Essays in applied structural macro-econometric modelling

by

Naveed Javed

Thesis submitted in fulfillment
of the requirements for the degree of

Doctor of Philosophy

September, 2023

Department of Economics

School of Economics and Public Policy

The University of Adelaide

Australia

Table of Contents

List of Figures **vi**

List of Tables **ix**

Abstract **x**

Declaration **xi**

Acknowledgements **xii**

Dedication **xiii**

Introduction **1**

References 6

1 Chapter 1: Dornbusch's overshooting and the systematic component of monetary policy in SOE-SVARs **8**

1.1 Introduction 10

1.2 Econometric strategy 15

1.2.1 Data 15

1.2.2 Model 15

1.2.3 Identification 16

1.3 Results 23

1.3.1 IRFs to a SOE contractionary monetary shock 23

1.3.2 IRFs to a US contractionary monetary shock 25

1.3.3 Importance of Restriction 4 (leaning against the RER) 27

1.3.4 Role of Restriction 2 (leaning against the credit spread) 29

1.3.5 Deviations from UIP after a monetary policy shock 30

1.3.6 Forecast error variance decomposition of the RER 32
1.4 Robustness Checks 33
1.4.1 Excluding ZLB episodes 33
1.4.2 Including the SOE credit spread 35
1.5 Conclusion 38

A Appendix to Chapter 1 **40**
A.1 Data Sources 40
A.2 Baseline 45
A.2.1 IRFs to a one standard deviation SOE monetary policy shocks . 45
A.2.2 IRFs to a one standard deviation US monetary policy shocks . . 48
A.2.3 Responses of US variables to US shocks 51
A.2.4 IRFs to a one standard deviation SOE monetary policy shocks with/ out $\psi_e > 0$ 52
A.2.5 IRFs to a one standard deviation SOE monetary policy shocks with/out $\psi_{cs^*} < 0$ 55
A.3 Robustness check # 1 58
A.3.1 IRFs to a one standard deviation SOE monetary policy shocks with/out $\psi_e > 0$ 58
A.4 Baseline 59
A.4.1 Contemporaneous coefficients in SOE monetary policy equations. 59
A.4.2 Contemporaneous coefficients in US monetary policy equations. 60
A.5 Robustness check # 1 61
A.5.1 Contemporaneous coefficients in SOE monetary policy equations. 61
A.5.2 Contemporaneous coefficients in US monetary policy equations. 62
A.6 Robustness check # 2 63
A.6.1 Contemporaneous coefficients in SOE monetary policy equations. 63
A.6.2 Contemporaneous coefficients in US monetary policy equations. 64

References **68**

2 Chapter 2: What Are the Effects of Commodity Supply and Demand Shocks on Australian Output and Trade Balance Fluctuations? **70**
2.1 Introduction 72
2.2 Data Description 76

2.3 Empirical Model 77
2.3.1 Model identification 79
2.4 Baseline Model Results 83
2.4.1 Shock to Commodity Prices 83
2.5 Extended-SVAR Model 86
2.5.1 Shock to Chinese Steel Production 89
2.5.2 Shock to Commodity Prices 91
2.6 Robustness Check: Extended-SVAR Model with Log-Linear Detrending 92
2.7 Variance Decomposition Analysis 95
2.8 Conclusion 97

B Appendix to Chapter 2 **99**
B.1 Data Sources 99
B.2 Plots of variables 101
B.2.1 Plots of variables for baseline model 101
B.2.2 Plots of variables for extended model 102
B.2.3 Plots of variables for robustness check 104

References **107**

3 Chapter 3: High-Frequency Identification of US monetary shocks using Proxy-SVAR models - A Survey of the Literature **109**
3.1 Introduction 111
3.2 Survey of the Literature 112
3.2.1 Proxy-SVAR literature 112
3.2.2 Literature on US monetary policy instruments 114
3.3 Application: Comparison between a standard and information-robust proxy 126
3.3.1 Data and Model 126
3.3.2 Results 128
3.4 Conclusion and future research 130

C Appendix to Chapter 3 **132**
C.1 Plots of proxies 132

References **135**

Conclusion **136**

References 141

List of Figures

Chapter 1 **8**

1.1 IRFs of SOE variables to a one standard deviation SOE contractionary monetary shock identified using block exogeneity and Restrictions 1 to 4 24

1.2 IRFs of SOE variables to a one standard deviation US contractionary monetary shock identified using block exogeneity and Restrictions 1 to 4 26

1.3 IRFs of SOE variables to a SOE contractionary monetary shock 28

1.4 IRFs of the RER to a US contractionary monetary shock 29

1.5 Deviations from UIP conditional on US and SOE monetary policy shocks 31

1.6 Forecast error variance decomposition of the SOE/US RER 32

1.7 IRFs of SOE variables to a one standard deviation SOE contractionary monetary shock using the shorter estimation sample period 1992:Q1 to 2008:Q3 34

1.8 Deviations from UIP when using the shorter estimation sample period 1992:Q1 to 2008:Q3 35

1.9 IRFs of SOE variables to a SOE contractionary monetary shock identified using block exogeneity and Restrictions 1 to 5 36

1.10 Deviations from UIP for Robustness check 2 (SOE credit spread) . . . 37

A.1 Australia - IRFs to a one standard deviation contractionary monetary policy shock 45

A.2 Canada - IRFs to a one standard deviation contractionary monetary policy shock 45

A.3 New Zealand - IRFs to a one standard deviation contractionary monetary policy shock 46

A.4 Norway - IRFs to a one standard deviation contractionary monetary policy shock 46

A.5 Sweden - IRFs to a one standard deviation contractionary monetary policy shock 47

A.6 UK - IRFs to a one standard deviation contractionary monetary policy shock . . . 47
A.7 Australia - IRFs to a one standard deviation US contractionary monetary policy shock . . . 48
A.8 Canada - IRFs to a one standard deviation US contractionary monetary policy shock . . . 48
A.9 New Zealand - IRFs to a one standard deviation US contractionary monetary policy shock . . . 49
A.10 Norway - IRFs to a one standard deviation US contractionary monetary policy shock . . . 49
A.11 Sweden - IRFs to a one standard deviation US contractionary monetary policy shock . . . 50
A.12 UK - IRFs to a one standard deviation US contractionary monetary policy shock . . . 50
A.13 IRFs to a one standard deviation US contractionary monetary policy shock for six SOEs . . . 51
A.14 Australia - IRFs to a one standard deviation contractionary monetary policy shock with/ out $\psi_e > 0$ . . . 52
A.15 Canada - IRFs to a one standard deviation contractionary monetary policy shock with/ out $\psi_e > 0$ . . . 52
A.16 New Zealand - IRFs to a one standard deviation contractionary monetary policy shock with/ out $\psi_e > 0$ . . . 53
A.17 Norway - IRFs to a one standard deviation contractionary monetary policy shock with/ out $\psi_e > 0$ . . . 53
A.18 Sweden - IRFs to a one standard deviation contractionary monetary policy shock with/ out $\psi_e > 0$ . . . 54
A.19 UK - IRFs to a one standard deviation contractionary monetary policy shock with/ out $\psi_e > 0$ . . . 54
A.20 Australia - IRFs to a one standard deviation contractionary monetary policy shock with/out $\psi_{cs^*} < 0$ . . . 55
A.21 Canada - IRFs to a one standard deviation contractionary monetary policy shock with/out $\psi_{cs^*} < 0$ . . . 55
A.22 New Zealand - IRFs to a one standard deviation contractionary monetary policy shock with/out $\psi_{cs^*} < 0$ . . . 56
A.23 Norway - IRFs to a one standard deviation contractionary monetary policy shock with/out $\psi_{cs^*} < 0$ . . . 56
A.24 Sweden - IRFs to a one standard deviation contractionary monetary policy shock with/out $\psi_{cs^*} < 0$ . . . 57

A.25 UK - IRFs to a one standard deviation contractionary monetary policy shock with/out $\psi_{cs^*} < 0$ 57
A.26 Robustness check # 1 - IRFs to a one standard deviation contractionary monetary policy shocks for six SOEs $\psi_e > 0$ 58

Chapter 2 **70**

2.1 Commodity Price Index 72
2.2 Baseline SVAR Model: IRFs to a one standard deviation Innovation in commodity prices 84
2.3 Extended SVAR Model: IRFs to a one standard deviation Innovation in Chinese steel production 90
2.4 Extended SVAR Model: IRFs to a one standard deviation Innovation in Commodity prices 91
2.5 Robustness check: IRFs to a one standard deviation Innovation in Chinese steel production 93
2.6 Robustness check: IRFs to a one standard deviation Innovation in Commodity prices 94
2.7 Forecast Error Variance Decomposition 95
2.8 Forecast Error Variance Decomposition to Chinese steel production shocks for extended model and robustness check 96

Chapter 3 **109**

3.1 IRFs to a one standard deviation contractionary monetary policy shock 128
3.2 IRFs to a one standard deviation contractionary monetary policy shock 130

C.1 Plots of proxies 132

List of Tables

Chapter 1 **8**

A.1 Contemporaneous coefficients in SOE monetary policy equations using full sample period from 1992:Q1 to 2019:Q4 59

A.2 Contemporaneous coefficients in US monetary policy equations using full sample period from 1992:Q1 to 2019:Q4 60

A.3 Contemporaneous coefficients in SOE monetary policy equations using shorter sample period from 1992:Q1 to 2008:Q3 61

A.4 Contemporaneous coefficients in US monetary policy equations using shorter sample period from 1992:Q1 to 2008:Q3 62

A.5 Contemporaneous coefficients in SOE monetary policy equations using sample period from 2000:Q1 to 2019:Q4 63

A.6 Contemporaneous coefficients in US monetary policy equations using sample period from 2000:Q1 to 2019:Q4 64

Chapter 2 **70**

2.1 Summary of the responses of Output and Trade Balance 85

Abstract

This thesis is a compilation of three diverse but self-contained chapters on the impact of monetary policy shocks on six Small-Open-Economies (SOEs) and the US and the impact of commodity demand and supply shocks on the Australian economy.

In the first chapter, we estimate Small-Open-Economy-Structural Vector Autoregression (SOE-SVAR) models for Australia, Canada, New Zealand, Norway, Sweden and the United Kingdom to measure the effects of SOE and US monetary policy shocks on bilateral SOE/US exchange rates. We find that a contractionary SOE (US) monetary shock triggers an immediate appreciation (depreciation) of the exchange rate followed by a reversion, in line with Dornbusch's overshooting and uncovered interest rate parity. SOE monetary impulses account for a greater portion of the short-run volatility of the exchange rate than US monetary shocks.

In the second chapter, I estimate SVAR models to measure the effects of commodity supply and commodity demand shocks on Australia's output and trade balance. I find that commodity supply and demand shocks emerge as a relatively minor and negligible sources of business cycle fluctuations in output and trade balance. I further find that output expands (contracts) in response to commodity demand (commodity supply) shocks. Interestingly, for both commodity supply and demand shocks the trade balance worsens substantially.

In the third chapter, I review the literature on proxy-SVAR models and document the evolution of various types of proxies for US monetary policy shocks. The chapter also contains an application of proxy-SVAR models using high-frequency monetary policy instruments for the US. I compare the two most recent US monetary instruments and find that information-robust monetary instrument produces different results compared to the instrument that does not take into account the information content of monetary policy announcements.

Declaration

I certify that this work contains no material which has been accepted for the award of any other degree or diploma in my name in any university or other tertiary institution and, to the best of my knowledge and belief, contains no material previously published or written by another person, except where due reference has been made in the text. In addition, I certify that no part of this work will, in the future, be used in a submission in my name for any other degree or diploma in any university or other tertiary institution without the prior approval of the University of Adelaide and where applicable, any partner institution responsible for the joint award of this degree.

I acknowledge that copyright of published works contained within this thesis resides with the copyright holder(s) of those works.

I also give permission for the digital version of my thesis to be made available on the web, via the University's digital research repository, the Library Search and also through web search engines, unless permission has been granted by the University to restrict access for a period of time.

I acknowledge the support I have received for my research through the provision of an Australian Government Research Training Program Scholarship and Baillieu Supplementary Scholarship.

Author's signatures: Date: 25 September 2023

Acknowledgements

First and foremost, I would like to thank my PhD supervisor, Prof. Nicolas Groshenny. I personally consider Nicolas Groshenny more than a supervior, he is my mentor who has left an indelible impression on my personality. I know him since 2018 when I took his course of "International Finance - IV" before he supervised me in my Masters thesis. In hindsight, I can safely say that he judged my potential to do research earlier than I, myself realized. The best part about Nicolas Groshenny is that he always remains patient, which I believe is the key to success in a long and arduous journey like a PhD. I can't thank him enough for his support and will always be indebted to him.

I would also like to thank Dr. Stephanie McWhinnie for her support on every occasion, to Dr. Raul Barreto and all the lecturers at the School of Economics and Public Policy. I would like to thank all the participants of all the conferences where I presented my work and to the participants of PhD workshop for their invaluable insights.

Last but not the least, I would like to thank my long-time friend, Zeeshan Javaid. He was the first one who encouraged me to study Economics and adopt it as a profession. I would not have been submitting this thesis today, had he not entered my life.

Dedication

This thesis is for my late father, Javed Butt, whom I miss a lot, for my beloved mother, Nasreen Javed, for my lovely wife, Atufa Dar, for the most loving daughter, Aayana Naveed, and for my son and best buddy, Abdul Hadi.

Introduction

Since the revolutionary contribution of Sims (1980), vector autoregressions (VARs) are in use to model the effects of structural shocks. His VAR methodology made it easier to impose identification assumptions, study the impulse response functions, and to do innovation accounting using forecast error decomposition (Ramey, 2016). There is a large literature that employs the VAR methodology to measure the effects of various types of macroeconomic shocks using a variety of identification schemes[1]. In this thesis, I compile three distinct research ideas for three chapters. However, the research ideas are closely related based on the following features: First, I use Structural Vector Autoregression (SVAR) models where I apply most recent and novel identification schemes. Second, in all three chapters the models are estimated using state-of-the-art Bayesian estimation techniques and finally, I focus on topical policy relevant questions relating to small-open-economy and US monetary shocks and Australian commodity shocks.

In the first chapter, we address the long-debated question in international finance: How does exchange rate respond to monetary disturbances in small-open-economies? Specially, we re-visit the Dornbusch's overshooting hypothesis (Dornbusch, 1976) using Bayesian SVAR models and show that Dornbusch's overshooting still holds true. In the second chapter, I address a policy relevant question for Australia: what is the impact of commodity supply and demand shocks on the Australian macroeconomy? I find that output expands (contracts) in response to commodity demand (commodity supply) shocks while the trade balance deteriorates significantly in response to both the shocks. In the third chapter, I focus on reviewing the literature on proxy-SVAR models, also known as SVAR-IV (Instrument Variable) models. Further, I document various types of US monetary policy instruments. I also provide an application that employs Bayesian Proxy-SVAR (BP-SVAR) models where I compare the results using a US monetary policy instrument that takes into account the information effect to an instrument that does not. I find that contraction in US output and a fall in prices are

[1] See Ramey (2016) for a review of literature on VAR modelling and identification assumption.

more pronounced when I use the information-robust proxy as compared to when I use the standard proxy.

Our first chapter tests the Dornbusch's overshooting hypothesis using SVAR models for six small-open-economies (SOEs): Australia, Canada, New Zealand, Norway, Sweden and the UK. Dornbusch's overshooting hypothesis tests the exchange rate puzzle and delayed overshooting puzzle. An exchange rate puzzle is defined as: when the exchange rate depreciates on impact rather than appreciating in response to a contractionary monetary shock. A delayed overshooting puzzle is defined as: when the exchange rate continues to appreciate rather than depreciating in response to a contractionary monetary shock. Our SVAR models for six SOEs are identified using the identification scheme recently proposed by Arias et al. (2019). A novel feature of the identification scheme developed by Arias et al. (2019) is that it applies a combination of exclusion and sign-restrictions directly on the structural parameters. We contribute towards the literature by building on this novel identification scheme by imposing a block exogeneity structure, following Cushman and Zha (1997). The block exogeneity identification scheme is essential to the SOE settings. We impose the exogeneity structure through a combination of exclusion restrictions and tight priors on the contemporaneous and lagged parameters, respectively. We further contribute towards the literature through an identification strategy where we simultaneously identify the systematic component of US and SOE monetary policies. Moreover, we achieve an agnostic identification that allows for a contemporaneous interaction between the exchange rate and policy rate. This identification approach is essential in addressing the puzzling responses that we observe in the empirical literature.

We find no evidence of exchange rate puzzle and delayed overshooting puzzle and thus, corroborate the Dornbusch's overshooting hypothesis. Our findings show an instantaneous appreciation of the exchange rate followed by a gradual depreciation in response to an SOE contractionary monetary shock. On the other hand, for all the six SOEs, a US monetary tightening results in an appreciation of the US dollar followed by a gradual depreciation. We contribute towards the SOE literature on exchange rates by pin-pointing that delayed overshooting is an artefact of incorrect identifying restrictions. Specifically, we conduct a sensitivity analysis which shows that the SOE central banks should cut the interest rate in response to a real appreciation of the exchange rate. We conclude that the response of SOE central banks to the exchange rate fluctuations is key to address the delayed overshooting puzzle. Moreover, we find little evidence of the forward discount puzzle. The responses of exchange rates to both the SOE and US monetary shocks are broadly consistent with UIP. Finally, the six SOEs and US monetary shocks explain about 20 and 10 percents of the short-run exchange rate volatility, respectively.

In the second chapter, I employ the SVAR methodology in Bayesian settings in order to quantify the impact of commodity supply and commodity demand shocks on Australian macroeconomy. My baseline model's specification draws from Drechsel and Tenreyro (2018), Schmitt-Grohé and Uribe (2018) and Di Pace et al. (2020). I quantify the impact of a (positive) commodity price shock that acts as a proxy for (negative) commodity supply shock. My SVAR model estimates the impact of exogenous disturbances in commodity prices on Australian output and the trade balance. I achieve identification through the Cholesky decomposition of the contemporaneous matrix with commodity prices ordered first. This ensures that commodity prices acts as an exogenous variable in the model. The exclusion restrictions are imposed using the techniques developed by Arias et al. (2019). For the lagged matrices, I achieve exogeneity of the commodity prices through the tight priors. The model is estimated using the Bayesian techniques of Arias et al. (2019). I find that a positive commodity price shock results in a contraction of output and a significant and persistent deterioration of the trade balance. Moreover, I find a negligible contribution of commodity price shocks to output fluctuations while these disturbances account for around 25% of the variance of the trade balance.

I build on my baseline model by introducing a *"foreign block"*. The *foreign block* helps in quantifying the impact of Chinese steel production on the Australian economy. In order to implement block exogeneity structure, I continue to impose Cholesky identification scheme on the contemporaneous matrix. The exogeneity structure on the lagged matrices is implemented through tight priors on the lagged parameters. I relax one restriction in the lagged matrices which implies that the exchange rate impacts the iron ore exports through the lags. I estimate my extended SVAR model to quantify the impact of (positive) commodity prices and (positive) Chinese steel production shocks on the Australian output and trade balance. The (positive) commodity prices and (positive) Chinese steel production act as proxies for (negative) commodity supply and (positive) commodity demand shocks, respectively. I find that Australian output expands in response to a commodity demand shock while it contracts to a commodity supply shock. The trade balance deteriorates significantly in response to both the commodity demand and supply shocks. I contribute towards the Australian macroeconomic literature by estimating the impact of Chinese demand for iron ore and coal on output and trade balance. My findings are useful for the policy-makers and contribute towards economic growth and trade related policies.

In the first part of my third chapter, I explore the literature on emerging tools that are employed to quantify the impact of monetary and fiscal policies. These tools are known in the literature as proxy-SVAR/SVAR-IV models. These models are gaining importance because of their point-estimation identification strategy. Proxy-SVAR model's identification approach follows from the exogeneity and relevance conditions common in the instrument variable literature. Essentially, the implication is that the instrument variable is correlated with the shock of interest (*relevance condition*) and uncorrelated with the rest of the shocks in the system (*exogeneity condition*). Another interesting feature of the proxy-SVAR models is the identification of multiple shocks with one or more instruments. The literature provides many interesting applications of using proxy-SVAR models that I discuss in this chapter. I also provide a review of various types of US monetary policy instruments that are used for identification in the proxy-SVAR literature.

In the second part of my third chapter, I use the proxy-SVAR model in Bayesian settings recently developed by Arias et al. (2021). These models are commonly known in the literature as Bayesian Proxy-SVAR (BP-SVAR) models. Further my BP-SVAR model uses two recently constructed high-frequency monetary policy instruments for the US measured within a 30-minutes (10 minutes before and 20 minutes after) tight-window around the Federal Open Market Committee (FOMC) announcements. The instrument constructed by Jarociński and Karadi (2020) purges the monetary policy announcements from the information effects. This is known in the literature as information-robust proxies. On the other hand, the instrument constructed by Gertler and Karadi (2015) does not take into account the information channel in the monetary announcement. My application highlights the differences between the results obtained from a standard US monetary instrument (Gertler and Karadi, 2015) and an information-robust instrument (Jarociński and Karadi, 2020). My findings suggest that the contraction in US output and a fall in prices is more pronounced when I identify the BP-SVAR model through the information-robust proxy of Jarociński and Karadi (2020) as compared to when I use the standard proxy of Gertler and Karadi (2015). These findings emphasis that results obtained from an information-robust proxy for the US monetary shocks are different from the ones obtained using the standard proxy.

In future, I intend to build on my work of chapter 3 by extending it to the SOE settings. Particularly, I will quantify the impact of US monetary shocks on small open macroeconomies. I intend to develop a hybrid identification scheme that would combine proxy-restrictions with short-run zero restrictions and sign-restrictions. I intend to use an instrument variable to identify the US monetary policy shocks in a BP-SVAR model. Further, a block exogeneity structure will be imposed using the short-run zero restrictions on structural parameters while the SOE monetary shock is identified by applying sign-restrictions directly on the structural parameters of the systematic component of SOE monetary policy rule. This work will contribute towards SOE monetary policy literature by providing insight into the spill-over effects of US monetary policy on the SOEs using a novel, hybrid identification scheme.

References

Arias, J. E., Caldara, D., and Rubio-Ramirez, J. F. (2019). The systematic component of monetary policy in SVARs: An agnostic identification procedure. *Journal of Monetary Economics*, 101:1–13.

Arias, J. E., Rubio-Ramírez, J. F., and Waggoner, D. F. (2021). Inference in Bayesian Proxy-SVARs. *Journal of Econometrics*, 225(1):88–106.

Bauer, M. D. and Swanson, E. T. (2023). A reassessment of monetary policy surprises and high-frequency identification. *NBER Macroeconomics Annual*, 37(1):87–155.

Beaudry, P. and Saito, M. (1998). Estimating the effects of monetary shocks: An evaluation of different approaches. *Journal of Monetary Economics*, 42(2):241–260.

Cushman, D. O. and Zha, T. (1997). Identifying monetary policy in a small open economy under flexible exchange rates. *Journal of Monetary economics*, 39(3):433–448.

Di Pace, F., Juvenal, L., and Petrella, I. (2020). Terms-of-trade shocks are not all alike.

Dornbusch, R. (1976). Expectations and exchange rate dynamics. *Journal of Political Economy*, 84(6):1161–1176.

Drechsel, T. and Tenreyro, S. (2018). Commodity booms and busts in emerging economies. *Journal of International Economics*, 112:200–218.

Gertler, M. and Karadi, P. (2015). Monetary policy surprises, credit costs, and economic activity. *American Economic Journal: Macroeconomics*, 7(1):44–76.

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

Ramey, V. A. (2016). Macroeconomic shocks and their propagation. *Handbook of macroeconomics*, 2:71–162.

Schmitt-Grohé, S. and Uribe, M. (2018). How important are terms-of-trade shocks? *International Economic Review*, 59(1):85–111.

Sims, C. A. (1980). Macroeconomics and reality. *Econometrica: journal of the Econometric Society*, pages 1–48.

Statement of Authorship

Title of Paper	Dornbusch's overshooting and the systematic component of monetary policy in SOE-SVARs
Publication Status	☐ Published ☐ Accepted for Publication ☐ Submitted for Publication ☑ Unpublished and Unsubmitted w ork w ritten in manuscript style
Publication Details	

Principal Author

Name of Principal Author (Candidate)	Naveed Javed		
Contribution to the Paper	Stating the research question, determining the research methodology, construction of database, coding of estimation, production of all results, interpretation and framing of results, writing.		
Overall percentage (%)	66%		
Certification:	This paper reports on original research I conducted during the period of my Higher Degree by Research candidature and is not subject to any obligations or contractual agreements with a third party that would constrain its inclusion in this thesis. I am the primary author of this paper.		
Signature		Date	12.09.2023

Co-Author Contributions

By signing the Statement of Authorship, each author certifies that:

i. the candidate's stated contribution to the publication is accurate (as detailed above);

ii. permission is granted for the candidate in include the publication in the thesis; and

iii. the sum of all co-author contributions is equal to 100% less the candidate's stated contribution.

Name of Co-Author	Nicolas Groshenny		
Contribution to the Paper	Stating the research question, determining the research methodology, interpretation and framing of results, writing.		
Signature		Date	12.09.2023

Name of Co-Author			
Contribution to the Paper			
Signature		Date	

Please cut and paste additional co-author panels here as required.

Chapter 1

Dornbusch's overshooting and the systematic component of monetary policy in SOE-SVARs*

Nicolas Groshenny[†] Naveed Javed[‡]

* We would like to thank Cristina Badarau, Kenza Benhima, Hilde Bjørnland, Efrem Castelnuovo, Federico Di Pace, Laurent Ferrara, Renée Fry-McKibbin, Francesco Furlanetto, James Graham, Eleonora Granziera, David Guerreiro, Michel Guillard, Qazi Haque, Eleni Iliopulos, Frédéric Karamé, Mariano Kulish, François Langot, Junior Maih, Jean-Baptiste Michau, James Morley, Lise Patureau, Jérémy Pépy, Matthew Read, Tim Robinson, Karsten Staehr, Jacek Suda, Olav Syrstad, Grégory Verdugo as well as conference and seminar participants at the University of Evry, the University of Paris VIII, CEF 2022 in Dallas, WAMS 2022 in Sydney, the 30 years anniversary GAINS conference in Le Mans, AFSE 2023 in Paris, CEF 2023 in Nice, ESAM 2023 in Sydney.

† Le Mans Université, GAINS-TEPP, CAMA. nicolas.groshenny@univ-lemans.fr.

‡ School of Economics, University of Adelaide, naveed.javed@adelaide.edu.au

Abstract

We estimate Small-Open-Economy SVAR models for Australia, Canada, New Zealand, Norway, Sweden and the United Kingdom to measure the effects of SOE and US monetary policy shocks on bilateral SOE/US exchange rates. Our identification strategy features block exogeneity and sign-restrictions imposed on the coefficients of the SOE and US monetary policy rules. Our approach leaves the response of the exchange rate to domestic and foreign monetary shocks unrestricted, while allowing for instantaneous interactions between the SOE policy rate and the exchange rate. We find that a contractionary SOE (US) monetary shock triggers an immediate appreciation (depreciation) followed by a reversion, in line with Dornbusch's overshooting and uncovered interest rate parity. SOE monetary impulses account for a greater portion of the short-run volatility of the exchange rate than US monetary shocks.

Keywords: Structural vector autoregressions, Small open economies, Monetary policy rules, Exchange rates, Spillovers of US monetary policy, Uncovered interest rate parity, Block exogeneity.

JEL codes: **C32**, **E52, F31, F41**.

1.1 Introduction

How does the exchange rate react to monetary policy shocks? This classic question in international finance is of practical importance to monetary policymakers in small open economies. Even though, this question largely remains unsettled. According to the benchmark theoretical result of Dornbusch (1976), a surprise monetary tightening causes the exchange rate to overshoot on impact, displaying an instantaneous appreciation followed by a gradual depreciation. On the other hand, a vast literature estimating Structural Vector Autoregressions often finds evidence of a gradual appreciation, typically lasting for more than a year, in response to a monetary shock. Such hump-shaped empirical responses are referred to as the delayed overshooting puzzle.[1] Yet, Faust and Rogers (2003) and Bjørnland (2009) argue that the delayed overshooting puzzle may be an artifact caused by incorrect identifying restrictions. In particular, identification schemes that do not allow for simultaneous interactions between money market variables and the exchange rate tend to produce delayed overshooting.

In this chapter, we estimate Bayesian SVAR models for six advanced small-open-economies with floating exchange rates (Australia, Canada, New Zealand, Norway, Sweden and the United Kingdom) to measure the effects of monetary shocks on bilateral SOE/US real exchange rates. We identify simultaneously SOE and US monetary impulses and compare the conditional dynamics of the exchange rate in response to these two disturbances. Our identification scheme hinges on two ingredients: i) Block exogeneity, a cornerstone of Small-Open-Economy SVARs (Cushman and Zha, 1997); ii) Sign restrictions imposed on the coefficients of the SOE and US monetary policy rules. Our approach has two main advantages: First, it is agnostic, in the sense that it leaves the response of the exchange rate to domestic and foreign monetary shocks unrestricted. Second, it allows for instantaneous interactions between the exchange rate and the SOE and US policy rates. These two features make our econometric strategy well suited to investigate the robustness of the delayed overshooting puzzle.

[1] See Eichenbaum and Evans (1995), Cushman and Zha (1997), Kim and Roubini (2000), Scholl and Uhlig (2008) and Kim et al. (2017) among others.

Our identification scheme builds on Arias et al. (2019). They identify the systematic component of monetary policy in the United States by imposing sign and exclusion restrictions on the coefficients of the Federal Reserve's interest-rate rule. We adapt their methodology to a SOE context in two ways. First, we impose a block-exogenous structure on the SVAR model, meaning that for each SOE, the variables in the model are classified into two blocks: a US (i.e., foreign) block and a SOE (i.e., domestic) block. The US block influences the SOE block (both contemporaneously and over time), whereas the SOE block has no effect on the US block. Second, we identify simultaneously the policy rule of the Federal Reserve and of the SOE central bank.To characterize the systematic component of US monetary policy, we require that the response of the federal funds rate be positive to US output and US inflation, and negative to the Baa credit spread (Caldara and Herbst, 2019).[2] For the SOE central bank, in line with Taylor (2001), we assume that it follows an augmented Taylor-type rule that reacts positively to output, inflation and the real SOE/USD exchange rate.[3] Hence, we require that the SOE's central bank never raises its policy rate in response to a real appreciation. The sign-restriction on the exchange-rate response embodies policymakers' rule of thumb that the central bank should lean against fluctuations in the real exchange rate, based on the wisdom that a real appreciation is an opportunity to ease monetary conditions (Obstfeld and Rogoff, 1995; Taylor, 2001).[4] This sign-restriction is supported by estimates of Taylor-type rules in DSGE models (Lubik and Schorfheide, 2007; Kam et al., 2009; Justiniano and Preston, 2010), and by findings from the SOE-SVAR literature (Bjørnland, 2009; Bjørnland and Halvorsen, 2014).

A distinctive feature of our study is to jointly identify SOE and US monetary policy shocks.[5] There is an obvious asymmetry between the Fed and the six SOE central

[2] Curdia and Woodford (2010, 2016) present a normative analysis justifying a negative systematic response of monetary policy to a tightening of credit conditions.

[3] Calvo and Reinhart (2002), Reinhart and Rogoff (2004), Obstfeld (2013) and Ilzetzki et al. (2019) provide evidence that central banks react to movements in the bilateral dollar exchange rate. Gopinath et al. (2020) and Gourinchas et al. (2019) document the central role played by the dollar in international trade and in the international monetary and financial system.

[4] Egorov and Mukhin (2023) show that, under dollar pricing (i.e., when prices are invoiced and sticky in dollars), it becomes desirable for non-US central banks to stabilize the dollar exchange rate.

[5] Several papers focus on US monetary shocks (Eichenbaum and Evans, 1995; Faust and Rogers,

banks that we consider. Put bluntly, we live in a dollar world (Gourinchas, 2021), and the Fed is the main driver of global funding costs (Miranda-Agrippino and Rey, 2020; Miranda-Agrippino and Nenova, 2022). Our approach enables us to investigate whether the responses of exchange rates to monetary shocks differ according to the origin of the shocks. We check the extent to which exchange rate responses to SOE and US shocks are consistent with Dornbusch's overshooting and UIP. Finally, we provide new empirical evidence on the spillover effects of US monetary policy on advanced SOEs, controlling for the endogenous response of SOE central banks.

We find no evidence of delayed overshooting. In the six SOEs, a domestic contractionary monetary shock triggers a strong and immediate appreciation of the exchange rate, followed quickly by a gradual depreciation. Symmetrically, a tightening of US monetary policy causes an instantaneous depreciation followed by an appreciation. In all cases, the peak response of the exchange rate occurs very quickly, on impact or shortly after. Our findings support the view that delayed overshooting is not a genuine stylised fact but rather the outcome of dubious identifying restrictions (Faust and Rogers, 2003; Bjørnland, 2009). Moreover, we find little evidence of the forward discount puzzle: the responses of exchange rates to both SOE and US monetary shocks are broadly consistent with UIP, and thus with Dornbusch's overshooting mechanism. These findings are consistent with Bjørnland (2009) and Rüth (2020).[6] SOE and US monetary shocks explain about 20 and 10 percents respectively of the short-run exchange rate volatility. The smaller contribution of US shocks may be due to the endogenous, exchange-rate stabilizing, responses of SOE central banks to US monetary policy (Rey, 2013). Turning to spillovers, a tightening by the Federal Reserve induces output and inflation to fall in the six SOEs. This result echoes the findings of Maćkowiak (2007) for emerging economies. It is also consistent with Gopinath et al.

2003; Scholl and Uhlig, 2008; Kim et al., 2017; Rüth, 2020; Castelnuovo et al., 2022). Others study the impacts of non-US monetary shocks (Cushman and Zha, 1997; Kim and Roubini, 2000; Bjørnland, 2009; Bjørnland and Halvorsen, 2014; Kim and Lim, 2018; Doko Tchatoka et al., 2022). Earlier papers consider relative money shocks without taking a stance on the origin of disturbances (Clarida and Gali, 1994; Rogers, 1999).

[6] Faust and Rogers (2003) and Scholl and Uhlig (2008) find that the forward discount puzzle is more robust than the delayed overshooting puzzle.

(2020) and Akinci et al. (2022) who show that a monetary tightening by the Fed can generate a global slump.

An early study documenting the delayed overshooting puzzle is Eichenbaum and Evans (1995). They employ a recursive identification scheme and find evidence of a gradual and persistent appreciation in both the nominal and real US exchange rates in response to a contractionary US monetary policy shock. Their findings contradict Dornbusch (1976) immediate overshooting hypothesis.[7] Further studies by Faust and Rogers (2003) and Scholl and Uhlig (2008) replace the controversial recursive identification scheme with sign restrictions on the impulse response functions. However, these studies again document puzzling responses with delays lasting for around 3 years. Kim et al. (2017) using sign restrictions similar to Scholl and Uhlig (2008) report findings consistent with Dornbusch's prediction except during Volcker's tenure as Fed Chair. Rüth (2020) uses surprises in Federal funds futures around policy announcements as external instruments to estimate a proxy-SVAR model and measure the effects of U.S. monetary policy shocks on various measures of U.S. exchange rates. His findings are consistent with Dornbusch's predictions, including during Volcker's tenure. Castelnuovo et al. (2022) follow an approach similar to ours to investigate the delayed overshooting puzzle for the United States. They identify their SVAR model by applying restrictions on the structural parameters of the systematic component of US monetary policy and find no evidence of delayed overshooting puzzle.[8]

A stream of the SVAR literature focuses on SOEs. Cushman and Zha (1997) and Kim and Roubini (2000) apply non-recursive zero restrictions to implement block exogeneity and identify monetary policy shocks. Bjørnland (2009) uses data from four SOEs to estimate an SVAR model combining short-run and long-run zero restrictions.

[7] With nominal rigidities, the responses of the real and nominal exchange rates are similar in the short run.

[8] Our study complements the work of Castelnuovo et al. (2022). Castelnuovo et al. (2022) concentrate on US monetary shocks and the US economy. Instead, we focus on SOEs, we implement block exogeneity, and we jointly identify SOE and US monetary policy shocks. In addition, Castelnuovo et al. (2022) restrict the IRFs of industrial production and prices to be negative on impact, while we do not impose any sign restriction on IRFs.

Her identification scheme allows for simultaneous interaction between monetary policy and the exchange rate, while requiring that monetary shocks have no impact on the real exchange rate in the long-run. She finds no evidence of delayed overshooting, suggesting that Dornbusch was right after all. Recently, Doko Tchatoka et al. (2022) estimate a time-varying SVAR model with stochastic volatility using the same data and identification scheme as Bjørnland (2009). Their results are in line with Bjørnland (2009). Other studies applying agnostic identification procedures to analyse the exchange-rate response to monetary shocks in SOEs include Bjørnland and Halvorsen (2014) and Kim and Lim (2018).[9] Bjørnland and Halvorsen (2014) consider six SOE and identify monetary disturbances by imposing a combination of sign and exclusion restrictions on IRFs. Unlike us, they impose a sign restriction on the impact response of the exchange rate, forcing an instantaneous appreciation, and thus ruling out the so-called exchange rate puzzle by construction (Eichenbaum and Evans, 1995). They do not find evidence of delayed overshooting. Kim and Lim (2018) consider four SOEs and achieve identification by imposing sign-restrictions on IRFs. They confirm the findings of Bjørnland and Halvorsen (2014). In contrast to these two studies, we apply the methodology developed by Arias et al. (2018) to impose some exclusion restrictions for block exogeneity, as well as some sign-restrictions on the coefficients of the systematic component of monetary policy.[10]

The rest of the chapter is structured as follows. Section 1.2 describes the data, the identification scheme and the Bayesian estimation. Section 1.3 presents our main results. Section 1.4 contains robustness checks and Section 1.5 concludes.

[9] See also Jääskelä and Jennings (2011) and Read (2023) for related studies focusing on Australia.

[10] Bjørnland and Halvorsen (2014) and Kim and Lim (2018) use the Penalty Function Approach proposed by Uhlig (2005). Arias et al. (2018) criticize the PFA approach for imposing additional restrictions that are not specified by the user.

1.2 Econometric strategy

1.2.1 Data

We consider six advanced small-open economies with flexible exchange rates, namely Australia, Canada, New Zealand, Norway, Sweden and the UK. We use quarterly data from 1992:Q1 to 2019:Q4.[11] The starting date corresponds broadly to the adoption of Inflation Targeting by the six SOEs considered here (Kim and Lim, 2018). The end date marks the onset of the COVID-19 pandemic.[12] Following Cushman and Zha (1997), we organize the variables into two blocks, a foreign one and a domestic one. The foreign block stands for the US economy, while the domestic block represents the SOE. The foreign block consists of four variables: US real GDP (y^*), US inflation (π^*), Moody's Baa corporate credit spread (cs^*), and the US shadow rate (r^*) constructed by Wu and Xia (2016). The domestic block includes real GDP (y), inflation (π) measured as the annualized quarterly rate of change in the consumer price index, the policy rate (r) proxied by the 3-month interbank rate, and the bilateral SOE/US real exchange rate (e).[13] All variables are expressed in log levels except the credit spread, inflation rates and policy rates, which are expressed in percentage points.

1.2.2 Model

Our structural model is given by:

$$\mathbf{y_t'A_0} = \sum_{l=1}^{p} \mathbf{y}_{t-l}'\mathbf{A}_l + \mathbf{c}' + \epsilon_\mathbf{t}' \ , \tag{1.1}$$

where $\mathbf{y_t'} = \begin{bmatrix}\mathbf{y_{1t}'} & \mathbf{y_{2t}'}\end{bmatrix}$, $\mathbf{y_{1t}'} = [y_t^*, \pi_t^*, cs_t^*, r_t^*]$ and $\mathbf{y_{2t}'} = [y_t, \pi_t, r_t, e_t]$. $\mathbf{y_{1t}}$ is a $(n_1 \times 1)$ vector of US variables and $\mathbf{y_{2t}}$ is a $(n_2 \times 1)$ vector of SOE variables, with $n = n_1 + n_2$

[11] Data sources are provided in Appendix A.

[12] Estimating the SVAR over a stable monetary policy regime helps solving the delayed overshooting puzzle. See Kim and Lim (2018), Kim et al. (2017) and Castelnuovo et al. (2022).

[13] For Sweden and the UK, we use shadow rates constructed by De Rezende and Ristiniemi (2023) and Wu and Xia (2016), respectively.

denoting the total number of variables. Similarly, the vector of structural shocks $\epsilon_{\mathbf{t}}$ is divided into two blocks, $\epsilon_{\mathbf{t}}{}' = \begin{bmatrix} \epsilon'_{1t} & \epsilon'_{2t} \end{bmatrix}$. $\mathbf{A_i}$, for $0 \leq i \leq p$, are $(n \times n)$ matrices of structural parameters, with $\mathbf{A_0}$ invertible. $\mathbf{c}$ is a $(n \times 1)$ vector of constants, p is the lag length, and T is the sample size. Conditional on past information and initial conditions $\mathbf{y}_0, ..., \mathbf{y}_{1-p}$, the vector $\epsilon_{\mathbf{t}}$ is Gaussian with mean zero and covariance matrix $\mathbb{I}_n$. Following Rubio-Ramirez et al. (2010), we can write the SVAR in compact form:

$$\mathbf{y_t}{}'\mathbf{A_0} = \mathbf{x_t}{}'\mathbf{A}_+ + \epsilon'_t \ , \tag{1.2}$$

where $\mathbf{x'_t} = [\mathbf{y'_{t-1}} \dots \mathbf{y'_{t-p}} \ \ 1]$ for $1 \leq \mathrm{t} \leq \mathrm{T}$. $\mathbf{A_0}$ and $\mathbf{A}_+ = [\mathbf{A'_1} \dots \mathbf{A'_p} \ \ \mathbf{c}']$ are matrices of structural parameters.

Post-multiplying Equation (1.2) by $\mathbf{A_0^{-1}}$, we obtain the reduced form VAR model:

$$\mathbf{y'_t} = \mathbf{x'_t}B + u'_t \ , \tag{1.3}$$

where $B = \mathbf{A}_+\mathbf{A_0^{-1}}, u'_t = \epsilon'_t\mathbf{A_0^{-1}}$ and $\mathbf{E}[u_t u'_t] = \Sigma = (\mathbf{A_0}\mathbf{A'_0})^{-1}$. B is the matrix of reduced-form coefficients and Σ is the variance-covariance matrix of reduced-form residuals.

1.2.3 Identification

Our strategy aims at identifying simultaneously SOE and US monetary shocks by bringing together two distinct approaches: sign restrictions on policy parameters (Arias et al., 2019) and block exogeneity (Cushman and Zha, 1997).

The first procedure, sign restrictions on policy parameters, offers an agnostic approach to identify the systematic component of monetary policy, and thereby monetary policy shocks.[14] The appeal of this method stems from its agnosticism and robustness as it hinges solely on a few qualitative and fairly uncontroversial restrictions on the

[14] Leeper et al. (1996) make explicit the link between identifying the systematic component of monetary policy and identifying monetary policy shocks.

structural coefficients of the monetary policy rule. This method only achieves set-identification. A caveat inherent in identification schemes based on sign-restrictions is the so-called multiple shocks problem: many shocks, different from the one we are trying to identify, may satisfy the set of sign-restrictions (Fry and Pagan, 2011). The fact that we impose sign restrictions simultaneously on both US and SOE policy parameters should in principle help to alleviate this partial-identification problem.

The second procedure, block exogeneity, is the hallmark of any SOE model: the SOE is influenced by foreign factors and has no impact on the Rest of the World. In practice, block exogeneity consists in imposing a set of non-recursive zero restrictions. In our setting, block exogeneity complements the minimal set of sign restrictions on policy parameters and augments the information content of our identification scheme. In other words, block exogeneity strengthens the identification of US and SOE monetary shocks through a set of highly plausible zero restrictions.

Our goal in this chapter is to assess the robustness of the delayed overshooting puzzle. Importantly for our purpose, our identification strategy allows for a simultaneous relationship between the exchange rate and the SOE policy rate and leaves the response of the exchange rate to SOE and US monetary shocks unrestricted at all horizons (Faust and Rogers, 2003; Bjørnland, 2009). We first present details about the way we implement block exogeneity. We then explain the identification of the systematic component of monetary policy in the US and in SOEs through sign restrictions on structural parameters.

Block exogeneity

We adapt the methodology of Arias et al. (2019) to incorporate block exogeneity (Cushman and Zha, 1997). Given the partition of $\mathbf{y_t}' = \begin{bmatrix} \mathbf{y_{1t}}' & \mathbf{y_{2t}}' \end{bmatrix}$, the matrix of contemporaneous relationships, $\mathbf{A_0}$, has the following structure:

$$\mathbf{A_0} = \begin{bmatrix} A_{0,11} & A_{0,12} \\ \\ A_{0,21} & A_{0,22} \end{bmatrix},$$

where $A_{0,11}$ is $(n_1 \times n_1)$, $A_{0,12}$ is $(n_1 \times n_2)$, $A_{0,21}$ is $(n_2 \times n_1)$, $A_{0,22}$ is $(n_2 \times n_2)$. To ensure that SOE variables in $\mathbf{y_{2t}}$ do not influence US variables in $\mathbf{y_{1t}}$ contemporaneously, we apply zero-restrictions on the block $A_{0,21}$:

$$\mathbf{A_0} = \begin{bmatrix} A_{0,11} & A_{0,12} \\ \\ 0 & A_{0,22} \end{bmatrix}.$$

We should also prevent SOE variables from influencing US variables in a dynamic fashion. Put differently, in line with Cushman and Zha (1997), we should impose a block of zero-restrictions on each lag matrix $\mathbf{A}_l$, $1 \leq l \leq p$, in Equation (1.1), so that:

$$\mathbf{A}_l = \begin{bmatrix} A_{11,l} & A_{12,l} \\ \\ A_{21,l} & A_{22,l} \end{bmatrix} = \begin{bmatrix} A_{11,l} & A_{12,l} \\ \\ 0 & A_{22,l} \end{bmatrix},$$

where $A_{11,l}$ is $(n_1 \times n_1)$, $A_{12,l}$ is $(n_1 \times n_2)$, $A_{21,l}$ is $(n_2 \times n_1)$, $A_{22,l}$ is $(n_2 \times n_2)$. Unfortunately, the procedure of Arias et al. (2019) only allows us to impose a maximum of $(n-k)$ zero restrictions per equation, where $k = 1, ..., n$, denotes the order of the k^{th} equation in the system. As a result, we cannot impose $A_{21,l} = \mathbf{0}$.

We bypass this issue by formulating a variant of Minnesota priors on the reduced-form VAR, where the priors for the coefficients governing the influence of lagged SOE variables on US variables are concentrated tightly around zero. To do that, we specify Independent $\mathcal{NIW}$ priors for $\beta = vec(B)$, the vector of reduced-form coefficients, and Σ, the variance-covariance matrix of reduced-form residuals:[15]

[15] Arias et al. (2019) specify Natural Conjugate $\mathcal{NIW}$ priors for the reduced-form parameters.

$$\beta \sim \mathcal{N}(\beta_0, \Omega_0), \tag{1.4}$$

$$\Sigma \sim \mathcal{IW}(S_0, \alpha_0).^{16} \tag{1.5}$$

We center the distribution of every first-order auto-regressive coefficient at 1, and 0 otherwise, as in standard Minnesota priors. The variance-covariance matrix Ω_0 contains the hyper-parameters that control the tightness of the distributions of reduced-form coefficients.[17] The elements of Ω_0 take the following form:

$$\sigma_{c_i}^2 = \sigma_i^2(\lambda_1\lambda_4)^2 \qquad \text{if constant} \tag{1.6}$$

$$\sigma_{ii}^2 = (\lambda_1/L^{\lambda_3})^2 \qquad \text{if } i = j \tag{1.7}$$

$$\sigma_{ij}^2 = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 \qquad \text{if } i \neq j \tag{1.8}$$

$$\sigma_{Ex_{ij}}^2 = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 \qquad \text{if } i \neq j \text{ and } e_x < j \leq n \tag{1.9}$$

where σ_i^2 and σ_j^2 denote the variances of OLS residuals of the auto-regressive models estimated for variables i and j. e_x denotes the number of exogenous variables. Equations (1.6), (1.7) and (1.8) describe the standard Minnesota priors. Equation (1.9) is key to implement block exogeneity: it only applies to the US block and features the additional hyper-parameter λ_5, which controls the tightness of the distributions of coefficients appearing in front of SOE variables in the US block (Dieppe et al., 2016). We set $\lambda_5 = 1e-8$ to obtain highly informative priors concentrated around zero. We select standard prior variances for the rest of the parameters ($\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 1$). To sum up, we implement block exogeneity through a combination of two ingredients:

Such Natural Conjugate priors are ill-fitted for our purpose: they feature a Kronecker structure for the variance-covariance matrix of the reduced-form parameters, so that variances are proportional to one another. Moreover, the Kronecker structure implies that every equation has the same set of explanatory variables, meaning that if we removed a variable in one equation, that variable would be removed from all equations. Imposing block exogeneity on one equation would then impose it on all equations (Dieppe et al., 2016; Koop et al., 2010). Fortunately, the techniques developed by Arias et al. (2018) work for any prior distributions.

[16] We set the hyperparameters of the inverse Wishart distribution in a conventional way: $\alpha_0 = n+1$ and $S_0 = \mathbb{I}_n$ (Dieppe et al., 2016).

[17] Unlike with Natural Conjugate ($\mathcal{NIW}$) priors, Ω_0 is independent of Σ.

i) we impose exclusion restrictions on the block $A_{0,21}$ in the structural-form matrix of contemporaneous relationships; ii) we formulate a special case of Independent $\mathcal{NIW}$ priors for the reduced-form VAR, where the reduced-form coefficients follow Minnesota priors with an additional hyper-parameter for block exogeneity.

Sign-restrictions on monetary policy parameters

Our identification scheme builds heavily on Arias et al. (2019). Using the techniques developed by Arias et al. (2018), they impose sign and exclusion restrictions on the coefficients of the Federal Reserve's interest-rate rule. As mentioned above, we adapt their methodology to a SOE context in two ways. First, we impose a block-exogenous structure on the SVAR model, meaning that for each SOE, the variables are classified into a US block and a domestic block, where the SOE block has no effect on the US block. Second, for each SOE, we identify simultaneously the interest-rate rule followed by the Federal Reserve and the SOE central bank. Like in Arias et al. (2019), our identification concentrates on the contemporaneous structural parameters. We abstract from the constant term and lags of the structural model in Equation (1.1):

$$\begin{bmatrix} y_t^* \ \pi_t^* \ cs_t^* \ r_t^* \ y_t \ \pi_t \ r_t \ rer_t \end{bmatrix} \begin{bmatrix} a_{0,11} & a_{0,12} & a_{0,13} & a_{0,14} & a_{0,15} & a_{0,16} & a_{0,17} & a_{0,18} \\ a_{0,21} & a_{0,22} & a_{0,23} & a_{0,24} & a_{0,25} & a_{0,26} & a_{0,27} & a_{0,28} \\ a_{0,31} & a_{0,32} & a_{0,33} & a_{0,34} & a_{0,35} & a_{0,36} & a_{0,37} & a_{0,38} \\ a_{0,41} & a_{0,42} & a_{0,43} & a_{0,44} & a_{0,45} & a_{0,46} & a_{0,47} & a_{0,48} \\ 0 & 0 & 0 & 0 & a_{0,55} & a_{0,56} & a_{0,57} & a_{0,58} \\ 0 & 0 & 0 & 0 & a_{0,65} & a_{0,66} & a_{0,67} & a_{0,68} \\ 0 & 0 & 0 & 0 & a_{0,75} & a_{0,76} & a_{0,77} & a_{0,78} \\ 0 & 0 & 0 & 0 & a_{0,85} & a_{0,86} & a_{0,87} & a_{0,88} \end{bmatrix} = \begin{bmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \\ \epsilon_{3,t} \\ \epsilon_{4,t} \\ \epsilon_{5,t} \\ \epsilon_{6,t} \\ \epsilon_{7,t} \\ \epsilon_{8,t} \end{bmatrix}'$$

We identify the first and the fifth shock in the SVAR model as the US and the SOE monetary policy shocks, respectively:

$$y_t^* a_{0,11} + \pi_t^* a_{0,21} + cs_t^* a_{0,31} + r_t^* a_{0,41} = \epsilon_{1,t} \tag{1.10}$$

$$y_t^* a_{0,15} + \pi_t^* a_{0,25} + cs_t^* a_{0,35} + r_t^* a_{0,45} + y_t a_{0,55} + \pi_t a_{0,65} + r_t a_{0,75} + rer_t a_{0,85} = \epsilon_{5,t} \tag{1.11}$$

2.3.2.1 The systematic component of US monetary policy: The US monetary policy equation, from Equation (1.10), is given by:

$$r_t^* = -a_{0,41}^{-1} a_{0,11} y_t^* - a_{0,41}^{-1} a_{0,21} \pi_t^* - a_{0,41}^{-1} a_{0,31} cs_t^* + a_{0,41}^{-1} \epsilon_{1,t}, \tag{1.12}$$

where $-a_{0,41}^{-1} a_{0,11} = \psi_{y^*}$, $-a_{0,41}^{-1} a_{0,21} = \psi_{\pi^*}$, $-a_{0,41}^{-1} a_{0,31} = \psi_{cs^*}$ and $a_{0,41}^{-1} = \sigma^*$.

To characterize the systematic component of US monetary policy, we impose the following two restrictions.

Restriction 1. The contemporaneous response of the US policy rate to US output and US inflation is positive: $\psi_{y^*} > 0$ and $\psi_{\pi^*} > 0$.

Restriction 2. The contemporaneous reaction of the US policy rate to the Baa corporate credit spread is negative: $\psi_{cs^*} < 0$.

Restriction 1 is motivated by Taylor (1993) and a large DSGE literature.[18] Restriction 2 is consistent with the SVAR evidence provided by Caldara and Herbst (2019).[19] Combining Restrictions 1 and 2, we obtain the following characterization of US monetary policy:

$$r_t^* = \underbrace{-a_{0,41}^{-1} a_{0,11}}_{\psi_{y^*} > 0} y_t^* \underbrace{-a_{0,41}^{-1} a_{0,21}}_{\psi_{\pi^*} > 0} \pi_t^* \underbrace{-a_{0,41}^{-1} a_{0,31}}_{\psi_{cs^*} < 0} cs_t^* \underbrace{+a_{0,41}^{-1}}_{\sigma^*} \epsilon_{1,t} \tag{1.13}$$

[18] Regarding the timing assumption implied by Restriction 1, where the policy rate, output and inflation interact simultaneously, we follow the argument of Arias et al. (2019): Monetary authorities crunch a large battery of real-time indicators to nowcast the current state of the economy.

[19] Curdia and Woodford (2010, 2016) present DSGE-based analysis justifying a negative systematic response of monetary policy to a worsening of credit conditions.

2.3.2.2 The systematic component of monetary policy in SOEs: The SOE monetary policy equation, from Equation (1.11), is given by:

$$\begin{aligned} r_t = & -a_{0,75}^{-1} a_{0,15} y_t^* - a_{0,75}^{-1} a_{0,25} \pi_t^* - a_{0,75}^{-1} a_{0,35} cs_t^* - a_{0,75}^{-1} a_{0,45} r_t^* \\ & -a_{0,75}^{-1} a_{0,55} y_t - a_{0,75}^{-1} a_{0,65} \pi_t - a_{0,75}^{-1} a_{0,85} rer_t + a_{0,75}^{-1} \epsilon_{5,t}, \end{aligned} \tag{1.14}$$

where $-a_{0,75}^{-1} a_{0,55} = \psi_y$, $-a_{0,75}^{-1} a_{0,65} = \psi_\pi$, $-a_{0,75}^{-1} a_{0,85} = \psi_e$ and $a_{0,75}^{-1} = \sigma$.

To identify the systematic component of SOE monetary policy, we impose the following two restrictions.

Restriction 3. The contemporaneous reaction of the SOE policy rate to domestic output and inflation is positive: $\psi_y > 0$ and $\psi_\pi > 0$.

Restrictions 4: The contemporaneous reaction of the SOE policy rate to the real bilateral SOE/US exchange rate is positive: $\psi_e > 0$.

Restrictions 3 and 4 leave the reaction of the SOE central bank to foreign variables unrestricted as in Cushman and Zha (1997). Restriction 4 means that SOE central bank usually leans against the real SOE/US exchange rate, cutting its policy rate in response to an appreciation of the domestic currency, and raising it in response to a depreciation. Restriction 4 is consistent with findings based on SVARs (Bjørnland, 2009; Bjørnland and Halvorsen, 2014) and Taylor-type rules embedded in DSGE models (Lubik and Schorfheide, 2007; Kam et al., 2009; Justiniano and Preston, 2010).[20] Taken together, Restrictions 3 and 4 imply that the SOE central bank follows a Taylor-type rule in line with Taylor (2001):

[20] Calvo and Reinhart (2002), Reinhart and Rogoff (2004), Obstfeld (2013) and Ilzetzki et al. (2019) find that many central banks react to the dollar exchange rate. Egorov and Mukhin (2023) argue that, under dollar pricing, stabilizing the dollar exchange rate can be desirable.

$$r_t = \underbrace{-a_{0,75}^{-1}a_{0,15}}_{unrestricted} y_t^* \underbrace{-a_{0,75}^{-1}a_{0,25}}_{unrestricted} \pi_t^* \underbrace{-a_{0,75}^{-1}a_{0,35}}_{unrestricted} cs_t^* \underbrace{-a_{0,75}^{-1}a_{0,45}}_{unrestricted} r_t^*$$

$$\underbrace{-a_{0,75}^{-1}a_{0,55}}_{\psi_y>0} y_t \underbrace{-a_{0,75}^{-1}a_{0,65}}_{\psi_\pi>0} \pi_t \underbrace{-a_{0,75}^{-1}a_{0,85}}_{\psi_e>0} rer_t \underbrace{+a_{0,75}^{-1}}_{\sigma} \epsilon_{5,t} \tag{1.15}$$

1.3 Results

This section presents our main results.[21] First, we discuss the IRFs to a tightening of SOE and US monetary policy. Second, we investigate the importance of Restriction 4 (*leaning against the RER*) and Restriction 2 (*leaning against credit frictions*) for the identification of, respectively, SOE and US monetary shocks. Third, we examine the deviations from UIP conditional on SOE and US monetary disturbances. Finally, we conduct a Forecast Error Variance Decomposition to evaluate the contribution of SOE and US shocks to the short-run volatility of the SOE/US real exchange rates.

1.3.1 IRFs to a SOE contractionary monetary shock

Figure 1.1 plots the IRFs of domestic variables for the six SOEs to one standard deviation contractionary monetary shock. The solid lines depict the point-wise posterior median responses while the grey shaded bands correspond to the 68% equal-tailed point-wise posterior probability bands. For all SOEs, we observe that the policy rate jumps on impact within a range of 15 to 40 basis points, and reaches its peak within the next three quarters. Except for Canada, the policy rate increase remains significant for several quarters. For all SOEs, the posterior median response of output displays an instantaneous contraction and stays below trend for several years after the shock. Except for the UK, the decline in output remains significant for several quarters after

[21] We set the lag order $p = 2$. he posterior distributions.

the shock. For the UK, the response of output is insignificant, although the bulk of the 68% probability bands lies in the negative region, suggesting that output contracts at least in the short run. For Canada New Zealand, the response of output stays significantly below trend throughout the entire five-year horizon. Hence, we do not observe any evidence of the output puzzle (Uhlig, 2005).

For all SOEs, inflation falls instantaneously and reverts back to its steady state quickly. The negative impact response of inflation is either significant or borderline significant across all SOEs.

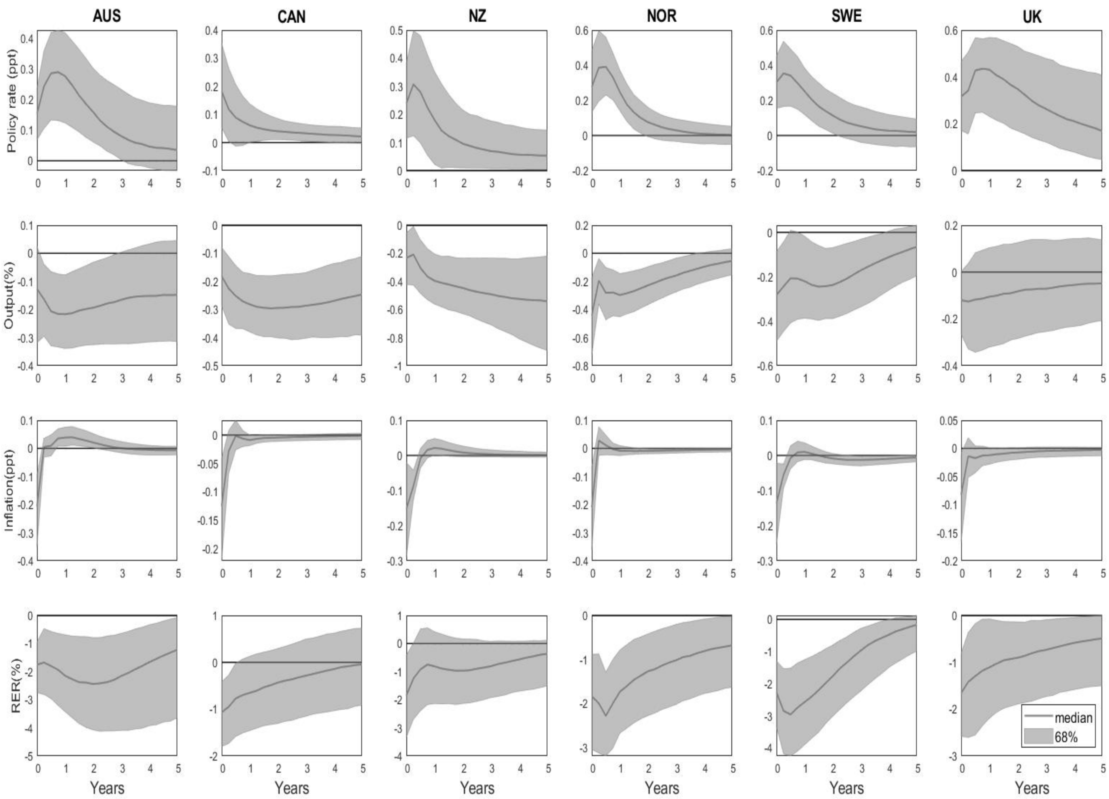


Figure 1.1: IRFs of SOE variables to a one standard deviation SOE contractionary monetary shock identified using block exogeneity and Restrictions 1 to 4. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 68% equal-tailed point-wise posterior probability bands.

Turning to our focus of interest, in all SOEs, the RER appreciates sharply and significantly on impact. This is a remarkable result: We do not find any evidence of the exchange rate puzzle (an immediate depreciation after the tightening of domestic monetary policy). For Canada, New Zealand and the UK, the instantaneous appreciation is immediately followed by a gradual and monotonous depreciation. For Norway and Sweden, the RER appreciation reaches its peak two quarters after the monetary tightening. Instead, the AUD/US RER displays a hump-shaped response, with the peak appreaciation occuring two years after the shock. Thus, except for Australia, we observe little evidence of the delayed overshooting puzzle (a gradual and persistent appreciation that reaches its peak roughly two years after the shock). The IRFs of the bilateral SOE/US RER to a SOE monetary shock appear broadly consistent with the Dornbusch's overshooting hypothesis. Our findings are in line with the SOE-SVAR studies by Bjørnland (2009), Bjørnland and Halvorsen (2014), Kim and Lim (2018) and Doko Tchatoka et al. (2022). Our findings reinforce the view that the exchange rate puzzle and the delayed overshooting puzzle may be artefacts caused by dubious identifying restrictions that hinder the simultaneous interactions between monetary policy and the exchange rate.

1.3.2 IRFs to a US contractionary monetary shock

Figure 1.2 shows the responses of SOE variables to a one standard deviation US contractionary monetary shock. The responses of US variables are reported in the Appendix. They are almost identical across the six SOE-SVARs due to the block-exogeneous structure of the model: the US policy rate jumps significantly on impact to around 20 basis points, US output contracts, US inflation falls and the Baa corporate credit spread increases.

Looking at the RER responses across the six SOEs, we observe that the US dollar appreciates significantly on impact in response to the US monetary tightening (i.e., no exchange rate puzzle). Moreover, the US dollar reaches its peak appreciation within the first quarter after the shock, and gradually depreciates afterwards (i.e., no delayed overshooting). Similar findings are reported in US SVAR monetary policy literature

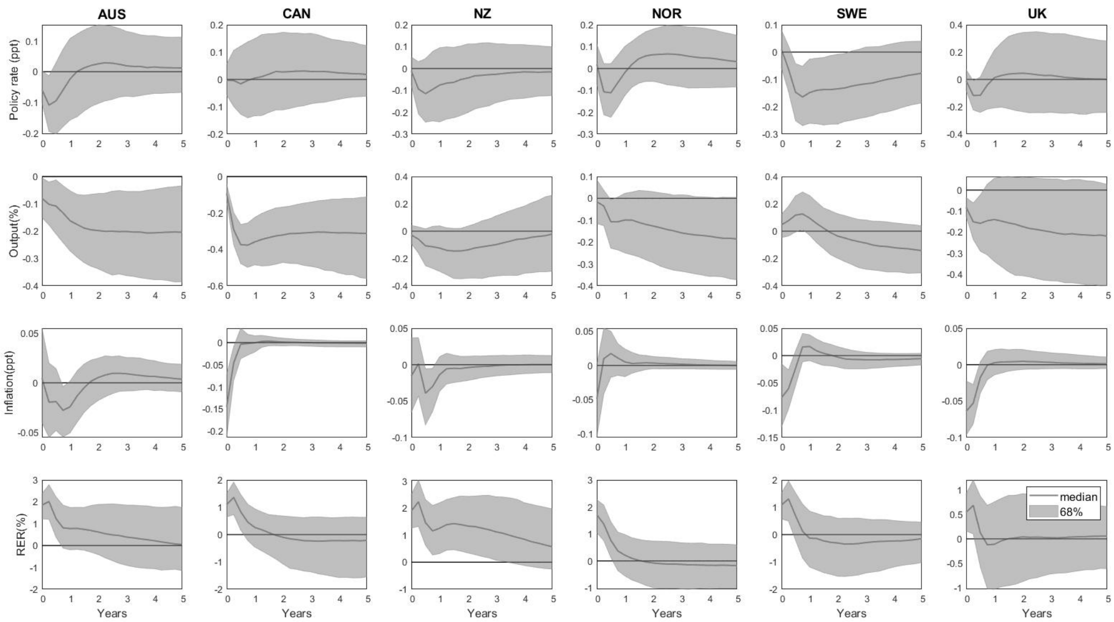


Figure 1.2: IRFs of SOE variables to a one standard deviation US contractionary monetary shock identified using block exogeneity and Restrictions 1 to 4. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 68% equal-tailed point-wise posterior probability bands.

(Kim et al., 2017; Rüth, 2020; Castelnuovo et al., 2022). A distinguishing feature of our study, however, is the inclusion of the zero lower bound (ZLB) period.[22]

Looking at the response of SOE output to US monetary tightening, we observe a protracted contraction for Australia, Canada, the UK and to a lesser extent Norway. The decline in output is particularly strong in Canada and Australia. The response of output in New Zealand and Sweden is more muted. Following the US monetary tightening, inflation falls in all SOEs. Except for Canada, SOE central banks lower their policy rate slightly to mitigate the negative spillovers of the US monetary tightening. The monetary easing is most visible in Sweden and Australia.

[22] Our results are robust to using a shorter sample period from 1992:Q1 to 2008:Q3 that excludes the ZLB period. See the section on robustness checks below.

1.3.3 Importance of Restriction 4 (leaning against the RER)

We now perform a sensitivity analysis to shed light on the importance of Restriction 4 ($\psi_e > 0$) in our identification scheme. We re-estimate the six SOE-SVAR models without Restriction 4 while keeping everything else unchanged. As a result, the contemporaneous response of the SOE policy rate to the real exchange rate is now left unrestricted. The systematic component of SOE monetary policy takes the following form:

$$
\begin{aligned}
r_t = & \underbrace{-a_{0,75}^{-1}a_{0,15}}_{unrestricted}\, y_t^* \underbrace{-a_{0,75}^{-1}a_{0,25}}_{unrestricted}\, \pi_t^* \underbrace{-a_{0,75}^{-1}a_{0,35}}_{unrestricted}\, cs_t^* \underbrace{-a_{0,75}^{-1}a_{0,45}}_{unrestricted}\, r_t^* \\
& \underbrace{-a_{0,75}^{-1}a_{0,55}}_{\psi_y>0}\, y_t \underbrace{-a_{0,75}^{-1}a_{0,65}}_{\psi_\pi>0}\, \pi_t \underbrace{-a_{0,75}^{-1}a_{0,85}}_{unrestricted}\, rer_t \underbrace{+a_{0,75}^{-1}}_{\sigma}\, \epsilon_{5,t}.
\end{aligned} \tag{1.16}
$$

.

Figure 1.3 compares the baseline IRFs (solid lines) of SOE variables to an SOE monetary shock, with Restriction 4 imposed as in Equation (1.15), to the sensitivity-analysis IRFs (dashed lines) without Restriction 4, as in Equation (1.16). The shaded regions are the 68% equal-tailed point-wise posterior probability bands obtained after relaxing Restriction 4. Figure 1.3 shows that relaxing Restriction 4 has virtually no effect on the IRFs of the policy rate, output and inflation. Instead, relaxing Restriction 4 greatly alters the IRFs of the RER. Most remarkable is the fact that, for all SOEs, the effects of monetary shocks on the RER are now insignificant, even in the short run. This finding clearly goes against the consensus view that monetary policy plays a role in accounting for the elevated short-run volatility typically observed in exchange rates. Beyond this striking observation, we also find that the evidence of other puzzles becomes stronger. For Canada, the posterior median response clearly indicates that the exchange rate depreciates instead of appreciating, consistent with the exchange rate puzzle. Taken together, these puzzling findings stand in stark contrast to Dornbusch (1976) overshooting hypothesis according to which a surprise tightening of monetary

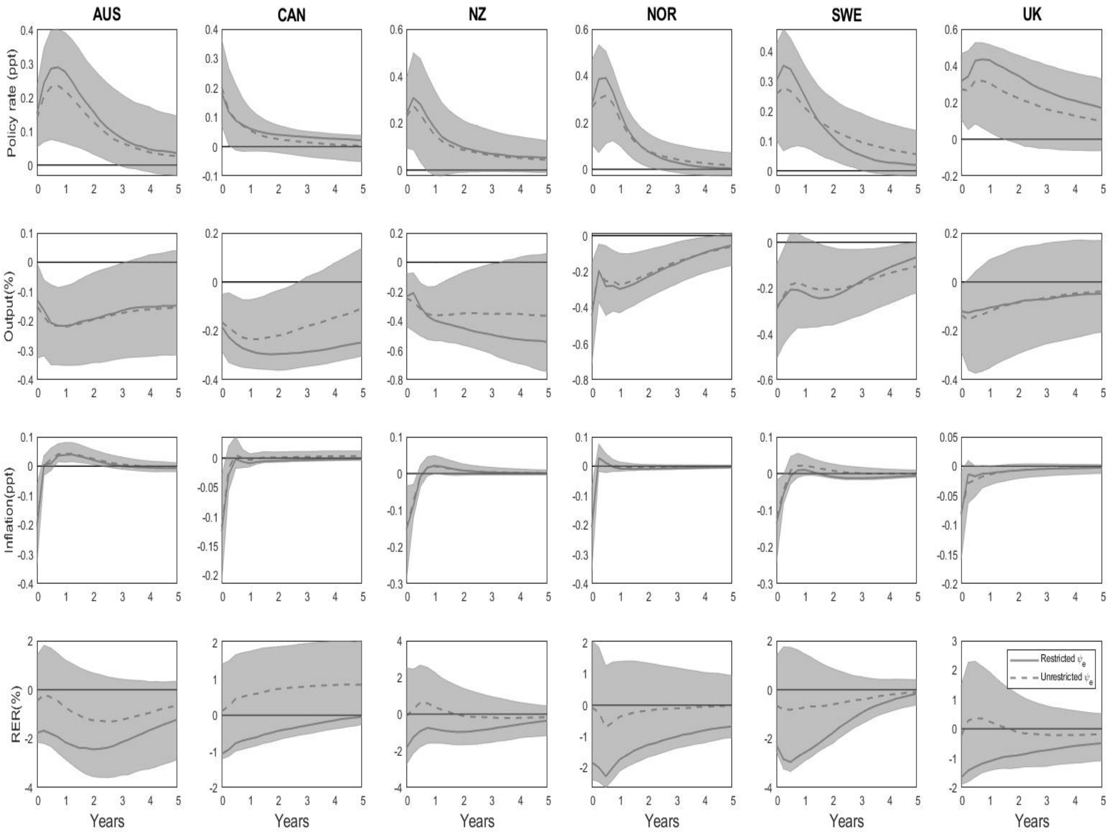


Figure 1.3: IRFs of SOE variables to a SOE contractionary monetary shock. *Note:* The solid lines are the point-wise posterior median responses under block exogeneity and Restrictions 1 to 4. The dashed lines are the posterior median responses, after relaxing Restriction 4, along with the corresponding 68% equal-tailed posterior probability bands.

policy at home causes an instantaneous appreciation of the domestic currency, immediately followed by a gradual depreciation back to the steady state. Considering the puzzling evidence obtained when relaxing Restriction 4 and the assorted motivations for imposing Restriction 4 found in various strands of the literature (Taylor, 2001; Bjørnland, 2009; Lubik and Schorfheide, 2007), we conclude that imposing Restriction 4 contributes usefully to a proper identification of the systematic behavior of SOE central banks.[23]

[23] The IRFs to a US monetary shock with and without Restriction 4 are reported in Appendix A. They show that Restriction 4 is irrelevant for identifying US monetary shocks.

1.3.4 Role of Restriction 2 (leaning against the credit spread)

We now evaluate the implications of Restriction 2 ($\psi_{cs^*} < 0$) for the identification of the systematic component of US monetary policy. We re-estimate the six SOE-SVARs without Restriction 2, keeping everything else unchanged. The contemporaneous response of the US policy rate to the credit spread is left unrestricted:

$$r_t^* = \underbrace{-a_{0,41}^{-1}a_{0,11}}_{\psi_{y^*}>0} y_t^* \underbrace{-a_{0,41}^{-1}a_{0,21}}_{\psi_{\pi^*}>0} \pi_t^* \underbrace{-a_{0,41}^{-1}a_{0,31}}_{unrestricted} cs_t^* \underbrace{+a_{0,41}^{-1}}_{\sigma^*} \epsilon_{1,t}. \tag{1.17}$$

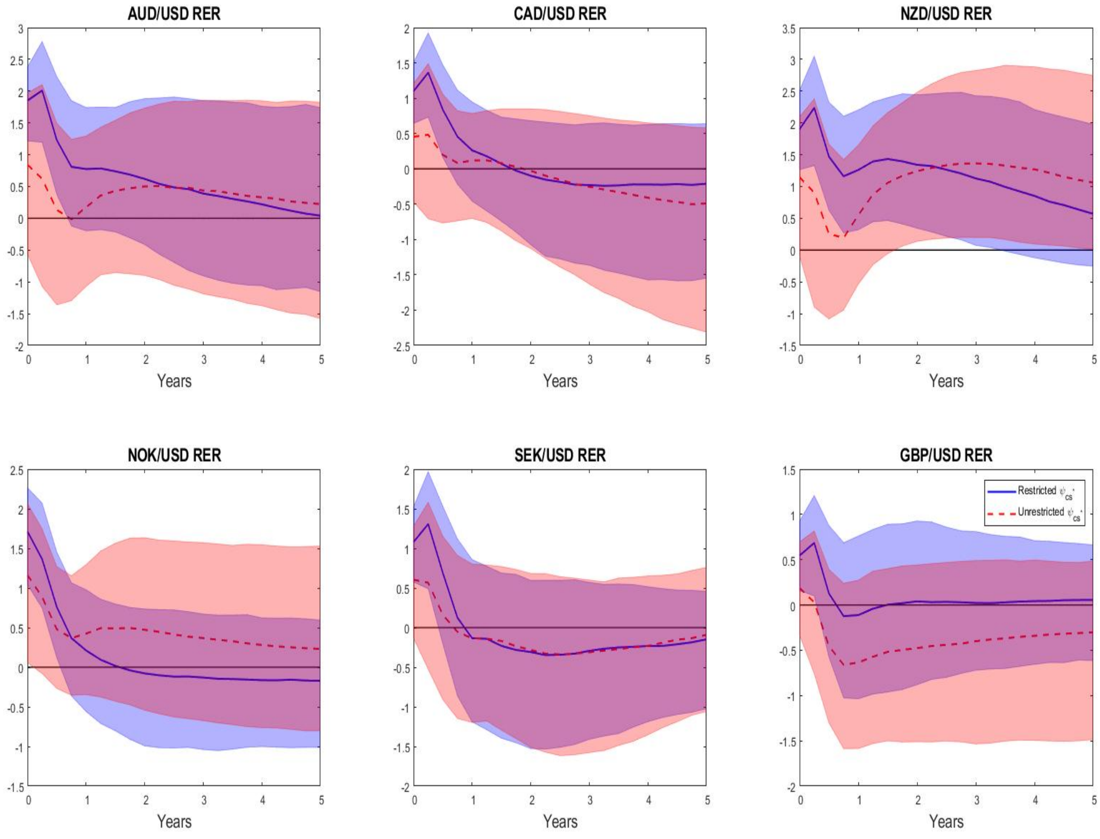


Figure 1.4: IRFs of the RER to a US contractionary monetary shock. *Note:* The solid lines are the point-wise posterior median responses under block exogeneity and Restrictions 1 to 4. The dashed lines are the posterior median responses after relaxing Restriction 2 ($\psi_{cs^*} < 0$). The shaded regions are the corresponding 68% equal-tailed posterior probability bands.

Figure 1.4 compares the IRFs of the SOE/US RER to a US contractionary monetary policy shock, with and without Restriction 2. As we can see, the main effect of relaxing Restriction 2 is that the 68% posterior probability bands become wider, to the extent that the short-run response of the RER becomes insignificant for the six SOEs. This finding, which suggests that US monetary policy shocks have no material effects on exchange rates even in the short run, goes against the conventional wisdom on the contribution of monetary disturbances to exchange rate volatility. Moreover, the fact that we are here talking about US monetary policy (and not about SOE monetary policy, as in the previous sensitivity analysis of Restriction 4), which is perceived as the main driver of the global financial cycle (Rey, 2013; Miranda-Agrippino and Rey, 2020), makes this finding look somewhat implausible. We generally observe that relaxing Restriction 2 shifts the posterior probability bands towards negative territory, meaning that the indicative evidence of a depreciation of the US dollar (instead of an appreciation, as we would have expected) builds up. In other words, relaxing Restriction 2 makes the exchange rate puzzle more visible (see in particular the IRFs of the GBP/USD, CAD/USD and SEK/USD). Overall, these dubious phenomena emphasize the added value of imposing Restriction 2 to correctly characterize the systematic behavior of the Federal Reserve and thereby identify genuine US monetary shocks.

1.3.5 Deviations from UIP after a monetary policy shock

The uncovered interest rate parity (UIP) condition postulates that a decline in the interest rate differential between the foreign and the domestic policy rates has to be quantitatively offset by an expected depreciation of the nominal exchange rate one period ahead. Examining violations of UIP conditional on monetary disturbances is central to our study as UIP is one of the key building blocks underpinning Dornbusch's overshooting hypothesis (Rüth, 2020), and more generally the New Open Economy Macroeconomics (Lane, 2001). Following Eichenbaum and Evans (1995) and Bjørnland

(2009), we compute the excess return measured in domestic currency, Λ_t, as:

$$\Lambda_t = r_t^* - r_t + 4 \times (\mathbb{E}_\mathbf{t}\{s_{t+1}\} - s_t), \tag{1.18}$$

where s_t is the nominal exchange rate.[24] According to UIP, the excess return Λ_t should be zero at all horizons:

$$\mathbb{E}_\mathbf{t}\{\Lambda_{t+j}\} = 0 \qquad \text{for all } j \geqslant 0,$$

where $\mathbb{E}_t$ denotes the conditional expectations operator.

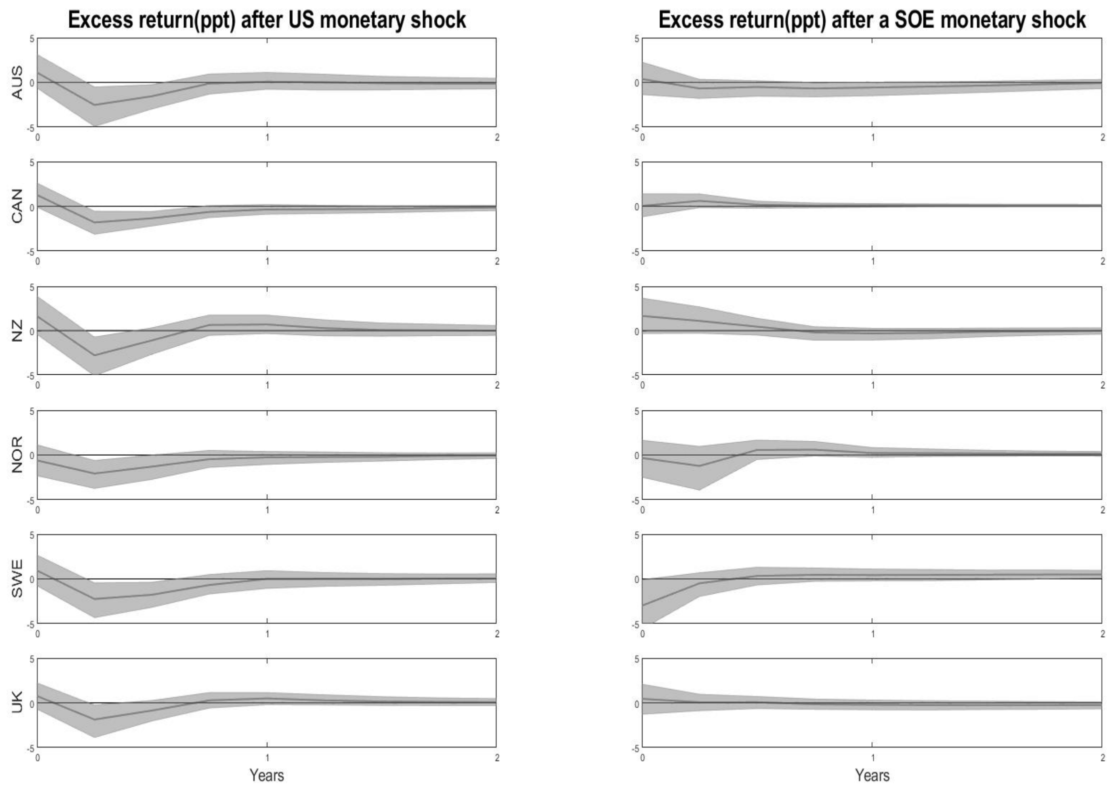


Figure 1.5: Deviations from UIP conditional on US (lhs) and SOE (rhs) monetary policy shocks. *Note:* The solid lines represent the point-wise posterior median estimates of excess returns. The shaded areas are the 68% posterior probability intervals.

Figure 1.5 reports the point-wise posterior median estimates of excess returns con-

[24] As our system includes SOE and US inflation rates along with the SOE/US RER, it is straightforward to construct the IRFs of the nominal exchange to a monetary shock.

ditional on US (lhs panel) and SOE (rhs panel) monetary policy shocks, along with the 68% posterior probability intervals. We do not find any evidence of UIP violations in response to SOE monetary shocks: Excess returns triggered by SOE disturbances are quantitatively modest and insignificant at all horizons. Deviations from UIP generated by US policy shocks are also moderate and insignificant, except during the quarter after the shock, when they are borderline significant. Thus, overall, the conditional dynamics of exchange rates following US and SOE monetary disturbances appear to be largely consistent with UIP. Our results are in line with Bjørnland (2009), who reports exchange rate movements broadly consistent with UIP conditional on SOE monetary disturbances, and with Rüth (2020) who finds little evidence of UIP violations conditional on US monetary shocks. Instead, Eichenbaum and Evans (1995), Faust and Rogers (2003) and Scholl and Uhlig (2008) report evidence of the forward discount puzzle, i.e. large and significant deviations from UIP, conditional on US monetary shocks.

1.3.6 Forecast error variance decomposition of the RER

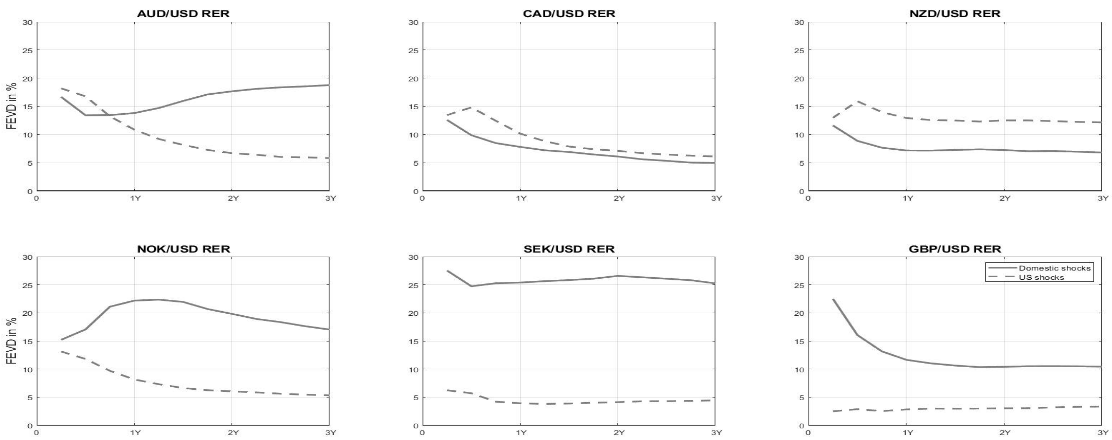


Figure 1.6: Forecast error variance decomposition of the SOE/US RER. *Note:* In each panel, the solid line represents the posterior median estimate of the contribution of SOE monetary shocks to the forecast-error variance of the RER, while the dashed line shows the contribution of US monetary shocks.

Figure 1.6 shows the forecast-error variance decomposition of the six SOE/US real exchange rates. With the exception of Canada and New Zealand, domestic monetary policy shocks account for a greater share of the RER volatility than US shocks. Depending on the country, domestic monetary policy shocks roughly explain 10 to 25 percents of the volatility of the exchange rate in the short run, while the share attributed to US shocks varies from 3 to 18 percents. Depending on the country, the joint contribution of US and SOE monetary disturbances to the short-run volatility of the SOE/US RER ranges from 25 to 35 percents.

1.4 Robustness Checks

We perform two robustness checks. In the first, we re-estimate the six SOE-SVARs over the shorter sample period 1992:Q1 - 2008:Q3 to exclude episodes of unconventional monetary policy and binding zero-lower-bound. In this exercise, we do not use any shadow rates to measure the stance of SOE and US monetary policy. In the second robustness check, we estimate a 9-variable SOE-SVAR model for the six SOEs. This larger model includes an SOE corporate credit spread. Due to limited data availability, the estimation period is restricted to 2000:Q1 - 2019:Q4. We impose an additional sign restriction which requires that the SOE central bank leans against the SOE credit spread. It is our hunch that adding a sign-restriction will sharpen our identification that correctly characterize the systematic component of SOE monetary policy. Caldara and Herbst (2019) characterize the US monetary policy component by including credit spread, we test if the same argument is true for SOE central banks.

1.4.1 Excluding ZLB episodes

Figure 1.7 plots the IRFs of SOE variables to a one standard deviation SOE contractionary monetary policy surprise, when using the sample period 1992:Q1 - 2008:Q3 in estimation.[25] The dashed lines are the posterior median IRFs based on the sample

[25] For the US, Sweden and the UK, we replace the shadow rate with the 3-month interbank rate.

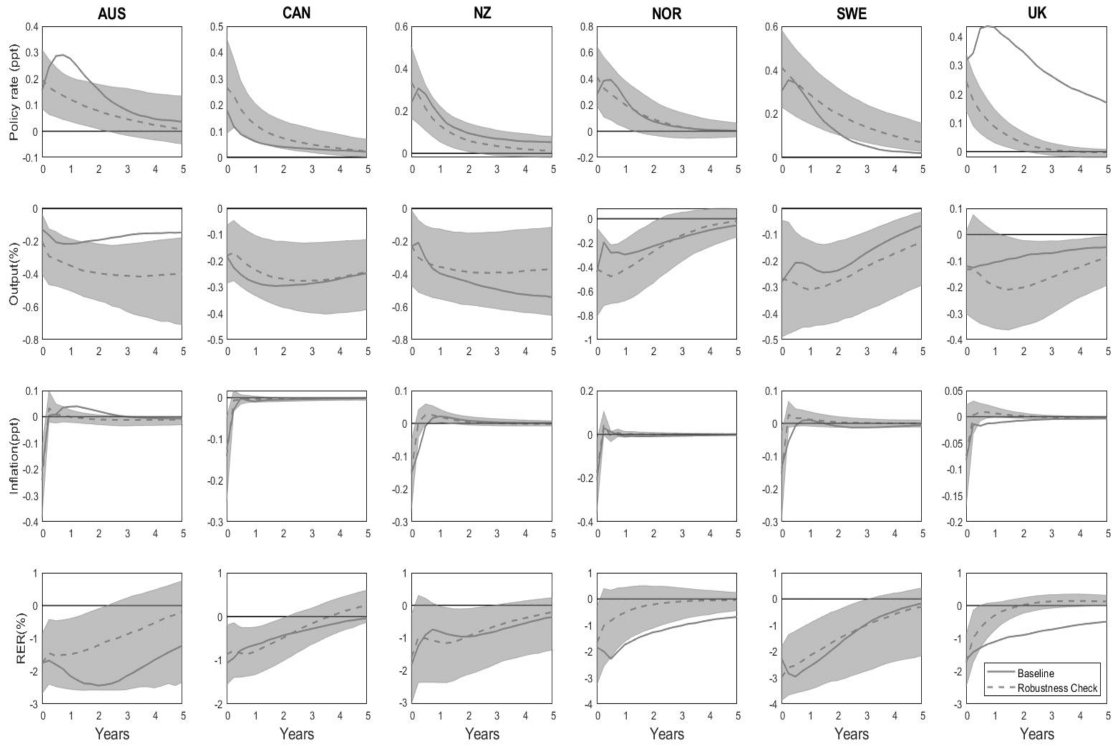


Figure 1.7: IRFs of SOE variables to a one standard deviation SOE contractionary monetary shock identified through block exogeneity and Restrictions 1 to 4, using the sample period 1992:Q1 to 2008:Q3 in estimation. *Note:* The dashed lines depict the point-wise posterior median responses for the sample period 1992:Q1 to 2008:Q3. The shaded regions represent the corresponding 68% equal-tailed point-wise posterior probability bands. The solid lines depict the posterior median responses for the full sample period 1992:Q1 to 2019:Q4.

1992:Q1 - 2008:Q3, and the shaded bands are the associated 68% posterior probability bands. The solid lines, instead, are the posterior median IRFs from the baseline estimation over the full sample 1992:Q1 - 2019:Q4. Figure 1.8 plots the deviations from UIP based on the estimation period 1992:Q1 - 2008:Q3. Looking at Figure 1.7 and Figure 1.8, we conclude that our main results are qualitatively and quantitatively robust to excluding the ZLB episodes.

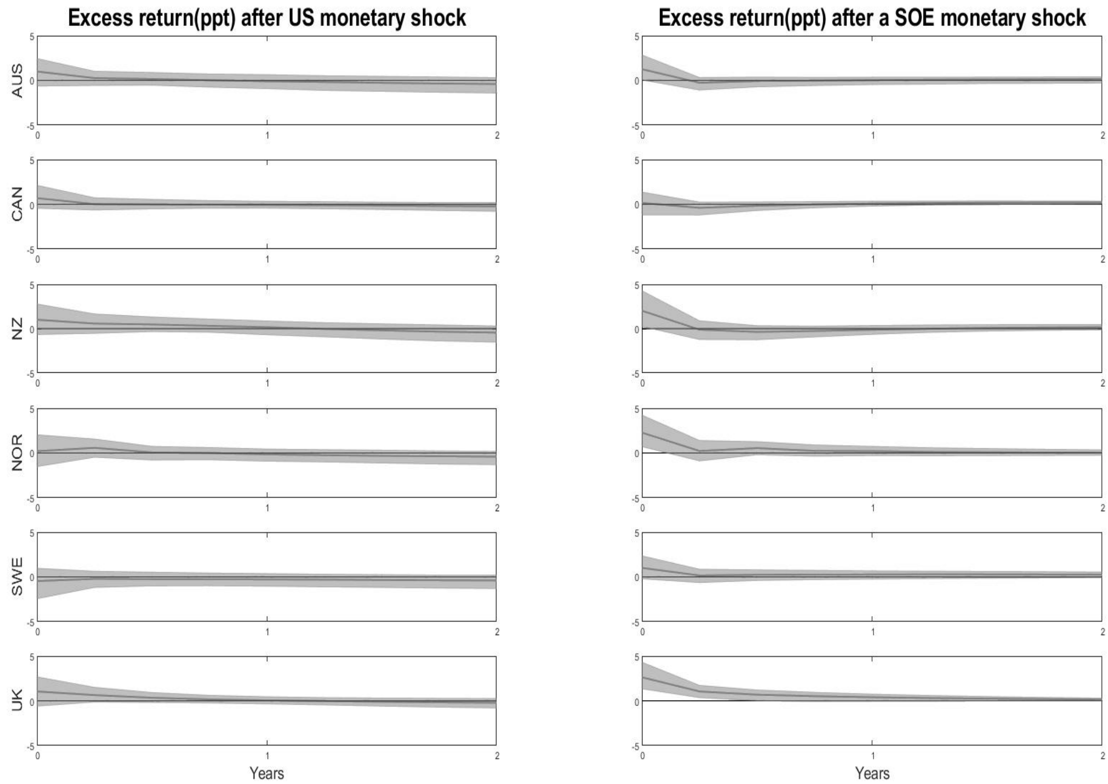


Figure 1.8: Deviations from UIP when using the shorter estimation sample period 1992:Q1 to 2008:Q3. *Note:* In each panel, the solid line represents the point-wise posterior median estimates of excess returns. The shaded areas are the 68% posterior probability intervals.

1.4.2 Including the SOE credit spread

We extend our baseline model to include an additional variable, a SOE corporate credit spread, in the domestic block. Due to data availability, the sample period used in estimation is 2000:Q1 to 2019:Q4.[26,27] Our motivation comes from Caldara and Herbst (2019) and Beckers et al. (2020), who show the importance of including a corporate credit spread measure in the systematic component of US and Australian monetary policy, respectively. For Canada, New Zealand, Sweden and the UK, we construct the SOE credit spread as the difference between the Standard & Poor's investment grade

[26] We set the lag order $p = 1$.

[27] We do not consider this specification as our baseline model due to the limited availability of SOE credit spread data. Due to less observations and an additional variable, we loose a lot of degrees of freedom. Thus, the IRFs are less precisely estimated.

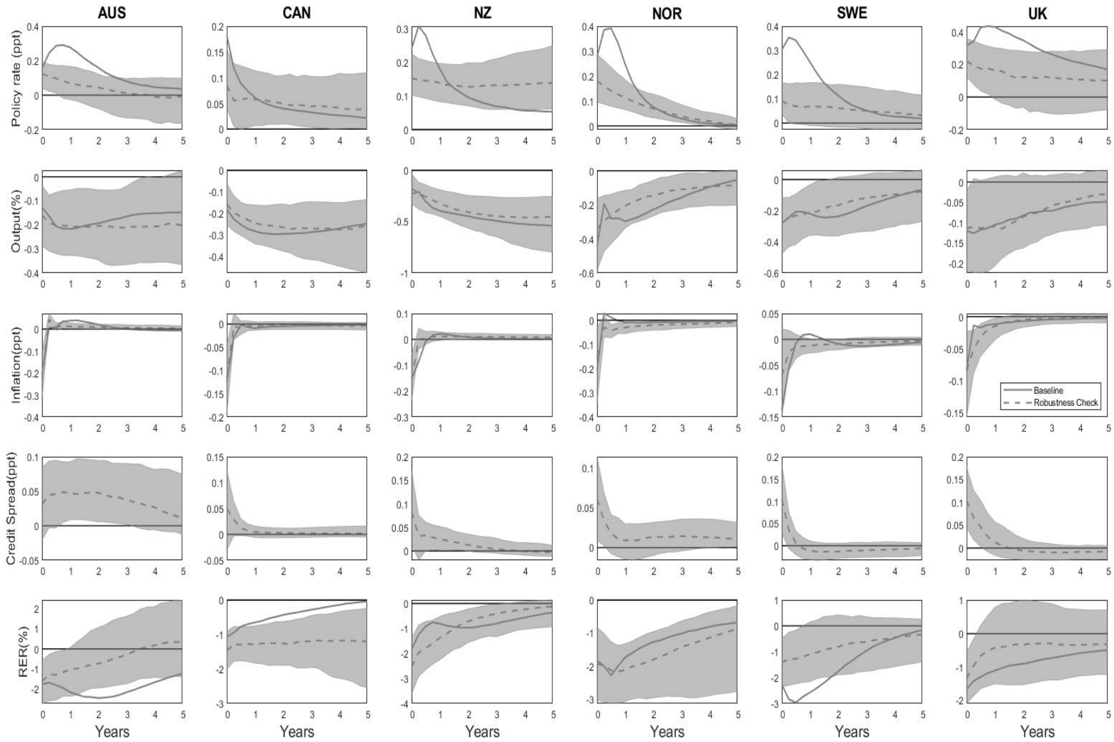


Figure 1.9: IRFs of SOE variables to a SOE contractionary monetary shock identified using block exogeneity and Restrictions 1 to 5. *Note:* Dashed lines depict pointwise posterior median IRFs when the SOE credit spread is included in the domestic block and Restriction 5 is imposed. The shaded regions represent the associated 68% posterior probability bands. The solid lines correspond to the baseline posterior median IRFs.

corporate bond yield and the government bond yield. For Australia, we use the credit spread measure by constructed by Beckers et al. (2020).[28] We extend our baseline identification scheme, based on block exogeneity and Restriction 1 to 4, by formulating a fifth restriction.

Restrictions 5. The contemporaneous response of the SOE policy rate to the domestic credit spread is negative: $\psi_{cs} < 0$.

Restriction 5 means that SOE central bank, guided by a concern for financial stability, typically cuts its policy rate in response to an increase in the domestic credit

[28] Special thanks to Olav Syrstad for help with the Norwegian credit spread data. See Appendix A for further details on SOE credit spread data.

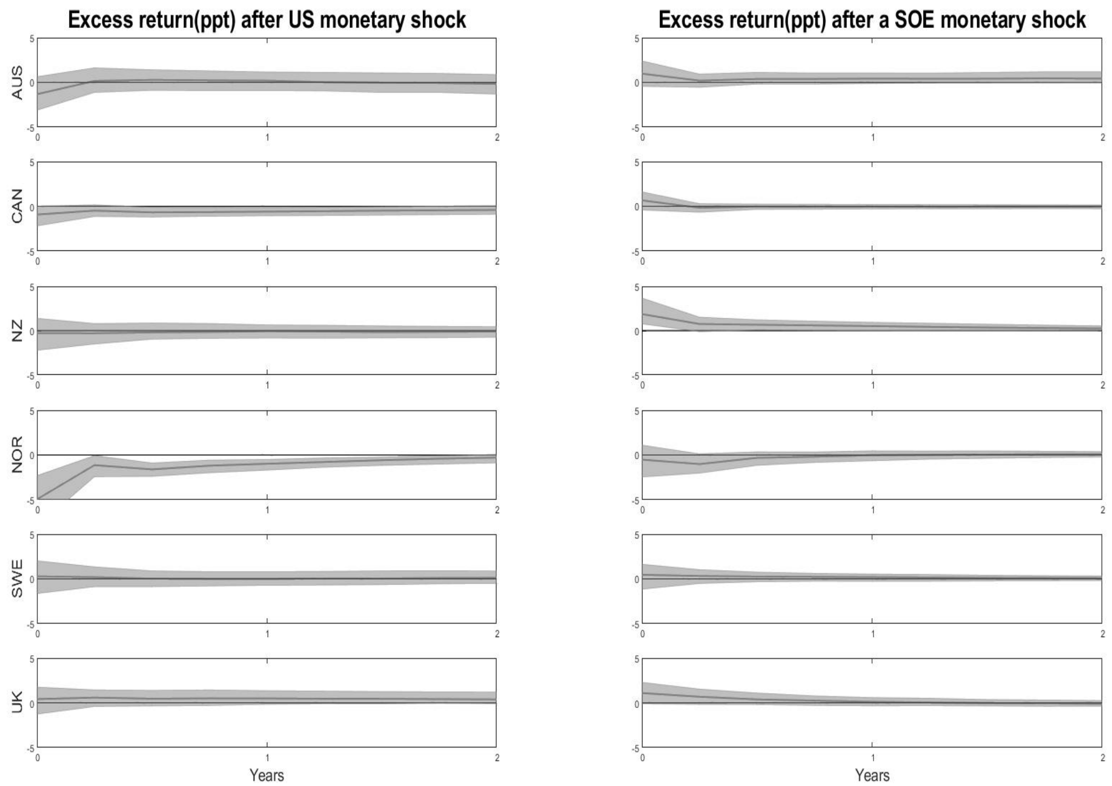


Figure 1.10: Deviations from UIP for Robustness check 2 (SOE credit spread). *Note:* Solid lines represent point-wise posterior median estimates of excess returns. Shaded areas are the 68% posterior probability intervals.

spread. The SOE monetary policy rule now reads:

$$
\begin{aligned}
r_t = & \underbrace{-a_{0,85}^{-1}a_{0,15}}_{unrestricted} y_t^* \underbrace{-a_{0,85}^{-1}a_{0,25}}_{unrestricted} \pi_t^* \underbrace{-a_{0,85}^{-1}a_{0,35}}_{unrestricted} cs_t^* \underbrace{-a_{0,85}^{-1}a_{0,45}}_{unrestricted} r_t^* \\
& \underbrace{-a_{0,85}^{-1}a_{0,55}}_{\psi_y>0} y_t \underbrace{-a_{0,85}^{-1}a_{0,65}}_{\psi_\pi>0} \pi_t \underbrace{-a_{0,85}^{-1}a_{0,75}}_{\psi_{cs}<0} cs_t \underbrace{-a_{0,85}^{-1}a_{0,95}}_{\psi_e>0} rer_t \underbrace{+a_{0,85}^{-1}}_{\sigma} \epsilon_{5,t}.
\end{aligned}
\tag{1.19}
$$

Figure 1.9 plots the IRFs (posterior median and 68% bands) of SOE variables to a contractionary SOE monetary policy shock for the second robustness check and compares them against the baseline posterior median. Figure 1.10 reports the deviations from UIP for the second robustness check. Keeping in mind the different sample periods used in baseline and in the second robustness check, Figure 1.9 and Figure 1.10 suggest that our main findings seem qualitatively robust to incorporating an SOE credit spread and imposing Restriction 5 (*the SOE central bank leans against the domestic credit spread*).

1.5 Conclusion

In this chapter, we estimate set-identified SVAR models for six small open economies to investigate the effects of domestic and US monetary policy shocks on the real exchange rate. Particularly, we revisit the Dornbusch (1976) overshooting hypothesis, which says that in response to a contractionary monetary shock the exchange rate appreciates instantaneously before gradually depreciating. On the other hand, some empirical literature finds exchange rate responses that are inconsistent with Dornbusch's overshooting mechanism. Two types of puzzling responses emerge: (i) an exchange rate puzzle which says that the exchange rate depreciates on impact rather than appreciating and (ii) a delayed overshooting puzzle which says that the exchange rate continues to appreciate rather than depreciating after a contractionary monetary shock.

We use the state-of-the-art Bayesian estimation techniques recently proposed by Arias et al. (2019) to estimate six SVAR models for Australia, Canada, New Zealand, Norway, Sweden and the United Kingdom. The identification scheme developed by Arias et al. (2019) applies a combination of exclusion and sign-restrictions directly on the structural parameters. We build on the work of Arias et al. (2019) by implementing a block exogeneity structure through a combination of exclusion restrictions and tight priors on the contemporaneous and lagged parameters, respectively. Further, we apply sign-restrictions to simultaneously identify the systematic component of monetary polices for the US and SOEs. Our agnostic identification scheme preserves the contemporaneous interaction between the exchange rate and the domestic policy rates, while leaving the response of the exchange rate to domestic and US monetary shocks unrestricted.

We find no evidence of exchange rate puzzle and delayed overshooting puzzle. In the six SOEs, a domestic contractionary monetary surprise results in an on impact appreciation of the exchange rate, followed by a gradual depreciation. Moreover, we find that a tightening of US monetary policy causes an instantaneous appreciation followed by a depreciation of the US dollar. Our findings support the view that delayed overshooting is an artefact of incorrect identifying restrictions. Specifically, we show that the response of SOE central banks to exchange rate fluctuations is the key to address the delayed overshooting puzzle. Moreover, we find little evidence of the forward discount puzzle: the responses of exchange rates to both SOE and US monetary shocks are broadly consistent with UIP, and thus with Dornbusch's overshooting hypothesis. Finally, the six SOEs and US monetary shocks explain about 20 and 10 percents of the short-run exchange rate volatility, respectively.

Appendix A

Appendix to Chapter 1

A.1 Data Sources

The dataset spans from 1992:Q1 to 2019:Q4.

United States

- **Real Gross Domestic Product** (Billions of Chained 2012 Dollars, Seasonally Adjusted Annual Rate)
  - Source: FRED Economic Data (GDPC1)
- **Consumer Price Index: All Items for the United States** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (USACPIALLMINMEI)
- **Federal Funds Effective Rate** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (FEDFUNDS)
- **Shadow rate** (Percent)
  - Source: Wu and Xia (2016)
- **Moody's Seasoned Baa Corporate Bond Yield Relative to Yield on 10-Year Treasury Constant Maturity** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (BAA10YM)

Australia

- **Real Gross Domestic Product for Australia** (Domestic Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (NGDPRSAXDCAUQ)
- **Consumer Price Index: All Items: Total: Total for Australia** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (AUSCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Australia** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01AUQ156N)
- **Exchange rate** [**]
  - Source: Reserve Bank of Australia (Refinitiv Datastream, AUUSDSP)
- **Credit spread**
  - Description: Credit market spread between Australian large business variable lending rate and 3-month Bank Accepted Bill (BAB) rate.
  - Source: Reserve Bank of Australia
    * https://www.rba.gov.au/publications/rdp/2020/2020-01/supplementary-information.html

Canada

- **Real Gross Domestic Product for Canada** (Domestic Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (NGDPRSAXDCCAQ)
- **Consumer Price Index: All Items: Total: Total for Canada** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (CANCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Canada** (Percent, Not Seasonally Adjusted)

[**] For each SOE, we calculate real exchange rate from the nominal exchange rate and the US and domestic price levels, such that, $e_t = s_t + p_t^* - p_t$; where e_t and s_t are the logs of real and nominal exchange rates, respectively and p_t^* and p_t are the logs of US and domestic consumer price indices, respectively.

  - Source: FRED Economic Data (IR3TIB01CAQ156N)

- **Exchange rate**
  - Source: Bank of Canada (Refinitiv Datastream, CNXRUSD)

- **Credit spread**
  - Description: Credit spread is calculated as the yield in the S&P Canada Investment Grade Corporate Bond Index (Code: SPFICAV) less the Canada, Government Bond Yield: 3 Year Benchmark (End Month)(Code: CNB14068; CANSIM-Statistic Canada).

New Zealand

- **Production-based gross domestic product (GDP)** (Real, NZD, Seasonally Adjusted)
  - Source: Statistics New Zealand (GDP06.Q.QT0.rs)

- **Consumer Price Index: All Items for New Zealand** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (NZLCPIALLQINMEI)

- **3-Month or 90-day Rates and Yields: Interbank Rates for New Zealand** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01NZQ156N)

- **Exchange rate**
  - Source: Reserve Bank of New Zealand (EXR.MS11.D06)

- **Credit spread**
  - Description: Credit spread is calculated as the yield in the S&P New Zealand Investment Grade Corporate Bond Index (Code: SPNZICZ) less the New Zealand Government Bond Yield, 2 Years (Code:NZGBY2Y; Reserve Bank of New Zealand).

Norway

- **Real Gross Domestic Product for Norway** (Millions of Chained 2010 National Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (CLVMNACSCAB1GQNO)

- **Consumer Price Index: All Items for Norway** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (NORCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Norway** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01NOQ156N)
- **Exchange rate**
  - Source: Norges Bank (Refinitiv Datastream, NWXRUSD)
- **Credit spread**
  - Description: Risk-premium new 5-years bond. Index made up of Norwegian industrial issuers. Percentage points over 3 month NIBOR. (RPREM.IND.M060)

Sweden

- **Real Gross Domestic Product for Sweden** (Millions of Chained 2010 National Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (CLVMNACSCAB1GQSE)
- **Consumer Price Index: All Items for Sweden** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (SWECPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for Sweden** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01SEQ156N)
- **Shadow rate** (Percent)
  - Source: De Rezende and Ristiniemi (2023)
- **Exchange rate**
  - Source: Sveriges Riksbank Bank (Refinitiv Datastream, SDXRUSD)
- **Credit spread**
  - Description: Credit spread is calculated as the yield in the S&P Sweden Investment Grade Corporate Bond Index (Code: SPSEICR) less the Swedish Government Bond, maturity 2 years (Code: SEGVB2Y; Sveriges Riksbank, Refinitiv Datastream).

United Kingdom

- **Real Gross Domestic Product for United Kingdom** (Millions of Chained 2010 National Currency, Seasonally Adjusted)
  - Source: FRED Economic Data (CLVMNACSCAB1GQUK)
- **Consumer Price Index of All Items in the United Kingdom** (Index 2015=100, Not Seasonally Adjusted)
  - Source: FRED Economic Data (GBRCPIALLQINMEI)
- **3-Month or 90-day Rates and Yields: Interbank Rates for the United Kingdom** (Percent, Not Seasonally Adjusted)
  - Source: FRED Economic Data (IR3TIB01GBQ156N)
- **Shadow rate** (Percent)
  - Source: https://sites.google.com/view/jingcynthiawu/shadow-rates
- **Exchange rate**
  - Source: Bank of England (Refinitiv Datastream, UKXRUSD)
- **Credit spread**
  - Description: Credit spread is calculated as the yield in the S&P U.K. Investment Grade Corporate Bond Index (Code: SPUKICG) less the Long-Term Government Bond Yields: 10-year: Main (Including Benchmark) for the United Kingdom (Code: IRLTLT01GBQ156N; FRED Economic Data).

A.2 Baseline

A.2.1 IRFs to a one standard deviation SOE monetary policy shocks

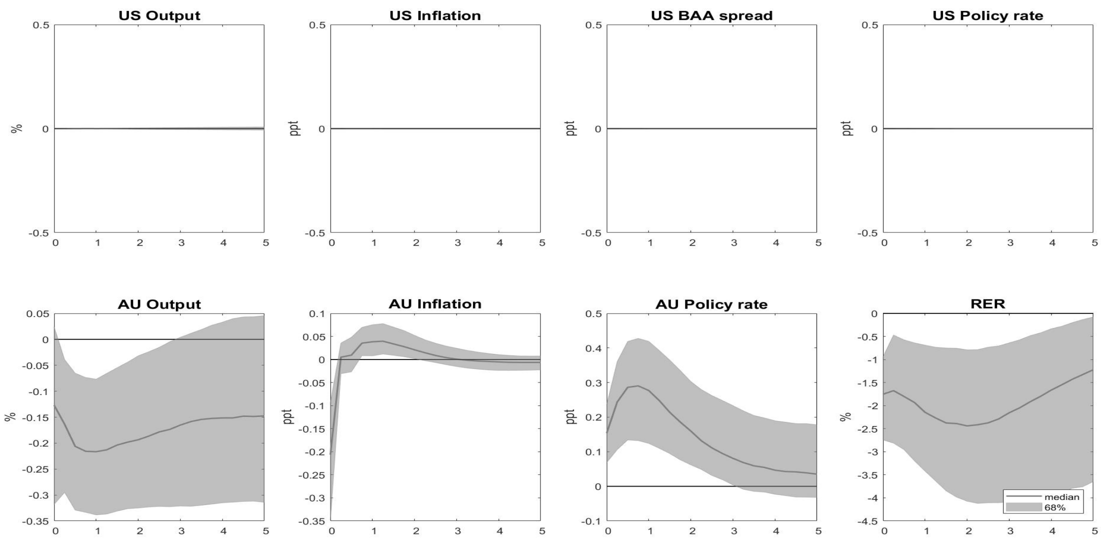


Figure A.1: Australia - IRFs to a one standard deviation contractionary monetary policy shock

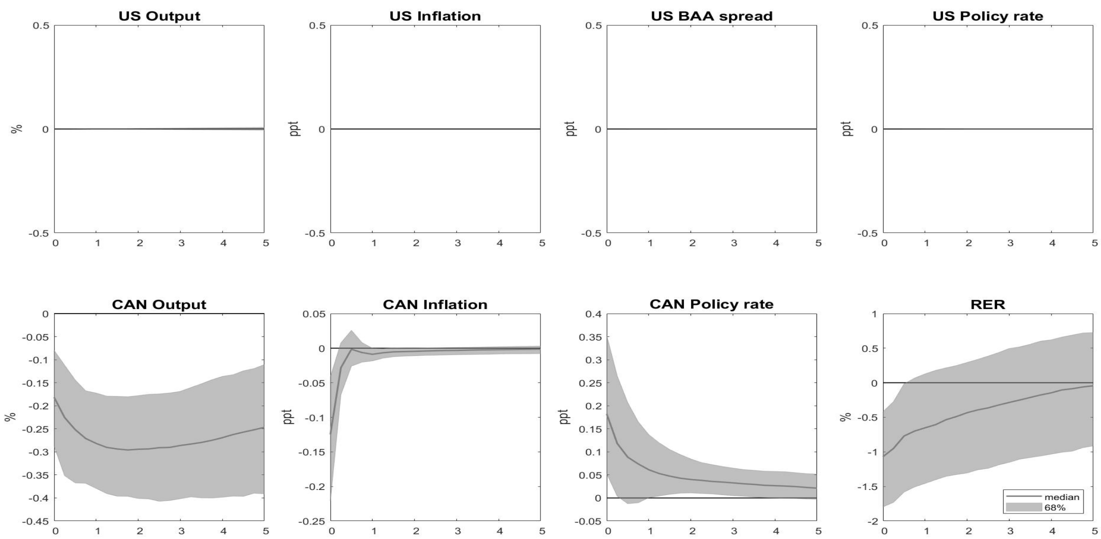


Figure A.2: Canada - IRFs to a one standard deviation contractionary monetary policy shock

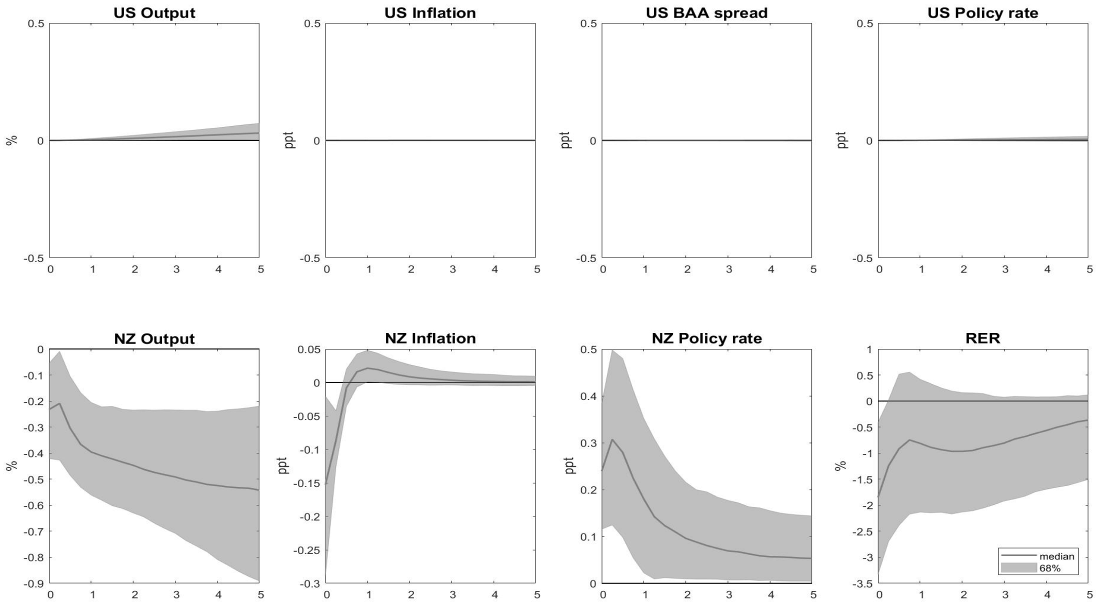


Figure A.3: New Zealand - IRFs to a one standard deviation contractionary monetary policy shock

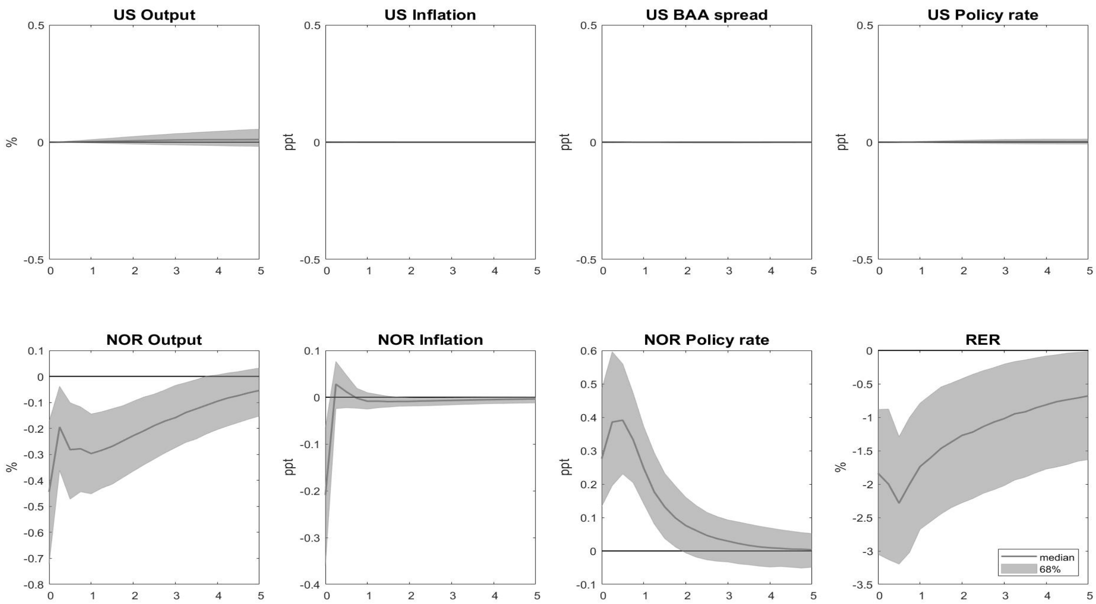


Figure A.4: Norway - IRFs to a one standard deviation contractionary monetary policy shock

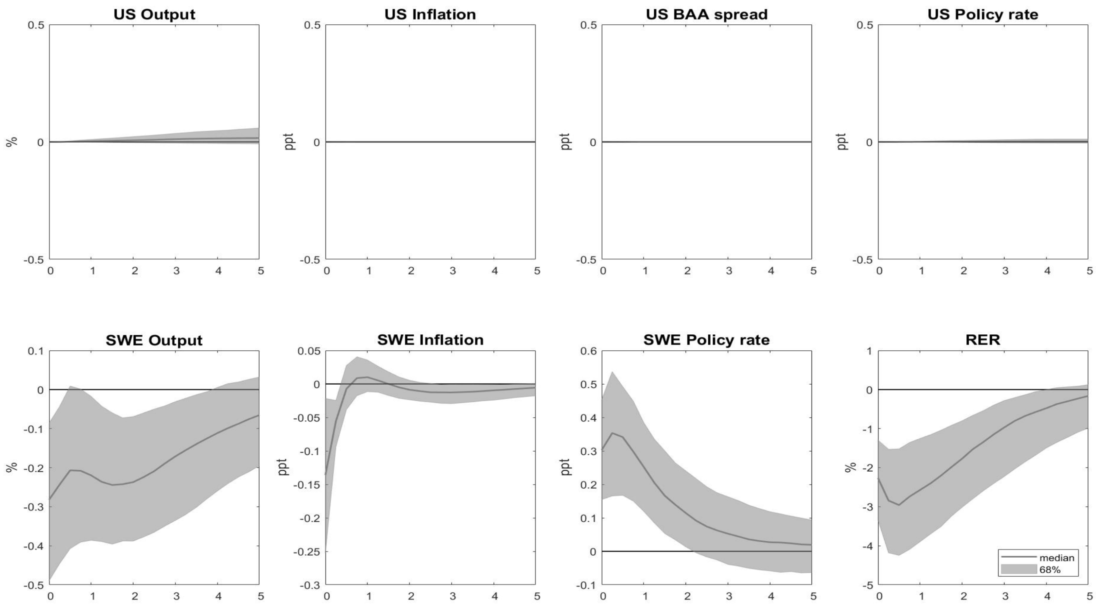


Figure A.5: Sweden - IRFs to a one standard deviation contractionary monetary policy shock

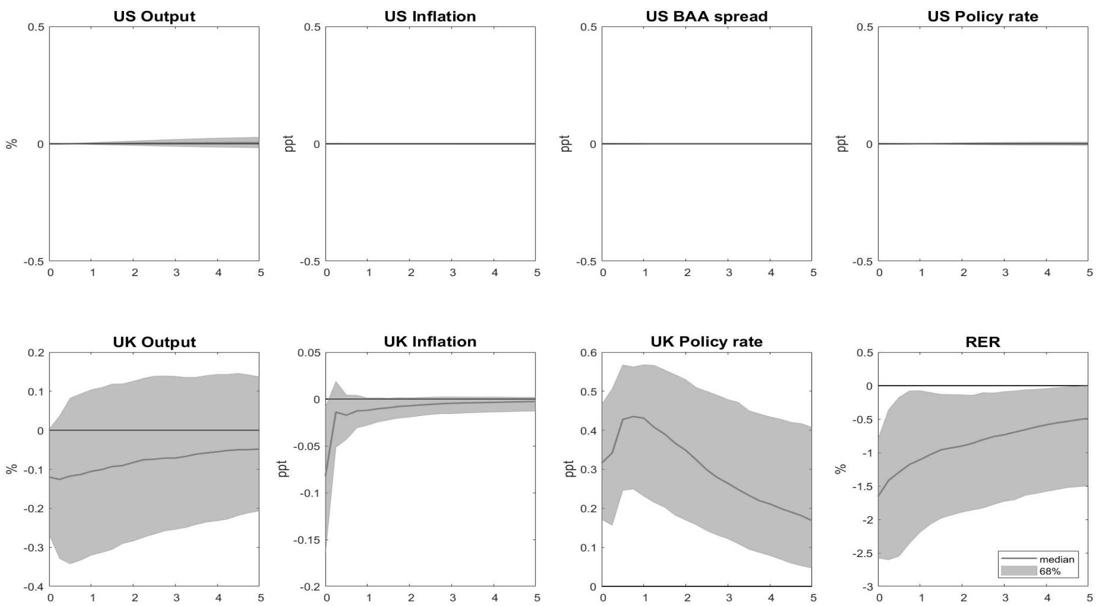


Figure A.6: UK - IRFs to a one standard deviation contractionary monetary policy shock

A.2.2 IRFs to a one standard deviation US monetary policy shocks

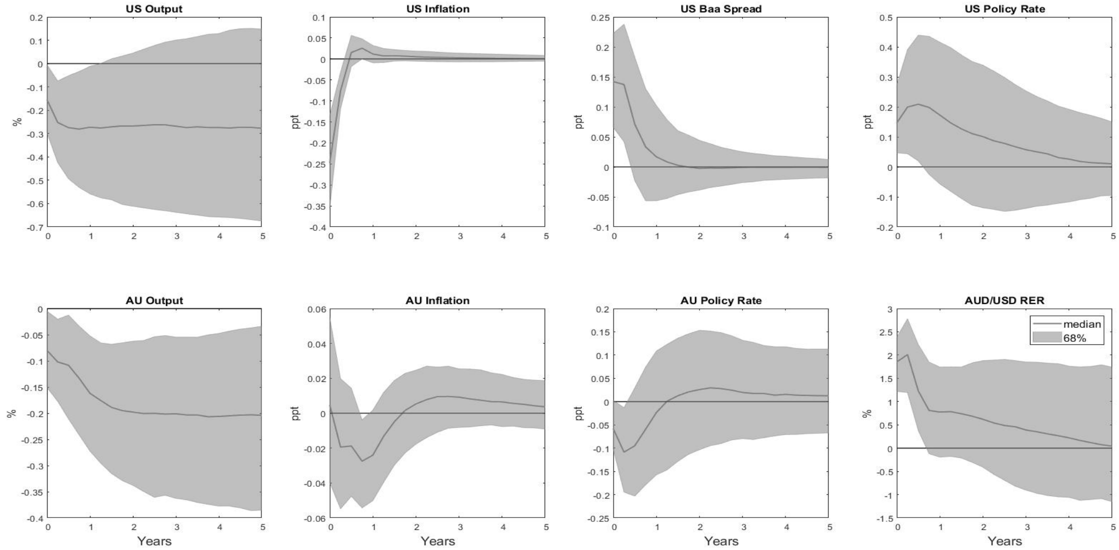


Figure A.7: Australia - IRFs to a one standard deviation US contractionary monetary policy shock

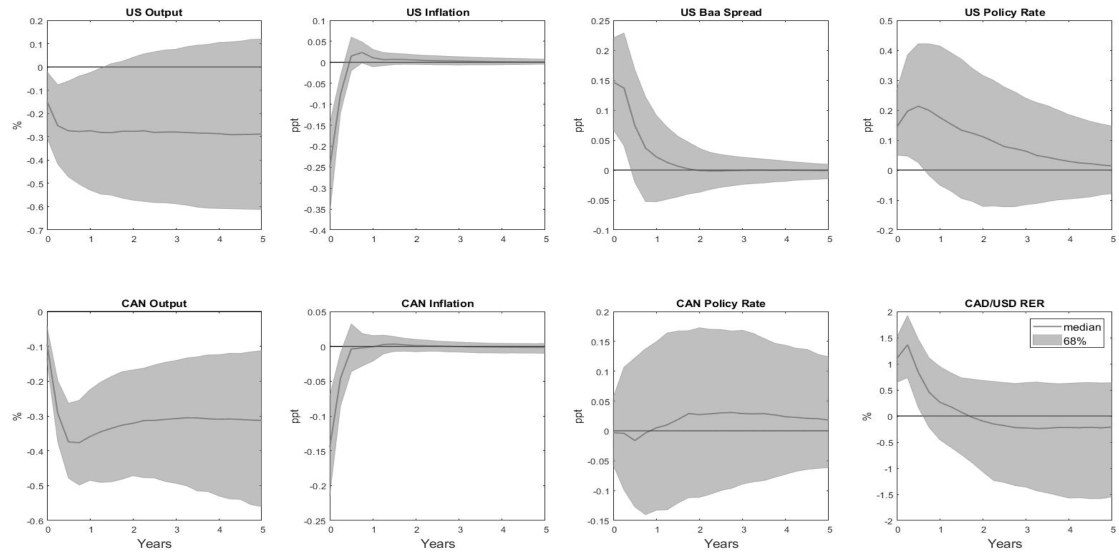


Figure A.8: Canada - IRFs to a one standard deviation US contractionary monetary policy shock

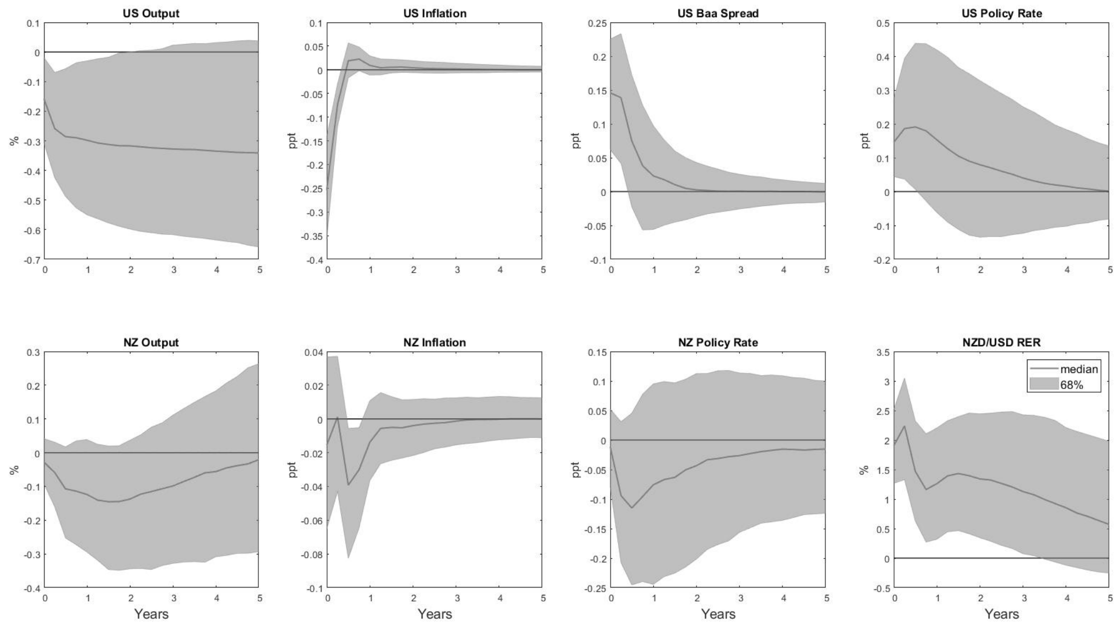


Figure A.9: New Zealand - IRFs to a one standard deviation US contractionary monetary policy shock

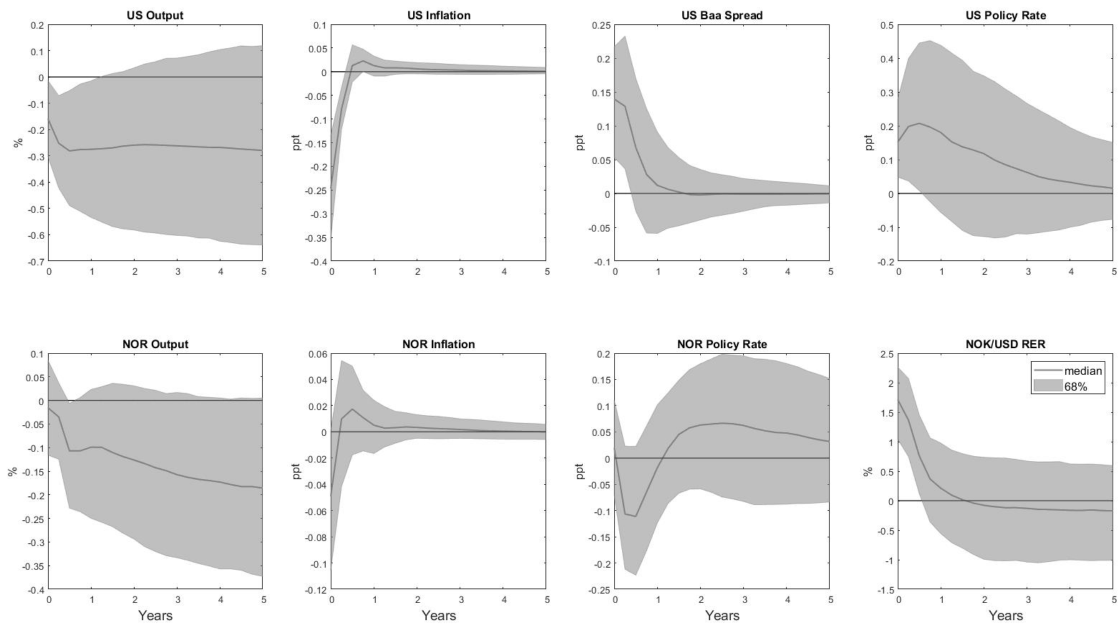


Figure A.10: Norway - IRFs to a one standard deviation US contractionary monetary policy shock

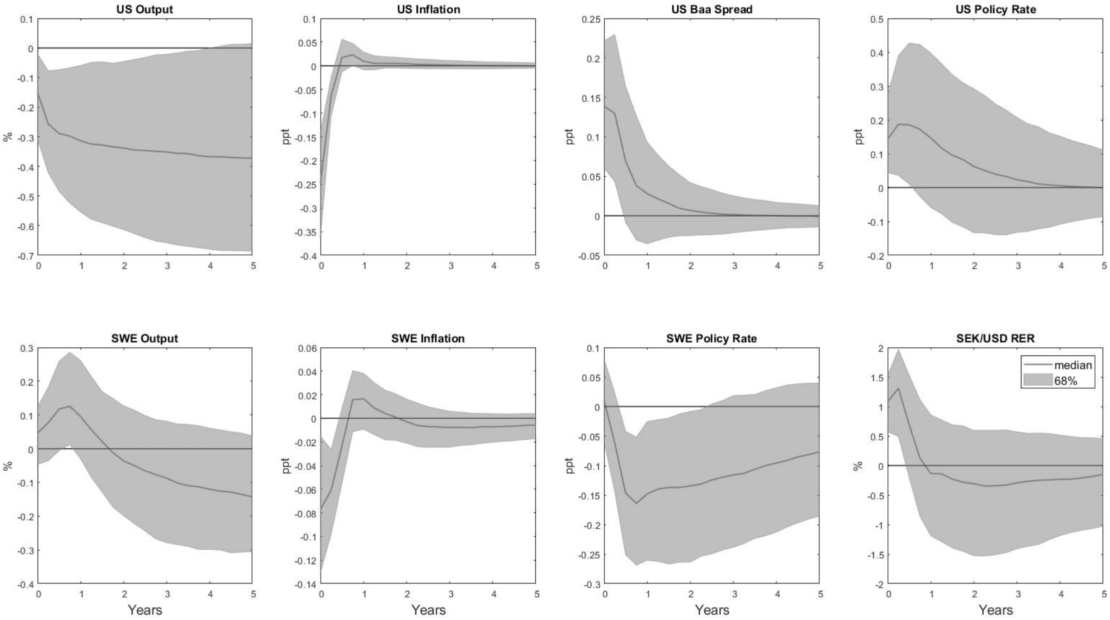


Figure A.11: Sweden - IRFs to a one standard deviation US contractionary monetary policy shock

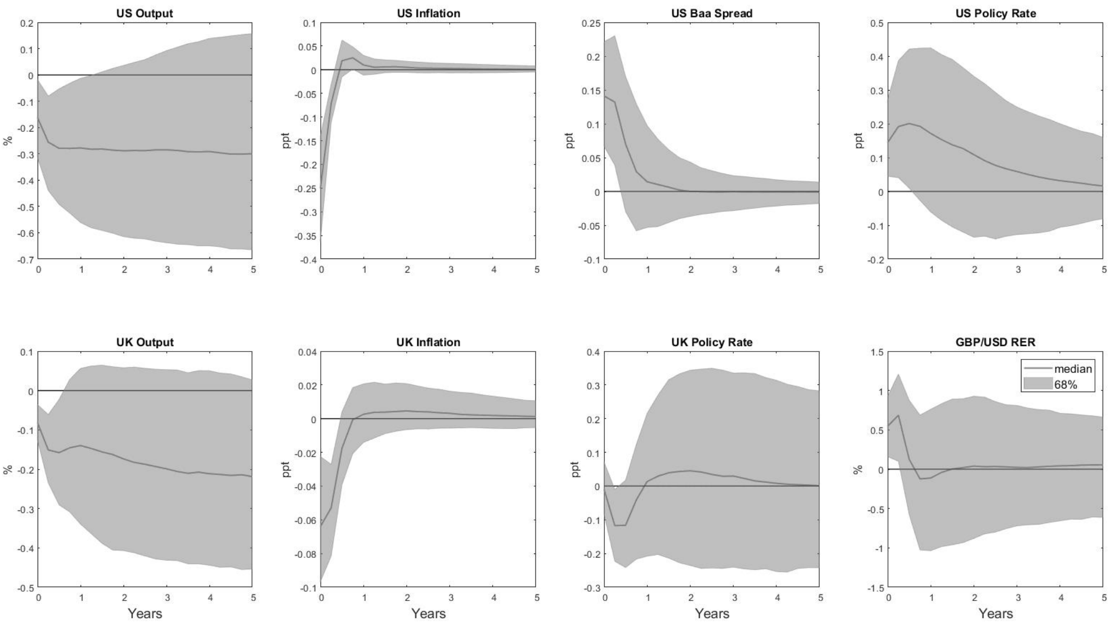


Figure A.12: UK - IRFs to a one standard deviation US contractionary monetary policy shock

A.2.3 Responses of US variables to US shocks

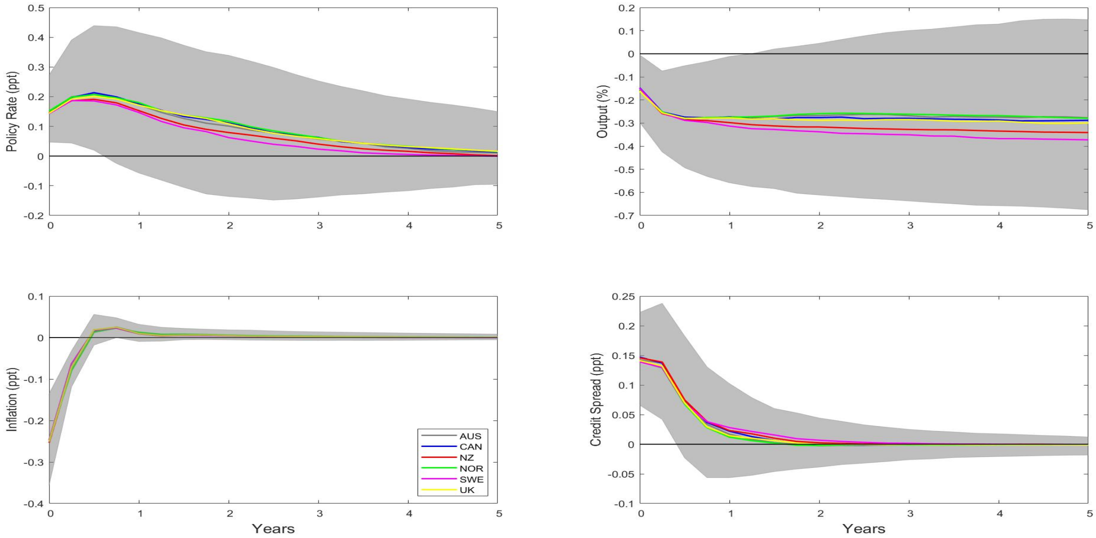


Figure A.13: IRFs to a one standard deviation US contractionary monetary policy shock for six SOEs

A.2.4 IRFs to a one standard deviation SOE monetary policy shocks with/ out $\psi_e > 0$

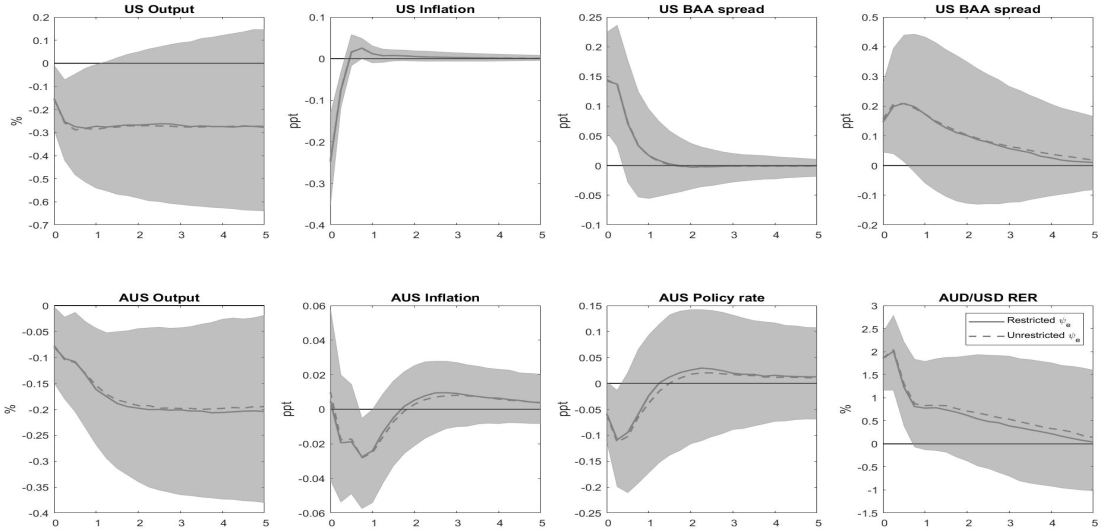


Figure A.14: Australia - IRFs to a one standard deviation contractionary monetary policy shock

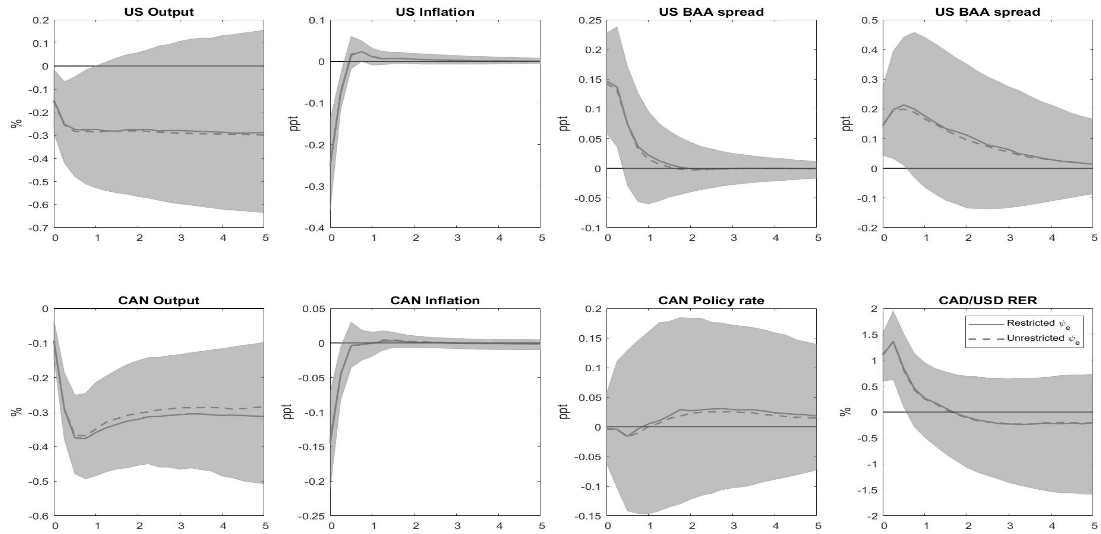


Figure A.15: Canada - IRFs to a one standard deviation contractionary monetary policy shock

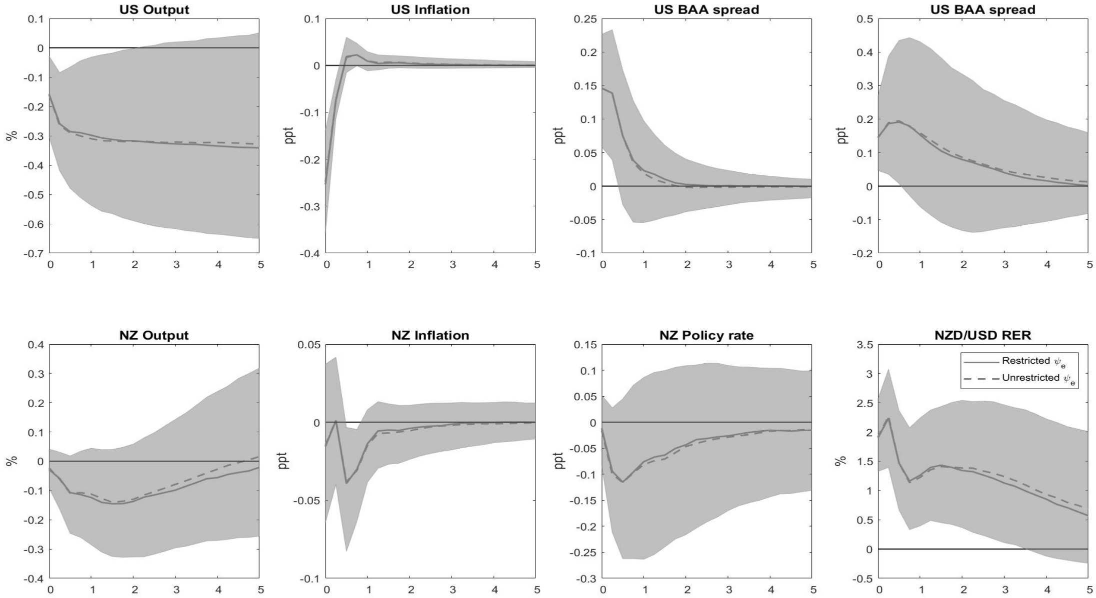


Figure A.16: New Zealand - IRFs to a one standard deviation contractionary monetary policy shock

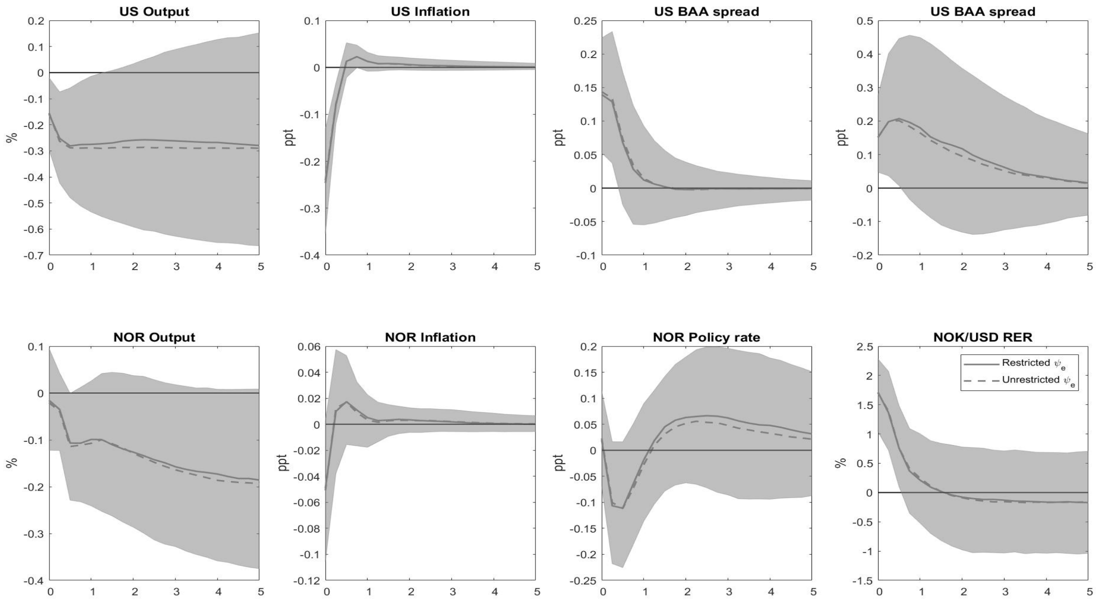


Figure A.17: Norway - IRFs to a one standard deviation contractionary monetary policy shock

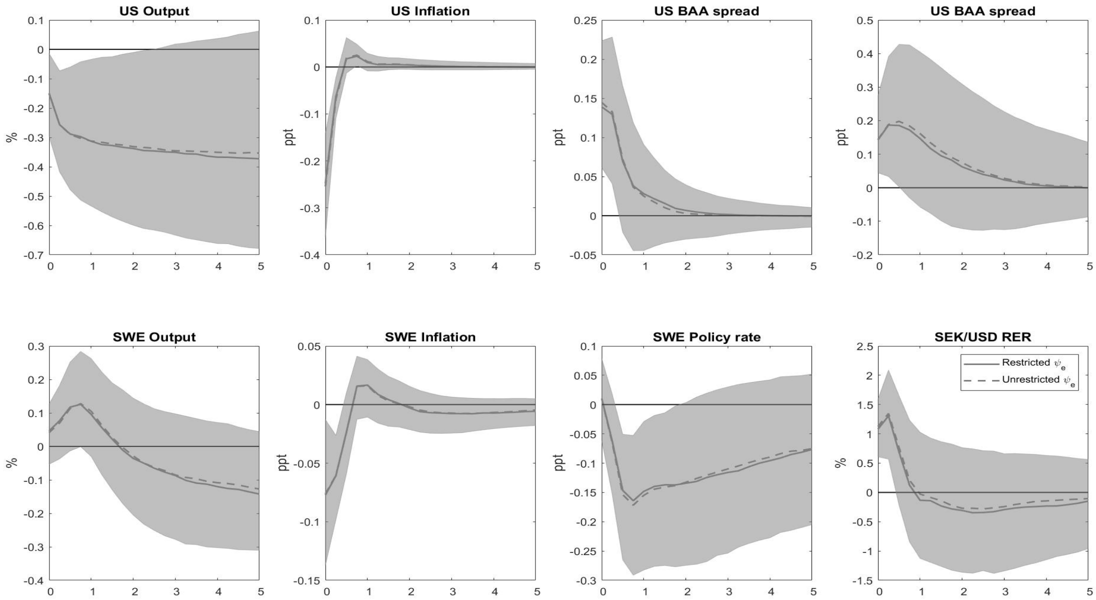


Figure A.18: Sweden - IRFs to a one standard deviation contractionary monetary policy shock

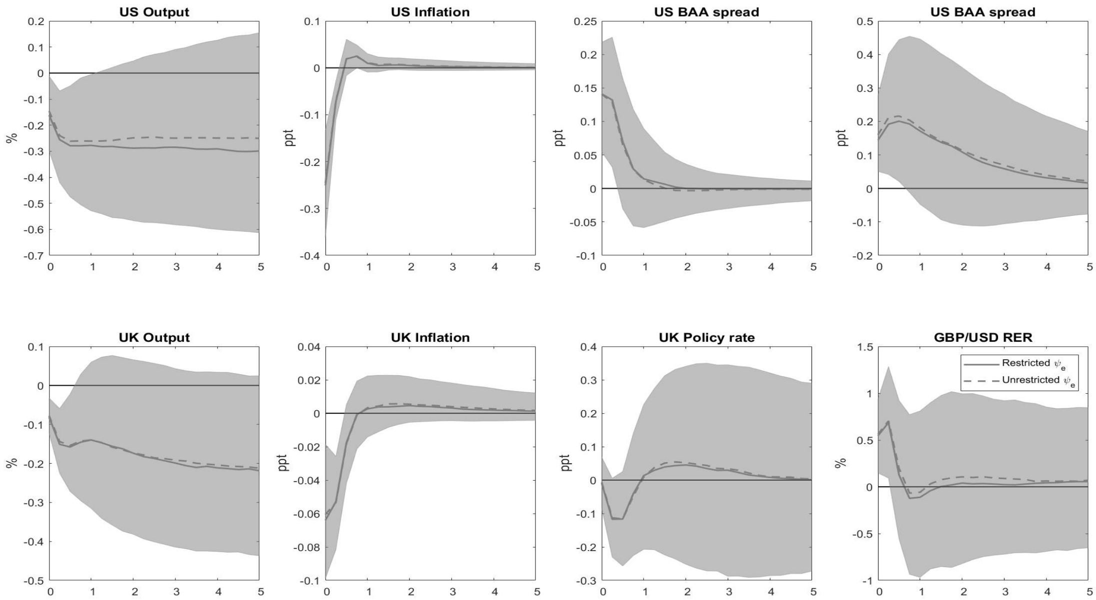


Figure A.19: UK - IRFs to a one standard deviation contractionary monetary policy shock

A.2.5 IRFs to a one standard deviation SOE monetary policy shocks with/out $\psi_{cs^*} < 0$

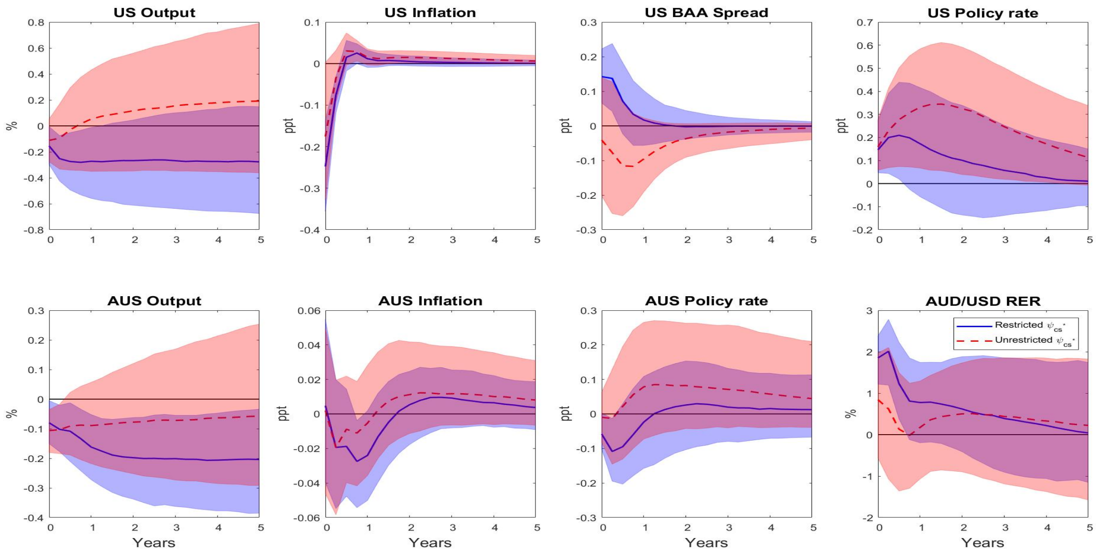


Figure A.20: Australia - IRFs to a one standard deviation contractionary monetary policy shock

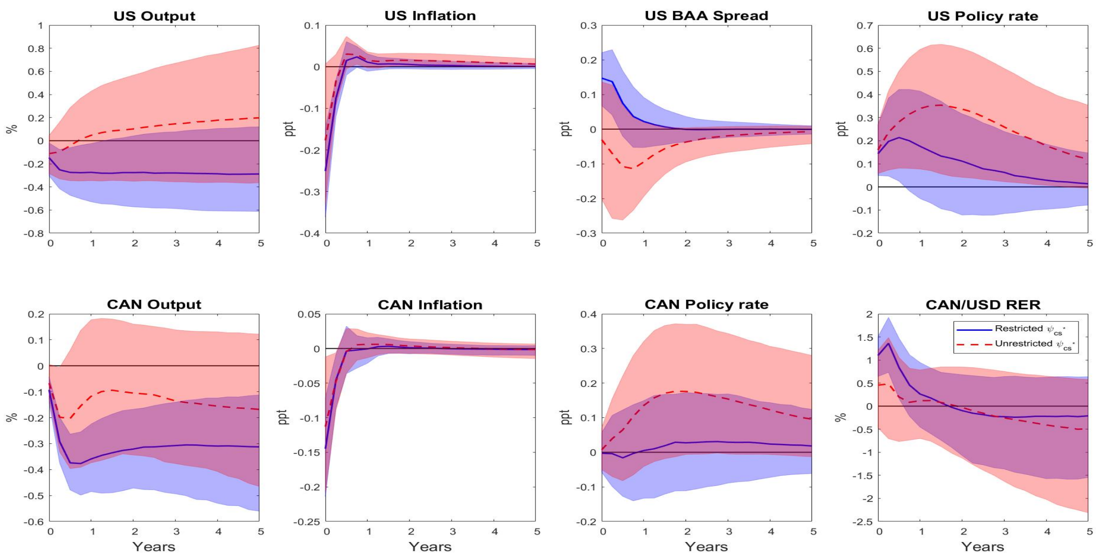


Figure A.21: Canada - IRFs to a one standard deviation contractionary monetary policy shock

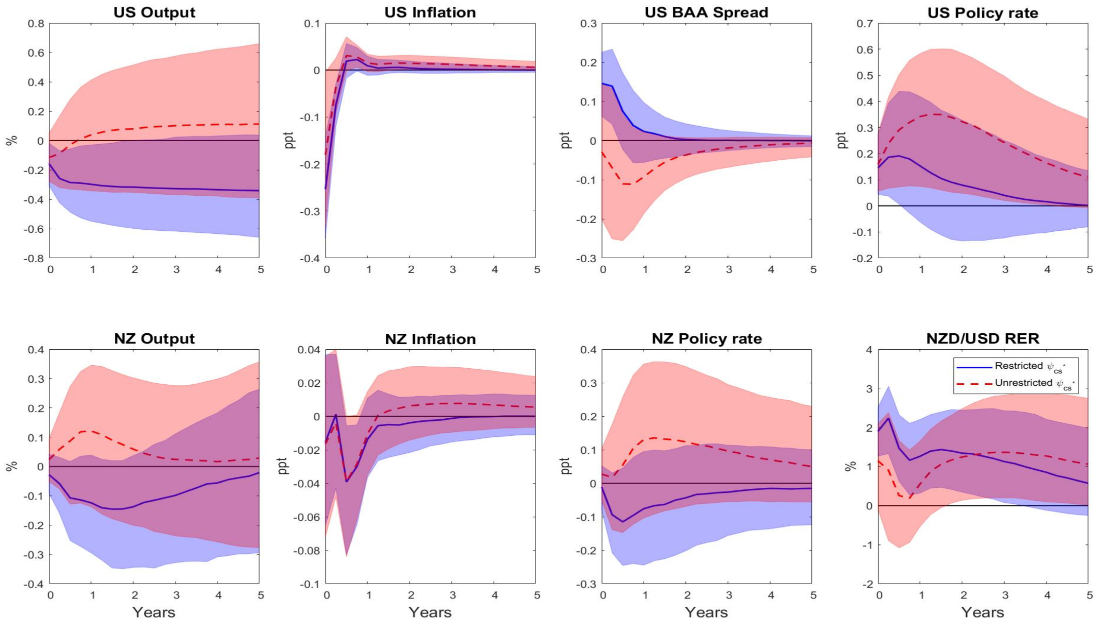


Figure A.22: New Zealand - IRFs to a one standard deviation contractionary monetary policy shock

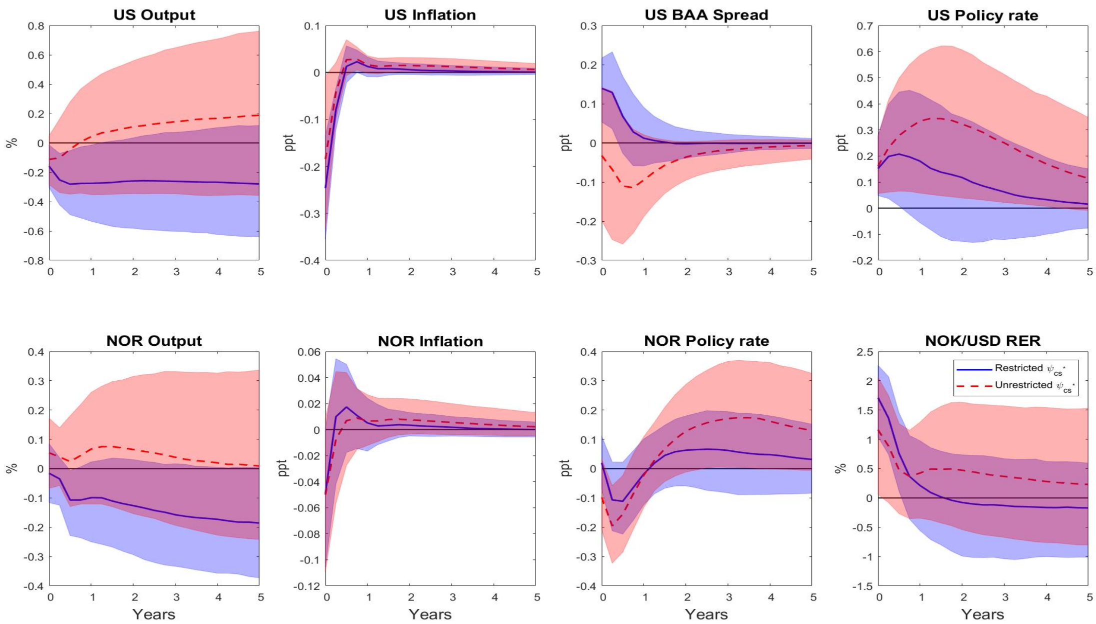


Figure A.23: Norway - IRFs to a one standard deviation contractionary monetary policy shock

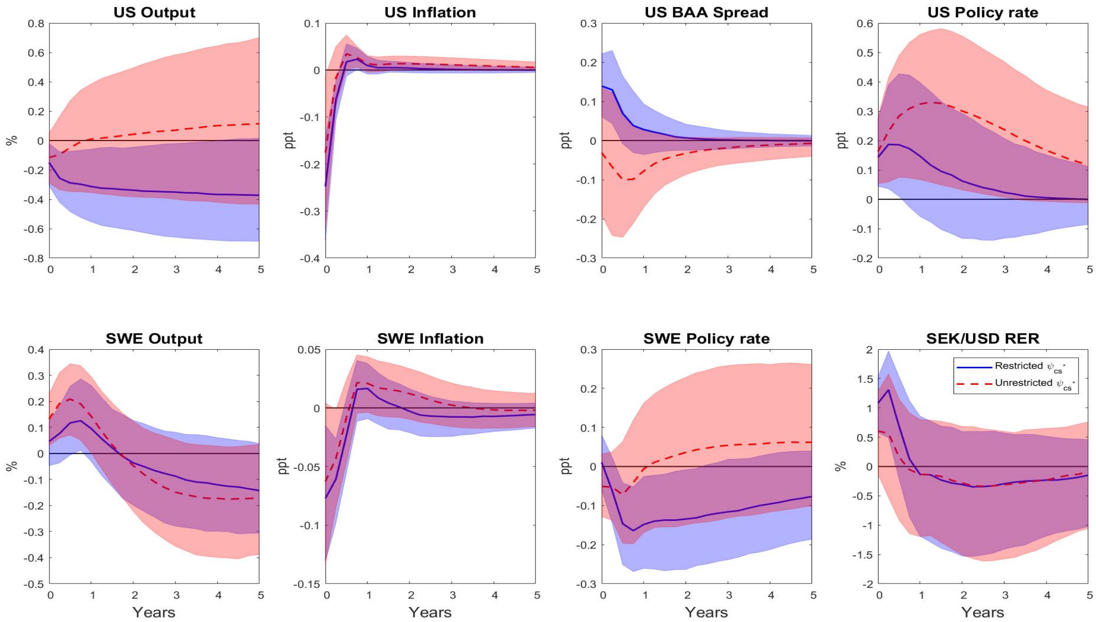


Figure A.24: Sweden - IRFs to a one standard deviation contractionary monetary policy shock

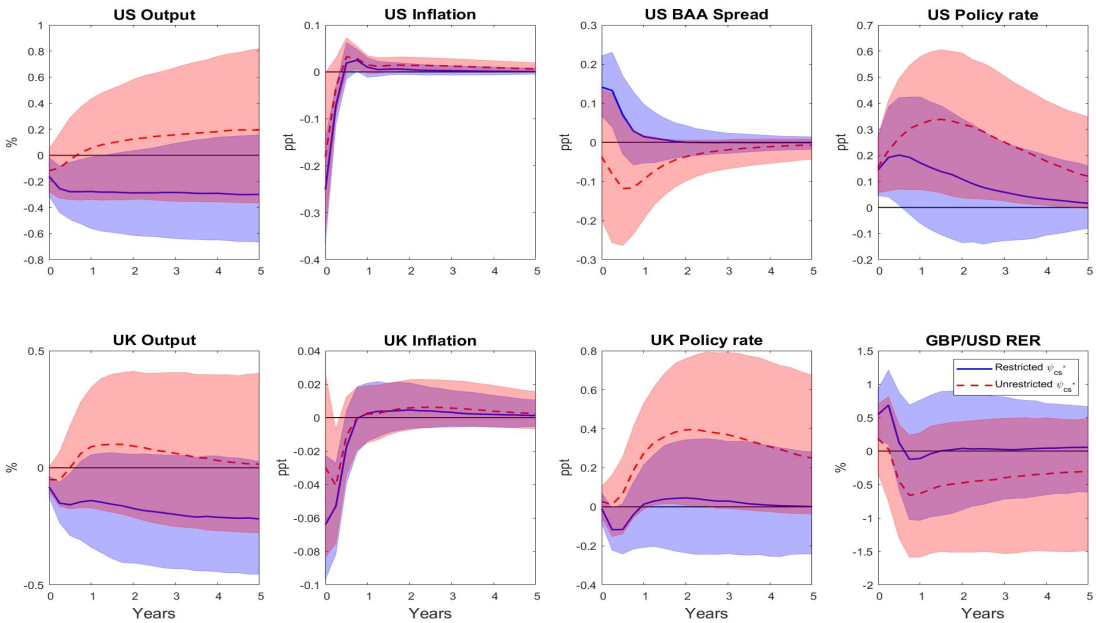


Figure A.25: UK - IRFs to a one standard deviation contractionary monetary policy shock

A.3 Robustness check # 1

A.3.1 IRFs to a one standard deviation SOE monetary policy shocks with/out $\psi_e > 0$

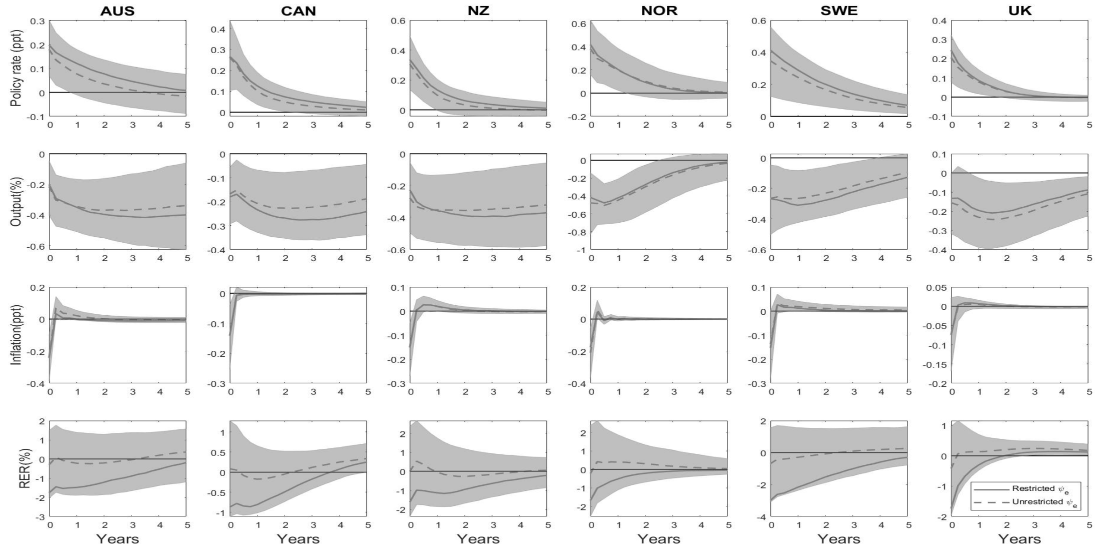


Figure A.26: IRFs to a one standard deviation contractionary monetary policy shocks for six SOEs

A.4 Baseline

A.4.1 Contemporaneous coefficients in SOE monetary policy equations.

	Coefficients		
SOEs	ψ_y	ψ_π	ψ_e
AUS	0.72 [0.24;2.20]	0.65 [0.17;2.75]	0.09 [0.02;0.35]
CAN	0.99 [0.24;3.10]	1.32 [0.37;3.63]	0.16 [0.05;0.49]
NZ	0.87 [0.24;3.99]	1.44 [0.36;5.49]	0.11 [0.03;0.46]
NOR	0.62 [0.16;2.73]	1.31 [0.34;5.80]	0.16 [0.04;0.69]
SWE	0.94 [0.21;4.01]	1.92 [0.58;8.11]	0.14 [0.04;0.64]
UK	1.85 [0.54;7.48]	3.64 [0.79;14.28]	0.18 [0.05;0.72]

Table A.1: Contemporaneous coefficients in SOE monetary policy equations using full sample period from 1992:Q1 to 2019:Q4. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

A.4.2 Contemporaneous coefficients in US monetary policy equations.

	Coefficients		
US/SOEs	ψ_{y^*}	ψ_{π^*}	ψ_{cs^*}
US/AUS	0.41	0.38	-0.57
	[0.11;0.87]	[0.10;1.00]	[-1.36;-0.19]
US/CAN	0.40	0.39	-0.59
	[0.14;0.90]	[0.11;1.00]	[-1.38;-0.17]
US/NZ	0.41	0.40	-0.57
	[0.12;0.99]	[0.12;0.99]	[-1.42;-0.15]
US/NOR	0.43	0.38	-0.55
	[0.13;0.91]	[0.10;0.97]	[-1.33;-0.14]
US/SWE	0.42	0.39	-0.54
	[0.12;0.98]	[0.10;1.09]	[-1.34;-0.16]
US/UK	0.43	0.39	-0.56
	[0.13;0.93]	[0.11;1.00]	[-1.33;-0.15]

Table A.2: Contemporaneous coefficients in US monetary policy equations using full sample period from 1992:Q1 to 2019:Q4. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

A.5 Robustness check # 1

A.5.1 Contemporaneous coefficients in SOE monetary policy equations.

	Coefficients		
SOEs	ψ_y	ψ_π	ψ_e
AUS	0.72 [0.19;2.77]	0.73 [0.16;2.56]	0.11 [0.02;0.45]
CAN	1.16 [0.30;4.00]	1.42 [0.46;4.36]	0.24 [0.06;0.70]
NZ	0.99 [0.29;3.80]	1.67 [0.45;6.31]	0.14 [0.04;0.56]
NOR	0.79 [0.21;3.06]	1.80 [0.46;6.93]	0.20 [0.05;0.79]
SWE	1.19 [0.31;4.65]	2.03 [0.55;7.90]	0.15 [0.03;0.72]
UK	1.21 [0.30;5.12]	2.62 [0.87;9.29]	0.14 [0.03;0.65]

Table A.3: Contemporaneous coefficients in SOE monetary policy equations using shorter sample period from 1992:Q1 to 2008:Q3. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

A.5.2 Contemporaneous coefficients in US monetary policy equations.

	Coefficients		
US/SOEs	ψ_{y^*}	ψ_{π^*}	ψ_{cs^*}
US/AUS	0.50 [0.14;1.23]	0.67 [0.18;1.74]	-0.99 [-2.02;-0.29]
US/CAN	0.51 [0.16;1.24]	0.73 [0.20;2.02]	-1.01 [-2.09;-0.31]
US/NZ	0.53 [0.16;1.26]	0.76 [0.20;2.05]	-0.98 [-2.11;-0.31]
US/NOR	0.51 [0.15;1.19]	0.75 [0.21;1.89]	-1.01 [-2.17;-0.30]
US/SWE	0.49 [0.14;1.20]	0.73 [0.20;1.84]	-1.01 [-2.12;-0.34]
US/UK	0.50 [0.15;1.21]	0.72 [0.21;1.88]	-0.96 [-2.12;-0.28]

Table A.4: Contemporaneous coefficients in US monetary policy equations using shorter sample period from 1992:Q1 to 2008:Q3. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

A.6 Robustness check # 2

A.6.1 Contemporaneous coefficients in SOE monetary policy equations.

	Coefficients			
SOEs	ψ_y	ψ_π	ψ_{cs}	ψ_e
AUS	0.64 [0.16;2.84]	0.70 [0.19;2.51]	-3.00 [-8.48;-0.91]	0.10 [0.02;0.50]
CAN	0.63 [0.19;1.56]	0.60 [0.19;1.86]	-0.90 [-2.65;-0.33]	0.08 [0.02;0.31]
NZ	0.94 [0.23;3.89]	0.81 [0.23;5.83]	-1.35 [-7.50;-0.32]	0.06 [0.01;0.40]
NOR	0.41 [0.13;1.22]	0.64 [0.16;2.52]	-2.47 [-10.21;-0.58]	0.11 [0.03;0.45]
SWE	0.27 [0.07;1.06]	0.94 [0.26;2.09]	-0.74 [-2.32;-0.23]	0.06 [0.02;0.19]
UK	1.19 [0.33;3.90]	2.25 [0.65;8.46]	-1.74 [-6.07;-0.57]	0.19 [0.04;0.44]

Table A.5: Contemporaneous coefficients in SOE monetary policy equations using sample period from 2000:Q1 to 2019:Q4. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

A.6.2 Contemporaneous coefficients in US monetary policy equations.

	Coefficients		
US/SOEs	ψ_{y^*}	ψ_{π^*}	ψ_{cs^*}
US/AUS	0.43 [0.14;1.05]	0.48 [0.16;1.61]	-0.55 [-1.38;-0.16]
US/CAN	0.47 [0.12;1.23]	0.45 [0.10;1.10]	-0.65 [-1.86;-0.21]
US/NZ	0.54 [0.15;1.39]	0.46 [0.09;1.38]	-0.60 [-1.44;-0.15]
US/NOR	0.40 [0.14;1.35]	0.46 [0.12;1.40]	-0.53 [-1.44;-0.17]
US/SWE	0.52 [0.19;1.14]	0.44 [0.11;1.17]	-0.62 [-1.72;-0.21]
US/UK	0.47 [0.11;1.39]	0.45 [0.10;1.39]	-0.59 [-1.70;-0.15]

Table A.6: Contemporaneous coefficients in US monetary policy equations using sample period from 2000:Q1 to 2019:Q4. *Note:* The entries in the table are the posterior median estimates and the entries in the brackets are the respective 68% probability intervals.

References

Akinci, O., Benigno, G., Pelin, S., and Turek, J. (2022). The Dollar's Imperial Circle. *Federal Reserve Bank of New York Staff Reports*, 1045.

Arias, J. E., Caldara, D., and Rubio-Ramirez, J. F. (2019). The systematic component of monetary policy in SVARs: An agnostic identification procedure. *Journal of Monetary Economics*, 101:1–13.

Arias, J. E., Rubio-Ramírez, J. F., and Waggoner, D. F. (2018). Inference based on structural vector autoregressions identified with sign and zero restrictions: Theory and applications. *Econometrica*, 86(2):685–720.

Beckers, B. et al. (2020). *Credit spreads, monetary policy and the price puzzle*. Reserve Bank of Australia.

Bjørnland, H. C. (2009). Monetary policy and exchange rate overshooting: Dornbusch was right after all. *Journal of International Economics*, 79(1):64–77.

Bjørnland, H. C. and Halvorsen, J. I. (2014). How does monetary policy respond to exchange rate movements? New international evidence. *Oxford Bulletin of Economics and Statistics*, 76(2):208–232.

Caldara, D. and Herbst, E. (2019). Monetary policy, real activity, and credit spreads: Evidence from Bayesian proxy SVARs. *American Economic Journal: Macroeconomics*, 11(1):157–192.

Calvo, G. A. and Reinhart, C. M. (2002). Fear of floating. *The Quarterly Journal of economics*, 117(2):379–408.

Castelnuovo, E., Pellegrino, G., Ranzato, G., et al. (2022). *Delayed Overshooting Puzzle: Does Systematic Monetary Policy Matter?* Department of Economics and Management "Marco Fanno", University of Padova.

Clarida, R. and Gali, J. (1994). Sources of real exchange rate fluctuations: how important are nominal shocks? *Carnegie-Rochester Conference Series on Public Policy*, 41:1–56.

Curdia, V. and Woodford, M. (2010). Credit spreads and monetary policy. *Journal of Money, Credit and Banking*, 42:3–35.

Curdia, V. and Woodford, M. (2016). Credit frictions and optimal monetary policy. *Journal of Monetary Economics*, 84:30–65.

Cushman, D. O. and Zha, T. (1997). Identifying monetary policy in a small open economy under flexible exchange rates. *Journal of Monetary economics*, 39(3):433–448.

De Rezende, R. B. and Ristiniemi, A. (2023). A shadow rate without a lower bound constraint. *Journal of Banking & Finance*, 146:106686.

Dieppe, A., Legrand, R., and Roye, B. V. (2016). The BEAR toolbox (ECB Working PaperNo. 1934). *European Central Bank*.

Doko Tchatoka, F., Haque, Q., and Terrell, M. (2022). Monetary policy shocks and exchange rate dynamics in small open economies. *Available at SSRN 4025608*.

Dornbusch, R. (1976). Expectations and exchange rate dynamics. *Journal of Political Economy*, 84(6):1161–1176.

Egorov, K. and Mukhin, D. (2023). Optimal Policy under Dollar Pricing. *SAFE Working Paper*, 377.

Eichenbaum, M. and Evans, C. L. (1995). Some empirical evidence on the effects of shocks to monetary policy on exchange rates. *The Quarterly Journal of Economics*, 110(4):975–1009.

Faust, J. and Rogers, J. H. (2003). Monetary policy's role in exchange rate behavior. *Journal of Monetary Economics*, 50(7):1403–1424.

Fry, R. and Pagan, A. (2011). Sign restrictions in structural vector autoregressions: a critical review. *Journal of Economic Literature*, 49(4):938–960.

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

Gourinchas, P.-O. (2021). The Dollar Hegemon? Evidence and Implications for Policymakers. *The Asian Monetary Policy Forum*, in Steven J. Davis, Edward S Robinson and Bernard Yeung (ed.), Ch.7:264–300.

Gourinchas, P.-O., Rey, H., and Sauzet, M. (2019). The International Monetary and Financial System. *Annual Review of Economics*, 11:859–893.

Ilzetzki, E., Reinhart, C. M., and Rogoff, K. S. (2019). Exchange arrangements entering the twenty-first century: Which anchor will hold? *The Quarterly Journal of Economics*, 134(2):599–646.

Jääskelä, J. P. and Jennings, D. (2011). Monetary policy and the exchange rate: Evaluation of VAR models. *Journal of International Money and Finance*, 30(7):1358–1374.

Justiniano, A. and Preston, B. (2010). Can structural small open-economy models account for the influence of foreign disturbances? *Journal of International Economics*, 81(1):61–74.

Kam, T., Lees, K., and Liu, P. (2009). Uncovering the hit list for small inflation targeters: A Bayesian structural analysis. *Journal of Money, Credit and Banking*, 41(4):583–618.

Kim, S. and Lim, K. (2018). Effects of monetary policy shocks on exchange rate in small open Economies. *Journal of Macroeconomics*, 56:324–339.

Kim, S. and Roubini, N. (2000). Exchange rate anomalies in the industrial countries: A solution with a structural VAR approach. *Journal of Monetary economics*, 45(3):561–586.

Kim, S.-H., Moon, S., and Velasco, C. (2017). Delayed overshooting: is it an'80s puzzle? *Journal of Political Economy*, 125(5):1570–1598.

Koop, G., Korobilis, D., et al. (2010). Bayesian multivariate time series methods for empirical macroeconomics. *Foundations and Trends® in Econometrics*, 3(4):267–358.

Lane, P. R. (2001). The new open economy macroeconomics: A survey. *Journal of International Economics*, 54(2):235–266.

Leeper, E. M., Sims, C., and Zha, T. (1996). What does monetary policy do? *Brookings Papers on Economic Activity*, 1996(2):1–78.

Lubik, T. A. and Schorfheide, F. (2007). Do central banks respond to exchange rate movements? A structural investigation. *Journal of Monetary Economics*, 54(4):1069–1087.

Maćkowiak, B. (2007). External shocks, US monetary policy and macroeconomic fluctuations in emerging markets. *Journal of Monetary Economics*, 54(8):2512–2520.

Miranda-Agrippino, S. and Nenova, T. (2022). A tale of two global monetary policies. *Journal of International Economics*, 136:103606.

Miranda-Agrippino, S. and Rey, H. (2020). U.S. Monetary Policy and the Global Financial Cycle. *Review of Economic Studies*, 87(6):2754–2776.

Obstfeld, M. (2013). Crises and the international system. *International Economic Journal*, 27(2):143–155.

Obstfeld, M. and Rogoff, K. (1995). The mirage of fixed exchange rates. *Journal of Economic Perspectives*, 9(4):73–96.

Read, M. (2023). Estimating the effects of monetary policy in Australia using sign-restricted Structural Vector Autoregressions. *Economic Record*, forthcoming.

Reinhart, C. M. and Rogoff, K. S. (2004). The modern history of exchange rate arrangements: a reinterpretation. *The Quarterly Journal of Economics*, 119(1):1–48.

Rey, H. (2013). Dilemma not trilemma: the global financial cycle and monetary policy independence. *Proceedings - Economic Policy Symposium - Jackson Hole*, Federal Reserve Bank of Kansas City:285–333.

Rogers, J. H. (1999). Monetary shocks and real exchange rates. *Journal of International Economics*, 49(2):269–288.

Rubio-Ramirez, J. F., Waggoner, D. F., and Zha, T. (2010). Structural vector autoregressions: Theory of identification and algorithms for inference. *The Review of Economic Studies*, 77(2):665–696.

Rüth, S. K. (2020). Shifts in monetary policy and exchange rate dynamics: Is Dornbusch's overshooting hypothesis intact, after all? *Journal of International Economics*, 126:103344.

Scholl, A. and Uhlig, H. (2008). New evidence on the puzzles: Results from agnostic identification on monetary policy and exchange rates. *Journal of International Economics*, 76(1):1–13.

Taylor, J. B. (1993). Discretion versus policy rules in practice. In *Carnegie-Rochester Conference Series on Public Policy*, volume 39, pages 195–214. Elsevier.

Taylor, J. B. (2001). The role of the exchange rate in monetary-policy rules. *American Economic Review*, 91(2):263–267.

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.

Wu, J. C. and Xia, F. D. (2016). Measuring the macroeconomic impact of monetary policy at the zero lower bound. *Journal of Money, Credit and Banking*, 48(2-3):253–291.

Statement of Authorship

Title of Paper	What Are the Effects of Commodity Supply and Demand Shocks on Australian Output and Trade Balance Fluctuations?
Publication Status	☐ Published ☐ Accepted for Publication ☐ Submitted for Publication ☑ Unpublished and Unsubmitted w ork w ritten in manuscript style
Publication Details	

Principal Author

Name of Principal Author (Candidate)	Naveed Javed		
Contribution to the Paper			
Overall percentage (%)	100%		
Certification:	This paper reports on original research I conducted during the period of my Higher Degree by Research candidature and is not subject to any obligations or contractual agreements with a third party that would constrain its inclusion in this thesis. I am the primary author of this paper.		
Signature		Date	12.09.2023

Co-Author Contributions

By signing the Statement of Authorship, each author certifies that:

i. the candidate's stated contribution to the publication is accurate (as detailed above);

ii. permission is granted for the candidate in include the publication in the thesis; and

iii. the sum of all co-author contributions is equal to 100% less the candidate's stated contribution.

Name of Co-Author			
Contribution to the Paper			
Signature		Date	

Name of Co-Author			
Contribution to the Paper			
Signature		Date	

Please cut and paste additional co-author panels here as required.

Chapter 2

What Are the Effects of Commodity Supply and Demand Shocks on Australian Output and Trade Balance Fluctuations?*

Naveed Javed†

* I wish to thank Efrem Castelnuovo and James Hansen for helpful comments.
† School of Economics, University of Adelaide, naveed.javed@adelaide.edu.au

Abstract

In this chapter, I estimate SVAR models to measure the effects of commodity supply and commodity demand shocks on Australia's output and trade balance. I use commodity prices as a proxy for (negative) commodity supply shocks and Chinese steel production as (positive) commodity demand shocks. I find that commodity supply and demand shocks emerge as a relatively minor and negligible sources of business cycle fluctuations in output and trade balance, except for the fluctuations in trade balance (35%) to commodity demand shocks. In response to a commodity demand shock, output expands however, output contracts in response to a commodity supply shock. Interestingly, for both commodity supply and demand shocks the trade balance worsens substantially.

Keywords: Commodity supply shock, Commodity demand shock, Trade balance, Business cycles, Structural vector autoregressions.

JEL codes: **C32, E32, F43**.

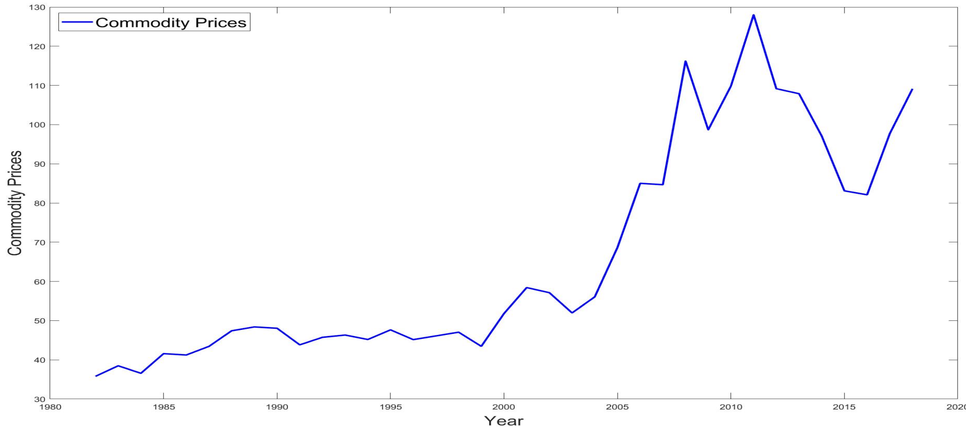


Figure 2.1: Commodity Price Index (Source: Reserve Bank of Australia)

2.1 Introduction

Australia's commodity price index, displayed in Figure 2.1, increased markedly over the past two decades, reaching a peak in September 2012[1]. The unprecedented surge in commodity prices that started in the late Nineties has been attributed to the strong economic growth of China and its demand for steel to produce infrastructure (Dungey et al., 2017). Given the importance of commodities in Australia's export base, wide swings in commodity prices would be expected to have a large impact on the Australian economy.[2]

My empirical strategy is inspired by recent studies by Drechsel and Tenreyro (2018), Schmitt-Grohé and Uribe (2018) and Di Pace et al. (2020). Drechsel and Tenreyro (2018) estimate a SVAR model to measure the macroeconomic effects of commodity price shocks for Argentina. They find a significant expansion in Argentina's GDP and a significant contraction in trade balance in response to a commodity price shock. Schmitt-Grohé and Uribe (2018) estimate a SVAR model to measure the macroeconomic effects of terms of trade shocks for a panel of 38 poor and emerging economies. According to their median estimates, terms of trade shocks account for roughly 10% of

[1] The commodity price index includes 21 major commodities exported by Australia, which accounted for over 90 percent of Australia's commodity export earnings in 2012.

[2] Share of commodities in Australia's export base is around 40% (https://wits.worldbank.org/CountryProfile/en/AUS).

both the variance of output and that of the trade balance. Moreover, output and the trade balance both expand on impact in response to a terms of trade shock. Di Pace et al. (2020) follow Schmitt-Grohé and Uribe (2018) in their choice for panel of countries. However, Di Pace et al. (2020) identify the terms of trade shocks as (positive) export prices and (negative) import prices shocks using sign and narrative restrictions. They find that output expands significantly in response to both the export and import prices shocks. However, trade balance deteriorates in response to the export prices shock and slightly improves in response to the import prices shock. They find that export and import prices shocks jointly account for around 40% of output and 32% of trade balance fluctuations.

I analyse the effects of commodity price shocks on macroeconomic fluctuations in Australia in my baseline model. Commodity price shocks serve as proxy for commodity supply shocks. A (positive) commodity price shock is interpreted as a (negative) commodity supply shock (Dungey et al., 2014). I estimate a structural vector autoregression (SVAR) model over the period 1983:Q3 to 2018:Q4 to quantify the impact of exogenous disturbances in commodity prices on output and the trade balance. I find that the contribution of commodity price shocks to output fluctuations is negligible. However, these disturbances account for around 25% of the variance of the trade balance. In response to a positive innovation in commodity prices, output contracts and trade balance deteriorates significantly and persistently by around 0.2%.

I extend my baseline model by introducing a large *"foreign block"* and thus, endogenising commodity prices. I estimate my extended SVAR model over the period 1988:Q1 to 2018:Q4 to quantify the impact of commodity supply and demand disturbances on output and trade balance.[3] I find a persistent expansion of output to commodity demand shocks, while the trade balance deteriorates significantly. On the other hand, I find an insignificant and persistent contraction of output and a significant and persistent deterioration of trade balance in response to commodity supply shocks.

[3] The starting date is restricted due to Chinese steel production data.

Despite the central role typically attributed to terms of trade and commodity price disturbances in narratives of Australian business cycles, there is no consensus in the literature on the effects of these shocks on output and the trade balance. Moreover, relatively few studies have considered the reaction of the trade balance to commodity price shocks. The literature on the impact of terms of trade and commodity price shocks on Australia's business cycle fluctuations can be split into three related streams. I see my work as an attempt to build a bridge between the recent contributions by Drechsel and Tenreyro (2018), Schmitt-Grohé and Uribe (2018) and these three strands of literature.

A first stream of papers applies SVAR models to search for evidence of the Harberger-Laursen-Metzler (HLM) effect in Australia among various commodity exporting countries (Ahmed and Park, 1994; Cashin and McDermott, 2002; Otto, 2003). The HLM effect predicts that an exogenous increase in a country's terms of trade leads to an improvement in that country's trade balance. Ahmed and Park (1994) use quarterly data from 1960:Q1 to 1987:Q4 for 7 OECD economies, including Australia, to estimate a SVAR model identified through long-run restrictions (Blanchard and Quah, 1989). For Australia, and in general, they do not find much empirical support for the HLM effect. Instead, Cashin and McDermott (2002) examine the period 1970:Q2 to 1997:Q2 for five countries, including Australia, and conclude that the HLM effect plays an important role in the dynamics of Australia's current account. Finally, in a seminal paper, Otto (2003) inspects the empirical robustness of the HLM effect by estimating a 3-variable (terms of trade, output, trade balance) SVAR model with annual data from 1960-1996 for a large panel of small open economies. Otto (2003) documents evidence for Australia of a positive relationship between the terms of trade and the trade balance, and concludes more generally that there is, across countries, strong support in the data for the existence of the HLM effect.

A second strand of the literature explores the role of terms of trade shocks in macroeconomic fluctuations using SVAR models (Dungey and Pagan, 2000, 2009; Liu, 2010). An early work on Australian terms of trade and SVAR models by Dungey and

Pagan (2000) estimate an 11-variable SVAR model. They use Australian quarterly data from 1980:Q1 to 1998:Q3, dividing the variables into two blocks: an exogenous foreign block and an endogenous domestic block. They find that in response to a terms of trade shock Australian GDP expands significantly for more than a year. Dungey and Pagan (2009) extend their previous work by considering a longer sample period from 1980:Q1 to 2006:Q4 and confirm their earlier empirical findings. Liu (2010) also employs Australian quarterly data from 1980:Q1 to 2006:Q1 to estimate a VAR model but imposes a set of sign restrictions to identify structural disturbances. In line with Dungey and Pagan (2000) and Dungey and Pagan (2009), Liu (2010) finds that a positive terms of trade shock leads to a temporary increase in output. To sum up, these earlier papers that focus on terms of trade shocks and consider data up until 2006, all tend to find evidence of a temporary boost in output.

Finally, a third stream of empirical studies using SVAR models considers sample periods from more recent years and deals explicitly with the identification of commodity price shocks, as opposed to terms of trade shocks (Dungey et al., 2020, 2017, 2014; Knop and Vespignani, 2014; Jääskelä and Smith, 2013). Jääskelä and Smith (2013) estimate a VAR model over the period from 1984:Q1 to 2010:Q2 and use a set of sign restrictions to unpack the terms of trade shock into a world demand shock, a commodity-market-specific shock and a globalisation shock. They find a significant and persistent expansion in output in response to a commodity price shock. Like Jääskelä and Smith (2013), Dungey et al. (2014) also unpack the terms of trade shock into a Chinese resource demand shock, a commodity price shock and a foreign output shock. They estimate their SVAR model with data from 1988:Q1 to 2011:Q2 and document an insignificant slowdown in output in response to an increase in both the Chinese steel production and commodity prices. (Dungey et al., 2017, 2020) extend the period of investigation to 2016 and confirm evidence of a slowdown in output. Finally, Knop and Vespignani (2014) estimate a VAR model over the period 1993:Q1 to 2013:Q1 and identify commodity price shocks through recursive zero short-run restrictions and report a significant decline in output growth rate in response to a commodity price shock. Overall, this set of more recent papers suggests that commodity price disturbances hitting the Australian economy may have a dampening effect on output growth.

The rest of this chapter is structured as follows. Section 2.2 describes the data. Section 2.3 outlines the econometric strategy and identification scheme for my baseline model. Section 2.4 presents the baseline results. Section 2.5 contains the extended model and discussion of results. Section 2.6 and 2.7 discuss the robustness check and variance decomposition analysis, respectively. Finally, Section 2.8 concludes.

2.2 Data Description

My baseline model includes six variables: commodity prices (cp), trade balance (tb), domestic output (yd), consumption (c), investment (i) and real exchange rate (RER).[4] Appendix B contains a full description of the data and their sources.

The commodity price index comes from the Reserve Bank of Australia (RBA). The commodity price index is computed by giving weights to different commodities: Bulk Commodities (54.0%), Rural Commodities (12.2%), Base Metals (4.2%) and Other resources (29.6%).

Domestic output is the real gross domestic product expressed in chain volume measure. Consumption is the final household consumption expenditure expressed in chain volume measure. Investment is gross fixed capital formation expressed in chain volume measure. Exports and imports are the exports of goods and services and imports of goods and services expressed in chain volume measure. Chain volume measures are derived by deflating the original current price series by specifically compiled measures of price change by the Australian Business Statistics (ABS). The reference year for chain volume measures is the year prior to the latest complete financial year incorporated into the time series. The reference year is updated with the completion of estimates for each financial year.

[4] $tb = x - m$; x = exports per capita, m = imports per capita.

The exchange rate is represented by the real exchange rate trade weighted index calculated by the RBA. The real exchange rate trade weighted index is a price in terms of a weighted average of a basket of 19 currencies, where Chinese renminbi is given the highest weight (27.7%). The real exchange rate-trade weighted index will therefore give a measure of whether the Australian dollar is rising or falling on average against the currencies of Australia's trading partners. This variable choice is consistent with Dungey et al. (2017, 2014), Dungey and Pagan (2000, 2009), Lawson et al. (2008) and Jääskelä and Smith (2013).

My definition of the variables closely follow Drechsel and Tenreyro (2018). All the variables are divided by the working age population to express them in real per-capita terms except cp and RER.[5] All the variables cp, yd, c, i and RER are in log-levels except the trade balance. The trade balance is computed as a ratio of exports minus imports to the trend component of per-capita gross domestic product.

The baseline SVAR model is estimated using quarterly data from 1983:Q3 to 2018:Q4, corresponding to the post-float period of the Australian dollar. Dungey and Pagan (2000, 2009), Dungey et al. (2014, 2017, 2020) and Jääskelä and Smith (2013) also use quarterly data beginning from around 1983. However, Otto (2003) uses annual data from 1960 to 1996. My data set has the advantage that it contains the critical six years period (2013-2018) after the recent boom in the commodity prices.

2.3 Empirical Model

My 6-variable SVAR model is motivated by Drechsel and Tenreyro (2018) and Schmitt-Grohé and Uribe (2018) in choice of variables while my structural model follows Arias et al. (2019). I adapt the methodology of Arias et al. (2019) to a small open economy context by implementing a block-exogeneity structure. The variables in the

[5] The working age population data is obtained from Australian Demographic Statistics and it is constructed by summing the total number of residents aged between 15 and 64.

SVAR are classified into two blocks: a *"foreign block"* and a *"domestic block"*. Block exogeneity restrictions imply that, *"foreign block"* influences the *"domestic block"* (both contemporaneously and through lags), while the *"domestic block"* has no effect on the *"foreign block"*. The structural model takes the form:

$$\mathbf{y_t'}\mathbf{A_0} = \sum_{L=1}^{p} \mathbf{y_{t-L}'}\mathbf{A_L} + \mathbf{c}' + \epsilon_t'. \tag{2.1}$$

I divide $\mathbf{y_t}$ into two blocks: $\mathbf{y_{1t}'} = [cp_t]$ and $\mathbf{y_{2t}'} = [tb_t, yd_t, c_t, i_t, RER_t]$.[6] $\mathbf{y_{1t}}$ is a $n_1 \times 1$ vector of exogenous (*foreign*) variables and $\mathbf{y_{2t}}$ is a $n_2 \times 1$ vector of endogenous (*domestic*) variables; where n_1 is 1 as I have one exogenous variable in my baseline model, n_2 is the number of endogenous variables and n is the total number of variables, implying $n_1 + n_2 = n$. Similarly, ϵ_t is divided into exogenous (ϵ_{1t}) and endogenous (ϵ_{2t}) blocks of $n_1 \times 1$ and $n_2 \times 1$ vectors of structural shocks, respectively. $\epsilon_{1t}' = [\epsilon^{cp}]$ and $\epsilon_{2t}' = [\epsilon^{tb}\ \epsilon^{yd}\ \epsilon^{c}\ \epsilon^{i}\ \epsilon^{RER}]$. $\mathbf{A_i}$ are $n \times n$ matrices of structural parameters for $0 \leq i \leq p$ with $\mathbf{A_0}$ invertible, $\mathbf{c}$ is a $n \times 1$ vector of parameters, p is the lag length, and T is the sample size. I set $p = 4$.[7] The vector ϵ_t, conditional on past information and the initial conditions $\mathbf{y}_0, ..., \mathbf{y}_{1-p}$, is Gaussian with mean zero and covariance matrix $\mathbb{I}_n$ (the $n \times n$ identity matrix). The SVAR described in Equation (2.1) can be written in compact form as in Rubio-Ramirez et al. (2010):

$$\mathbf{y_t}'\mathbf{A_0} = \mathbf{x_t}'\mathbf{A}_+ + \epsilon_t' \ , \tag{2.2}$$

where $\mathbf{A}_+ = [\mathbf{A_1'} \dots \mathbf{A_p'} \ \ \mathbf{c}']_{(np+1)\times n}$ and $\mathbf{x_t'} = [\mathbf{y_{t-1}'} \dots \mathbf{y_{t-p}'} \ \ 1]$ for $1 \leq \text{t} \leq \text{T}$. I post-multiply Equation (2.2) by $\mathbf{A_0}^{-1}$ to obtain the reduced form:

$$\mathbf{y_t'} = \mathbf{x_{t-L}'}B + u_t' \ , \tag{2.3}$$

where $B = \mathbf{A}_+\mathbf{A_0}^{-1}, u_t' = \epsilon_t'\mathbf{A_0}^{-1}$ and $\mathbf{E}[u_t u_t'] := \Sigma :=: (\mathbf{A_0}\mathbf{A_0'})^{-1}$. Here B is a matrix of

[6] Appendix B contains the plots of variables as these enter in the baseline SVAR model.
[7] The lag length specification tests suggest that either four (Hannan-Quinn information criteria, Akaike Information Criteria, Likelihood Ratio) or three (Schwartz Bayesian Information Criteria, Final Prediction Error) lags should be used.

reduced-form VAR coefficients and Σ is the residual variance-covariance matrix. $\mathbf{A_0}$ and $\mathbf{A_+}$ are matrices of structural parameters.[8]

2.3.1 Model identification

Exogeneity restrictions

As a small open economy I assume that Australian production decisions have no impact on commodity prices. Consequently, I model commodity price fluctuations as a univariate autoregressive process.[9] I impose the Cholesky decomposition by applying exclusion restrictions directly on the structural parameters in $\mathbf{A_0}$ matrix, which makes cp_t contemporaneously exogenous. The contemporaneous matrix takes the form:

$$\mathbf{A_0} = \begin{bmatrix} a_{0,11} & a_{0,12} & a_{0,13} & a_{0,14} & a_{0,15} & a_{0,16} \\ 0 & a_{0,22} & a_{0,23} & a_{0,24} & a_{0,25} & a_{0,26} \\ 0 & 0 & a_{0,33} & a_{0,34} & a_{0,35} & a_{0,36} \\ 0 & 0 & 0 & a_{0,44} & a_{0,45} & a_{0,46} \\ 0 & 0 & 0 & 0 & a_{0,55} & a_{0,56} \\ 0 & 0 & 0 & 0 & 0 & a_{0,66} \end{bmatrix}$$

Further, I impose exogeneity restriction on $\mathbf{A_L}$ matrices through informative priors, such that lags of endogenous (*"domestic"*) variables do not impact cp_t. The next two subsections contain the technical details of how I have implemented block exogeneity through Independent Normal Inverse Wishart prior.

[8] I estimate the model through 1,00,000 draws from posterior distribution.

[9] Schmitt-Grohé and Uribe (2018) find that an F-test against the alternative that the terms of trade depend on lagged values of the trade balance, output, consumption, and investment is rejected at the 5% level for 32 out of the 38 countries in the sample.

Implementing block exogeneity

I extend the methodology of Arias et al. (2019) in order to implement block exogeneity on lagged matrices. Consider matrix $\mathbf{A_L}$ from Equation (2.1):[10]

$$\mathbf{y_t'} = \begin{bmatrix} \mathbf{y_{1t}'} & \mathbf{y_{2t}'} \end{bmatrix}, \quad \mathbf{A_L} = \begin{bmatrix} A_{11,L} & A_{12,L} \\ \\ A_{21,L} & A_{22,L} \end{bmatrix}, \quad \epsilon_t' = \begin{bmatrix} \epsilon_{1t}' & \epsilon_{2t}' \end{bmatrix}$$

where $A_{11,L}$ is $n_1 \times n_1$, $A_{12,L}$ is $n_1 \times n_2$, $A_{21,L}$ is $n_2 \times n_1$, $A_{22,L}$ is $n_2 \times n_2$, $\mathbf{y_{1t}'}$ is $1 \times n_1$, $\mathbf{y_{2t}'}$ is $1 \times n_2$, ϵ_{1t}' is $1 \times n_1$ and ϵ_{2t}' is $1 \times n_2$; $n_1 + n_2 = n$.

The seminal work of Cushman and Zha (1997) implement block exogeneity on the lagged matrices by restricting $A_{21,L} = 0$. This restriction means that the domestic block ($\mathbf{y_{2t}'}$) does not influence the dynamics of the foreign block ($\mathbf{y_{1t}'}$) through lags. However, when using the methodology developed by Arias et al. (2019), one can only impose $(n-k)$ zero restrictions per equation, where k is the k^{th} equation in the system (Arias et al., 2021; Kilian and Lütkepohl, 2017). As a result, I cannot simply impose $A_{21,L} = 0$. Instead, I use a special case of Independent Normal Inverse Wishart ($\mathcal{NIW}$) priors where I adopt conventional Minnesota priors for the reduced-form VAR coefficients, β.[11] Specifically to implement block exogeneity on the lagged matrices, I use informative priors with distribution centered at zero for all the reduced-form coefficients of domestic variables appearing in the foreign block. In my case, these informative priors correspond to imposing block exogeneity restrictions of Cushman and Zha (1997).[12] Thus, the lagged matrices take the from:[13]

[10] For simplicity, I omit constant terms in the exposition of the model/methodology.
[11] $\beta = vec(B)$.
[12] The block exogeneity structure would be more explicit in section 2.5 when I include more variables in the *foreign block*.
[13] I apply informative priors on the elements in $\mathbf{A_L}$ matrices, highlighted in *bold*.

$$\mathbf{A_L} = \begin{bmatrix} a_{L,11} & a_{L,12} & a_{L,13} & a_{L,14} & a_{L,15} & a_{L,16} \\ \boldsymbol{a_{L,21}} & a_{L,22} & a_{L,23} & a_{L,24} & a_{L,25} & a_{L,26} \\ \boldsymbol{a_{L,31}} & a_{L,32} & a_{L,33} & a_{L,34} & a_{L,35} & a_{L,36} \\ \boldsymbol{a_{L,41}} & a_{L,42} & a_{L,43} & a_{L,44} & a_{L,45} & a_{L,46} \\ \boldsymbol{a_{L,51}} & a_{L,52} & a_{L,53} & a_{L,54} & a_{L,55} & a_{L,56} \\ \boldsymbol{a_{L,61}} & a_{L,62} & a_{L,63} & a_{L,64} & a_{L,65} & a_{L,66} \end{bmatrix}$$

The exogeneity restrictions, through tight priors, are imposed on the first column of $\mathbf{A_L}$ matrics except the first elements $a_{L,11}$. Hence, the first equation of the structural model in Equation (2.2) represents the law of motion of commodity prices which is given by:[14]

$$cp_t a_{0,11} = cp_{t-1}a_{1,11} + cp_{t-2}a_{2,11} + cp_{t-3}a_{3,11} + cp_{t-4}a_{4,11} + \epsilon_t^{cp} \quad (2.4)$$

$$cp_t = cp_{t-1}\underbrace{a_{1,11}a_{0,11}^{-1}}_{\psi_{cp_{t-1}}} + cp_{t-2}\underbrace{a_{2,11}a_{0,11}^{-1}}_{\psi_{cp_{t-2}}} + cp_{t-3}\underbrace{a_{3,11}a_{0,11}^{-1}}_{\psi_{cp_{t-3}}} + cp_{t-4}\underbrace{a_{4,11}a_{0,11}^{-1}}_{\psi_{cp_{t-4}}} + \epsilon_t^{cp}\underbrace{a_{0,11}^{-1}}_{\sigma} \quad (2.5)$$

where $a_{L,11}$ and $a_{0,11}$ denote the elements (1,1) of $\mathbf{A_L}$ and $\mathbf{A_0}$, respectively. As a consequence, the first element of ϵ_t, denoted ϵ_t^{cp} has the interpretation of a commodity price shock, because it is the only innovation that affects commodity prices. I estimate the reduced form VAR in Equation (2.3) using the methodology developed by Arias et al. (2018) and Arias et al. (2019).[15]

Independent ($\mathcal{NIW}$) priors and posteriors

Arias et al. (2019) use Natural Conjugate ($\mathcal{NIW}$) priors which are not suited for the purpose of implementing block exogeneity.[16] Fortunately, the techniques developed by

[14] I abstract from the lags of domestic variables as an implication of exogeneity restrictions applied through informative priors.

[15] Following Arias et al. (2019), I normalize the IRFs by imposing sign-restrictions that commodity prices respond positively on impact in response to a positive innovation in commodity prices and that $\sigma > 0$.

[16] Natural Conjugate ($\mathcal{NIW}$) priors employ Kronecker structure for the variance-covariance matrix of the reduced-form parameters. Hence, the variances are proportional to one another. As a result, imposing block exogeneity on one equation would impose it on all equations (Dieppe et al., 2016; Koop et al., 2010). Secondly, the structure implied by the Kronecker product requires that every equation

Arias et al. (2018) can be used for any prior distributions. Hence, I use Independent ($\mathcal{NIW}$) priors over the reduced-form parameters to implement block exogeneity. The prior distributions of the reduced-form parameters from Equation (2.3) take the form:

$$\beta \sim \mathcal{N}(\beta_0, \Omega_0), \tag{2.6}$$

$$\Sigma \sim \mathcal{IW}(S_0, \alpha_0), \tag{2.7}$$

where the residual variance-covariance matrix Σ has an inverse Wishart distribution with scale matrix S_0 and degrees of freedom α_0.[17]

I follow standard Minnesota priors for the prior distribution of the coefficients (β) in Equation (2.6), where the distribution is centered at 1 for the coefficients of own 1st lags and for the rest the distributions are centered at 0 including the constants (Kilian and Lütkepohl, 2017; Dieppe et al., 2016). Ω_0 is an arbitrary variance-covariance matrix of β and, unlike the Natural Conjugate ($\mathcal{NIW}$) priors, it is independent of residual variance-covariance matrix Σ. Moreover, Ω_0 is a diagonal matrix containing the hyperparameters that control the variance of the distributions. The diagonal elements of Ω_0 take the form:

$$\sigma^2_{c_i} = \sigma^2_i(\lambda_1\lambda_4)^2 \qquad \text{if constant} \tag{2.8}$$

$$\sigma^2_{ii} = (\lambda_1/L^{\lambda_3})^2 \qquad \text{if } i = j \tag{2.9}$$

$$\sigma^2_{ij} = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2/L^{\lambda_3})^2 \qquad \text{if } i \neq j \tag{2.10}$$

$$\sigma^2_{Ex_{ij}} = (\sigma_i/\sigma_j)^2(\lambda_1\lambda_2\lambda_5/L^{\lambda_3})^2 \qquad \text{if } i \neq j \text{ and } e_x < j \leq n \tag{2.11}$$

where σ^2_i and σ^2_j denote the OLS residual variance of the auto-regressive models estimated for variables i and j. e_x denotes the number of exogenous variables.

Equations (2.8), (2.9) and (2.10) are the standard Minnesota priors. I follow Dieppe

has the same set of explanatory variables (Koop et al., 2010), meaning that if I remove a variable in one equation, that variable would be removed from all equations.

[17] I set the hyperparameters of inverse Wishart distribution in a conventional way: $\alpha_0 = n + 1$ and $S_0 = \mathbb{I}_n$ (Dieppe et al., 2016).

et al. (2016) and introduce hyperparamter (λ_5) in Equation (2.11). λ_5 is applied only on the foreign block and controls the tightness of the distributions of domestic variables in the foreign block. I set $\lambda_5 = 1e\ {-8}$. This small value imposes highly informative priors on the parameters of the domestic variables in the foreign block, where the distribution is centered at zero. I select standard prior variances for the rest of the parameters ($\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 1$). Thus, I implement block exogeneity through a special case of Independent ($\mathcal{NIW}$) priors, where the reduced-form coefficients follow Minnesota priors with an additional hyperparameter. Finally, if the prior distribution over the reduced-form parameters is $\mathcal{NIW}(\alpha_0, \beta_0, S_0, \Omega_0)$, then the posterior distribution over the reduced-form parameters is $\mathcal{NIW}(\tilde{\alpha}, \tilde{\beta}, \tilde{S}, \tilde{\Omega})$, where

$$\begin{aligned}
\tilde{\alpha} &= T + \alpha_0, \\
\tilde{\Omega} &= [\Omega_0^{-1} + Z'VZ]^{-1}, \\
\tilde{\beta} &= \tilde{\Omega}[\Omega_0^{-1}\beta_0 + Z'V\mathbf{Y}] \\
\tilde{S} &= S_0 + (\mathbf{Y} - \mathbf{X}\hat{B})'(\mathbf{Y} - \mathbf{X}\hat{B}),
\end{aligned}$$

where $\mathbf{Y} = [\mathbf{y_1} \cdots \mathbf{y_T}]'$, $\mathbf{X} = [\mathbf{x_1} \cdots \mathbf{x_T}]'$, $V = \Sigma^{-1} \otimes \mathbb{I}_T$, $Z = \mathbb{I}_n \otimes X$ and $\hat{B} = \tilde{\beta} + chol(\tilde{\Omega})' * RAND((n*p+1)*n, 1)$.

2.4 Baseline Model Results

2.4.1 Shock to Commodity Prices

Figure 2.2 plots the impulse response functions to a one-standard deviation (SD) shock to commodity prices. The solid lines are the point-wise posterior median responses while the gray shaded regions are the 68% equal-tailed point-wise posterior probability bands. In response to a positive disturbance in commodity prices, output contracts. Although the median response of domestic output is insignificant a substantial mass of the posterior distribution lies on the negative side, highlighting the transmission channel of a negative commodity supply shock. This finding is in line with the findings of Dungey et al. (2014, 2017, 2020). Moreover, the trade balance deteriorates and the median response stays significantly negative throughout the 5 years horizon period.

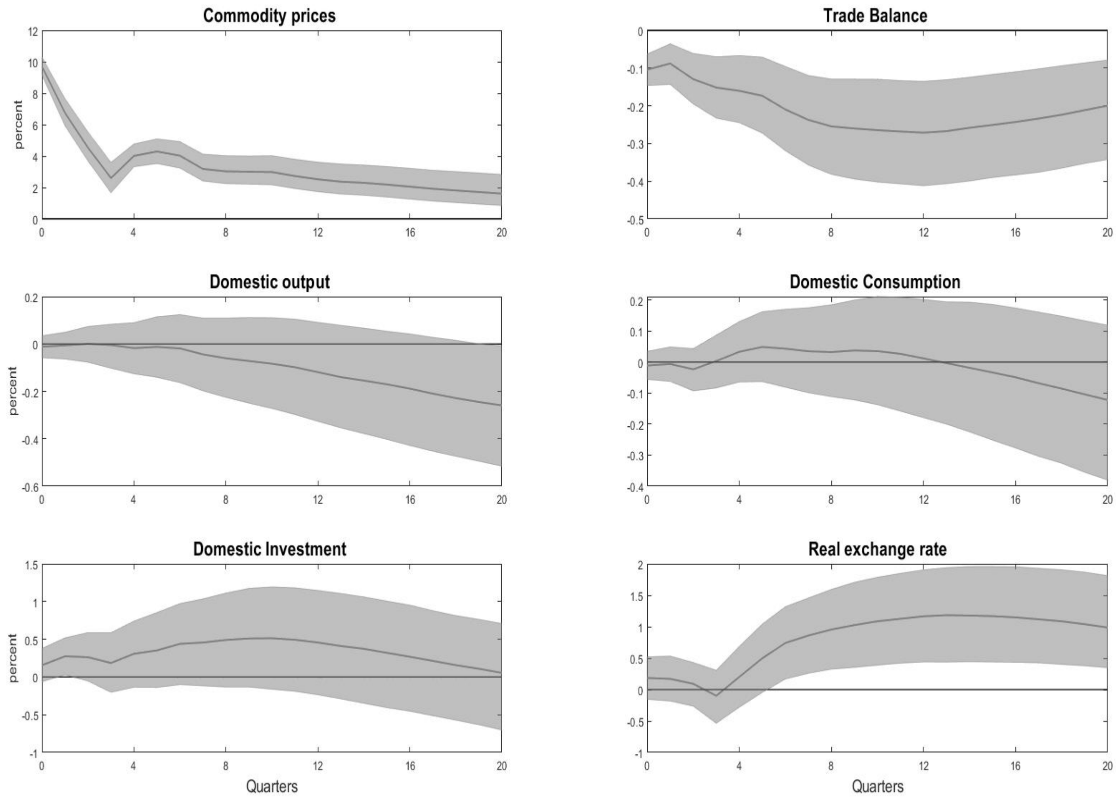


Figure 2.2: Baseline SVAR Model: IRFs to a one standard deviation Innovation in commodity prices. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 66% equal-tailed point-wise posterior probability bands.

Table 2.1 shows where my findings fit into the Australian literature. My finding of a persistent and insignificant contraction in domestic output aligns with Dungey et al. (2014, 2017). Dungey et al. (2020) find persistent and significant contraction in output. These studies use commodity price shocks and study relatively similar periods, capturing the recent commodity price boom. Knop and Vespignani (2014) find a statistically significant contraction in output growth in response to a commodity price shock. Generally, my finding that domestic output contracts in response to positive innovation in commodity prices is in line with the recent Australian literature.

Importantly, my study contributes towards the findings in the literature through the model's trade balance response.

Paper	**Domestic Output**	**TB**
Dungey and Pagan, 2000 (TOT)	$+*$	NA
Dungey and Pagan, 2009 (TOT)	$+$	NA
Liu, 2010 (TOT)	$+$	NA
Jääskelä and Smith, 2013 (CP)	$+ * @$	NA
Dungey et al., 2014 (CP)	$-@$	NA
Dungey et al., 2017 (CP)	$-@$	NA
Dungey et al., 2020 (CP)	$- * @$	NA
Vespignani and Knop, 2014 (CP)	$-*$	NA
Otto, 2003 (TOT)	$+@$	$+(HLM)$
Cashin and McDermott, 2002 (TOT)	NA	$+(HLM)$
Ahmed and Park, 1994 (TOT)	$+@$	neither $+$ nor $-$
My findings	$-@$	$- * @$

Note: $+(-)$ denotes positive (negative) response, $*$ denotes significant, @ denotes highly persistent. TOT and CP stand for terms of trade and commodity prices, respectively.

Table 2.1: Summary of the responses of Output and Trade Balance

I find that in response to a positive commodity price shock, the trade balance deteriorates significantly and persistently. This finding disagrees with the findings of Otto (2003) and Cashin and McDermott (2002). They find evidence of the HLM effect. These studies however do not include the most recent terms of trade boom, which was characterized by it's size and persistence. The evidence of the HLM effect depends crucially on the persistence of the terms of trade shock and likely explains this discrepancy.

In the next section, I extend my baseline SVAR model by including more variables in the *"foreign block"* to make it more in line with the Australian literature. Further, I report the impact brought about by a large *"foreign block"* on Australian output and trade balance.

2.5 Extended-SVAR Model

Most of the Australian literature on the impact of terms of trade and commodity price shocks on Australian macroeconomic fluctuations include a large *"foreign block"* (Dungey and Pagan, 2000, 2009; Jääskelä and Smith, 2013; Dungey et al., 2014, 2017, 2020). As the baseline results differ somewhat from the literature, it is of interest to extend this study by making it more in line with the Australian literature. I draw motivation from Dungey et al. (2014, 2017, 2020) and extend the baseline model. I endogenise the commodity prices by placing it in the *"foreign block"* along with Chinese steel production (csp_t), world output (yw_t) and iron ore exports ($iron_t$). I extend the system in Equation (2.1) from the baseline model, such that y'_{1t} and ϵ'_{1t} become:

$$y'_{1t} = \begin{bmatrix} csp_t & cp_t & yw_t & iron_t \end{bmatrix}$$

$$\epsilon'_{1t} = [\epsilon^{csp} \; \epsilon^{cp} \; \epsilon^{yw} \; \epsilon^{iron}]$$

The extended SVAR model is estimated on the sample period from 1988:Q1 to 2018:Q4. I continue to use data in log-levels.[18] Chinese steel production is used as a proxy for Chinese resource demand because there is no direct measure of Chinese resource demand (Dungey et al., 2014, 2017, 2020). Chinese steel production serves as a demand shock in the system. On the other hand, (positive) commodity prices shock continues to serve as a (negative) supply shock. The world output variable is measured as real GDP (export weighted) of Australia's major trading partners including China. This is the same measure used in Jääskelä and Smith (2013) and Dungey et al. (2014, 2017, 2020). The real value of Australia's iron ore exports is the fourth variable in the *"foreign block"*. This variable in measured as sum of metalliferous ores and metal scrap. This is the same measure as used in Dungey et al. (2020).[19]

Chinese steel production is placed first in the extended-SVAR model. This ordering

[18] Appendix B contains the plots of variables as these enter in the extended SVAR model.
[19] See Appendix B for data description and sources.

helps to capture the impact of demand shocks on the Australian economy as the most recent boom in commodity prices is attributed to the increased demand for iron ore and coal by China. Moreover, this ordering gives Chinese steel production the greatest chance to be an important source of macroeconomic fluctuation as I continue to impose Cholesky decomposition on the contemporaneous identification matrix. This identification strategy continues to reflect the assumption of small open economy nature of the Australian economy, where the Australian variables do not affect the foreign variables but the foreign variables contemporaneously and over time affect the Australian domestic variables. Moreover, the variables in the *"foreign block"* affect each other with a lag but are not affected by the variables in the domestic block. The only exception is iron ore exports, which does get affected by real exchange rate with a lag. I set the lag length to 2. Formally, the identification restrictions on $\mathbf{A_0}$ and $\mathbf{A_L}$ take the form:

$$\mathbf{A_0} = \begin{bmatrix} a_{0,11} & a_{0,12} & a_{0,13} & a_{0,14} & a_{0,15} & a_{0,16} & a_{0,17} & a_{0,18} & a_{0,19} \\ 0 & a_{0,22} & a_{0,23} & a_{0,24} & a_{0,25} & a_{0,26} & a_{0,27} & a_{0,28} & a_{0,29} \\ 0 & 0 & a_{0,33} & a_{0,34} & a_{0,35} & a_{0,36} & a_{0,37} & a_{0,38} & a_{0,39} \\ 0 & 0 & 0 & a_{0,44} & a_{0,45} & a_{0,46} & a_{0,47} & a_{0,48} & a_{0,49} \\ 0 & 0 & 0 & 0 & a_{0,55} & a_{0,56} & a_{0,57} & a_{0,58} & a_{0,59} \\ 0 & 0 & 0 & 0 & 0 & a_{0,66} & a_{0,67} & a_{0,68} & a_{0,69} \\ 0 & 0 & 0 & 0 & 0 & 0 & a_{0,77} & a_{0,78} & a_{0,79} \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & a_{0,88} & a_{0,89} \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & a_{0,99} \end{bmatrix}$$

$$\mathbf{A_L} = \begin{bmatrix} A_{11,L} & A_{12,L} \\ \\ A_{21,L} & A_{22,L} \end{bmatrix}$$

where $A_{11,L}$ is $n_1 \times n_1$, $A_{12,L}$ is $n_1 \times n_2$, $A_{21,L}$ is $n_2 \times n_1$, $A_{22,L}$ is $n_2 \times n_2$, $\mathbf{y'_{1t}}$ is $1 \times n_1$, $\mathbf{y'_{2t}}$ is $1 \times n_2$, ϵ'_{1t} is $1 \times n_1$ and ϵ'_{2t} is $1 \times n_2$; $n_1 + n_2 = n$.

I apply block exogeneity restrictions through informative priors on all elements of $A_{21,L}$ matrices except $a_{L,94}$. Relaxing the restriction on $a_{L,94}$ implies that fluctuations in real exchange rate impacts the iron ore exports though lags (Dungey et al., 2020). $\mathbf{A_L}$ is further expanded below:[20]

$$\mathbf{A_L} = \begin{bmatrix} a_{L,11} & a_{L,12} & a_{L,13} & a_{L,14} & a_{L,15} & a_{L,16} & a_{L,17} & a_{L,18} & a_{L,19} \\ a_{L,21} & a_{L,22} & a_{L,23} & a_{L,24} & a_{L,25} & a_{L,26} & a_{L,27} & a_{L,28} & a_{L,29} \\ a_{L,31} & a_{L,32} & a_{L,33} & a_{L,34} & a_{L,35} & a_{L,36} & a_{L,37} & a_{L,38} & a_{L,39} \\ a_{L,41} & a_{L,42} & a_{L,43} & a_{L,44} & a_{L,45} & a_{L,46} & a_{L,47} & a_{L,48} & a_{L,49} \\ \boldsymbol{a_{L,51}} & \boldsymbol{a_{L,52}} & \boldsymbol{a_{L,53}} & \boldsymbol{a_{L,54}} & a_{L,55} & a_{L,56} & a_{L,57} & a_{L,58} & a_{L,59} \\ \boldsymbol{a_{L,61}} & \boldsymbol{a_{L,62}} & \boldsymbol{a_{L,63}} & \boldsymbol{a_{L,64}} & a_{L,65} & a_{L,66} & a_{L,67} & a_{L,68} & a_{L,69} \\ \boldsymbol{a_{L,71}} & \boldsymbol{a_{L,72}} & \boldsymbol{a_{L,73}} & \boldsymbol{a_{L,74}} & a_{L,75} & a_{L,76} & a_{L,77} & a_{L,78} & a_{L,79} \\ \boldsymbol{a_{L,81}} & \boldsymbol{a_{L,82}} & \boldsymbol{a_{L,83}} & \boldsymbol{a_{L,84}} & a_{L,85} & a_{L,86} & a_{L,87} & a_{L,88} & a_{L,89} \\ \boldsymbol{a_{L,91}} & \boldsymbol{a_{L,92}} & \boldsymbol{a_{L,93}} & a_{L,94} & a_{L,95} & a_{L,96} & a_{L,97} & a_{L,98} & a_{L,99} \end{bmatrix}$$

Hence, the first two equations of the structural model in Equation (2.2) correspond to the commodity demand and commodity supply shocks, respectively:[21]

$$\begin{aligned} csp_t = csp_{t-1} \underbrace{a_{1,11}a_{0,11}^{-1}}_{\psi_{csp_{t-1}}} + csp_{t-2} \underbrace{a_{2,11}a_{0,11}^{-1}}_{\psi_{csp_{t-2}}} + cp_{t-1} \underbrace{a_{1,21}a_{0,11}^{-1}}_{\psi_{cp_{t-1}}} + cp_{t-2} \underbrace{a_{2,21}a_{0,11}^{-1}}_{\psi_{cp_{t-2}}} + \\ yd_{t-1} \underbrace{a_{1,31}a_{0,11}^{-1}}_{\psi_{yd_{t-1}}} + yd_{t-2} \underbrace{a_{2,31}a_{0,11}^{-1}}_{\psi_{yd_{t-2}}} + iron_{t-1} \underbrace{a_{1,41}a_{0,11}^{-1}}_{\psi_{yd_{t-1}}} + \\ iron_{t-2} \underbrace{a_{2,41}a_{0,11}^{-1}}_{\psi_{yd_{t-2}}} + \epsilon_t^{csp} \underbrace{a_{0,11}^{-1}}_{\sigma_{csp}} \end{aligned} \tag{2.12}$$

[20] I apply extremely tight priors on the elements in $\mathbf{A_L}$ matrices, highlighted in *bold*.

[21] I abstract from the lags of domestic variables as an implication of exogeneity restrictions applied through tight priors. I also omit constant terms for simplicity.

$$
\begin{aligned}
cp_t = -csp_t \underbrace{a_{0,12}a_{0,22}^{-1}}_{\psi_{csp_t}} + csp_{t-1} \underbrace{a_{1,12}a_{0,22}^{-1}}_{\psi_{csp_{t-1}}} + csp_{t-2} \underbrace{a_{2,12}a_{0,22}^{-1}}_{\psi_{csp_{t-2}}} + cp_{t-1} \underbrace{a_{1,22}a_{0,22}^{-1}}_{\psi_{cp_{t-1}}} + \\
cp_{t-2} \underbrace{a_{2,22}a_{0,22}^{-1}}_{\psi_{cp_{t-2}}} + yd_{t-1} \underbrace{a_{1,32}a_{0,22}^{-1}}_{\psi_{yd_{t-1}}} + yd_{t-2} \underbrace{a_{2,32}a_{0,22}^{-1}}_{\psi_{yd_{t-2}}} + \\
iron_{t-1} \underbrace{a_{1,42}a_{0,22}^{-1}}_{\psi_{iron_{t-1}}} + iron_{t-2} \underbrace{a_{2,42}a_{0,22}^{-1}}_{\psi_{iron_{t-2}}} + + \epsilon_t^{cp} \underbrace{a_{0,22}^{-1}}_{\sigma_{cp}}
\end{aligned} \tag{2.13}
$$

where $a_{L,11}$ and $a_{L,22}$ denote the elements (1,1) and (2,2) of $\mathbf{A_L}$, while σ_{csp} and σ_{cp} denote the elements (1,1) and (2,2) of $\mathbf{A_0}$, respectively. As a consequence, the first two elements of ϵ_t, denoted by ϵ_t^{csp} and ϵ_t^{cp} have the interpretation of a commodity demand and supply shocks, respectively.[22]

2.5.1 Shock to Chinese Steel Production

From Figure 2.3, we see that a one-SD innovation in the Chinese steel production proxied as a demand shock results in a highly persistent and significant increase in steel production in China, commodity prices, world output and iron ore exports. The demand shock acts as a stimulus to domestic output and investment. My main finding is a persistent but insignificant expansion in domestic output. However, a major mass of the posterior distribution lies on the positive side, which shows an expansionary response of domestic output in response to the demand shock. The positive Chinese resource demand is transmitted through the investment channel that results in an expansion of domestic output. The expansion in domestic investment may be attributed to an increase in investment in the booming mining industry. This could result in the phenomenon called the "Dutch Disease", where the resources in the economy are transferred from the the lagging sector (exports not part of the booming sector and import-competing goods and services) and the non-tradeable sector to the booming

[22] Following Arias et al. (2019), I normalize the IRFs by imposing sign-restrictions such that Chinese steel production and commodity prices respond positively on impact in response to positive innovation in Chinese steel production and commodity prices, respectively, and that $\sigma_{csp}, \sigma_{cp} > 0$.

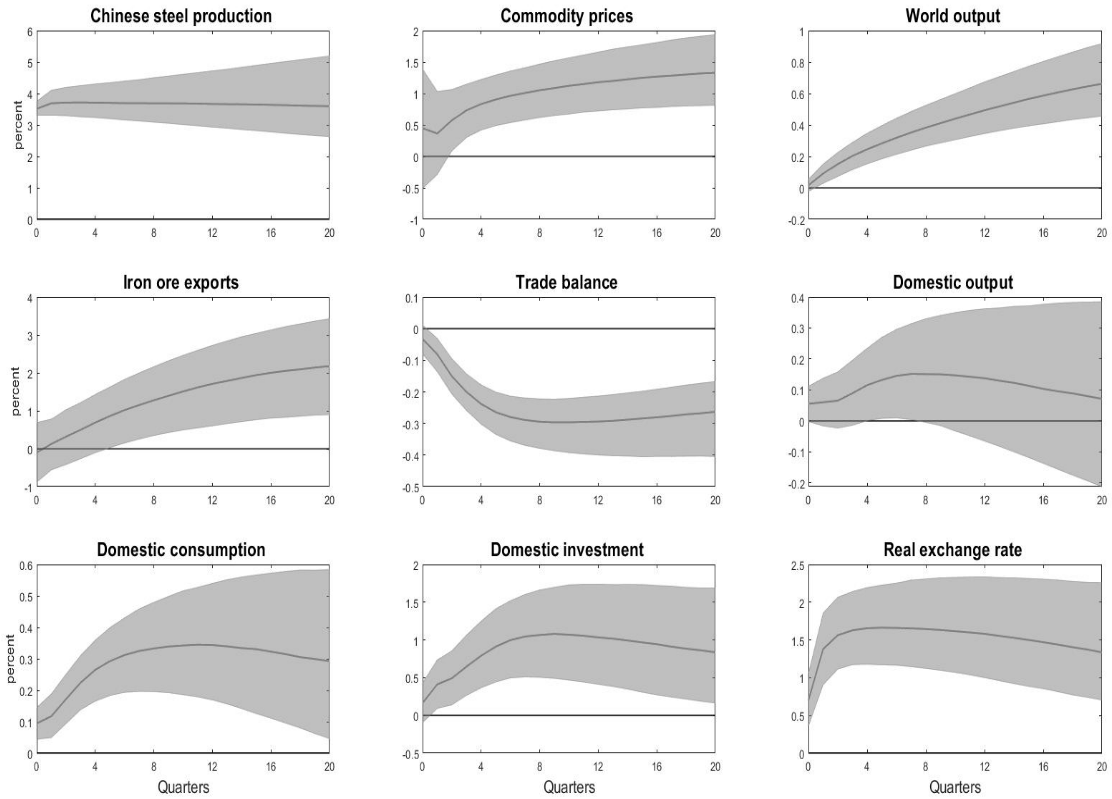


Figure 2.3: Extended SVAR Model: IRFs to a one standard deviation Innovation in Chinese steel production. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 66% equal-tailed point-wise posterior probability bands.

sectors due to the real exchange rate appreciation (Corden, 2012).[23,24]

My other important finding is a persistent and significant deterioration in trade balance through the exchange rate transmission mechanism. The exchange rate appreciates because of the increase in the commodity prices and iron ore exports while the trade balance deteriorates. The result suggests that an increase in the iron ore exports is offset by an increase in imports due to strengthening of the Australian dollar, thus, we observe a deterioration in the trade balance. As the exchange rate depreciates and

[23] At this point, my work does not focus on the sector/industry level data, however, I intend to extend this body of work to address the interesting question of Dutch Disease for Australia.

[24] Bjørnland et al. (2019) show that a resource boom resulting from increased oil production (for instance, in case of Norway) also significantly increases productivity in other industries, including manufacturing.

the real value of the iron ore exports persistently increases, we see signs of improvement in the trade balance. However, the trade balance remains negative, highly persistent and significant throughout the horizon.

2.5.2 Shock to Commodity Prices

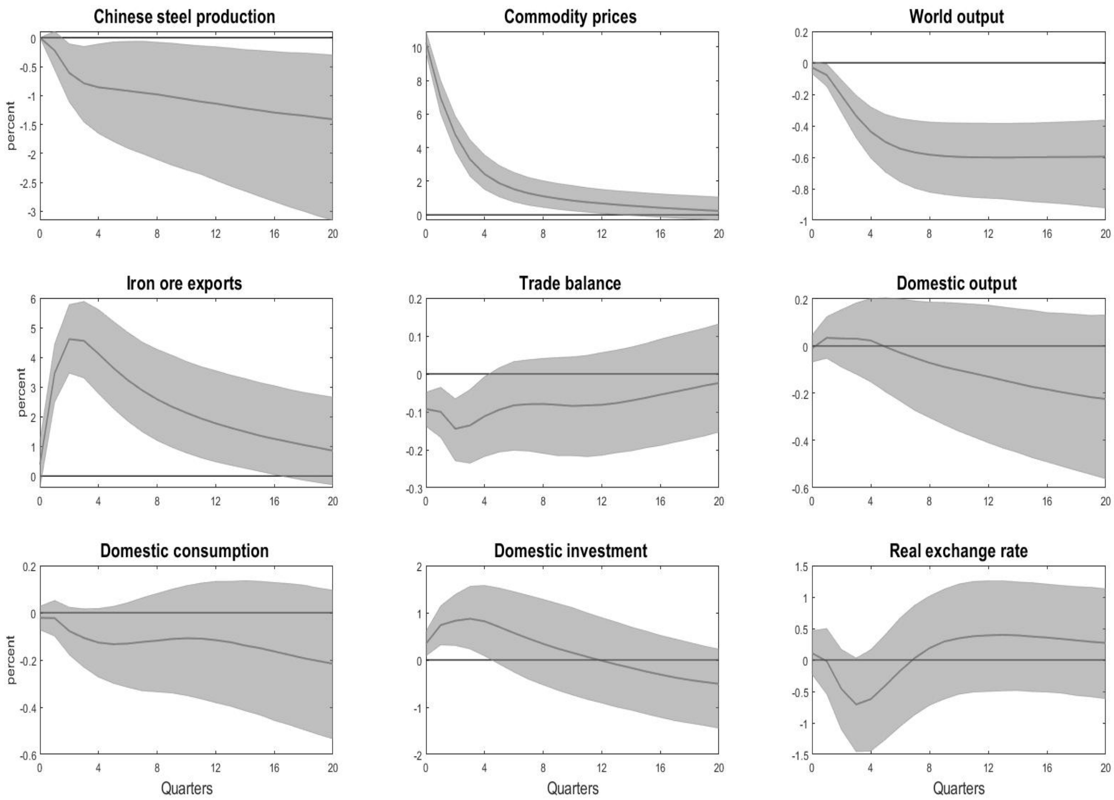


Figure 2.4: Extended SVAR Model: IRFs to a one standard deviation Innovation in Commodity prices. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 66% equal-tailed point-wise posterior probability bands.

From Figure 2.4, we see that a one-SD innovation in the commodity prices proxied as a (negative) supply shock results in a significant increase in Australian commodity prices and a significant and persistent decrease in Chinese steel production and world output. Further, I find a contemporaneous increase in the real value of the iron ore exports and domestic investment. Both the iron ore exports and domestic investment peak in the first year before falling towards the baseline. The short-lived increase in domestic investment, unlike the significant and persistent increase in domestic investment to the demand shocks in the previous section, is attributed to the less persistent incearse in commodity prices compared to the response of commodity prices to demand shocks.

The real exchange rate depreciates during the first year because of substantial decrease in Chinese steel production and world output. This puts a downward pressure on the demand of Australian dollar offsetting the impact of increasing commodity prices and iron ore exports. However, the real exchange rate appreciates 1 year after the shock which coincides with a fall in the real value of the iron ore exports. The combined effect of an appreciation in the real exchange rate and a decrease in the real value of the iron ore exports results in a significant deterioration of the trade balance.

The (negative) commodity supply shock results in an insignificant contraction in domestic output. The contraction in domestic output is very small, around 0.2% below the baseline. Dungey et al. (2014, 2017, 2020) also find a contraction in domestic output. This result is attributed to an expansion in output of resource sector (iron ore) which results in the movement of factors of production from non-resource to resource sector, where higher prices are paid for labour and capital. The fall in non-resource sector output is not fully offset by the rise in resource sector output (Dungey et al., 2014).

2.6 Robustness Check: Extended-SVAR Model with Log-Linear Detrending

This robustness check is motivated by Australian literature (Dungey and Pagan, 2000, 2009; Dungey et al., 2014, 2017, 2020). I log-linearly detrend all the variables except the trade balance. Trade balance is divided by the trend component of domestic

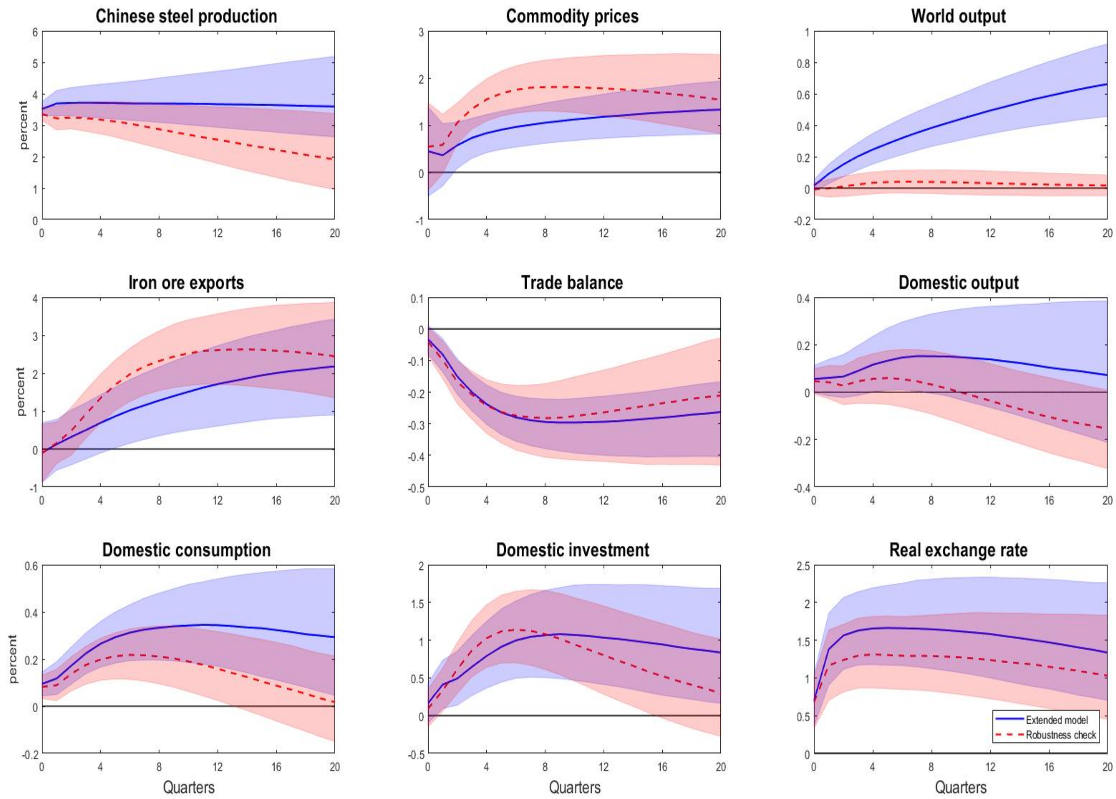


Figure 2.5: Robustness check: IRFs to a one standard deviation Innovation in Chinese steel production. *Note:* The blue (solid) and red (dashed) lines depict the point-wise posterior median responses of extended model and robustness check respectively. The shaded bands represent the 66% equal-tailed point-wise posterior probability bands.

output.[25] I continue to use the same sample from 1988:Q1-2018:Q4 with lag length of $p = 2$. Further, I continue with the same identification strategy as explained in section 2.3.

From Figure 2.5, we see a highly persistent and significant deterioration of the trade balance in response to a commodity demand shock. However, the response of domestic output is less persistent as compared to the response of domestic output in my extended model in section 2.5. On impact we see a negligible and insignificant expansion in domestic output.

[25] Appendix B contains the plots of variables as these enter in the SVAR model for robustness analysis.

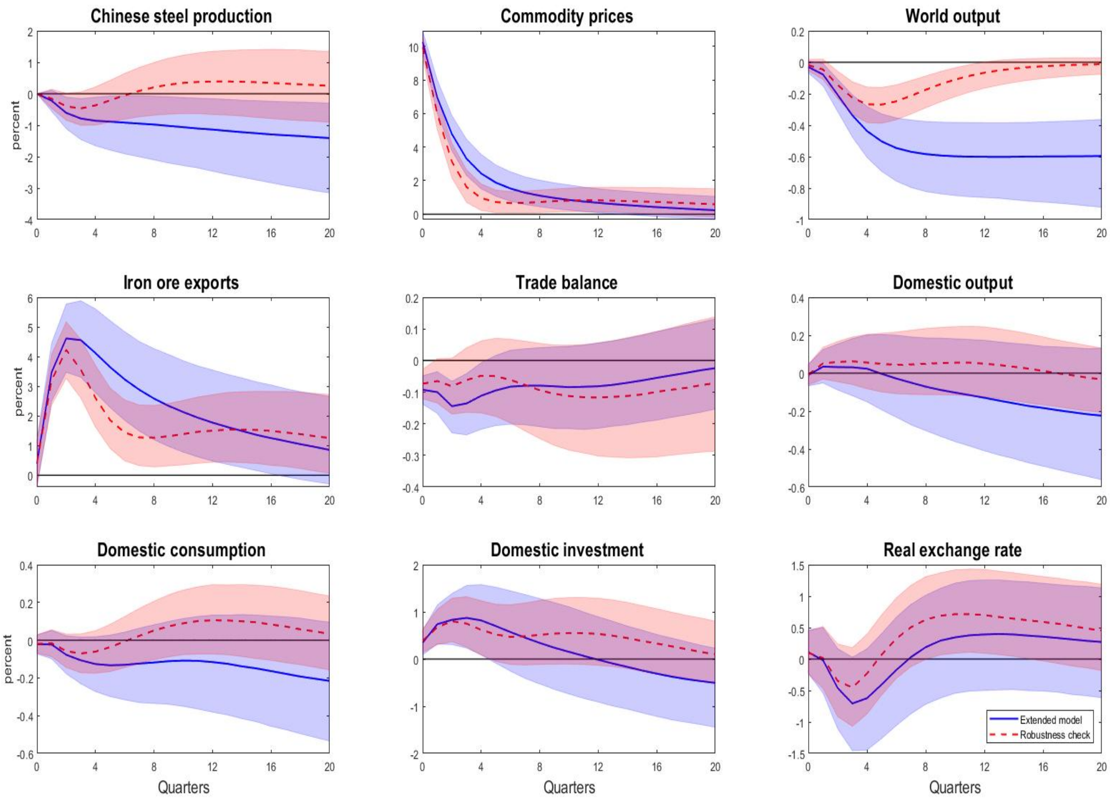


Figure 2.6: Robustness check: IRFs to a one standard deviation Innovation in Commodity prices. *Note:* The blue (solid) and red (dashed) lines depict the point-wise posterior median responses of extended model and robustness check respectively. The shaded bands represent the 66% equal-tailed point-wise posterior probability bands.

Figure 2.6 shows that in response to the commodity supply shock, trade balance persistently deteriorates with a substantial mass of posterior distribution lying on the negative side. On the other hand, the response of domestic output is muted compared to the response of domestic output in my extended model in section 2.5.

The responses of output and trade balance and other variables to commodity demand and supply shocks suggest that my results of section 2.5 are generally robust to alternative detrending method.

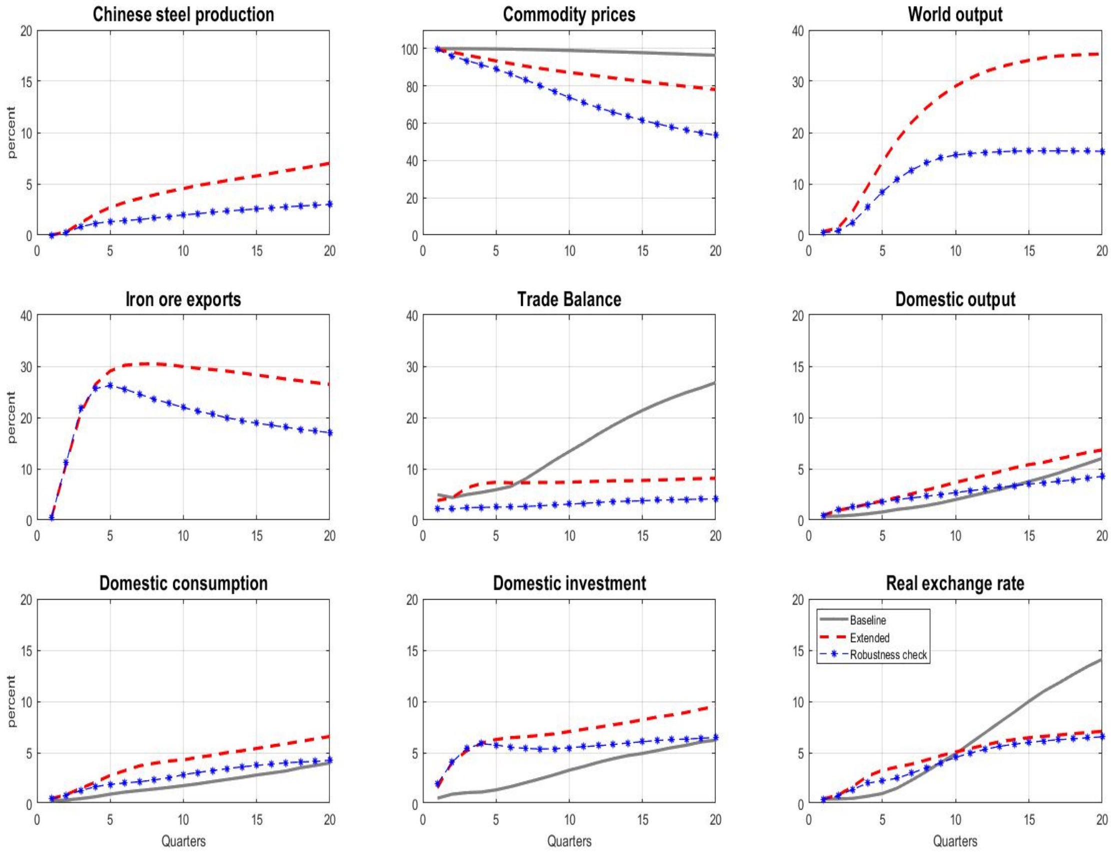


Figure 2.7: Forecast Error Variance Decomposition to commodity price shocks for baseline model, extended model and robustness check.

2.7 Variance Decomposition Analysis

In this section, I consider the contribution of commodity supply and commodity demand disturbances to macroeconomic fluctuations through variance decomposition analysis. Figure 2.7 shows the variance decomposition for the baseline model, extended model and the robustness check where results are reported for 20 quarters forecast horizon. The main findings are that the commodity price shocks contribute between 4% and 7% of the variation in domestic output. On the other hand, commodity price shocks contribute around 25% of the variation in trade balance in the baseline model. In the extended SVAR model, shocks to commodity prices contribute around 8% of the variation in trade balance. According to Schmitt-Grohé and Uribe (2018) median estimates, terms of trade shocks account for roughly 10% of both the variance of output

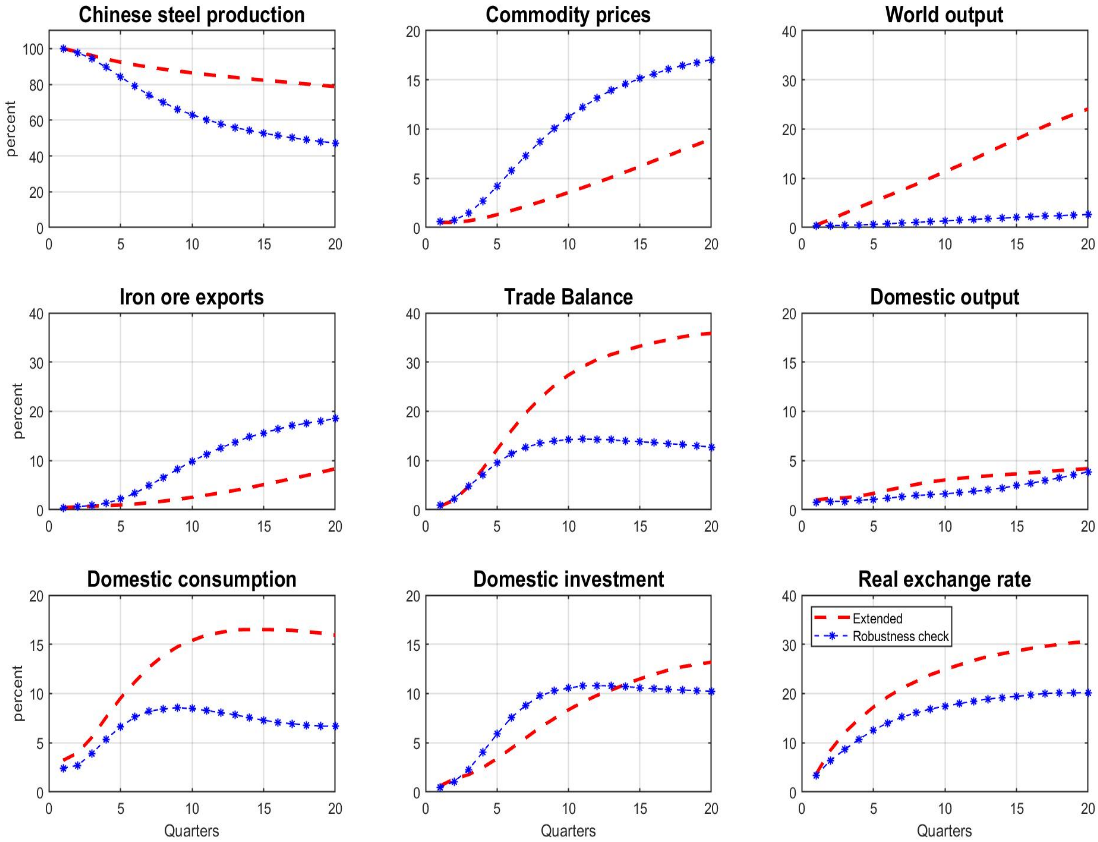


Figure 2.8: Forecast Error Variance Decomposition to Chinese steel production shocks for extended model and robustness check.

and that of the trade balance. Jääskelä and Smith (2013) and Dungey et al. (2017, 2020) report variance in domestic output ranges between 5% and 9% in response to a (negative) commodity supply shock. Moreover, from Figure 2.8 we see that a (positive) commodity demand shock contributes around 12% to 35% of the variations in trade balance while its contribution towards domestic output is around 4%. Dungey et al. (2014, 2017, 2020) report variance in domestic output ranges between 3% and 5% in response to a commodity demand shock.

The variance decomposition analysis shows that, generally, (negative) commodity supply and (positive) commodity demand shocks contribute only a small amount to output and trade balance fluctuations. The only exception is the trade balance variance (35%) to Chinese resource demand disturbances. These are interesting findings particularly considering the popular wisdom on the impact of the commodities boom on the Australian macroeconomy.

2.8 Conclusion

The Australian literature presents a mixed view on the impact of commodity prices and terms of trade shocks on macroeconomic fluctuations. In this chapter, I quantify the impact of commodity demand and supply shocks on macroeconomic fluctuations in Australia. In my baseline model a (positive) commodity price shock acts as a proxy for (negative) commodity supply shock. My model specification is inspired by Drechsel and Tenreyro (2018), Schmitt-Grohé and Uribe (2018) and Di Pace et al. (2020). I estimate a structural vector autoregression (SVAR) model over the period 1983:Q3 to 2018:Q4 to quantify the impact of exogenous disturbances in commodity prices on output and the trade balance using the Bayesian estimation techniques of Arias et al. (2019). I achieve identification through Cholesky decomposition of the contemporaneous matrix with commodity prices ordered first. I apply exclusion restrictions on the structural parameters of the contemporaneous matrix. Moreover, I apply extremely tight priors on the lagged parameters of the commodity price equation. My identification scheme that combines Cholesky decomposition and highly informative priors ensure that the commodity prices acts exogenous to the system. For my baseline model, I find that a positive commodity price shock results in a contraction of output and a significant and persistent deterioration of the trade balance. Moreover, I find a negligible contribution of commodity price shocks to output fluctuations while these disturbances account for around 25% of the variance of the trade balance.

I extend my baseline model by introducing a large *"foreign block"*. I also, extend the exogeneity structure of my baseline model by applying additional exclusion restrictions on the contemporaneous matrix and highly informative priors on the lagged matrix. The combination of exclusion restrictions and informative priors help is implementing block exogeneity in my model. One exception in the block exogeneity structure is relaxing the restriction where the exchange rate impacts the iron ore exports through lags. I estimate my extended SVAR model over the period 1988:Q1 to 2018:Q4 to quantify the impact of commodity supply and demand disturbances on output and trade balance. Similar to the baseline model, (positive) commodity price shocks act as (negative) commodity supply shocks while (positive) Chinese steel production shocks act as (positive) commodity demand shocks. I find that in response to the commodity demand and supply shocks, the responses of domestic output does not deviate significantly from baseline while trade balance deteriorates on impact and remains below trend for the subsequent 5 years. These findings provide further evidence on the response of the Australian macroeconomy to international shocks and allow policymakers to consider other contributions to Australian economic growth over the past three decades.

Appendix B

Appendix to Chapter 2

B.1 Data Sources

The data comes from the Australian Bureau of Statistics (ABS), Reserve Bank of Australia (RBA), the World Bank, and Datastream.

Commodity Prices (cp_t)
Index of commodity prices in Australian dollar-All items. *Source*: Reserve Bank of Australia, Statistical Table I2.

Population
The working age population is constructed by summing the total number of residents between age 15 and 64. *Source*: Australian Demographics Statistics, ABS Cat. No. 3101.0.

Domestic Output (yd_t)
Real gross domestic product, expressed in chain volume measure. *Source*: Australian National Accounts: National Income, Expenditure and Product, ABS Cat. No. 5206.0.

Consumption (c_t)
Final households consumption, expenditure expressed in chain volume measure. *Source*: Australian National Accounts: National Income, Expenditure and Product, ABS Cat. No. 5206.0.

Investment (i_t)
Gross fix capital formation, expressed in chain volume measure. *Source*: Australian National Accounts: National Income, Expenditure and Product, ABS Cat. No. 5206.0.

Exports (x_t) and **Imports** (m_t)
Exports of goods and services and imports of goods and services, expressed in chain volume measure. *Source*: Australian National Accounts: National Income, Expenditure and Product, ABS Cat. No. 5206.0.

Real exchange rate-Trade weighted index (RER_t)
Australian dollar trade-weighted exchange rate index, adjusted for relative consumer price levels. *Source*: Reserve Bank of Australia, Statistical Table F15.

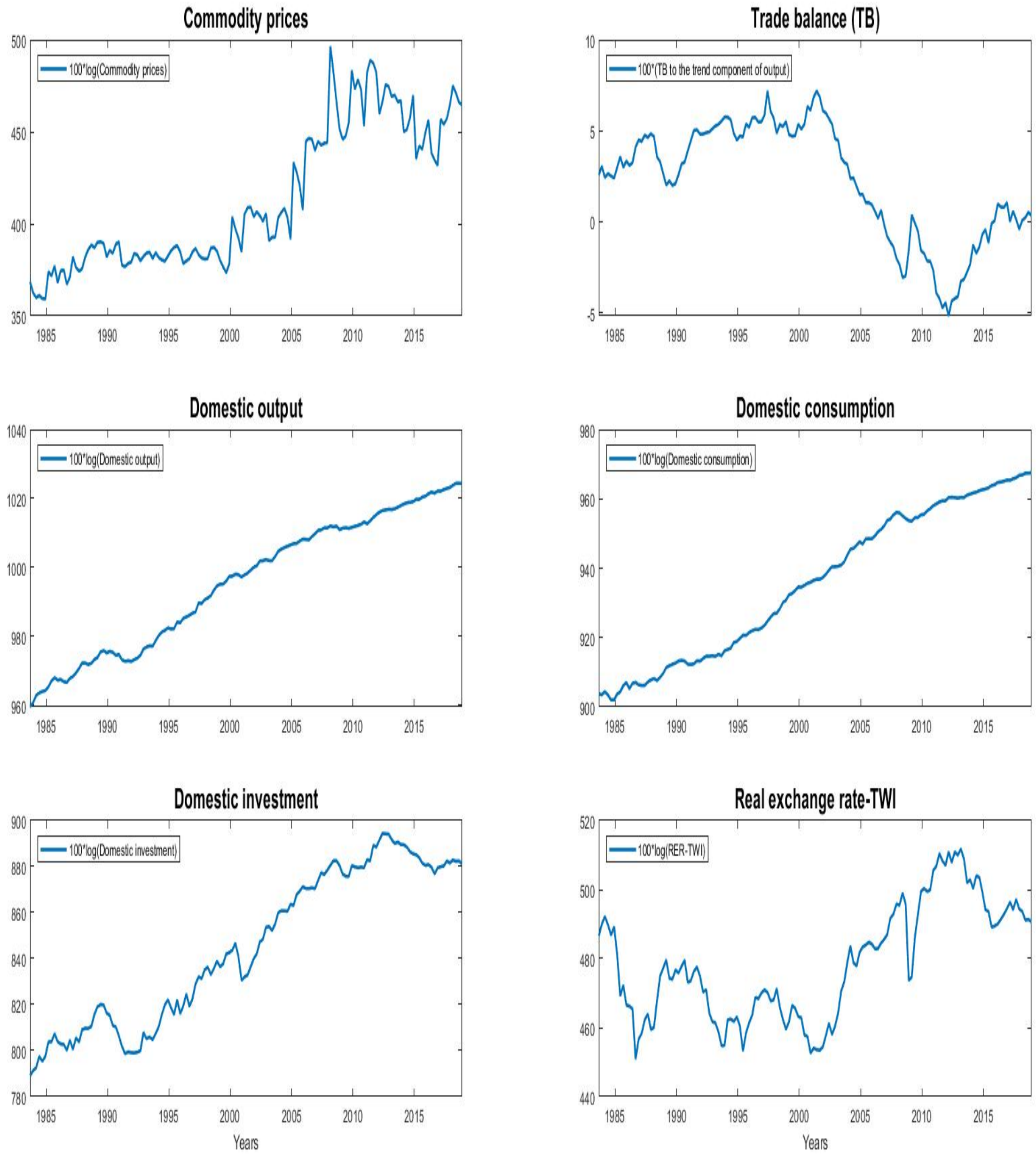
Chinese Steel Production (csp_t)
Source: Datastream (code: CHVALSTLH), Seasonally Adjusted using Census X-13 EViews.

Iron Ore Exports ($Iron_t$)
Iron ore exports is constructed using sum of metalliferous ores and metal scrap (*Source*: International Trade in Goods and Services, ABS Cat 5368.0) deflated by the consumer price index (*Source*: Reserve Bank of Australia), Seasonally Adjusted using Census X-13 EViews.

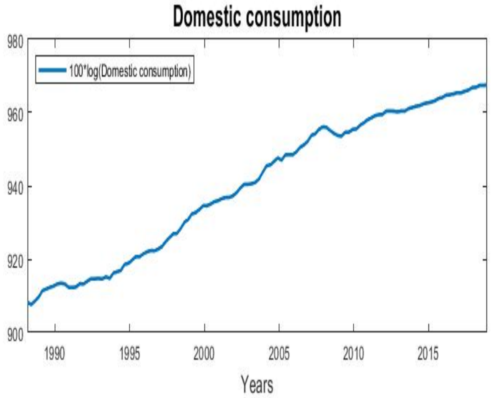
World Output (yw_t)
Real GDP (export weighted) of Australia's Major Trading Partners. (*Source*: Reserve Bank of Australia).

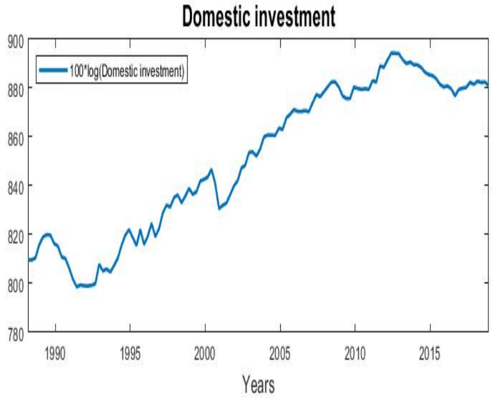
B.2 Plots of variables

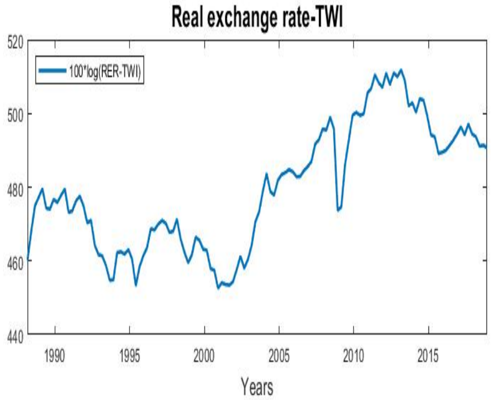
B.2.1 Plots of variables for baseline model



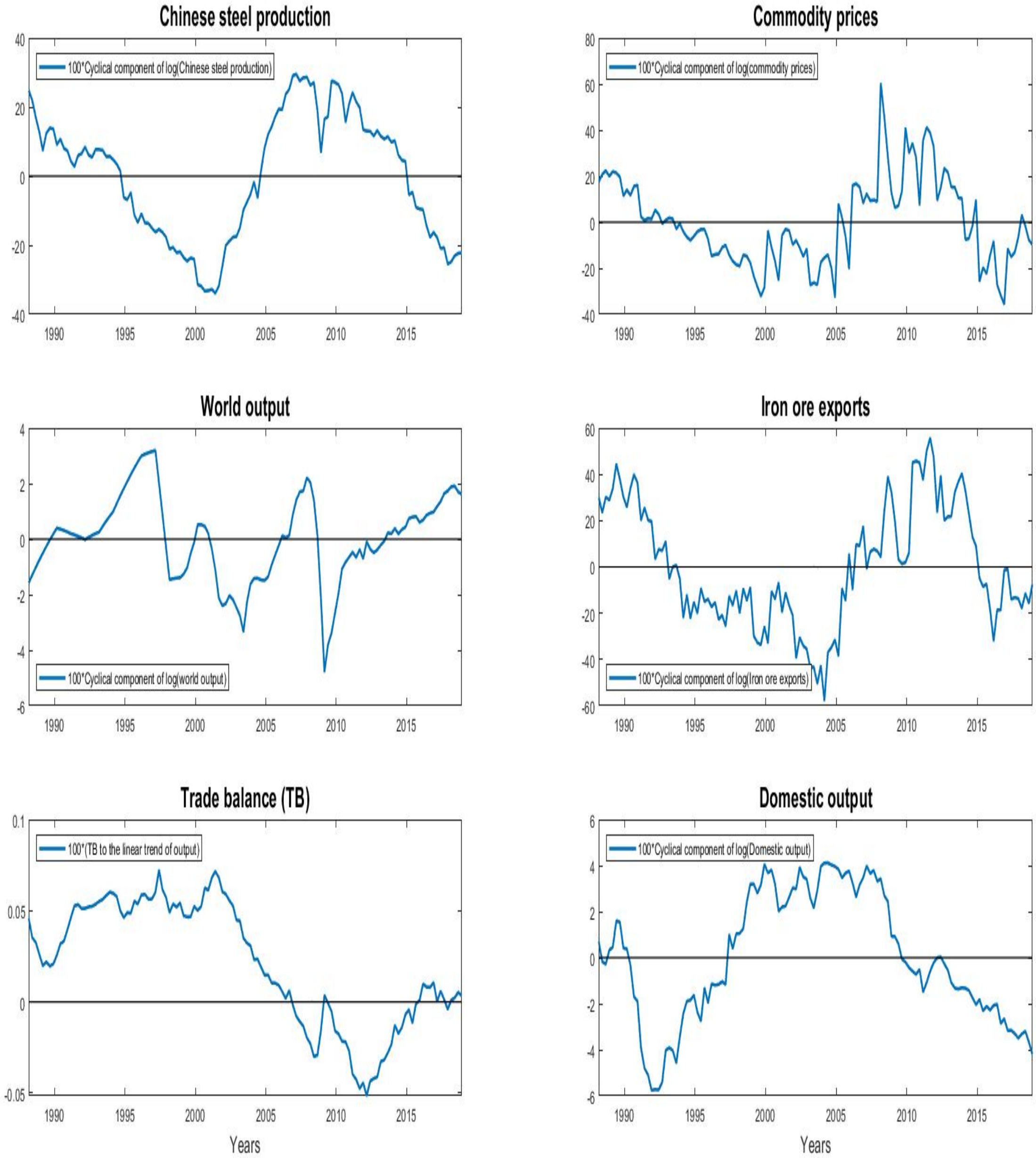
B.2.2 Plots of variables for extended model

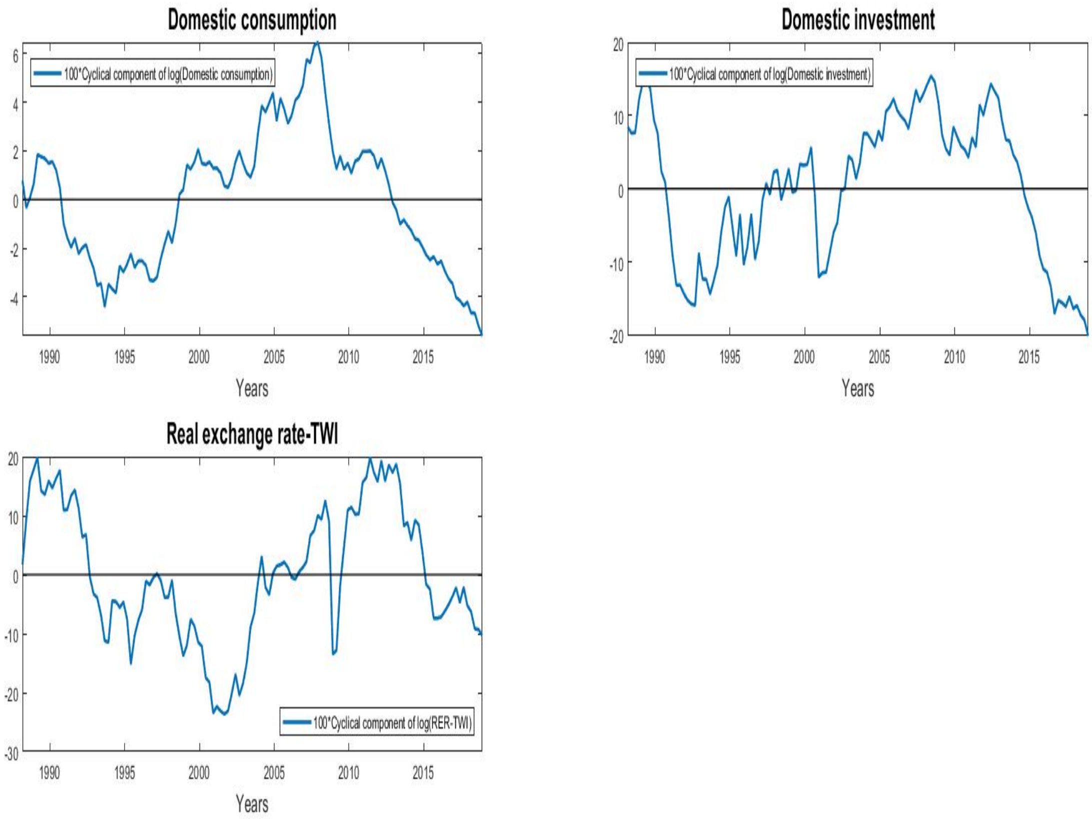

Domestic consumption
100*log(Domestic consumption)
980
960
940
920
900
1990
1995
2000
2005
2010
2015
Years


Domestic investment
100*log(Domestic investment)
900
880
860
840
820
800
780
1990
1995
2000
2005
2010
2015
Years


Real exchange rate-TWI
100*log(RER-TWI)
520
500
480
460
440
1990
1995
2000
2005
2010
2015
Years

B.2.3 Plots of variables for robustness check




Domestic consumption
100*Cyclical component of log(Domestic consumption)
Years
Domestic investment
100*Cyclical component of log(Domestic investment)
Years
Real exchange rate-TWI
100*Cyclical component of log(RER-TWI)
Years

References

Ahmed, S. and Park, J. H. (1994). Sources of macroeconomic fluctuations in small open economies. *Journal of Macroeconomics*, 16(1):1–36.

Arias, J. E., Caldara, D., and Rubio-Ramirez, J. F. (2019). The systematic component of monetary policy in SVARs: An agnostic identification procedure. *Journal of Monetary Economics*, 101:1–13.

Arias, J. E., Rubio-Ramírez, J. F., and Waggoner, D. F. (2018). Inference based on structural vector autoregressions identified with sign and zero restrictions: Theory and applications. *Econometrica*, 86(2):685–720.

Arias, J. E., Rubio-Ramírez, J. F., and Waggoner, D. F. (2021). Inference in Bayesian Proxy-SVARs. *Journal of Econometrics*, 225(1):88–106.

Bjørnland, H. C., Thorsrud, L. A., and Torvik, R. (2019). Dutch disease dynamics reconsidered. *European Economic Review*, 119:411–433.

Blanchard, O. J. and Quah, D. (1989). The dynamic effects of aggregate demand and aggregate supply. *The American Economic Review*, 79(4):655–673.

Cashin, P. and McDermott, C. J. (2002). Terms of trade shocks and the current account: Evidence from five industrial countries. *Open Economies Review*, 13:219–235.

Corden, W. M. (2012). Dutch disease in Australia: policy options for a three-speed economy. *Australian Economic Review*, 45(3):290–304.

Cushman, D. O. and Zha, T. (1997). Identifying monetary policy in a small open economy under flexible exchange rates. *Journal of Monetary economics*, 39(3):433–448.

Di Pace, F., Juvenal, L., and Petrella, I. (2020). Terms-of-trade shocks are not all alike.

Dieppe, A., Legrand, R., and Roye, B. V. (2016). The BEAR toolbox (ECB Working PaperNo. 1934). *European Central Bank*.

Drechsel, T. and Tenreyro, S. (2018). Commodity booms and busts in emerging economies. *Journal of International Economics*, 112:200–218.

Dungey, M., Fry-McKibbin, R., and Linehan, V. (2014). Chinese resource demand and the natural resource supplier. *Applied Economics*, 46(2):167–178.

Dungey, M., Fry-Mckibbin, R., and Volkov, V. (2020). Transmission of a resource boom: the case of Australia. *Oxford Bulletin of Economics and Statistics*, 82(3):503–525.

Dungey, M. and Pagan, A. (2000). A structural VAR model of the Australian economy. *Economic record*, 76(235):321–342.

Dungey, M. and Pagan, A. (2009). Extending a SVAR model of the Australian economy. *Economic Record*, 85(268):1–20.

Dungey, M. H., Fry-McKibbin, R., Todoroski, V., and Volkov, V. (2017). Recovery from dutch disease. *CAMA Working Paper No. 69/2017*.

Jääskelä, J. P. and Smith, P. (2013). Terms of trade shocks: what are they and what do they do? *Economic Record*, 89(285):145–159.

Kilian, L. and Lütkepohl, H. (2017). *Structural vector autoregressive analysis*. Cambridge University Press.

Knop, S. J. and Vespignani, J. L. (2014). The sectorial impact of commodity price shocks in Australia. *Economic Modelling*, 42:257–271.

Koop, G., Korobilis, D., et al. (2010). Bayesian multivariate time series methods for empirical macroeconomics. *Foundations and Trends® in Econometrics*, 3(4):267–358.

Lawson, J., Rees, D., et al. (2008). Sensitivity Analysis— RDP 2008-01: A Sectoral Model of the Australian Economy. *Reserve Bank of Australia Research Discussion Papers*, (May).

Liu, P. (2010). The effects of international shocks on Australia's business cycle. *Economic Record*, 86(275):486–503.

Otto, G. (2003). Terms of trade shocks and the balance of trade: there is a Harberger-Laursen-Metzler effect. *Journal of International Money and Finance*, 22(2):155–184.

Rubio-Ramirez, J. F., Waggoner, D. F., and Zha, T. (2010). Structural vector autoregressions: Theory of identification and algorithms for inference. *The Review of Economic Studies*, 77(2):665–696.

Schmitt-Grohé, S. and Uribe, M. (2018). How important are terms-of-trade shocks? *International Economic Review*, 59(1):85–111.

Statement of Authorship

Title of Paper	High-Frequency Identification of US monetary shocks using Proxy-SVAR models - A Survey of the Literature
Publication Status	☐ Published ☐ Accepted for Publication ☐ Submitted for Publication ☑ Unpublished and Unsubmitted w ork w ritten in manuscript style
Publication Details	

Principal Author

Name of Principal Author (Candidate)	Naveed Javed		
Contribution to the Paper			
Overall percentage (%)	100%		
Certification:	This paper reports on original research I conducted during the period of my Higher Degree by Research candidature and is not subject to any obligations or contractual agreements with a third party that would constrain its inclusion in this thesis. I am the primary author of this paper.		
Signature		Date	12.09.2023

Co-Author Contributions

By signing the Statement of Authorship, each author certifies that:

i. the candidate's stated contribution to the publication is accurate (as detailed above);

ii. permission is granted for the candidate in include the publication in the thesis; and

iii. the sum of all co-author contributions is equal to 100% less the candidate's stated contribution.

Name of Co-Author			
Contribution to the Paper			
Signature		Date	

Name of Co-Author			
Contribution to the Paper			
Signature		Date	

Please cut and paste additional co-author panels here as required.

Chapter 3

High-Frequency Identification of US monetary shocks using Proxy-SVAR models - A Survey of the Literature*

Naveed Javed†

* I would like to thank Jonas Arias for providing me with the MATLAB codes of Arias et al. (2021).
† School of Economics, University of Adelaide, naveed.javed@adelaide.edu.au

Abstract

This chapter contains a literature review of SVAR-IV (instrument variable) models, also known as proxy-SVAR models. This chapter also includes an application of a proxy-SVAR model using high-frequency monetary policy instruments for the US. The literature survey is organised into two parts. The first part provides an overview of the literature on proxy-SVAR models. This part summarises the theoretical aspects and different methodologies used in the literature to identify the proxy-SVAR models. The second part explores the instrument variables used in the literature to identify the US monetary policy shocks. I document how the various types of high-frequency and information-robust monetary policy instruments evolved overtime. Finally, I culminate with an application comparing the two most recent US monetary policy instruments using proxy-SVAR models in Bayesian settings. The application emphasises the superiority of an information-robust monetary instrument.

Keywords: Structural vector autoregressions, External instruments, Monetary policy.

JEL codes **C32, E52, E58**.

3.1 Introduction

Proxy-structural vector autoregression (SVAR) models are gaining importance in the modern empirical macro literature. Proxy-SVAR models are used to quantify the impact of monetary and fiscal shocks on macro-economic variables by achieving identification through external instrument(s). One or more external instruments are used to identify a proxy-SVAR model. The identification scheme in an SVAR-IV model is based on two conditions: i) *relevance condition:* the external instruments are correlated with the structural shocks of interest and ii) *exogeneity condition:* uncorrelated with the rest of the structural shocks (Miranda-Agrippino and Ricco, 2023; Giacomini et al., 2022; Arias et al., 2021; Swanson, 2021; Mertens and Montiel Olea, 2018; Stock and Watson, 2018; Caldara and Kamps, 2017; Lunsford, 2015; Mertens and Ravn, 2014, 2013). Moreover, recent studies have constructed high-frequency external instruments for the US monetary shocks to achieve identification in proxy-SVAR models (Bauer and Swanson, 2023a,b; Sastry, 2022; Miranda-Agrippino and Ricco, 2021; Jarociński and Karadi, 2020; Lunsford, 2020; Caldara and Herbst, 2019; Nakamura and Steinsson, 2018; Gertler and Karadi, 2015).

This chapter reviews both the methodological literature on proxy-SVARs and the literature on various types of external instruments used to identify US monetary policy shocks. I also provide an application of proxy-SVAR model in Bayesian setting using the recent methodology developed by Arias et al. (2021). I achieve identification using the most recent proxy constructed by Jarociński and Karadi (2020) that accounts for the information content in US monetary policy shocks. Further, I compare the results obtained from a proxy-SVAR model using a standard proxy that does not adjust for the information effects (Gertler and Karadi, 2015) and an information-robust proxy (Jarociński and Karadi, 2020). I find that, in response to a contractionary US monetary shock, the fall in real GDP and prices is more pronounced when I use the information-robust proxy as compared to when I use the standard proxy. Thus, my findings emphasis that information-robust proxy is a better instrument for US monetary shocks as compared to the standard proxy.

The rest of the chapter is divided into three sections. Section 3.2 surveys the literature on proxy-SVARs and provides an overview of different proxies used to measure US monetary shocks. Section 3.3 contains the empirical application and section 3.4 concludes.

3.2 Survey of the Literature

3.2.1 Proxy-SVAR literature

The idea of achieving SVAR identification through external instrument is conceived by Stock and Watson (2008), which is subsequently put into practice by a large stream of empirical macro literature. The idea implies that in a multi-variable SVAR model with, for instance, one instrument and one shock of interest, the identification is achieved through *relevance condition*: the instrument is correlated with the shock of interest, and *exogeneity restrictions*: the instrument is uncorrelated with the rest of the shocks in the system. Subsequently, Stock and Watson (2012) applies a method similar to SVAR-IV when they use dynamic factor model (DFM) to examine the macroeconomic dynamics of the US economy during the global financial crisis and the subsequent recovery. Similar to SVAR-IV identification method, they identify six shocks to oil markets, monetary policy, productivity, uncertainty, liquidity and financial risk, and fiscal policy, using seventeen external instruments.

To my knowledge, the earliest study that applies external instrument identification approach in an SVAR model is by Mertens and Ravn (2013). They exploit the features of both SVARs and narrative identification approach to estimate the impact of personal and corporate income tax shocks on US macroeconomy for a quarterly sample from 1950:Q1 to 2006:Q4. They use narrative measures of tax changes as the proxy in a SVAR model, where the proxies are correlated with the tax shocks and uncorrelated with other structural shocks. They construct the narrative measures of tax changes by decomposing the Romer and Romer (2010) account of changes in federal

US tax liabilities into changes in personal and corporate income tax liabilities. They achieve identification with two proxies for two structural shocks, however, their model can handle multiple proxies for multiple structural shocks. Moreover, they argue that in case of one proxy, the zero restrictions that implement exogeneity restrictions are sufficient. However, when there are more than one proxies, we would require additional exclusion restrictions on covariance matrix to achieve point identification of the impulse responses. Finally, they construct an additional structure in order to formally measure the statistical reliability of the process as measurement of the shock of interest. This helps in assessing the relevance of the shock. They appreciate that narratively identified shocks suffer from measurement errors. They propose a structure based on some additional assumptions to formally measure the reliability of the narrative series as a proxy of the shock of interest. Low values of these reliability statistics show that the proxies may not contain much information useful for identification. In a subsequent study, Mertens and Ravn (2014) identify SVAR-IV model where a single proxy is correlated with two structural shocks. Moreover, they show that it is not required for the proxy to be uncorrelated with past structural shocks. Caldara and Kamps (2017) adopt the proxy-SVAR approach proposed by Mertens and Ravn (2013) to compute the sign and magnitude of fiscal multiplier.

Lunsford (2015), like Mertens and Ravn (2013), uses multiple proxies to identify multiple shocks. Lunsford (2015) contributes towards the literature by providing a test for a weak proxy variable based on F-statistic. He computes the F-statistic through the linear projection of the proxy on the VAR errors. That is, the proxy is the dependent variable when computing the F-statistic. Literature prior to Lunsford (2015) test the proxy strength using the F-statistic where one of the VAR errors is projected onto the proxy variable (Montiel Olea et al., 2021; Gertler and Karadi, 2015). In the application, Lunsford (2015) measures the impact of consumption TFP shock and investment TFP shock using the consumption TFP and investment TFP proxies for US quarterly data from 1948:Q1 to 2015:Q2. These proxies are based on Fernald (2014)'s consumption and investment TFP series.

Mertens and Montiel Olea (2018) follows Mertens and Ravn (2013), Mertens and Ravn (2014) and Lunsford (2015) in terms of identification scheme. Mertens and Montiel Olea (2018) achieve point-identification with multiple proxies for multiple structural shocks using additional zero restrictions. They contribute to the macro literature by developing the narrative identification approach for marginal rather than average tax shocks (Mertens and Ravn, 2013) and by analyzing responses along the income distribution to changes in marginal tax rates rather than in taxable income. In a recent contribution to SVAR-IV literature, Stock and Watson (2018) show that SVAR-IV method is more efficient asymptotically than Local-Projection(LP)-IV under strong instrument asymptotics, and this method does not require lead-lag exogeneity.

Unlike the literature discussed so far that applies frequentist estimation techniques to estimate SVAR-IV models, Giacomini et al. (2022) and Arias et al. (2021) develop algorithms to estimate the proxy-SVAR models in Bayesian settings. Arias et al. (2021) contribute towards literature by achieving identification though a combination of IV, sign-restrictions on structural parameters and additional exclusion restrictions in order to identify multiple structural shocks using multiple proxies.[1] The additional sign and exclusion restrictions are imposed on the covariances (in case of more than one proxies) between the proxies and the structural shocks of interest. In addition to these restrictions, Giacomini et al. (2022) develop their algorithms further and allow for additional short-run zero restrictions (Sims, 1980; Christiano et al., 1999), long-run zero restrictions (Blanchard and Quah, 1989), sign restrictions on impulse responses (Uhlig, 2005), and zero or sign restrictions on the contemporaneous structural parameters (Arias et al., 2019).

3.2.2 Literature on US monetary policy instruments

Beaudry and Saito (1998) is the earliest paper that questions as to why the choice of an appropriate monetary policy indicator is crucial even if we have a valid instrument.

[1] I explain the methodology developed by Arias et al. (2021) in more detail in section 3.3.

They further highlight that relevance and exogeneity conditions are not the only sufficient requirement to identify the shock of interest. There is an additional requirement that the monetary policy variable used as an indicator must capture all the direct effect of the monetary innovation. Moreover, their discussion shows that choice of instrument varies according to the particular system of variables. They are the first ones to use constructed shocks as an instrument in a SVAR. They use dummy variables constructed by Romer and Romer (1989) and its lags as their set of instruments for non-borrowed reserves, where non-borrowed reserves are used as a measure of monetary policy. They find very little evidence against the IV procedure. However, they find that Cholesky decomposition is likely to provide biased estimates of the effects of monetary shocks.

In an early work relating to the information channel of monetary policy, Romer and Romer (2000) find that Federal Reserve has information about future inflation that private sector forecasters do not have. They highlight the asymmetry of information between the Fed and private sector inflation forecasts. Monetary policy provides a signal (information channel) on the basis of which the private sector revise their forecast. The signal reflects the internal forecast of the Fed about the state of the economy. Romer and Romer (2000) argue that a rise in long-term interest rates in response to monetary tightening may be attributed to the Fed holding more information on future inflation than the private sector forecasters. A plausible explanation is that markets would perceive a monetary tightening as a signal of future inflation. Thus, the private sector would update its forecast of expected inflation upwards which results in a rise in long-term interest rates. Specifically, Romer and Romer (2000) carry out regression analysis of inflation on Fed and private sector forecasts. They find coefficient on private sector forecast to be small and insignificant while coefficient on Fed forecast is large and highly significant. This concludes that Fed's internal forecast about future path of inflation contains additional information that is not accessible to the private sector at that time. Contrary to Romer and Romer (2000), Campbell et al. (2012) find no evidence that Fed announcements contain significant information about inflation. However, Campbell et al. (2012) find evidence of information effect in the private sector's forecast of unemployment because the private sector revise its unemployment downward in response to a monetary tightening.

Faust et al. (2004) identify SVAR model using high-frequency federal funds futures contracts. They construct high-frequency data using the daily change in fed funds futures rates on the Federal Open Market Committee (FOMC) meeting days and other announced changes in the Fed's target federal funds rate. Specifically, their idea is based on regressing the FOMC-day change in the contracts for horizons 1–5 on the target surprise. The coefficient estimates from these regressions are used as impulse responses of fed funds rates to policy shocks. They replicate the 6-variable SVAR model of Christiano et al. (1999) and find that high-frequency identification helps remove price puzzle compared with standard recursive identification. In another study, Gürkaynak et al. (2005) carry out a high-frequency event study analysis to measure the impact of monetary policy actions and statements on asset prices: bond yields and stock price index. They construct a new data set that captures changes in asset prices in a thirty-minute window around the FOMC announcement. Further, they use two factors of monetary policy announcements: "current federal funds rate target" associated with changes in the current federal funds rate target (policy actions) and "future path to policy" associated with changes in futures rates to horizons of one year that are independent of changes in the current funds rate target (policy statements). They find that FOMC statements have a much greater impact on long-term Treasury yields. Specifically, around 75% to 90% of the variations in 5-year and 10-year Treasury yield is due to the factor associated with FOMC statements rather than changes in federal funds rate target.

In a recent stream of literature on monetary policy instrument, Gertler and Karadi (2015) quantify the impact of US monetary policy shocks on both economic and financial variables using a proxy-SVAR model in frequentist settings. They achieve identification through a high-frequency monetary policy instrument measured within a 30-minutes (10 minutes before and 20 minutes after) tight window around the FOMC announcement. Further, following Gürkaynak et al. (2005), Gertler and Karadi (2015) account for forward guidance as their policy instrument is constructed as a change in the three month ahead future while they use one-year government bond rate as policy indicator. Their motivation behind employing the high-frequency identification (HFI) approach is to address the issue of simultaneity: the policy shifts influence financial variables and get influenced by them as well, unlike in a standard recursive Cholesky identification scheme. They exploit the HFI approach to test the impact of monetary policy shocks on credit costs through the responses of term premia and credit spread. Gertler and Karadi (2015) find conventional responses of output and price index when they identify the model through external instrument and include term premia and credit spread in the model. On the other hand, they find puzzling responses of output and price index when they employ the recursive Cholesky identification scheme.[2]

Following the work of Gertler and Karadi (2015), Caldara and Herbst (2019) quantify the impact of US monetary shocks in a Bayesian Proxy-SVAR (BP-SVAR) approach adopted from Stock and Watson (2012) and Mertens and Ravn (2013). Caldara and Herbst (2019) take into account the changes in corporate credit spread while measuring the systematic component of monetary policy. In order to construct the proxy for monetary policy shocks, Caldara and Herbst (2019) apply the event-study methodology developed by Kuttner (2001), which uses high frequency financial data to construct monetary policy surprises associated with FOMC announcements. Specifically, Caldara and Herbst (2019) use the surprises in spot months contracts based on the current month funds rate in a tight 30-minutes window around the FOMC announcement while they use the fed funds rate as the policy indicator. Their policy indicator and instrument are different from that of Gertler and Karadi (2015) because their sample does not include the zero-lower bound (ZLB) period unlike that of Gertler and Karadi (2015). Caldara and Herbst (2019) find large and conventional economic effects of monetary policy shocks when they include corporate credit spread as compared to when they do not. Finally, their findings corroborate that of Gertler and Karadi (2015): responses obtained from HFI identification schemes do not show evidence of output puzzle and price puzzle unlike the responses obtained from standard Cholesky

[2] Plots of recently constructed proxies are in Appendix C.

identified SVAR models.

In a recent theoretical study, Melosi (2017) develops a dynamic general equilibrium model to quantitatively study the signalling channel of monetary transmission for the US economy. In the model, the central bank is assumed to have imperfect information and can make errors in forecasting the targeted macroeconomic aggregates. Moreover, the central bank's policy rate provides public information (signals) to price-setting firms regarding the central bank's view on current inflation and the output gap, which directly influences firm's beliefs about macroeconomic fundamentals. The signalling channel of monetary transmission largely depends on price-setting firm's interpretation of changes in the policy rate. A monetary tightening might be interpreted as the central bank responding to an exogenous deviation from its monetary policy rule. In this case, a monetary tightening (easing) puts downward (upward) pressure on inflation and firm's inflation expectations. On the other hand, a higher interest rate could also be interpreted as the response of the central bank to inflationary non-policy shocks, which, in the model, are an adverse aggregate technology shock or a positive demand shock. In this case, raising (cutting) the policy rate induces firms to expect higher (lower) inflation, and hence inflation rises (falls). The model finds signalling effect as a reason of persistent inflation in the US in 1970s and a sluggish downward trend of inflation expectation in 1980s. Melosi (2017) also uses a VAR model in order to further evaluate the empirical performance of the theoretical model. He shows that the theoretical model can closely replicate the response of expectations to monetary shocks implied by a VAR model.

Lunsford (2020) tests forward guidance as a tool to influence private sector's expectations through communication about the likely future course of interest rates and economic conditions. He uses high-frequency federal funds rate surprises in a 30-minute window to measure two policy surprises: a change in the current federal funds rate and a change in the expected path of the federal funds rate, as proxy for the forward guidance surprise. He also uses changes in financial market variables during FOMC meeting days and changes in private forecasts during FOMC meeting months on the two policy surprises. Specifically, Lunsford (2020) studies the private sector responses against the nature of FOMC's forward guidance language in the meetings held from February 2000 to May 2006. He measures the impact of conventional effect versus the information effect through this sample period using an event-study regression. He makes use of the fact that FOMC's statements differed in its language. FOMC only used "economic-outlook" forward guidance from February 2000 to June 2003 and later it added "policy-inclination" forward guidance from August 2003 to May 2006. He finds evidence of information effect from February 2000 to August 2003 when Fed uses "economic-outlook" forward guidance announcements. However, when Fed uses "policy-inclination" forward guidance from August 2003 to May 2006, the information effects weaken and standard effects become more relevant.

More recent literature on external instruments further highlights the importance of an "information channel" in the transmission of monetary policy shocks. The "information channel" rises from asymmetric information between the private sector and Central bank. For instance, an informationally constrained private sector would either conceive a policy rate hike as a contractionary monetary shock or an endogenous response of the Central bank to a stronger than expected future economic outlook (information effect). Nakamura and Steinsson (2018) take into account the "Fed information effect" inherent in the FOMC announcements while constructing monetary shocks. They use high-frequency identification of monetary non-neutrality measured through unexpected changes in interest rates and movements in bond prices around a short 30-minute window surrounding FOMC meetings. Specifically, they estimate the parameter of interest in order to measure the effect of FOMC announcement on change in bond prices relative to its effect on the change in policy indicator in a 30-minute window. Using their high-frequency based identification they find information channel to be more dominant than the conventional channel of monetary policy: two-third of the response of real interest rates to FOMC announcements are due to the information channel while one-third is from the more conventional channel of monetary policy. They find large amount of monetary non-neutrality as an increase in interest rate results in around one-for-one increase in both nominal and real interest rates. Inflation, measured through the slope of the Philips curve, shows relatively smaller decrease. The "Fed information effect" plays a larger role in impacting the expected output growth. Nakamura and Steinsson (2018) find that a tightening of monetary policy leads to an increase in expected output growth rather than a decrease. The unconventional response of output growth shows that information channel in the FOMC announcement influences the private sector's belief about the state of the macroeconomy. For instance, private sector's belief on positive economic fundamentals leads to an increase in output growth in response to monetary tightening.

Further to the work of Nakamura and Steinsson (2018), Jarociński and Karadi (2020) and Miranda-Agrippino and Ricco (2021) construct two high-frequency instruments for US monetary policy that take into account these informational rigidities, unlike Gertler and Karadi (2015) and Caldara and Herbst (2019) who assume that private sector enjoys full information. These two high-frequency instruments are considered complementary (Miranda-Agrippino and Nenova, 2022) and crucial in understanding some puzzles in the US monetary literature. Jarociński and Karadi (2020) disentangle the Central bank announcements into i) information about monetary policy and ii) Central bank's assessment of the economic outlook. They construct two proxies: a proxy for monetary policy shocks and a proxy for Central bank information shocks. The intuition behind the construction of these proxies is that each month is characterised either as a pure monetary shock or a pure Central bank information shock. To construct these proxies, they make use of high-frequency co-movements in interest rates and stock prices in a 30 minutes window around the FOMC announcements. Particularly, if interest rates and stock prices co-move negatively (positively), it is reflected as a pure monetary policy (pure Central bank information) shock. They use three month federal funds rate and Standard and Poor's (S&P) 500 index to measure short-term interest rate expectations and stock prices, respectively. For constructing the proxies, they use the months of fed funds futures surprises where the stock price surprise has the opposite sign compared with the fed funds futures surprise as the proxy for pure monetary policy shock (the proxy is zero otherwise). The fed funds futures surprises in the remaining months, where the stock price surprise has the same sign as that of fed funds futures surprise, is characterized as a proxy for pure Central bank information shock (again, the proxy is zero otherwise). This procedure is dubbed as "poor man's sign-restrictions". Jarociński and Karadi (2020) find that a key difference from the standard HFI of monetary policy shocks that fail to account for the Central bank information effect is that their purged monetary policy shock induces a more pronounced price-level decline. My results in section 3.2 re-emphasize these findings when I compare the information-robust proxy of Jarociński and Karadi (2020) and the standard proxy of Gertler and Karadi (2015).

Miranda-Agrippino and Ricco (2021) construct their monetary policy instrument using high-frequency market-based surprises in the fourth federal funds futures (FF4) around the FOMC announcements. They proceed in three steps: in the first step, they regress FF4 on Greenbook forecasts and forecast revisions for real output growth, inflation and the unemployment rate (Romer and Romer, 2004). The residual from the above regression is obtained after controlling for the transfer of information that implicitly happens at the time of the FOMC announcements. This residual acts as an instrument for monetary policy shocks (MPI_m). In the second step, they use the aggregate method (Stock and Watson, 2012; Caldara and Herbst, 2019) to sum up the daily (MPI_m) into a monthly instrument. The final step takes into account the slow absorption of information by the private sector by removing the autoregressive component from the monthly instrument obtained in the second step. The resulting residual from step 3 is the informationally robust monetary policy instrument. In the empirical exercise, using a six variable BP-SVAR model, Miranda-Agrippino and Ricco (2021) compare the results from their informationally robust proxy with the proxy of Gertler and Karadi (2015) and narrative measure of Romer and Romer (2004). The impulse responses obtained from informationally robust proxy of Miranda-Agrippino and Ricco (2021) show no evidence of price and output puzzles in response to a contractionary monetary shock to a one-year rate. Moreover, the unemployment rate rises significantly in line with theory. On the other hand, industrial production contracts and unemployment rate falls on impact when using Gertler and Karadi (2015) proxy and Romer and Romer (2004) narrative measure while the price level increases when using the Romer and Romer (2004) narrative measure.

Two recent empirical studies by Degasperi et al. (2023) and Miranda-Agrippino and Nenova (2022) use the information-robust proxies of Miranda-Agrippino and Ricco (2021) and Jarociński and Karadi (2020), respectively. Degasperi et al. (2023) study the transmission of US monetary policy shocks using data from 30 economies. They identify their BP-SVAR model using the instrument variable constructed by Miranda-Agrippino and Ricco (2021). They find strong financial spillovers of the US monetary shocks and establish the role of Fed as the global central bank. In an other study, Miranda-Agrippino and Nenova (2022) quantify the macroeconomic and financial spillovers of unconventional monetary policy of the Fed and the European central bank (ECB) using daily projections approach of Jordà (2005). They make use of "poor man's" identification of Jarociński and Karadi (2020) to isolate the Fed and ECB announcements in which the high-frequency stock market response is in line with economic theory. They find that integration of international financial markets led to a dominant role of the US and US dollar as an important driver of the global financial cycle (GFC).

Following Jarociński and Karadi (2020), Jarociński (2022) constructs proxies for ECB monetary shocks and ECB information shocks based on high-frequency comovement of interest rates and stock prices around the policy announcements. He uses monthly VAR models to measure the impact of ECB monetary shocks and information shocks on the US macro and financial variables. He finds a stronger impact of ECB information shocks on the US economy as compared to ECB monetary shocks. Specifically, he finds that ECB surprises that generate information effect at home spillovers to the US economy as well, which leads to an easing of US economic conditions. The impact is particularly significant on US financial variables: stock prices, corporate bond spreads and the dollar exchange rate followed by an expansion in US real activity and higher prices. Moreover, he applies same methodology to study the effects of Fed shocks on the euro area. He finds stronger spillovers of the US monetary shocks than the information shocks.

In another related work to Jarociński and Karadi (2020), Cieslak and Schrimpf (2019) decompose Central bank communication into monetary news and non-monetary news defined as news about economic growth and news affecting financial risk premia. They exploit high-frequency co-movement of stock prices and interest rates combined with monotonicity restrictions across maturities in the yield curve. Specifically, monetary policy shock affects the real rate and induces a negative co-movement of stocks and yields while both growth and risk premia shocks induce a positive co-movement of stocks and yields. Moreover, the growth shocks have greater affect on the short-to-intermediate segment of the yield curve while risk premium shocks affect more on the long segment. They show that non-monetary news drives a significant part of financial markets reaction during the financial crisis and in the early recovery, while monetary news gains importance since 2013.

A paper related to the information-robust instruments by Miranda-Agrippino and Ricco (2023) test for the robustness of various monetary policy instruments based on "lead-lagged exogeneity" condition. They argue that other than the standard relevance and contemporaneous exogeneity restrictions, necessary for the identification of SVAR-IV, the instrument must satisfy a 'limited lead-lag exogeneity condition' which ensures that the VAR innovations and the instrument correlate only via the contemporaneous structural shock of interest. They find that Romer and Romer (2004)'s narrative IV and Swanson (2021)'s IV for forward guidance and quantitaive easing shocks fail both the contemporaneous and lag exogeneity conditions. Moreover, they test for three variants of high-frequency IVs proposed in Gertler and Karadi (2015), Caldara and Herbst (2019) and Miranda-Agrippino and Ricco (2021) and find that only the IV that controls for information effect (Miranda-Agrippino and Ricco, 2021) passes the test for lagged conditional exogeneity.

A recent work by Sastry (2022) develops a theoretical and empirical model in order to analyse the disagreement between policymakers and private sector regarding beliefs about economic fundamentals and the response of monetary policy. He exploits three possible mechanisms that could cause this disagreement: i) asymmetric information between Fed and private sector about the state of the economy, ii) asymmetric beliefs about the policymaker's reaction to public signals, and iii) asymmetric confidence in those signals. Sastry (2022) distinguishes between two reasons for the surprisingly strong reaction of the Fed to economic news: First, changes in the state of the economy can cause the Fed to change the interest rate by more than the private sector expected, and second, economic news can cause the Fed to revise its estimate of the state of the economy by more than the private sector expected. His paper provides evidence in support of both of these phenomena, with underreaction of the private sector to economic news being particularly important. He finds essentially no role for a Fed information effect in the data.

Bauer and Swanson (2023a) construct a new more refined measure of monetary policy instrument that is both more *relevant* and more *exogenous* than those used in previous studies. Bauer and Swanson (2023a) address the questions of instrument *exogeneity* (raised in Cieslak, 2018) and *relevance* (raised in Ramey, 2016), in two steps: first, they improve the *relevance* of monetary policy surprises through the inclusion of press conferences, speeches, and testimony by the Federal Reserve Chair and, thus, expand the set of FOMC announcements. Secondly, in order to address the *exogeneity* issue, Bauer and Swanson (2023a) remove the predictable component of the monetary policy surprises from this newly expanded set of monetary announcements, by taking into account the "Fed response to new" channel (Bauer and Swanson, 2023b). Moreover, they orthogonalize the monetary surprises with respect to macroeconomic and financial data. They find biased estimates while measuring the impact of monetary surprises on macroeconomic variables while employing an SVAR model with the high-frequency monetary instruments used in the previous literature. The biased estimates are attributed to the correlation between the instrument and macro variables. This issue is resolved when they use their more exogenous orthogonal high-frequency instrument that produces impulse responses with no evidence of puzzling responses. Further, they find reliable estimates while measuring the impact of monetary surprises on financial variables using either of the two approaches: traditional unadjusted monetary policy instrument and the newly constructed one by Bauer and Swanson (2023a).

Finally, a most recent paper by Bauer and Swanson (2023b), unlike Lunsford (2020), propose an alternative explanation of the "Fed information effect" that they find for the period from February 2000 to August 2003. Bauer and Swanson (2023b) call this effect the "Fed information channel". Moreover, they argue that the unconventional response of output growth found in Nakamura and Steinsson (2018) is due to the Fed information channel. However, conventional responses are obtained when they take into account this new channel of monetary transmission. Therefore, they disentangle the "Fed response to news" channel and the "Fed information effect" channel. The "Fed response to news" channel transmits through incoming, publicly available news on the health of the economy. This is measured using the difference between Fed's actual policy response function based on the state of the economy and the private sector's ex-ante estimate of that function. The essence of the news channel proposed by Bauer and Swanson (2023b) is attributed to a more stronger than private sector's expected response of the Fed to business cycle fluctuations. Their empirical findings strongly favour the Fed response to news channels compared with the Fed information effect. They argue that although Campbell et al. (2012) and Nakamura and Steinsson (2018) find that coefficients on the policy surprises have "wrong" sign these coefficient's statistical significance is fragile. Bauer and Swanson (2023b) argue that this "wrong" sign is due to the omitted variable bias. Further, they find that the sign of these coefficients is not robust to sample period and to the choice of the variables being forecast. On the other hand, they show that economic news released in the days leading up to an FOMC announcement does have a major impact on the responses of variable forecast. For instance, including employment report as an indicator of economic news in regression analysis reverses the sign on the coefficients of the monetary surprises, thus, making the response of forecast variables conventional.

In the next section, I provide an empirical application using BP-SVAR model of Arias et al. (2021) that compares the standard (Gertler and Karadi, 2015) versus information-robust proxy (Jarociński and Karadi, 2020).

3.3 Application: Comparison between a standard and information-robust proxy

In this section, I employ the most recent SVAR-IV methodology developed by Arias et al. (2021) and study the dynamic effects of US monetary shocks on US macroeconomic and financial variables using the proxies constructed by Jarociński and Karadi (2020) and Gertler and Karadi (2015). This section contains: data and model description, replication results of the "poor man's sign-restrictions" of Jarociński and Karadi (2020) using the pure monetary policy proxy and the methodology developed by Arias et al. (2021). Finally, I compare the results obtained from the standard proxy (Gertler and Karadi, 2015) and information-robust proxy (Jarociński and Karadi, 2020).

3.3.1 Data and Model

I adopt the SVAR specification and the pure monetary policy proxy of Jarociński and Karadi (2020). Accordingly, the endogenous variables are: one-year government bond yield, S&P 500 index, real GDP, GDP deflator and excess bond premium (EBP).[3]

I follow the most recent BP-SVAR methodology developed by Arias et al. (2021). The structural model takes the form:[4]

$$\mathbf{\tilde{y}}_\mathbf{t}'\mathbf{\tilde{A}_0} = \sum_{L=1}^{p} \mathbf{\tilde{y}}_\mathbf{t-L}'\mathbf{\tilde{A}_L} + \mathbf{\tilde{c}}' + \tilde{\epsilon}_t' \tag{3.1}$$

[3] The data set is publicly available and follows the baseline specification of Jarociński and Karadi (2020).

[4] Model is estimated through 5,00,000 draws from posterior distribution.

where $\mathbf{\tilde{y}_t'} = [\mathbf{y_t'} \quad \mathbf{m_t'}]$, $\tilde{n} = n + k$, $\mathbf{\tilde{A}_i}$ is an $\tilde{n} \times \tilde{n}$ matrix for $0 \leq i \leq p$ with $\mathbf{\tilde{A}_0}$ is invertible, $\mathbf{\tilde{c}}$ is a $1 \times \tilde{n}$ row vector, and $\tilde{\epsilon}_t$ is conditionally standard normal. $\mathbf{y_t}$ is an $n \times 1$ vector of endogenous variables, $\mathbf{m_t}$ is a $k \times 1$ vector of instruments. Equation (3.1) can be written as:

$$\mathbf{\tilde{y}_t'} \mathbf{\tilde{A}_0} = \mathbf{\tilde{x}_t'} \mathbf{\tilde{A}}_+ + \tilde{\epsilon}_t' \ , \tag{3.2}$$

where $\mathbf{\tilde{A}}_+ = [\mathbf{\tilde{A}_1'} \ldots \mathbf{\tilde{A}_p'} \quad \mathbf{\tilde{c}'}]$ and $\mathbf{\tilde{x}_t'} = [\mathbf{\tilde{y}_{t-1}'} \ldots \mathbf{\tilde{y}_{t-p}'} \quad 1]$ for $1 \leq \mathrm{t} \leq \mathrm{T}$.

Let $\tilde{\epsilon}_t' = [\epsilon_t' \quad \nu_t']$, where ϵ_t is $n \times 1$ and ν_t is $k \times 1$ and ν_t is uncorrelated with ϵ_t. In my case, I have one instrument in my application, so $k = 1$. A proxy SVAR imposes that $\mathbf{y_t}$ evolves according to $\mathbf{y_t'} \mathbf{A_0} = \mathbf{x_t'} \mathbf{A}_+ + \epsilon_t'$ for $1 \leq \mathrm{t} \leq \mathrm{T}$, where $\mathbf{A}_+ = [\mathbf{A_1'} \ldots \mathbf{A_p'} \quad \mathbf{c'}]$ and $\mathbf{x_t'} = [\mathbf{y_{t-1}'} \ldots \mathbf{y_{t-p}'} \quad 1]$, with $\mathbf{A_i}$ an $n \times n$ matrix for $0 \leq \mathrm{i} \leq p$, $\mathbf{A_0}$ invertible, and $\mathbf{c}$ a $1 \times n$ row vector. ϵ_t are the structural shocks and ν_t are other shocks that affect the proxy. Hence,

$$\mathbf{\tilde{A}_i} = \begin{bmatrix} \mathbf{A_i} & \mathbf{\Gamma}_{i,1} \\ & \\ \mathbf{0}_{k \times n} & \mathbf{\Gamma}_{i,2} \end{bmatrix}$$

where $\mathbf{\Gamma}_{i,1}$ is $n \times k$ and $\mathbf{\Gamma}_{i,2}$ is $k \times k$ for $0 \leq \mathrm{i} \leq p$ and $\mathbf{0}_{k \times n}$ is a $k \times n$ matrix of zeros. These zero restrictions on $\mathbf{\tilde{A}}_0$ and $\mathbf{\tilde{A}}_+$ are the block restrictions.

Finally, the instrument satisfies the i) relevance condition: the external instrument is contemporaneously correlated with the structural shock of interest ($\mathbb{E}[\mathbf{m}_t \quad \epsilon_{1t}] = \psi$) and ii) exogeneity restriction: the external instrument is contemporaneously uncorrelated with the remaining shocks in the system ($\mathbb{E}[\mathbf{m}_t \quad \epsilon_{2t}] = 0$). Here, ϵ_{1t} is the structural shock of interest (in my case it is the monetary policy shocks), ϵ_{2t} are other shocks and ψ is the $k \times k$ covariance matrix of the k proxies (in my case $k = 1$).

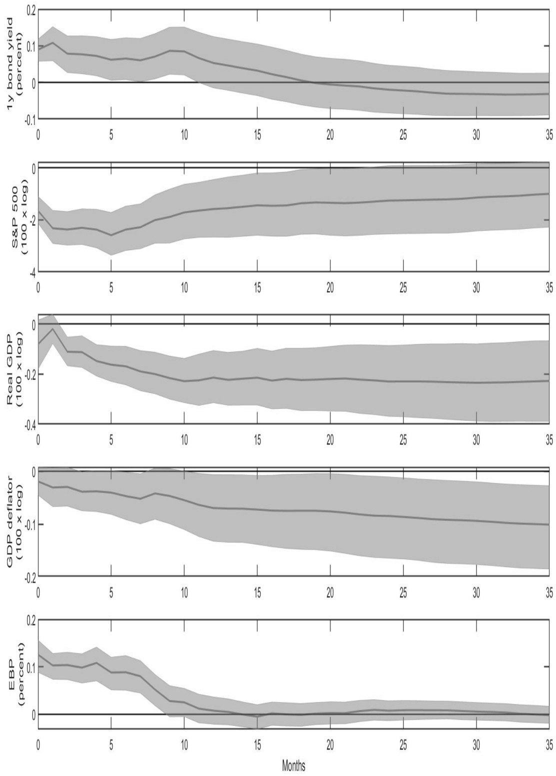


Figure 3.1: IRFs to a one standard deviation contractionary monetary policy shock identified using Jarociński and Karadi (2020) proxy and sample size from 1984:M1 to 2016:M12. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 68% equal-tailed point-wise posterior probability bands.

3.3.2 Results

In the first step, I replicate the left panel in Figure 3 of Jarociński and Karadi (2020). Figure 3.1 below shows the responses of a 1-SD contractionary shock to US 1-year government bond yield over the monthly sample period 1984:M1 to 2016:M12. The solid lines are the point-wise posterior median responses and the shaded bands represent the 68% equal-tailed point-wise posterior probability bands. I achieve identification using the proxy-SVAR methodology developed by Arias et al. (2021). Further, I employ the “poor man’s proxy” for pure US monetary policy constructed by Jarociński and Karadi (2020) as an instrument for US 1-year government bond yield. The IRF’s are qualitatively similar to those of Jarociński and Karadi (2020), however, I find some quantitative differences in magnitude of the responses. These differences may be attributed to the identification schemes and sampling techniques: left two panels of Figure 2 in Jarociński and Karadi (2020) use Gibbs sampler to draw from the posterior, while Arias et al. (2021) use Importance sampler. Moreover, Figure 3 in Jarociński and Karadi (2020) use Cholesky identification with proxy ordered first. Specifically, in my results of Figure 3.1 below, we see that 1y gov bond yield rises by 0.1 percent on impact as compared to 0.05 percent in Jarociński and Karadi (2020). We see gradual decrease in 1y gov bond yield which becomes insignificant in around 2 years. The responses of other variables are in line with theoretical findings: persistent and significant contraction in real GDP, fall in GDP deflator and S&P stock market index while the EBP increases reflecting tightening financial market.

In the second step, I compare the responses obtained from the BP-SVAR models using the proxies of Jarociński and Karadi (2020) and Gertler and Karadi (2015). I adopt the SVAR specification from the first step. Figure 3.2 shows the responses of a 1-SD contractionary shock to US 1-year government bond yield over the sample perod 1990:M1 to 2012:M12.[5] The blue solid posterior median responses along with 68% posterior probability bands are obtained using Jarociński and Karadi (2020) proxy while the red dashed lines are from Gertler and Karadi (2015) proxy. The impact responses of 1y gov bond yield and S&P index are more pronounced when I use the Gertler and Karadi (2015) proxy as compared to when I use Jarociński and Karadi (2020) proxy. Moreover, I find a persistent increase in 1y gov bond yield in both the models as compared to the corresponding response from the first step. The key findings are related to real GDP and GDP deflator where we can see the effect of information-robust proxy of Jarociński and Karadi (2020). The real GDP contracts and GDP deflator falls significantly on impact when I use the information-robust instrument as compared to when I use the instrument of Gertler and Karadi (2015) that does not take into account the information effect. Most importantly, the responses of real GDP and

[5] My sample period corresponds to the proxy series constructed by Gertler and Karadi (2015). However, unlike Gertler and Karadi (2015), I do not extend the proxy series by using the residuals of reduced form VAR because I employ the SVAR specification of Jarociński and Karadi (2020).

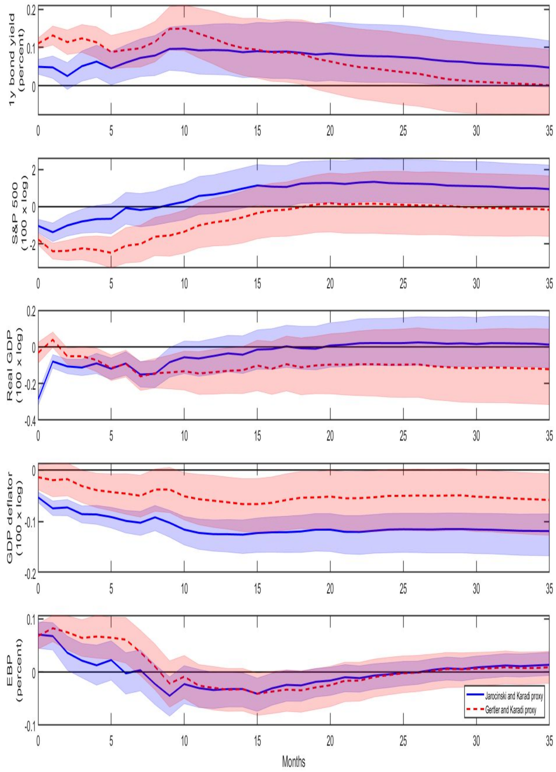


Figure 3.2: IRFs to a one standard deviation contractionary monetary policy shock identified using Jarociński and Karadi (2020) proxy (blue solid lines) and Gertler and Karadi (2015) (red dashed lines) and sample size from 1990:M1 to 2012:M6. *Note:* The solid lines depict the point-wise posterior median responses and the shaded bands represent the 68% equal-tailed point-wise posterior probability bands.

GDP deflator are more pronounced with Jarociński and Karadi (2020) instrument as compared to Gertler and Karadi (2015) one. This shows that controlling for information effect in the monetary policy instrument has a substantial impact on the responses of macroeconomic variables.

3.4 Conclusion and future research

Over the last decade, proxy-SVAR models have become an effective tool to identify the fiscal and monetary shocks. The idea in proxy-SVAR models is to make use of an instrument variable that is correlated with the shock of interest (relevance condition) and uncorrelated with the rest of the shocks (exogeneity condition). This identification scheme results in obtaining point estimated IRFs. Further, one or more external instruments can be used to identify a proxy-SVAR model which makes the estimation more robust. One challenge in the proxy-SVAR identification scheme is to construct a strong instrument. Recent literature constructs high-frequency external instruments for the US monetary shocks which makes the use of proxy-SVAR models more common.

In this chapter, I review the literature on proxy-SVAR methodology and various types of US monetary policy instruments. Further, I provide an application using BP-SVAR model recently developed by Arias et al. (2021). I compare the results obtained from an information robust proxy of Jarociński and Karadi (2020) with that of a standard proxy that does not adjust for the information effects (Gertler and Karadi, 2015). I find that, in response to a contractionary US monetary shock, the fall in real GDP and prices is more pronounced when I use the information-robust proxy as compared to when I use the standard proxy. Thus, my findings emphasis that results obtained from an information-robust proxy for the US monetary shocks are different from the ones obtained using the standard proxy.

In future, I intend to explore the spillover effects of US monetary shocks on small-open-economies (SOEs). I would achieve identification through a combination of proxy, exclusion and sign restrictions. Methodologically, I intend to build on the work of Arias et al. (2021) to the SOE settings. Specifically, I would introduce block exogeneity in the BP-SVAR model of Arias et al. (2021) through a combination of exclusion restrictions and informative priors on structural parameters. Further, I would impose sign-restrictions on the structural parameters that are inline with macroeconomic and financial literature. In this way, I will contribute towards the empirical macro SOE monetary literature using a novel identification scheme that is more robust than used in the literature.

Appendix C

Appendix to Chapter 3

C.1 Plots of proxies

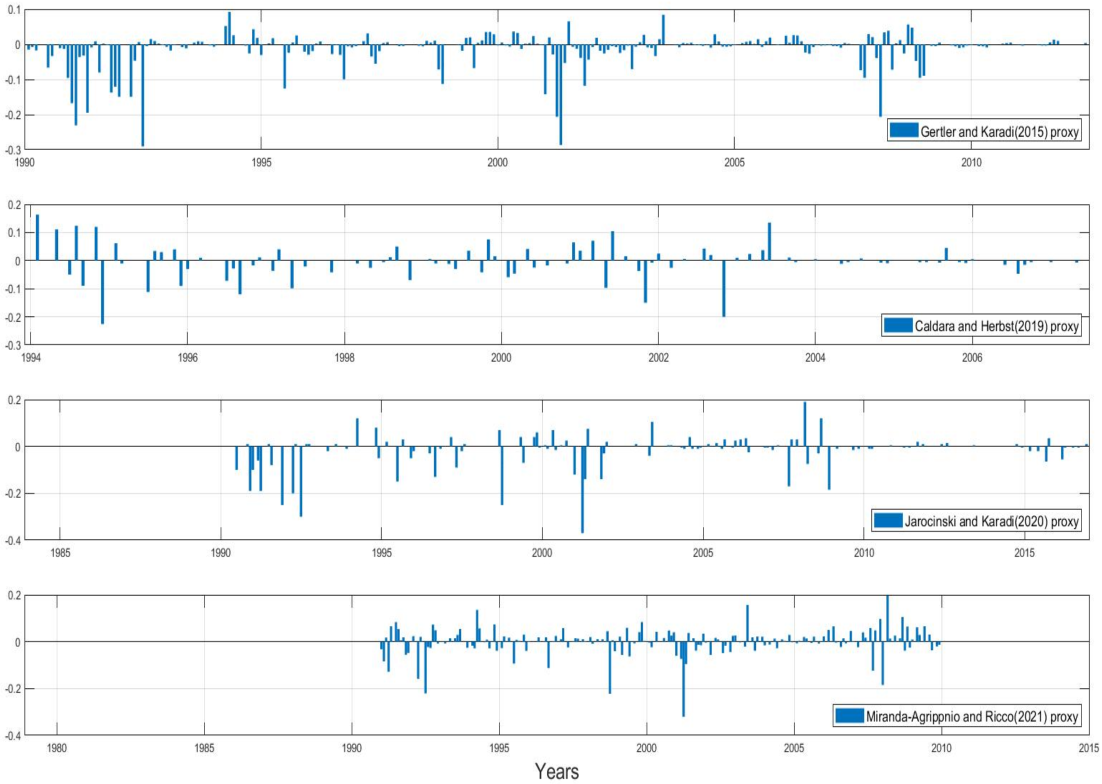


Figure C.1: Plots of proxies

References

Arias, J. E., Caldara, D., and Rubio-Ramirez, J. F. (2019). The systematic component of monetary policy in SVARs: An agnostic identification procedure. *Journal of Monetary Economics*, 101:1–13.

Arias, J. E., Rubio-Ramírez, J. F., and Waggoner, D. F. (2021). Inference in Bayesian Proxy-SVARs. *Journal of Econometrics*, 225(1):88–106.

Bauer, M. D. and Swanson, E. T. (2023a). A reassessment of monetary policy surprises and high-frequency identification. *NBER Macroeconomics Annual*, 37(1):87–155.

Bauer, M. D. and Swanson, E. T. (2023b). An alternative explanation for the “fed information effect”. *American Economic Review*, 113(3):664–700.

Beaudry, P. and Saito, M. (1998). Estimating the effects of monetary shocks: An evaluation of different approaches. *Journal of Monetary Economics*, 42(2):241–260.

Blanchard, O. J. and Quah, D. (1989). The dynamic effects of aggregate demand and aggregate supply. *The American Economic Review*, 79(4):655–673.

Caldara, D. and Herbst, E. (2019). Monetary policy, real activity, and credit spreads: Evidence from Bayesian proxy SVARs. *American Economic Journal: Macroeconomics*, 11(1):157–192.

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.

Campbell, J. R., Evans, C. L., Fisher, J. D., Justiniano, A., Calomiris, C. W., and Woodford, M. (2012). Macroeconomic effects of federal reserve forward guidance [with comments and discussion]. *Brookings papers on economic activity*, pages 1–80.

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.

Cieslak, A. (2018). Short-rate expectations and unexpected returns in treasury bonds. *The Review of Financial Studies*, 31(9):3265–3306.

Cieslak, A. and Schrimpf, A. (2019). Non-monetary news in central bank communication. *Journal of International Economics*, 118:293–315.

Degasperi, R., Hong, S., and Ricco, G. (2023). The global transmission of us monetary policy. *Center for Research in Economics and Statistics, Working paper*.

Faust, J., Swanson, E. T., and Wright, J. H. (2004). Identifying VARs based on high frequency futures data. *Journal of Monetary Economics*, 51(6):1107–1131.

Fernald, J. (2014). A quarterly, utilization-adjusted series on total factor productivity. Federal Reserve Bank of San Francisco.

Gertler, M. and Karadi, P. (2015). Monetary policy surprises, credit costs, and economic activity. *American Economic Journal: Macroeconomics*, 7(1):44–76.

Giacomini, R., Kitagawa, T., and Read, M. (2022). Robust Bayesian inference in proxy SVARs. *Journal of Econometrics*, 228(1):107–126.

Gürkaynak, R. S., Sack, B., and Swanson, E. T. (2005). Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements. *The International Journal of Central Banking*, 54(8):2291–2304.

Jarociński, M. (2022). Central bank information effects and transatlantic spillovers. *Journal of International Economics*, 139:103683.

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

Jordà, Ò. (2005). Estimation and inference of impulse responses by local projections. *American economic review*, 95(1):161–182.

Kuttner, K. N. (2001). Monetary policy surprises and interest rates: Evidence from the Fed funds futures market. *Journal of monetary economics*, 47(3):523–544.

Lunsford, K. G. (2015). Identifying structural VARs with a proxy variable and a test for a weak proxy.

Lunsford, K. G. (2020). Policy language and information effects in the early days of Federal Reserve forward guidance. *American Economic Review*, 110(9):2899–2934.

Melosi, L. (2017). Signalling effects of monetary policy. *The Review of Economic Studies*, 84(2):853–884.

Mertens, K. and Montiel Olea, J. L. (2018). Marginal tax rates and income: New time series evidence. *The Quarterly Journal of Economics*, 133(4):1803–1884.

Mertens, K. and Ravn, M. O. (2013). The dynamic effects of personal and corporate income tax changes in the United States. *American economic review*, 103(4):1212–1247.

Mertens, K. and Ravn, M. O. (2014). A reconciliation of SVAR and narrative estimates of tax multipliers. *Journal of Monetary Economics*, 68:S1–S19.

Miranda-Agrippino, S. and Nenova, T. (2022). A tale of two global monetary policies. *Journal of International Economics*, 136:103606.

Miranda-Agrippino, S. and Ricco, G. (2021). The transmission of monetary policy shocks. *American Economic Journal: Macroeconomics*, 13(3):74–107.

Miranda-Agrippino, S. and Ricco, G. (2023). Identification with external instruments in structural VARs. *Journal of Monetary Economics*, 135:1–19.

Montiel Olea, J. L., Stock, J. H., Watson, M. W., et al. (2021). Inference in structural vector autoregressions identified with an external instrument. *Journal of Econometrics*, 225(1):74–87.

Nakamura, E. and Steinsson, J. (2018). High-frequency identification of monetary non-neutrality: the information effect. *The Quarterly Journal of Economics*, 133(3):1283–1330.

Ramey, V. A. (2016). Macroeconomic shocks and their propagation. *Handbook of macroeconomics*, 2:71–162.

Romer, C. D. and Romer, D. H. (1989). Does monetary policy matter? A new test in the spirit of Friedman and Schwartz. *NBER macroeconomics annual*, 4:121–170.

Romer, C. D. and Romer, D. H. (2000). Federal Reserve information and the behavior of interest rates. *American economic review*, 90(3):429–457.

Romer, C. D. and Romer, D. H. (2004). A new measure of monetary shocks: Derivation and implications. *American economic review*, 94(4):1055–1084.

Romer, C. D. and Romer, D. H. (2010). The macroeconomic effects of tax changes: estimates based on a new measure of fiscal shocks. *American Economic Review*, 100(3):763–801.

Sastry, K. (2022). Disagreement about monetary policy. *Available at SSRN 3421723*.

Sims, C. A. (1980). Macroeconomics and reality. *Econometrica: journal of the Econometric Society*, pages 1–48.

Stock, J. H. and Watson, M. W. (2008). Lecture 7: Recent Developments in structural VAR modeling. *Presentd at the NBER Summer Institute Econometrics Lectures: What's New in Economics-Time Series, July*, 15.

Stock, J. H. and Watson, M. W. (2012). Disentangling the Channels of the 2007-2009 Recession. Technical report, National Bureau of Economic Research.

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.

Swanson, E. T. (2021). Measuring the effects of federal reserve forward guidance and asset purchases on financial markets. *Journal of Monetary Economics*, 118:32–53.

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.

Conclusion

This thesis is a compilation of three distinct chapters that explores the impact of monetary policy and commodity prices with a particular focus on SOEs. Although the three chapters address three different research questions there are some commonalities. All the three chapters employ SVAR models, use Bayesian estimation techniques and focus on policy relevant research questions. The first chapter of my thesis employs Bayesian SVAR models for six SOEs to address a long-debated policy relevant question of Dornbusch's overshooting hypothesis. The findings suggest that the results are in line with the Dornbusch's overshooting hypothesis. The second chapter quantifies the impact of commodity supply and demand shocks for Australia in a Bayesian SVAR settings. Given the importance of Australia as a major exporter of iron ore and coal, this question is of specific importance to the policy makers. I find that in response to the commodity demand and supply shocks, the responses of domestic output does not deviate significantly from baseline while trade balance deteriorates significantly. In my third chapter, I review the literature on proxy-SVARs and various monetary policy instruments. I also provide an application that employs Bayesian proxy-SVAR model to compare a proxy that takes into account the information effect in the Fed's monetary announcements to a proxy that does not. I find that US output contracts and prices fall more when I use the information robust proxy as compared to when I use the standard proxy. All the three chapters contribute towards the literature by employing state-of-the-art empirical macro-econometric techniques to address the most relevant policy-related questions.

In the first chapter, we test the overshooting hypotheses proposed by Dornbusch (1976). Overshooting hypothesis is based on the response of exchange rate to a contractionary monetary shock. Specifically, the exchange rate appreciates on impact and then gradually depreciates, in response to the contractionary monetary shocks. Some empirical literature finds exchange rate responses that are not inline with Dornbusch's overshooting hypothesis. There are two types of puzzles reported in the literature. An exchange rate puzzle where the exchange rate depreciates on impact rather than appreciating and a delayed overshooting puzzle where the exchange rate continues to appreciate rather than depreciating.

We test the overshooting hypothesis using SVAR models for six SOEs: Australia, Canada, New Zealand, Norway, Sweden and the UK. We employ the identification scheme recently proposed by Arias et al. (2019). The identification scheme combines exclusion and sign-restrictions with a novel feature where these restrictions are imposed directly on the structural parameters. Our novel identification scheme combines the sign-restrictions on structural parameters (Arias et al., 2019) and block exogeneity restrictions (Cushman and Zha, 1997). Our block exogeneity structure, which is essential in SOE settings, is imposed through a combination of exclusion restrictions and tight priors on the contemporaneous and lagged parameters, respectively. Our identification strategy simultaneously identify the US and SOE monetary shocks. We leave the response of real exchange rate agnostic which helps us preserve the contemporaneous interaction between the exchange rate and policy rates. This identification approach is essential in addressing the puzzling responses that we observe in the empirical literature.

Our results are consistent with the Dornbusch's overshooting hypothesis as we find no evidence of exchange rate puzzle and delayed overshooting puzzle. We find that a domestic contractionary monetary shock results in an on impact appreciation of the exchange rate followed by a gradual depreciation. We find consistent responses of exchange rates in response to a US monetary tightening such that, the US dollar appreciates on impact before it gradually depreciates. Further our sensitivity analysis shows that delayed overshooting is an artefact of incorrect identifying restrictions. Specifically, we show that the response of SOE central banks to exchange rate fluctuations is the key to address the delayed overshooting puzzle. Moreover, the responses of exchange rates to both the SOE and US monetary shocks are broadly consistent with UIP and thus, we find little evidence of the forward discount puzzle. Finally, the six SOEs and US monetary shocks explain about 20 and 10 percents of the short-run exchange rate volatility, respectively.

The second chapter of my thesis deals with the impact of commodity prices on Australian macroeconomic fluctuations. I employ the SVAR methodology to quantify the impact of commodity demand and supply shocks on macroeconomic fluctuations in Australia. In my baseline model I quantify the impact of a (positive) commodity price shock that acts as proxy for (negative) commodity supply shocks. My model specification is inspired by Drechsel and Tenreyro (2018), Schmitt-Grohé and Uribe (2018) and Di Pace et al. (2020). I estimate an SVAR model for Australia to quantify the impact of exogenous disturbances in commodity prices on output and the trade balance. I achieve identification through the Cholesky decomposition of the contemporaneous matrix. The exclusion restrictions are imposed using the techniques developed by Arias et al. (2019). The commodity prices is ordered first which ensures that it acts as exogenous to the system. Moreover, for the lagged matrices, I achieve exogeneity of the commodity prices through highly informative priors. The model is estimated using the Bayesian techniques of Arias et al. (2019). For my baseline model, I find that a positive commodity price shock results in a contraction of output and a significant and persistent deterioration of the trade balance. Moreover, I find a negligible contribution of commodity price shocks to output fluctuations while these disturbances account for around 25% of the variance of the trade balance.

In my extended model, I introduce a *"foreign block"* by implementing a block exogeneity structure by applying additional exclusion restrictions on the contemporaneous matrix and tight priors on the lagged matrices. I relax one restriction in the block exogeneity structure which implies that the exchange rate impacts the iron ore exports through the lags. I estimate my extended SVAR model to quantify the impact of (positive) commodity prices and (positive) Chinese steel production shocks on output and the trade balance. The (positive) commodity prices and (positive) Chinese steel production act as proxies for (negative) commodity supply and (positive) commodity demand shocks, respectively. I find that in response to a commodity demand shock, output expands however, output contracts in response to a commodity supply shock. On the other hand, the trade balance deteriorates significantly in response to both the commodity supply and demand shocks. My findings contribute towards the literature by quantifying the responses of the Australian macroeconomy to Chinese steel production shocks (an increase in foreign demand of Australian iron ore and coal). This would help the policy-makers to better understand how the Australian macroeconomy reacts to the foreign commodity demand shocks.

In the third chapter, I explore the literature on proxy-SVAR models and the US monetary policy instruments. The proxy-SVAR models have emerged as an important tool to estimate the impact of fiscal and monetary shocks because of their non-controversial, point estimation identification approach. Proxy-SVAR models, also known as SVAR-IV models, rely on the satisfaction of traditional instrument variable conditions: relevance and exogeneity conditions. The relevance condition ensures that the instrument variable is correlated with the shock of interest while the exogeneity condition ensures that the instrument is uncorrelated with the rest of the shocks. Moreover, identification achieved through more than one external instruments makes the estimates sharper and more robust. I also provide a review of the evolution of US monetary policy instrument from the first instrument constructed by Beaudry and Saito (1998) to the most recent one by Bauer and Swanson (2023).

I also provide an application that highlights the differences between using a standard US monetary instrument (Gertler and Karadi, 2015) and an information-robust instrument (Jarociński and Karadi, 2020). Further, I employ a BP-SVAR model recently developed by Arias et al. (2021) to quantify the impact of US monetary shocks on the economic and financial variables. I find that, in response to a contractionary US monetary shock, the fall in real GDP and prices is more pronounced when I use the information-robust proxy as compared to when I use the standard proxy. Thus, my findings emphasis that results obtained from an information-robust proxy for the US monetary shocks are different from the ones obtained using the standard proxy.

In the future, I intend to extend my work of chapter 3 by quantifying the impact of US monetary shocks on small-open-economies (SOEs). My identification scheme would combine proxy-restrictions with exclusion and sign-restrictions. I intend to build on the methodological work of Arias et al. (2021) to the SOE settings by introducing a block exogeneity structure in the BP-SVAR model. Further, I would impose sign-restrictions on the structural parameters that are inline with macroeconomic and financial literature. My extended work will contribute towards the empirical macro SOE monetary literature by developing a novel, hybrid identification scheme that is sharper and more robust than the ones currently in use in the literature.

References

Arias, J. E., Caldara, D., and Rubio-Ramirez, J. F. (2019). The systematic component of monetary policy in SVARs: An agnostic identification procedure. *Journal of Monetary Economics*, 101:1–13.

Arias, J. E., Rubio-Ramírez, J. F., and Waggoner, D. F. (2021). Inference in Bayesian Proxy-SVARs. *Journal of Econometrics*, 225(1):88–106.

Bauer, M. D. and Swanson, E. T. (2023). A reassessment of monetary policy surprises and high-frequency identification. *NBER Macroeconomics Annual*, 37(1):87–155.

Beaudry, P. and Saito, M. (1998). Estimating the effects of monetary shocks: An evaluation of different approaches. *Journal of Monetary Economics*, 42(2):241–260.

Cushman, D. O. and Zha, T. (1997). Identifying monetary policy in a small open economy under flexible exchange rates. *Journal of Monetary economics*, 39(3):433–448.

Di Pace, F., Juvenal, L., and Petrella, I. (2020). Terms-of-trade shocks are not all alike.

Dornbusch, R. (1976). Expectations and exchange rate dynamics. *Journal of Political Economy*, 84(6):1161–1176.

Drechsel, T. and Tenreyro, S. (2018). Commodity booms and busts in emerging economies. *Journal of International Economics*, 112:200–218.

Gertler, M. and Karadi, P. (2015). Monetary policy surprises, credit costs, and economic activity. *American Economic Journal: Macroeconomics*, 7(1):44–76.

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

Schmitt-Grohé, S. and Uribe, M. (2018). How important are terms-of-trade shocks? *International Economic Review*, 59(1):85–111.