monetary policy. [1] When I assess the leeway for improvement of these tools in the three chapters below, I try to pay special attention to the aspects of predictability (low uncertainty) of outcomes and homogeneity of impact (low inequality) with macroeconomic policies. Low uncertainty fosters expectation formation and prevents precautionary economic inaction. Homogeneous

[1] In the euro area, we have the tendency to think that fiscal policy takes care of the real side of the economy, whereas monetary policy seeks to stabilize prices.

effects are desirable for ethical reasons, but also because they are relevant to economic outcomes. Heterogeneous responses to policies are a particular concern for the outcomes of monetary policy measures in the euro area, because there is no pan-European fiscal authority that could offset the redistributive effects or inefficiencies of monetary policy across member states.

Nussbaum (2019) highlights how emotions like envy and fear can trigger aggression in societies and erode political systems built on trust and care. By keeping inequality and uncertainty to a low level, policymakers can contribute to a cooperative and peaceful society. Because macroeconomic policies are usually sizable, small changes can make great differences. The following chapters try to provide some directives for this. By understanding structural relations and incrementally overcoming the friction discovered, we can set possibilities for experimentation free and find solutions to unprecedented challenges.

Overview of chapters

The leading questions of the subsequent co-authored chapters are:
When and under what circumstances is fiscal spending an effective tool for stabilizing the economy?
Did the euro decrease or increase the capability of monetary policy to stabilize prices and business cycles from the perspective of the individual member countries?
How does monetary policy affect different countries, companies and individuals, and which channels are at play?

The first chapter evaluates the efficiency of fiscal spending policies, using US data with a novel identification approach and an endogenous selection of states of high and low uncertainty. We find that the average output effects of exogenous spending increases are positive, and that the multiplier is significantly and persistently above one. We also analyze the fiscal transmission mechanism holistically. Beyond estimating fiscal spending output multipliers, we discuss the important role of private demand (in particular, private consumption) and investment, in order to better understand our main multiplier estimates. Our finding, that a significant increase in both demand components following a fiscal spending shock, provides fresh evidence for the still-unsettled debate concerning whether a rise in government spending crowds in or crowds out private demand. Besides fiscal policy effects on demand, we highlight the significant impact of fiscal spending shocks on labor market outcomes (like hours worked and employment) and inflation. We present additional evidence showing that the different samples are mainly responsible for the divergent multiplier estimates: more specifically, lower multipliers are implied by the inclusion of war episodes or a reliance on military spending for identification. While the existing literature on uncertainty and fiscal policy relies on constructed uncertainty indices and defines uncertainty regimes based on threshold values, our approach is more data-driven, because it determines the states of uncertainty endogenously. It is more agnostic about the state determination and reduces the risk of misspecification. Overall, our state-dependent findings highlight the important role of uncertainty in shaping the fiscal policy transmission mechanism.

The second chapter evaluates the introduction of the euro from the perspective of the individual euro area countries. To do this, we measure monetary policy stress - i.e., a deviation from stabilizing rule-based monetary policy - in each country before and after the introduction of the euro. We are the first to benchmark the euro areas and their performance by synthetically constructing a doppelganger of the euro area which is sampled from industrialized non-euro area countries. Within the global trend of declining monetary policy stress, the average euro area country has performed better than its synthetic benchmark. Monetary policy stress is a concept that is related to uncertainty but much more narrowly than concepts such as monetary policy uncertainty (Swanson, 2006) and news-based uncertainty (Baker et al., 2016).

Most studies that have proxied monetary policy uncertainty have made statements about the probability density of the interest rate. Such measures do not take into account the fact that part of the variation of policy rates (and therefore its uncertainty) can be explained by supply and demand factors. Our measure is identified and can be linked to rule-based policy. This distinction is particular helpful when discussing the 'one-size-fits-none' reasoning, which is frequently used in policy discussion. This reasoning makes the argument that the unification of interest rate setting is problematic, because heterogeneous business cycles and inflation rates have prevailed while national central banks have been abolished. This reasoning suggests that monetary policy before the introduction of the euro was indeed designed to be capable of stabilizing national business cycles. Proponents of this view seem to over-emphasize the costs of giving up individual currencies while ignoring the important favorable developments associated with the euro. The 'one-size-fits-none' reasoning does not consider how some countries had to respond to foreign exchange developments because their currencies were dominated by the D-Mark or U.S. dollar in the sense of the dominant currency paradigm Gopinath et al. (2020). Consequently, giving up the euro would likely result in a situation where the domestic economy would face the adverse effects of globally dominant currencies.

The third chapter evaluates how firms readjust their labor share, labor costs, and value added after a monetary policy shock. In the basic New Keynesian model, the labor share of value added is expected to decrease after monetary tightening. We are the first to validate this proposition with firm data. With a very granular dataset covering over 2 million firms, we can also study heterogeneity across firms. We hypothesize that heterogeneity in the factor input cost structure matters for monetary policy transmission at firm-level: The higher the firm's average labor share, the more their costs are driven by the payroll. After a monetary policy shock occurs, labor-intensive firms alter their total costs by altering the costs of employees. Conversely, firms characterized by a high leverage ratio primarily react by adjusting production and value added – because they have a larger and more diverse balance sheet, they can alter their cost and earning structure by altering this balance sheet. This heterogeneity is important for policymakers, because firm heterogeneity can be a catalyst for redistributive effects and the efficiency of monetary policy. We can confirm our hypothesis, and find that a firm's average labor share and the leverage ratio are important factors with which to differentiate between firms. Since value added and payroll expenses do not react synchronously on average and in addition heterogeneously across companies, monetary policy has a redistributive effect. Monetary policy affects the redistribution of income between workers and capital. Because the heterogeneity of firms is clustered at the country-level monetary policy has likely a redistributive effect on countries: At the macro-level, understanding wage setting is central to understanding inflation dynamics and unemployment. A weak wage response can translate into a weak inflation rate response following a monetary policy shock. Heterogenous wage setting across firms can therefore exacerbate macroeconomic imbalances if they are clustered as they are across the regions of the euro area. We conclude that more direct transmission channels could alleviate the redistributive effect of monetary policy. In addition, unified European labor rights and a deeper integration of capital and labor markets could improve monetary policy transmission to the euro area countries.

Limits and scope of this dissertation

This dissertation contributes to a growing field of macroeconomic literature by using innovative methods and studying new datasets. Nevertheless, it is important to highlight the current limits of empirical work in this context. The greatest limitations on my dissertation are data availability and quantifiability. For some policy measures, no data is available to assess them empirically. For example, there is no European fiscal authority whose efficiency could be assessed. Quantifiability is a limit for both uncertainty and inequality. Aggregate uncertainty is a vague concept and relatively difficult to measure. Inequality is easier to measure, but redistributive effects must be traded off against stimulus - which usually requires a normative statement.

Limits to data are a particular concern for designing programs to tackle challenges that have not yet been overcome, such as the looming climate catastrophe. However, data constraints are also an issue for this thesis. For example, most top-tier studies [2] on fiscal policy look at US data, because of the great

[2] see the literature review in Ramey (2019)

availability of historical data and the number of studies that provide benchmarks for analytical work. Consequently, there is a path dependence, because we can better understand drivers of the results across studies if we base studies on the same data. Nevertheless, with a US-centric perspective we are likely to miss relevant features and be biased when we try to generalize the results. An artifact in this context, which we have to better account for in the future, is the sheer military power and activity of the United States. An important share of the literature on fiscal spending cited in the first chapter relies on military spending data to identify government spending shocks. However, as military expenses, investments and education are usually not comparable to their civilian equivalents, we need to better distinguish between them to understand which kind of policy is most efficient.[3] With the availability of better data, we will be hopefully able to answer questions like: How should public spending programs be structured to be most efficient?

Uncertainty is in most cases quantified by calculating the (expected) higher moments of a probability distribution of observables.[4] So far, there is no consensus in the literature, what proxy of uncertainty is most relevant for macroeconomics. Among others, the forecast error density of one key variable, the total variation of a system of equations, or the amount of uncertainty related articles in newspapers have been suggested to measure degrees of macroeconomic uncertainty.[5] In a Knightian sense, quantifying uncertainty is almost impossible. Knight (1921) refers to uncertainty as unpredictability of events. When quantifiable risks turn in to unpredictable threats, fear and instincts can take over and guide the actions of economic agents. Such scenarios are likely not well to be described by linear or linearized systems, which are mostly used in empirical applications. Instead, we must resort to qualitative analysis.[6] However, even when we can approximate uncertainty by higher moments of the data, its causal identification is complicated. Ideally, we would like to study a shock and its underlying uncertainty independently of each other. Otherwise, large shocks translate to high uncertainty. For monetary policy, this would mean disentangling identified monetary policy shock from identified monetary policy uncertainty shocks – something which has not been achieved. Most studies focus on the identification of one of the shocks, whereas the other is investigated in a more cursory manner.[7] In Chapter 2, we define the monetary shocks according to the needs of our research question - as causal drivers of monetary stress, such that they characterize pronounced deviations from a simple stabilizing rule. This means, however, that they cannot easily be compared with the monetary policy shocks of other applications that study the stimulative effect of monetary policy such as Jarociński and

[3] For example, there is likely a difference between a vehicle that is used for public transportation and warfare, e.g., one could conjecture that the expected depreciation rates or network effects would be different.

[4] an exception in a theoretical application is Burzoni et al. (2021), who define an order instead of a probability distribution. However, in empirical application ordering outcomes seems impractical.

[5] Bloom (2014) discusses a variety of these measures.

[6] see for example Nussbaum (2019)

[7] see for example Husted et al. (2020)

Karadi (2020). In the latter applications, one might think of a large monetary policy shock to eliminate certain degrees of uncertainty. A famous example is the "whatever it takes" statement by Mario Draghi.[8]

For inequality, measurement is usually straightforward, but the data requirements are enormous and policy conclusions are harder to derive. Because macroeconomics usually studies aggregates (such as inflation or GDP), the field has an inherent tendency to abstract from inequality.[9] In the past, this tendency was reinforced by data limitations. To assess the heterogeneity and inequality of macroeconomic effects or outcome, micro data – e.g., on firm, plant, product, household or individual level - which should cover multiple decades is required. It should include multiple economic business cycles and be representative at the macro-level. There is also a moral and cultural dimension to the desirability of certain types of inequality. A macroeconomic policy that is designed to stimulate the economy will not affect everybody equally. The acceptable trade-off between stimulus and inequality must be part of a democratic process, and cannot be defined by economists. While some might find it reasonable to make an individual worse off for the greater good, others become skeptical when some benefit significantly more than others. Eventually, empirical macro studies will be able to point out structural factors that lead to inequality or, in the best case, fine-tune policy instruments to the degree that they are less redistributive, while maintaining their stimulating impact.

[8]At the Global Investment Conference in London on 26th of July 2012, Mario Draghi, who was president of the European Central Bank at that time, said "Within our mandate, the ECB is ready to do whatever it takes to preserve the euro. And believe me, it will be enough." (European Central Bank, 2012)

[9]The recent emergence of heterogeneous agent models that show how inequality can affect the transmission of macro policies and macro aggregates might change this paradigm (Kaplan et al., 2018).

Chapter One

Government Spending Multipliers in (Un)certain Times

This chapter is based on joint work with Mathias Klein and Malte Rieth, which has been submitted in a revision to the Journal of Public Economics. A previous version of this project was published as DIW Discussion Paper No. Discussion Papers 1901 in 2020 under the name same name

1 Government Spending Multipliers in (Un)certain Times

We estimate the dynamic effects of government spending shocks, using time-varying volatility in US data modeled through a Markov switching process. We find that the average government spending multiplier is significantly and persistently above one, driven by a crowding-in of private consumption and non-residential investment. We rationalize the results empirically through a contemporaneously countercyclical response of government spending and an efficient weighting of observations inversely to their error variance. We then show that the multiplier is significantly smaller when volatility is high, consistent with theories predicting reduced effectiveness of fiscal interventions in uncertain times.

Keywords: Fiscal policy, government spending multiplier, uncertainty,
structural vector autoregressions, heteroskedasticity.
JEL classifications: C32, E62, H50.

1.1 Introduction

There is renewed interest among researchers and policy makers in the effects of fiscal policy on macroeconomic activity. This topic is especially important when interest rates are close to zero. In this case, fiscal policy becomes the instrument of last resort. Despite the importance of the topic, there is a wide range of different estimates for the size of the government spending multiplier; that is, by how much does aggregate output rise following an exogenous increase in government spending. In addition, there is a discussion whether the effectiveness of fiscal interventions depends on the state of the economy (Auerbach and Gorodnichenko (2012); Ramey and Zubairy (2018); Tagkalakis (2008)).

In this article, we use an agnostic identification approach to estimate the dynamic effects of government spending shocks in the United States. For identification we exploit time-variation in the volatility of US data, which we model through a Markov switching in heteroskedasticity process following Herwartz and Lütkepohl (2014). Several studies show that changes in volatility are a main feature of macroeconomic time series (Stock and Watson (2002), Justiniano and Primiceri (2008), Carriero et al. (2016), Diebold et al. (2017)). We use these changes in volatility as a 'probabilistic instrument' (Rigobon, 2003) for government spending shocks to study how fiscal policy influences the macroeconomy.

Compared to other identification schemes frequently used in the fiscal policy literature, identification through heteroskedasticity offers three main advantages. First, unlike the recursive identification scheme (Blanchard and Perotti, 2002), it does not impose a timing restriction and allows for a contemporaneous response of government spending to output. Second, contrary to the narrative identification scheme (Barro and Redlick, 2011; Ramey, 2011), it does not require the availability of a valid external instrument for post-WWII data. Third, different to linear models, it generalizes the estimation by weighting observations inversely to their sampling uncertainty, thereby increasing efficiency.

Furthermore, because the model is nonlinear, it allows for studying the state-dependent effects of government spending shocks across different volatility regimes. The theoretical analysis of Bloom (2009) and Bloom et al. (2018) predicts that fiscal policy is less effective when uncertainty, modeled through time-varying structural shock variances, is high, because firms then postpone hiring and capital decisions. Similarly, Bernanke (1983) and McDonald and Siegel (1986) show that higher uncertainty increases the real option value of waiting before making investment decisions. Additionally, in uncertain times there is a stronger precautionary savings motive by consumers (Fernandez-Villaverde et al. (2011), Basu and Bundick (2017)). By allowing the effects of government spending shocks to depend on the volatility regime, our model enables us to test whether these theoretical predictions are supported by the data.

We find that exogenous changes in government spending are, on average, an effective tool to stimulate the economy. A positive government spending shock leads to a significant increase in output and private consumption. The crowding-in of private consumption is driven by a persistent and significant rise in both non-durable and durable consumption. Moreover, the unexpected fiscal stimulus leads to a significant rise in inflation, short-term interest rates, employment, and hours worked.

We estimate a cumulative two-year government spending multiplier of 1.5. The multiplier is significantly larger than one for about two years. This estimate is in line with recent evidence (Ben Zeev and Pappa, 2017), but larger than estimates typically found when applying timing-restrictions or exogenous changes in defense spending. We show that there are three main reasons for the differences across identification schemes: two economic and one technical reason. First, we do not restrict the contemporaneous response of government spending to output to zero but estimate this elasticity. We document a significantly negative response. As the size of the government spending multiplier is negatively related to this elasticity, a countercyclical spending response implies a larger multiplier, as shown formally by Caldara and Kamps (2017). While their approach requires the availability of non-fiscal instruments to estimate the government spending-output elasticity, our methodology allows estimating this relationship by making use of a natural feature of macroeconomic time series, changes in volatility. Second, military spending shocks produce smaller government spending multipliers than general government spending shocks. Third, accounting for heteroskedasticity generalizes the estimation. It gives more weight to observations with low error variance compared to observations with high error variance. This increases efficiency and affects point estimates.

Regarding state-dependency, we find evidence for the hypothesis that the impact of government spending shocks varies across volatility regimes. The multiplier is significantly larger in the low than in the high volatility state. This result is consistent with the theoretical predictions and it suggests that the level of uncertainty in the economy affects the effectiveness of government spending policy. Moreover, it rationalizes our finding of an average multiplier above one as the generalized model attaches larger weight to the more precisely estimated larger multiplier of the low volatility regime.

Related literature. This paper is closely related to two studies that also use changes in volatility to identify fiscal policy shocks. Bouakez et al. (2014) model conditional heteroskedasticity through a GARCH process, whereas we use a Markov switching framework. Lütkepohl and Schlaak (2018) show that the latter is usually the better choice when the data generating process is unknown because modeling time-varying volatility through a latent variable gives more voice to the data, yielding more precise estimates of the structural parameters. Moreover, Bouakez et al. (2014) pre-define one break point in 1979 for general, unspecified changes in the effects of fiscal policy. Instead, our state-dependent approach is agnostic about the change points, which can be multiple and are determined endogenously, and, at the same time, more specific about the nonlinearities, which are due to changes in uncertainty.

Lewis (2021) concentrates on the econometric theory of identification through heteroskedasticity. He presents the methodology in an extensive and rigorous manner and then applies it to fiscal policy by estimating aggregate tax and government spending multipliers. We depart in three dimensions, which are especially important from a policy perspective. First, while that paper focuses on estimating aggregate fiscal multipliers, we study the transmission mechanism in detail. We find a significant increase in both private consumption and private investment following an expansionary fiscal spending shock, providing fresh evidence on the unsettled debate whether a rise in government spending crowds in or crowds out private demand. Furthermore, we document a significant positive impact of fiscal spending shocks on labor market outcomes, like hours worked and employment, and on inflation. Second, we highlight the role of efficient observation weighting in explaining differences in estimated multipliers. Third, we extend the methodology to a state-dependent setting and show that the government spending multiplier is significantly lower when uncertainty is high, that is, when fiscal stimulus is probably needed most.

Our paper also complements studies on the state-dependent effects of fiscal policy. Ricco et al. (2016) document that more disagreement among US professional forecasters about future government spending reduces the efficacy of government spending news shocks in a Bayesian threshold SVAR. Alloza (2017) shows that recursively identified government spending shocks, or military spending news shocks, have smaller effects on output when stock market volatility is high. He uses a SVAR interacted with a dummy variable indicating these periods. Wolff and Jerow (2020) employ forecast errors in survey data on spending expectations to identify government spending multipliers. They first show using local projections that the effects of the shocks are smaller when macroeconomic uncertainty is high and then construct a medium-scale DSGE model to rationalize this finding. Our analysis differs from these papers regarding the identification strategy, the state variable, and the empirical model. Whereas they either rely on a recursive identification or instruments for spending shocks, we use time-varying volatility for identification. Moreover, while they use constructed uncertainty indices and define regimes ex ante based on some cut-off rule, our approach is more data-driven. The Markov switching in heteroskedasticity model is more agnostic about the state determination, which reduces the risk of misspecification of the transition variable, function, or points. Relative to threshold models it is more flexible as it allows for mixtures of states.

The rest of the paper is structured as follows. Section 2 describes the methodology and data. Section 3 presents our baseline results and shows robustness of the estimated average multiplier. In Section 4, we discuss our findings in the light of the literature. In Section 5, we allow for state-dependent effects of government spending shocks and investigate how uncertainty affects the fiscal transmission mechanism. Section 6 concludes.

1.2 Empirical methodology and data

In this section, we first present the empirical model and the identification strategy. We then describe the data and the estimation procedure.

1.2.1 The MS-SVAR model

The general M state, p lag reduced form Markov switching in heteroskedasticity structural vector autoregressive (MS-SVAR) model with n variables is

$$y_t = c + A_1 y_{t-1} + \cdots + A_p y_{t-p} + u_t. \tag{1}$$

In the baseline model $y_t = [g_t, x_t, r_t]'$, with g_t government spending, x_t output, and r_t the three-month Treasury Bill rate. Below, we conduct an extensive sensitivity analysis, showing among others that our main findings are robust to including tax revenues. Further, c is a vector of constants, A_i are parameter matrices with $i = 1, \ldots, p$, and u_t is a vector of zero-mean reduced form errors. In standard SVAR models, a linear transformation is used to obtain the structural shocks, ε_t, as $\varepsilon_t = B^{-1} u_t$ or $B\varepsilon_t = u_t$. Usually, $\varepsilon_t \sim (0, I_n)$ and the reduced form covariance matrix is decomposed as $E[u_t u_t'] = \Sigma_u = BB'$.

In our analysis, we follow Herwartz and Lütkepohl (2014) and assume that $E[u_t u_t'] = \Sigma_u(S_t)$. S_t is a first order discrete valued Markov process that can take on M different values, $S_t = 1, \ldots, M$, with transition probabilities given by $p_{kl} = P(S_t = l | S_{t-1} = k), k, l = 1, \ldots, M$. For estimation with maximum likelihood (ML), we assume that u_t is normally and independently distributed conditional on a given state:

$$u_t | S_t \sim \mathrm{NID}(0, \Sigma_u(S_t)). \tag{2}$$

The normality assumption is not critical for the empirical analysis. If conditional normality is not fulfilled, then the estimation will simply be pseudo ML.

While the model is linear in a given state, it is nonlinear as a whole. Equations (1) and (2) imply that only the reduced form covariance matrix switches between states, as we are interested in the heteroskedasticity features of the data for identification purposes. Thus, we impose more regularity on the model than in the MS-SVAR model used by Sims and Zha (2006). Furthermore, although there is a finite number of states in practice, the model captures smooth transitions between them as the actual volatility is described as a mixture of states, each weighted with a certain probability that may be less than one.

We exploit these changes in the covariance matrix for the identification of structural shocks that are consistent with the statistical properties of the data. We consider two states of the world – a high-volatility state and a low-volatility state. Then, we can decompose the reduced form covariance matrices as:

$$\Sigma_u(1) = BB' \quad \text{and} \quad \Sigma_u(2) = B\Lambda B', \tag{3}$$

The main idea underlying this identification strategy is to use additional moments from the data to ensure that the order condition holds. With only one state-independent estimate of Σ_u, we would have only six estimable moments for $n = 3$ but nine unknown parameters in B, if the latter is unrestricted. With two volatility regimes, we have 12 moments and 12 unknowns as $\Lambda = diag(\lambda_1, \lambda_2, \lambda_3)$ contains only three additional (positive) elements on the main diagonal and otherwise zeros. Lanne et al. (2010) show that if these elements are all distinct, the rank condition holds, that is, the decomposition in (3) is unique apart from changes in the signs of the shocks and permutations of the $\lambda_j, j = 1,\dots,n$ and corresponding orderings of the columns of B. In summary, if we order the λ_js, assume that the structural shocks are orthogonal, have the same impact effects across states, and are normalized to have unit variance in the first state, we can uniquely identify the structural shocks through the linear transformation $\varepsilon_t = B^{-1}u_t$.

The crucial assumption for point-identification of the full B-matrix is that the λ_js are all different. As they can be interpreted as variance shifts of the structural shocks relative to the benchmark state (in our case the low-volatility state), having distinct λ_js means that the volatility shifts are not the same for all shocks. This assumption can be checked after estimation, which is an advantage over more conventional just-identifying assumption that cannot be assessed. Furthermore, zero restrictions on B can become over-identifying in the presence of heteroskedasticity and, hence, testable. Finally, for the specific decomposition (3), we assume that B is state-independent. We relax this assumption in Section 1.5.

Figure 1.1 illustrates how identification through heteroskedasticity works. Both panels present artificial data describing the relationship between government spending and output under the assumption that government spending negatively depends on output (government spending is countercyclical) and output increases with government spending (government spending multiplier is positive). It is impossible to identify the output response to variations in government spending from the left panel alone, because every line fitted to the system will match the data equally badly. However, once the variance of the spending shock increases, as shown in the right panel, shifts of the spending equation increase relative to shifts of the output equation. The circle turns into an ellipse that is centered around the output equation such that the latter can be estimated from the data. The important condition is that there are no proportional variance changes as this would simply imply a widening of the circle.

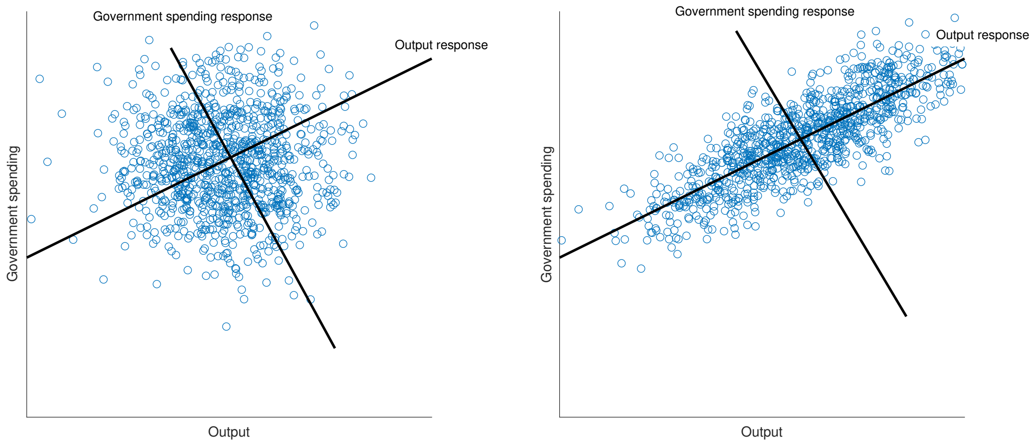


Figure 1.1: Graphical illustration of identification through heteroskedasticity.

Notes: Left panel shows state 1, the right panel state 2.

Relying on this identification approach is also convincing from an economic point of view. Several studies document that changes in volatility are a main feature of macroeconomic data (Stock and Watson (2002), Justiniano and Primiceri (2008), Carriero et al. (2016), Diebold et al. (2017)). One prominent example is the change associated with the Great Moderation that started around the 1980s. Macroeconomic aggregates are more volatile before and fluctuate less during that period. Our approach makes use of such data features to identify structural shocks. This is less restrictive than using timing, sign, or external instrument restrictions. It does not restrict the size or the sign of the impact, nor does it require the availability of a valid instrument for post-WWII data.

A related statistical approach to identify structural shocks and estimate fiscal multipliers is proposed by Guay (2020). He makes use of higher moments of the reduced form innovations to achieve identification either of the entire structural system or subsystems thereof, and provides a testing procedure for identification prior to estimation. He applies this new methodology to estimate tax and spending multipliers.

1.2.2 Estimation, data, and model selection

We estimate the parameters in (1) by means of the expectation maximization (EM) algorithm developed for structural models by Herwartz and Lütkepohl (2014). Crucial for the analysis is to incorporate the regime-switching nature of the covariance matrix described in (2), given the restrictions in (3). We use the following concentrated out log likelihood function in the maximization step of the EM algorithm, which weighs observations according to their sampling uncertainty:

$$\mathscr{L}(B,\Lambda_m) = \frac{1}{2}\sum_{m=1}^{M}\left[\widehat{T}_m \log(\det(\Sigma_u(m))) + \mathrm{tr}\left((\Sigma_u(m))^{-1}\sum_{t=1}^{T}\widehat{\xi}_{mt|T}\widehat{u}_t\widehat{u}_t'\right)\right], \tag{4}$$

where $\xi_{mt|T}, m = 1,\ldots,M,\ t = 1,\ldots,T$ are the model smoothed probabilities, $T_m = \sum_{t=1}^{T}\xi_{mt|T}$, and the hat denotes estimated parameters from the previous iteration.

Once the EM algorithm converges, we obtain standard errors of the point estimates of the parameters through the inverse of the negative Hessian matrix evaluated at the optimum. We use them as a statistic to determine whether the estimated variances change significantly and by differing amounts across states. This is a requirement for identification. For the dynamic analysis, we compute bootstrapped impulse responses, following Herwartz and Lütkepohl (2014). Given the heteroskedasticity, classical residual bootstrapping may be problematic for generating reliable confidence intervals. Re-sampling needs to preserve the second order characteristics of the data. Therefore, we use a fixed design wild bootstrap with $u_t^\star = \varphi_t\widehat{u}_t$, where φ_t is a random variable independent of y_t following a Rademacher distribution. φ_t is either 1 or –1 with probability 0.5.

Our baseline model consists of three endogenous variables, namely government spending (the sum of government consumption and government investment at the federal, state and local level), GDP, and the three-month Treasury Bill rate. Several studies show that the conduct of monetary policy influences the macroeconomic effects of fiscal policy (Canova and Pappa (2011); Davig and Leeper (2011)). Specifically, the extent to which monetary policy reacts to the consequences of fiscal spending shocks is an important determinant of their size. Therefore, we explicitly control for the stance of monetary policy. In Section 1.3.4, we show that our main qualitative results are robust when controlling for tax policy, fiscal foresight and when excluding the period during which the interest rate has reached its lower bound.

All series are from Ramey and Zubairy (2018). Nominal government spending and output are divided by potential output such that they are measured in dollar equivalents. The baseline sample covers the period 1954Q1-2015Q4. The starting date avoids the years from 1945 to the Korean war, commonly considered to be special within post-WWII data from a fiscal point of view (Monacelli et al., 2010). In Section 1.4, we discuss the importance of the sample for understanding our main multiplier estimate and the differences to other studies. We use five lags to account for potential seasonal patterns in the quarterly variables. This also increases the confidence in results for longer response horizons.[1]

Table 1.1 shows some specification statistics for the two-state Markov switching model. It is clearly

[1]The qualitative results are robust when changing the lag length.

Table 1.1: Model selection.

Notes: L_T denotes the likelihood function evaluated at the optimum, $AIC = -2log(L_T) + 2f$, and $SC = -2log(L_T) + log(T)f$, with f the number of free parameters.

Model	$log(L_T)$	SC	AIC
Reduced form linear VAR(5)	1791.8	-3335.5	-3493.6
Reduced form MS(2)-VAR(5)	1899.8	-3459.1	-3675.7

preferred over a standard linear VAR according to the log-likelihoods and both types of information criteria. The latter are shown to work well to judge the performance of MS models (Psaradakis and Spagnolo, 2006), whereas standard tests are problematic for this purpose as some parameters might not be identified under the null hypothesis of a smaller number of states than under the alternative (Hansen, 1992). The evidence against the linear model is strong. This suggests notable changes in volatility in the sample and that the identification of the structural shocks can be achieved by relying on the heteroskedasticity property of the data.

1.3 Government spending shocks and multipliers

1.3.1 Estimated volatility regimes and identification

We first analyze the estimated state-dependent reduced form covariances to see whether the model captures and separates the changes in volatility apparent in the data, according to Table 1.1, which are crucial for our identification strategy. This information also helps us interpret our endogenously identified regimes. Table 1.2 presents the estimated state covariances of the MS(2)-SVAR(5) model. It shows strong increases in volatility in state 2. The variances of the reduced form errors in the equations for public spending, output, and the interest rate increase by factors of 3, 5, and 17, respectively. We read this as further evidence that the sample is characterized by strong changes in volatility.

State 1: $\Sigma_u(1)$	State 2: $\Sigma_u(2)$
$\begin{bmatrix} 0.024 & . & . \\ 0.033 & 0.228 & . \\ 0.001 & 0.028 & 0.062 \end{bmatrix}$	$\begin{bmatrix} 0.079 & . & . \\ 0.073 & 1.164 & . \\ -0.016 & 0.280 & 1.055 \end{bmatrix}$

Table 1.2: Estimated state covariances of MS(2)-SVAR(5) model with $y_t = [g_t, x_t, r_t]'$.

The table also shows that the covariances increase substantially (in absolute value) in state 2; by about similar multiples as the variances relative to state 1. These changes in the covariances illustrate the idea behind identification through heteroskedasticity. In a period of high government spending volatility, we learn more about the relation between government spending and output as the covariance between both variables temporarily increases. Government spending shocks are then more likely to occur and can be used as a 'probabilistic instrument' (Rigobon, 2003) to trace out the response of output.

To achieve identification from a statistical point of view, we need significant and differential changes in the volatility of the structural shocks. Table 1.3 shows the estimated structural variances in state 2. As the ordering of the λ_js is arbitrary, we simply order them from smallest to largest. All estimates are significantly larger than one, implying that all structural shock variances increase when switching from state 1 to state 2. Thus, we label state 2 the high volatility state. Identification requires that the variance shifts are all distinct from each other. This is the case according to their standard errors, which do not overlap, and suggests that the decomposition in (3) is unique.

Table 1.3: Estimates and standard errors of relative variances of the MS(2)-SVAR(5) model.

Notes: The standard errors are obtained from the inverse of the negative Hessian evaluated at the optimum.

Parameter	Estimate	Standard error
λ_1	3.25	0.89
λ_2	6.05	1.38
λ_3	17.28	3.75

Furthermore, we test for identification more formally, following Lütkepohl et al. (2021). We use the smoothed state probabilities to determine the fraction of observations in state 1 τ, the total number of quarterly observations T, and the estimated changes in the structural shock variances λ_js in the following test statistic:

$$Q_r\left(\tilde{\kappa}_1, \tilde{\kappa}_2\right) = c\left(\tau, \tilde{\kappa}_1, \tilde{\kappa}_2\right)^2 \left[-T \sum_{k=s+1}^{s+r} \log\left(\lambda_k\right) + Tr \log\left(\frac{1}{r} \sum_{k=s+1}^{s+r} \lambda_k\right)\right],$$

where

$$c\left(\tau, \tilde{\kappa}_1, \tilde{\kappa}_2\right)^2 = \left(\frac{1+\tilde{\kappa}_1}{\tau} + \frac{1+\tilde{\kappa}_2}{1-\tau}\right)^{-1},$$

r is the number of restrictions and $s \in \{0,1\}$. We set the kurtosis parameters $\tilde{\kappa_1} = \tilde{\kappa_2} = 0$, given the Gaussianity assumption.

We perform three tests, summarized in Table 1.4. The first tests the equality of all three structural shock variances in state 2, while the other two test two λ_js against each other. In all cases, the null hypothesis is that the shock variances are equal. The asymptotic distribution of $Q_r\left(\tilde{\kappa}_1, \tilde{\kappa}_2\right)$ under the null is χ^2 with $\frac{1}{2}(r+2)(r-1)$ degrees of freedom. In all three cases, the null hypothesis of equality of the structural shock variances is rejected. The associated p-values are essentially zero for the first and the third test. The second test rejects the null hypothesis at the 10% level. These results support the evidence from the estimated standard errors, and are in line with Lütkepohl et al. (2021) who show that both approaches lead to similar conclusions. Together, the evidence indicates that the baseline model is statistically identified.

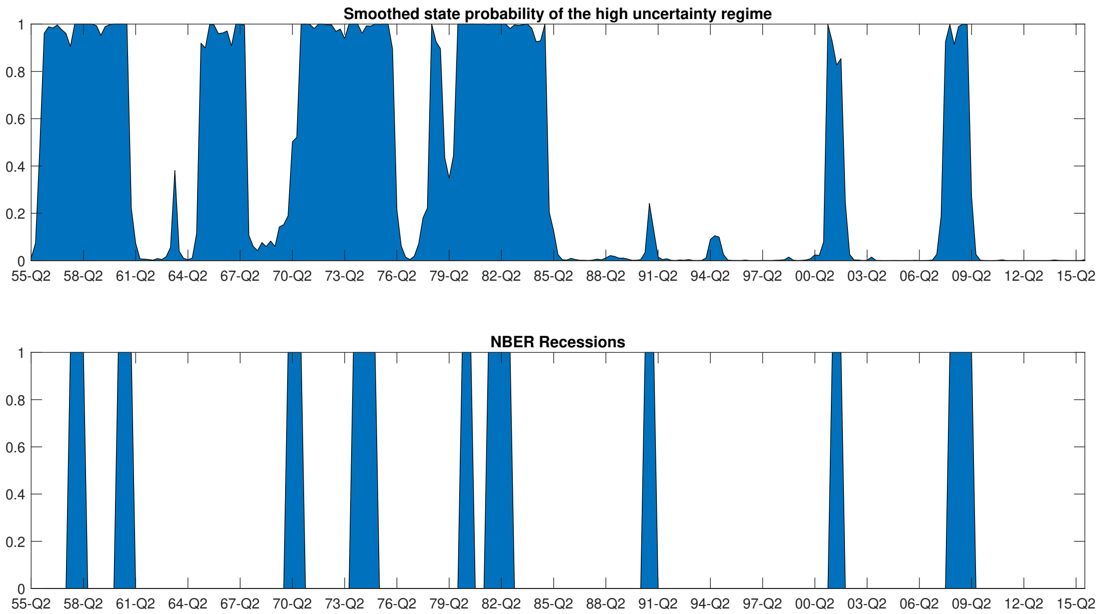
H_0	$Q_r(\tilde{\kappa}_1, \tilde{\kappa}_2)$	Degrees freedom	p-value
$\lambda_1 = \lambda_2 = \lambda_3$	41.142	5	8.785×10^{-8}
$\lambda_1 = \lambda_2$	5.444	2	0.066
$\lambda_2 = \lambda_3$	15.152	2	5.127×10^{-4}

Table 1.4: Identification tests.

Figure 1.2 presents the estimated smoothed state probabilities to develop a notion about the economic drivers of the regimes. The upper part shows the smoothed probabilities for the high volatility regime. This regime prevails in the first part of the sample, the low-volatility state dominates the second part. The pattern might reflect the changes associated to the relatively tranquil times of the Great Moderation in the 1990s and 2000s with high growth and low inflation. In addition, many of the transitions to the high volatility state are associated with specific events in the economic history of the US. There are peaks around the OPEC oil prices shocks in the late 1960s and the beginning of the 1970s. Moreover, there is a spike around the Energy Crisis, the subsequent economic recession, as well as the chairmanship of the Federal Reserve of Paul Volcker at the end of the 1970s and the first half of the 1980s. In the second part of the sample, there are peaks around the burst of the Dot-com Bubble in 2001 and the Great Recession in 2009. Overall, this short narrative suggests that the endogenously determined regimes capture relevant developments in the US economy.

Figure 1.2: Smoothed state probabilities.

Notes: The figure shows the smoothed state probabilities $\hat{\xi}_{mt|T}$ for the high volatility state ($m = 2$) in the upper panel and recessions defined by the NBER in the lower panel.



The estimated probabilities and shock variances can also be rationalized from a fiscal policy perspective. As shown by Davig and Leeper (2011), fiscal policy during the pre-Volcker period is characterized by a high degree of instability. In particular, they find that during that period, fiscal policy deviated from a policy rule that induces government debt sustainability. This fiscal instability can affect private agents' consumption and investment decision as it complicates the projection of future policy adjustments. Consequently, an unstable fiscal policy environment might increase economic uncertainty reflected in larger fluctuations of aggregate variables. Moreover, other studies detect a change in the fiscal transmission mechanism at the beginning of the 1980s (Bilbiie et al., 2008; Perotti, 2005) which, in light of our smoothed probabilities, seems to be associated with moving from the mainly high uncertainty regime to the mainly low uncertainty regime. Besides the specific conduct of fiscal policy, the smoothed probabilities spike around specific war episodes in the US history, such as the Lebanon Crisis, the Vietnam War, and the Afghanistan War. Because military spending is one main part of overall government spending, substantial increases in defense spending induced by military interventions lead to larger movements in government spending, which finally implies more volatility in public expenditures.

The bottom panel shows recessions, as defined by the NBER. The graph shows that state 2 has some commonalities with recessions but measures something different. The correlation between both series is 0.43.

Having labeled the regimes, we now turn to the interpretation of the structural shocks. As mentioned above, our data-driven identification scheme is not based on *a priori* economic reasoning about the model-economy and a corresponding ordering of the columns of B or the respective λ_j. Thus, to attach an economic label to the statistically identified shocks, we inspect the forecast error variance decomposition and call the shock that explains the largest part of the variance in government spending a government spending shock. This idea is based on the identification scheme of Blanchard and Perotti (2002), which implies that government spending is contemporaneously exclusively driven by public spending shocks. We relax this assumption and allow for the possibility that other shocks affect government spending as well instantaneously, but we maintain the presumption that government spending shocks are the main driver at the quarterly frequency.

Table 1.5 shows the variance decompositions at various horizons and in both states. As before, shocks are ordered according to the size of their variances. Given our labeling scheme, there is little ambiguity in finding the structural government spending shock ε_t^g. It is the first shock, which explains more than 95% of the variance of government spending at all horizons and in both states. This indicates that our government spending shocks are similar to recursively identified ones, but not identical.

We confirm this impression in Figure 1.3, which compares the estimated government spending shocks to those of a linear model identified through a Cholesky decomposion with government spending ordered first. The correlation between both series is 0.85, suggesting that the shocks are closely related

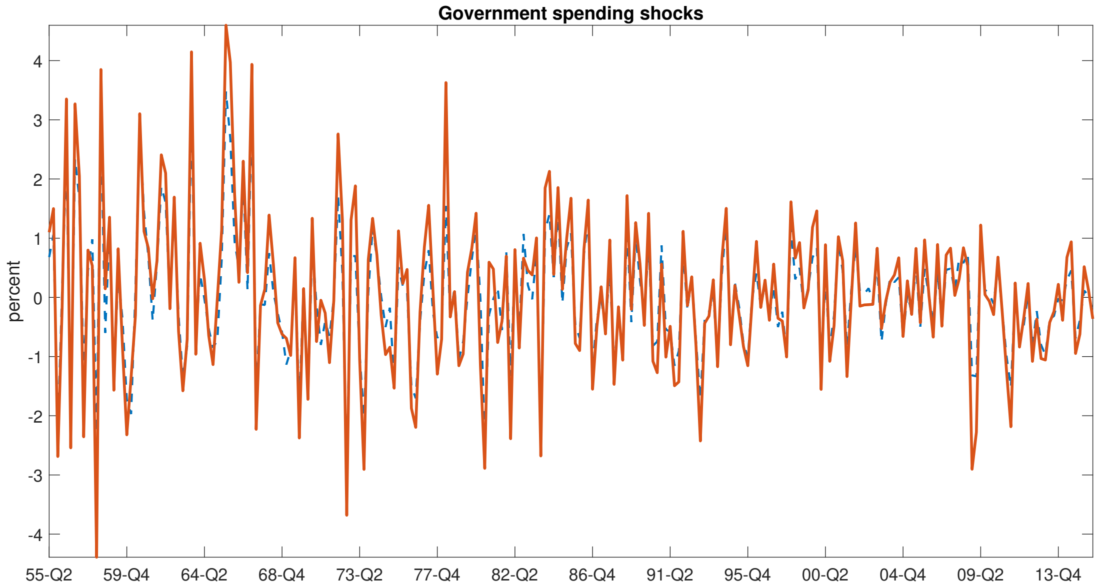
Horizon	Variable	Shock in state 1 ε_t^g	ε_t^x	ε_t^r	Shock in state 2 ε_t^g	ε_t^x	ε_t^r
1 quarter	Spending	0.96	0.04	0.00	0.93	0.07	0.00
	GDP	0.37	0.62	0.01	0.24	0.73	0.03
	T-Bill rate	0.01	0.02	0.98	0.00	0.01	0.99
4 quarters	Spending	0.96	0.04	0.00	0.91	0.07	0.02
	GDP	0.38	0.61	0.01	0.24	0.73	0.03
	T-Bill rate	0.04	0.08	0.88	0.01	0.03	0.96
12 quarters	Spending	0.97	0.03	0.01	0.92	0.05	0.03
	GDP	0.39	0.58	0.02	0.24	0.68	0.08
	T-Bill rate	0.07	0.20	0.73	0.02	0.09	0.90
20 quarters	Spending	0.98	0.02	0.00	0.95	0.03	0.02
	GDP	0.38	0.56	0.06	0.22	0.61	0.17
	T-Bill rate	0.11	0.23	0.66	0.03	0.10	0.87

Table 1.5: Forecast error variance decompositions.

but different. These differences will turn out to be important quantitatively, as we show below. The shocks show a clear pattern of heteroskedasticity, being generally more volatile during the first part of the sample and fluctuating less in the second part. This coincides with the estimated volatility states in Figure 1.2, which are mostly in the high volatility regime before the mid-1980s. The MS model accounts for the volatility changes of the underlying data.

Figure 1.3: Comparison of government spending shocks.

Notes: The figure shows the estimated government spending shocks from the unrestricted Markov switching model (solid line) and from a linear-recursive model (dashed line).



Finally, the other two shocks resemble an unspecified output shock, ε_t^x, and a interest rate shock, ε_t^r. But the exact interpretation and labeling are of no specific interest for the following. The second shock explains more than half of the variation in output in both states and at all horizons. The third shock explains nearly all of the unpredictable changes in the T-Bill rate in the first year, and at least 66% after five years.

1.3.2 Impulse responses for baseline model

Figure 1.4 shows the impulse responses. The left column presents the effects of the government spending shock on the endogenous variables in rows. Shaded areas indicate 90% confidence bands.[2] A shock of one standard deviation leads to a significant contemporaneous increase in government spending by about 0.15%. The response is highly persistent, reaching a maximum after three years.[3] The unexpected fiscal expansion has an immediate and substantial positive effect on output, which increases significantly by 0.29%. GDP rises further for the first three quarters, before gradually returning to the level where it would have been without the shock. The T-Bill rate increases significantly for the first year. Overall, the expansionary effect induced by the fiscal stimulus leads to an increase in interest rates by roughly 10 basis points.[4]

The second column summarizes the impact of the second shock, which is of interest because it shows the response of government spending to exogenous output variations. Output increases significantly by 0.38% in the same quarter the shock hits, and then in a hump-shaped manner. The peak response is at three quarters. Government spending falls upon impact. Although this effect is small and only borderline significant, the countercyclical behavior of government spending is of central importance for the size of the government spending multiplier (Caldara and Kamps, 2017). The government spending response is insignificant for the remaining horizon.

A key concept in the fiscal policy literature is the government spending multiplier. It is also of great importance for the policy debate because it gauges the effectiveness of surprise fiscal stimuli. It measures by how many dollars aggregate output increases when government spending increases exogenously by one dollar. We focus on the dynamics of the cumulative government spending multiplier, which measures the cumulative change in output relative to the cumulative change in government spending from the time of the government expenditure shock to a given period of the forecast horizon. We compute it as the ratio between the cumulative output and the cumulative government spending response as both variables are already normalized by potential output.

Figure 1.5 shows the estimated cumulative multiplier together with 90% confidence bands. The

[2]The responses are calculated for the low volatility state. As the impact effects do not vary across states, the responses in the high volatility state are the same up to scaling. The confidence bands are partially asymmetric as they are based on the empirical distributions generated through bootstrap methods instead on the asymptotic distribution of the impulse response coefficients, which are by construction symmetric but often poor approximations in finite samples. Asymmetry might reflect small-sample bias or non-Gaussianity in the errors, which are used for resampling. GDP growth, for example, is often left skewed. Interest rate changes may have fat tails, especially if volatility is time-varying.

[3]The spending responses returns to steady state at longer horizons. This rules out a permanent increase in public spending and suggests that the intertemporal government budget constraint is satisfied.

[4]As shown below, the government spending shock also leads to an increase in inflation, such that the interest rate response can be reconciled with a monetary authority following a standard Taylor-rule.

Figure 1.4: Impulse responses of the baseline specification.

Notes: The left column shows the effects of a government spending shock on the endogenous variables in rows. The right column shows the effects of an unspecified output shock. 90% confidence bands are constructed by a wild bootstrap.

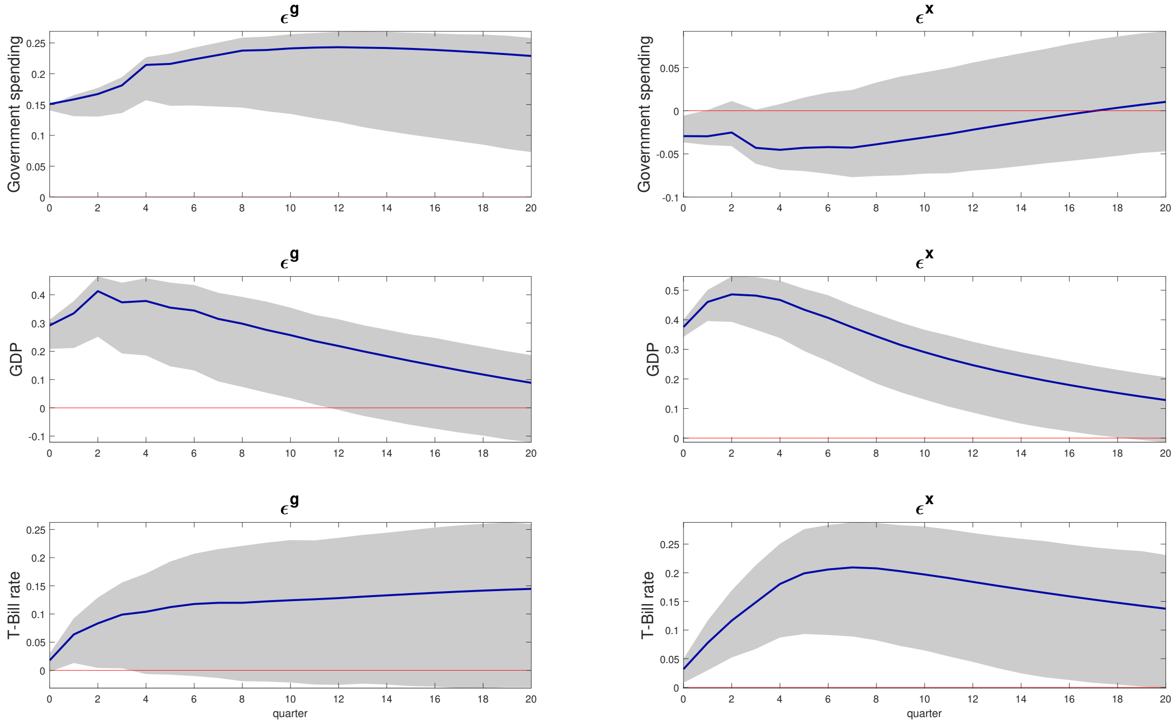
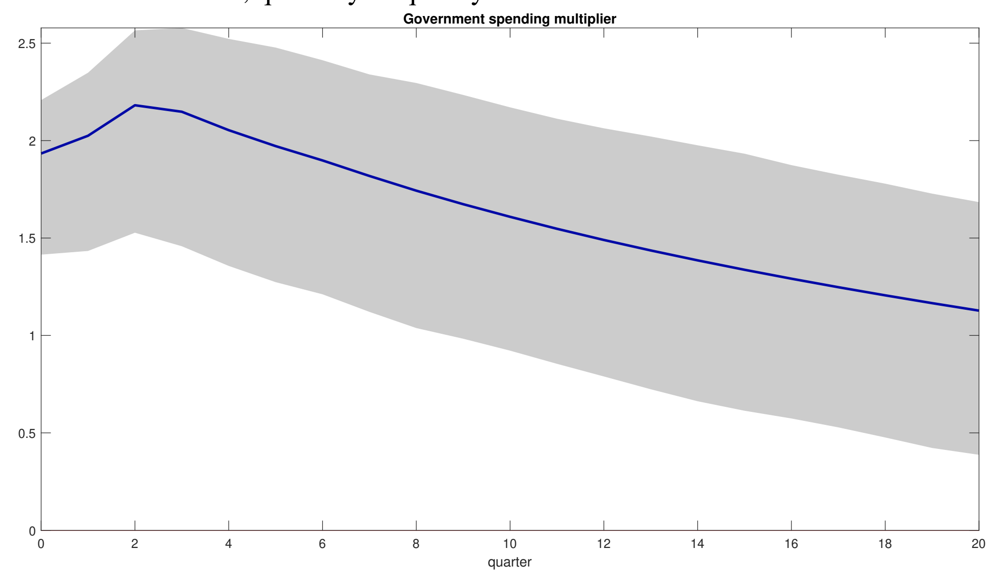


Figure 1.5: Baseline estimate of average government spending multiplier.

Notes: 90% confidence bands; quarterly frequency.



multiplier is larger than one over the whole forecast horizon, and statistically larger than unity for about ten quarters. It is 1.9 upon impact and it peaks two quarters after the shock. Subsequently, it decreases gradually and it is 1.1 after five years. Other studies find similar multiplier values (Ben Zeev and Pappa, 2017), but our estimate is larger than those typically found by applying timing-restrictions or using narrative accounts (Blanchard and Perotti, 2002; Ramey, 2011; Ramey and Zubairy, 2018). In the following, we provide several empirical reasons for this difference.

One of the reasons is based on the insights of Caldara and Kamps (2017). The authors show that the size of the multiplier depends negatively on the contemporaneous elasticity of government spending to output. The intuition for this relation can be summarized as follows. There exists a positive comovement between government spending and output in the data, which any identification approach decomposes into a fraction explained by government spending shocks and a fraction explained by the remaining shocks of the SVAR. The timing assumption of Blanchard and Perotti (2002) implies that government spending does not contemporaneously respond to any other shock. Therefore, all the positive comovement in the data must be explained by the government spending shock. This leads to a spending multiplier of about one. If the systematic response of government spending increases, the remaining shocks explain a larger part of the positive comovement and the spending multiplier decreases. In contrast, if the systematic response decreases, that is, turns negative, the multiplier increases.

Our estimate suggests a significant contemporaneous response of government spending to output of -0.08 and, thereby, a relatively large multiplier. Caldara and Kamps (2017) estimate an impact reaction of spending to output of -0.13 and obtain spending multipliers larger than one, as well.[5] They explain the negative reaction by nominal government spending being not fully indexed to inflation, implying that real government spending falls in response to an increase in inflation induced by the surprised output expansion. Thus, the negative elasticity should not be interpreted as an undertaken policy action but rather as a mechanical adjustment implied by price changes.

While Caldara and Kamps (2017) rely on external instruments to estimate the government spending to output elasticity and the corresponding multiplier, we use the heteroskedasticity in the data for estimating these relationships. We see our approach as complementing theirs. At the same time, ours gives full voice to the data, as it does not have to assume instrument exogeneity. Furthermore, it is not plagued by weak instruments problem for post-WWII data, it does not require the time-consuming construction of narrative instruments, and it can be readily applied to other countries.

1.3.3 Economy-wide effects

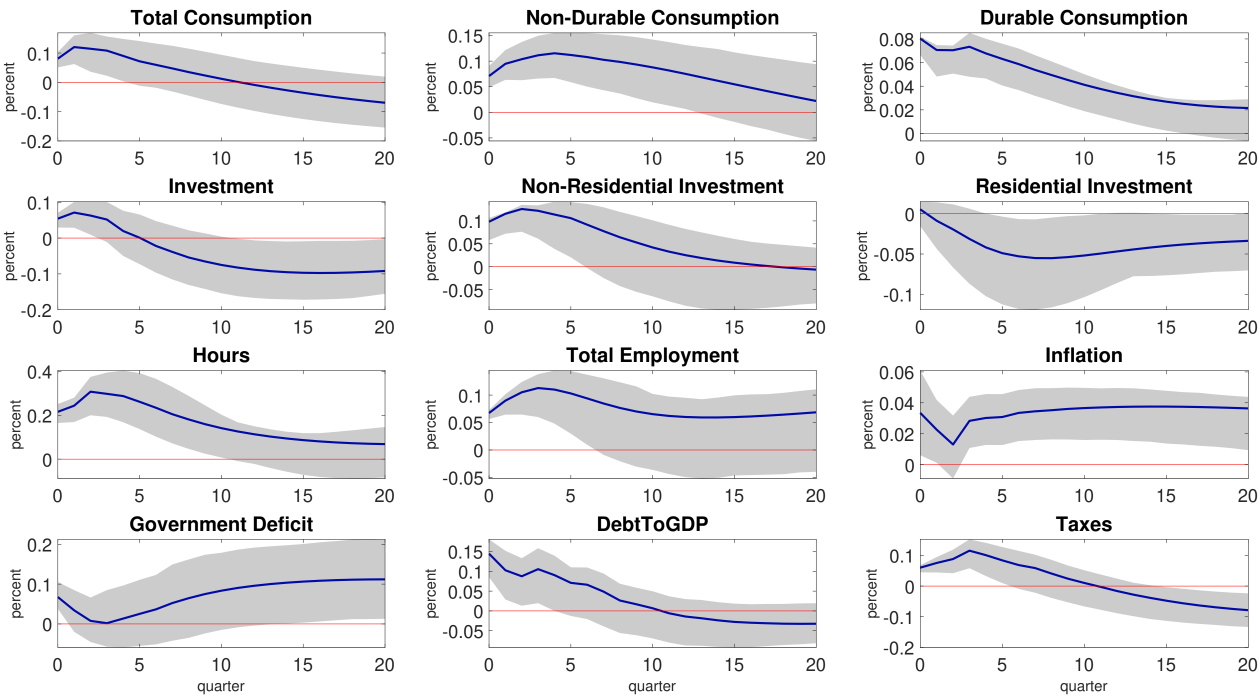
In this subsection, we analyze how government spending shocks affect consumption and investment to understand which components of private demand drive our multiplier estimates. Moreover, we

[5]Caldara and Kamps (2017) estimate an impact multiplier between 0.9 and 1.2 depending on the specific form of the underlying fiscal rule. Similar to our estimates, they find that the spending multiplier peaks at values between 1.3 and 1.7 three-quarters after the shock.

discuss the effects on hours worked, employment, inflation, and other important fiscal variables like the government deficit, the public debt-to-GDP ratio, and tax revenues.[6] For this, we augment our baseline model by one of the additional variables at a time and summarize their impulse responses in Figure 1.6.

Figure 1.6: Impulse responses of other variables of interest to a government spending shock.

Notes: 90% confidence bands constructed by a wild bootstrap.



Private consumption expenditures increase significantly. The maximum response is reached two quarters after the shock occurred. This increase in consumption is driven by a significant and quantitatively similar rise in both of its components. Households purchase more non-durable and durable goods. The increase in private consumption following an exogenous increase in public consumption is difficult to rationalize using standard textbook models of the real business cycle or the New Keynesian paradigm. In these models, a government spending increase that is financed through higher lump-sum taxes in the future induces a negative wealth effect such that households increase savings and reduce consumption expenditures. However, extensions of the baseline model, such as limited asset market participation, countercyclical markups, consumption-leisure complementarity or different monetary-

[6] We use the Ramey and Zubairy (2018) data set for the government deficit and the GDP deflator. Inflation is calculated as the quarterly percentage change of the GDP deflator. Total consumption expenditures are calculated as the sum of the FRED codes PCND, PCESV, PCDG, where, again, the first two are used to calculate the non-durable consumption expenditures and the last one, PCDG, is our measure of durable consumption expenditures. For investment we use the codes FPI and PNFI, PRFI for (non-)residential investment, respectively. Data on GDP, tax revenues, government debt, government deficit, total hours worked, and employment in the nonfarm business sector are taken from Valerie Rameys' homepage. We normalize private consumption, investment expenditures and tax revenues by potential output. Hours worked and employment are normalized by total population. The government deficit and debt are expressed relative to nominal GDP.

fiscal policy interactions might overturn these effects such that households consume more in response to a fiscal stimulus (Forni et al., 2009; Gali et al., 2007; Leeper et al., 2017). Our evidence supports these theoretical modifications of the textbook model that limit the wealth effect due to an increase in government spending.[7]

Private Investment rises significantly only in the first couple of quarters. After about one year, the response is insignificant and turns significantly negative – although only borderline – after about three years. The negative investment response at longer horizons is mainly due to a significant fall in residential investment, whereas non-residential investment is above trend for about three years and significantly so for six quarters. Although hard to reconcile with theory, the (short-lived) positive investment response is in line with the empirical evidence presented by Ben Zeev and Pappa (2017). Taken together, the large multiplier in the first year of the surprise fiscal expansion seems to be driven by a crowding-in of household consumption and firm non-residential investment.

Furthermore, the government spending shock leads to a significant increase in hours worked, which reaches a maximum after two quarters. Besides its effect on the intensive margin, the spending shock also affects the extensive labor margin. Total employment increases significantly and the response shows a similar pattern as the one of hours worked. These responses are in line with the predictions of textbook models in which households supply more labor to compensate for future tax increases due to the fiscal expansion. The spending shock also has a positive impact on inflation, which increases significantly and persistently. The inflationary effect rationalizes the persistent increase in the T-Bill rate in response to a government spending shock documented in Figure 1.4 as an endogenous response of monetary policy. The public deficit and the government debt-to-GDP ratio increase significantly, suggesting that the identified fiscal surprise expansions are mainly deficit-financed. The debt-to-GDP ratio converges back to its pre-shock level, which rules out explosive debt dynamics.[8] Although tax revenues rise slightly following the fiscal spending shock, the government spending increase outweighs the tax increase implying a higher fiscal deficit.

1.3.4 Robustness

We now show that our results are robust to various alterations of the model and the data.

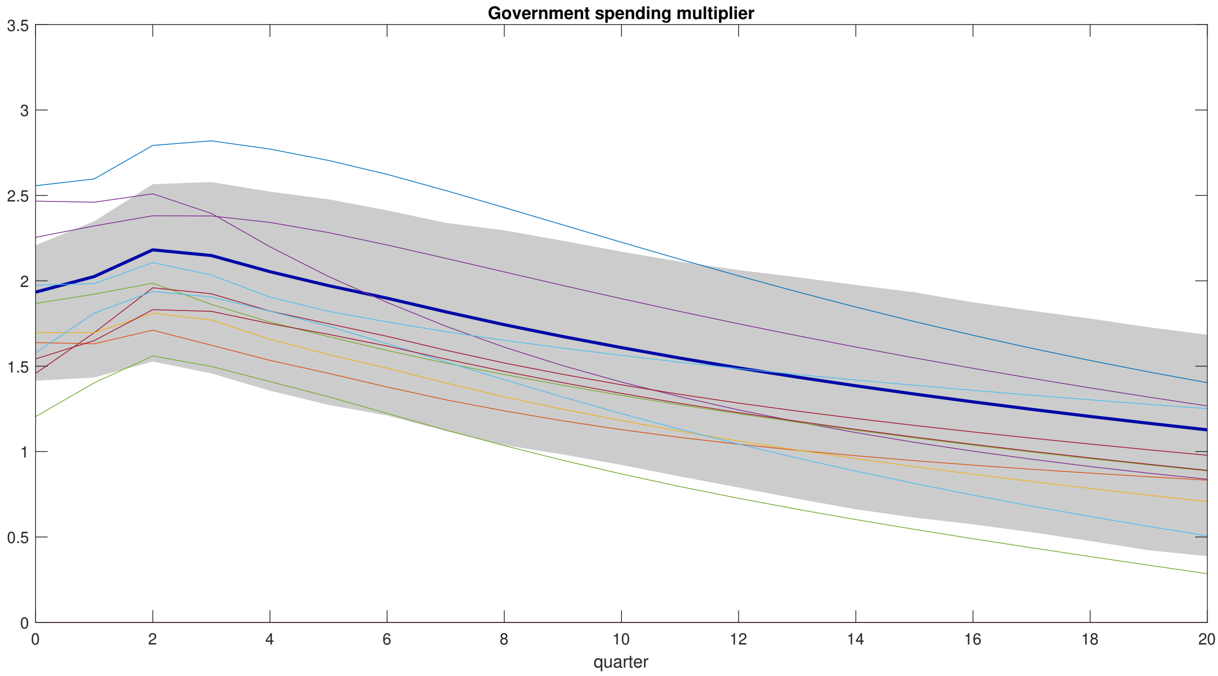
[7] In general, the response of consumption to government spending shocks is the subject of a considerable debate with different results emerging from different identification schemes based on short-run restrictions and narrative identification schemes (Gali et al., 2007; Ramey, 2011).

[8] Debt and the intertemporal government budget constraint are typically omitted from empirical investigations of the effects of fiscal shocks (Favero and Giavazzi, 2012). Chung and Leeper (2007) have analyzed an empirical model that explicitly considers the government intertemporal budget constraint via cross-equation restrictions derived from a log-linearized intertemporal budget constraint.

Augmented models. Figure 1.7 contains the multiplier corresponding to each of the models underlying Figure 1.6. It shows that adding other variables to the model leaves the multiplier largely within the 90% confidence bands of the baseline estimate. Moreover, all multipliers are larger than one, suggesting a crowding-in of private demand for at least 10 quarters.

Figure 1.7: Multipliers of augmented models.

Notes: The figure shows the point estimates for the dynamic government spending multiplier of the augmented models (thin solid lines) and of the baseline model (thick line) and the 90% confidence bands for the baseline estimate.



Tax policy. As the financing side of the fiscal stimulus might affect the size of the government spending multiplier, we control for tax policy in the estimation. In particular, the vector of observables now additionally includes real net tax revenues. The asterisked line in Figure 1.8 presents the corresponding estimate and compares it to the baseline multiplier point estimate (solid line) and confidence bands (shaded area). It shows that the main finding is not affected when controlling for tax policy, although the multiplier is slightly smaller compared to the baseline estimate. The cumulative one-year multiplier is around 1.7, similar to the value found by Caldara and Kamps (2017) who also include tax revenues as control variable.

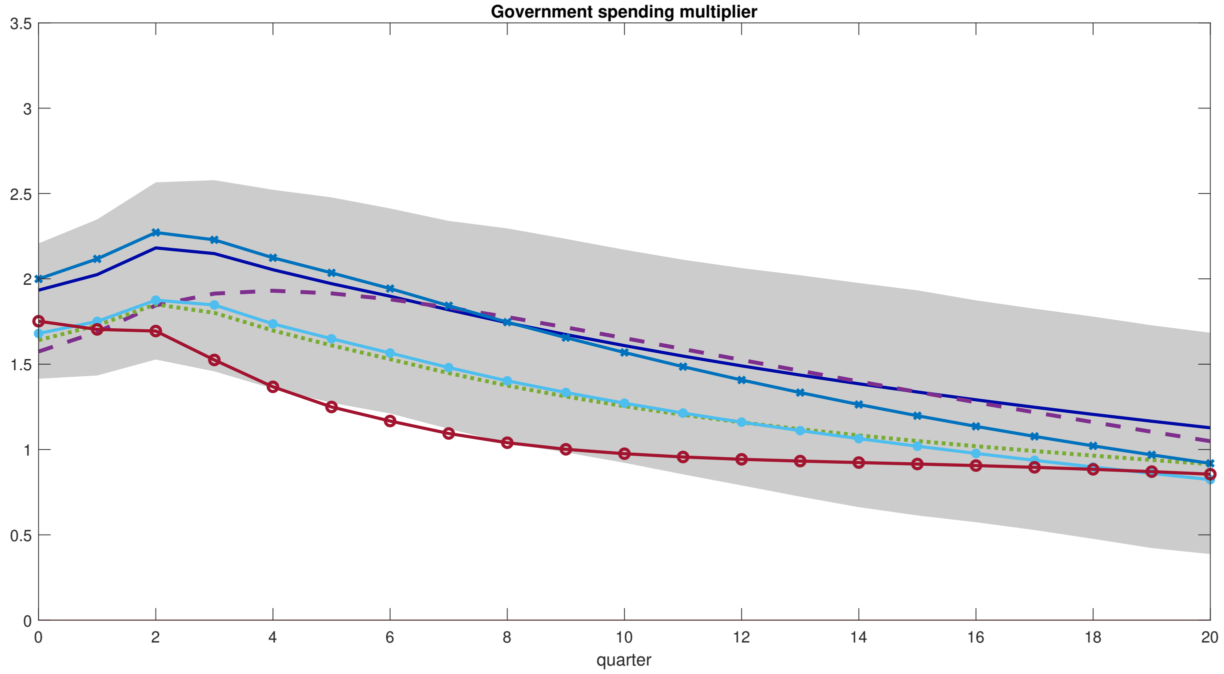
Another important estimate in this model is the output elasticity of tax revenues, that is, by how much tax revenues change when GDP exogenously increases by 1%. For the baseline sample, we estimate a value of 1.2, which is similar but somewhat smaller than the one reported by Blanchard and Perotti (2002); Caldara and Kamps (2017). Our sample differs from theirs in that it includes the Great

Recession and the subsequent years which makes a direct comparison difficult.[9]

As an additional robustness test, we re-estimate the augmented four-variable models underlying Figure 1.6, but with taxes replacing the T-Bill rate in the set of the three baseline variables that are kept constant across models. For each modified extended model, we compute the government spending multiplier and compare it to the one of the three-variable model including taxes instead of the T-Bill rate. Figure 1.13 in the Appendix 1.A shows that the multipliers are greater than one. They tend to be smaller than the ones in the specifications with the T-Bill rate. Taken together, we conclude from these two sensitivity tests, that the implicit assumption in our baseline model, which excludes taxes, that government spending is largely orthogonal to tax policy at the quarterly frequency seems to be a reasonable approximation.

Figure 1.8: Robustness of the estimated government spending multiplier.

Notes: The figure shows the cumulative spending multipliers of the baseline model (thick line) together with its 90% confidence bands and of a model that controls for fiscal foresight using forecast errors (dashed line), for the zero lower bound excluding the data after 2007 (dotted line), for taxes (asterisked line), for defense stocks excess returns (line with circles), and for a model that uses federal spending only (line with x).



Fiscal foresight. As alternative sensitivity test, we account for fiscal foresight. This can arise when private agents do not just react to actual spending increases, but also to news about impending future spending plans. Then, the econometrician might not be able to recover the true unexpected spending

[9] When we restrict the sample to the post 1986 period, we find a larger tax elasticity of 1.7. Intuitively, the Tax Reform Act implemented by the Reagan administration generally lowered tax rates and thus the revenues-to-GDP ratio such that tax revenues become more elastic with respect to GDP changes.

shock because the information sets of agents and the econometrician are misaligned (Leeper et al., 2013). We address this issue by including a variable into the baseline SVAR that captures expectations about future policy actions. Specifically, we use real-time professional forecasts for government spending, which is a spliced series of government spending forecasts provided by the Greenbook (1966-2004) and the Survey of Professional Forecasters (1982-2016). We extend the series provided by Auerbach and Gorodnichenko (2012), which covers 1966 to 2008, to include the Great Recession and the following years. Alternatively, we add the Fisher and Peters (2010) series on defense stock market excess returns to the baseline SVAR. Figure 1.8 shows that the corresponding estimated multipliers (dashed line and line with circles) are similar to the baseline estimate. In particular, they are larger than one. This analysis suggests that our baseline findings are quantitatively not strongly affected by fiscal foresight.

Zero lower bound. The baseline sample includes the Great Recession and the subsequent zero lower bound period. To investigate whether the results are robust to excluding these years, we estimate the baseline model ending the sample in 2007Q4. As the dotted line shows, the results are hardly affected. The estimated multiplier is within the confidence interval of the baseline estimate.

Omitted variables. A more general concern with the relatively small model is that it omits relevant variables and is potentially noninvertible (Stock and Watson, 2018). The limitation to smaller models is a well-known drawback of MS models, which are computationally intensive given their strong nonlinearity and treatment of the transition variable as latent. At the same time, this agnostic modeling of heteroskedasticity gives full voice to the data as it is not based on *a priori* definitions of transition variables, points or functions. Moreover, the sensitivity analysis in this subsection, which adds many control variables to the baseline model and shows that the main results are unchanged, suggests that our estimates do not suffer from omitted variable bias.

Further checks. Our main results are also robust to replacing total government spending by federal government spending. The line with x in Figure 1.8 shows that the multiplier estimate is barely affected by this change. We also report the results of the identification test for the main robustness checks in Table 1.10 in the Appendix 1.A. The tests show that the alternative 3-variables models are all statistically fully identified. For the augmented models, the tests usually reject the assumption that all 4 shock variances and a subset of 3 of them are equal, but for some pairs of shock variances does not reject their equality. Nevertheless, the impulse responses of these partially identified models are still informative (Lütkepohl et al., 2021).

1.4 Results in perspective

In this section, we compare our results to the literature and analyze the implications of our modeling choices for the size of the estimated multiplier. Our framework differs from Blanchard and Perotti (2002) and Ramey (2011) along two important dimensions. First, we use a different identification scheme, which is based on heteroskedasticity, whereas they use either a timing assumption or news on military spending as instrument. Second, our identification strategy implies that we estimate a nonlinear model that takes into account changes in the variances of the error terms, while they use linear models.

To quantify the importance of each dimension for explaining our results, we firstly concentrate on the timing assumption. It is usually justified by the fact that policy makers need time to decide on, approve, and implement changes in fiscal policy. To assess whether it is supported by the data, we set the contemporaneous response of government spending to exogenous output increases to zero, that is, we set $b_{12} = 0$ in restriction $R1$, and estimate this model against one with unrestricted impact matrix B. Table 1.6 displays the p-value of the corresponding likelihood-ratio test. The data do not reject $R1$. However, the test result needs to be treated with caution as the power of the test is unclear in our sample. Moreover, the evidence in Table 1.6 conflicts with Figure 1.4, which shows a borderline significantly negative response of government spending to output shocks. Taken together, we thus conclude that the data do not speak clearly in favor or against the timing assumption, in line with the evidence in Bouakez et al. (2014) and Lewis (2021).

Table 1.6: Likelihood-ratio tests.

Notes: The table shows the null and alternative hypothesis for tests of $R1$ and $R2$ on the B-matrix and the associated degrees of freedom (df) and p-values.

Null hypothesis	Alternative hypothesis	df	p-value
$R1 : b_{12} = 0$	B unrestricted	1	0.382
$R2 : b_{12} = b_{13} = b_{23} = 0$	B unrestricted	3	0.573
$R2 : b_{12} = b_{13} = b_{23} = 0$	$R1 : b_{12} = 0$	2	0.539

In the next step, we impose two more restrictions that help us approaching a just-identified model in a linear setting; that is, without heteroskedasticity, against which we can compare the effect of modeling changes in volatility explicitly. In $R2$, we impose additional zeros on the impact reaction of government spending and output to identify a monetary policy shock, that is, we set $b_{12} = b_{13} = b_{23} = 0$. This implies a Cholesky decomposition. The additional restrictions are motivated by a literature on the identification of monetary policy shocks (Christiano et al., 1999), which assumes that prices are rigid and that the real economy is affected by such shocks with a time lag. The data do not reject $R2$, neither against the alternative of unrestricted B, nor against $R1$ (Table 1.6).

Table 1.7 shows the estimated cumulative output multiplier corresponding to the baseline model with unrestricted B and to the model under $R1$ and $R2$, respectively, for selected years of the response horizon. In line with the argument of Caldara and Kamps (2017), the estimated multiplier decreases for all years under $R1$. The decline is substantial. The estimate for the first year drops from 2.15 to 1.43. For year 5, the unrestricted model suggests a crowding-in, while MS-R1 implies a crowding-out of private demand with a multiplier below unity. Thus, the relationship between the multiplier and the spending elasticity is highly nonlinear and particularly steep around the Blanchard and Perotti (2002) value of 0. In contrast, the two additional exclusion restrictions in MS-R2 have essentially no extra effect on the dynamic multiplier. Therefore, together with the evidence from the likelihood-ratio tests, we maintain $R2$ and estimate a just-identified linear (homoskedastic) model. The estimates shown in the next row drop further when neglecting the heteroskedasticity. The multiplier declines to one in the first year and falls below unity thereafter.

Table 1.7: Comparison of total government spending multipliers.

Notes: The table shows the cumulative total government spending multiplier for Markov switching (MS) models and a linear model for different model specifications and identification schemes. In the MS baseline model, the impact matrix B is unrestricted. For $R1 : b_{12} = 0$ and for $R2 : b_{12} = b_{13} = b_{23} = 0$.

	Response horizon in years				
Model	1	2	3	4	5
Government spending shock					
MS unrestricted	2.15	1.82	1.55	1.34	1.17
MS-R1	1.43	1.22	1.06	0.92	0.81
MS-R2	1.40	1.21	1.06	0.93	0.82
Linear-R2	1.01	0.91	0.83	0.76	0.69
Military spending shock					
MS unrestricted	0.63	0.51	0.46	0.42	0.37

This finding deserves further attention. It suggests that the estimated multiplier does not only depend on the identification scheme but also on the way the models deal with heteroskedasticity. Specifically, it seems that the generalized model, which accounts for time-varying volatility, implies a larger estimate than the linear model, which ignores heteroskedasticity. In analogy to standard regression analysis, this outcome is intuitive as generalized models weigh observations differently than homoskedastic models. While the former give more weight, or likelihood (see Eq. 4), to observations with low error variance than to observations with high error variance, the latter treat them equally. This difference not only affects inference but also point estimates.

The reasoning opens the door for two possible explanations for the larger multipliers implied by the heteroskedastic models. First, it could be that the multiplier is the same in both volatility regimes and that the larger estimates for the MS models are only due to sampling error in the high volatility regime, which erroneously implies a lower multiplier for these observations but, luckily, the estimate is down-weighted because of the associated higher uncertainty. If this is the case, it seems reasonable to use the estimates from the more efficient MS model. Lütkepohl and Schlaak (2018) show that heteroskedastic VAR models typically outperform linear models in terms of estimation precision of the structural parameters when the data feature changes in volatility.

The second reason may be that the multiplier is state-dependent. In particular, if the multiplier is larger in the low than in the high volatility regime, a Markov switching in heteroskedasticity model with state-independent impact matrix would yield a larger average multiplier than a linear model. It would give more weight to the larger multiplier in the low volatility state as this estimate is associated with lower sampling uncertainty. We investigate this possibility in the next section.

In the bottom row of Table 1.7, we estimate the total government spending multiplier associated with military spending shocks. Several studies use a narrative approach (Ramey (2011), Barro and Redlick (2011) and Ramey and Zubairy (2018)). In this literature, wars and the associated expenses are assumed to be exogenous to current economic conditions. Instead, we identify military spending shocks through heteroskedasticity. For this, we augment the three-variable MS model with military spending and compute the cumulative total government spending multiplier as the ratio of the response of output to the response of total government spending following a military spending shock. The latter is identified though the forecast error variance decomposition. It is the shock that explains 90% of the weighted impact variation of military spending across regimes.

The estimates show that the total government spending multiplier induced by military spending shocks is smaller than the total government spending multiplier generated by total government spending shocks, and below one throughout the full response horizon. As discussed in the literature, this might be due to the different nature of these shocks, which are typically not tailor-made to stimulate the economy, to the observation that large increases in military spending are often accompanied with other macroeconomic policies, like price controls, or to other confounding developments such as rising patriotism (Nakamura and Steinsson, 2014).

Finally, we compare our estimates to those of Lewis (2021) who also uses time-varying volatility to identify government spending shocks. Our point estimates broadly lie within the confidence bands documented in this study, which include multipliers around 1.5 on impact and above 2 for the one year horizon. However, our point estimates are substantially larger than the ones reported there, which are 0.6-0.7 in the first two years after the shock. Our baseline specification differs in three ways from Lewis (2021): sample, control variables, and multiplier construction. Specifically, we exclude the Korean War, include the interest rate instead of tax revenues, and normalize government spending and GDP by potential GDP.

We re-estimate our model adjusting for these differences, one at a time. Table 1.8 shows the corresponding impact multipliers. The dynamics are roughly the same across specifications and shifted proportionally with the impact multiplier. The main change in the point estimate occurs when modifying the sample, whereas controlling for tax revenues and normalizing by GDP plays a minor role.[10] The impact multiplier drops significantly when using the sample 1950Q1-2006Q4 and becomes similar to the one of Lewis (2021).

Table 1.8: Comparison of impact multipliers.

Notes: The table shows the impact total government spending multiplier for the MS(2) model with alternative endogenous variables and samples.

Specification	Impact spending multiplier
Baseline estimate	1.93
No normalization by potential GDP	1.89
Interest rate replaced by taxes	1.68
Sample 1950Q1-2006Q4	0.83
Caldara and Kamps (2017) 1954Q1-2006Q4	1.69

Perotti (2008) argues that during the Korean War not only government spending increased significantly but also the tax rate was raised considerably to balance the fiscal budget. This makes it more difficult to differentiate between exogenous spending and tax shocks. In 1953, the US economy also entered a recession during which the unemployment rate more than doubled. This is another confounding event to the military spending buildup. In addition, the increase in military spending associated with this episode appears to be a permanent one (Fisher and Peters, 2010). Military spending increased more than sixfold during the Korean War and was maintained at elevated levels thereafter. In contrast, later waves of military spending appear to be temporary fluctuations around a permanently higher level. Several studies argue along the same lines and exclude the Korean War from their sample when investigating the effects of government spending shocks (Fisher and Peters, 2010; Gali et al., 2007; Monacelli et al., 2010).

Another piece of evidence that highlights the importance of the sample is given in the bottom row of Table 1.8. There, we use the identification strategy and model of Caldara and Kamps (2017) but restrict the sample to exclude the Korean War. This entails identifying a proxy-SVAR with with non-fiscal instrumental variables to estimate the parameters of the fiscal spending rule and the government spending multiplier. The multiplier is similar to our baseline estimate with a value above unity (for the whole horizon).

Finally, the difference in point estimates between samples might also reflect an econometric issue. Our identification strategy works better on our baseline sample 1954Q1-2015Q4 than on the alternative sample 1950Q1-2006Q4. For the latter sample, we can reject the null hypotheses that $\lambda_1 = \lambda_2 = \lambda_3$ and

[10]When not normalizing by potential GDP, the multiplier is calculated as the ratio of the GDP to the government spending response times the average inverse share of government spending to GDP over the sample.

$\lambda_1 = \lambda_2$ with p-values of 0.006 and 0.063, respectively, which are similar to those in Table 1.4 for the baseline sample. But the data do not reject the assumption that $\lambda_2 = \lambda_3$ with p-value of 0.204, such that the model is statistically not fully identified. In sum, we document the important role of the Korean War episode to understand the difference between our results and the findings of Lewis (2021), and we provide arguments that may justify our complementary sample choice.

1.5 The state-dependent effects of government spending shocks

In this section, we relax the assumption of constant impact effects across volatility regimes. We allow for state-dependent effects to analyze whether the dynamic impact of government spending shocks depends on the level of volatility of an economy. This is a central implication of seminal theoretical work, which shows that demand policy is less effective when structural shock variances increase. In a situation of higher volatility, firms might postpone hiring and investment decisions (Bloom, 2009). In Bloom et al. (2018) uncertainty is modeled through a Markov process for the structural shock variances. Similarly, Bernanke (1983); McDonald and Siegel (1986) show that higher uncertainty raises the real option value of waiting before making investment decisions. Moreover, Basu and Bundick (2017); Fernandez-Villaverde et al. (2011) point out that when the time-varying second moments of structural shocks increase, there is a stronger precautionary savings motive by consumers. These theoretical models typically interpret such an increase in shock variances as an increase in uncertainty.

The Markov switching in heteroskedasticity model provides a natural way to test these prediction as it allows for time-varying shock variances. Moreover, the framework provides advantages over models with exogenously determined regimes, such as threshold or smooth transition models, as it treats the potential transition variable(s) as latent. It is more agnostic about the state determination and reduces the risk of misspecification of the transition variable, function, or points. Relative to threshold models it is more flexible as it allows for mixtures of states.[11]

1.5.1 Volatility regimes and identification of state-dependent model

To implement the dependency of the multiplier on the volatility regime, we generalize the decomposition in (3) as in Bacchiocchi and Fanelli (2015) to

$$\Sigma_u(1) = BB' \quad \text{and} \quad \Sigma_u(2) = (B+Q)\Lambda(B+Q)', \tag{5}$$

[11] A related literature uses threshold or smooth transition models to analyze whether the impact of fiscal policy depends on the slack in the economy or on the level of the monetary policy rate (Auerbach and Gorodnichenko (2012); Ramey and Zubairy (2018)).

where Q is a $n \times n$ coefficient matrix that is added to the impact matrix B. Now, the observable changes in the reduced form covariance matrix between regimes can be explained by changes in the impact effects, by changes in volatility of the structural shocks (as before), or by a combination of both. The decomposition in (3) is a special case of (5) with the restrictions $Q = 0_{n\times n}$. As this special case was just-identifying, it follows that the decomposition (5) is not (locally) unique and we need to place some restrictions on the elements of B, Q and Λ. Specifically, the decomposition contains $n(n+1)$ symmetry restrictions but $2n^2 + n$ elements, such that we need n^2 further restrictions.

These are driven by our research question. As we are particularly interested in the state-dependent effects of government spending shocks, we let their impact effects change across regimes and assume others to be state-invariant. In detail, we estimate q_{21}, which is the main parameter of interest. Under the assumption that the first shock is the spending shock, this coefficient measures the state-dependent impact of exogenous increases in government spending on output. Moreover, we allow the impact of ε_t^g on government spending to change by q_{11} when switching to state 2. As before, we estimate λ_1, λ_2 and λ_3, letting all structural variances potentially change across regimes. For ε_t^g-shocks, these assumptions imply that both their size and their relative impact on government spending and output can change across regimes.

We set all other elements of Q to zero. Moreover, we impose $b_{13} = b_{23} = 0$, based on the likelihood-ratio tests in Table **??**. In contrast, we leave b_{12} unconstrained. Although the likelihood-ratio test does not reject the zero assumption for this parameter, the confidence bands of the impulse response of government spending do (Figure 1.4). Moreover, Table 1.7 and Caldara and Kamps (2017) show that small differences in this parameter can have large effects on the size of the multiplier. To summarize, we impose the following restrictions that just-identify the structural model:

$$B = \begin{bmatrix} * & * & 0 \\ * & * & 0 \\ * & * & * \end{bmatrix}, \quad Q = \begin{bmatrix} * & 0 & 0 \\ * & 0 & 0 \\ 0 & 0 & 0 \end{bmatrix}, \quad \Lambda = \begin{bmatrix} * & 0 & 0 \\ 0 & * & 0 \\ 0 & 0 & * \end{bmatrix}, \tag{6}$$

where asterisks denote unrestricted elements. These restrictions assure that the order condition is satisfied.

To see whether these restrictions locally identify the structural model, we follow Bacchiocchi and

Fanelli (2015) and check whether the matrix[12]

$$(I_2 \otimes D_K^*) \begin{pmatrix} (B \otimes I_K) & 0_{K^2 \times K^2} & 0_{K^2 \times K} \\ 2(B+Q)\Lambda \otimes I_K & 2(B+Q)\Lambda \otimes I_K & ((B+Q) \otimes (B+Q))U_K' \end{pmatrix} \begin{pmatrix} S_B & 0_{K^2 \times a_Q} & 0_{K^2 \times a_\Lambda} \\ 0_{K^2 \times a_B} & S_Q & 0_{K^2 \times a_\Lambda} \\ 0_{K^2 \times a_B} & 0_{K^2 \times a_Q} & S_\Lambda \end{pmatrix}$$

has full rank using 100,000 matrices drawn from the uniform distribution on the interval between -10 and 10. We find that the rank condition is satisfied for all draws. Finally, we use four lags to reduce the number of estimable parameters for this highly nonlinear model and to ensure that the maximized likelihood is independent of starting values.

1.5.2 Results for state-dependent model

While *a priori* we again do not know the ordering of the shocks, we can determine them after estimation based on the decomposition of the forecast error variances. Table 1.9 shows that the ordering of the shocks is the same as before. The government spending shock is ordered first, explaining on average 92% of the variability of government spending across both states (using the smoothed state probabilities as weights) upon impact. The other two structural shocks follow with variance contributions to GDP and the T-Bill rate of 62% and 91%, respectively, on average across states.

	Shock		
Variable	ε_t^g	ε_t^x	ε_t^r
Gov. spending	0.92	0.08	0.00
GDP	0.37	0.62	0.00
T-Bill rate	0.00	0.09	0.91

Table 1.9: Weighted impact forecast error variance decomposition.

Figure 1.9 shows the smoothed state probabilities of the state-dependent model for the low and high volatility state in the upper and lower panel, respectively. The state probabilities are similar to those of the baseline model (Figure 1.2). The state-dependent model is more often in the high volatility state than the baseline model but the correlation between either of the smoothed probabilities between both models is 0.76. Again, the figure clarifies that the volatility states are not the same as recessions or the period of the zero lower bound on interest rates (although the partially overlap with these episodes). This is important for interpreting our estimated state-dependent effects of government spending shocks as both, recessions and the zero lower bound, are shown to increase the effectiveness of fiscal interventions (Auerbach and Gorodnichenko, 2012; Christiano et al., 2011), whereas higher uncertainty is hypothesized to lower it

[12] D_K^* denotes the is the Moore-Penrose inverse of the duplication matrix, S_B, S_Q, S_Λ denote selection matrices on the parameters B, Q and Λ, respectively, U_K denotes a $K \times K^2$ full row rank matrix with the property that $U_K' diag(M) = vec(M)$ for a diagonal matrix M and a_B, a_Q, a_Λ denote the free parameters in S_B, S_Q, S_Λ, respectively

(Bloom et al., 2018).

Figure 1.9: Smoothed state probabilities of state-dependent model.

Notes: The figure shows the smoothed state probabilities of the state-dependent model for the low and high volatility state in the upper and lower panel, respectively.

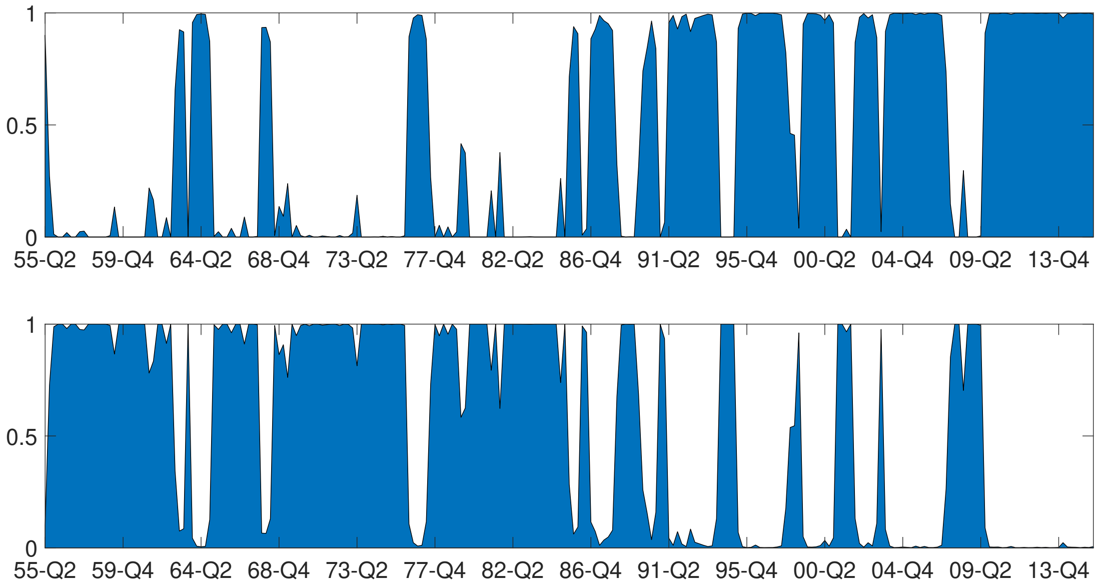


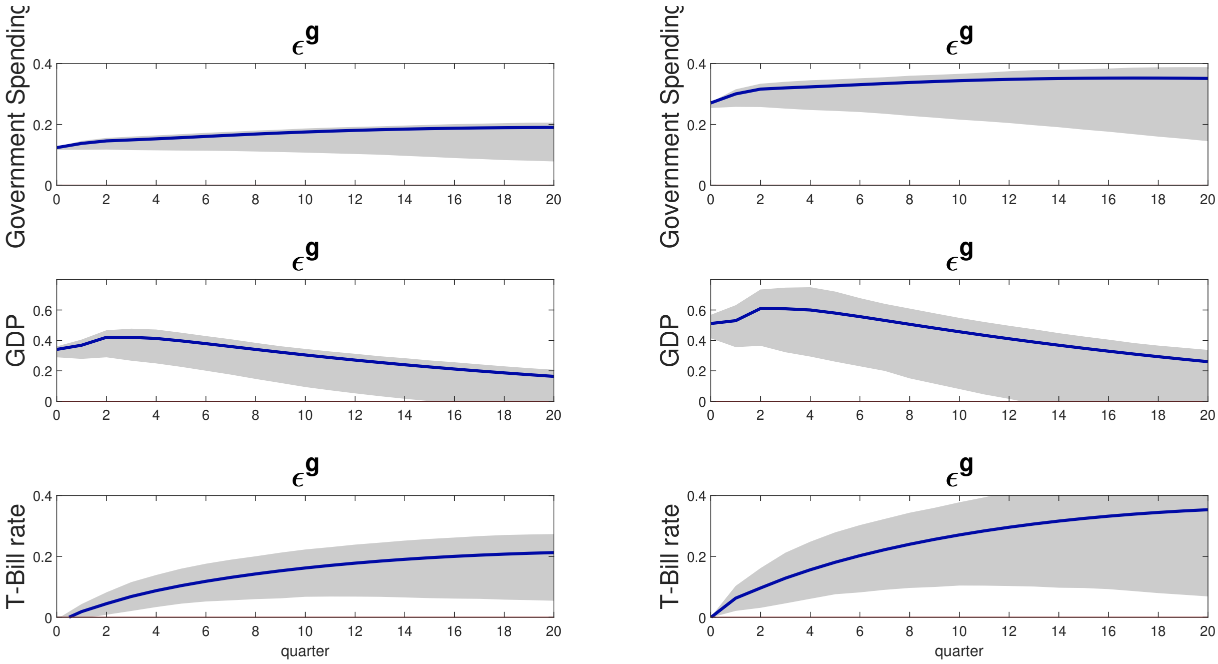
Figure 1.10 presents the state-dependent impulse responses to a government spending shock of one standard deviation in the low volatility regime (left column) and in the high volatility regime (right column).[13] Qualitatively, the dynamics are similar in both states. Government spending gradually and persistently increases. GDP responds positively upon impact, peaks at three quarters, and slowly returns to the level where it would have been without the shock. The output responses are significantly larger than zero for about three years. Similarly, the T-Bill rate increases significantly in both states. Quantitatively, however, there are substantial differences in both the absolute and relative responses of government spending and output between regimes. The increase in spending is about twice as large in the high volatility regime, consistent with the earlier evidence that government spending shocks are larger during these episodes. In contrast, the output response is only mildly larger in state two. Taken together, the point estimates suggest that the government spending multiplier is lower in the high volatility regime.

This is confirmed by Figure 1.11. The solid lines show the cumulative government spending multiplier in both regimes, together with 90% confidence bands. In the low volatility regime (left panel), the multiplier is above 2 upon impact and increases slightly in the subsequent quarters. It reaches a maximum three quarters after the shock occurred. Then, it gradually declines but is still larger than one after five years. The multiplier is substantially smaller in the high volatility regime. It is below 2 on

[13]The model produces reasonable dynamics following the monetary policy shock, which are qualitatively similar to the responses to the third shock in the baseline model. An exogenous increase in the T-Bill rate leads to a persistent decline in output and government spending. Moreover, the output shock triggers a gradual increase in the T-Bill rate, whereas government spending declines.

Figure 1.10: State-dependent responses to government spending shock.

Notes: The figure shows the responses to a government spending shock of one standard deviation in the low volatility regime (left column) and in the high volatility regime (right column). 90% confidence bands constructed by a wild bootstrap.



impact and the effect is not significantly larger than one at the end of impulse response horizon. The formal comparison between regimes in Figure 1.12 shows that the multiplier is statistically significantly larger in the low volatility regime for the full horizon.

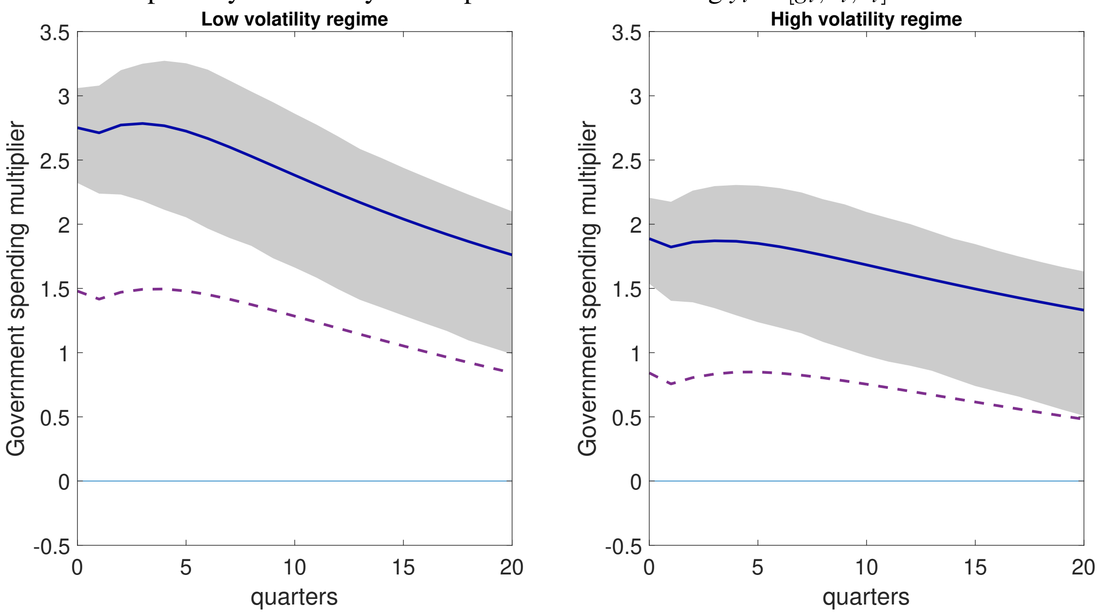
To assess the sensitivity of the state-dependency to the identification scheme, the dashed lines in Figure 1.11 show regime-specific multipliers using a recursive identification in both states. Different to (6), this identification strategy imposes the following restrictions on the decomposition (5):

$$B = \begin{bmatrix} * & 0 & 0 \\ * & * & 0 \\ * & * & * \end{bmatrix}, \quad Q = \begin{bmatrix} * & 0 & 0 \\ * & * & 0 \\ * & * & * \end{bmatrix}, \quad \Lambda = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix},$$

which imply that the change in variances is shifted to Q, that the elasticity of government spending to output is zero in both regimes, and that the response of the interest rate to spending and output shocks is state-dependent. Figure 1.11 shows that the multiplier is larger in the low than in the high volatility regime (Figure 1.12, we omit the error bands for visibility). Moreover, in both states the point estimate is below the one implied by the generalized decomposition, in line with the argument that the timing assumption of Blanchard and Perotti (2002) produces lower multipliers.

Figure 1.11: State-dependent government spending multipliers.

Notes: The figure shows the government spending multiplier in the low volatility regime (left panel) and in the high volatility regime (right panel). The solid line and shaded area correspond to the point estimate and 90% confidence bands, respectively, for decomposition (5) and the dashed line refers to the point estimate implied by a Cholesky decomposition with ordering $y_t = [g_t, x_t, r_t]'$.



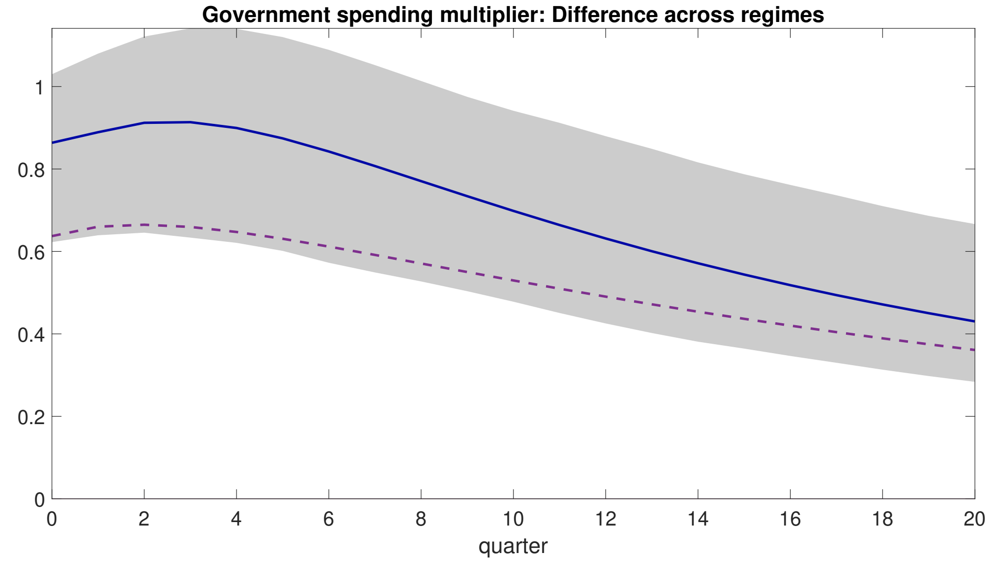
Summarizing, we find evidence for state-dependent effects of fiscal policy. The government spending multiplier is significantly larger when volatility is low, consistent with the theoretical predictions of Bloom et al. (2018). The findings also rationalize that we find larger average multipliers in the heteroskedastic than in the linear models. These differences reflect differences in the weighting of the state-dependent effects. The higher multiplier estimates for the low uncertainty regime obtain more weight in the MS models than in the linear models, producing larger average estimates.

These findings contain important policy implications. Combined with the smoothed state probabilities (Figure 1.9), they imply that fiscal policy tended to be less effective in the first part of the sample because there was a larger probability of being in the high volatility regime. In contrast, fiscal policy was more effective during the Great Moderation and the years that followed. This might be driven by the more stable policy environment in the second part of the sample. By following a well-defined and transparently communicated rule-based behavior, such a policy framework can reduce uncertainty and simplify expectation formation by private agents. When future policy interventions and output fluctuations are easier to predict, the precautionary savings motive of households is reduced and firms' region of inaction shrinks.

Moreover, because the low volatility regime also includes the period of constrained monetary policy at the end of the sample, our results support the view that fiscal policy is more effective when interest rates are at the zero lower bound, in line with recent empirical evidence by Miyamoto et al. (2018); Ramey

Figure 1.12: Difference between state-dependent government spending multipliers.

Notes: The figure shows the difference between the government spending multiplier in the low volatility regime (left panel of Figure 1.11) and in the high volatility regime (right panel of Figure 1.11). The solid line and shaded area correspond to the point estimate and 90% confidence bands, respectively, for decomposition (5) and the dashed line refers to the difference between the state-dependent point estimates implied by a Cholesky decomposition with ordering $y_t = [g_t, x_t, r_t]'$ as displayed in Figure 1.11.



and Zubairy (2018). In contrast, as high volatility states partially coincide with NBER recessions, our findings suggest that fiscal policy may be less effective during periods of economic slack.

1.6 Conclusion

We estimate the macroeconomic effects of government spending shocks in the US. We identify the shocks through an agnostic identification scheme that exploits time-varying volatility in quarterly post-WWII data modeled through a Markov switching approach.

We find that the average output effects of exogenous spending increases are positive and that the multiplier is significantly and persistently above one. This is mostly due to a crowding-in of private consumption. The estimated average multiplier is larger than multipliers estimated based on timing-restrictions or exogenous increases in military spending. There are three main reasons that explain these differences. First, we find that government spending responds contemporaneously countercyclical to exogenous changes in output. As the size of the government spending multiplier is negatively affected by the size of this elasticity (Caldara and Kamps, 2017), estimating a negative elasticity implies a larger multiplier. Second, we find that exogenous changes in defense spending produce smaller total government spending multipliers than exogenous changes in total government spending. Third, accounting for the heteroskedasticity in US macroeconomic data implies an efficient weighting of observations. The generalized model puts more weight on observations with low error variance than on observations with high error variance, producing a different estimate than models with time-constant variance and weighting.

In the second part of the paper, we exploit that the Markov switching model naturally lends itself to an evaluation of the state-dependent effects of government spending shocks. We show that the multiplier is considerably smaller in the high volatility regime than in the low volatility regime. This finding is consistent with an increased option value of waiting, or region of inaction, that reduces the effectiveness of fiscal stimulus in these periods. It supports theoretical predictions of models that imply lower general equilibrium effects of macroeconomic policy when the time-varying second moments of structural shocks increase (Bernanke (1983), Bloom et al. (2018)). Typically, such an increase is interpreted as ex ante uncertainty of about the future state of the economy (Basu and Bundick, 2017). Overall, our state-dependent findings highlight the important role of uncertainty in shaping the fiscal policy transmission mechanism.

1.A Appendix to Chapter One

Figure 1.13: Multipliers of augmented models including taxes.

Notes: The figure shows the point estimates for the dynamic government spending multiplier of augmented models (thin solid lines) and of the underlying three-variable model, which includes taxes instead of the T-Bill rate (thick line), together with 90% confidence bands for the three-variable model (shaded area).

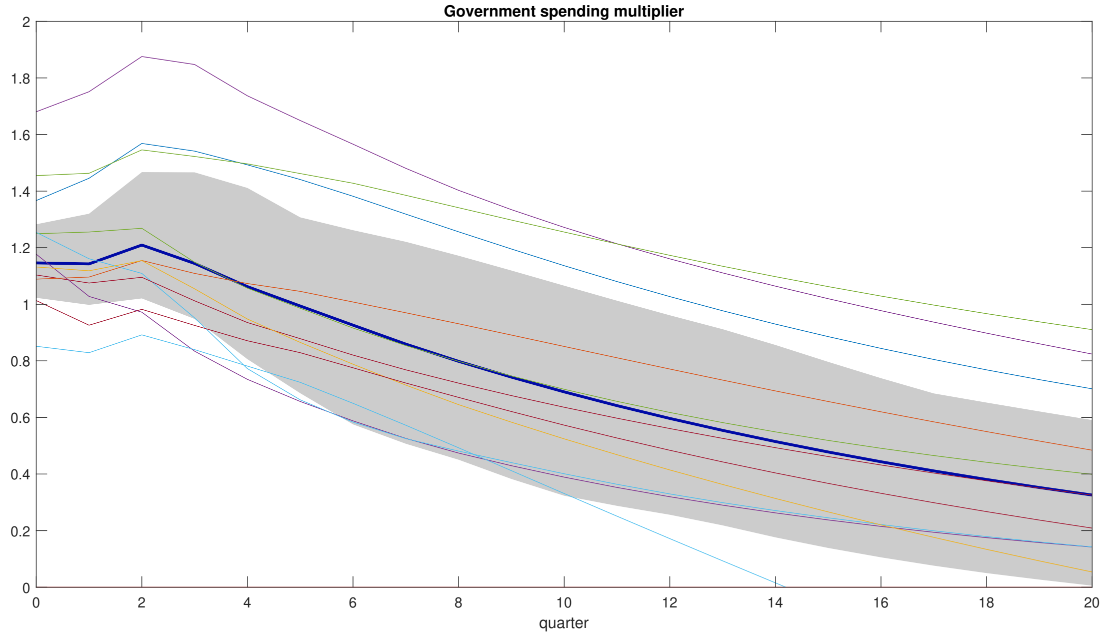


Table 1.10: Identification tests for sensitivity analysis.

Notes: The table shows the p-values for the null hypothesis that $\lambda_j = \cdots = \lambda_k$ for $j \neq k$.

Specification	H_0 : equality of $\lambda_{j,k}$				
	$\lambda_{1,2,3}$	$\lambda_{1,2}$	$\lambda_{2,3}$		
Zero Lower Bound	3.92E-07	0.002	0.015		
Federal Spending	5.15E-12	0.001	0.000		
	$\lambda_{1,2,3,4}$	$\lambda_{1,2,3}$	$\lambda_{2,3,4}$	$\lambda_{1,2}$	$\lambda_{3,4}$
Forecast errors	1.59E-20	1.84E-09	5.15E-11	0.142	0.004
Taxes	1.58E-11	1.69E-08	0.022	1.88E-06	0.180
Fisher Peters	0.000	0.407	0.001	0.888	0.022

Chapter Two

Better off without the Euro?

A Structural VAR Assessment of European Monetary Policy

This chapter is based on joint work with Patrick Christian Harms. A previous version of this project was published as DIW Discussion Paper No. 1907 in 2020 under the same name. In 2019, a previous (policy) version of this project was published under the name ”20 years of common European monetary policy: Reasons to celebrate” in the DIW Weekly Report 9 (20)

2 Better off without the Euro? - A Structural VAR Assessment of European Monetary Policy

Modern OCA theory has developed different conclusions on when forming a currency union is beneficial. An important pragmatic question in this context is: Did delegating monetary policy to the ECB increase stress in the individual euro area countries? An SVAR analysis reveals that monetary stress has declined more in the euro area than in the euro areas' doppelganger. The synthetic doppelganger is composed of other OECD countries. This result is independent of the identification strategy (sign restrictions/heteroskedasticity/Cholesky). The results can be rationalized by more formalized central banking and the euro becoming a dominant currency.

Keywords: Economic and Monetary Union, ECB, euro area, structural vector autoregressions, monetary policy stress, sign restrictions, heteroskedasticity, dominant currency
JEL classifications: C32, E42, E52, F45

2.1 Introduction

Did delegating monetary policy to the supra-national level increase monetary stress in the individual countries? Economic theory yields contradicting answers to this question. Twenty years after the introduction of the euro, this study assesses the performance of monetary policy from the perspective of the founding members of the European Monetary Union (EMU) in an empirical framework. We measure *monetary policy stress* as the variance of identified monetary shocks. The monetary shocks are deviations from stabilizing and rule-based policy from the individual countries' perspective.

The '*Impossible Trinity'* – rooted in the seminal work of Mundell (1963) and Fleming (1962) – dictates that you cannot have stabilizing monetary policy, a fixed exchange rate, and capital mobility at the same time. Following this reasoning the euro has often been characterized as a currency that impedes stabilizing monetary policy at the national level. This conclusion is premature.

Before the introduction of the euro, the European Exchange Rate Mechanism coordinated exchange rates among European countries and restricted monetary autonomy at the national level. Moreover, the presence of monetary spillovers (Iacoviello and Navarro, 2019) and the dominant role of the US dollar (Gopinath et al., 2020) are empirically well documented de facto limits for the monetary autonomy of small open economies. Consequently, choosing a free-floating regime instead of the euro, might have come at the risk of being dominated by a global reserve currency.

Stabilizing monetary policy requires an independent central bank. Today, the ECB is considered the most independent central bank worldwide (Nergiz Dincer and Eichengreen, 2014). Chari et al. (2019) show how delegating the monetary competence to a supranational institution can have beneficial welfare effects by strengthening the central bank's commitment to its mandate, even if the economies have heterogeneous macroeconomic shocks.

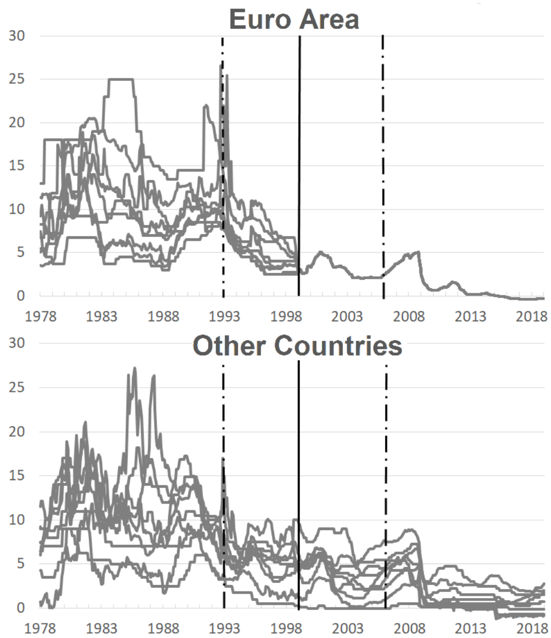
After all, there is no consensus about which of the positive and negative effects is dominant. Evidence on the performance of the ECB relative to international benchmarks is still scarce. This study aims to close this gap. We measure monetary policy stress as the variance of monetary shocks, which are defined as deviations from stabilizing policy rules. This benchmark definition of good policy as rule-based policy allows us to compare the pre-EMU sample with the post-EMU sample. Put simply, we conduct the thought experiment that since the ECB took over, it conducted monetary policy for all countries individually. This allows us to compare the performance of the national central banks prior to the introduction of the common currency with the ECB's performance thereafter.

Conceptually, our empirical approach measuring monetary stress and evaluating policy rules is related to Clarida et al. (1998), Sturm and Wollmershäuser (2008), and Quint (2016). While those studies look at reduced form residuals from single equation estimates, we identify structural shocks and use a synthetic control method to obtain a benchmark for the euro area. While there is a general trend of decreasing stress from monetary policy over time, it is more pronounced in the euro area than in the synthetic doppelganger country. This result holds even after conducting several robustness checks. In addition, we rationalize our results with regressions inspired by the dominant currency paradigm for all the countries. We find that prior to the introduction of the euro most countries' monetary policy stress was related to U.S. dollar and D-Mark fluctuations. Countries had to adjust their monetary policy according to exchange rate fluctuations, which caused monetary stress. This result vanishes for all euro area members following the introduction of the euro.

Our results are highly policy relevant for three main reasons. First, they allow to render the frequently used term 'one size fits none' as misleading.[1] Proponents of this view seem to over-emphasize

[1] This or a similar reasoning is for example used in Berger and De Haan (2002), Enderlein (2005), Enderlein et al. (2013), Sapir et al. (2015) and Wyplosz (2016).

Figure 2.1: Time series of interest rates i_t



Notes: Interest rates for euro area countries (top panel) and non-euro OECD countries (lower panel), the solid line is 1999, the date of the introduction of the common currency, the dashed lines represent the sample modifications applied in the robustness exercise.

the costs of giving up individual currencies while ignoring important favorable developments. Second, our results confirm that joining the EMU and abandoning the European Monetary System (EMS) was beneficial for most of the member countries and the average euro area country. Third, our results provide some evidence that leaving the euro or choosing an independent currency in the first place might (have) come at the cost of being dominated by the D-Mark or the U.S. dollar.
The remainder of this paper is structured as follows: Section 2.1.1 explains the chosen empirical approach, section 2.2 presents the results for the average euro area country and the individual countries as well as how monetary stress may be related to exchange rate fluctuations. In section 2.3, we show that our results are robust when we change the country sample and the time sample, employ various specifications of the doppelganger or a welfare-based measure. Section 2.4 provides possible interpretations of the results and section 2.5 concludes.

2.1.1 Empirical approach

The empirical approach in this paper tackles the question from two different methodological perspectives: First, and in line with the literature, single equation Taylor (1993)-rules are estimated and the level of monetary policy stress is calculated in a similar way to the original reference Clarida et al. (1998). Second, the factor of improvement of the monetary policy stress is discussed for the euro area and the other countries on the basis of structural residuals. The reason we add results based on identified structural vector autoregressions (SVAR) is that the measured deviations from the estimated rules - strictly speaking - are reduced form residuals. In fact, this type of stress estimate may capture demand and supply shocks instead of genuine monetary disturbances.

In the SVAR exercise, the identifying assumptions make sure that the residuals can be labeled as deviations from rule based stabilizing interest rate policy. Those results are provided for 10 euro area countries and 8 non-euro but OECD countries and are summarized using synthetic control method (SCM) doppelgangers as proposed by Abadie and Gardeazabal (2003).[2]

The research question whether individual countries would have been better off without adopting the euro from a monetary policy perspective drives our conceptual framework. The outstanding feature in figure 2.1 is at the heart of the empirical investigation: while the two panels of monetary policy target interest rates are similar for the time before 1999 (solid vertical line), the euro area countries have started to only use one policy rate for all countries afterwards.[3] By comparing the factor of change of deviations from estimated policy rules before 1999 and after we try to measure the effect of this particular structural break. This *unification* of policy rate setting can be seen as a treatment, which only euro area countries received, while the other country group is untreated.[4]

Monetary policy stress describes deviations from a policy that is stabilizing from the perspective of a prototypical closed economy. This implies that - for instance - the stability of the foreign exchange rate as a goal for monetary policy is excluded right away. The rationale for this choice is twofold: first, there is no counterpart in the ECB's policy function to the goal of exchange rate stabilization that many of the individual members followed before the monetary unification. Second there is no compelling theoretical case for combining macroeconomic with exchange rate stabilization. In fact there is a consensus view that the stabilization of the domestic economy is the primary goal of monetary policy.[5] Thus being *better off* refers to receiving relatively more domestic macroeconomic stabilization and nothing else.

2.1.2 Single equation Taylor rules

In this first part of the empirical analysis, the equation

$$i_t = r_t^n + E_t[\pi_{t+k}] + \phi_\pi(E_t[\pi_{t+k}] - \pi^*) + \phi_y y_t^{gap} + \varepsilon_{t,MP} \tag{7}$$

is estimated, from which the measure of stress $\varepsilon_{t,MP}$ can be derived. Since $\varepsilon_{t,MP}$ by definition has $E[\varepsilon_{t,MP}] = 0$, its square is used as the preferred stress measure throughout the text:

$$Stress_t = \varepsilon_{t,MP}^2. \tag{8}$$

Equation 7 is estimated using standard least squares separately for 1980-1998 (before the euro was introduced) and after (1999-2018). The sample of countries follows from table 2.1, where the euro area sample consists of Germany, Belgium, Spain, France and Italy mainly due to data availability. In the estimation logic, we follow Clarida et al. (1998): For the euro area countries who joined in 1999, we used the estimated Bundesbank rule as the policy rule for the pre-euro sample and the estimated ECB rule afterwards.[6] For all other countries, national policy rules were estimated for the whole sample.[7] The equations were estimated for all available countries separately and estimates for r_t^N and $E_t[\pi_{t+k}]$ were

[2] We use all founding members of the EMU except Luxembourg, which previously was part of a currency union with Belgium. In section 2.3, we include Luxembourg, Greece and three additional OECD-countries, which we kept out of the sample due to data quality or availability.

[3] This is where the term 'one size fits all/none' [(Issing, 2001), (Enderlein, 2005)] comes from.

[4] Here, the suspicion of 'one size fits none' would be that heterogeneous policy rates would be preferable over unified monetary policy if the business cycles and price setting dynamics are heterogeneous as well.

[5] For a theoretical discussion in the framework of equilibrium models see Gali and Monacelli (2005) and Faia and Monacelli (2008).

[6] Due to the dominant role the Bundesbank played in the EMS.

[7] For all data sources please find a precise list in the data Appendix 2.A.1

plugged in for the constant to derive the stress level.[8] In calculating $\varepsilon_{t,MP}$ for the euro area countries, we follow Quint (2016) instead of Sturm and Wollmershäuser (2008) and use the difference between the observed interest rate series from the rule-implied country-specific interest rate.[9] The results in table 2.1 reflect this approach: The factor in the table reflects the stress level before the euro was introduced

Table 2.1: Ratios of Taylor rule stress estimates

	Euro Area (weighted)	US	UK	CA
Factor	1.89	1.28	0.80	3.31

Notes: Results based on single equation Taylor rule estimates. The reported factor is $[\frac{\Sigma_{i=t_1}^{T_1}\varepsilon_{t,MP}^2/T_1}{\Sigma_{i=t_2}^{T_2}\varepsilon_{t,MP}^2/T_2}]$

over the stress level afterwards so that values larger than one point toward an improvement, while values smaller than one imply a worsening.

Based on those results using a similar methodology as Quint (2016) and Sturm and Wollmershäuser (2008) one would argue that the level of weighted monetary policy stress has decreased since the introduction of the euro. This is also true for the US - to a lesser extent - and for Canada - to a larger extent. According to this measure, only the stress level of the UK has, in fact, increased after 1999 compared to before.

The results from the single-equation analysis are broadly in line with the results in Quint (2016), which already led to the conclusion that, compared to other *federations*, euro area countries are not subject to a large level of monetary stress. However, we do not want to stop the analysis here: The residual term $\varepsilon_{t,MP}$ does not have a structural interpretation, which makes it hard to defend the interpretation as *monetary stress*. Due to the method and data availability, our sample of euro area countries is insufficient. Further, while the US, UK, and Canada appear to be sensible economies for comparison, the choice appears somewhat arbitrary. In the following analysis, we tackle these two issues by first basing our results on structural VAR models and second by broadening our country and time sample.

2.1.3 SVAR analysis

Consider the SVAR(p) model

(9) $$y_t = c + B_1 y_{t-1} + ... + B_p y_{t-p} + B_0 s_t$$

and its reduced-form

(10) $$y_t = c + B_1 y_{t-1} + ... + B_p y_{t-p} + r_t,$$

where $y_t = [\tilde{y}_t, \tilde{p}_t, \tilde{i}_t]'$ is the vector of endogenous variables consisting of the output-gap, detrended prices and detrended interest rates. The difference between the two expressions is the *structure* on B_0 and the

[8] r_t^N is taken from Holston et al. (2017) and extended with own estimates for the single euro area countries and $E_t[\pi_{t+k}]$ are backward-looking annual averages of the inflation rate before 1990 and *Ifo World Economic Survey*-data thereafter due to availability.

[9] So that $\varepsilon_{t,MP} = i_t^j - i_t^{*,j}$. Sturm and Wollmershäuser (2008) calculate the stress level as the difference between the euro area wide rule-implied rate and the country-specific interest rate implied by the same rule or $i_t^{*,EA} - i_t^{*,j}$. Quint (2016)'s approach can be extended to the sample before the euro was introduced.

fact that Σ_r is of full rank so that the r_ts are correlated across equations while Σ_s is diagonal so that the s_ts are orthogonal. While there is no general agreement on the right way to identify structural models related to monetary policy, sign restrictions and identification via heteroskedasticity are often used as alternatives to the Cholesky-ordering. We base our results on all three methods to ensure that the identification does not drive our results qualitatively.

When disentangling the effects of QE from conventional policy or when the researchers are interested in obtaining a precise estimate of the impact of unanticipated policy changes, high frequency instruments have merged as a prominent way to identify SVARs. Because those instruments are not available for all the countries, we cannot identify our SVARs in that fashion. As our research question is not centered around unanticipated shocks and the related effects of monetary stimulus for the economy this is not a major drawback. We are interested in capturing deviations from rules that aim to stabilize the economy. For example, contractionary policy with the purpose of supporting a Foreign Exchange (FX) intervention is something that we want to capture as a deviation from macroeconomic stabilization.[10] Hence, we label the shock as a monetary stress shock.

This study deliberately estimates a rule that ignores the fact that the national central banks in the EMS had to set interest rates in such a way that the exchange rate remained stable. This is needed to examine the ability of monetary policy to stabilize prices and real economic developments before and after the introduction of the euro. The error term will exactly capture the fact that national banks had to deviate from a stabilizing rule in order to keep exchange rates within the corridor. Equally, the fact that the ECB sets interest rates for the euro area as a whole is also ignored. The rules are estimated in such a way that they only contain two factors, inflation and output of the domestic economy, which are justified from a theoretical perspective (Clarida et al., 2001; Kydland and Prescott, 1977; Taylor, 1993).[11] Our approach to measuring monetary policy's ability to stabilize, enables us to compare the systems. In order to make a fair comparison, we must treat all countries equally. This is a delicate undertaking because of the heterogeneity of central bank statutes around the world. We follow the argument of Taylor (1993) that rules of central banks will eventually not be algebraically describable but some combination of inflation and output is a good approximation of most of the rules.[12] We allow for a structural break at the introduction of the euro,[13] which takes different forms depending on the model and identification we use.

The SVAR analysis in this paper is based on three different ways to obtain the structural form of the VAR, which are described in the following.

Sign restrictions At least since Canova and De Nicolo (2002) and Uhlig (2005) sign restrictions are a well established method to identify SVARs. This type of identification results in a whole set of admissible models and does not yield a consistent point estimate. We follow this general idea with a few modifications: Since we are interested in the variance estimates attached to each model we impose[14] $diag(B_0) = [1,1,1]$, so that Σ_s is not the identity matrix but carries the variance estimates of the different

[10] Since the macroeconomic trilemma dictates that a central bank can either pursue macroeconomic stabilization or stabilization of exchange rates once there are free capital markets (Obstfeld et al., 2005).

[11] Of course, interest rates are an endogenous variable in the VAR and, thus, its lagged values are also included in the reduced-form estimation. While this may be seen as a deviation from stability oriented monetary policy, Woodford (2003) emphasizes the importance of monetary policy's history-dependence, which provides a clear rationale for interest rate smoothing.

[12] Every central bank will retain a bit of leeway in order to be able to respond to particular situations with a certain degree of flexibility. For the general public and for policy makers, it is more important to understand this general approximation than the exact formula.

[13] The first observation of the second part of the sample is always January 1999.

[14] After the identification has taken place.

shocks on its diagonal. The sign patterns used for identification are summarized in the matrix

$$B_0 = \begin{pmatrix} 1 & ? & - \\ ? & 1 & ? \\ + & + & 1 \end{pmatrix}, \tag{11}$$

implying that the immediate response of the interest rate to output and inflation innovations is positive and that output indeed falls as a response to a monetary policy shock.[15] Since the set identification results must be further summarized, the median value of the set of variance estimates is used as a measure of stress.[16] When we use sign restrictions, we estimate and identify the SVAR for the period 1980-1998 and for 1999-2018, separately.

Identification using heteroskedasticity The approach of Rigobon (2003) uses the changes in the variances of the variables to identify monetary policy shocks. As we specifically want to study the changes in variances of structural innovations, this identification approach is particularly well suited to identify a monetary policy stress shock. We use the following SVAR model and estimate it using a Feasible Generalized Least Squares (FGLS) for the whole sample:

$$y_t = c_z + B_{z1} y_{t-1} + ... + B_{zp} y_{t-p} + B_0 \Lambda_z^{1/2} s_t, \tag{12}$$

with structural errors $s_t \sim N(0, I)$ and the normalization of the structural impact matrix $diag(B_0) = I$. We allow the reduced form parameters $B_{z1} \ldots B_{zp}$ to vary across the regimes $z = 1, 2$ – i.e. pre- and post-euro introduction. Furthermore, as we are interested in studying the variances of the same kind of shock across the regimes, we leave the B_0^{-1} matrix constant across time but let the standard deviation of the shocks, denoted by the diagonal matrix $\Lambda_z^{1/2}$, vary across the two regimes z of interest. The reduced form covariance matrices can be written as

$$\Sigma_{u1} = B_0 \, \Lambda_1^{1/2} \Lambda_1^{1/2\prime} B_0' \tag{13}$$

for the first state and as

$$\Sigma_{u2} = B_0 \, \Lambda_2^{1/2} \Lambda_2^{1/2\prime} B_0' \tag{14}$$

for the second state. Having a total number of 12 structural parameters in B_0, $\Lambda_1^{1/2}$ and $\Lambda_2^{1/2}$ our system is exactly identified with the 12 degrees of freedom in the two reduced form covariance matrices. Because the identification is purely driven by data and not by economic assumptions, the identified and orthogonal shocks do not have an inherent economic label. However, for our purpose, we can derive an adequate labeling. We use the forecast error variance decomposition (FEVD) to determine the shock, which is responsible for most of the variance of the interest rate. As this shock is the main driver of the uncertainty in the interest rate, it can easily be interpreted as a monetary policy stress shock. This comes close to a Cholesky ordering, where the zero restrictions enforce the same assumption on the monetary shock.

Identification using timing restrictions By using zero restrictions, it is assumed that only the interest rate reacts to the monetary shock contemporaneously and that the other variables need time

[15]Practically, we used an algorithm close to the original Canova and De Nicolo (2002) approach, which is based on Given's rotations across the space of orthogonal matrices. This results in a different number of admissible models for every application and specification of the step size of the rotations.

[16]Typically, the distributions of this parameter estimate may be interpreted as versions of the χ^2 distribution.

to factor in monetary developments; therefore, the shock always explains 100 percent of the on impact FEVD. In addition it is assumed that because of price rigidity inflation does not react to demand shocks contemporaneously. The following zero restrictions identify our system and allow us to estimate a diagonal covariance Σ_s matrix of the structural shocks for the pre- and post-euro sample separately.[17]

$$B_0 = \begin{pmatrix} 1 & 0 & 0 \\ ? & 1 & 0 \\ ? & ? & 1 \end{pmatrix} \tag{15}$$

2.1.4 Data

Frequency of the data is monthly, for the construction of the output-gap and the detrended price level - based on the price deflator - the Chow and Lin (1971) interpolation technique was used.[18] All time series are expressed as deviations from flexible trends as proposed by Hamilton (2017). The sample of euro area countries includes Germany (DE), Belgium (BE), Spain (ES), Finland (FI), France (FR), Ireland (IE), Italy (IT), the Netherlands (NL), Austria (AT), Portugal (PT). The set of non-euro OECD countries is Australia (AU), Canada (CA), Denmark (DK), Japan (JP), Norway (NO), Switzerland (CH), United Kingdom (UK), and the United States (US).[19] Our baseline sample covers 1980m12-2018m12.

2.1.5 Synthetic control and the weighting scheme

Finding a way to summarize the euro area results is simple: The ECB targets prices and supports economic activity for the currency union as a whole and does not apply a specific weighting scheme to the countries. Hence, nominal GDP weights are the most obvious choice. We chose to apply the weights based on the levels of nominal GDP from the period of the sample split, which is 1999.

For the control countries, the research question requires a more sophisticated approach since there is no obvious counterpart to the nominal GDP weights. SCM is found to be useful in macroeconomic applications (Born et al., 2019a,b). The idea is to construct a *doppelganger* of the unit under treatment and to then measure the effect of an intervention by comparing the unit of interest to the doppelganger after the intervention. In the case of this application, the variable of interest is monetary policy stress. It is common practice to add different measures to the pool of variables, which may further describe outstanding features of the unit under treatment. In this application, six additional covariates are chosen[20]: the country size, measured by nominal GDP itself, the level of central bank independence,[21] and the level of economic development, measured as GDP per capita. We also try to control for macroeconomic performance prior to the introduction of the euro by using averages of GDP growth, inflation, and the interest rate from the beginning of the sample until 1998.[22] Since the monetary stress

[17] To be consistent with the notation in the literature we re-order the vector of endogenous variables to $y_t = [\tilde{p}_t, \tilde{y}_t, \tilde{i}_t]'$.

[18] Industrial production was used to construct monthly GDP series and the CPI was used to construct monthly deflator series. For further information on the data sources please consult the data Appendix 2.A.1.

[19] In section 2.3 we exclude the period after the effective lower bound as well as the pre-Volcker period. There we also discuss results including Greece (GR), Luxembourg (LU), Sweden (SE), Mexico (MX), and New Zealand (NZ). Other OECD cannot be included due to either data-limitations and/or the fact that they adopted the euro only several years after 1999.

[20] In section 2.3 we construct the doppelganger with several more parsimonious specifications to ensure that our results remain robust.

[21] The Central Bank independence index from Garriga (2016) is based on a de jure measure of independence. The history of the index goes further back than the Nergiz Dincer and Eichengreen (2014) measure and is therefore our preferred measure.

[22] We try to match pre-sample averages instead of time series in order to circumvent any autocorrelation in the matching equation. This is also consistent with the reporting of our results, where we also focus on the pre-to-post change in the variance of the stress shock. Collapsing time series data to averages works well in SCM, when the number of states is small (Bertrand et al., 2004).

series stem from a monthly model and this part of the analysis does not have a particular interest in the monthly timing of these shocks, the SVAR variance estimate for the first part of the sample - representing pre-euro stress - is used as variable number 7. Thus, these variables describe the matrix X_0 - which corresponds to the non-EMU countries and the vector x_1 represents the euro area in equation 16.

$$\min_{w}(x_1 - X_0 w)' V (x_1 - X_0 w) \tag{16}$$

subject to

$$\textstyle\sum_{n=0}^{N} w_n = 1, w_n \geq 0. \tag{17}$$

Equation 16 reminds of a weighted least squares problem, with V being the weighting matrix. The idea of the method is to minimize the *square* distance between a set of average euro area characteristics (x_1) and a weighted counterpart of non-euro area countries ($X_0 \times w$) with respect to the *optimal* set of weights summarized by w subject to the obvious restriction that the sum of w's elements w_n is one and that all weights are non-negative. Since the elements in x_1 and X_0 are not of the same unit of measurement, the choice of V is crucial in this respect. Without prior knowledge of potential off-diagonal elements, we restrict V to be diagonal. Its diagonal elements are chosen to be $1/\hat{\sigma}_c$ of the variables, where $\hat{\sigma}_c$ is the standard deviation.

2.2 Results

This section consists of four parts: First, we display the results for all individual countries, where it turns out that all countries in the baseline sample were able to reduce the stress stemming from monetary policy. Second, we compare an average euro area country with non-euro area countries weighted according to the SCM method. The results show that the euro area average outperforms this synthetic doppelganger country across identification methods. Third, we repeat the matching for all euro area countries individually and find that most, but not all, countries separately outperform their individual doppelganger. Fourth, we provide evidence that the level of monetary policy stress before the introduction of the euro was related to FX fluctuations, which is no longer true after 1999. While this result uniformly holds for all euro area countries, this is not true for the non-euro area countries in our sample.

2.2.1 Single country results

The results from the SVAR analysis are presented in tables 2.2 and 2.3. For all countries, the ratio of the variance for the first - pre-euro - and the second - post-introduction - part of the sample ($\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$, where MP implies that this is the identified monetary component of the shocks) are reported. Again, the factor in the table reflects the stress level before the euro was introduced over the stress level afterwards so that values larger than one point toward an improvement, while values smaller than one imply a worsening.

Table 2.2: Factor of improvement of monetary stress for the individual euro area countries

	DE	**BE**	**ES**	**FI**	**FR**	**IE**	**IT**	**NL**	**AT**	**PT**
SR	10.1	21.2	26.1	3.4	24.1	21.2	45.2	29.5	4.0	127.3
IH	13.9	20.6	20.7	5.9	20.5	21.5	34.7	14.2	3.4	238.4
Cholesky	16.1	24.3	25.3	5.0	24.1	18.8	42.7	17.1	4.2	113.7

Notes: The table displays the ratio of the post-euro to pre-euro monetary stress $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for a sample of 10 euro area countries. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH) and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

Table 2.3: Factor of improvement of monetary stress for the individual non-euro area countries

	AU	**CA**	**DK**	**JP**	**NO**	**CH**	**UK**	**US**
SR	24.9	11.6	12.8	36.1	6.9	13.5	16.9	6.2
IH	19.1	15.1	9.2	27.2	7.7	7.9	19.8	7.5
Cholesky	21.3	15.7	8.1	34.3	6.9	8.0	18.3	7.1

Notes: The table displays the ratio of the post-euro to pre-euro monetary stress $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for a sample of 8 non-euro area countries. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH) and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

The first important note is that the countries exhibit a high degree of heterogeneity across all the measures. The reported factor of improvement in the level of stress stemming from monetary policy takes on very low values in countries like Finland, Austria, Norway and the US. These countries tend to have a high level of economic development and an advanced level of macroeconomic stability. At the other hand of the spectrum, we find countries such as Portugal and Italy or Japan. However, it seems to be generally unproductive to draw deeper conclusions from this type of results to answer the research question.

First, as the sample consists of 18 countries with 3 different results across identification methods, the flow of information is large. Second, the research question is on the euro area's performance compared to the pre-euro phase. Hence, the results are further summarized in the following. On the other hand, when discussing the summarized results the individual countries are helpful to identify potential drivers in the weighting scheme and to point out potential biases in this regard.

2.2.2 The average euro area country

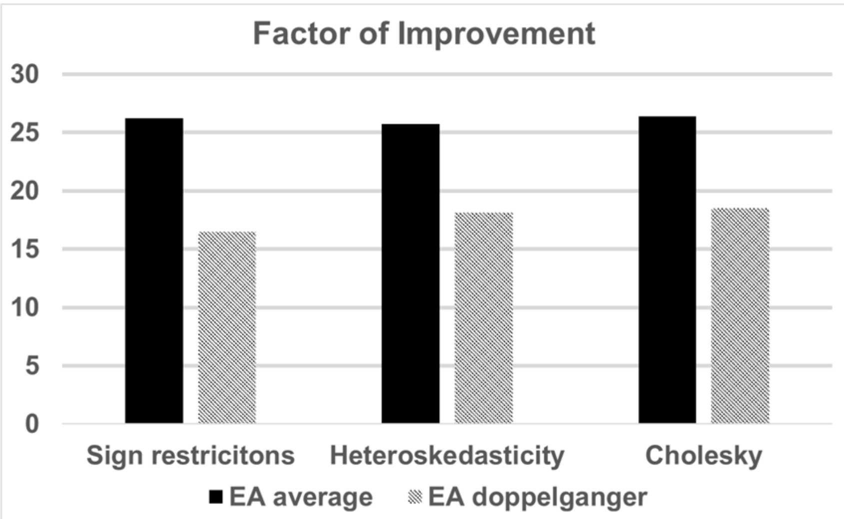
Table 2.11 in the Appendix 2.A.2 shows that the weights to replicate the average euro area country in the SCM exercise are predominantly distributed across 6 countries: The UK, which across specifications receives the highest weight, Switzerland, Norway, Canada, Japan, and - to a lesser extent - the US. Australia and Denmark, on the other hand, receive 0 weight across specifications.[23]

[23] For Australia this is because of a very high estimate for the pre-1999 level of monetary stress and for Denmark it is likely due to the combination of very low average inflation, interest rates, the small country size, and high GDP per capita.

The SCM seems to do a good job at replicating the pre-euro average for almost all metrics as tables 2.13 and 2.12 in the Appendix 2.A.2 show: Total GDP and GDP per capita are matched perfectly and similarly accurate are the estimates for the average interest and GDP growth rate. Where the method consistently fails is the central bank independence index: On average, euro area central banks seem to have been more independent than the non-euro sample. For the interpretation, this should not be a problem since - if anything - the bias in the results would go in favor of the non-euro doppelganger as lower independence, thus, leaves more room for improvement.

After the discussions of the individual country results and the empirical implementation of the SCM, we apply the resulting weights to summarize the above results.
For the important question whether the level of monetary policy stress has been reduced in the euro area, the results in figure 2.2 are consistent across identification methods.[24] While the change factor of the monetary policy stress measure takes on a value of about 25.7 - 26.4 for the average euro area country, its doppelganger country estimate ranges from 16.5 to 18.5 so that even the lowest weighted estimate for the euro area is still strongly above the value for the control country. Thus, for both country groups, we find a strong reduction in the level of stress stemming from monetary policy. Of course, this could be due to a general tendency around the industrialized world toward better central banking.[25] The doppelganger is constructed precisely to control for this type of trend.

Figure 2.2: Factor of improvement of monetary stress for the average euro area country and its doppelganger



Notes:The figure displays the post-euro to pre-euro ratio of the monetary stress measure $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for the euro area (EA) average and its doppelganger after applying the SCM country weights w_i to the individual country factor of improvement as in the tables 2.3 and 2.2. The identification assumptions are sign restrictions, identification using heteroskedasticity, and zero restrictions (Cholesky), following the recursive ordering described in equation 15. The values are displayed in table 2.14 in the Appendix 2.A.3

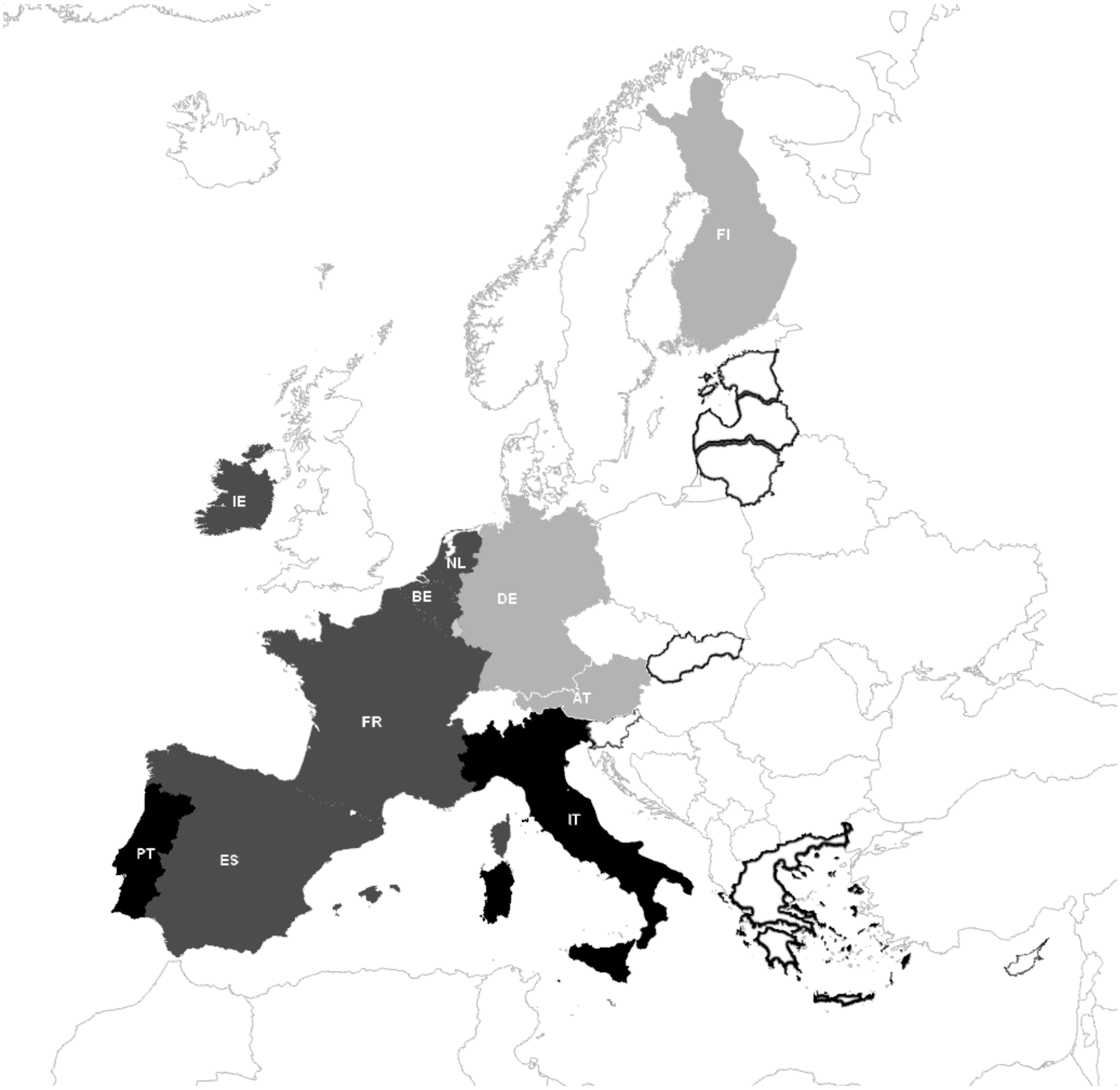
From this section, we conclude that the euro area has experienced a stronger improvement of the fit of monetary policy than the weighted control countries. Interestingly, across sample specifications, the euro area's factor of improvement is about 40 to 50% larger than in the doppelganger country.

[24]The exact numbers are displayed in table 2.14 in the Appendix 2.A.3. In section 2.3 we refer to results for a changing country and time sample. The result that the euro area outperforms its doppelganger does not depend on those choices.

[25]Svensson (2010) documents the historical convergence toward inflation targeting and Garriga (2016) and Dincer and Eichengreen (2013)'s indices clearly show an upward trend around the world, implying more independent and transparent central banks.

2.2.3 Doppelganger results from the single country perspective

Figure 2.3: Doppelganger results from the single country perspective.



Notes: The figure displays the improvements from a single country perspective relative to their doppelganger. While the countries in light grey improve but are outperformed by their doppelganger, the countries in dark grey and black outperform their respective doppelganger. The countries in black even outperform their doppelganger by more than 100 percent on average, across different identification assumptions. Moreover, the stability of these results across the three identification assumptions holds for every single county. The results are displayed in table 2.15 in the Appendix 2.A.3. The countries with thick boarders are those that adopted the euro after 1999.

While the results of the last section already provide an answer to the research question, the analysis is extended with another formulation of the problem. In particular, while the 'one size fits none' reasoning may not hold for the average country, it may very well hold for individual countries. Further, it appears of particular interest to identify those countries that drive the positive result for the average euro area. Thus, in this section, we un-do the euro area weighting and perform the same analysis from the perspective of every individual country. This allows us to compare every single country to its own *doppelganger*. For this exercise we use the same set of variables to construct the weighting matrix as in the last section. Figure 2.3 and table 2.15 in the Appendix 2.A.3 show the results.

A general remark is required for the results in this section. While the SCM method worked well for the purpose of replicating the average euro area country, its performance is weaker for each individual country.[26] However, there is one takeaway from this set of results: there is a small group of countries that, across identification methods, exhibits a lower factor of improvement than its doppelganger: Austria, Finland, and Germany. All other countries outperform their doppelganger. Table 2.15 in the Appendix 2.A.3 shows that this margin is already large for the Netherlands and Spain. Italy and Portugal double the performance of their doppelganger. Belgium, France, and Ireland still outperform their SCM counterparts, but by smaller margins. Thus, the group that did not perform better than its individual doppelganger only includes northern or core countries of the euro area.[27] These results hold true across all three identification assumptions.

2.2.4 Exchange rate fluctuations and monetary stress

In this section, we examine the relationship between monetary stress and exchange rates. Figures 2.5, 2.6, and 2.7, in the Appendix 2.A.4, show time-varying coefficients for the relationship between country-level monetary stress and the exchange rate between national currencies and the D-Mark, national currencies, the euro, and the U.S. dollar. Formally, we employ a Kalman-Filter as in 18 and 19.[28]

$$\omega_t^{MON-POL} = \beta_t \Delta E_t + v_t \tag{18}$$

$$\beta_t = \beta_{t-1} + \eta_t. \tag{19}$$

Figure 2.5 presents the time-varying relationship between national currencies and the D-Mark prior to the introduction of the euro. In all cases - except for Austria and the Netherlands - they are significantly different from zero for extended periods of time.[30] This result implies that monetary policy prior to 1999 reacted to D-Mark movements in a way that is unrelated to national price and output stabilization. Figure 2.6 repeats the exercise for all euro area countries, but now with the nominal U.S. dollar exchange rate. Extending the sample to the time period after 1999 shows that while the U.S. dollar has had an impact on most countries' monetary policy before 1999, no such effect is found after 1999.[31] This result implies that since the ECB conducts monetary policy, the dollar's influence on monetary policy stress is no longer observed and statistically insignificant in all countries. Thus, we can conclude that the joint currency provided some additional freedom from external influences following the introduction of the euro. Figure 2.7 shows that this is not the case when the exercise for the U.S. dollar is repeated for the non-euro area countries. At least in Australia, Norway, and Switzerland there is an influence of the dollar exchange rate on national monetary policy even after 1999.[32] Interestingly, an influence from the dollar

[26] Note that it is generally easier to match any mean observation compared to individual observations that are not located at the center of a given distribution.

[27] Here, the qualification is particularly important since for Germany - for instance - the SCM method performed poorly for important measures such as the level of central bank independence.

[28] Equation 18 is the observation equation and 19 is the state equation. The time-varying parameter β_t links the observed monetary policy shock $\omega_t^{MON-POL}$ to the exchange rate,[29] which we express as the first difference of the log, since nominal exchange rates are known to be very likely integrated of order one (Meese and Rogoff, 1983). Additionally, equation 19 shows that we assume that the process for generating the time-varying parameter follows β_t a random-walk.

[30] The fact that one cannot show a relationship between Austrian and Dutch monetary policy to the D-Mark exchange rate is due to the very strong relationship to the D-Mark, showing almost no variance in the nominal exchange rates.

[31] Note that for the countries that lost the competition against their individual doppelganger, the impact of the dollar is relatively small: In Germany, the effect decays after the reunification, when most of the international influence was lost, for Austria and Finland the effects are insignificant and small throughout the whole sample.

[32] While Switzerland has publicly announced exchange rate targets in the recent past, Bergsten and Gagnon (2017) count Norway as one of the most prominent currency manipulators globally.

on monetary policy is also observed for the UK. This section shows that the euro and the centralization of monetary policy free many countries from their obligations to take exchange rates into account when conducting monetary policy. The fact that this result holds for all euro area members, but not for all other countries, is evidence that this is a genuine advantage of joining the common European currency.

2.3 Robustness

2.3.1 Sample adjustments

The baseline sample covers two major economic crises, which are particularly important for the analysis. First, the global financial crisis led to a global decrease in policy rates, in many cases very close to the effective lower bounds. Second, in 1992 the EMS experienced a major crisis[33], which caused Italy and the UK leave the EMS. Moreover, the EMS crisis triggered some reforms of the EMS and its member states. Thus, in a first robustness exercise, the sample only covers 1993m1-2006m12 to exclude both incidents. This period has the additional advantage of a broad consensus about the goals of monetary policy and that Taylor (1993)-type inflation targeting was broadly established. The results are reported in tables 2.16-2.18 of the Appendix 2.A.5. While the per country results are more heterogeneous and most countries even experience a decrease in one of the three identified models, the average euro area country still outperforms its doppelganger by a 55-86 percent margin. For the short time sample, data is also available for Greece (GR), Luxembourg (LU), Sweden (SE), Mexico (MX) and New Zealand (NZ). Tables 2.19 and 2.20 in Appendix 2.A.5 report the results of the analysis with the increased country sample from 1993-2006. The euro area outperforms its doppelganger with an even greater margin. As Portugal is the only country that outperforms the average euro area by an order of magnitude, we make sure that this does not drive the results and exclude it in the calculation of the average euro area country in table 2.21 in Appendix 2.A.5. From this section we conclude that it is not a specific choice of the country sample, the time sample or a potential statistical outlier that drives the results.

2.3.2 Alternative doppelganger construction

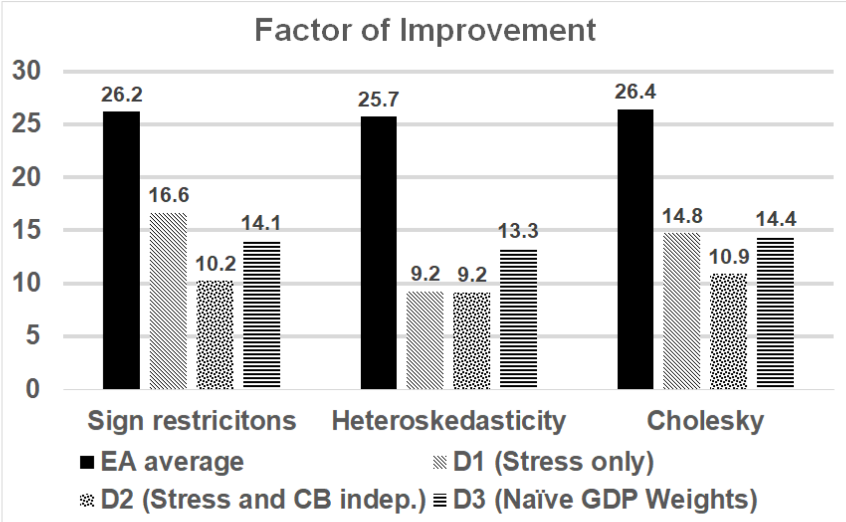
The doppelganger in our baseline specification is constructed matching six additional covariates, apart from the stress measure. Figure shows that our results remain valid for more parsimonious estimations of the doppelganger. The specifications of the doppelganger D1 and D2 are motivated by matching only variables that are tightly related to monetary stress, which is our measure of interest. While the doppelganger D1 matches only the monetary policy stress prior the introduction of the common currency, D2 includes also the independence of the central bank. Doppelganger D3 is constructed using a naïve weighting, analogously to the average euro area. It represents the average (gdp-weigthed) non-euro area country. The alternative doppelgangers have a tendency to be outperformed by the average euro area country by a even greater margin, than the baseline results.

2.3.3 Alternative measure

In theoretical models monetary policy is often evaluated according to its effects on welfare. In this context, welfare losses induced by certain outcomes that result from the objectives and the rule implemented by the central bank are expressed as a loss function. Excess inflation and excess output fluctuations are inefficiencies in the New Keynesian literature and thus reduce welfare. From this type

[33]Often referred to as *Black Wednesday*; see also the Appendix 2.A.7.

Figure 2.4: Factor of improvement of monetary stress for the average euro area country and three alternative doppelgangers



Notes: The figure displays the post-euro to pre-euro ratio of the monetary stress measure $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for the euro area (EA) average and the alternatives doppelgangers D1-D3 after applying the country weights w_i to the individual country factor of improvement as in the tables 2.3 and 2.2. The identification assumptions are sign restrictions, identification using heteroskedasticity, and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

of welfare analysis, an optimal outcome can be derived and a rule can be designed, which approximates this outcome as close as possible. Galí (2015) uses a loss function of the form

$$L = \frac{1}{2}\left[\left(\sigma + \frac{\phi+\alpha}{1-\alpha}\right)var(\tilde{y}_t) + \frac{\varepsilon}{\lambda}var(\pi_t)\right], \tag{20}$$

to evaluate simple policy rules, where the parameters stem from a simple representative agent model.[34] The variance of the output gap and of the inflation rate both induce welfare losses with respect to the optimal outcome.

As an additional robustness check, we have repeated parts of our analyses based on L. Instead of using monetary policy stress derived from econometric models, we use the observed welfare losses through the lens of the loss function.. Even though the concepts are different - monetary policy stress represents deviations from empirically estimated rules with no judgment about monetary policy optimality and welfare while L is directly related to a theoretical welfare concept - the comparison should lead to similar results. It is fairly obvious that other forces than monetary policy are probably at play when determining the variance of output and inflation such as supply, demand and technology shocks. This is precisely why the variance of the identified monetary policy shock has been our preferred measure so far. Also it makes clear what the SCM method is important for the comparison as it offers the possibility of controlling for potential trends in the model variables.

Figure 2.8 in the Appendix summarizes the results from the repeated SCM exercise based on the welfare losses derived from the loss function above.[35] For both samples, the figure tells the same story: the average euro area - EA 10 and EA 12 - country outperforms its doppelganger - based on 8 or 11 non-

[34] In the loss function L, $\Theta = \frac{1-\alpha}{1-\alpha+\alpha\varepsilon}$ and $\lambda = \frac{(1-\theta)(1-\beta\theta)}{\theta}\Theta$. Table 2.22 in the Appendix 2.A.6 summarizes the parameter values from the reference and their interpretation.

[35] L is a period-by-period loss function. In our empirical analysis we used the variance of inflation and the output-gap the two sample periods.

EMU countries - in both comparisons.
From this robustness check we conclude that even without a stochastic model, euro area monetary policy has improved by more than a synthetic doppelganger and we are still unable to detect evidence in favor of the one-size-fits-none reasoning.

2.4 Interpretation

The empirical exercise delivers four important results, which this section puts into context. First, we find a worldwide tendency toward better monetary policy. Second, the average euro area country outperforms its doppelganger. Third, despite some heterogeneity, individual countries mostly outperform their doppelganger. Fourth, the deviations from the policy rule in the euro area are not correlated with the foreign exchange rate.

2.4.1 The general tendency toward better monetary policy

The professionalization and formalization of monetary policy between the 1970s and 1990s clearly explain the overall trend of better monetary policy. Trivially, central banks are likely to have become better in monetary policy implementation over time. Clarida et al. (1998) offer the failure to accurately forecast reserve demand as a potential interpretation for monetary policy shocks. That is, whenever a central bank has a problem with setting its *operational target*, which correctly represents its monetary policy stance, this would show up as an unexpected innovation, which is orthogonal to the inflation and output-gap in our SVAR model. Bindseil (2014) and Bindseil (2016) argue that by 2007, monetary policy implementation approaches by most central banks were "well-focused and transparent compared to the 1920-1990 period."
(Svensson, 2010) provides an overview of how central banks adopted explicit goals for inflation over time.[36] While the Banca d'Italia ended being a *branch* of the Italian treasury in 1981 (Passacantando, 2013), it took the Bank of England until the Blair years in 1997 to become independent (Andréadès, 2013). All three - professionalization, independence, and the adoption of explicit targets - will push a central bank toward a strategy that brings it closer to following the objectives of stability of inflation and/or output. In the SVAR-model we use, this would imply that anything unrelated to the *new* objective of stabilization - for example interest rate setting in order to support the treasury - would end up as a residual in the reduced form, ultimately implying a higher variance of the identified shock.
In most estimated policy rules (such as those used in Clarida et al. (2000)), some form of the inflation target or long-term inflation expectations are incorporated in the intercept terms of the policy function. In our estimation, we assume that the inflation target and long-term expectations are stable throughout the sample periods. However, there is evidence that in the pre-1999 period, this assumption might be violated (Bomfim and Brayton, 1998; Cogley and Sargent, 2005). This would show up as unexplained variance in the VARs in the pre-1999 period.

2.4.2 The advantages of adopting the euro

Within the global trend, the average euro area country has performed better than its doppelganger. As discussed, and despite some heterogeneity across the countries, the factors driving the global trend seem to be particularly strong before 1999 both within and outside the euro area. Therefore, the SVAR should

[36] He counts New Zealand as the first country to embrace explicit inflation targeting (1989/1990).

be a fair approximation for all the countries in our sample and for some individual heterogeneity to be averaged out. The typical 'one size fits none' reasoning is that the unification of interest rate setting is problematic because national central banks were abolished while heterogeneous business cycles and inflation rates prevailed. Implicitly, this reasoning implies that monetary policy before the introduction of the euro was indeed designed to stabilize national business cycles. If those assumptions were correct, improvements in the fit of monetary policy should not have taken place; in particular compared to non-euro area countries.

The empirical results from the last section challenge this reasoning. In fact, one can find empirical evidence against many of the assumptions of the 'one size fits none' reasoning: Campos et al. (2017) assess the synchronization of business cycles across the world. They find that business cycles generally more synchronized since 1999 than before and find a significantly stronger tendency toward convergence in the euro area. Similarly, Franks et al. (2018) provide empirical evidence for a high degree of convergence of inflation rates in the euro area.

The 'one-size-fits-none'-reasoning does not take into account that some countries had to respond to foreign exchange developments because their currencies were dominated by the D-Mark or U.S. dollar in the sense of Gopinath et al. (2020). In particular, the EMS de jure and de facto constrained euro members - by a varying degree - in their ability to implement monetary policy according to their national needs. In fact, many decisions to change interest rates can be traced back to either the stabilization of the exchange rate system or political reasons.[37] A very homogeneous result across the euro area countries is that none of them experienced monetary stress because of exchange rates after adopting the euro. This is evidence that creating a common currency offered protection against being dominated by the D-Mark or the U.S. dollar. Nevertheless, in terms of conduct of monetary policy, some countries might have benefited more than others.

2.4.3 Heterogeneity of the single country estimates

The results from a single country perspective are well in line with the predictions of Chari et al. (2019), who argue that while a subset of countries might have joined the EMU in order to obtain more central bank independence, others might have profited from the improved coordination. In particular, Germany might have had motives beyond solely improving its already well-functioning monetary policy. Germany's persistent current account surplus is likely to be associated with its persistent decline in the real effective exchange rate since the introduction of the euro.[38] Table 2.2 and the application of the SCM in table 2.15 reveals that there is heterogeneity in the absolute improvement and that there may be some heterogeneity in the relative improvement of monetary policy fit in the euro area.

To a large extent, the heterogeneity in absolute improvements of the euro area countries reflects the state of development of the national economies and, in particular, their monetary authorities and their position in the EMS. Austria, Finland, and Germany tend to have relatively low factors of improvement while France, Italy, Portugal, and Spain have relatively high factors of improvement. In particular, Austria and Germany appear to have had a level of monetary policy quality already before the euro's introduction, which was unmatched in the whole sample of OECD countries. This is reflected in the failure of the doppelgangers to replicate the data in terms of central bank independence, average inflation, the average nominal interest rate, and the pre-euro stress estimate itself. Thus, when the factor of improvement of those countries is compared to their individual doppelgangers, the comparison is unfair due to the general trend toward better monetary policy making observed around the world.

At the same time, the positive performance of other countries - such as France, Italy, Portugal, and Spain

[37] Appendix 2.A.7 provides a short description of the mechanisms and the history of the EMS.

[38] Engler et al. (2014) discuss how the euro area countries can create and offset such imbalances in a currency union.

- compared to their doppelganger countries can be seen as just the other side of the same coin: The introduction of the euro allowed those countries that had no chance to implement independent monetary policy in the EMS to participate in the formalization and improvement trend in monetary policy making. Those countries - constrained by their inferior position in the EMS (Giavazzi and Giovannini, 1987) - simply had much more to gain from an improvement in central bank policies than those countries that were already able to implement inflation targeting-type policies in the past.

2.5 Conclusion

Increased central bank credibility, the conduct of more rule-based policy and becoming a global reserve currency have made the euro a success. We identify a global trend of declining monetary stress due to more formalized, transparent, and experienced monetary policy. Within this trend, the average euro area country outperforms its non-euro area doppelganger.

Following its creation, the common currency protected all euro area countries from receiving monetary stress due to foreign exchange fluctuations. This is not true for all our benchmark countries. In Australia, Norway, Switzerland and the UK, US-Dollar fluctuations still correlate with monetary stress. We interpret this as evidence that the beneficial effects of the common currency prevail and delegating monetary policy to the ECB did not cause stress.

The interpretation that the countries *lost* their individual interest rates to stabilize the economy is not consistent with our findings for two reasons. First, the leeway to stabilize the economy was small prior to the introduction of the euro, as the countries had to import the monetary policy of other countries and set interest rates according to the needs of exchange rate stabilization. Second, the reasoning neglects the positive aspects of central bank coordination/commitment and the size effect of the euro, which is studied by Chari et al. (2019) and Gopinath et al. (2020). Our results are robust across time samples, country samples, and identification strategies. For the individual countries only Austria, Finland, and Germany are outperformed by their doppelganger. However, these countries had little leeway to improve their central banks performance. Moreover, Germany, in particular, might not necessarily have joined the euro area to improve its monetary policy but rather to achieve a higher level of real exchange rate stability.

The euro area would benefit from a constructive discussion on how to prevent future crisis and further synchronize business cycles. A more stable union may further ease the conduct of monetary policy. Therefore, a common debt instrument, stronger automatic stabilizers, as well as the completion of the banking union and the capital markets union should be prioritized in policy discussions. Further, policymakers should be aware of the importance of improved central bank credibility, the conduct of rule-based policy and the dominance of the euro, when considering joining or leaving the currency union.

2.A Appendix to Chapter Two

2.A.1 Data Appendix

This Data Appendix describes the complete Data Sources used in all sections or subsections in this paper.

Data for section 2.1.2

Output-Gap and Inflation data follow from the quarterly series reported in Appendix 2.A.1 (tables 2.5 and 2.5). The quarterly interest rates are quarterly averages for the particular countries also in Appendix 2.A.1 (table 2.6).
Natural Interest rates follow from Holston et al. (2017) and can be downloaded from https://www.newyorkfed.org/research/policy/rstar. For the unavailable countries (single Euro Area countries) the same codes were used and calculated by the authors of this paper.
Inflation Expectations data are from Boumans and Garnitz (2019). For the time before the start of the sample therein, moving averages of 4 quarters of past inflation were used to approximate adaptive expectation formation.

Data for section 2.2.1

Table 2.4: Sources for quarterly real GDP time series

Measure: Real GDP		**Frequency: Quarterly**	
Country	**DS Mnemonic**	**Source**	**Comment**
Austria	OEOEXO03D	OECD Quarterly National Accounts	2015=100
Belgium	BGOEXO03D		
Finland	FNOEXO03D		
France	FROEXO03D		
Germany	BDOEXO03D		
Greece	GROEXO03D		
Ireland	IROEXO03D		
Italy	ITOEXO03D		
Luxembourg	LXOEXO03D		
Netherlands	NLOEXO03D		
Portugal	PTOEXO03D		
Spain	ESOEXO03D		
Australia	AUOEXO03D		
Canada	CNOEXO03D		
Denmark	DKOEXO03D		
Japan	JPOEXO03D		
Mexico	MXOEXO03D		
New Zealand	NZOEXO03D		
Norway	NWOEXO03D		
Sweden	SDOEXO03D		
Switzerland	SWOEXO03D		
United Kingdom	UKOEXO03D		
United States	USOEXO03D		

Table 2.5: Sources for quarterly nominal GDP time series

Measure: Nominal GDP		**Frequency: Quarterly**	
Country	**DS Mnemonic**	**Source**	**Comment**
Austria	OEOEXA03B	OECD Quarterly National Accounts	Current Prices, Annual Levels
Belgium	BGOEXA03B		
Finland	FNOEXA03B		
France	FROEXA03B		
Germany	BDOEXA03B		
Greece	GROEXA03B		
Ireland	IROEXA03B		
Italy	ITOEXA03B		
Luxembourg	LXOEXA03B		
Netherlands	NLOEXA03B		
Portugal	PTOEXA03B		
Spain	ESOEXA03B		
Australia	AUOEXA03B		
Canada	CNOEXA03B		
Denmark	DKOEXA03B		
Japan	JPOEXA03B		
Mexico	MXOEXA03B		
New Zealand	NZOEXA03B		
Norway	NWOEXA03B		
Sweden	SDOEXA03B		
Switzerland	SWOEXA03B		
United Kingdom	UKOEXA03B		
United States	USOEXA03B		

Table 2.6: Sources for monthly interest rate series

Measure: Interest rates		**Frequency: monthly**	
Country	**DS Mnemonic**	**Source**	**Comment**
Austria	OEprate.	European Central Bank	Policy Rate
Belgium	BGprate.	European Central Bank	Policy Rate
Finland	FNOIR030R	OECD Main Economic Indicators	Money Market Rate
France	FRINTER3	OECD Main Economic Indicators	Money Market Rate
Germany	BDINTER3	OECD Main Economic Indicators	Money Market Rate
Greece	GRprate.	European Central Bank	Policy Rate
Ireland	IRprate.	European Central Bank	Policy Rate
Italy	ITINTER3	OECD Main Economic Indicators	Money Market Rate
Luxembourg	LXI60L..	International Financial Statistics	Start: 1985
Netherlands	NLINTER3	OECD Main Economic Indicators	Money Market Rate
Portugal	PTprate.	European Central Bank	Policy Rate
Spain	ESINTER3	OECD Main Economic Indicators	Money Market Rate
Australia	AUI60...	International Financial Statistics	Money Market Rate
Canada	CNBCBPR	Datastream	Policy Rate
Denmark	DKBCBPR	Datastream	Policy Rate
Japan	JPprate.	Bank of Japan	Policy Rate
Mexico	MXMIR060R	OECD Main Economic Indicators	Money Market Rate
New Zealand	NZMIR076R	OECD Main Economic Indicators	Money Market Rate
Norway	NWI60…; nwprate.	International Financial Statistics; Norges Bank	Money Market rate until 2017; From 2017 Policy Rate
Sweden	SDprate.	Sveriges Riksbank	Policy Rate
Switzerland	SWINTER3	OECD Main Economic Indicators	Money Market Rate
United Kingdom	UKprate.	Bank of England	Policy Rate
United States	USINTER3	Refinitiv	Money Market Rate
Euro Area Countries	EMINTER3	European Central Bank	All from 1999: Money Market Rate

Table 2.7: Sources for monthly consumer price index series (used for frequency conversion by interpolation)

Measure: CPI **Country**	**DS Mnemonic**	**Frequency: Monthly** **Source**	**Comment**
Austria	OECONPRCF	National Statistical Office	
Belgium	BGCONPRCF	National Statistical Office	
Finland	FNCONPRCF	National Statistical Office	
France	FROCP009F	OECD Main Economic Indicators	
Germany	BDCONPRCF	National Statistical Office	
Greece	GRCONPRCF	National Statistical Office	
Ireland	IRCONPRCF	National Statistical Office	
Italy	ITCONPRCF	National Statistical Office	
Luxembourg	LXOCP009F	OECD Main Economic Indicators	
Netherlands	NLCONPRCF	National Statistical Office	
Portugal	PTCONPRCF	National Statistical Office	
Spain	ESCONPRCF	National Statistical Office	
Australia	AUCCPI..E	National Statistical Office/Refinitiv	
Canada	CNCONPRCF	National Statistical Office	
Denmark	DKCONPRCF	National Statistical Office	
Japan	JPCONPRCF	National Statistical Office	
Mexico	MXCONPRCF	National Statistical Office	
New Zealand	NZCCPI..E	National Statistical Office/Refinitiv	
Norway	NWCONPRCF	National Statistical Office	
Sweden	SDCONPRCF	National Statistical Office	
Switzerland	SWCONPRCF	National Statistical Office	
United Kingdom	UKOCP009F	OECD Main Economic Indicators	
United States	USCONPRCF	Bureau of Labor Statistics	

Table 2.8: Sources for monthly industrial production series (used for frequency conversion by interpolation)

Measure: Industrial Production		**Frequency: Monthly**	
Country	**DS Mnemonic**	**Source**	**Comment**
Austria	OEOPRI35G	OECD Main Economic Indicators	
Belgium	BGOPRI35G	OECD Main Economic Indicators	
Finland	FNOPRI35G	OECD Main Economic Indicators	
France	FROPRI35G	OECD Main Economic Indicators	
Germany	BDOPRI35G	OECD Main Economic Indicators	
Greece	GROPRI35G	OECD Main Economic Indicators	
Ireland	IROPRI35G	OECD Main Economic Indicators	
Italy	ITOPRI35G	OECD Main Economic Indicators	
Luxembourg	LXOPRI35G	OECD Main Economic Indicators	
Netherlands	NLOPRI35G	OECD Main Economic Indicators	
Portugal	PTOPRI35G	OECD Main Economic Indicators	
Spain	ESOPRI35G	OECD Main Economic Indicators	
Australia	AUCIND..G	National Statistical Office/Refinitiv	
Canada	CNOPRI35G	OECD Main Economic Indicators	
Denmark	DKOPRI35G	OECD Main Economic Indicators	
Japan	JPOPRI35G	OECD Main Economic Indicators	
Mexico	MXOPRI35G	OECD Main Economic Indicators	
New Zealand	NZCUNP..O	National Statistical Office/Refinitiv	Unemployment Rate
Norway	NWOPRI35G	OECD Main Economic Indicators	
Sweden	SDOPRI35G	OECD Main Economic Indicators	
Switzerland	SWCIND..G;	National Statistical Office/Refinitiv;	Constructed from both series
	SWI66..XR	International Financial Statistics	
United Kingdom	UKOPRI35G	OECD Main Economic Indicators	
United States	USOPRI35G	OECD Main Economic Indicators	

Notes: In Mexico data collection for industrial production only starts in January 1980, thus causing a delay of the sample start due to the trend extraction exercise, In New Zealand data collection for the unemployment rate only starts in March 1986, thus causing a further delay of the sample start

Data for section 2.2.2

Central Bank Independence index taken from Garriga (2016) and can be downloaded from https://sites.google.com/site/carogarriga/cbi-data-1
Average interest rate is the unweighted average of the monthly interest rate series for the respective time periods.
Average inflation rate is the unweighted average of the growth rate in the respective time period based on the quarterly GDP deflator series derived from the ratio of nominal to real GDP

Table 2.9: Sources for SCM weight calculation covariates II

	All Frequencies: Annually (1999)							
	Measure: Total GDP			Measure: Population Size			Measure: GDP per Capita	
Country	**DS Mnemonic**	**Source**	**Comment**	**DS Mnemonic**	**Source**	**Comment**	**DS Mnemonic**	**Source**
Austria	OEAUVGDP	DG ECFIN AMECO		OEOCFTPP	OECD Economic Outlook		OEWDUGY7C	World Bank WDI
Belgium	BGAUVGDP	DG ECFIN AMECO		BGOCFTPP	OECD Economic Outlook		BGWDUGY7C	World Bank WDI
Finland	FNAUVGDP	DG ECFIN AMECO		FNOCFTPP	OECD Economic Outlook		FNWDUGY7C	World Bank WDI
France	FRAUVGDP	DG ECFIN AMECO		FROCFTPP	OECD Economic Outlook		FRWDUGY7C	World Bank WDI
Germany	BDAUVGDP	DG ECFIN AMECO		BDOCFTPP	OECD Economic Outlook		BDWDUGY7C	World Bank WDI
Greece	GRAUVGDP	DG ECFIN AMECO		GROCFTPP	OECD Economic Outlook		GRWDUGY7C	World Bank WDI
Ireland	IRAUVGDP	DG ECFIN AMECO		IROCFTPP	OECD Economic Outlook		IRWDUGY7C	World Bank WDI
Italy	ITAUVGDP	DG ECFIN AMECO		ITOCFTPP	OECD Economic Outlook		ITWDUGY7C	World Bank WDI
Luxembourg	LXWDLGSKA	World Bank WDI	/1000000000	LXPOPTOT.	Statistics Luxemburg	/1000	LXWDUGY7C	World Bank WDI
Netherlands	NLAUVGDP	DG ECFIN AMECO		NLOCFTPP	OECD Economic Outlook		NLWDUGY7C	World Bank WDI
Portugal	PTAUVGDP	DG ECFIN AMECO		PTOCFTPP	OECD Economic Outlook		PTWDUGY7C	World Bank WDI
Spain	ESAUVGDP	DG ECFIN AMECO		ESOCFTPP	OECD Economic Outlook		ESWDUGY7C	World Bank WDI
Australia	AUAUVGDP	DG ECFIN AMECO		AUOCFTPP	OECD Economic Outlook		AUWDUGY7C	World Bank WDI
Canada	CNAUVGDP	DG ECFIN AMECO		CNOCFTPP	OECD Economic Outlook		CNWDUGY7C	World Bank WDI
Denmark	DKAUVGDP	DG ECFIN AMECO		DKOCFTPP	OECD Economic Outlook		DKWDUGY7C	World Bank WDI
Japan	JPAUVGDP	DG ECFIN AMECO		JPOCFTPP	OECD Economic Outlook		JPWDUGY7C	World Bank WDI
Mexico	MXAUVGDP	DG ECFIN AMECO		MXOCFTPP	OECD Economic Outlook		MXWDUGY7C	World Bank WDI
New Zealand	NZAUVGDP	DG ECFIN AMECO		NZOCFTPP	OECD Economic Outlook		NZWDUGY7C	World Bank WDI
Norway	NWAUVGDP	DG ECFIN AMECO		NWOCFTPP	OECD Economic Outlook		NWWDUGY7C	World Bank WDI
Sweden	SDAUVGDP	DG ECFIN AMECO		SDOCFTPP	OECD Economic Outlook		SDWDUGY7C	World Bank WDI
Switzerland	SWAUVGDP	DG ECFIN AMECO		SWWD8FD7P	World Bank WDI	/1000	SWWDUGY7C	World Bank WDI
United Kingdom	UKAUVGDP	DG ECFIN AMECO		UKOCFTPP	OECD Economic Outlook		UKWDUGY7C	World Bank WDI
United States	USAUVGDP	DG ECFIN AMECO		USOCFTPP	OECD Economic Outlook		USWDUGY7C	World Bank WDI

Data for section 2.2.4

Table 2.10: Sources for monthly exchange rate series

Measure: DM FX rates		Frequency: Monthly	
Country	**DS Mnemonic**	**Source**	**Comment**
Austria	BDWU5015A	Deutsche Bundesbank	
Belgium	BDWU5001A		
Finland	BDWU5002A		
France	BDWU5012A		
Ireland	BDWU5017A		
Italy	BDWU5007A		
Netherlands	BDWU5000A		
Portugal	BDWU5004A		
Spain	BDWU5006A		
Measure: US-Dollar FX rates		**Frequency: Monthly**	
Austria	OEXRUSD.	Bank of England	
Belgium	BGXRUSD.		
Finland	FNXRUSD.		
France	FRXRUSD.		
Germany	BDXRUSD.		
Ireland	IRXRUSD.		
Italy	ITXRUSD.		
Netherlands	NLXRUSD.		
Portugal	PTXRUSD.		
Spain	ESXRUSD.		
Australia	AUXRUSD.		
Canada	CNXRUSD.		
Denmark	DKXRUSD.		
Japan	JPXRUSD.		
New Zealand	NZXRUSD.		
Norway	NWOCC016		1/NWOCC016
Sweden	SDXRUSD.		
Switzerland	SWXRUSD.		
United Kingdom	UKXRUSD.		
United States	1/BDXRUSD.		1/BDXRUSD.

2.A.2 SCM statistics

Table 2.11: SCM weights

	SR	**IH**	**Cholesky**
AU	0%	0%	0%
CA	0%	0%	21%
DK	0%	0%	0%
JP	2%	0%	9%
NO	6%	1%	4%
CH	5%	12%	3%
UK	86%	86%	63%
US	1%	1%	0%

Notes: SCM weighting vectors for the baseline specification, different identification assumptions: Sign restrictions (SR), Identification using heteroskedasticity (IH) and zero restrictions (Cholesky) following the recursive ordering described in equation 15

Table 2.12: Monetary policy stress in the euro area (EA) and its doppelganger

	Stress EA	Stress EA doppelganger
SR	28.55	29.25
IH	0.007	0.001
Cholesky	135.35	138.01

Notes: Average monetary policy stress $\hat{\sigma}^2_{1,MP}$ in the euro area (EA) and its doppelganger replication following from as the last column $X_0 w$ in equation 16. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

Table 2.13: Key characteristics of the euro area average and its doppelganger

	EA average	D SR	D IH	D CHOL
CB Independence	0.48	0.32	0.30	0.30
GDP	1203.06	1161.08	1202.73	1203.51
i	9.24	9.49	9.35	9.08
GDP Growth	2.62	3.02	2.97	3.18
Inflation	8.58	6.85	7.67	6.77
GDP per capita	34501.86	34187.25	34501.46	34501.22

Notes: The average euro area (EA) country and its doppelganger (D) replications following from X_0w as in equation 16. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH) and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

2.A.3 Supplementary results

Table 2.14: Factor of improvement of monetary stress for the average euro area country and its doppelganger

	EA average	EA doppelganger
IH	25.7	18.1
SR	26.2	16.5
Cholesky	26.4	18.5

Notes: The table displays the post-euro to pre-euro ratio of the monetary stress measure $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for the euro area average and its doppelganger after applying the SCM country weights w_i to the individual country factor of improvement as in the tables 2.3 and 2.2 . The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

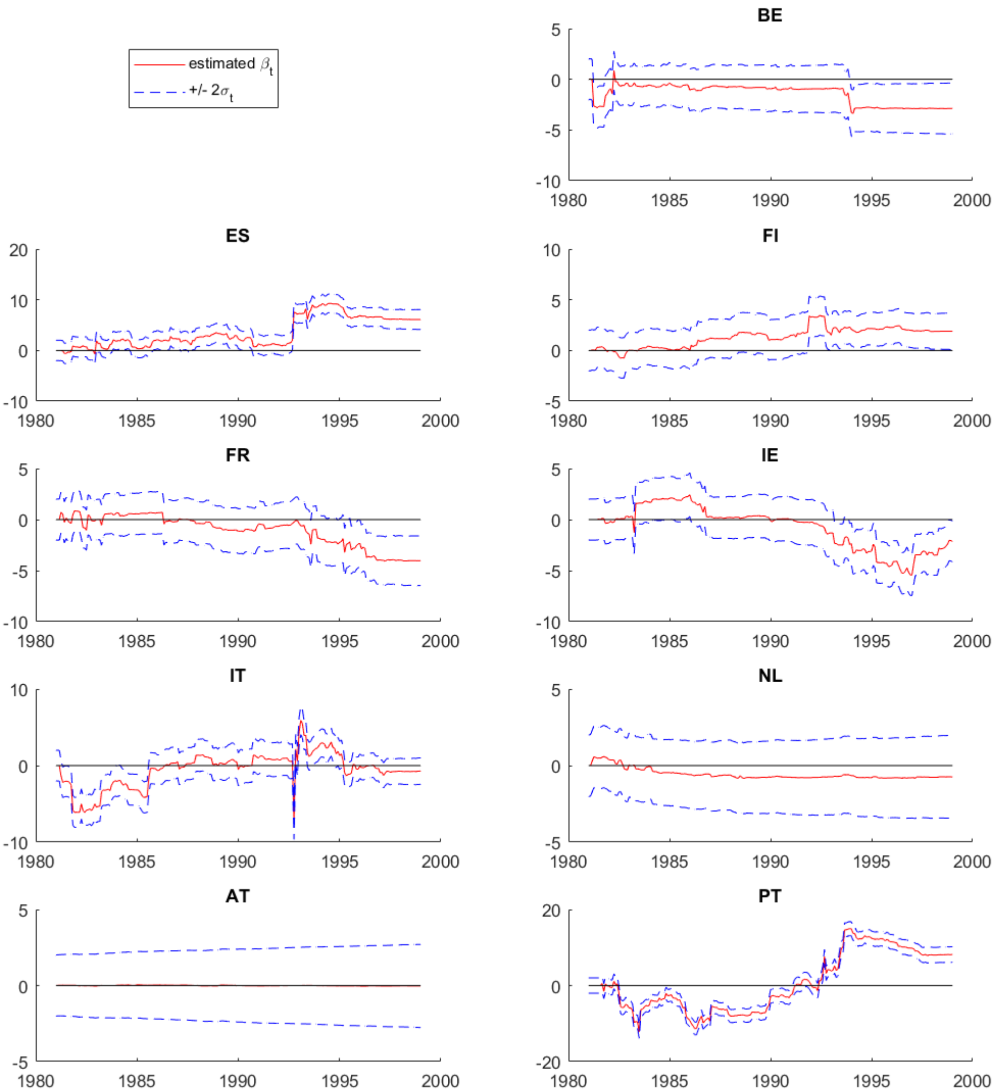
Table 2.15: Factor of improvement of monetary stress for individual euro area countries and their doppelgangers

	A SR	D SR	A HET	D IH	A Chol	D Chol
DE	*10*	*24*	*14*	*19*	*16*	*22*
BE	**21**	**20**	**21**	**17**	**24**	**17**
ES	**26**	**17**	**21**	**20**	**25**	**18**
FI	*3*	*22*	*6*	*19*	*5*	*17*
FR	**24**	**18**	**21**	**19**	**24**	**20**
IE	**21**	**21**	**22**	**16**	**19**	**16**
IT	**45**	**19**	**35**	**18**	**43**	**20**
NL	**29**	**13**	**14**	**12**	**17**	**11**
AT	*4*	*13*	*3*	*14*	*4*	*13*
PT	**127**	**17**	**238**	**20**	**114**	**18**

Notes: The table displays the post-euro to pre-euro ratio of the monetary stress measure $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for every country (A) compared to an estimate for a doppelganger (D) for every individual country. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Chol), following the recursive ordering described in equation 15.

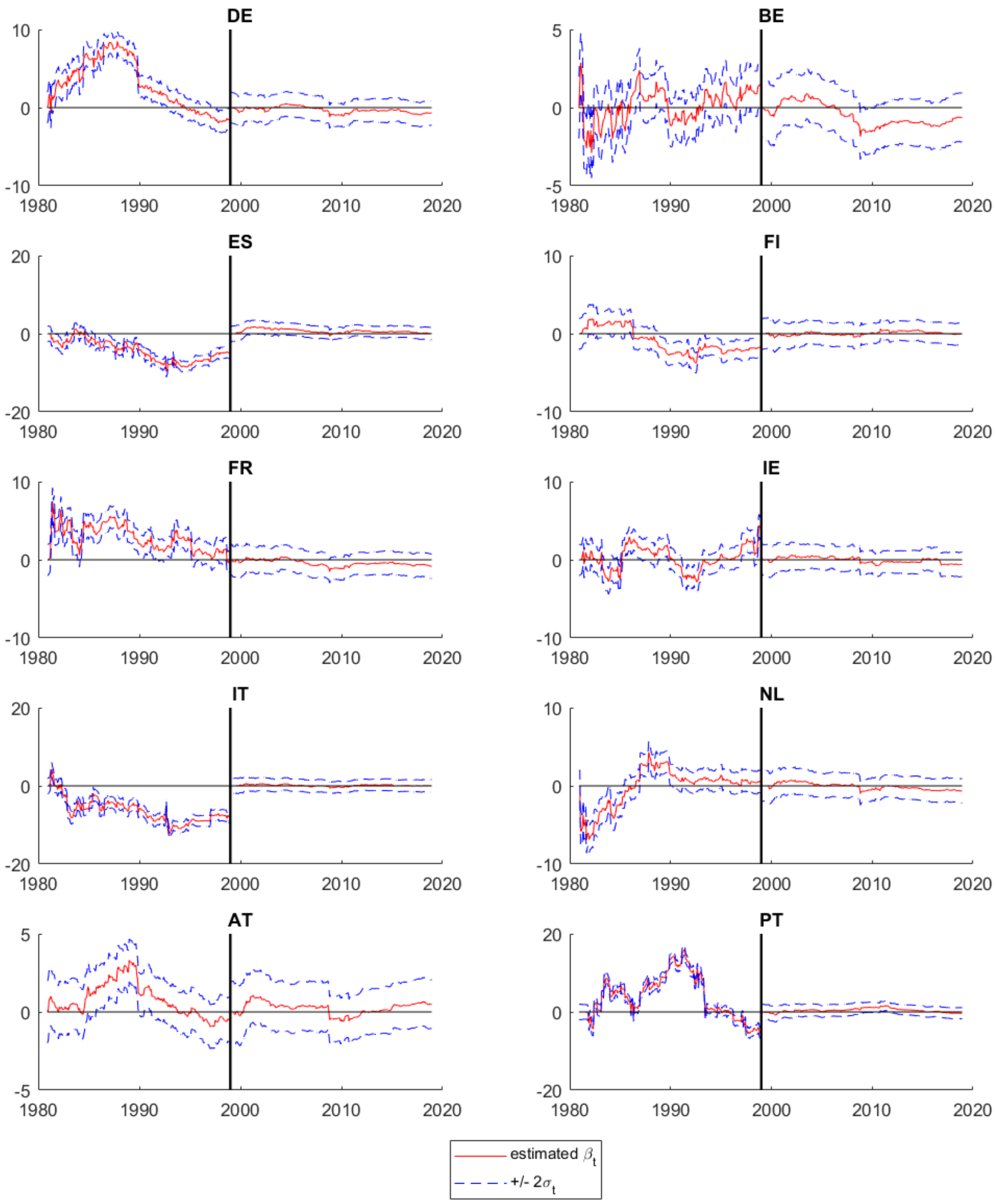
2.A.4 Graphs: Exchange rate fluctuations and monetary stress

Figure 2.5: Monetary stress and the D-Mark



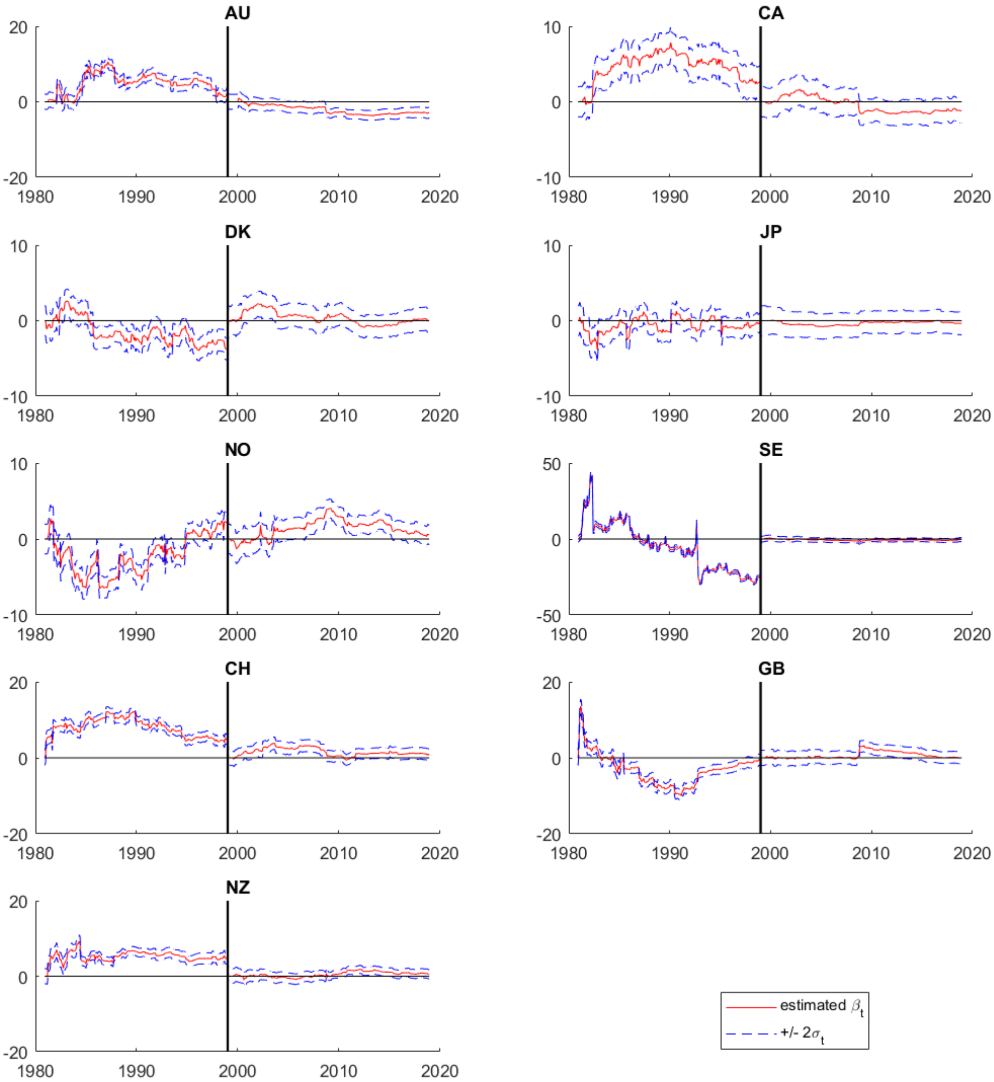
Notes: Time-varying impact of ΔE_t (change of D-Mark/national currency) on $\omega_t^{MON-POL}$, Belgium (1), Spain (2), Finland (3), France (4), Ireland (5), Italy (6), Netherlands (7), Austria (8), Portugal (9)

Figure 2.6: Monetary stress and the U.S. dollar, euro area countries



Notes: Time-varying impact of ΔE_t (change of Dollar/national currency) on $\omega_t^{MON-POL}$, Germany (1), Belgium (2), Spain (3), Finland (4), France (5), Ireland (6), Italy (7), Netherlands (8), Austria (9), Portugal (10)

Figure 2.7: Monetary stress and the U.S. dollar, non-euro area countries



Notes: Time-varying impact of ΔE_t (change of Dollar/national currency) on $\omega_t^{MON-POL}$, Australia (1), Canada (2), Denmark (3), Japan (4), Norway (5), Sweden (6), Switzerland (7), United Kingdom (8), New Zealand (9)

2.A.5 Robustness exercises

Shortening of the time sample

Table 2.16: Key results for the euro area and its doppelganger using a shorter time sample

Time Sample: 1993 - 2006

	EA average factor	EA doppelganger factor		
IH	2.52	1.62		
SR	15.17	8.69		
Cholesky	8.15	4.36		
	EA Average	D SR	D IH	D Cholesky
CB Independence	0.48	0.33	0.32	0.35
GDP	1203.06	1203.11	1203.02	1200.11
i	7.75	7.92	7.87	5.45
GDP Growth	2.33	2.27	2.28	1.56
Inflation	3.44	3.34	3.37	3.36
GDP per capita	34501.86	34502.02	34502.17	34506.68
	Stress EA	Stress EA doppelganger		
SR	10.23	10.25		
IH	0.13	0.05		
Cholesky	16.45	13.58		

Notes: Results for the euro area (EA) average and its doppelganger (D) after applying the SCM country weights w_i to the individual country factor of improvement in monetary policy stress $\frac{\hat{\sigma}^2_{1,MP,i}}{\hat{\sigma}^2_{2,MP,i}}$ for the time sample 1993 - 2006. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15. Below, the attributes of the doppelganger for the different identification assumptions.

Table 2.17: Factor of improvement of the individual euro area countries using a shorter time sample

	DE	**BE**	**ES**	**FI**	**FR**	**IE**	**IT**	**NL**	**AT**	**PT**
SR	0.68	0.74	0.55	2.48	5.36	1.02	41.50	2.83	0.21	102.41
IH	3.07	3.05	2.38	0.74	3.53	2.56	0.38	4.95	2.15	0.26
Cholesky	1.48	0.51	0.77	2.30	5.16	1.66	26.20	2.23	0.15	43.63

Notes: Results for the time sample 1993 - 2006. $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for a sample of 10 euro area countries. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

Table 2.18: Factor of improvement of the individual non-euro area countries using a shorter time sample

	AU	**CA**	**DK**	**JP**	**NO**	**CH**	**UK**	**US**
SR	7.50	8.86	16.37	0.85	3.70	0.63	0.24	0.34
IH	2.63	1.41	0.37	2.08	0.35	1.56	1.79	8.44
Cholesky	11.20	9.16	8.87	2.17	5.38	0.29	0.36	1.11

Notes: Results for the time sample 1993 - 2006. $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for a sample of 8 non-euro area countries. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

Increased country sample for 1993-2006

Table 2.19: Key results of the euro area and its doppelganger using a shorter time sample and more countries

Expansion of the Country Sample for 1993-2006

	EA	EA Doppelganger		
Het	2.45	1.54		
SR	12.15	1.53		
Cholesky	8.33	1.17		
	EA Average	D SR	D IH	D Cholesky
CB Independence	0.48	0.34	0.33	0.33
GDP	1171.73	1171.63	1171.72	1171.60
i	6.21	5.79	9.51	5.53
GDP Growth	2.35	2.94	2.95	2.24
Inflation	2.34	2.52	10.35	2.66
GDP per capita	34362.96	34363.18	34362.96	34363.22
	Stress EA	Stress Doppelganger		
SR	4.91	4.84		
IH	0.13	0.28		
Cholesky	16.92	17.80		

Notes: Results for the euro area (EA) average and its doppelganger (D) after applying the SCM country weights w_i to the individual country factor of improvement in monetary policy stress $\frac{\hat{\sigma}^2_{1,MP,i}}{\hat{\sigma}^2_{2,MP,i}}$ for the time sample 1993 - 2006 with five additional countries. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15. Below, the attributes of the doppelganger for the different identification assumptions.

Table 2.20: Factor of improvement of the additional countries

	GR	**LU**	**SE**	**MX**	**NZ**
SR	0.95	0.16	7.50	5.74	0.21
IH	0.31	0.90	0.05	1.98	0.52
Cholesky	15.91	0.16	6.79	8.92	0.24

Notes: Results for the time sample 1993 - 2006. $\frac{\hat{\sigma}^2_{1,MP}}{\hat{\sigma}^2_{2,MP}}$ for two additional euro area and 3 additional non-euro area countries. The identification assumptions are Sign restrictions (SR), Identification using heteroskedasticity (IH), and zero restrictions (Cholesky), following the recursive ordering described in equation 15.

Exclusion of Portugal

Table 2.21: Factor of improvement of the euro area excluding Portugal

Exclusion of Portugal

	EA
IH	20.23
SR	23.56
Cholesky	24.11

Notes: Results for the euro area (EA) average (excluding Portugal) factor of improvement in monetary policy stress $\frac{\hat{\sigma}^2_{1,MP,i}}{\hat{\sigma}^2_{2,MP,i}}$ for the baseline country/time sample.

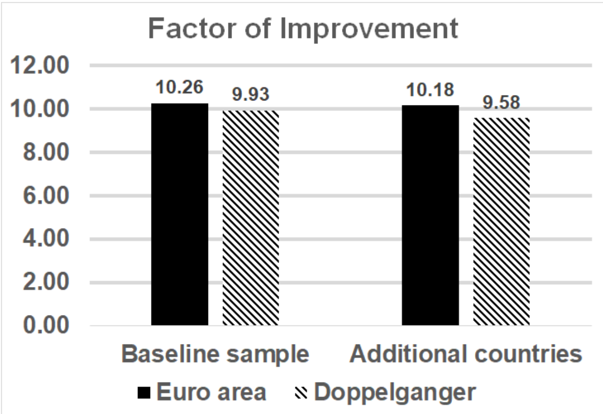
2.A.6 Results based on loss functions

Table 2.22: Loss function parameters

Parameter	Value	Interpretation
β	0.99	Household's discount factor
σ	1	Intertemp. subst. elasticity of consumption
ϕ	1	Labor supply elasticity
α	$\frac{1}{3}$	Capital share of output
ε	6	Substitution elasticity of consumption
θ	$\frac{2}{3}$	Calvo probability
λ	0.0425	Impact of marg. costs on inflation

Notes: Parametrization for the loss function L from Galí (2015)

Figure 2.8: Factor of improvement based on a loss function



Notes: Results for the euro area (EA) average factor of improvement in loss $\frac{L^{pre}}{L^{post}}$ for the baseline sample (10 EA countries vs. 8 non-EA countries) and the augmented sample (12 EA countries vs. 11 non-EA countries). The doppelganger is constructed, matching 6 the six baseline covariates and L^{pre}

2.A.7 EMS, monetary policy, and crises

This Appendix was published as in Box 2 of Fritsche and Harms (2019)

The EMS, which existed from 1979 until the introduction of the euro, consisted of two elements: the European Exchange Rate Mechanism (ERM) and the European Currency Unit (ECU), which served as an accounting unit.[39] The economies participating in the ERM set central rates in relation to the ECU currency basket and limit exchange rate fluctuations to $\pm$ 2.5 percent[40] around this rate.[41] The international foreign exchange markets determined the fluctuations between these upper and lower limits while central rate adjustments were the result of political negotiations and required the approval of all participants. The participating central banks were obliged to defend the upper and lower limits by buying and selling their own currencies as well as foreign currencies. They also could act providently within the fluctuation margins.[42]

In order to counter structural economic divergences, such as wage, inflation, and foreign trade developments, many adjustments to central rates took place, particularly in the early 1980s. Generally, some countries tended to devalue their currencies more often (France and Italy) and others (Germany and the Netherlands) only appreciated them (Höpner and Spielau, 2018). Therefore, Germany and, in particular, its Bundesbank played a dominant role in the EMS.

The role of monetary policy as it is understood today is not easy to identify in this system. The EMS was, on the one hand, a fixed exchange rate system, but on the other, it offered the possibility of discretionary adjustments. If central banks have to operate to a large extent on the foreign exchange markets by buying or selling their own currency, it affects the supply of liquidity to the financial system and, thus, the interest rate. If, for example, the Bundesbank was exposed to an extremely high demand for the Deutsche Mark and, thus, to high revaluation pressure, it would have to increase the supply of the Deutsche Mark just as drastically in order to counteract that pressure. In most cases, such stabilization is not possible without affecting the interest rate. Conversely, a change in the interest rate motivated by monetary policy (such as a rise in interest rates to combat inflation) can trigger devaluation or revaluation pressure in another country. If the other country does not want to adjust the exchange rate but has already exhausted the means to intervene in the foreign exchange market, the only remaining option is an interest rate increase. Both cases are examples of interest rate changes that clearly do not contribute to national macroeconomic stabilization.

Such economically unjustifiable interest rate decisions regularly occurred in the EMS. As early as the beginning of the 1980s, many other central banks copied a surprising three percentage point interest rate hike by the Bundesbank in order to prevent a devaluation.[43] This problem was exacerbated by the gradual abolition of capital controls from 1987 onwards under the Single European Act.

Many economists believe that the largest crisis of the EMS is a direct consequence of the fall of the Berlin Wall and the Bundesbank's reaction. Reunification and the resulting costs acted as a major economic stimulus package in Germany, while large parts of the EU struggled with recession or weak growth. When the inflation rate exceeded the five percent mark in 1992, the Bundesbank decided to raise interest rates several times. After the abolition of capital controls, the pressure exerted by the financial markets increased significantly. There was great uncertainty regarding how long the central banks of the other countries would be able to keep up with the Bundesbank and maintain their commitment to the Deutsche Mark, despite widely diverging economic trends.

[39] The ERM is the central element of the EMS, which is why it is the focus of this box. The ERM still exists today as ERM II and serves as an official system for countries of the European Union. Countries interested in adopting the euro must participate in ERM II for two years. Since most Eastern European countries interested in the euro have already introduced it, Denmark is currently the only participating country.

[40] From the outset, Belgium, Denmark, France, Germany, Ireland, and the Netherlands participated in the system and used these 2.5 percent as a fluctuation margin. Italy was granted a larger margin of $\pm$ six percent until 1990, as were Spain, which joined the EMS in 1989, the United Kingdom (1990), and Portugal (1992).

[41] The EMS was already largely de facto abolished over the course of the EMS crisis in 1992/1993 when the fluctuation margins were increased to $\pm$ 15 percent.

[42] Through the "Very Short Term Financing Facility," each currency was available to the countries at short notice in a theoretically unlimited volume on the condition that the foreign currency loans were repaid after 45, and later 75, days.

[43] Between March 1979 and February 1980, the Bundesbank increased the discount rate from four to seven percent.

In Scandinavia, which was first attacked by currency speculation in early September 1992, the Swedish Riksbank attempted to stabilize its exchange rate by temporarily raising interest rates as high as 500 percent. Later, speculation also hit the EMS. The Bank of England drastically raised the key interest rate on September 16, 1992, despite the United Kingdom's weak economy, as did the Bank of Italy. Ultimately, monetary policy was unable to counter speculative pressure and both countries left the EMS.

Chapter Three

Zooming in on Monetary Policy - The Labor Share and Production Dynamics of Two Million Firms

This chapter is based on joint work with Lea Steininger A previous version of this project is forthcoming in the DIW Discussion Paper Series and is submitted to the SSRN working paper series, parts of it will be published in a forthcoming DIW Weekly Report

3 Zooming in on Monetary Policy - The Labor Share and Production Dynamics of Two Million Firms

Conditional on a contractionary monetary policy shock, the labor share of value added is expected to decrease in the basic New Keynesian model. By providing firm-level evidence, we are first to validate this proposition. Using local projections and high dimensional fixed effects, we show that a one standard deviation contractionary monetary policy shock decreases firms' labor share by 0.4 percent, on average. However, reactions are heterogeneous along two dimensions: The labor share is most informative to discriminate firms by their response in payroll expenses, firms' leverage is most informative to discriminate by their response in value added. We inform the policy debate on transmission and redistribution effects of monetary policy.

Keywords:Monetary policy, firm heterogeneity, labor share, financial frictions, DSGE model validation.
JEL classifications: D22, D31, E23, E32, C52

3.1 Introduction

Our study provides firm-level evidence on the response of key economic variables to a monetary policy (MP) shock: value added, wages, and the labor share[1]. Both value added and wages constitute a cornerstone of corporate decision-making. Value added measures a company's contribution to economic output. Wages, on the other hand, are the main component in income for the majority of households, they are central to both investment and consumption behavior. From the perspective of the firm, labor compensation is typically a vital cost component. At the macro-level, understanding wage-setting is key for understanding inflation dynamics and unemployment (Christiano et al., 2016; Galí et al., 2012).[2] The labor share of value added is not only a key indicator for the distribution of income (Piketty, 2015) but directly linked to the mark-up and, by implication, central to the pricing behavior of firms (Cantore et al., 2020; Galí, 2015; Nekarda and Ramey, 2020). Notwithstanding these factor's centrality, and despite the fact that the New Keynesian (NK) model features clear implications, empirical literature of their response to MP shocks at firm-level is virtually nonexistent.

Our paper is thus first to test for the consistency of the NK model by investigating the change of the labor share at the highest resolution across euro area (EA) member states. Moreover, we dig deeper and propose the hypothesis that firm heterogeneity in cost structure is a decisive factor to consider for the labor share response to MP. We inform the current debate on the evolution of the redistribution channel of MP and firm heterogeneity (Auclert, 2019; Cantore et al., 2020; Kehrig and Vincent, 2021; Ottonello and Winberry, 2020). In order to study the nuances in the labor share, wages and value added as well as empirically explore our hypotheses, we analyze a micro-panel from 1999-2017 covering over 2.1 million firms.

We find a significant, highly robust and pronounced negative reaction of labor costs, value added and the labor share after a contractionary MP shock. The literature on transitory effects of firm heterogeneity focuses on fundamentals such as age (Cloyne et al., 2018), size (Gertler and Gilchrist, 1994) or balance sheet related measures (Jeenas, 2019; Ottonello and Winberry, 2020). We find that measures related to factor input costs matter predominantly for determining heterogeneity of cyclical behavior. In particular, while firms with a high labor share are affected most strongly by MP, firms with a low labor share operate more independently. Moreover, firms with high leverage ratios are also relatively more responsive to MP. However, across these two dimensions – labor share and leverage – results are driven by fundamentally different channels. These are, firms are most meaningfully discriminated according to what constitutes the main component in their factor input cost structure: labor-intensive firms react by making payroll amendments, highly leveraged firms react by altering their production. For the latter, value added drives the results.

We use information-neutral shocks developed by Jarociński and Karadi (2020) for identified changes in MP. These shocks may be interpreted as exogenous with respect to credit conditions insofar as they do not contain any central bank information on the state of the economy. Following Ottonello and Winberry (2020), we break down the euro-area wide MP shocks to the micro-level by employing the firm's idiosyncratic leverage ratio as an exposure measure. In a local projections regression framework, we then estimate the labor share's reaction, controlling for a high dimensional combination of fixed effects.

Results from this research project help design more effective governmental and central bank policies. The paper informs the discussion about fundamental questions in the field of monetary economics, including: how do low interest rates affect the distribution of income between wages and capital from the perspective

[1] labor compensation divided by value added

[2] Together with price-setting, wage-setting determines the response of the economy to MP shocks (Galí, 2011). Galí (2011) notes that, *'despite the central role of the wage-setting block in the NK model, the amount of work aimed at assessing its empirical relevance has been surprisingly scant'*.

of the firm? Are widely used NK models able to jointly match the response of wages and the labor share? How do firm-specific characteristics such as financial constraints, firm age or balance sheet composition affect the transmission of MP, and what channels are at play?
The remainder of the paper is structured as follows. The subsequent section explains the theory of our working hypothesis. Section 3.3 explains our data set. Section 3.4 contains the empirical results and robustness exercises. Section 3.5 discusses and interprets the findings, puts them into context with the existing literature and derives some policy implications. Section 3.6 concludes.

3.2 The labor share in the NK-Model, firm heterogeneity and working hypotheses

This section outlines the theoretical background of our empirical working hypothesis. We combine insights from the literature on firm-level frictions such as capital irreversibility (Lanteri, 2018; Ramey and Shapiro, 2001) and cash-in-advance (Carlstrom and Fuerst, 1995; Lucas and Stokey, 1985) with insights from the literature on macro-fluctuations – i.e. heterogeneous wage and price Philips curves (Cantore et al., 2020; Galí, 2015) - to motivate our hypothesis on the response of the firm level labor share to a MP shock.

In standard NK models, the labour share is related to the real marginal cost (i.e. the inverse of the price markup[3]). According to stylized textbook set-ups (Cantore et al., 2020; Galí, 2015), the labor shares' deviation from the steady state is defined as

$$ls_t = w_t + h_t - y_t \tag{21}$$

where ls_t is the labor share w_t is the real wage, h_t is hours worked, and y_t is the output; all at time t. Using a simple wage and price Philips curve as well as the assumption that MP is able to track the natural rate of interest, Cantore et al. (2020) show that log deviation of real wages can be expressed as

$$w_t = \frac{\lambda^p - \lambda^w}{1 + \kappa_p + \kappa_w} \varepsilon_t^{MP} \tag{22}$$

where λ^p, λ^w, κ_p, and κ_w are (positive) slope coefficients of the Phillips curves. ε_t^{MP} is a MP shock. In response to monetary tightening, real wages are expected to fall if wages are more flexible than prices - i.e. $\lambda^p < \lambda^w$. With respect to the production function, frictions in the labor market and financial constraints can influence the labor share in equation 21 - suggesting that the labor share decreases after monetary tightening under reasonable specifications of the NK model (Cantore et al., 2020).

The NK model, however, is a model of the aggregate economy and thus agnostic to the business models of different kinds of companies. Equation 22 illustrates that wage policies of specific firms influence the labor share at the firm-level.

MP shocks impact the costs of a company by altering borrowing costs.[4] To investigate the response of a firm to a MP shock, we hence suggest it is most informative to look at variables closely related to the company's cost structure. Typically, in the academic macroeconomic realm, the production process is best characterized via a production function that describes how output is produced as a combination of its factor inputs, capital and labor. Taken together, we consider it reasonable to employ variables that contain information about the composition of and expenses on these factor inputs, in order to discriminate meaningfully between firms, and to map out their response to MP. These variables are the (firm-average)

[3]the labour share equals the real marginal costs (i.e. $ls_t = mc_t$) if labor is compensated according its marginal product.
[4]Some firms may be affected more indirectly, e.g. if they have no or very low borrowing costs.

leverage ratio, and the labor share.
Leverage, on the one hand, is a commonly used measure to approximate the sensitivity of firms to monetary policy (Jeenas, 2019; Ottonello and Winberry, 2020). The firm-average labor share, on the other hand, is closely related to the production function of companies, serves as an indicator to distinguish companies between industries and business models, and is highly informative about labor costs a company is faced with. Surprisingly, the degree to which a firm employs labor during production is seldom discussed in macroeconomic literature. Due to the fact that labor has its own regulation, accounting standards and physical limits, the degree to which a company employs it serves as an important dimension to discriminate between firms, both within[5] and between[6] industries.
The literature on firm heterogeneity discusses two other variables used to discriminate firms: size (Gertler and Gilchrist, 1994), and age (Cloyne et al., 2018). Clearly, however, to discriminate between firms (or industries) along the lines of (a combination of) age or size does not result in the intuitive description of firms (or industries) that the labor share (often combined with leverage) provides.[7] Therefore, we argue, they are inferior to the labor share when it comes to complexity reduction and theoretical abstraction. Complexity reduction is necessary for economic modelling, and an indispensable analytical tool when disentangling the transmission channels of monetary policy.
With this in mind, it is important to note that the wage share and leverage are not two sides of the same coin. The correlation coefficient between leverage and wage share is very low[8], highlighting that both dimensions convey specific information that is not available when looking at only at them individually.
To illustrate further, consider two stylized companies with vastly dissimilar business models as an example: First, a consulting agency, characterized by a high labor share. Second, an aerospace plant with planes, airport infrastructure and other tangible assets financed by loans, characterized by high leverage.

We hypothesize that when short term credit conditions tighten, the two types of companies have fundamentally different prospects in reacting to the shock: while both companies face a cash in advance constraint[9], their cost structure and risk management are fundamentally different. As such, we expect that a company with a high labor share and its costs[10] primarily determined by the payroll alters its costs mostly via costs of employees. The consulting firm in our example thus reacts to the shock with large amendments in total costs by making redundancies, imposing hiring freezes or decreasing bonuses. The areospace plant, on the other hand, faces substantial risk of capital irreversibility (Lanteri, 2018).[11] That is to say, this type of company puts more emphasis on managing assets and investments compared to less leveraged ones, because their cost structure is less determined by payroll expenses but by external finance. The considered aerospace plant thus may alter its balance sheet primarily by reducing leverage and covering its capital costs.[12] This can be achieved in different ways e.g. by reducing investments or

[5]Consider, for instance, two companies selling the same good, e.g. translation services: these companies can go about their business either very labor-intensively (by employing a high amount of polyglot workers), or rather labor-unintensively (by using specialized computer software) – pointing toward striking *within*-industry firm heterogeneity in the labor share.

[6]Table 3.2 provides evidence that the labor share helps discriminate *between* industries, especially when combined with balance sheet indicators. The industry-level correlation coefficient between leverage and wage share is very low - .24 in our sample.

[7]In our data set, an industry-level example with high labor share and low leverage is *'scientific research and development'* and, conversely, *'manufacturing of tobacco products'* (see Table 3.2 in the Appendix 3.A.2) distinguishes itself as an industry with eminently high leverage and a low labor share. Age or size appear strikingly less handy when it comes to an intuitive distinction between these industries.

[8].14 in our data set

[9]i.e. they have to finance their respective payroll and long-lived capital assets with short-term debt (Carlstrom and Fuerst, 1995; Lucas and Stokey, 1985)

[10]including cost of short-term debt

[11]Ramey and Shapiro (2001) provides evidence from aerospace plant shut-down where capital is sold at large discounts on the secondary market.

[12]this latter transmission mechanism is known as the investment channel of MP

in case of short-term frictions for investment reduction, the aerospace plant may try to sell more planes - e.g. by offering discounts - to reduce the stock of inventory on its balance sheet and increase its revenue.

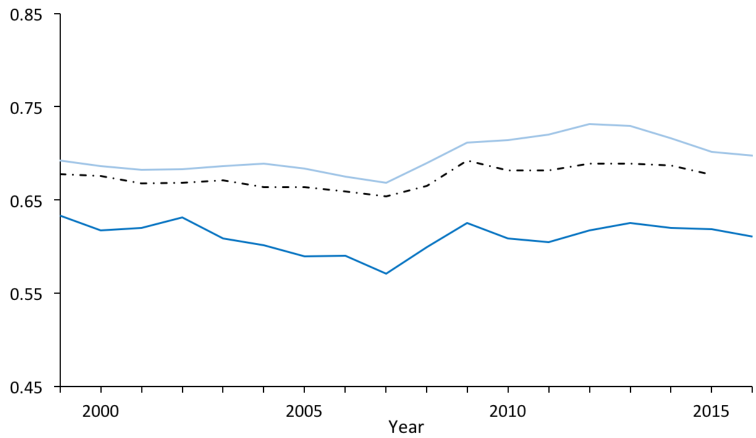
Based on the considerations laid out above, we deduct the following hypothesis:
All else equal, ...

- ... a MP tightening shock is followed by a decline in the firm-level labor share, on average.
- ... firms with large labor shares adapt to MP primarily by altering payroll expenses.
- ... firms with high leverage adapt to MP primarily by altering their value added.

3.3 Data

We base our analysis on an annual corporate-sector micro panel of EA companies from 1999 until 2017, and conduct the baseline estimations based on the eleven EA founding members to abstract from a potential sample selection bias at the country dimension. Figure 3.1 highlights that the EA aggregate labor share, in contrast to the US labor share, is relatively stable (Gutierrez and Piton, 2020) and thus not characterized by a decline described by Autor et al. (2020). Therefore, it serves as an excellent basis for the analysis of cyclical labor share components.

Figure 3.1: Euro area labor Shares - BvD Amadeus/EU KLEMS comparison



Notes: The dark dashed line in the middle displays the wage share according to *EU KLEMS* for EA-11 (defined as ratio of labor compensation to gross value added; data are from July 2018 release). The upper (light blue) line is the *BvD Amadeus* median firm-level wage share for EA-11. The bottom (dark blue) line is the *BvD Amadeus* total sum of costs of employees divided by the total sum of value added for EA-11 (all calculations based on our sample).

While the sample excludes the public sector, freelancers, and financial companies, it is highly representative at the macro-level. Figure 3.1 demonstrates the representativeness of our data, benchmarked by the EU KLEMS EA-11 labor share[13]. The solid lines (sample mean and median) capture the dynamics as well as the level of well-established statistics (dashed line) very well.

3.3.1 Firm-level data

The data for our main industry-level analysis come from the Bureau van Dijk's (BvD) Amadeus commercial database for European firms, which is a subset of the BvD Orbis dataset for global firms. This rich database comprises employment statistics, detailed balance sheet information and industrial

[13]EU KLEMS stands for EU level analysis of capital (K), labor (L), energy (E), materials (M) and service (S) inputs.

industry affiliation for SMEs and large firms, reported with annual frequency. It covers all EA countries and is thus, despite some noteworthy shortcomings[14], the best publicly available dataset for comparing firm data across Europe over time (Gopinath et al., 2017; Kalemli-Özcan et al., 2019).

Crucially, for our purpose, BvD's Amadeus provides firm reporting of total assets, equity, outstanding loans, sales, value added and cost of employees, the latter two of which are the basis to calculate firm-level labor shares over the period from the introduction of the euro in 1999 until 2017. We compute firm-specific leverage ratios as the share of total liabilities[15] over total assets. For the analysis we consider unconsolidated firm statements across the full range of corporate firms and industries. Appendix 3.A.1 describes our sample selection and data cleaning operations which we base on Kalemli-Özcan et al. (2019); Ottonello and Winberry (2020) and Belenzon et al. (2017). Following their suggestions, we drop observations with negative values of total assets, value added, number of employees or sales. In addition, we drop observations where the costs of employees or the labor share is below zero. We winsorize the labor share at the 99.5th and .05th percentile. To assure consistency across estimations, we only consider observations with non-missing labor shares. Our baseline sample includes the eleven EA founding members[16]. Table 3.1 displays the summary statistics of our data set.

Table 3.1: Summary statistics

Variable	***N***	**Mean** (Std. Dev.)	**p5**	**Median**	**p95**
MP shock$*LR_{t-1}$	13,966,175	0.02 (0.07)	-0.07	0.00	0.15
Cost of Employees/VA	14,453,176	72.63 (23.73)	28.71	75.68	102.98
Cost of Employees	14,453,176	1,310,734.65 (30,167,410.27)	18,539	182,000	3,298,416
Cost of Employees (g)	14,453,176	0.08 (0.23)	-0.22	0.04	0.53
Value added	14,453,176	2,155,078.55 (55,296,745.40)	30,361	259,857.50	5,011,055
Value added (g)	14,453,176	0.08 (0.28)	-0.30	0.04	0.63
TA	14,453,176	9,333,298.15 (361,352,962.85)	58,000	622,832	15,406,928
Equity	14,453,176	3,563,096.65 (14,3336,202.53)	10,067	172,142	5,853,024
LR	14,453,176	0.63 (0.25)	0.16	0.67	0.96
Cash/TA	14,052,516	0.16 (0.19)	0.00	0.09	0.57
Working capital/TA	14,290,750	0.27 (0.27)	-0.10	0.24	0.76
Sales	13,700,288	8,336,940.06 (220,991,484.74)	77,817	773,409.50	17,869,286
Cost of Employees/Sales	13,626,334	29.46 (19.66)	4.91	25.82	67.37
№ Employees	11,218,478	34.92 (562.05)	1.00	6.00	98.00
Age	2,120,040	12.89 (11.36)	2.67	9.52	34.34

Notes: This table presents summary statistics for the variables used in the empirical tests. All statistics are based on annual frequency. The '*MP shock*$*LR_{t-1}$' variable captures the annual sum of ECB monetary shocks provided by Jarociński and Karadi (2020), broken down to the firm-level via the exposure measure, the LR. '*Cost of Employees/VA*' and '*Cost of Employees/Sales*' are the firm level labor shares defined as $\frac{Cost of Employees}{Value Added}$ and $\frac{Cost of Employees}{Sales}$, respectively. Similarly, '*Cost of Employees (g)*'; '*Value Added (g)*' are the growth rates of the name-giving balance sheet items. '*TA*' is the total balance sheet size in euros. '*Equity*' are shareholder's funds. '*LR*' stands for leverage ratio and is defined as $\frac{Total Liabilities}{Total Assets}$. '*Cash/TA*' are firm's cash reserves divided by total assets. '*Working capital/TA*' is a given firm's working capital divided by total assets. '*№ Employees*' are firm's number of employees. '*Sales*' is any given firm's total revenue. '*Age*' is the average firm age in years. All balance sheet items at firm level are provided by Bureau van Dijk's Amadeus database. The sample includes 2,139,347 firms in 79 industries.

[14] such as increasing sample size over time and non-uniform national reporting requirements across countries

[15] that is, total assets minus equity

[16] These are, Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain.

3.3.2 Industry-level data

In order to extend our analysis to the industry-level, we collapse the micro-data variables to the country-industry[17] level and recompute the labor share.[18] As an exposure variable, in order to break down the EA-wide MP shocks to the industry-level, we employ the well-established external financial dependence ratio as in Rajan and Zingales (1998), recomputed at the two-digit NACE Rev. 2 level using data from Compustat. It is defined as the industry median fraction of capital expenditures financed by external funds for mature Computstat companies over the period 1999-2017.

3.3.3 Monetary policy shocks

Our employed MP shocks for the EA as a whole are identified by and retrieved from Jarociński and Karadi (2020). These quarterly shocks are relatively new, but at the same time quickly becoming the gold standard for identified EA MP surprises because they can be interpreted as unanticipated changes in credit conditions. Jarociński and Karadi (2020) provide evidence that it is not uncommon for the stock market to depreciate after markets are surprised with lower than expected policy rates (e.g., a surprisingly strong cut may sends the signal that the economy is in worse condition than previously expected by market participants). The authors thus use sign restrictions on the joint high frequency reaction of interest rates and stock market prices, disentangling information conveyed in the shocks about the ECB's assessment of the economic outlook as well as its MP decisions. These news about the state of the economy from changes in financing conditions are important for our paper because we specifically study the effect of changing credit conditions rather than the effect of changes in the state of the economy as a whole. We sum up the quarterly shocks in order to obtain annual data along the lines of Holm et al. (2021).

3.4 Empirical results

We estimate a significant and negative reaction of the labor share after a contractionary MP shock at the firm-level. We break down EA-wide policy surprises borrowed from Jarociński and Karadi (2020) to the firm level by multiplying them with the leverage ratio of the firms. Consequently, firms with a higher leverage ratio will experience a more pronounced shock, and those with little liabilities will not be strongly affected by monetary policy. Since the MP shocks are limited to the measurement of credit conditions and are thus information neutral, firms with more leverage are naturally more exposed to MP. Local projections allow us to estimate how a firm's labor share over horizon $j > 0$ responds to MP shocks conditional on the firm's leverage ratio, and to compute the corresponding impulse responses (Plagborg-Møller and Wolf, 2021). This motivates our baseline local projections framework (Jordà, 2005) as depicted by the following equation:

$$ls_{f,t+h,s,c} = \alpha_{f,h} + \beta_h LR_{f,t-1} MP_t + \Gamma_h' X_{f,t-1} + \delta_{t,h}\zeta_{s,h} + \delta_{t,h}\kappa_{c,h} + \varepsilon_{f,t+h,s,c} \tag{23}$$

where $ls_{f,t+h,s,c}$ denotes the labour share of firm f at time t in industry s and country c.[19]. β_h are the coefficients of interest that measure the impact of the EA-wide monetary policy shock MP_{t+h} on the firm-level for every horizon. The shock is broken down to the firm level by $LR_{f,t-1}$ which denotes the (lagged) firm-specific leverage ratio. $\alpha_{f,h}$, $\delta_{t,h}$, $\zeta_{s,h}$ and $\kappa_{c,h}$ are firm-, time-, industry- and country-

[17] that is, two-digit NACE Rev. 2 industry classification

[18] We keep only observations for which both payroll expenses and value added are reported.

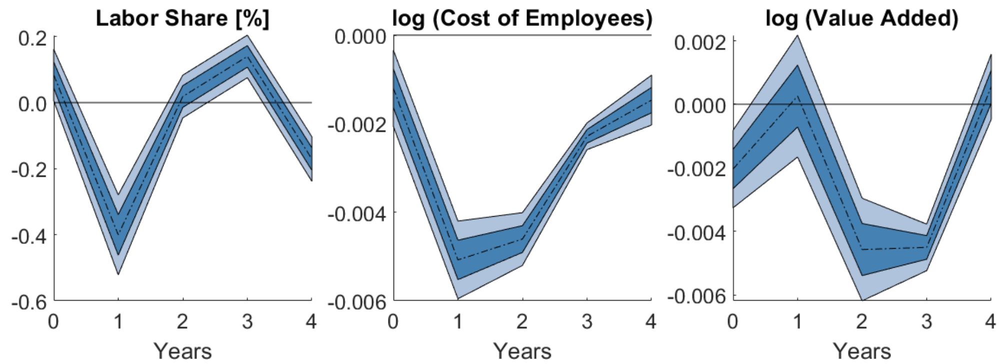
[19] The labor share is defined as $\frac{CostofEmployees}{ValueAdded}$.

fixed effects, respectively. $X_{f,t-1}$ is a vector of firm-specific and macro-economic controls. It contains $LR_{f,t+h-1}$, $LR^2_{f,t+h-1}$, $log(total\ assets)$, $\frac{working\ capital}{total\ assets}$, $\frac{cash}{total\ assets}$, $\frac{loans}{total\ assets}$, and an interaction term between $LR_{f,t+h-1}$ and (lagged) country GDP growth. $\varepsilon_{f,t+h,s,c}$ denotes the error term. For above impact horizons ($h > 0$) we also include $\varepsilon_{f,t+h-1,s,c}$. To investigate its components, we run the same specification as in Equation 23 on the dependent variables $log(value\ added)$ and $log(costs\ of\ employees)$. This enables us to elaborate the transmission channel through which MP affects the labor share. Finally, we cluster standard errors at the industry and firm-level to account for potential correlation within the unit where the shock takes place: clustering by firm allows to control for error correlation at the most granular level. The additional clustering at industry-level is conservative, given that we already include time-industry fixed effects but might help to control for potentially correlated shocks across firms in a given industry.

3.4.1 Results of the baseline specification

Figure 3.2 displays the results of our baseline specification. We find that a contractionary MP shock results in a significant decline in the labor share, and that this effect is stronger for firms with higher leverage ratios. On average, a one standard deviation MP contraction leads to a .4 percentage points decline in the labor share. This decline is driven by decreasing (log) costs of employees, which is small on impact but more pronounced after one year, and is still slightly negative after four years. (Log) value added declines slightly on impact, but react most pronounced after two years - which brings the labor share back to its initial level, after four years.

Figure 3.2: Labor Share and its components response to monetary tightening



Notes: The figure shows the responses to a one standard deviation MP tightening shock in i) the labor share; ii) $log(costs\ of\ employees)$; and iii) $log(value\ added)$ at firm-level. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm-level.

3.4.2 Ex-ante firm heterogeneity

Our analysis extends beyond the average firm: given that heterogeneity matters for the firm-specific reaction to MP, our research question is tailored around variation in business models, cost structure, and balance sheet composition – which are most meaningfully distinguished along the lines of firm-average labor share and leverage heterogeneity (see Section 3.2). As strand of literature also suggests size and age as relevant dimensions (Cloyne et al., 2018; Gertler and Gilchrist, 1994). We address firm heterogeneity empirically by splitting the data into quantile-bins according to (the firm average) size, age, leverage ratio, and labor share, then repeat the set of estimations depicted in Equation 23 on these quantiles of the data.

For each of these four dimensions of interest, data are split by these six quantiles: the lowest 10%, the 10-25%, the 25-50%, the 50-75%, 75-90%, and the top 10 (i.e. the 90-100%).[20] As a next step, we carry out regressions for these six quantiles-bins of firm characteristics. By applying this method, we not only control for *ex-ante* firm heterogeneity, but are also able to provide one estimate conditional on the average leverage ratio of the quantile.[21] [22]

Figure 3.3 displays this exercise's estimates. Thereby, each line represents a variable and each column represents a quantile-bin. We see that most heterogeneity across estimates is concentrated in the first two lines (that are, leverage ratio and labor share), where, with respect to the labor share sample spilt, we see the strongest reaction in the upper quantile and a less pronounced reaction in lower quantiles. Low labor share companies appear almost non-responsive to MP.[23] However, estimates of leverage-bins are to be taken with a grain of salt: since the leverage ratio is also our exposure variable, this variable features heterogeneity by construction.[24] Across the quantile-bins of size and age in the bottom two rows of Figure 3.3 we see much less heterogeneity, only very young firms exhibit a pronounced response with wide confidence bands, and very large and very old firms react with a negligible increase in the labor share.

When it comes to the components of the labor share –value added and costs of employees– we find substantially different reactions depending on how we group our firms. Looking at firms grouped from high to low leverage - based on the firm average across time - the reaction of value added varies strongly. However, when we group the same firms according to labor share, the reaction of costs of employees varies more strongly while heterogeneity in value added is mild (see Figure 3.11 and Figure 3.12 in Appendix 3.A.3). This finding suggests that indeed both dimensions are crucial to understand heterogeneous responses at firm-level because different frictions are at play.
While the breakdown by leverage ratio has more explanatory power for heterogeneity in the production process and value added, the labor share breakdown is more informative about the developments of labor utilization.[25] The low correlation coefficient of quantile-bins lets us already infer that we are not looking at two sides of the same coin. In the next section, motivated by the low correlation, we examine firms dominated by either high labor shares *or* high leverage ratios closely to disentangle channels more rigorously.[26]

[20]Quantiles are calculated based on sample means. For each individual firm we then calculate the sample average of the four dimensions and categorize it by putting it into one of the six quantile-bins.

[21]Our interpretation of leverage is different from Ottonello and Winberry (2020), who interpret the within-firm variation of leverage as a measure the distance to default – i.e. all else equal, a company taking on more leverage becomes more risky. Our interpretation of leverage, on the other hand, is more related to the business model: the more leverage a company has - and hence the more interest rates and debt-rollovers play a role for the overall costs of a company - the more likely it will make its strategic decision dependent on MP.

[22]Ottonello and Winberry (2020) show how the fixed effects estimator might yield biased estimates of the coefficient of interest in case of permanent differences in how firms respond to the aggregate shock. They propose to use the within-firm variation of the exposure variable as a measure to break down a macroeconomic shock to the firm-level. We discuss this issue further below in section 3.4.4, and show that our results remain robust.

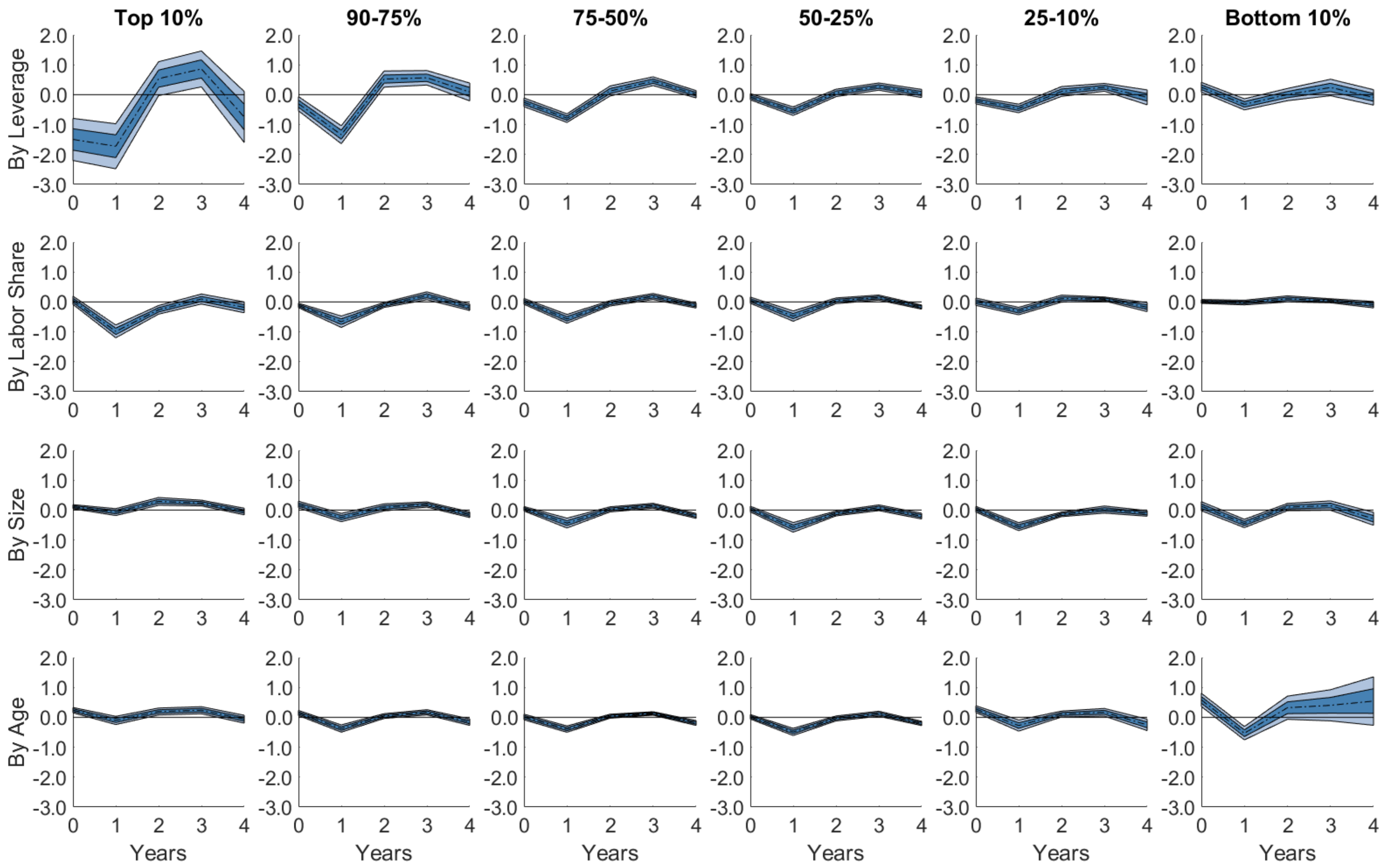
[23]Note that the correlation among groups preserves the sample correlation perfectly: while the correlation between leverage and labor share in the sample is 0.13 at the bin-level, the variables exhibit a correlation coefficient of 0.13.

[24]Below we address this issue by conditioning on high and low leverage (Section 3.4.3), and by taking the first difference of the exposure variable (Section 3.4.4).

[25]In the top leverage ratio group, for instance, payroll expenses does not react significantly on impact, but log value added increases significantly. Thus, for highly leveraged firms, the reaction in production is more central than the reaction in payroll expenses. For highest labor share firms, the converse is true: These firms are the only ones to reduce their costs of employees already on impact significantly. They alter their costs easiest by making redundancies, hiring freezes, or amending compensation schemes.

[26]Note, for example, that even the top quantile-bin by leverage ratio contains companies with very high and very low labor share.

Figure 3.3: Labor share response to monetary tightening - sample splits



Notes: Figure shows the firm-level labor share response (in percent) to a one standard deviation monetary tightening shock. The sample is split into six quantile-bins for each of the dimensions of firm heterogeneity: leverage ratio, labor share, size, and age. A firm is put into a bin based its sample mean. Quantile-bins of respective groups are depicted from left to right in a descending order. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

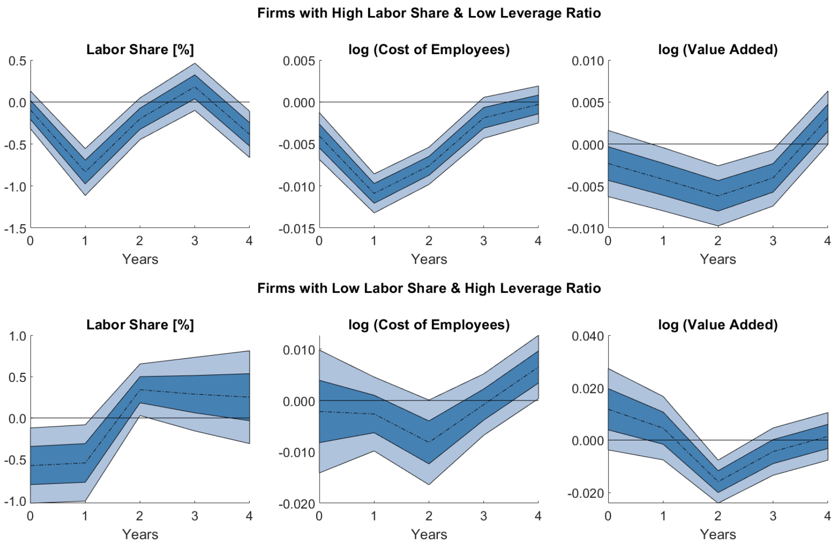
3.4.3 Firm heterogeneity: The role of firm-average leverage and labor share

Theoretical considerations of Section 3.2 highlight how firms with structurally high labor shares restructure their payroll expenses in response to MP. In addition, it lays out how firms with structurally high leverage adapt to MP by altering their production/value added. This motivates our study of the response of two very distinct types of firms: First, firms with high (average) labor shares that, at the same time, exhibit (average) low leverage ratios and, second, firms that exhibit high leverage ratios with low labor shares.[27]

i) Firms with both high labor share and low leverage ratio. Figure 3.4 depicts the responses of the labor share, costs of employees, and value added for firms where a high labor share is dominant. The labor share drops to one percent point one year after the shock.[28] This response is driven by a significant and pronounced decrease in costs of employees, which is significant until three years after the shock and reaches its minimum one year after the shock occurs. Value added declines significantly but less pronounced and less swiftly.

ii) Firms with both a low leverage ratio and a high labor share. The lower three impulse responses in Figure 3.4 depict these firms' reaction to MP. Already on impact, the labor share declines by about 0.6 percentage points. This decline can be decomposed into a strong increase in value added on impact and a small decline in costs of employees. After two years, value added decreases significantly and lets the labor share rebound. Overall, for the highly leveraged group, confidence intervals are much wider.

Figure 3.4: Labor share and its components response to monetary tightening - conditional sample splits



Notes: The upper row displays the response in the labor share of firms characterized by both a high labor share (top 25% quantile) and low leverage ratio (bottom 25% quantile) to a one standard deviation monetary tightening shock. The bottom row displays the response in the labor share of firms characterized by both a low labor share (bottom 25% quantile) and high leverage ratio (top 25% quantile) to a one standard deviation monetary tightening shock. Firms are put into a quantile-bin based on their sample means. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level

Note that a low leverage ratio is, by construction, associated with a reduced responsiveness to MP shocks – which corresponds to the credit channel in the literature. Therefore, it is particularly surprising

[27]In this subsection, we define high and low as the top and bottom 25% quantile of firm averages, respectively

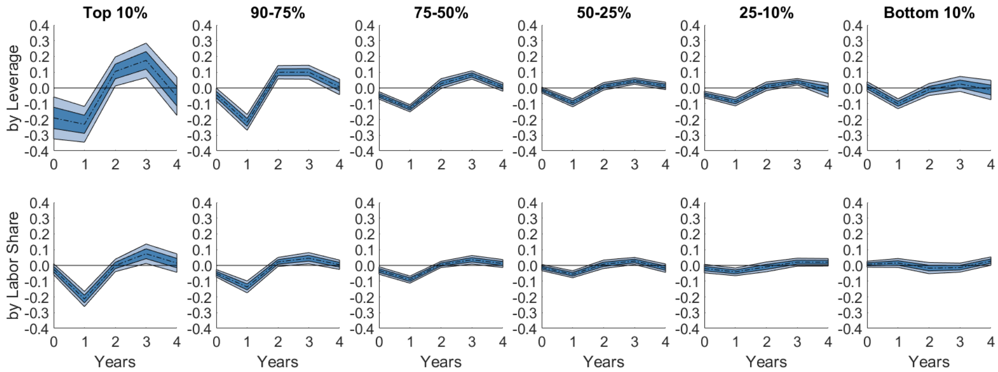
[28]Which is comparable to the reaction of firms in the top 10% leverage or labor share quantile; see figure 3.3

that the high labor share group reacts in such a pronounced and significant way while exhibiting low leverage. Thus, in order to control for possible bias in the results along the leverage-dimension, we employ the leverage ratio's within-firm variation as an exposure measure in the next subsection.

3.4.4 Alternative exposure measure: the leverage ratio's within-firm variation

Our baseline analysis employs the previous year's (that is, lagged) leverage ratio as an exposure measure to break down MP shocks to the firm-level. Ottonello and Winberry (2020) propose the leverage ratio's within-firm variation[29] as an exposure measure, where this within-firm variation is meant to also control for permanent differences across firms.

Figure 3.5: Labor Share response to monetary tightening - sample splits & alternative exposure



Notes: The figure shows the baseline estimation results (i.e. the response in the firm-level labor share after a one standard deviation monetary policy tightening shock) based on the leverage ratio's within-firm variation as an alternative exposure measure. The sample is split into six quantile-bins by the dimensions labor share and leverage ratio. Firms are put into bins based on the sample mean. Quantile-bins of respective groups are depicted from left to right in a descending order. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

Figure 3.5 displays the results when we carry out the estimation by quantile-bins as described in Section 3.4.2. We see a very similar pattern as before: The most pronounced labor share reaction is in the top quantiles of leverage and labor share.[30] Figure 3.6 depicts the results when we estimate separately for firms with high labor share (leverage ratio) and a low leverage ratio (labor share). We find that the still highly significant decline in the labor share of labor-intensive firms is again driven by a decline in costs of employees. Highly leveraged firms decrease their labor share while increasing value added. However, in the latter case, results are insignificant.

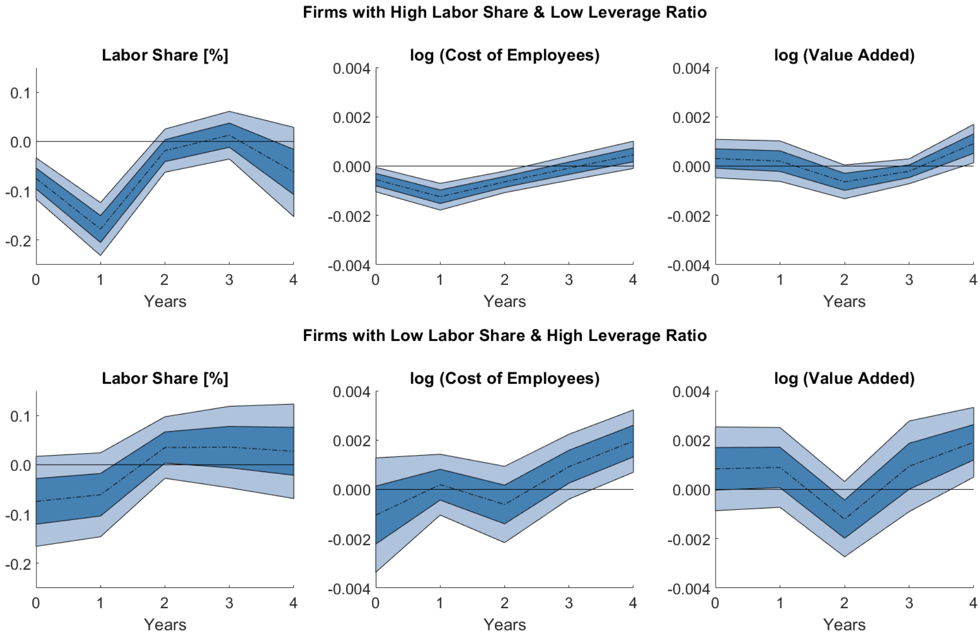
3.4.5 Industry-level aggregate estimation

As firm-level research is a relatively young, fast-evolving field in macroeconomic research, there is no established standard approach for measuring firms' financial sensitivity: The choice of exposure measure as well as the efficiency and unbiasedness of resulting estimators hinges on assumptions about the data generating process. At the industry-level, however, external financial dependence is a well-established, mature measure for industries' idiosyncratic financial sensitivity since the seminal work of Rajan and Zingales (1998). Hence, we use their exposure measure to provide industry-level evidence on

[29] i.e. $LR_{t,f} - E_f\{LR_{f,t}\}$

[30] Figure 3.13 in Appendix 3.A.4 depicts the sub-components of the labor share. For the responses of log value added, we see most heterogeneity across the quantiles of leverage, whereas the heterogeneity in the response of costs of employees is most pronounced across the quantiles of the labor share.

Figure 3.6: Labor share response to monetary tightening - conditional sample splits & alternative exposure



Notes: The figure shows estimation results based on the leverage ratio's within-firm variation as an alternative exposure measure. The upper row displays the response in the labor share of firms characterized by both a high labor share (top 25% quantile) and low leverage ratio (bottom 25% quantile) to a one standard deviation monetary tightening shock. The bottom row displays the response in the labor share of firms characterized by both a low labor share (bottom 25% quantile) and high leverage ratio (top 25% quantile) to a one standard deviation monetary tightening shock. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

the robustness of our findings. We do so by collapsing the firm-level dataset to the industry-level.[31] This step facilitates comparability with other macro-studies such as Cantore et al. (2020). By collapsing the dataset, we also render our estimation more robust to firms entering and exiting the market as well as potential sample-selection bias in our data set. Thus, we estimate the following industry-level equation:

$$ls_{t+h,s,c} = \alpha_{s,h}\kappa_{c,h} + \beta_h EFD_s MP_t + \gamma_h EFD_s \Delta GDP_{t-1,c} + \delta_{t,h}\kappa_{c,h} + \varepsilon_{t+h,s,c} \tag{24}$$

where $ls_{t+h,s,c}$ denotes the labor share in industry s at time t and country c for horizon h. EFD_s denotes a time-invariant measure of industry-specific external financial dependence[32]. $\alpha_{s,h}$ are industry-fixed effects, $\kappa_{c,h}$ are country-fixed effects, $\delta_{t,h}$ are time-fixed effects. β_h are the main coefficients of interest for horizon h, measuring the impact of a EA-wide monetary policy shock MP_t, depending on the extent of reliance on external finance EFD_s of industry s. γ_h aims to control for the effect of real economic activity on the labor share in industries with different external finance dependence. For impact horizons $h > 0$, we also include $\varepsilon_{t+h-1,s,c}$. We cluster standard errors at the industry and country-level.

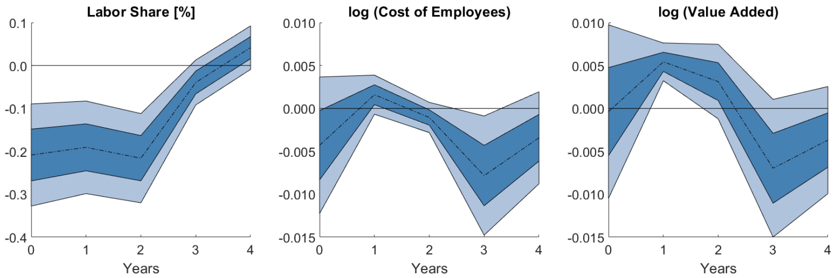
Figure 3.7 displays the main estimation results. In line with the firm-level findings, we report a significantly negative reaction of the aggregate labor share at the industry-level after a contractionary MP shock. The labor share has a significantly negative reaction on impact and remains negative for the following three years. Payroll expenses have a significantly negative reaction three years after the shock

[31] Two-digit NACE Rev. 2 industry classification

[32] which is the industry-median fraction of capital expenditures financed by external funds for mature Compustat companies over the 1999-2017 period, as proposed by Rajan and Zingales (1998).

occurs. Value added displays a humped-shaped pattern with a significantly positive reaction after one year, before turning negative, albeit insignificantly.

Figure 3.7: Labor share and its components response to monetary tightening at industry-level



Notes: This figure depicts the industry-level response to a one standard deviation monetary tightening shock in the labor share, log of cost of employees, and log of value added, from left to right. The exposure measure to break down the EA-wide monetary policy shocks is the time-invariant industry-median fraction of capital expenditures financed by external funds for mature Computstat companies over the period 1999-2017. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

3.5 Interpretation

Our empirical results support the hypotheses stated in Section 3.2. In line with NK theory, we find that, on average, firms react to a contractionary MP shock with a decline in the labor share. Declines are more pronounced for companies with higher labor shares and for companies with higher leverage. Looking at companies mainly characterized by a high labor share, we see that the labor share decline is driven relatively more by a decrease in costs of employees. In contrast, highly leveraged companies exhibit a pronounced reaction in value added, which drives the response of the labor share. Below we discuss how the results square with the existing literature and derive policy implications.

3.5.1 Relation to the literature

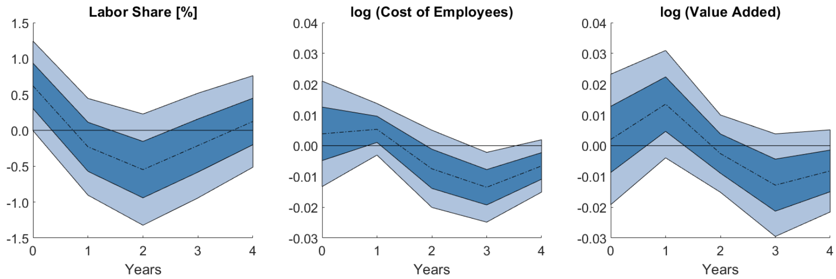
Our econometric strategy is related to approaches put forward by Ottonello and Winberry (2020) as well as Jeenas (2019), who deploy financial exposure variables to break MP shocks down to the firm-level. Notwithstanding, the research questions they tackle are fundamentally different from ours. Our findings complement their reasoning and results.[33] In contrast to these studies, however, we base our analysis on a data set that is not limited to listed firms. Below we discuss the particularities of listed firms and liquidity as an alternative measure of firm sensitivity to MP.

Listed firms Listed firms are regulated fundamentally different compared to non-listed ones, they tend have a more dispersed ownership structure, to operate more globally, and they are in a minority. In our dataset, less than 0.1% of firms are listed. A range of other studies (Autor et al., 2020; Jeenas, 2019; Ottonello and Winberry, 2020) focuses on listed firms when it comes to MP transmission or labor share analysis because of these firms' unique reporting and ownership characteristics. In order to relate to their research, we repeat the baseline estimation for the sample of listed firms exclusively. Figure 3.8 displays these results. We find that, on average, listed firms decrease their labor share by about 0.7 percentage points after an MP tightening shock. However, heterogeneity across listed firms is large, rendering the response insignificant. Costs of employees react significantly negative three years after the monetary

[33]Especially that the impact of MP is most pronounced after one year is a result that all three studies have in common.

tightening occurs.[34] The response in value added is insignificant and on impact close to zero. Given that most listed companies in Europe operate globally, are strong exporters, and have ample liquidity Sharpe and Suarez (2015), it is not surprising that they are less sensitive to MP shocks at home – they are more exposed to the global financial and monetary cycle.

Figure 3.8: Labor Share and its components response to monetary tightening - listed firms



Notes: The figure depicts the responses to a one standard deviation monetary tightening shock in i) the labor share, ii) $log(costs\ of\ employees)$ and iii) $log(value\ added)$ (from left to right) of listed firms. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

Liquidity vs. Leverage Sharpe and Suarez (2015) point out that, when it comes to making new investments, listed firms' chief financial officers tend to care more about the liquidity than borrowing costs.[35] Jeenas (2019) builds a model on these arguments. He argues that corporations face fixed treasury costs[36] to motivate his use of liquidity as the firm-level exposure variable. Inspired by his research, we conduct our analysis with two alternative exposure variables related to firms' liquidity conditions. First, gross liquidity measured by $\frac{cash}{total\ assets}$ and, second, net liquidity, measured by $\frac{cash - current\ liabilities}{current\ liabilities}$.[37] Results are displayed in the upper row of Figure 3.9. For both variables, we find a significant decrease in the wage share. For listed firms (lower row of the same figure), only gross liquidity leads to a significant decrease in the labor share.

Micro vs. macro In Section 3.7, we show that our firm-level findings are robust to the industry-level. However, these industry results contrast Cantore et al. (2020) and Nekarda and Ramey (2020) who report a rise in the labor share after a monetary tightening occurs.

3.5.2 The labor share is a catalyst for monetary policy

Our research question is inspired by insights from the literature on firm-level frictions (Carlstrom and Fuerst, 1995; Lanteri, 2018; Lucas and Stokey, 1985; Ramey and Shapiro, 2001) as well as the macroeconomic literature on MP transmission to the labor share (Cantore et al., 2020; Galí, 2015). Our results are in line with a cash-in-advance reasoning for companies with a big payroll and fit in with the main arguments of capital irreversibility (Lanteri, 2018; Ramey and Shapiro, 2001). In addition, our results are in line with the theoretical predictions of the textbook NK model (Cantore et al., 2020; Galí, 2015) — i.e. that monetary tightening leads to a decline in the labor share. The empirical validity of this

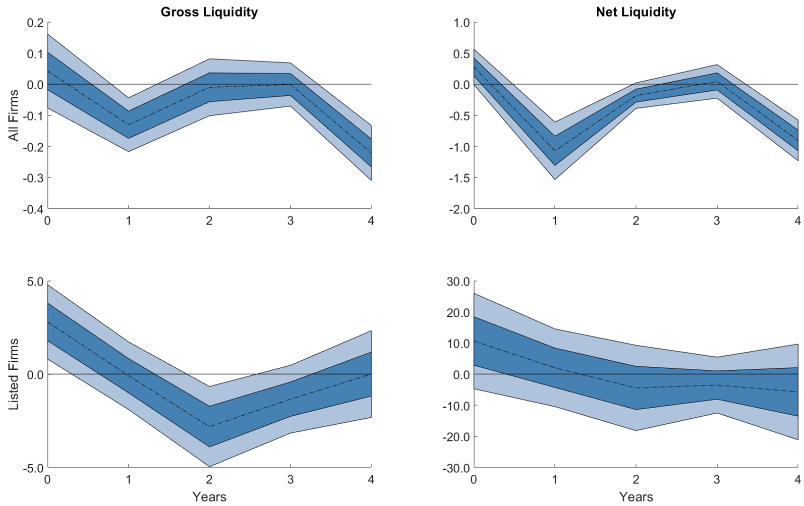
[34]The less timely reaction in costs of employees makes sense because European labor market frictions are comparatively large (Swanson, 2020).

[35]When CFOs were asked why they would not care about changes in their borrowing costs (which only a minority does), they stated that they have ample cash or cash flow to finance investments.

[36]e.g. costs related to issuing debt, changing maturity structure and managing cash reserves

[37]We winsorize net liquidity at the 99th and 1th percentile, gross liquidity is capped to be on the interval of 0 and 1 based on the sample selection in Appendix 3.A.1.

Figure 3.9: Labor Share response to monetary tightening - alternative exposures with listed firms



Notes: The figure shows the baseline estimation result, i.e., the response in the firm-level labor share after a one standard deviation monetary tightening shock. Left-hand side figures are based on *gross* liquidity as an alternative exposure measure. Right-hand side figures are based on *net* liquidity as an alternative exposure measure. The upper row displays the response all firms in the sample. The bottom row displays the response of listed firms. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm levels.

theoretical prediction, however, is challenged by the findings of Cantore et al. (2020) and Nekarda and Ramey (2020), who report a puzzling positive reaction in the labor share after a monetary tightening [38]. Our dataset enables us to control for firm-level dynamics, we deploy identified MP shocks by Jarociński and Karadi (2020) and analyze firms in the EA where no considerable downward trend in the labor share is reportedly present (Gutierrez and Piton, 2020). Therefore, our results are not explained by the rise of superstar firms (Autor et al., 2020) or other trends that affect both mark-up and labor share in the long-run. Thus, our paper is the first to robustly substantiate the theoretical predictions of the NK model about the response in the labor share to an MP shock under the microscope.

3.5.3 Policy implications

This study informs policy and academic discussions about the effectiveness and potential redistributive effects of MP as well as about European integration. Through the lens of the NK model, short-term effects of MP on the mark-up and the labor share are a cornerstone of effective policy. We find that the heterogeneity of firms affects the transmission mechanism. Firms with a higher labor share react more strongly to MP, firms with a lower labor share react significantly less - highlighting that policymakers need to think of ways to transmit MP to firms with low labor shares. This is of particular relevance given that firm structures are heterogeneous not just across countries and regions, but also over time.
Figure 3.10 illustrates substantial labor share heterogeneity across different jurisdictions, and suggests that regional disparities are strongly correlated within national borders. Given our results, uniform MP in the EA might be more potent at stabilizing business cycles in some member states compared to others.[39] A prioritization and serious consideration of both a banking union and a capital markets union in policy discussion and design is vital, as these will help alleviate heterogeneous effects of MP.

[38] Nekarda and Ramey (2020) discuss the price markup, which they measure by the inverse labor share.

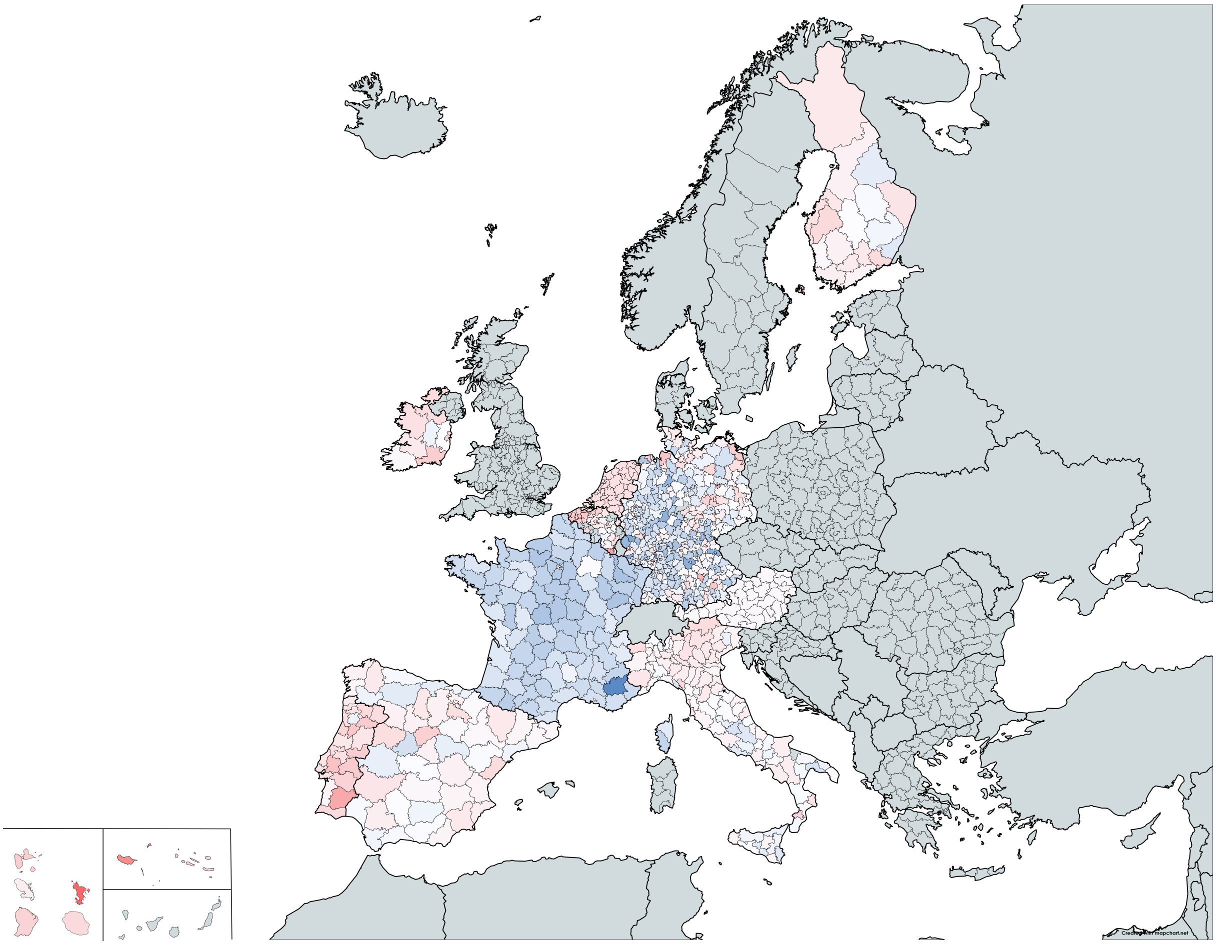
[39] For instance, insolvency laws, banking practices, labor rights, and social security schemes differ across member states. Digitalization and the development of digital infrastructure evolves at different speeds across EA countries.

The redistributive effect between capital and labor is an interesting avenue to study at a more granular level, which we leave to future research: as indicated by our results, MP might lead to redistribution in areas that rely more heavily on services and labor-intensive production, such as urban, densely populated spaces. These areas are already prone to valuation effects of MP due to rising real estate and property prices.

The decline of the labor share is the subject of extensive research. Digitalization and the rise of superstar firms are highlighted, among others, as reasons for a decline in the US. If such a decline also materializes in the EA in the future, it is likely that MP will become less effective, according to our research.

From the perspective of a central bank, redistribution might be a promising tool to steer inflation rates closer to target. Apart from MP, fiscal authorities and policymakers can monitor the redistributive effects of MP to determine if they are in line with the political agenda. Potential conflicts between the goals of the institutions can either be addressed by enhancing automatic stabilizers (such as a European unemployment insurance or basic income) or a legislative update of the ECB's toolkit toward transmission channels that are more independent from firm and bank heterogeneity.

Figure 3.10: Labor Shares across euro area regions



Notes: The image displays sample averages of aggregated labor shares in EA NUTS-3 regions of eleven founding members (EA-11). (Dark) blue colors mark regions with (very) high labor shares, (dark) red areas mark regions with (very) low labor shares. Teal-colored areas are missing data or not part of the EA at its inception.

3.6 Conclusion

Based on widely used structural models, we derive the prediction that a contractionary MP shock leads to a decline in the typical firms' labor share of value added. Refining and extending this hypothesis, we are first to propose and empirically validate that heterogeneity in the factor input cost structure matters most for the labor shares' response to MP at firm-level: The higher the labor share, the more firms' costs are driven by the payroll. Consequently, after an MP shock occurs, such firms react by altering the costs of employees. Conversely, firms characterized by a high leverage ratio primarily react by adjusting production and value added to alter their cost structure.

Consistent with this simple prediction, our article documents empirical evidence by employing Amadeus BvD data of firm-level reporting across the EA. As for the MP surprises, identified EA-wide shocks by Jarociński and Karadi (2020) are broken down to the firm-level using an exposure measure approach suggested by Ottonello and Winberry (2020) and Jeenas (2019).

We show a set of highly robust and cohesive findings that we believe should be accommodated in the discussion about firm-level effects of MP generally but especially with respect to their labor share. These include: the labor share and the leverage ratio, as two very distinct dimensions of firm heterogeneity; a difference demonstrated by the low empirical correlation (.13). After borrowing costs change due to MP shocks, these are both crucial for understanding firm-level dynamics and, thus, the labor share.

While firms characterized by both low leverage and high labor share are very responsive, the opposite does not hold true: Firms with high leverage yet low labor shares are decisively less responsive. Overall, we find that these dimensions are more informative about firms' response than age and size. Therefore, we emphasize the role of labor share when it comes to MP transmission.

Our article complements research that addresses the redistributive effects of MP by providing micro-level evidence, thus enabling us to investigate transmission channels more closely. We yield important policy conclusions and point out that future work needs to incorporate the labor share when analyzing MP transmission.

3.A Appendix to Chapter Three

3.A.1 Sample selection

Our sample selection operations are carried out along the lines of Kalemli-Özcan et al. (2019); Ottonello and Winberry (2020), and Belenzon et al. (2017).

- We keep only corporate industry firms (and thus drop banks, financial companies, foundations and research institutes, insurance companies, mutual and pension funds, trusts, private equity firms, public authorities, states, governments and venture capital firms).
- We keep only unconsolidated firm statements (i.e. ”U1” and ”U2”).
- We drop observations with missing dates, missing firm identifiers, and duplicates.
- We drop observations with negative costs of employees, negative value added, negative sales, negative total assets, or negative equity.
- We replace negative values of cash (and cash equivalent) with missing values.
- For each given year, we trim growth variables of value added and cost of employees at the 5 and 95 percentiles.[40] This treatment is helps eliminate outliers unrelated to monetary policy. At the same time, we do not have to eliminate large or small values of the main dependent variables of interest i.e. the wage share, costs of employees, or value added.[41]
- We replace cash to total assets with missing if larger than 1.
- We replace working capital to total assets with missing if larger 1 or smaller -1.
- We compute the labor share by dividing costs of employees by value added, keeping only observations with non-missing labor shares.

[40] Ottonello and Winberry (2020) drop observations with growth rates larger than 1 or smaller than -1, which has numerical advantages but removes values of fast growing firms disproportionately. A growth rate of larger than 1 is just doubling costs, which is not rare among small firms; a growth rate of -1 means going to zero, not an entirely uncommon phenomenon.

[41] Winsorization of the latter two would lead to a disproportionate elimination of very small and very large companies. Extreme growth rates in any direction, however, come in handy to detect outliers, because not only are they unlikely to be related to monetary policy decisions, but they judge observations relative to another observation by the same firm.

3.A.2 Labor share and leverage by industry

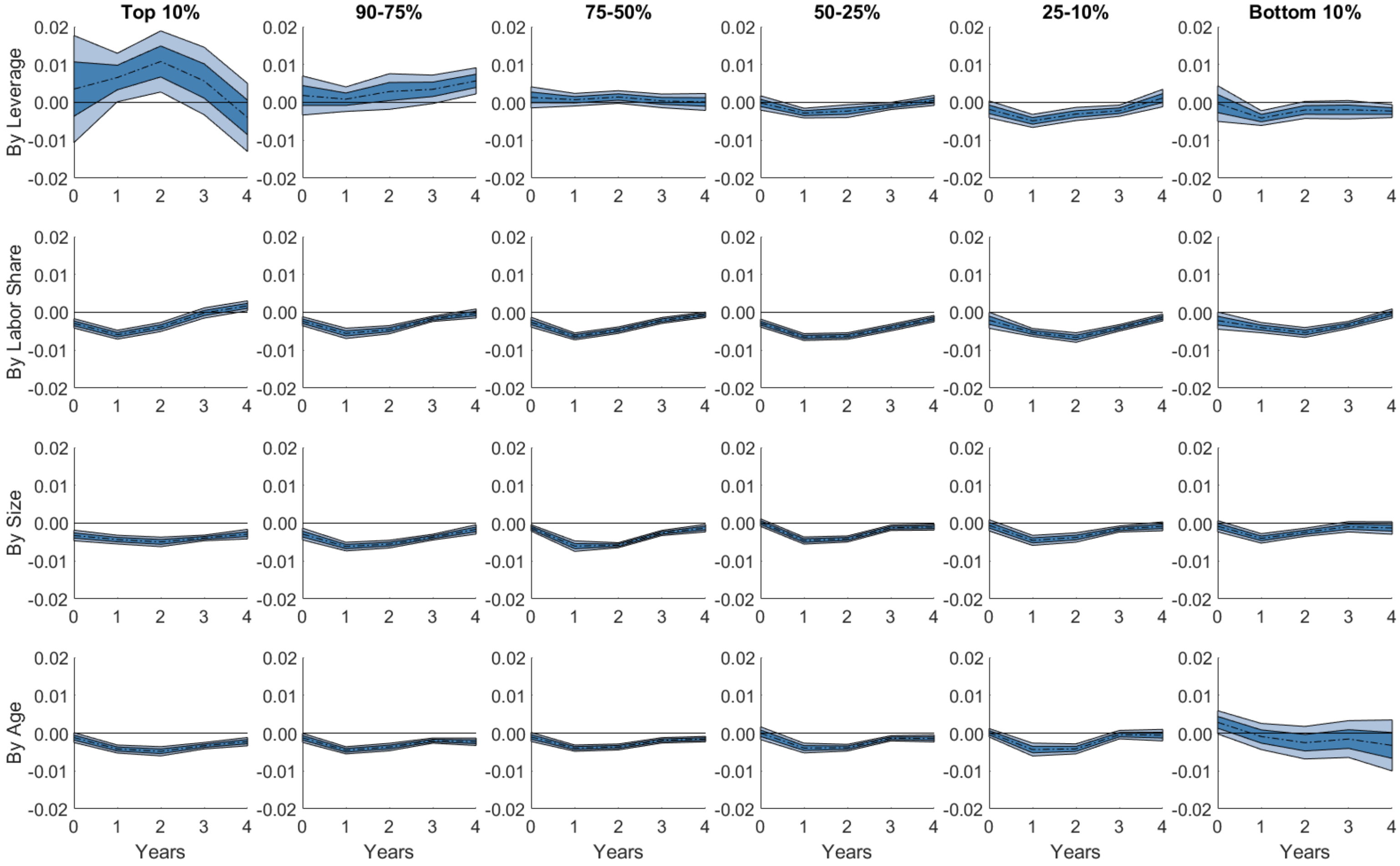
Table 3.2: Labor Share and leverage ratio by two-digit NACE code

Wage share	Leverage ratio	NACE code	Industry	Wage share	Leverage ratio	NACE code	Industry
0.23	0.38	65	Insurance, reinsurance and pension funding	0.53	0.51	47	Retail trade, except of motor vehicles and motorcycles
0.25	**0.72**	**12**	**Manufacture of tobacco products**	0.54	0.51	23	Manufacture of other non-metallic mineral products
0.25	0.44	19	Manufacture of coke and refined petroleum products	0.54	0.37	16	Manufacture of wood and of products of wood and cork
0.26	0.48	68	Real estate activities	0.54	0.53	56	Food and beverage service activities
0.28	0.44	94	Activities of membership organisations	0.54	0.48	13	Manufacture of textiles
0.31	0.57	6	Extraction of crude petroleum and natural gas	0.54	0.54	66	Activities auxiliary to financial service
0.32	0.56	35	Electricity, gas, steam and air conditioning supply	0.54	0.54	50	Water transport
0.32	0.58	42	Civil engineering	0.55	0.62	71	Architectural and engineering activities
0.32	0.48	60	Programming and broadcasting activities	0.55	0.56	93	Sports activities and amusement
0.33	0.34	64	Financial service activities, except insurance and pension funding	0.55	0.49	25	Manufacture of fabricated metal products
0.36	0.52	49	Land transport and transport via pipelines	0.55	0.51	81	Services to buildings and landscape activities
0.36	0.40	17	Manufacture of paper and paper products	0.55	0.52	18	Printing and reproduction of recorded media
0.37	0.27	43	Specialised construction activities	0.56	0.29	96	Other personal service activities
0.39	0.45	59	Motion picture, video and television programme production	0.56	0.52	28	Manufacture of machinery and equipment n
0.41	0.34	38	Waste collection, treatment and disposal activities	0.56	0.51	32	Other manufacturing
0.42	0.60	7	Mining of metal ores	0.57	0.58	74	Other professional, scientific and technical activities
0.43	0.47	52	Warehousing and support activities for transportation	0.57	0.47	46	Wholesale trade, except of motor vehicles and motorcycles
0.43	0.52	24	Manufacture of basic metals	0.58	0.30	91	Libraries, archives, museums and other cultural activities
0.43	0.66	41	Construction of buildings	0.58	0.50	73	Advertising and market research
0.44	0.50	11	Manufacture of beverages	0.59	0.48	21	Manufacture of basic pharmaceutical products
0.44	0.48	9	Mining support service activities	0.59	0.47	2	Forestry and logging
0.45	0.56	87	Residential care activities	0.59	0.68	39	Remediation activities and other waste management services
0.46	0.69	69	Legal and accounting activities	0.60	0.46	27	Manufacture of electrical equipment
0.46	0.60	1	Crop and animal production, hunting and related service activities	0.60	0.60	70	Activities of head offices; management consultancy activities
0.47	0.59	77	Rental and leasing activities	0.60	0.43	26	Manufacture of computer, electronic and optical products
0.47	0.38	8	Other mining and quarrying	0.61	0.50	86	Human health activities
0.47	0.45	14	Manufacture of wearing apparel	0.64	0.55	79	Travel agency, tour operator
0.48	0.48	22	Manufacture of rubber and plastic products	0.64	0.54	82	Office administrative, office support
0.49	0.46	92	Gambling and betting activities	0.66	0.66	51	Air transport
0.49	0.49	20	Manufacture of chemicals and chemical products	0.66	0.51	55	Accommodation
0.50	0.52	33	Repair and installation of machinery and equipment	0.68	0.54	84	Public administration and defence; compulsory social security
0.50	0.45	45	Wholesale and retail trade and repair of motor vehicles	0.69	0.46	58	Publishing activities
0.51	0.44	15	Manufacture of leather and related products	**0.72**	**0.30**	**72**	**Scientific research and development**
0.51	0.56	30	Manufacture of other transport equipment	0.72	0.46	62	Computer programming, consultancy and related activities
0.51	0.51	61	Telecommunications	0.75	0.56	85	Education
0.52	0.50	10	Manufacture of food products	0.80	0.61	90	Creative, arts and entertainment activities
0.53	0.48	63	Information service activities	0.80	0.63	53	Postal and courier activities
0.53	0.60	29	Manufacture of motor vehicles, trailers and semi-trailers	0.81	0.61	80	Security and investigation activities
0.53	0.63	36	Water collection, treatment and supply	0.87	0.57	78	Employment activities
				0.90	0.59	31	Manufacture of furniture

Notes: The table displays the sample average wage share and leverage ratio by ”Statistical Classification of Economic Activities in the European Community” (NACE) Revision 2 for SIC 2-digit industries in percent. It is ordered from low to high wage share industries.

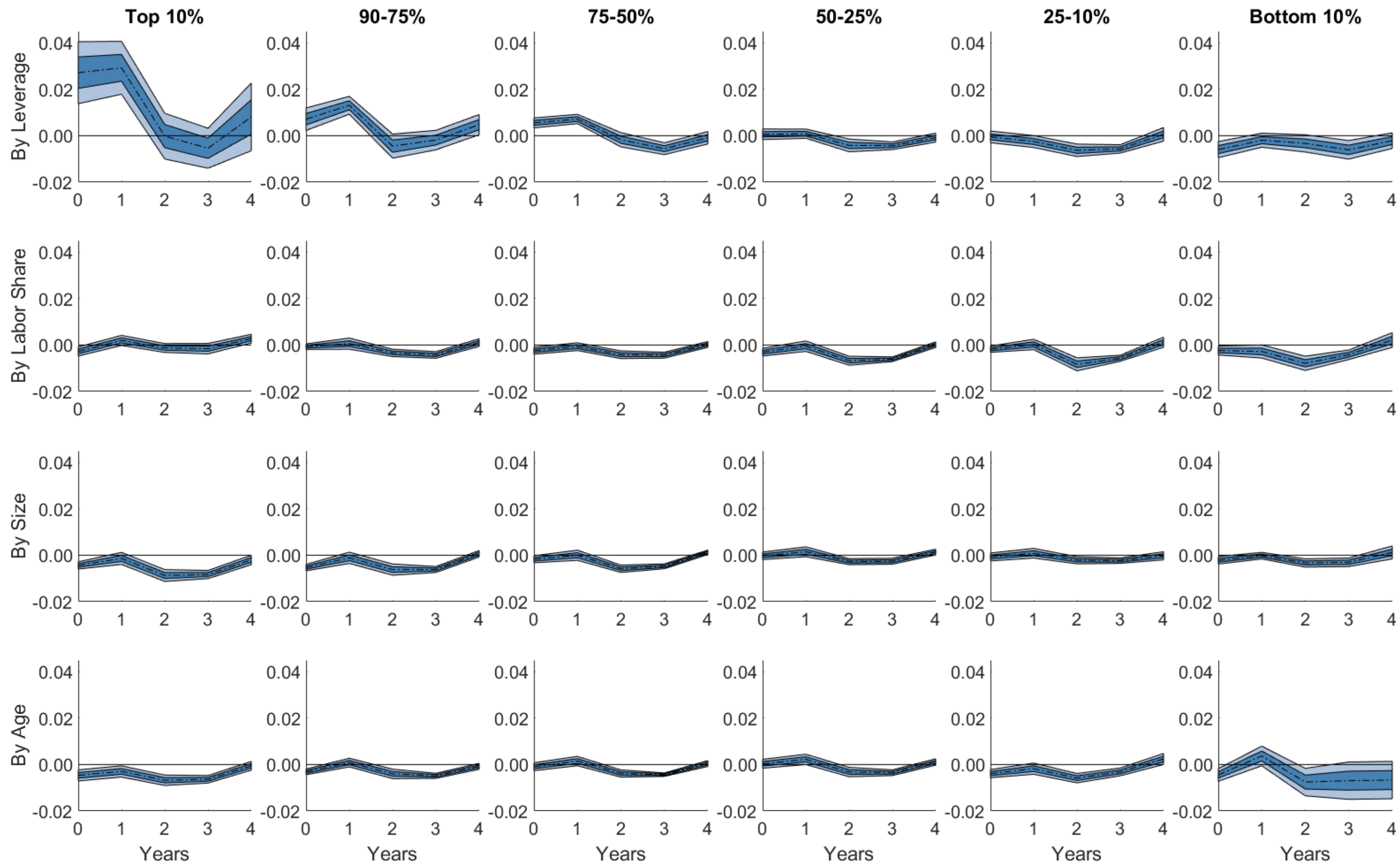
3.A.3 Responses of the components of the labor share

Figure 3.11: Payroll response to monetary tightening - sample splits



Notes: Figure shows the firm-level response of payroll expenses to a one standard deviation monetary tightening shock. The sample is split into six quantile-bins for each of the dimensions of firm heterogeneity: leverage ratio, labor share, size, and age. A firm is put into a bin based on its sample mean. Quantile-bins of respective groups are depicted from left to right in a descending order. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

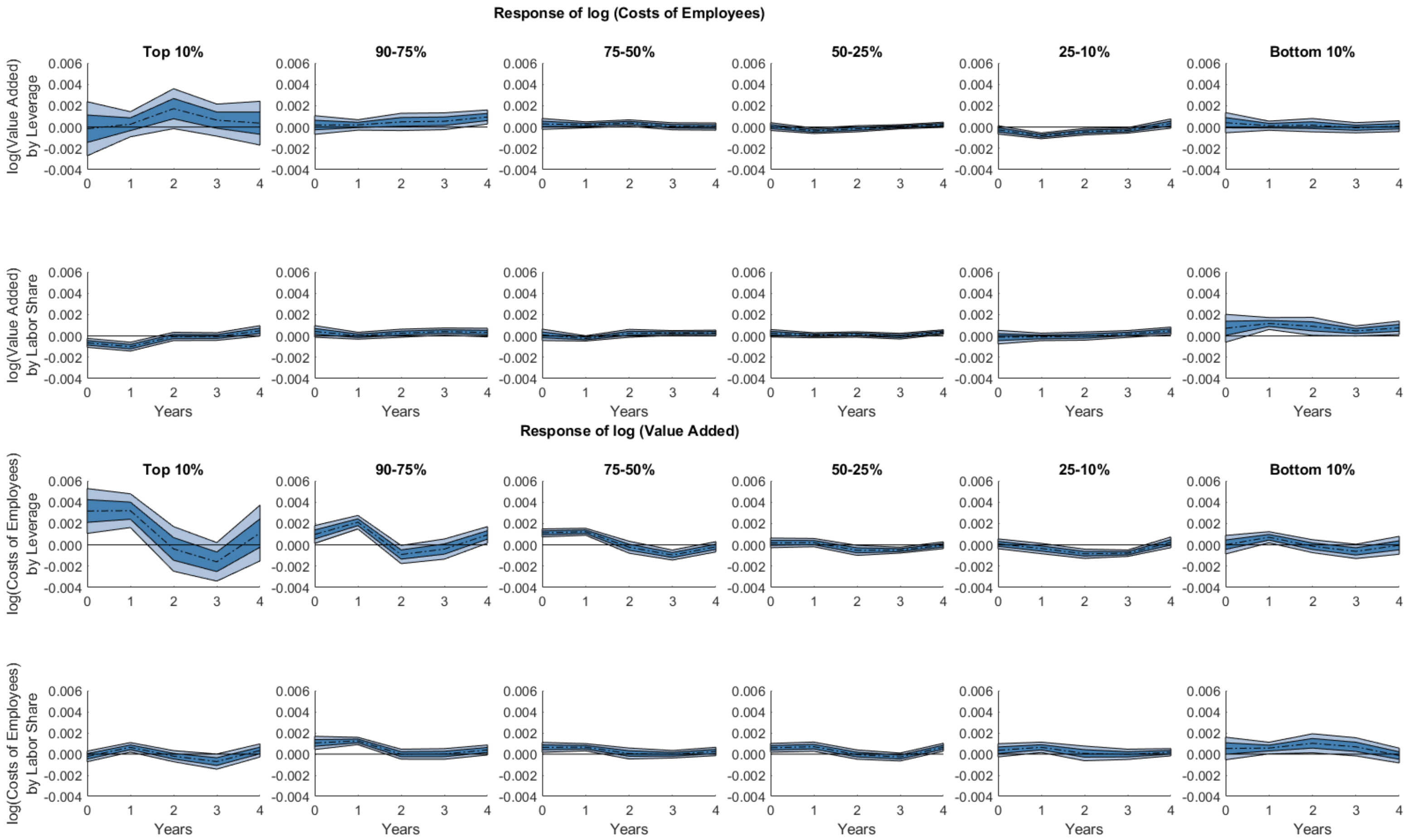
Figure 3.12: Value added response to monetary tightening - sample splits



Notes: Figure shows the firm-level response of value added to a one standard deviation monetary tightening shock. The sample is split into six quantile-bins for each of the dimensions of firm heterogeneity: leverage ratio, labor share, size, and age. A firm is put into a bin based on its sample mean. Quantile-bins of respective groups are depicted from left to right in a descending order. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

3.A.4 Responses of labor share components - within variation

Figure 3.13: Value added & payroll response to monetary tightening - sample splits & within variation



Notes: Figure shows the firm-level response of value added (upper panel) and costs of employees (lower panel) to a one standard deviation monetary tightening shock. The exposure measure is the within variation of leverage. The sample is split into six quantile-bins for the leverage ratio and the labor share. A firm is put into a bin based on its sample mean. Quantile-bins of respective groups are depicted from left to right in a descending order. We plot 95 (68) percent (dark) blue confidence bands calculated from standard errors clustered at the industry and firm level.

References

Abadie, A. and Gardeazabal, J. (2003). The economic costs of conflict: A case study of the basque country. *American economic review*, 93(1):113–132.

Alloza, M. (2017). Is fiscal policy more effective in uncertain times or during recessions? Technical report, Banco de Espana.

Altig, D., Baker, S., Barrero, J. M., Bloom, N., Bunn, P., Chen, S., Davis, S. J., Leather, J., Meyer, B., Mihaylov, E., et al. (2020). Economic uncertainty before and during the covid-19 pandemic. *Journal of Public Economics*, 191:104274.

Andréadès, A. M. (2013). *History of the Bank of England*. Routledge.

Auclert, A. (2019). Monetary policy and the redistribution channel. *American Economic Review*, 109(6):2333–2367.

Auerbach, A. J. and Gorodnichenko, Y. (2012). Measuring the output responses to fiscal policy. *American Economic Journal: Economic Policy*, 4:1–27.

Autor, D., Dorn, D., Katz, L. F., Patterson, C., and Van Reenen, J. (2020). The fall of the labor share and the rise of superstar firms. *The Quarterly Journal of Economics*, 135(2):645–709.

Bacchiocchi, E. and Fanelli, L. (2015). Identification in Structural Vector Autoregressive Models with Structural Changes, with an Application to US Monetary Policy. *Oxford Bulletin of Economics and Statistics*, 77:761–779.

Baker, S. R., Bloom, N., and Davis, S. J. (2016). Measuring economic policy uncertainty. *The quarterly journal of economics*, 131(4):1593–1636.

Barro, R. J. and Redlick, C. J. (2011). Macroeconomic Effects From Government Purchases and Taxes. *The Quarterly Journal of Economics*, 126(1):51–102.

Basu, S. and Bundick, B. (2017). Uncertainty shocks in a model of effective demand. *Econometrica*, 85:937–958.

Belenzon, S., Chatterji, A., and Daley, B. (2017). Eponymous entrepreneurs. *American Economic Review*, 107(6):1638–55.

Ben Zeev, N. and Pappa, E. (2017). Chronicle of a war foretold: The macroeconomic effects of anticipated defence spending shocks. *The Economic Journal*, 127:1568–1597.

Berger, H. and De Haan, J. (2002). Are small countries too powerful within the ecb? *Atlantic Economic Journal*, 30(3):263–282.

Bergsten, C. F. and Gagnon, J. E. (2017). *Currency conflict and trade policy: A new strategy for the United States*. Columbia University Press.

Bernanke, B. (1983). Irreversibility, uncertainty, and cyclical investment. *The Quarterly Journal of Economics*, 98:85–106.

Bertrand, M., Duflo, E., and Mullainathan, S. (2004). How much should we trust differences-in-differences estimates? *The Quarterly journal of economics*, 119(1):249–275.

Bilbiie, F., Meier, A., and Müller, G. J. (2008). What accounts for the changes in u.s. fiscal policy transmission? *Journal of Money, Credit and Banking*, 40:1439–1470.

Bindseil, U. (2014). *Monetary policy operations and the financial system*. OUP Oxford.

Bindseil, U. (2016). Evaluating monetary policy operational frameworks. In *Speech at the Jackson Hole conference on*, volume 31.

Blanchard, O. and Perotti, R. (2002). An empirical characterization of the dynamic effects of changes in government spending and taxes on output. *The Quarterly Journal of Economics*, 117(4):1329–1368.

Bloom, N. (2009). The impact of uncertainty shocks. *Econometrica*, 77(3):623–685.

Bloom, N. (2014). Fluctuations in uncertainty. *Journal of Economic Perspectives*, 28(2):153–76.

Bloom, N., Floetotto, M., Jaimovich, N., Eksten, I. S., and Terry, S. J. (2018). Really Uncertain Business Cycles. *Econometrica*, 86:1031–1065.

Bomfim, A. and Brayton, F. (1998). Long-run inflation expectations and monetary policy. *Dillen, H. dan Hopkins, E., Forward Interest Rates and Inflation Expectations: The Role of Regime Shift Premia and Monetary Policy, BIS*.

Born, B., Müller, G. J., Schularick, M., and Sedláček, P. (2019a). The costs of economic nationalism: evidence from the brexit experiment. *The Economic Journal*, 129(623):2722–2744.

Born, B., Müller, G. J., Schularick, M., and Sedlacek, P. (2019b). Stable genius? the macroeconomic impact of trump.

Bouakez, H., Chihi, F., and Normandin, M. (2014). Measuring the effects of fiscal policy. *Journal of Economic Dynamics and Control*, 47:123–151.

Boumans, D. and Garnitz, J. (2019). Ifo world economic survey database–an international economic expert survey. *Jahrbücher für Nationalökonomie und Statistik*, 237(1):71–80.

Burzoni, M., Riedel, F., and Soner, H. M. (2021). Viability and arbitrage under knightian uncertainty. *Econometrica*, 89(3):1207–1234.

Caldara, D. and Kamps, C. (2017). The analytics of svars: A unified framework to measure fiscal multipliers. *The Review of Economic Studies*, 84(3):1015–1040.

Campos, N. F., Fidrmuc, J., and Korhonen, I. (2017). Business cycle synchronisation in a currency union: Taking stock of the evidence. *Bank of Finland Research Discussion Paper*, (28).

Canova, F. and De Nicolo, G. (2002). Monetary disturbances matter for business fluctuations in the G-7. *Journal of Monetary Economics*, 49(6):1131–1159.

Canova, F. and Pappa, E. (2011). Fiscal policy, pricing frictions and monetary accommodation. *Economic Policy*, 26(68):555–598.

Cantore, C., Ferroni, F., and León-Ledesma, M. (2020). The missing link: monetary policy and the labor share. *Journal of the European Economic Association*.

Carlstrom, C. and Fuerst, T. (1995). Why isn't investment more sensitive to interest rates: Evidence from surveys. *Journal of Monetary Economics*, 36(2):247–267.

Carriero, A., Clark, T. E., and Marcellino, M. (2016). Common drifting volatility in large bayesian vars. *Journal of Business & Economic Statistics*, 34(3):375–390.

Chari, V. V., Dovis, A., and Kehoe, P. J. (2019). Rethinking optimal currency areas. *Journal of Monetary Economics*.

Chow, G. C. and Lin, A.-l. (1971). Best linear unbiased interpolation, distribution, and extrapolation of time series by related series. *The review of Economics and Statistics*, pages 372–375.

Christiano, L., Eichenbaum, M., and Rebelo, S. (2011). When is the government spending multiplier large? *Journal of Political Economy*, 119:78–121.

Christiano, L. J., Eichenbaum, M., and Evans, C. L. (1999). Monetary policy shocks: What have we learned and to what end? *Handbook of Macroeconomics*, 1:65–148.

Christiano, L. J., Eichenbaum, M. S., and Trabandt, M. (2016). Unemployment and business cycles. *Econometrica*, 84(4):1523–1569.

Chung, H. and Leeper, E. M. (2007). What has financed government debt? NBER Working Papers 13425.

Clarida, R., Galı, J., and Gertler, M. (1998). Monetary policy rules in practice: Some international evidence. *European Economic Review*, 42(6):1033–1067.

Clarida, R., Gali, J., and Gertler, M. (2000). Monetary policy rules and macroeconomic stability: evidence and some theory. *The Quarterly journal of economics*, 115(1):147–180.

Clarida, R., Galí, J., and Gertler, M. (2001). Optimal monetary policy in open versus closed economies: An integrated approach. *American Economic Review*.

Cloyne, J., Ferreira, C., Froemel, M., and Surico, P. (2018). Monetary policy, corporate finance and investment. *NBER Working Paper*, w25366.

Cogley, T. and Sargent, T. J. (2005). The conquest of US inflation: Learning and robustness to model uncertainty. *Review of Economic dynamics*, 8(2):528–563.

Corradin, S., Grimm, N., and Schwaab, B. (2021). Euro area sovereign bond risk premia during the covid-19 pandemic. *ECB Working Paper*.

Davig, T. and Leeper, E. M. (2011). Monetary-fiscal policy interactions and fiscal stimulus. *European Economic Review*, 55(2):211–227.

Diebold, F. X., Schorfheide, F., and Shin, M. (2017). Real-time forecast evaluation of DSGE models with stochastic volatility. *Journal of Econometrics*, 201:322–332.

Dincer, N. N. and Eichengreen, B. (2013). Central bank transparency and independence: updates and new measures.

Enderlein, H. (2005). One size fits none. *Central Banking*, 15(4):24–8.

Enderlein, H., Spiess, J., Guttenberg, L., and Vitorino, A. (2013). *Blueprint for a cyclical shock insurance in the euro area*. Notre Europe Brussels.

Engler, P., Ganelli, M. G., Tervala, J., and Voigts, S. (2014). *Fiscal devaluation in a monetary union*. Number 14-201. International Monetary Fund.

European Central Bank (2012). Verbatim of the remarks made by mario draghi.

European Commission (2020a). Employment-income loss and wage compensation during the pandemic: preliminary estimates. *Employment and Social Developments Quarterly Review – December 2020*.

European Commission (2020b). Labour market and wage developments in europe - annual review.

Faia, E. and Monacelli, T. (2008). Optimal monetary policy in a small open economy with home bias. *Journal of Money, credit and Banking*, 40(4):721–750.

Favero, C. and Giavazzi, F. (2012). Measuring tax multipliers: The narrative method in fiscal vars. *American Economic Journal: Economic Policy*, 4:69–94.

Fernandez-Villaverde, J., Guerron-Quintana, P., Rubio-Ramirez, J. F., and Uribe, M. (2011). Risk matters: The real effects of volatility shocks. *The American Economic Review*, 101:2530–2561.

Fischer, G. and Filauro, S. (2021). Income inequality in the eu: Trends and policy implications. *VOX, CEPR Policy Portal*.

Fisher, J. D. and Peters, R. (2010). Using stock returns to identify government spending shocks. *The Economic Journal*, 120:414–436.

Fleming, J. M. (1962). Domestic financial policies under fixed and under floating exchange rates. *Staff Papers*, 9(3):369–380.

Forni, L., Monteforte, L., and Sessa, L. (2009). The general equilibrium effects of fiscal policy: Estimates for the euro area. *Journal of Public Economics*, 93(3-4):559–585.

Franks, M. J. R., Barkbu, M. B. B., Blavy, M. R., Oman, W., and Schoelermann, H. (2018). *Economic Convergence in the Euro Area: Coming Together or Drifting Apart?* International Monetary Fund.

Fritsche, J. P. and Harms, P. C. (2019). 20 years of common european monetary policy: Reasons to celebrate. *DIW Weekly Report*, 9(20):179–187.

Galí, J. (2011). The return of the wage phillips curve. *Journal of the European Economic Association*, 9(3):436–461.

Galí, J. (2015). *Monetary policy, inflation, and the business cycle: an introduction to the new Keynesian framework and its applications*. Princeton University Press.

Gali, J., Lopez-Salido, J. D., and Valles, J. (2007). Understanding the effects of government spending on consumption. *Journal of the European Economic Association*, 5:227–270.

Gali, J. and Monacelli, T. (2005). Monetary policy and exchange rate volatility in a small open economy. *The Review of Economic Studies*, 72(3):707–734.

Galí, J., Smets, F., and Wouters, R. (2012). Unemployment in an estimated new keynesian model. *NBER macroeconomics annual*, 26(1):329–360.

Garriga, A. C. (2016). Central bank independence in the world: A new data set. *International Interactions*, 42(5):849–868.

Gertler, M. and Gilchrist, S. (1994). Monetary policy, business cycles, and the behavior of small manufacturing firms. *The Quarterly Journal of Economics*, 109(2):309–340.

Giavazzi, F. and Giovannini, A. (1987). Models of the ems: is europe a greater deutschmark area? In *Global macroeconomics: Policy conflict and cooperation*, pages 237–272. Springer.

Gopinath, G., Boz, E., Casas, C., Díez, F. J., Gourinchas, P.-O., and Plagborg-Møller, M. (2020). Dominant currency paradigm. *American Economic Review*, 110(3):677–719.

Gopinath, G., Kalemli-Özcan, S., Karabarbounis, L., and Villegas-Sanchez, C. (2017). Capital allocation and productivity in south europe. *The Quarterly Journal of Economics*, 132(4):1915–1967.

Guay, A. (2020). Identification of structural vector autoregressions through higher unconditional moments. *Journal of Econometrics*, forthcoming.

Gutierrez, G. and Piton, S. (2020). Revisiting the global decline of the (non-housing) labor share. *American Economic Review: Insights*, 2(3):321–38.

Hamilton, J. D. (2017). Why you should never use the hodrick-prescott filter. *Review of Economics and Statistics*, (0).

Hansen, B. E. (1992). The likelihood ratio test under nonstandard conditions: testing the markov switching model of gnp. *Journal of Applied Econometrics*, 7(S1).

Herwartz, H. and Lütkepohl, H. (2014). Structural Vector Autoregressions with Markov Switching: Combining Conventional with Statistical Identification of Shocks. *Journal of Econometrics*, 183(1):104–116.

Holm, M., Paul, P., and Tischbirek, A. (2021). The transmission of monetary policy under the microscope. *Journal of Political Economy (forthcoming)*.

Holston, K., Laubach, T., and Williams, J. C. (2017). Measuring the natural rate of interest: International trends and determinants. *Journal of International Economics*, 108:S59–S75.

Höpner, M. and Spielau, A. (2018). Better than the euro? the european monetary system (1979–1998). *New Political Economy*, 23(2):160–173.

Husted, L., Rogers, J., and Sun, B. (2020). Monetary policy uncertainty. *Journal of Monetary Economics*, 115:20–36.

Iacoviello, M. and Navarro, G. (2019). Foreign effects of higher us interest rates. *Journal of International Money and Finance*, 95:232–250.

IMF (2020). Fiscal monitor: Policies for the recovery.

Issing, O. (2001). The single monetary policy of the european central bank: One size fits all. *International finance*, 4(3):441–462.

Jarociński, M. and Karadi, P. (2020). Deconstructing monetary policy surprises — the role of information shocks. *American Economic Journal: Macroeconomics*, 12(2):1–43.

Jeenas, P. (2019). Firm balance sheet liquidity, monetary policy shocks, and investment dynamics. *Work in progress*.

Jordà, O. (2005). Estimation and inference of impulse responses by local projections. *American Economic Review*, 95(1):161–182.

Justiniano, A. and Primiceri, G. E. (2008). The time-varying volatility of macroeconomic fluctuations. *American Economic Review*, 98(3):604–41.

Kalemli-Özcan, c., Sørensen, B., Villegas-Sanchez, C., Volosovych, V., and Yeşiltaş, S. (2019). How to construct nationally representative firm leveldata from the orbis global database:new facts and aggregate implications. *Tinbergen Institute Discussion Paper*, TI 2015-110/IV.

Kaplan, G., Moll, B., and Violante, G. L. (2018). Monetary policy according to hank. *American Economic Review*, 108(3):697–743.

Kehrig, M. and Vincent, N. (2021). The micro-level anatomy of the labor share decline. *The Quarterly Journal of Economics*, 136(2):1031–1087.

Knight, F. H. (1921). *Risk, uncertainty and profit*, volume 31. Houghton Mifflin.

Kydland, F. E. and Prescott, E. C. (1977). Rules Rather than Discretion: The Inconsistency of Optimal Plans. *Journal of Political Economy*.

Lanne, M., Lütkepohl, H., and Maciejowska, K. (2010). Structural vector autoregressions with markov switching. *Journal of Economic Dynamics and Control*, 34(2):121–131.

Lanteri, A. (2018). The market for used capital: Endogenous irreversibility and reallocation over the business cycle. *American Economic Review*, 108(9):2383–2419.

Leeper, E. M., Traum, N., and Walker, T. B. (2017). Clearing up the fiscal multiplier morass. *American Economic Review*, 107:2409–54.

Leeper, E. M., Walker, T. B., and Yang, S. S. (2013). Fiscal foresight and information flows. *Econometrica*, 81(3):1115–1145.

Lewis, D. (2021). Identifying shocks via time-varying volatility. *The Review of Economic Studies*, page forthcoming.

Lucas, R. and Stokey, N. (1985). Money and interest in a cash-in-advance economy. Technical report, NBER Working Paper.

Lütkepohl, H., Meitz, M., Netšunajev, A., and Saikkonen, P. (2021). Testing identification via heteroskedasticity in structural vector autoregressive models. *The Econometrics Journal*, 24:1–22.

Lütkepohl, H. and Schlaak, T. (2018). Choosing between different time-varying volatility models for structural vector autoregressive analysis. *Oxford Bulletin of Economics and Statistics*, 80(4):715–735.

McDonald, R. and Siegel, D. (1986). The value of waiting to invest. *The Quarterly Journal of Economics*, 101(4):707–728.

Meese, R. A. and Rogoff, K. (1983). Do they fit out of sample? *Journal of international economics*, 14:3–24.

Miyamoto, W., Nguyen, T. L., and Sergeyev, D. (2018). Government spending multipliers under the zero lower bound: Evidence from japan. *American Economic Journal: Macroeconomics*, 10:247–77.

Monacelli, T., Perotti, R., and Trigari, A. (2010). Unemployment fiscal multipliers. *Journal of Monetary Economics*, 57(5):531–553.

Mundell, R. A. (1963). Capital mobility and stabilization policy under fixed and flexible exchange rates. *Canadian Journal of Economics and Political Science/Revue canadienne de economiques et science politique*, 29(4):475–485.

Nakamura, E. and Steinsson, J. (2014). Fiscal stimulus in a monetary union: Evidence from us regions. *American Economic Review*, 104:753–92.

Nekarda, C. J. and Ramey, V. A. (2020). The cyclical behavior of the price-cost markup. *Journal of Money, Credit and Banking*, 52(S2):319–353.

Nergiz Dincer, N. and Eichengreen, B. (2014). Central bank transparency and independence: Updates and new measures. *International Journal of Central Banking*.

Nussbaum, M. C. (2019). *The monarchy of fear: A philosopher looks at our political crisis*. Simon & Schuster.

Obstfeld, M., Shambaugh, J. C., and Taylor, A. M. (2005). The trilemma in history: tradeoffs among exchange rates, monetary policies, and capital mobility. *Review of Economics and Statistics*, 87(3):423–438.

Ottonello, P. and Winberry, T. (2020). Financial heterogeneity and the investment channel of monetary policy. *Econometrica*, 88(6):2473–2502.

Passacantando, F. (2013). Building an institutional framework for monetary stability: the case of italy (1979-1994). *PSL Quarterly Review*, 49(196).

Perotti, R. (2005). Estimating the Effects of Fiscal Policy in OECD Countries. CEPR Discussion Papers 4842.

Perotti, R. (2008). In search of the transmission mechanism of fiscal policy. In *NBER Macroeconomics Annual 2007, Volume 22*, NBER Chapters, pages 169–226. National Bureau of Economic Research, Inc.

Piketty, T. (2015). Putting distribution back at the center of economics: Reflections on capital in the twenty-first century. *Journal of Economic Perspectives*, 29(1):67–88.

Plagborg-Møller, M. and Wolf, C. (2021). Local projections and vars estimate the same impulse responses. *Econometrica*, 89(2):955–980.

Psaradakis, Z. and Spagnolo, N. (2006). Joint determination of the state dimension and autoregressive order for models with markov regime switching. *Journal of Time Series Analysis*, 27(5):753–766.

Quint, D. (2016). Is it really more dispersed? *International Economics and Economic Policy*, 13(4):593–621.

Rajan, R. and Zingales, L. (1998). Financial dependence and growth. *American Economic Review*, 88(3):559–586.

Ramey, V. A. (2011). Identifying government spending shocks: It's all in the timing*. *The Quarterly Journal of Economics*, 126:1–50.

Ramey, V. A. (2019). Ten years after the financial crisis: What have we learned from the renaissance in fiscal research? *Journal of Economic Perspectives*, 33(2):89–114.

Ramey, V. A. and Shapiro, M. D. (2001). Displaced capital: A study of aerospace plant closings. *Journal of political Economy*, 109(5):958–992.

Ramey, V. A. and Zubairy, S. (2018). Government spending multipliers in good times and in bad: evidence from us historical data. *Journal of Political Economy*, 126(2):850–901.

Ricco, G., Callegari, G., and Cimadomo, J. (2016). Signals from the government: Policy disagreement and the transmission of fiscal shocks. *Journal of Monetary Economics*, 82:107–118.

Rigobon, R. (2003). Identification through heteroskedasticity. *Review of Economics and Statistics*, 85(4):777–792.

Sapir, A., Wolff, G. B., et al. (2015). Euro-area governance: what to reform and how to do it. Technical report.

Sharpe, S. A. and Suarez, G. (2015). Interest rate rules vs. money growth rules a welfare comparison in a cash-in-advance economy. *Management Science*, 67(2):720–741.

Sims, C. A. and Zha, T. (2006). Were there regime switches in us monetary policy? *American Economic Review*, 96(1):54–81.

Stock, J. H. and Watson, M. W. (2002). Has the business cycle changed and why? *NBER macroeconomics annual*, 17:159–218.

Stock, J. H. and Watson, M. W. (2018). Identification and estimation of dynamic causal effects in macroeconomics using external instruments. *The Economic Journal*, 128(610):917–948.

Sturm, J.-E. and Wollmershäuser, T. (2008). The stress of having a single monetary policy in europe.

Svensson, L. E. (2010). Inflation targeting. In *Handbook of monetary economics*, volume 3, pages 1237–1302. Elsevier.

Swanson, E. (2020). Implications of labor market frictions for risk aversion and risk premia. *American Economic Journal: Macroeconomics*, 12(2):194–240.

Swanson, E. T. (2006). Have increases in federal reserve transparency improved private sector interest rate forecasts? *Journal of Money, Credit and Banking*, pages 791–819.

Tagkalakis, A. (2008). The effects of fiscal policy on consumption in recessions and expansions. *Journal of Public economics*, 92(5-6):1486–1508.

Taylor, J. B. (1993). Discretion versus policy rules in practice. In *Carnegie-Rochester conference series on public policy*, volume 39, pages 195–214. Elsevier.

Uhlig, H. (2005). What are the effects of monetary policy on output? results from an agnostic identification procedure. *Journal of Monetary Economics*, 52(2):381–419.

Wolff, J. and Jerow, S. (2020). Fiscal policy and uncertainty. *Available at SSRN 3602980*.

Woodford, M. (2003). Interest and prices: Foundations of a theory of monetary policy, princeton university.

Wyplosz, C. (2016). The six flaws of the eurozone. *Economic Policy*, 31(87):559–606.

Auflistung der enthaltenen Papiere der Dissertation

Government Spending Multipliers in (Un)certain Times
Gemeinsam mit Mathias Klein und Malte Rieth
Revise and Resubmit beim Journal of Public Economics
Erschienen als DIW Discussion Papers Nummer 1901
https://www.diw.de/de/diw_01.c.799546.de/publikationen/diskussionspapiere/2020_1901/government_spending_multipliers_in__un_certain_times.html
Anteil am Papier: 1/3

Better off without the Euro? A Structural VAR Assessment of European Monetary Policy
Gemeinsam mit Patrick Harms
Erschienen als DIW Discussion Papers Nummer 1907
https://www.diw.de/de/diw_01.c.801198.de/publikationen/diskussionspapiere/2020_1907/better_off_without_the_euro__a_structural_var_assessment_of_european_monetary_policy.html
Anteil am Papier: 1/2

Zooming in on Monetary Policy - The Labor Share and Production Dynamics of Two Million Firms
Gemeinsam mit Lea Steininger
Workingpaperversion veröffentlicht unter:
https://papers.ssrn.com/sol3/papers.cfm?abstract_id=3890084
Erscheinend als DIW Wochenbericht und als DIW Discussion Paper
Anteil am Papier: 2/3

Erklärung gemäß Promotionsordnung §9 Absatz 1 sowie §12 Absatz 4

Zum Erstellen dieser Dissertation habe ich folgende Hilfsmittel und Hilfen genutzt: R, Rstudio, Stata, Latex, MATLAB, Mathematica, Overleaf.com, mapchart.net, Texstudio, sowie MS Office.

Ich bezeuge durch meine Unterschrift, dass meine Angaben über die bei der Abfassung meiner Dissertation benutzten Hilfsmittel, über die mir zuteil gewordene Hilfe sowie über frühere Begutachtungen meiner Dissertation in jeder Hinsicht der Wahrheit entsprechen.

Jan Philipp Fritsche,
Berlin, den 30.07.2021