FISCAL AND MONETARY POLICIES IN SMALL OPEN ECONOMIES

BY

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Abstract

Apart from the introduction, this thesis includes three empirical chapters focusing on fiscal and monetary policies in small open economies.

The first chapter is titled "*Fiscal Space and Government-Spending & Tax-Rate Cyclicality Patterns: A Cross-Country Comparison, 1960–2016*". In this chapter, I compare fiscal cyclicality across advanced and developing countries, geographic regions as well as income levels over the 1960–2016 period, then identify factors that explain countries' government spending and tax-policy cyclicality. Public debt/tax base ratio provides a more robust explanation for government-spending cyclicality than public debt/output ratio but the reverse is true when capital investment is accounted for in government spending. On average, a more indebted (relative to tax base) government spends more in good times and cuts back spending indifferently compared with a low-debt country in bad times. I also find that country's sovereign wealth fund has a countercyclical effect in our estimation. Finally, the analysis depicts a significant economic impact of an enduring interest-rate rise on fiscal space, that is, a 10% increase in public debt/tax base ratio is associated with an upper bound of 5.9% increase in government-spending procyclicality.

The second chapter is titled "*Global Commodity-Price Shocks and Inflation Targeting in Emerging and Developing Countries*". This chapter examines if the inflation targeting regime makes a difference in the output and inflation responses when global commodity-price shocks take place. I apply the traditional SVAR with Cholesky decomposition approach for 99 emerging and developing countries over the 1990Q1-2016Q4 period and compute the median impulse responses of GDP growth and inflation for the IT and the non-IT countries. Following symmetric price shocks, I find that only the IT countries display persistent improvements in GDP growth, with cumulative responses remaining significant at least for six quarters after the shocks. The non-IT countries show insignificant responses in GDP growth, however. The analysis of asymmetric shocks also indicates that the IT countries are more resilient to the negative price shocks with long-lasting increases in GDP growth compared to the non-IT countries. In any case, the inflationary responses are transitory, similarly across both groups. In addition, the variance decomposition shows a modest role played by global commodity-price shocks in explaining the variations of output and inflation, with the fuel-price shock having the largest effects than the agriculture-price and metal-price shocks.

The third chapter is titled "*The Effect of Monetary Policy on the New Zealand Dollar: a Bayesian SVAR Approach*". This chapter uses the Bayesian SVAR approach introduced by Baumeister and Hamilton (2015) to examine the effect of New Zealand monetary policy shocks on exchange rate over the 1999-2020 period. I bring stock prices to the estimation and employ the co-movements of interest rates and stock prices to untangle the unexpected monetary policy shocks from other shocks that simultaneously affect interest rates and exchange rate. By choosing the priors consistently with the existing studies, this study is explicit about the influence of priors on posterior distributions and impulse response functions. The results show that, following an unexpected New Zealand monetary contraction, the value of New Zealand dollar against the US dollar increases immediately and even remains stronger in the long-run. There is no evidence of "delay overshooting" at least for one year in the estimation.

Chapter 1 Introduction

Fiscal and monetary policies are the two cornerstone macroeconomic policies in every country to stabilize the macroeconomy. Understanding how these policies can help secure a stable and safe economy in response to unexpected shocks from both external and internal economic conditions is crucially important to policymakers. In this thesis, I revisit some of these old but still very important issues about fiscal and monetary policies in small open economies.

Chapter 2 is titled "*Fiscal Space and Government-Spending & Tax-Rate Cyclicality Patterns: A Cross-Country Comparison, 1960–2016*" and was published in the Journal of Macroeconomics in May 2019 (Aizenman, Jinjarak, Nguyen, & Park, 2019). This chapter is motivated by the traditional wisdom that the fiscal capacity of countercyclical policy is the key resilience to unexpected shocks by mitigating business cycles and preventing a prolonged depression in the aftermath of financial crises (Auerbach, 2011; Ostry, Ghosh, Kim, & Qureshi, 2010). While the ratio of public debt to GDP is used frequently in policy discussions, Aizenman and Jinjarak (2011) argue that the ratio of public debt to tax base accounting for tax revenue may provide a more informative measure of the fiscal burden associated with the stock of public debt. Therefore, this chapter not only compares fiscal cyclicality across countries and regions over 1960-2016 but also and more importantly identifies the factors that explain countries' fiscal cyclicality patterns (including both government spending and tax-policy cyclicality), focusing on fiscal capacity proxies. I show that the public debt/tax base ratio provides a more robust explanation than the public debt/GDP ratio for government-spending cyclicality but the reverse is true when capital investment is accounted for in government spending. On average, a more indebted (relative to tax base) government spends more in good times and cuts back spending indifferently compared with a low-debt country in bad times. Other factors including economic structure and political risks also play a part in explaining fiscal cyclicality across countries. I also find that country's sovereign wealth fund has a countercyclical effect in the estimation. Finally, the analysis depicts a significant economic impact of an enduring interest-rate rise on fiscal space, that is, a 10% increase in public debt/tax base ratio is associated with an upper bound of 5.9% increase in government-spending procyclicality. Details of other extensions and robustness checks can be found in the Online Appendix (see sciencedirect.com/science/article/pii/S0164070418303938).

Chapter 3 is titled "*Global Commodity-Price Shocks and Inflation Targeting in Emerging and Developing Countries*". Given the increasing exposure to global commodity-price shocks due to globalization, I would like to examine if the inflation targeting (IT) regime makes a difference in the output and inflation responses in emerging and developing countries. Despite the empirical evidence favoring IT –i.e., IT helps anchor inflation and reduce inflation volatility without lowering output, the methodologies that fail to solve the endogeneity of monetary policy will leave those findings ambiguous. In this context, my study contributes to the abovementioned literature, not only revealing the impacts of global fuel, agriculture, and metal prices on a larger sample of less developed economies but also using the structural vector autoregression (SVAR) approach to better control the endogeneity of multiple macroeconomic variables, and thus providing more robust results. I apply the traditional SVAR with Cholesky decomposition method for 99 emerging and developing countries over the 1990Q1-2016Q4 period and compute the median impulse responses of GDP growth and inflation for the IT and the non-IT countries. Exercises for both symmetric and asymmetric price shocks are included. Following symmetric price shocks, I find that only the IT countries display persistent improvements in GDP growth, with cumulative responses remaining significant at least for six quarters after the shocks. The non-IT countries show insignificant responses in GDP growth, however. In the analysis of asymmetric shocks, I also find strong evidence indicating that the IT group is more resilient to the negative price shocks with long-lasting improvements in the GDP growth compared to the non-IT group. The impacts of the positive price shocks are generally silent. In any case, the inflationary responses are transitory, similarly across both groups. In addition, the variance decomposition shows a modest role played by global commodity-price shocks in explaining the variations of output and inflation, with the fuel-price shock having the largest effects than the agriculture-price and metal-price shocks. Finally yet importantly, I find significant evidence on the improved terms-of-trade, higher output growth, and lower inflation for the commodity net exporters following the shocks while the reverse is true for the commodity net importers.

The third chapter is titled "*The Effect of Monetary Policy on the New Zealand Dollar: a Bayesian SVAR Approach*". I am motivated by the fact that New Zealand is a small open trade-driven economy running a flexible exchange rate regime and hence the movements in the exchange rate may greatly contribute to New Zealand's domestic prices and economic activities. Understanding the responses of exchange rate to monetary policy shocks is very important to the monetary policymakers. Along with pursuing price stability as a key objective, the Policy Targets Announcement introduced in December 1999 thus called the Reserve Bank of New Zealand to focus on achieving stability not only in output and interest rate but also in the exchange rate. However, the literature of monetary policy – exchange rate analyses for New Zealand – is rather scant. This chapter greatly contributes to the literature, using the Bayesian SVAR approach to re-examine the impact of monetary policy shocks on New Zealand dollar (NZD) exchange rate. By applying this approach, I am transparent about the influence of the chosen priors on posterior distributions and impulse response functions, avoiding being too dogmatic as the traditional SVARs. I estimate a system of five variables including the US and New Zealand short-term interest rates and stock prices, and NZD exchange rate against the USD using the weekly data from 03/01/1999 to 18/09/2020. I employ stock prices to disentangle the monetary policy shocks from other shocks that jointly drive interest rates and exchange rate. The monetary policy shocks are then identified as unexpected changes in short-term interest rates. The results show that, to an unexpected increase in New Zealand short-term interest rate, the NZD appreciates immediately and keeps appreciating without a sign of "delay overshooting" at least for one year in the estimation. The findings are in line with other empirical studies with significant evidence that contradict the uncovered interest parity theory's prediction. The New Zealand monetary policy shocks, however, contribute very modestly to the variances of NZD exchange rate.

I am very delighted to share the data and codes of these chapters for your interest and/or replication. Those can be found publicly following my Github link (github.com/HienNguyenTK).

Chapter 2 Fiscal Space and Government-Spending & Tax-Rate Cyclicality Patterns: A Cross-Country Comparison, 1960–2016[1]

2.1 Introduction

The Global Financial Crisis (GFC) focused attention on unsustainable leverage growth as a key contributing factor in growing financial fragility associated with "bubbly" dynamics. Essentially a prolonged appreciation of financial and real estate markets increases the vulnerability to sharp asset valuation corrections. A deep enough correction may trigger banking crises and fire sales dynamics, potentially pushing the economy into a prolonged depression and a growing exposure to social and political instability.[2] Concerns about reliving the 1930s Great Depression explain the complex set of policies implemented by the US and other affected countries in the aftermath of the GFC: a massive infusion of liquidity in support of financial and banking systems and bailing out systemic banks and prime creditors. The forced deleverage of private borrowers, and the growing fear of a prolonged recession, induced higher household savings and lower investment, further deepening recessionary forces.

To counter these forces, many countries, therefore, experimented with fiscal stimuli aimed at mitigating the deepening recessions. Stabilizing the banking and financial systems, in addition to the stimuli, ended up sharply raising countries' public debt/GDP ratio, pushing advanced countries towards a public debt/GDP ratio of above 100%. Similar trends applied to emerging market economies (EMEs), driving their ratio of public debt to GDP upward, with some reaching well above 50%.[3] Notwithstanding the fact that the average public debt/GDP ratio of EMEs is below that of OECD countries, EMEs' lower tax base/GDP ratios, as well as the

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[1] This chapter was co-authored with Joshua Aizenman, Yothin Jinjarak, and Donghuyn Park and was published in the Journal of Macroeconomics in 2019. My individual contribution accounts for approximately 60% of the work, including collecting the data, doing analysis, and writing the paper.

[2] See Minsky (1992) for the financial instability hypothesis, which analyzes financial market fragility over the life cycle of an economy with speculative investment bubbles endogenous to financial markets. Rajan (2006) pointed out that banking deregulation during the 1980s–2000s increased leverage and risk-taking, contributing to greater exposure to financial stability associated with tail risks. Schularick and Taylor (2012) and Jordà, Schularick, and Taylor (2013) provided empirical evidence linking leverage, business cycles, and crises.

[3] Examples of advanced economies facing a public debt/GDP ratio of 100% or above during recent years include Belgium, Greece, Italy, Japan, Portugal, and Singapore. For merging markets, these include Cabo Verde, Eritrea, Jamaica, St. Kitts and Nevis, and Lebanon. A high level of debt (either public debt or external debt) is associated with lower growth, lower investment, and deteriorating macroeconomic policy environment (see Pattillo, Poirson, and Ricci (2002), Reinhart, Rogoff, and Savastano (2003), and Reinhart and Rogoff (2010)).

higher interest rates paid on their debt (due to sovereign risk premia), imply a rising fragility of EMEs compared with OECD countries. As such, while the ratio of public debt to GDP is used frequently in policy discussions, the ratio of public debt to tax base accounting for tax revenue may provide a more informative measure of the fiscal burden associated with the stock of public debt (Aizenman & Jinjarak, 2011). Henceforth, I refer to this fiscal measure as *limited fiscal space*.[4]

Importantly, the post-GFC trajectory failed to deal with leverage concerns: "At $164 trillion – equivalent to 225% of global GDP – global debt continues to hit new record highs almost a decade after the collapse of Lehman Brothers. Compared with the previous peak in 2009, the world is now 12% of GDP deeper in debt, reflecting a pickup in both public and nonfinancial private sector debt after a short hiatus. All income groups have experienced increases in total debt, but, by far, EMEs are in the lead." (International Monetary Fund, 2018). In other words, stabilizing a crisis triggered by an unsustainable leverage growth, in turn, contributed to a potentially untenable increase in leverage/GDP ratios.

For the past decade, the monetary easing associated with the US Federal Reserve (FED) and the European Central Bank policies in the aftermath of the GFC led to an unprecedented decline of policy interest rates and risk premia. These developments markedly reduced the flow costs of serving the rising public and private debt, thereby masking the increasing fragility associated with the rising aggregate leverage/GDP. This period has now passed: the (so far) robust recovery of the US, the gradual unwinding of the FED's balance sheet, the projected upward trajectory of the FED's funds rate, and the recovery of the Eurozone will impose growing fiscal challenges that will test countries' fiscal space and their ability to cope with projected higher interest rates by raising their resilience.

A key resilience margin is securing fiscal space – the fiscal capacity of a countercyclical policy aimed at mitigating business cycles and preventing a prolonged depression in the aftermath of financial crises (Auerbach, 2011; Ostry et al., 2010); see also Gavin, Hausmann, Perotti, and Talvi (1996) on the identification of fiscal procyclicality as a major amplifier of developing countries' vulnerability to shocks. Remarkably, over the last two decades leading to the GFC, a growing share of fiscal policies in developing countries and EMEs had graduated from

[4] The euro crisis provided a vivid example of how focusing on public debt/GDP below a certain threshold caused a failure to recognize the large heterogeneity of the tax base/GDP in the Eurozone (Aizenman, Hutchison, & Jinjarak, 2013). Similarly, the interest expense needed to serve the public debt as a share of tax revenue may provide a robust measure of the burden of serving the public debt and be more informative than the interest cost of the public debt/GDP ratio.

procyclicality and become countercyclical (see Frankel (2011) and Frankel, Végh, and Vuletin (2013)). Cross-country studies offer several explanations. Woo (2009) presented some evidence showing that social polarization, as measured by income and educational inequality, is consistently and positively associated with fiscal procyclicality, controlling for other determinants; there is also a robust negative impact of fiscal procyclicality on economic growth. Aizenman and Jinjarak (2012) found that higher income inequality is strongly associated with a lower tax base, lower de facto fiscal space, and higher sovereign spreads. Végh and Vuletin (2015) find that tax policy is less procyclical (more countercyclical) in countries with the better institutional quality and more financially integrated; tax and spending policies are conducted in a symmetric way over the business cycle. For brevity, Table 2.1 provides a summary of the related literature.[5]

It has been common wisdom that when there is a decline in economic activity of uncertain duration, government intervention should take the form of an expansionary fiscal policy. Yet, the viability of this option hinges on the degree to which the government has elastic enough access to borrowing. In the past, this has been the case for most OECD countries, allowing them to contemplate both tax cuts and fiscal expansions in recessionary times. In contrast, developing countries and emerging markets with the low tax base and large public debt burden have much more limited fiscal space, facing thereby more complex trade-offs. Specifically, cutting taxes or increasing government expenditure in recessionary times will increase the interest rate on their public debt, dampening thereby the stimulus and increasing the cost of serving the debt overhang. Gaining more fiscal space for these countries requires practicing more countercyclical policies in good times, allowing them to repay more of their debt in expansions times, increase thereby their fiscal space in recessions.

The rapid increase in public debt/GDP of most OECD countries in recent years suggests that the old fiscal dichotomy between the fiscal space of OECD and the rest of countries is blurred by now. This was vividly illustrated by the growing challenges facing Italy, Iceland, Ireland, and other OECD indebted members. Thereby, greater pro-cyclicality in good times may be the key to secure greater fiscal space in future rescissions. Considering these developments, this

[5] Related strands of the literature examine fiscal multipliers: see Ramey and Zubairy (2018), Leeper, Traum, and Walker (2017), and Ilzetzki, Mendoza, and Végh (2013); fiscal rules: see Budina, Kinda, Schaechter, and Weber (2012); and large fiscal adjustments: see Alesina, Favero, and Giavazzi (2015). Empirically, fiscal cyclicality, fiscal multipliers, fiscal rules, and large fiscal adjustments are intertwining issues; their relationships remain an open question and a challenge to address altogether in one go.

chapter focuses mostly on the cyclicality patterns exhibited in recent decades, and the projected challenges associated with possible future interest rate hikes.

Against this background, I assess definitions and empirical measures of fiscal cyclicality, compare fiscal cyclicality across Asia, Latin America, the OECD, and other regions, then identify factors accounting for spending- and tax-policy cyclicality patterns. I link the capacity of countercyclical policy to the fiscal space and the stage of economic and institutional development, as both are associated with the servicing capabilities of domestic and foreign debt. My analysis focuses on differences across the country groups and examines the role of economic structure (commodity versus manufacturing outputs), financial openness, as well as institutional and socio-economic factors (political risks, polarization, and ethnic polarization). The chapter concludes with an analysis of possible scenarios and suggested policies aiming at increasing the resilience of EMEs.

This study reveals a mixed fiscal scenery, where more than half of the countries are characterized by limited fiscal space, and fiscal policy is either pro- or acyclical. More limited fiscal capacity, as measured by the ratio of public debt to 3-years moving-average tax revenue and its volatility are positively associated with fiscal cyclicality, while public debt/GDP is statistically significant in several cases, suggesting that public debt/tax base ratio provides a robust fiscal-space explanation for studying government-spending and tax-rate cyclicality.[6] I calculate the impact of an enduring interest-rate rise on fiscal space, rank countries and regions by the fragility of their fiscal space to such an environment, and discuss policies to increase fiscal resilience. The rest of the chapter is organized as follows. Section 2.2 reports the empirical analysis with the baseline estimation and robustness checks. Section 2.3 provides the conclusion.

[6] Public debt/tax base in public finance is akin to the net debt to earnings before interest depreciation and amortization ratio in the corporate sector (aka Debt/EBITDA). Net debt to earnings ratio is a measurement of leverage, how many years it would take for a company to pay back its debt if net borrowing is zero, and EBITDA are held constant; used frequently by credit rating agencies. Ratios higher than 4 or 5 typically set off alarm bells; see S&P Ratings Direct (2013) https://www.spratings.com/scenario-builder-portlet/pdfs/CorporateMethodology.pdf and https://www.economist.com/business/2018/03/08/americas-companies-have-binged-on-debt-a-reckoning-looms.

2.2 Empirical analysis

2.2.1 Fiscal cyclicality: panel estimation

While there are many papers trying to estimate the cyclical sensitivity of fiscal policy in OECD countries and developing ones, either assessing by group or by individual country, the results are sometimes conflicting.[7] In this section, I simply compare cyclical patterns of fiscal instruments including government spending and tax rates in OECD and non-OECD countries and across income groups by exploiting as much data available as possible over the 1960–2016 period to estimate the following panel regression:

$$FISCAL_{it} = \alpha + \beta * \Delta logRGDP_{it} + \varepsilon_{it} \tag{2.1}$$

where *i* and *t* denote country and year, α is a constant term, ε_{it} is an error term, *FISCAL* is measured by either growth rate of real general government final consumption (*RGS*) or a tax rate, and *RGDP* is real gross domestic product (GDP).[8] I deflated the nominal government spending and GDP using the GDP deflator. For the government spending, the estimated $\hat{\beta}$ is the measure of spending-policy cyclicality: a positive and statistically significant coefficient indicates fiscal procyclicality; a negative and statistically significant coefficient indicates fiscal countercyclicality and a statistically insignificant coefficient indicates fiscal acyclicality. I also use Végh and Vuletin (2015)'s novel dataset of tax rates including value-added tax – VAT, personal income tax – PIT, and corporate income tax – CIT. However, the interpretations of the signs of tax-rate cyclicality coefficient are opposite those of spending-policy estimates. I leave in panel estimation the potential bias that is likely due to endogeneity and omitted variables, which will be addressed later in the baseline 2-step model below.

Table 2.2 shows government-spending cyclicality in the OECD and non-OECD countries with pooled-Ordinary Least Squares (Pooled-OLS) and Fixed Effects (controlling for country and year effects) specifications with robust standard errors. During 1960–2016, the non-OECD countries are more procyclical than the OECD countries, which is in line with Alesina, Campante, and Tabellini (2008). When it comes to tax-rate cyclicality as shown in Table 2.3, I find fairly consistent results as of Végh and Vuletin (2015) that OECD countries are fiscally

[7] See, for example, Lane (2003), Aghion and Marinescu (2007), Ilzetzki and Végh (2008), Bénétrix and Lane (2013), and Riera-Crichton, Végh, and Vuletin (2015).
[8] See the Online Appendix for more details on the data source.

procyclical in VAT, but countercyclical in CIT and PIT; whereas non-OECD countries are acyclical in VAT and associated with tax procyclicality in CIT and PIT.

As studying with panel data estimation of the cyclical patterns of government spending across income groups, the results shown in Table 2.4 come with no surprise as expected by Ilzetzki and Végh (2008). I find that higher-income countries are least fiscally procyclical, followed by upper-middle-income countries, lower-middle-income countries, and low-income countries.

2.2.2 Baseline empirical analysis

This section reports the empirical patterns of fiscal cyclicality across geographic regions, the OECD and non-OECD countries, and different income groups by estimating fiscal cyclicality by country. I then explore the determinants of countries' capacities in conducting countercyclical fiscal policy, focusing on tax base, public debt, economic structure, financial openness, as well as institutional and socio-economic factors.

My choice of explanatory variables takes into consideration the factors associated with fiscal capacity in conducting countercyclical policy – credit constraints, institutional quality, and tax-base variability (these factors are by no means exhaustive and their inclusion is subject to data availability). First, the credit constraints. The shape of the supply of funds facing the public sector in recessions is a key determinant of fiscal space. A flatter supply of funds implies an easier countercyclical policy funded by borrowing, which in turn is affected by the presence of buffers (international reserves, sovereign wealth funds), possibly managed by a fiscal rule that allows for more countercyclicality during recessions. Furthermore, low external and internal private and public debt-GDP ratios, as well as the ability to borrow in domestic currency, are associated with greater fiscal space, thereby allowing for cheaper borrowing in bad times. Second, the quality of institutions. Institutional quality is among the factors that are associated with fiscal space, which also include default history and inflation, and terms-of-trade volatility. In particular, the collection efficiency of tax revenue is impacted by the maturity of institutions and the spectrum of taxes (e.g. VAT and income taxes that are properly enforced). Greater political and ethnic polarization, inequality, and corruption may reduce a population's cooperation to pay their "fair share", thereby making tax collection harder, increasing the country's sovereign spreads, and leading to lower fiscal space. Public procyclicality may also be weaker in countries with more progressive taxes and transfers, as well as more countercyclical infrastructure expenditure, such as the use of infrastructure and housing investment as a countercyclical policy by the People's Republic of China. Third, the tax-base variability. The magnitude of revenue procyclicality depends on the production structure. A higher commodity share in GDP may be associated with greater exposure to the procyclicality of government revenues. Increased urbanization and international trade are associated with the easier collection of taxes, implying that tax compliance is higher and may result in tax-revenue procyclicality.

2.2.2.1 Empirical specification

To estimate the empirical patterns of fiscal-policy cyclicality and its determinants, I start by using a benchmark framework in the literature; see, for example, Woo (2009). Specifically, I proceed with the empirical analysis in two-step estimation:

Step 1: I run the following time-series regressions to measure the cyclicality of fiscal policy (spending, tax rates) by country over the 1960–2016 period:

$$FISCAL_{i,t} = \alpha_i + \beta_i * \Delta logRGDP_{i,t} + \varepsilon_{i,t} \tag{2.2}$$

where *FISCAL* again reflects either the spending side in growth rate of *RGS* or the tax side in VAT, PIT, and CIT respectively although I crucially aim to explain government-spending cyclicality more than cyclical tax-rate behavior. In the baseline model, I use a standard two-step Prais-Winsten regression to correct for the first-order autocorrelation in the residuals (AR(1)). The cyclical behavior of government spending and each tax rate by country is interpreted according to the signs of estimated $\hat{\beta}$ as in section 2.2.1.

There is some variation in the estimation of fiscal cyclicality in the literature where output-gap is commonly used.[9] The output gap is the deviation of real output from the potential series by applying filtering tools, for example, Hodrick-Prescott filter, Baxter-King filter, and Kalman filter. However, I use real GDP growth in baseline estimation as it is unlikely that any of the potential output estimation and filtering are commonly applicable across countries, though this controversial proxy is employed later in a robustness check. As a bottom line, I aim for an empirical framework that is straightforward and as easy to replicate as possible in a cross-country or panel sample. As constructing the sample for time-series estimation, I replace 33 countries with insufficient data from *World Development Indicators* (WDI) with *International Financial Statistics* (IFS) data and keep ones with at least 25 years of observations.

Step 2: I then study the determinants of fiscal (spending, tax rates) cyclicality for the 1960–2016 period by estimating the following cross-country regressions, in which I focus on the

[9] See Table 2.1 for some common estimation methods of fiscal-policy cyclicality.

measure of limited fiscal capacity, macroeconomic and socio-economic, as well as institutional variables:

$$\hat{\beta}_i = \alpha_0 + \theta_k * X_{ki} + \gamma_l * CONTROL_{li} + e_i \quad (2.3)$$

where *i* denotes country, the estimated $\hat{\beta}$ from the left hand side are either $\widehat{\beta_{GS}}$, $\widehat{\beta_{VAT}}$, $\widehat{\beta_{PIT}}$, and $\widehat{\beta_{CIT}}$ respectively which are estimated from equation (2.2), X_{ki} includes main variables of interest (limited fiscal capacity, export structure, and country risks), and $CONTROL_{li}$ includes macroeconomic variables, averaged over 1960–2016 period, including inflation, trade openness, financial openness, government size (i.e. its consumption share in the GDP), and political constraints. I estimate equation (2.3) by OLS with the White robust standard error (RSE).

A brief explanation of my selection of the determinants is needed. To calculate the ratio of public debt to tax revenue, I use general government tax including social contributions. To capture its second moments, I also calculate the volatility of limited fiscal capacity, using its standard deviation. As the size of tax base is persistent in the short- to medium-run, I also add an alternative measure of limited fiscal capacity, using the ratio of public debt to the 3-years moving-average of tax revenue.[10] In the estimation, I compare the public debt/tax base with the public debt/GDP ratios, as fiscal space is a multidimensional concept, exemplified in several fiscal indicators (International Monetary Fund, 2016). To account for socio-economic and institutional quality, I use several composite risk indicators, including financial, economic, and political conditions from *International Country Risk Guide*. I also control for political constraints – the *extent* to which the *executives face political constraints* in *implementing* their policy – drawn from Henisz (2002).[11]

2.2.2.2 Government-spending cyclicality and its determinants

Table 2.5 reports the summary of government-spending cyclicality for the 1960–2016 period based on the country-specific coefficients ($\widehat{\beta_{GS}}$).[12] Looking across geographic regions, the government-spending cyclicality of Sub-Saharan Africa is the highest among the estimates (0.89; most procyclical), followed by Latin American and the Caribbean (0.77), the Middle East and North Africa (0.69), East Asia and Pacific (0.46), Europe and Central Asia (0.41),

[10] Note that our results are robust to using different time-average for the denominator of the fiscal-capacity variable (for example, 2-years moving average and 5-years moving average of tax revenues). The detailed estimations will be provided upon request.
[11] See the Online Appendix for the descriptive summary statistics of the variables in our baseline sample.
[12] Empirical patterns of government-spending cyclicality by country will be provided upon request.

South Asia (0.35), while North America has negative and the lowest estimates (-0.25; most countercyclical). For OECD and non-OECD countries, the latter group, on average, is more fiscally procyclical (0.74) than the former one (0.19). Looking across income levels, the degree of procyclicality is negatively associated with income level, i.e. the low-income countries are most fiscally procyclical (0.93) followed by lower-middle-income countries (0.78), upper-middle-income countries (0.69), and the high-income group (0.32). Overall, the findings from country-specific regressions are consistent with those in panel estimation.

What might explain the cross-country differences? Table 2.6 reports the estimation of government-spending cyclicality coefficients ($\widehat{\beta_{GS}}$) on socio-economic and institutional variables over the 1960–2016 period. The main findings are as follows. Political constraints (*polcon*) are negatively associated with government-spending procyclicality, implying a greater degree of political constraints preventing policy discretions, which in turn limits fiscal procyclicality. Inflation *(inf)* is positively associated with fiscal procyclicality, suggesting the role of macroeconomic instability, seigniorage, and passive monetary policy. Trade openness *(trade)* and financial openness *(TAL)* are negatively associated with fiscal cyclicality, implying that the countries are less likely to conduct procyclical fiscal policy if they are more trade and financially open; fiscal multipliers are smaller for more open economies. Government size, as measured by its consumption share in GDP *(gs)*, is statistically insignificant in explaining fiscal-policy procyclicality.

More limited fiscal capacity, as measured by public debt/tax base ratio *(fiscal, lfiscap)* and its volatility *(fiscal_vol, lfiscap_vol)* are positively associated with fiscal procyclicality, while public debt/GDP ratio *(debt)* and its volatility *(debt_vol)* are statistically insignificant, suggesting that the ratio of public debt to tax base provides a robust explanation for government-spending procyclicality for the 1960–2016 period. Manufacturing export share *(manu)* is negatively associated with while natural resource export share *(nare)* is positively associated with fiscal procyclicality. The composite risk and three component risk indices (economic, financial, and political), as well as eight out of twelve political component risk indices (including social economic conditions, investment profile, internal conflict, corruption, military in politics, law and order, ethnic tensions, and bureaucracy quality), are negatively associated with fiscal procyclicality, thus indicating that higher country risk is associated with higher fiscal procyclicality.

2.2.2.3 Tax-rate cyclicality and its determinants

Similarly, only countries with at least 25 years of tax-rate observations are taken, hence I am left with 35 countries with VAT, 62 countries with PIT, and 62 countries with CIT. However, because of the infrequent adjustment of tax rates and the non-convergence of the AR(1) coefficient, the estimated coefficients of some countries cannot be obtained in the two-step Prais-Winsten procedure and the sample size thus becomes smaller.[13]

After obtaining tax-rate cyclicality coefficients ($\widehat{\beta_{VAT}}$, $\widehat{\beta_{PIT}}$, and $\widehat{\beta_{CIT}}$), I then regress them on socio-economic and institutional variables with different sets of control variables in each case, using OLS (RSE) whose results are summarized in the Online Appendix. I find that $\widehat{\beta_{VAT}}$ is negatively associated with economic and financial risk indices, i.e. VAT becomes more fiscally procyclical as economic and financial risk decreases. I also find that personal income tax rate is more procyclical with: more limited fiscal space (*lfiscap*) and its volatility (*fiscal_vol, lfiscap_vol*), lower manufacturing export share (*manu*), higher socio-economic and political risks (*CRI, ERI, FRI, PRI, socecon*, *inconflict*, *exconflict*, *ethnic*, and *democracy*), and lower institutional quality (*corrupt*, *law*, and *bureau*). The latter is consistent with the findings of Végh and Vuletin (2015). Among the socio-economics and institutional variables, only natural resource export share (*nare*) and religious tensions (*religious*) are associated with CIT-cyclicality. Corporate income tax rate becomes more procyclical with higher natural resource export share and lower religious risks.

2.2.2.4 Economic significance on fiscal cyclicality

To derive the economic impact of explanatory variables on fiscal cyclicality, I calculate and rank their economic significance by multiplying their (sample) standard deviation with their estimated coefficient from corresponding regression, thereby approximating the impact of one standard deviation change of that explanatory variable on the degree of fiscal cyclicality. For government-spending cyclicality, Figure 2.1 shows the economic impact of natural resource export share (positive), limited fiscal space (positive), socio-economic and institutional risks (negative), and manufacturing export share (negative). In Figure 2.2, the economic significance of public/tax base is calculated for each geographic region. The finding suggests that East Asia & the Pacific and Sub-Saharan Africa are more exposed to the positive association between fiscal-spending procyclicality and fiscal space than the other regions.

[13] Empirical patterns of tax-rate cyclicality by country will be provided upon request.

Regarding tax-rate cyclicality, the economic impacts of the explanatory variables on each tax-rate cyclicality vary considerably. For VAT-cyclicality, most socio-economic, financial, and institutional risks have negative and higher economic impacts than limited fiscal capacity whereas natural resource share has the largest positive economic impact. Nonetheless, the economic impact patterns of these variables on PIT-cyclicality are quite opposite: limited fiscal capacity (negative and large), natural resource share (negative), socio-economic and institutional risks (positive), and manufacturing export share (positive). This may suggest that the cyclicality patterns of VAT differ significantly from those of PIT. Interestingly for CIT-cyclicality, I find significant effects of the economic significance of manufacturing export share (positive), limited fiscal capacity (positive), religious tensions (negative and largest), natural resource share (negative), while the impacts of other institutional variables are mixed and insignificant.[14]

2.2.3 Robustness checks

The fiscal-cyclicality literature suggests that the reduced-form relation between government spending and output (equation (2.2)) is the appropriate framework to study fiscal cyclicality and there is "no strong reason to exclude any equilibrium feedback from fiscal policy to the level of output" (Lane, 2003). As the reverse causality in equation (2.2) could result in the endogeneity bias, I addressed this potential issue below.

I note that the residuals in the Prais-Winsten approach are assumed to follow AR(1) process; they are unobservable. I conduct robustness checks by re-estimating fiscal cyclicality by country in the 1st step estimation, correcting for heteroscedasticity and serial correlation (up to AR(2)). The estimation is done with OLS (RSE) and OLS with Newey-West standard errors. Furthermore, I use the Two-Stage Least Squares (2SLS) to address potential endogeneity issue of real GDP growth rate in equation (2.2). Following the literature, I use global liquidity shock (*SHOCKGL*) measured as the real return on 6-month Treasury bills weighted by countries' de jure financial openness using Chinn and Ito (2006) index as an excluded instrument; this variable is a proxy for country's exposure to global liquidity. In addition, an external shock captured in the weighted real GDP growth of trading partners (*SHOCKJP*) and the US business cycle (*KAUS*) defined by the National Bureau of Economic Research (NBER) is also used as alternative excluded instruments (IVs). The shocks from the US business cycle is also weighted by countries' de jure financial openness using Chinn and Ito (2006) index. I check for each

[14] See the Online Appendix for the economic significance of variables dealing with tax-rate cyclicality.

country the relevance and exogeneity of the instruments, and the over-identification tests; subject to data availability, the validity of these instruments varies across countries. Then I re-estimate the 2nd step on the cross-country regression, using the Weighted Least Squares (WLS) with the weight being the inverse of standard errors of $\hat{\beta}$ from the 1st step.

Table 2.7 summarizes the robustness checks for government-spending cyclicality using instrumental variables: the results are supportive of the findings in the baseline model. Public debt/tax base ratio and its volatility are positive and significant in most of the specifications but public debt/GDP ratio and its volatility are less so, suggesting the ratio of public debt to tax base as a more informative variable for understanding the cyclical government-spending pattern over 1960–2016 period. The impacts of other variables are consistent with the baseline findings, including natural resource share (positive), manufacturing export share (negative), socio-economic and institutional risks (negative).

I next look into the choice of the output used in estimating the fiscal cyclicality. The robustness check on the country-specific government-spending cyclicality is done by regressing the change of real government spending on the output gap (deviation of real GDP from its Hodrick-Prescott trend); in order to compare with the baseline model using the real GDP growth. The smoothness parameter is set to 6.25 for the annual data following Ravn and Uhlig (2002). Across the specifications: Prais-Winsten, OLS with Newey-West standard errors, or 2SLS in the 1st step and WLS in the 2nd step, I find consistent results with the baseline model. The ratio of public debt to tax base and its volatility are positive and significant while the ratio of public debt to GDP and its volatility are not, and institutional variables are negative and significant.[15]

Note that I have so far used the 1st-step estimated coefficients regardless of the statistical significance; their qualitative and quantitative variations reflect the fiscal countercyclicality, procyclicality, or acyclicality in the sample. That is, I do not normalize/set them to zero if they are statistically insignificant in the baseline estimation. This practice is consistent with the existing studies using the 2-stage estimation, i.e. the average time-series estimates as in Lane (2003) and Woo (2009) or the x-year rolling window estimates as in Nerlich and Reuter (2015) and Guerguil, Mandon, and Tapsoba (2017). However, to check robustness of the estimation, I re-do the analysis by setting insignificant estimated $\hat{\beta}$ to zero for both Prais-Winsten and OLS

[15] Subject to the availability of higher-frequency data (i.e. quarterly), separating cycles from trends would be a useful extension, especially for the emerging markets. Bashar, Bhattacharya, and Wohar (2017) disentangle the correlations in cycles from the correlations in slopes of the relevant variables for 11 OECD countries. They find that the growth (slope) of government spending is positively associated with real GDP in several countries.

estimates; while though some variation is inevitable, I do not find this change to overturn the main findings in the baseline estimation.[16]

I note that some of the association between tax-rate cyclicality and the control variables is sensitive to the choice of econometric specifications; coefficients of several variables are weak statistically in the cross-country regressions when 2SLS is used in the 1st step. The Online Appendix summarizes the 2nd-step regressions for tax-rate cyclicality under alternative estimations. I find that VAT-cyclicality is associated with the volatility of debt/GDP ratio in several cases (negative), natural resource share (positive), manufacturing export share (negative), and different institutional risk indices (negative). In contrast, PIT-cyclicality is shown to be associated with limited fiscal capacity and its volatility (negative), natural resource share (negative), manufacturing export share (positive), and various socio-economic and political risks proxies (positive). CIT-cyclicality is associated with natural resource share (negative), manufacturing export share (positive), and a range of institutional risk indices (positive); when estimated by the Prais-Winsten specification, it is also negatively associated with religious tensions.

2.2.4 Fiscal cyclicality at good times and bad times

Recent studies point to the asymmetry of fiscal cyclicality in good times vis-à-vis bad times. Alesina, Barbiero, Favero, Giavazzi, and Paradisi (2017) use the narrative-identified exogenous fiscal stabilizations (i.e. the stabilization is not supposed to be correlated with the economic cycle) to show that, for 16 OECD countries, the cuts in government spending and transfers are much less harmful than tax hikes. Auerbach and Gorodnichenko (2017) show that for G-7 countries, government-spending shocks do not lead to persistent increases in debt to GDP ratios or costs of borrowing, especially during periods of economic weakness. While these lessons from the advanced economies are informative, this study is concerned with both industrial and developing countries. I note that the estimated $\hat{\beta}$'s so far (from equation (2.2)) provide the interesting patterns of government-spending and tax-rate cyclicality; I could delve further by separating the fiscal reactions in good times from those in bad times. Define good times as the periods with positive real GDP growth rate and bad times as the periods with negative real GDP growth rate, the regression equation is as follows:

$$FISCAL_{i,t} = \alpha_i + \gamma_i * \Delta logRGDP_{i,t} + \lambda_i * D_{i,t} + \theta_i * (\Delta logRGDP_{i,t} * D_{i,t}) + v_{i,t} \quad (2.4)$$

[16] These additional robustness results are available in the Online Appendix.

where $D_{i,t}$ = 0 if good times (strong economic growth in country i at time t), $D_{i,t}$ = 1 if bad times (weak economic growth), and θ_i tests the asymmetric response of government spending in bad times compared to good times for country i. To obtain Prais-Winsten estimators, I regress the following separately:

$$\text{D = 0: } FISCAL_{i,t} = \alpha_i + \gamma_i * \Delta logRGDP_{i,t} + u_{i,t} \quad (2.4a)$$

$$\text{D = 1: } FISCAL_{i,t} = (\alpha_i + \lambda_i) + (\gamma_i + \theta_i) * \Delta logRGDP_{i,t} + \omega_{i,t} \quad (2.4b)$$

I find that the empirical patterns of fiscal cyclicality (either spending or tax) vary significantly across specifications (Prais-Winsten and OLS). Essentially, I have mixed findings of asymmetries in government-spending cyclicality patterns and the tax-rate policy cyclicality across good times and bad times.[17] When I re-estimate the government-spending estimators on the determinants, I find that the associations between the government-spending procyclicality $\widehat{\beta_{GS}}$ and explanatory variables during good times are largely similar to the baseline model: positive with limited fiscal capacity and its volatility as well as natural resource share of exports, and negative with manufacturing share of exports and country risks (see Table 2.8, columns 1–2). The volatility of public debt/GDP ratio is also positively associated with government-spending procyclicality in good times. In bad times, the volatility of limited fiscal capacity and investment profile are statistically significant and negatively associated with government-spending cyclicality (see column 3). Hence, it seems that in bad times, public debt, tax base, and investment confidence play a larger role in the government-spending cyclicality. This implies that a more indebted (relative to tax base) government spends more in good times and cuts back indifferently compared with a low-debt country in bad times.

The results so far suggest that, for both government-spending and tax-rate cyclicality, there is no one-size-fits-all explanation for all (OECD/developing) countries at all (good/bad) times. In essence, fiscal space, trade and financial openness, the export shares of natural resource and manufacturing, inflation, and institutional risks are associated with the cross-country patterns of fiscal cyclicality.

2.2.5 Sovereign wealth funds and government-spending cyclicality

I delve deeper into fiscal behavior by looking at the role of sovereign wealth funds on government-spending cyclicality by regressing the following:

[17] The detailed estimated coefficients at good times and bad times by country will be provided upon request.

$$\hat{\beta}_i = \alpha_0 + \gamma_k * CONTROL_{ki} + \rho * SWF_i + \delta_1 * fiscap_i + \delta_2 * (SWF_i * fiscap_i) + \theta_1 * CRI_i + \theta_2 * (SWF_i * CRI_i) + \varepsilon_i \quad (2.5)$$

, where the dummy *SWF* = 1 if the country has a sovereign wealth fund in operation starting at any point during the 1960–2016 period; *SWF* = 0 otherwise. Focusing on the fiscal space and institutional risks, I include their interactions with the *SWF* variable. I regress equation (2.5) using the WLS estimation with real GDP (at 2010 US$) as the weight. Table 2.9 reports the estimation results for the full sample (1960–2016) and a sub-sample of good times. The estimates for bad-times are qualitatively similar but statistically insignificant. The negative coefficients of *SWF* interactions with public debt/tax ratio and institutional quality suggest that the existence of sovereign wealth funds has a negative association with the government-spending procyclicality. Essentially, the findings point to the benefit of investing in sovereign wealth funds as the countercyclical fiscal buffers in good times to mitigate tax revenue shortfalls in bad times, thereby increasing the availability of countercyclical spending policy.

2.2.6 Excluding social contributions from tax base

Tax base has several components, among which social contributions play an important role in many countries' budgets. What would happen if I repeat the estimation using tax base without social security contributions? However, I do not find much difference in the regression results as well as the economic significance of each explanatory variables to $\widehat{\beta_{GS}}$, $\widehat{\beta_{VAT}}$, $\widehat{\beta_{PIT}}$, and $\widehat{\beta_{CIT}}$ in the whole sample period and sub-periods.

2.2.7 Government-spending cyclicality with capital investment

To further check for the robustness, this section provides the estimation results with alternative government spending series. I check whether the main findings would hold if the capital investment is accounted for in the government spending. I re-estimate both the panel estimation and the 2-step cross-country estimations using *World Economic Outlook (*WEO*)*'s general government total-expenditure; the government spending is defined as total expense plus the net acquisition of nonfinancial assets for the 1980–2016 period. The net acquisition of nonfinancial assets equals gross fixed capital formation less consumption of fixed capital plus changes in inventories and transactions in other nonfinancial assets. Using this alternative data set, I find that, overall, the rankings of spending-policy cyclicality across OECD and non-OECD countries, as well as geographic regions and income levels based on both panel and country-specific time-series estimations, are in line with the baseline data set (that is, the government spending without the capital investment). However, there are some discernible differences, as summarized in Table 2.10. Public debt/GDP ratio and its volatility are significantly and positively associated with $\widehat{\beta_{GS}}$ but public debt/tax base ratio and its volatility are no longer significant except in several cases. Manufacturing export share remains negatively associated with government-spending procyclicality while natural resources export share is insignificant in some regressions. Composite risk, economic risk, government stability, socioeconomic conditions, and law and order indices are consistently and negatively associated with fiscal procyclicality as in the baseline model, and other proxies for financial risk, political risk, investment profile, internal conflict, external conflict, corruption, military in politics, ethnic tensions, and bureaucracy are significant in several estimations.

The findings on the cyclicality of government spending with the capital expenditure suggest that it may be useful to look into not only the size but also the composition of government expenditures (i.e. healthcare, education, defense) to study which components of the spending drive the fiscal cyclicality. Given the heterogeneous population and income inequality, it is quite likely that the composition of government spending is influenced by trade and financial openness, political economy consideration, the availability of social safety nets, and fiscal capacity.[18]

2.2.8 External debt as an alternative indicator of fiscal constraint

In addition to public debt, external debt has also been studied in the literature as a measure of countries' fiscal constraints. A range of empirical studies found non-linear negative effects of external debt on growth (see Pattillo et al. (2002) and Reinhart and Rogoff (2010)).

I also verify the external debt/GDP as an alternative to the public debt/GDP for measuring and calculating alternative indicators of fiscal capacity and its volatility. Data on the total external debt is from WDI (defined as the debt owed to nonresidents repayable in currency, goods, or services). Total external debt is the sum of public, publicly guaranteed, and private nonguaranteed long-term debt, use of IMF credit, and short-term debt. Short-term debt includes all debt having an original maturity of one year or less and interest in arrears on long-term debt. Overall, the additional estimation results do not change the main conclusions. I find that the alternative indicators of fiscal capacity and fiscal capacity volatility based on the external debt

[18] Shelton (2007) studies the size and composition of government expenditure across countries from 1970–2000. It is likely that the spending composition is time-varying, especially after the GFC, and because of the growing concerns over income inequality across industrial and developing countries in recent years.

are not significant in the baseline sample (WDI data for the 1960–2016 period) but are positively associated with government spending procyclicality in several estimations using WEO data accounting for capital investment, though these results are not robust across specifications. Regarding the external debt/GDP ratio and its volatility, most of their coefficients are positively significant in the WEO sample but insignificant in the WDI sample. That is, as I account for capital investment, both public debt and external debt are significantly and positively associated with government expenditure cyclicality. Overall, the results confirm that, in the baseline sample, fiscal capacity and its volatility measured by public debt/tax revenue are informative in explaining government spending behavior, while those measured by external debt/tax revenue are not.

2.2.9 Determinants of government-spending cyclicality by region

It is clear that the degrees of fiscal cyclicality differ markedly across countries and regions. Given the differences in the economic development and institutions, it is unlikely that I can come up with a sweeping explanation, but at least I can try. In order to examine the economic significance of each explanatory variable on government-spending cyclicality on the regional basis, I repeat the 2^{nd}-step estimation by region. North America and South Asia are dropped due to insufficient data. Hence, I study in details five geographic regions: East Asia and the Pacific, Europe and Central Asia, Latin America and the Caribbean, the Middle East and North Africa, and Sub-Saharan Africa. Figure 2.3 shows the economic impacts by region, focusing closely on the associations of public debt, export structure, and country risks with the government-spending cyclicality.

East Asia and the Pacific: limited fiscal capacity has positive and significant impacts while governance and institutional quality, as measured by most of the country risk indices, have large and negative effects on fiscal procyclicality. Europe and Central Asia: manufacturing export share and most of the institutional indices have a negative association with government-spending procyclicality; however, public debt/GDP ratio has a statistically significant and negative association with the government-spending cyclicality (i.e. lower debt/GDP ratio is associated with more fiscally procyclical). Latin America and the Caribbean: better institutional quality, more stable politics, a smaller share of natural resource exports, and lower public debt/GDP ratio are associated with lower government-spending procyclicality. The Middle East and North Africa: somewhat intriguing as good scores on some socio-economic and political-stability variables are negatively associated with fiscal procyclicality as expected, but the democracy risk index is positively associated with fiscal procyclicality. Sub-Saharan African countries: interestingly some evidence of better institutional quality positively associating with procyclicality, yet the positive association of limited fiscal space and negative association of manufacturing export share with government-spending procyclicality are the most obvious in this region.

2.2.10 Fiscal space in a deteriorating macro environment

An enduring rise in the global interest rate will result in increasing the cost of borrowing and servicing public debt. I now look closely at the economic significance of limited fiscal capacity on government-spending cyclicality, using the public debt/3-years moving-average tax base ratios, shown in Figure 2.4, and calculate to see what would happen if fiscal capacity drops by 10%. It is, specifically, 0.1*(Regional-specific estimated coefficient of public debt/tax base)*(Actual average ratio of regional-specific public debt/tax base over the 1960–2016 period). The top panel in Figure 2.4 shows the limited fiscal capacity, as measured by the actual ratios of public debt/tax base and public debt/3-years moving-average tax base respectively, average over 2010–2016 period. East Asia and the Pacific and the Middle East and North Africa have on average lower fiscal capacity compared to Latin America and the Caribbean, Sub-Saharan Africa, and Europe and Central Asia. However, as shown in the bottom panels, Sub-Saharan Africa is distinctly fragile fiscally, being exposed to large government-spending procyclicality if the macroeconomic environment and its fiscal space deteriorate. Based on the calculation, a 10% increase in public debt/tax base ratio is associated with an upper bound of 5.9% increase in government-spending procyclicality.

In addition, I look at the economic impact of deteriorating fiscal space, i.e. if a fiscal capacity drops by 10%, what would happen to the government-spending cyclicality. Specifically, I calculate: 0.1*(Actual average country-specific public debt/tax base)*(Regional-specific estimated coefficient of public debt/tax base). I use regional-specific coefficient in place of country-specific coefficient as there is insufficient country-level data to estimate a country-specific 2nd-step regression (i.e., equation (2.3); $\widehat{\beta_{GS}}$ = f(public debt/tax base, control variables)). As shown in the upper panel of Figure 2.5 (using 3-year tax base; the Online Appendix provides the calculation using 1-year tax base), Iraq, Japan, Singapore, Egypt, Greece, Libya, Yemen, Jamaica show limited fiscal capacity based on the 2010–2016 data, accumulating public debt four to eight times larger than their tax base (Iraq has public debt approximately forty times higher than its tax revenue). According to the calculation, shown in the lower half of Figure 2.5, fiscally fragile countries are mostly in Sub-Saharan Africa (Republic of Congo, Nigeria, Rwanda Seychelles,) and a few cases in East Asia and the Pacific (Vietnam, Indonesia, Cambodia; and Japan, which is rather an exceptional case[19]).

2.3 Concluding remarks

This study reveals a mixed fiscal environment in which more than half of the countries in the study are characterized by limited fiscal space and fiscal policy is either pro- or acyclical. It confirms that OECD and higher-income countries are more fiscally countercyclical than non-OECD and lower-income countries. I also find that, compared to public debt/GDP, the ratio of public debt to tax base is a robust measure of limited fiscal space and provides a more robust explanation for government-spending cyclicality but the reverse is true when capital investment is accounted for in government spending. Moreover, the cyclicality is asymmetric: on average, a more indebted (relative to tax base) government spends more in good times and cuts back indifferently compared with a low-debt country in bad times. Lastly, the analysis predicts that a 10% enduring increase of interest-rate is associated with an upper bound of 5.9% increase in government-spending procyclicality.

Considering the sizable increase in total leverage/GDP in the aftermath of the GFC, countries could use the global recovery as an opportune time to invest in greater fiscal space, which could be done by increasing the tax base. Countries could also benefit by investing in countercyclical fiscal buffers, including the accumulation of sovereign wealth funds in good times to mitigate tax revenue shortfalls in bad times (e.g., Chile, Norway); indeed, it is shown that countries' sovereign wealth funds have a countercyclical effect in the estimation. Likewise, a deeper safety net will add a countercyclical buffer that mitigates the adverse income effects of recessions, thus reducing income inequalities over time.

A limitation of this study is that, due to data constraints, I focus on the general government and thereby overlook the contribution of local and state government in a federal union system to cyclicality patterns. Chances are that controlling for these issues, I would find deeper pro- or acyclical patterns (e.g., in the US, state governments are frequently forced to apply procyclical expenditure patterns, which means cutting budgets at the time of deep and prolonged recessions). Furthermore, while it is widely agreed that procyclical fiscal policy should be mitigated as much as possible (International Monetary Fund, 2017), there is no consensus on

[19] This is partly due to the historically massive distribution and subscription of government bonds through the Yū-cho Ginkō (less true today) and the preference of Japanese bondholders to the domestic government bonds.

the practical approach, i.e. which spending components receive priority, and the fiscal rules to achieve such optimal degree of fiscal cyclicality.

As different governments face a wide range of political pressures and several targets (i.e. allocation efficiency, redistribution, debt stabilization, and structural reforms) with various ranking priorities, fiscal challenges are mostly context-specific without one-size fitting for all countries at all times. The cross-country findings suggest that I need a better understanding of the mixes of (i) components of government spending, public debt, and tax base; (ii) fiscal policy, monetary policy, socio-economics, and institutions; and (iii) the role of central banks and quasi-government entities (e.g. sovereign wealth funds, state-owned enterprises). I study these monetary-fiscal-political economy interactions in my follow-up.

Table 2.1 Empirical literature on estimation of fiscal-policy cyclicality

Studies	Methodology	Measurement of fiscal cyclicality	Sample	Key findings
Lane (2003)	$\Delta\, log(G_{it}) = \alpha_i + \beta_i * \Delta\, log(Y_{it}) + \varepsilon_{it}$ (1) $\widehat{\beta_i} = \alpha_0 + \alpha_1 Z_i + \varepsilon_i$ (2) *G*: various components of government spending *Y*: real GDP *Z*: control variables (1): Country regression using OLS procedure with a correction for AR(1) in the residuals; (2): WLS.	$\beta_i > 0$: procyclicality $\beta_i < 0$: countercyclicality	22 OECD countries 1960–1998	The level of procyclicality varies across spending categories and countries. Volatile output and dispersed political power are associated with government spending procyclicality.
Kaminsky, Reinhart, and Végh (2004)	$\rho(GS,OG)$, $\varphi(inflationtax,OG)$ ρ, φ: country correlation coefficient *GS*: cyclical government spending; *OG*: output gap. The cyclical series are estimated by the Hodrick-Prescott filter method.	$\rho > 0$: procyclicality $\rho < 0$: countercyclicality $\varphi > 0$: countercyclicality $\varphi < 0$: procyclicality	104 countries 1960–2003	Most OECD countries have countercyclical fiscal policy while most of developing countries have procyclical fiscal policy.
Talvi and Végh (2005)	$\rho(FC,OG)$, $\varphi(inflationtax,OG)$ ρ, φ: country correlation coefficient *FC*: cyclical government consumption, cyclical revenue; *OG*: output gap. The cyclical series are estimated by the Hodrick-Prescott filter method.	$\rho > 0$: procyclicality $\rho < 0$: countercyclicality $\varphi > 0$: countercyclicality $\varphi < 0$: procyclicality	56 countries 1970–1994	Fiscal revenues are procyclical in both developing and industrial countries. Government consumption in the G7 countries is acyclical when that in non-G7 industrial countries and developing countries is procyclical. Inflation tax rate is countercyclical in industrial countries and procyclical in developing countries.
Aghion and Marinescu (2007)	$\frac{b_{1t}-b_{i,t-1}}{y_{it}} = -a_{1it}y_{gap,it} + a_{2it} + \varepsilon_{it}$ (1) *b*: gross government debt *y*: GDP y_{gap} is computed using Hodrick-Prescott filter (1): 10-year centered rolling window; local Gaussian-weighted OLS; AR(1) Markov Chain Monte Carlo process	$a_{1it} > 0$: countercyclical $a_{1it} < 0$: procyclical	19 OECD countries 1961–2005	The budget deficit has become increasingly countercyclical in most OECD countries over the past 20 years. However, this trend has been significantly less pronounced in the EMU.
Alesina et al. (2008)	$\Delta\, F_{it} = \alpha_i + \beta_i * OG_{it} + \gamma X_{it} + \lambda F_{it-1} + v_t + \varepsilon_{it}$ (1) *F*: government surplus or public spending; *OG*: output gap, *X*: control variables. *OG* is estimated by the Hodrick-Prescott filter method. (1): Fixed Effects where *OG* of country *i* is instrumented by *OG* of the region of country *i*.	β_i is interpreted depending on the fiscal policy variable	83 countries 1960–2003	Fiscal policy is procyclical in many developing countries. Political distortion (i.e. corruption) is positively correlated with procyclicality of fiscal policy.

	Alternatively, (1) is estimated by country to get $\widehat{\beta_i}$ and then run cross-country regression of $\widehat{\beta_i}$ on X_i.			
Ilzetzki and Végh (2008)	$\Delta\ log(GS_{it}) = \alpha_i + \beta_i * \Delta\ log(Y_{it}) + \varepsilon_{it}$ (1) Y: output, GS: government spending, or its components (1) is regressed using alternative methods include 2SLS, GMM, OLS estimation of simultaneous equations, Granger causality tests, VAR.	$\beta_i > 0$: procyclicality $\beta_i < 0$: countercyclicality	49 countries 1960–2006	Fiscal policy is always procyclical in developing countries and acyclical/procyclical in high-income countries.
Woo (2009)	$\Delta\ log\ GS_{it} = \alpha_i + \beta_i * \Delta\ log\ Y_{it} + \varepsilon_{it}$ (1) $\widehat{\beta_i} = \alpha_0 + \alpha_1(\text{Social polarization})_i + \phi X_i + \varepsilon_i$ (2) GS: real general government spending Y: real GDP X: control variables (1): Country regression using Prais-Winsten procedure; (2): OLS, WLS.	$\beta_i > 0$: procyclicality $\beta i < 0$: countercyclicality	96 countries 1960–2003	Developing countries are more procyclical than OECD countries. Latin America is the most fiscally procyclical region, followed by Sub-Saharan Africa and East Asian. Income inequality and educational inequality is positively associated with fiscal procyclicality.
Végh and Vuletin (2015)	$Tax_{it} = \alpha_i + \beta_i * OG_{it} + \varepsilon_{it}$ (1) $\Delta Taxrate_{it} = \alpha_i + \beta_i * \Delta\ log(RGDP_{it}) + \varepsilon_{it}$ (2) *Tax*: Inflation tax, cyclical component of revenues, and Revenues/GDP *OG*: output gap *Taxrate*: VAT, PIT, CIT, Tax index The cyclical series are estimated by the Hodrick-Prescott filter method. (1): Fixed Effects (2): Fixed Effects, Instrumental Variables	β_i is interpreted depending on the fiscal policy variable	62 countries 1960–2013	Tax policy is acyclical in industrial countries but mostly procyclical in developing countries. Better institutional quality (less corruption and more bureaucratic quality) and more financially integration are associated with less procyclical/more countercyclical fiscal policy.
Guerguil et al. (2017)	$\Delta logG_{it} = \alpha_{it} + \delta_{it}*\Delta logG_{it-1} + \beta_{it}*\Delta logY_{it} + \gamma_{it}*X_{it} + \varepsilon_{it}$ (1) Y: real GDP G: public spending (total spending or investment spending) X: control variables (1): Local Gaussian-Weighted OLS	$\beta_{it} < 0$: countercyclicality $\beta_{it} > 0$: procyclicality	167 countries 1990–2012	Total public spending was countercyclical in both fiscal-rule countries and non-fiscal rule countries but the degree of countercyclicality is more pronounced in the former group. In contrast, investment spending was procyclical in both groups and it is more procyclical in the fiscal-rule countries.

Table 2.2 Fiscal behavior of government spending of OECD and non-OECD countries, 1960–2016

Dependent variable: Percentage change of real government spending

VARIABLE	OECD			Non-OECD		
	OLS	**FE**	**FE**	**OLS**	**FE**	**FE**
Percentage change of real GDP	0.537***	0.486***	0.508***	0.714***	0.698***	0.706***
	(0.057)	(0.086)	(0.101)	(0.055)	(0.059)	(0.060)
Constant	0.022***	0.023***	0.046***	0.014***	0.015***	0.041***
	(0.002)	(0.003)	(0.011)	(0.003)	(0.002)	(0.013)
Number of countries		35	35		161	161
Observations	1,692	1,692	1,692	6,368	6,368	6,368
R-squared	0.114	0.088	0.259	0.085	0.076	0.101
Country Fixed Effects		YES	YES		YES	YES
Year Fixed Effects			YES			YES

Note: Ordinary least squares and fixed effects with robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.3 Fiscal behavior of tax rates of OECD and non-OECD countries, 1960–2016

Dependent variable: Tax rate

VARIABLE	VAT				PIT				CIT			
	OECD		**Non-OECD**		**OECD**		**Non-OECD**		**OECD**		**Non-OECD**	
Real GDP growth rate	-0.149***	-0.149***	-0.009	-0.009	0.486**	0.484**	-0.191**	-0.191**	0.315**	0.313***	-0.148**	-0.147**
	(0.045)	(0.044)	(0.033)	(0.033)	(0.204)	(0.202)	(0.087)	(0.087)	(0.121)	(0.120)	(0.065)	(0.065)
Constant	17.294***	16.818***	14.739***	14.268***	48.061***	46.268***	30.831***	30.974***	33.220***	32.213***	32.774***	32.205***
	(0.114)	(1.062)	(0.116)	(0.653)	(0.538)	(1.943)	(0.309)	(2.330)	(0.341)	(1.387)	(0.238)	(1.244)
Number of countries	26	26	42	42	27	27	49	49	27	27	49	49
Observations	926	926	958	958	1,097	1,097	1,661	1,661	1,200	1,200	1,740	1,740
R-squared	0.031	0.028	0.001	0.002	0.012	0.000	0.008	0.001	0.014	0.011	0.009	0.000
Fixed Effect	YES		YES		YES		YES		YES		YES	
Random Effect		YES		YES		YES		YES		YES		YES

Note: Fixed effects and random effects with robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.4 Fiscal behaviour of government spending by income level, 1960–2016

	Dependent variable: Percentage change of real government spending							
	HICs		**UMCs**		**LMCs**		**LICs**	
Percentage change of real GDP	0.517***	0.586***	0.715***	0.725***	0.639***	0.632***	0.877***	0.866***
	(0.079)	(0.080)	(0.055)	(0.064)	(0.156)	(0.159)	(0.141)	(0.147)
Constant	0.023***	0.062***	0.014***	0.021	0.016**	0.042**	0.011**	0.036
	(0.003)	(0.014)	(0.002)	(0.022)	(0.006)	(0.018)	(0.004)	(0.026)
Number of countries	62	62	52	52	52	52	30	30
Observations	2,576	2,576	2,133	2,133	2,063	2,063	1,288	1,288
R-squared	0.078	0.163	0.112	0.164	0.052	0.091	0.077	0.108
Country FE	YES	YES	YES	YES	YES	YES	YES	YES
Year FE		YES		YES		YES		YES

Note: HICs = high-income countries, UMCs = upper-middle-income countries, LMCs = lower-middle-income countries, LICs = lower-income countries. Fixed effects with robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.5 Government-spending cyclicality by region and income

	Mean	SD	Minimum	Maximum
Region				
East Asia and Pacific	0.46	0.72	-0.98	1.84
Europe and Central Asia	0.41	0.55	-1.36	1.47
Latin America and Caribbean	0.77	0.54	-0.13	2.42
Middle East and North Africa	0.69	0.35	0.16	1.36
North America	-0.25	0.36	-0.50	0.01
South Asia	0.35	1.02	-0.67	2.08
Sub-Saharan Africa	0.89	0.93	-2.90	3.44
Level				
High income	0.32	0.53	-1.36	1.56
Low income	0.93	1.13	-2.90	3.44
Lower-middle income	0.78	0.67	-0.98	2.08
Upper-middle income	0.69	0.50	-0.54	2.42
OECD group				
OECD	0.19	0.55	-1.36	1.36
non-OECD	0.74	0.72	-2.90	3.44
Total countries		**170**		
Entire sample	**0.64**	**0.72**	**-2.90**	**3.44**

Note: $\widehat{\beta_{GS}}$ is the estimated coefficient from equation (2.2) using Prais-Winsten approach to measure government-spending cyclicality. Higher $\widehat{\beta_{GS}}$ indicates greater procyclicality (lesser countercyclicality).

Table 2.6 Cross–country regression of government-spending cyclicality using Prais–Winsten estimates, 1960–2016

	Dependent variable: Government-spending cyclicality $\widehat{\beta_{GS}}$												
VARIABLE	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**	**(10)**	**(11)**	**(12)**	**(13)**
polcon	-1.951***	-1.744***	-1.743***	-1.744***	-1.744***	-1.961***	-1.947***	-1.783***	-1.433***	-1.673***	-1.720***	-1.778***	-1.709***
	(0.551)	(0.567)	(0.568)	(0.567)	(0.568)	(0.547)	(0.553)	(0.557)	(0.515)	(0.607)	(0.607)	(0.611)	(0.617)
inf	0.135**	0.120	0.119	0.120	0.119	0.138**	0.127**	0.107	0.115*	0.082	0.092	0.104	0.089
	(0.064)	(0.082)	(0.082)	(0.082)	(0.082)	(0.062)	(0.062)	(0.065)	(0.061)	(0.065)	(0.071)	(0.068)	(0.061)
trade	-0.335***	-0.213*	-0.213*	-0.213*	-0.213*	-0.330***	-0.339***	-0.280**	-0.296**	-0.134	-0.148	-0.170	-0.134
	(0.120)	(0.125)	(0.125)	(0.125)	(0.125)	(0.120)	(0.118)	(0.119)	(0.131)	(0.102)	(0.107)	(0.107)	(0.103)
TAL	-0.002	-0.003**	-0.003**	-0.003**	-0.003**	-0.002*	-0.002	-0.002	-0.003**	-0.002*	-0.003**	-0.003**	-0.002
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
gs	1.070	-0.099	-0.098	-0.096	-0.098	1.091	1.064	1.346	0.935	1.284	0.785	0.609	1.411
	(1.363)	(1.590)	(1.591)	(1.590)	(1.591)	(1.384)	(1.367)	(1.389)	(1.337)	(1.917)	(1.990)	(1.962)	(1.907)
fiscap		0.001***											
		(0.000)											
fiscap_vol			0.001***										
			(0.000)										
lfiscap				0.002***									
				(0.000)									
lfiscap_vol					0.001***								
					(0.000)								
debt						-0.048							
						(0.194)							
debt_vol							0.068						
							(0.220)						
nare								0.499**					
								(0.228)					
manu									-0.804***				
									(0.241)				
CRI										-0.018***			
										(0.005)			
ERI											-0.030**		
											(0.012)		
FRI												-0.024**	
												(0.011)	
PRI													-0.016***
													(0.004)
Constant	1.420***	1.367***	1.368***	1.366***	1.368***	1.445***	1.404***	1.064***	1.448***	2.380***	2.254***	2.139***	2.188***
	(0.327)	(0.324)	(0.324)	(0.324)	(0.324)	(0.318)	(0.331)	(0.343)	(0.326)	(0.527)	(0.547)	(0.594)	(0.472)
Number of countries	144	94	94	94	94	144	144	143	143	117	117	117	117
R-squared	0.133	0.178	0.177	0.178	0.177	0.133	0.134	0.162	0.190	0.190	0.167	0.161	0.196
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

(continued).

VARIABLE	(14)	(15)	(16)	(17)	(18)	(19)	(20)	(21)	(22)	(23)	(24)	(25)
polcon	-2.086***	-1.627***	-1.532**	-1.890***	-1.898***	-1.610**	-1.823***	-1.985***	-1.785***	-1.994***	-1.669**	-1.596***
	(0.653)	(0.600)	(0.616)	(0.643)	(0.645)	(0.618)	(0.617)	(0.635)	(0.653)	(0.624)	(0.646)	(0.606)
inf	0.122*	0.081	0.068	0.112*	0.127*	0.085	0.106	0.141**	0.095	0.121**	0.114*	0.093
	(0.066)	(0.064)	(0.065)	(0.064)	(0.065)	(0.053)	(0.064)	(0.065)	(0.061)	(0.057)	(0.063)	(0.066)
trade	-0.162	-0.133	-0.114	-0.126	-0.179	-0.209**	-0.132	-0.195*	-0.170	-0.163	-0.223**	-0.185*
	(0.114)	(0.099)	(0.102)	(0.121)	(0.122)	(0.101)	(0.114)	(0.117)	(0.104)	(0.112)	(0.111)	(0.099)
TAL	-0.002	-0.001	-0.002*	-0.003**	-0.003**	-0.000	-0.003**	-0.003**	-0.002	-0.003**	-0.002**	-0.002*
	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
gs	0.644	1.038	1.229	0.725	0.215	1.830	1.431	0.390	1.766	0.378	0.802	1.364
	(2.046)	(1.957)	(1.897)	(1.962)	(2.001)	(2.052)	(1.943)	(1.924)	(2.035)	(1.966)	(1.887)	(2.020)
govstab	-0.116											
	(0.083)											
socecon		-0.118***										
		(0.028)										
invest			-0.155***									
			(0.043)									
inconflict				-0.076**								
				(0.038)								
exconflict					-0.044							
					(0.041)							
corrupt						-0.219***						
						(0.046)						
military							-0.110***					
							(0.041)					
religious								-0.050				
								(0.055)				
law									-0.156***			
									(0.044)			
ethnic										-0.124**		
										(0.060)		
democracy											-0.088	
											(0.055)	
bureau												-0.184***
												(0.044)
Constant	2.292***	1.842***	2.255***	1.982***	1.861***	1.761***	1.602***	1.670***	1.720***	1.917***	1.640***	1.556***
	(0.706)	(0.397)	(0.505)	(0.476)	(0.563)	(0.353)	(0.397)	(0.491)	(0.371)	(0.491)	(0.433)	(0.366)
Number of countries	117	117	117	117	117	117	117	117	117	117	117	117
R-squared	0.151	0.204	0.205	0.164	0.146	0.223	0.182	0.146	0.189	0.177	0.156	0.191
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

Note: OLS specification with robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 2.7 Cross-country regressions of government-spending cyclicality, 1960-2016: robustness checks

Dependent variable: Government-spending cyclicality $\widehat{\beta_{GS}}$								
Main variables	**Model (1)**	**Model (2)**	**Model (3)**	**Model (4)**	**Model (5)**	**Model (6)**	**Model (7)**	**Model (8)**
fiscap	0.002***	0.001***	0.001***	0.001***	0.006***	0.136*	0.027***	0.166**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.073)	(0.002)	(0.069)
fiscap_vol	0.001***	0.001**	0.001***	0.001***	0.005***	0.073	0.020***	0.095*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.051)	(0.002)	(0.051)
lfiscap	0.002***	0.001***	0.001***	0.001***	0.007***	0.146**	0.031***	0.172***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.062)	(0.002)	(0.058)
lfiscap_vol	0.001***	0.001**	0.001***	0.001***	0.005***	0.072**	0.022***	0.089**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.034)	(0.002)	(0.037)
debt	-0.038	-0.079	-0.016	-0.013	0.617	0.226	0.489	0.448
	(0.146)	(0.187)	(0.128)	(0.124)	(0.450)	(0.356)	(0.535)	(0.313)
debt_vol	0.074	0.050	0.070	0.074	0.241	0.406	1.051	0.641**
	(0.197)	(0.200)	(0.175)	(0.174)	(0.594)	(0.355)	(0.821)	(0.293)
nare	0.604***	0.395*	0.544***	0.544***	0.602	0.394	1.132**	0.581
	(0.167)	(0.229)	(0.153)	(0.154)	(0.609)	(0.461)	(0.536)	(0.406)
manu	-0.676***	-0.571**	-0.562***	-0.548***	-1.449***	-0.493	-1.477**	-0.623
	(0.187)	(0.230)	(0.150)	(0.152)	(0.493)	(0.453)	(0.635)	(0.415)
CRI	-0.021***	-0.012**	-0.016***	-0.016***	-0.040***	-0.028**	-0.074***	-0.032***
	(0.004)	(0.005)	(0.004)	(0.004)	(0.014)	(0.014)	(0.020)	(0.011)
ERI	-0.036***	-0.019*	-0.023**	-0.021**	-0.079**	-0.062*	-0.122**	-0.063**
	(0.010)	(0.011)	(0.009)	(0.009)	(0.036)	(0.033)	(0.053)	(0.026)
FRI	-0.030***	-0.012	-0.019**	-0.019**	-0.077***	-0.038	-0.107***	-0.048**
	(0.008)	(0.011)	(0.008)	(0.008)	(0.027)	(0.025)	(0.038)	(0.021)
PRI	-0.018***	-0.011**	-0.015***	-0.015***	-0.029**	-0.022*	-0.055***	-0.024***
	(0.004)	(0.004)	(0.003)	(0.003)	(0.013)	(0.011)	(0.014)	(0.009)
govstab	-0.145**	-0.087	-0.125*	-0.119*	-0.331	-0.274	-0.678**	-0.398**
	(0.069)	(0.077)	(0.064)	(0.064)	(0.218)	(0.223)	(0.314)	(0.189)
socecon	-0.123***	-0.086***	-0.091***	-0.089***	-0.193**	-0.128*	-0.345***	-0.192***
	(0.023)	(0.026)	(0.022)	(0.022)	(0.085)	(0.075)	(0.118)	(0.062)
invest	-0.166***	-0.111***	-0.130***	-0.128***	-0.267**	-0.275***	-0.401**	-0.300***
	(0.031)	(0.041)	(0.030)	(0.031)	(0.110)	(0.087)	(0.157)	(0.074)
inconflict	-0.116***	-0.048	-0.094***	-0.096***	-0.121	-0.126	-0.297***	-0.118*
	(0.027)	(0.035)	(0.024)	(0.024)	(0.088)	(0.081)	(0.108)	(0.063)
exconflict	-0.060*	-0.019	-0.052*	-0.053*	-0.016	-0.142	-0.356**	-0.114
	(0.032)	(0.038)	(0.029)	(0.030)	(0.114)	(0.096)	(0.147)	(0.077)
corrupt	-0.186***	-0.164***	-0.144***	-0.143***	-0.287**	-0.211*	-0.576***	-0.216**
	(0.038)	(0.043)	(0.035)	(0.035)	(0.133)	(0.116)	(0.184)	(0.091)
military	-0.131***	-0.071*	-0.102***	-0.103***	-0.243**	-0.127	-0.368***	-0.169**
	(0.028)	(0.040)	(0.029)	(0.029)	(0.105)	(0.084)	(0.116)	(0.067)
religious	-0.088**	-0.019	-0.079**	-0.085**	-0.218	-0.107	-0.081	-0.084
	(0.040)	(0.053)	(0.036)	(0.037)	(0.148)	(0.107)	(0.124)	(0.083)
law	-0.181***	-0.114***	-0.137***	-0.135***	-0.298**	-0.239**	-0.537***	-0.240**
	(0.036)	(0.040)	(0.033)	(0.034)	(0.127)	(0.111)	(0.167)	(0.092)
ethnic	-0.133***	-0.087	-0.120***	-0.125***	-0.013	-0.138	-0.276**	-0.145*
	(0.041)	(0.057)	(0.037)	(0.038)	(0.114)	(0.109)	(0.131)	(0.087)
democracy	-0.124***	-0.040	-0.103***	-0.101***	-0.238	-0.055	-0.277**	-0.140
	(0.036)	(0.054)	(0.035)	(0.035)	(0.184)	(0.109)	(0.126)	(0.093)
bureau	-0.193***	-0.123***	-0.157***	-0.153***	-0.383**	-0.197	-0.389**	-0.280***
	(0.040)	(0.043)	(0.040)	(0.041)	(0.164)	(0.127)	(0.181)	(0.099)

Note: Model (1): 1st step by Prais-Winsten estimation, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step).
Model (2): 1st step and 2nd step by OLS estimation with robust standard errors.
Model (3): 1st step by OLS estimation with Newey-West standard errors to correct heteroscedasticity and AR(1) of the residuals, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step).

Model (4): 1st step by OLS estimation with Newey-West standard errors to correct heteroscedasticity and AR(2) of the residuals, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step).
Model (5): 1st step by Two-Stage Least Squares estimation (excluded instrument is *lag.SHOCKGL*), 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step). In the 1st step, there are 48/135 countries having F-partial statistics (1st stage) >= 2; and 122/135 countries having p-value (Durbin test for endogeneity) > 5%.
Model (6): 1st step by Two-Stage Least Squares estimation (excluded instruments are *lag6.SHOCKGL and lag5.SHOCKJP*), 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step). In the 1st step, there are 33/124 countries having F-partial statistics (1st stage) >= 2; 117/124 countries having p_value (Durbin test for endogeneity) > 5%; and 122/124 countries having p-value (Sargan test for overidentification) > 5%.
Model (7): 1st step by Two-Stage Least Squares estimation (excluded instrument is *lag.KAUS*), 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step). In the 1st step, there are 40/135 countries having F-partial statistics (1st stage) >= 2; and 119/135 countries having p-value (Durbin test for endogeneity) > 5%.
Model (8): 1st step by Two-Stage Least Squares estimation (excluded instruments are *lag.KAUS, lag6.SHOCKGL, and lag5.SHOCKJP*), 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step). In the 1st step, there are 51/130 countries having F-partial statistics (1st stage) >= 2; 122/130 countries having p-value (Durbin test for endogeneity) > 5%; and 124/130 countries having p-value (Sargan test for overidentification) > 5%.
The same set of control variables (*polcon, inf, trade, TAL, gs*) is used in each cross-country regression. Their estimators are not shown in the table.
Robust standard errors are in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

Table 2.8 Cross-country regressions on government-spending cyclicality at good times versus bad times, 1960-2016

Dependent variable: Government-spending cyclicality $\widehat{\beta_{GS}}$			
	Good times		**Bad times**
Main variables	**Model (1)**	**Model (2)**	**Model (3)**
fiscap	0.001***	0.001*	-0.021
	(0.000)	(0.000)	(0.017)
fiscap_vol	0.001***	0.001*	-0.090**
	(0.000)	(0.000)	(0.044)
lfiscap	0.002***	0.001*	-0.023
	(0.000)	(0.000)	(0.020)
lfiscap_vol	0.001***	0.001*	-0.242*
	(0.000)	(0.000)	(0.133)
debt	0.037	0.014	-0.426
	(0.146)	(0.142)	(0.431)
debt_vol	0.292**	0.243**	-0.710
	(0.115)	(0.103)	(0.579)
nare	0.373*	0.226	-1.572
	(0.223)	(0.206)	(1.337)
manu	-0.733***	-0.468**	1.467
	(0.253)	(0.227)	(1.404)
CRI	-0.018***	-0.011	-0.027
	(0.007)	(0.007)	(0.032)
ERI	-0.047***	-0.032**	-0.093
	(0.015)	(0.014)	(0.100)
FRI	-0.034***	-0.021*	-0.051
	(0.012)	(0.012)	(0.072)
PRI	-0.012**	-0.006	-0.019
	(0.006)	(0.006)	(0.027)
govstab	-0.075	-0.032	-0.565
	(0.104)	(0.101)	(0.490)
socecon	-0.099***	-0.063**	-0.190
	(0.032)	(0.030)	(0.223)
invest	-0.105*	-0.051	-0.602*
	(0.062)	(0.062)	(0.349)
inconflict	-0.071	-0.036	0.124
	(0.053)	(0.052)	(0.180)
exconflict	-0.086	-0.058	0.080
	(0.062)	(0.058)	(0.257)
corrupt	-0.118*	-0.067	0.219
	(0.063)	(0.061)	(0.404)
military	-0.045	-0.020	0.001
	(0.052)	(0.050)	(0.225)
religious	-0.062	-0.031	-1.286
	(0.056)	(0.056)	(1.189)
law	-0.095	-0.045	-0.096
	(0.068)	(0.064)	(0.235)
ethnic	-0.078	-0.035	0.169
	(0.058)	(0.060)	(0.405)
democracy	-0.047	0.003	-0.493
	(0.066)	(0.067)	(0.508)
bureau	-0.094	-0.033	-0.041
	(0.071)	(0.067)	(0.267)

Note: Model (1): 1st step by Prais-Winsten estimation, 2nd step by OLS estimation with robust standard errors. Model (2) and model (3): 1st step and 2nd step by OLS estimation with robust standard errors.
The same set of control variables (*polcon*, *inf*, *trade*, *TAL*, *gs*) is used in each cross-country regression. Their estimators are not shown in the table.
Robust standard errors are in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

Table 2.9 Sovereign wealth funds and government-spending cyclicality

	Dependent variable: Government-spending cyclicality $\widehat{\beta_{GS}}$			
VARIABLES	**Full sample**		**Good times**	
	(1)	**(2)**	**(3)**	**(4)**
polcon	-0.817	-0.826	-0.311	-0.327
	(0.751)	(0.748)	(0.671)	(0.669)
inf	-0.091	-0.090	-0.041	-0.040
	(0.080)	(0.080)	(0.060)	(0.060)
trade	0.058	0.066	0.033	0.043
	(0.132)	(0.131)	(0.146)	(0.146)
TAL	-0.004	-0.004	-0.001	-0.000
	(0.003)	(0.003)	(0.003)	(0.003)
fiscap	0.121***		0.103*	
	(0.041)		(0.054)	
SWF*fiscap	-0.125***		-0.105*	
	(0.041)		(0.054)	
lfiscap		0.119***		0.103**
		(0.038)		(0.049)
SWF*lfiscap		-0.125***		-0.106**
		(0.038)		1.592**
CRI	-0.019**	-0.019**	-0.016*	-0.017*
	(0.008)	(0.008)	(0.009)	(0.009)
SWF*CRI	-0.046***	-0.046***	-0.024*	-0.024*
	(0.011)	(0.011)	(0.013)	(0.013)
Constant	1.778***	1.810***	1.560**	1.592**
	(0.660)	(0.653)	(0.675)	(0.669)
Number of countries	81	81	80	80
R-squared	0.584	0.586	0.367	0.372
p-value	0.000	0.000	0.000	0.000

Note: WLS specification, the weight is real GDP (2010 US$) by country averaged over the full period in full sample, over good times in good-times sub-sample.
Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

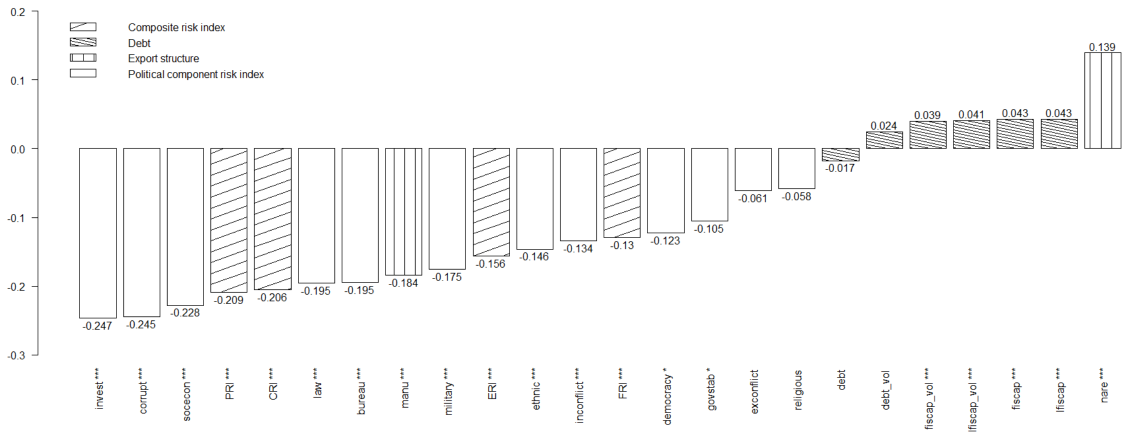
Table 2.10 Cross-country regressions on government-spending cyclicality, 1980-2016

	Dependent variable: Government-spending cyclicality $\widehat{\beta_{GS}}$					
Main variables	**Model (1)**	**Model (2)**	**Model (3)**	**Model (4)**	**Model (5)**	**Model (6)**
fiscap	0.068	0.136	0.089	0.044	0.042	0.187**
	(0.071)	(0.088)	(0.060)	(0.058)	(0.058)	(0.084)
fiscap_vol	0.110	0.292*	0.144	0.106	0.112	0.168
	(0.119)	(0.161)	(0.102)	(0.107)	(0.109)	(0.136)
lfiscap	0.079	0.141	0.099	0.047	0.044	0.210**
	(0.074)	(0.087)	(0.063)	(0.059)	(0.060)	(0.081)
lfiscap_vol	0.126	0.312*	0.159	0.120	0.126	0.183
	(0.125)	(0.157)	(0.107)	(0.112)	(0.114)	(0.140)
debt	0.707**	0.684***	0.747**	0.743***	0.715***	1.376***
	(0.288)	(0.237)	(0.303)	(0.265)	(0.253)	(0.384)
debt_vol	1.411**	1.564***	1.587***	1.490**	1.475**	2.957***
	(0.557)	(0.565)	(0.534)	(0.620)	(0.595)	(0.884)
nare	0.648	0.597*	0.698*	0.424	0.425	-0.724
	(0.416)	(0.344)	(0.415)	(0.338)	(0.333)	(0.611)
manu	-1.342***	-1.090***	-1.370***	-0.815**	-0.841**	-1.215*
	(0.380)	(0.340)	(0.340)	(0.330)	(0.326)	(0.624)
CRI	-0.030*	-0.033***	-0.030*	-0.031***	-0.031***	-0.072***
	(0.016)	(0.010)	(0.016)	(0.010)	(0.009)	(0.019)
ERI	-0.071*	-0.085***	-0.066*	-0.080***	-0.079***	-0.157***

	(0.037)	(0.026)	(0.037)	(0.026)	(0.025)	(0.046)
FRI	-0.055	-0.066***	-0.055	-0.069***	-0.068***	-0.121***
	(0.034)	(0.021)	(0.034)	(0.021)	(0.020)	(0.041)
PRI	-0.016	-0.016*	-0.017	-0.015*	-0.015**	-0.047***
	(0.012)	(0.009)	(0.012)	(0.008)	(0.007)	(0.012)
govstab	-0.443**	-0.343***	-0.358**	-0.320**	-0.316**	-0.703***
	(0.173)	(0.124)	(0.172)	(0.130)	(0.120)	(0.235)
socecon	-0.209***	-0.224***	-0.220***	-0.247***	-0.248***	-0.402***
	(0.072)	(0.055)	(0.068)	(0.054)	(0.052)	(0.117)
invest	-0.159	-0.221***	-0.146	-0.212***	-0.215***	-0.483***
	(0.114)	(0.059)	(0.115)	(0.058)	(0.055)	(0.100)
inconflict	-0.050	-0.045	-0.052	-0.043	-0.042	-0.244**
	(0.065)	(0.054)	(0.067)	(0.045)	(0.045)	(0.099)
exconflict	0.067	0.021	0.064	-0.007	-0.011	-0.213*
	(0.104)	(0.070)	(0.101)	(0.053)	(0.052)	(0.108)
corrupt	-0.293**	-0.141	-0.297**	-0.125	-0.120	-0.227
	(0.144)	(0.090)	(0.146)	(0.076)	(0.074)	(0.143)
military	-0.039	-0.093	-0.039	-0.116*	-0.120*	-0.413***
	(0.096)	(0.067)	(0.097)	(0.066)	(0.063)	(0.125)
religious	0.069	-0.004	0.047	-0.028	-0.033	-0.204*
	(0.113)	(0.065)	(0.119)	(0.057)	(0.058)	(0.117)
law	-0.285**	-0.225**	-0.261**	-0.200**	-0.195**	-0.356**
	(0.114)	(0.094)	(0.112)	(0.097)	(0.092)	(0.151)
ethnic	-0.062	-0.024	-0.086	-0.087	-0.091	-0.253*
	(0.102)	(0.068)	(0.107)	(0.068)	(0.069)	(0.143)
democracy	0.087	0.022	0.046	-0.044	-0.047	-0.306
	(0.187)	(0.092)	(0.194)	(0.084)	(0.080)	(0.187)
bureau	-0.179	-0.172*	-0.212*	-0.190**	-0.192**	-0.594***
	(0.127)	(0.100)	(0.123)	(0.094)	(0.091)	(0.167)

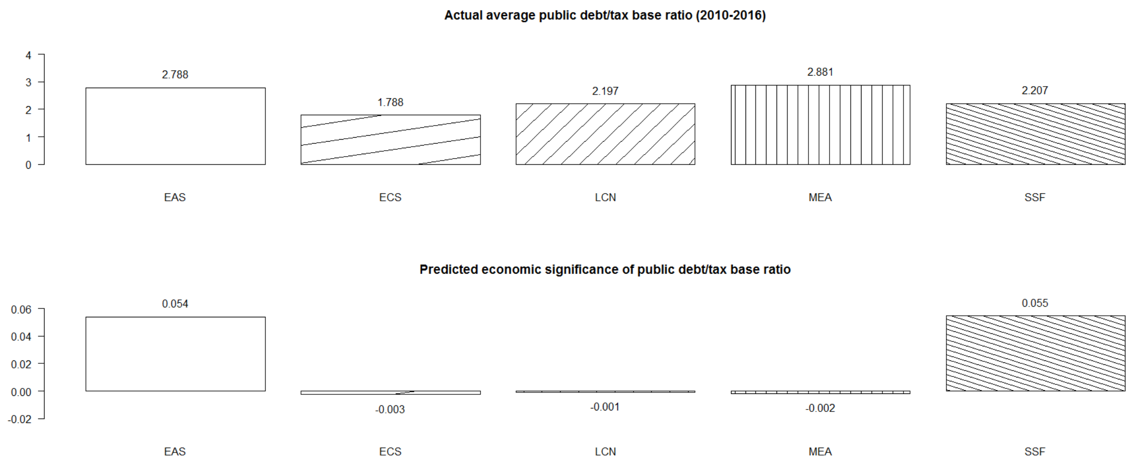
Note: Model (1): 1st step by Prais-Winsten estimation, 2nd step by OLS with robust standard errors.
Model (2): 1st step by Prais-Winsten estimation, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step).
Model (3): 1st step and 2nd step by OLS estimation with robust standard errors.
Model (4): 1st step by OLS estimation with Newey-West standard errors to correct heteroscedasticity and AR(1) of the residuals, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step).
Model (5): 1st step by OLS estimation with Newey-West standard errors to correct heteroscedasticity and AR(2) of the residuals, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step).
Model (6): 1st step by Two-Stage Least Squares estimation (excluded instruments are *lag5.SHOCKJP, lag.KAUS)*, 2nd step by Weighted Least Squares estimation (weight is the inverse of standard errors of $\hat{\beta}$ in the 1st step). In the 1st step, there are 37/84 countries having F-partial statistics (1st stage) >= 2; 76/84 countries having p (Durbin test for endogeneity) > 5%; and 81/84 countries having p-value (Sargan test for overidentification) > 5%.
The same set of control variables (*polcon, inf, trade, TAL, gs*) is used in each cross-country regression. Their estimators are not shown in the table.
Robust standard errors are in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

Figure 2.1 Economic significance of variables to government-spending cyclicality, 1960–2016



Note: $\widehat{\beta_{GS}}$ by country is estimated from equation (2.2) using Prais-Winsten approach. *** p<0.05, ** p<0.01, * p<0.2.

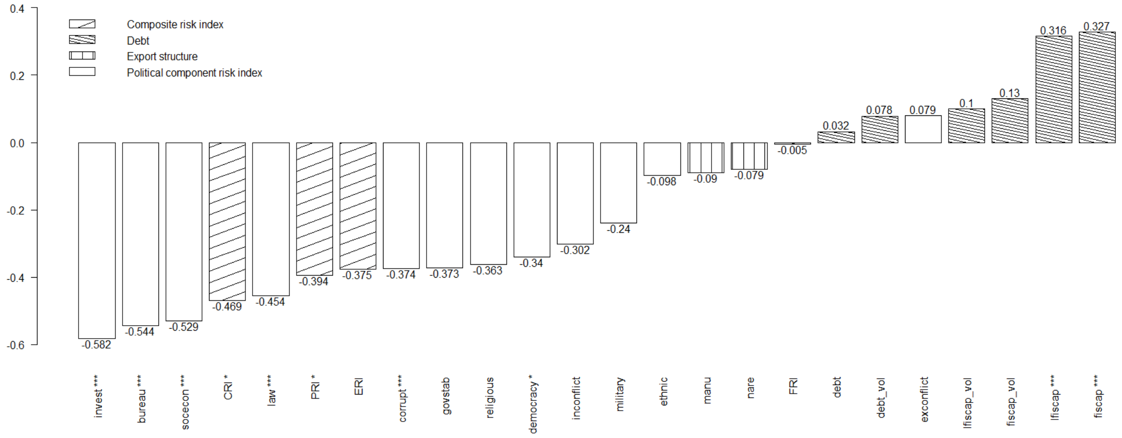
Figure 2.2 Economic significance of public debt/tax base to government-spending cyclicality by region



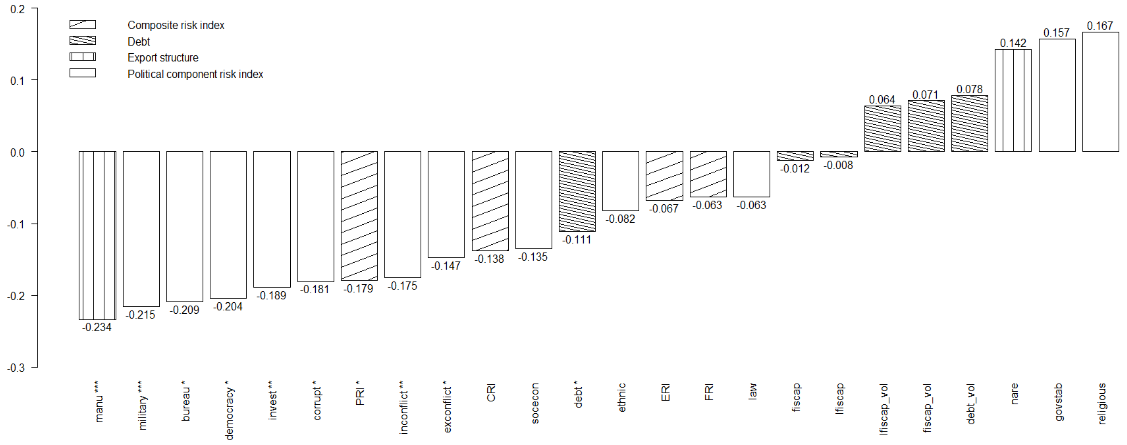
Note: $\widehat{\beta_{GS}}$ by country is estimated from equation (2.2) using Prais-Winsten approach. EAS: East Asia & Pacific; ECS: Europe & Central Asia; LCN: Latin America & Caribbean; MEA: Middle East & North Africa; SSF: Sub-Saharan Africa. The countries are grouped according to World Bank regions.

Figure 2.3 Economic significance of variables to government-spending cyclicality by region

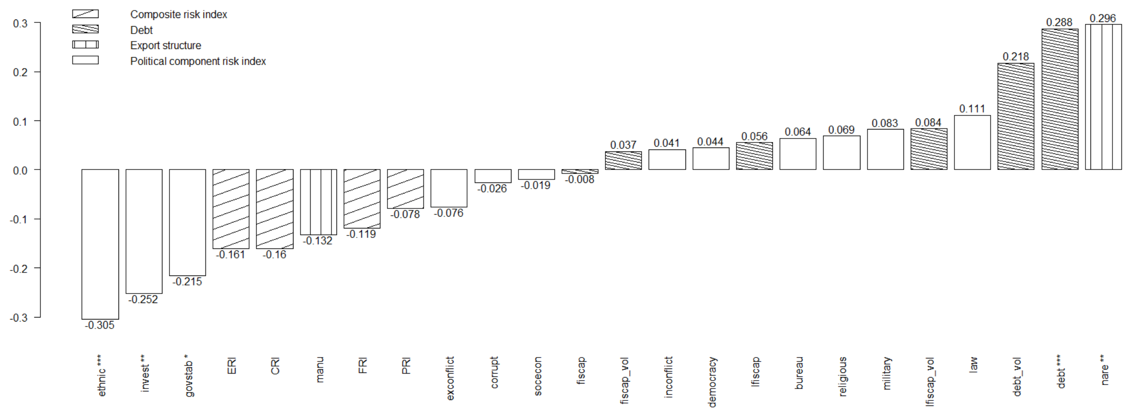
a. East Asia and Pacific



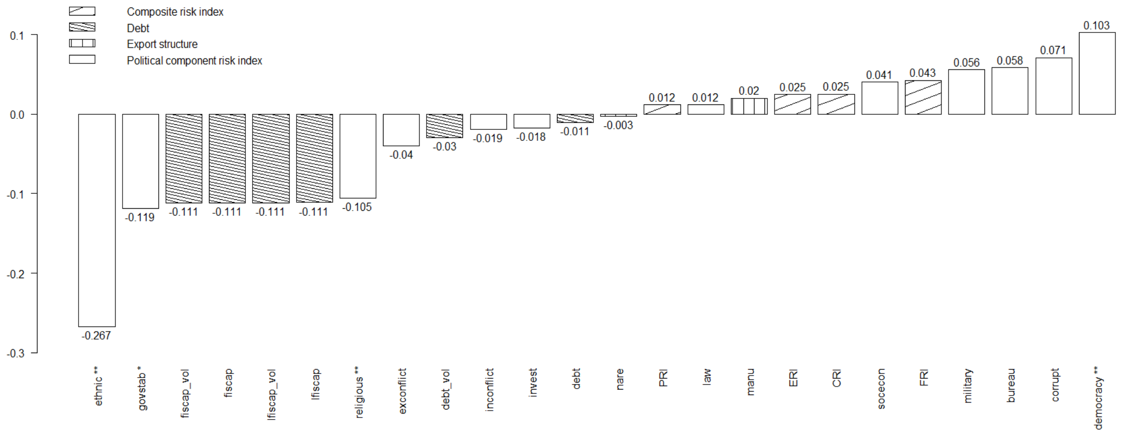
b. Europe and Central Asia



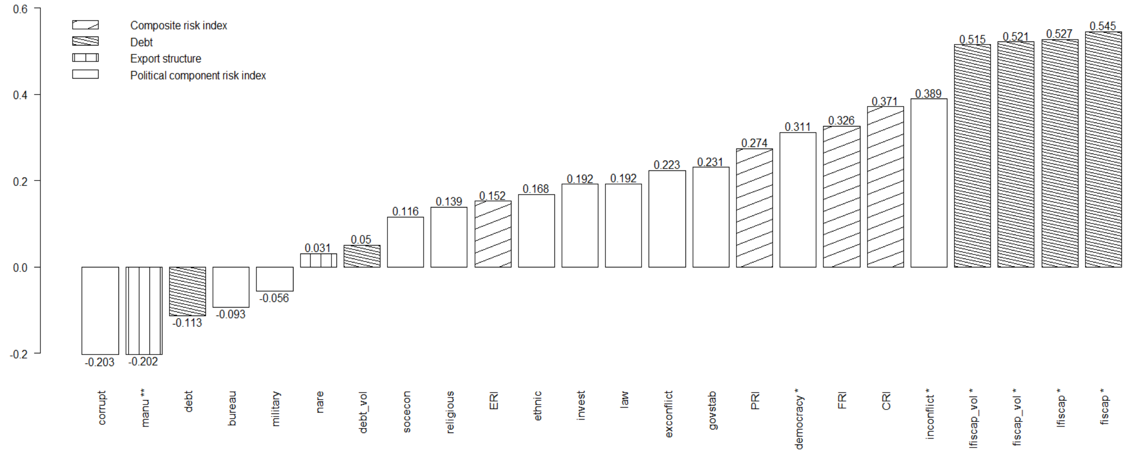
c. Latin America and Caribbean



d. Middle East and North Africa

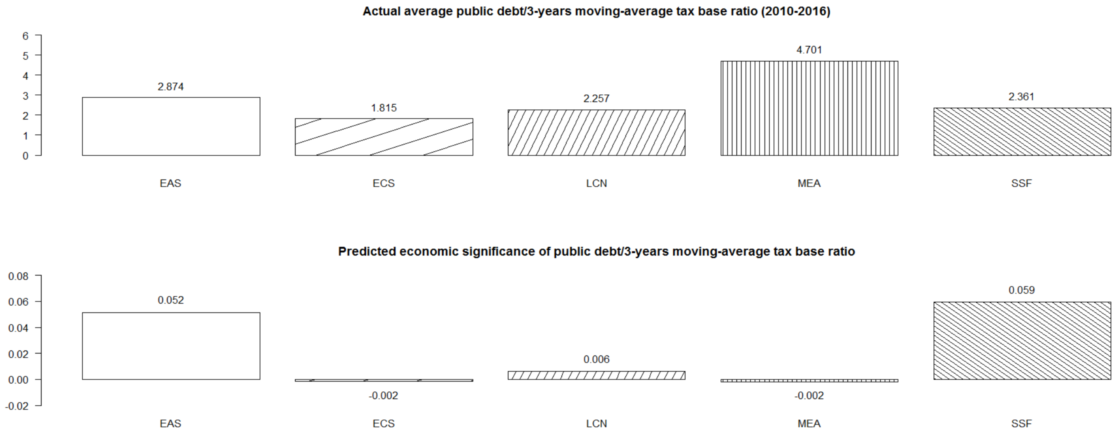


e. Sub-Saharan Africa



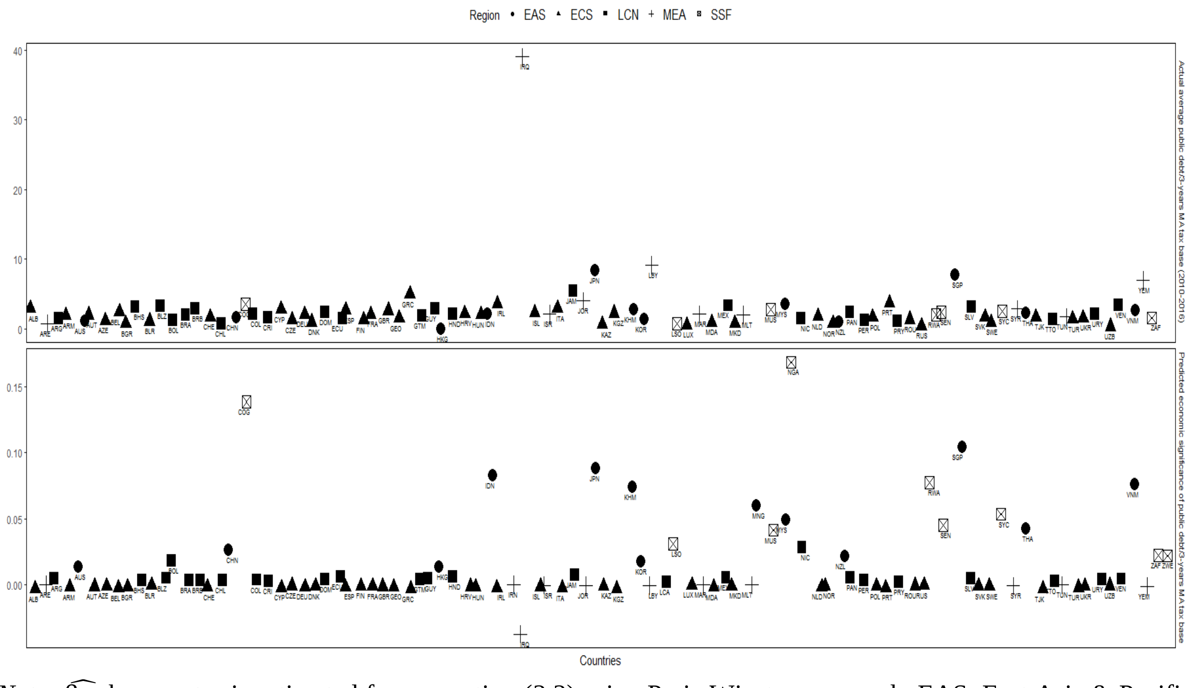
Note: $\widehat{\beta_{GS}}$ by country is estimated from equation (2.2) using Prais-Winsten approach. The economic significance of each explanatory variable in each region is calculated by multiplying its corresponding standard deviation with its estimated coefficient from cross-country regression for that region (similar to equation (2.3)) to approximate the effect of its one standard deviation increase on the fiscal cyclicality. The countries are grouped according to World Bank regions. *** p<0.05, ** p<0.01, * p<0.2.

Figure 2.4 Economic significance of public debt/3-years moving-average tax base to government-spending cyclicality by region



Note: $\widehat{\beta_{GS}}$ by country is estimated from equation (2.2) using Prais-Winsten approach. EAS: East Asia and Pacific; ECS: Europe and Central Asia; LCN: Latin America and Caribbean; MEA: Middle East and North Africa; SSF: Sub-Saharan Africa. The countries are grouped according to World Bank regions.

Figure 2.5 Economic significance of public debt/3-years moving-average tax base to government-spending cyclicality by country



Note: $\widehat{\beta_{GS}}$ by country is estimated from equation (2.2) using Prais-Winsten approach. EAS: East Asia & Pacific; ECS: Europe & Central Asia; LCN: Latin America & Caribbean; MEA: Middle East & North Africa; SSF: Sub-Saharan Africa. The countries are grouped according to World Bank regions. The upper half shows the public debt/(3-year tax base) ratio. The lower half shows the economic impact of deteriorating fiscal space, i.e. if a fiscal capacity drops by 10%, what would happen to the government-spending cyclicality. Specifically, I calculate: 0.1*(Actual average country-specific public debt/(3-year tax base))*(Regional-specific estimated coefficient of public debt/(3-year tax base)). I use regional-specific coefficient in place of country-specific coefficient as there is insufficient country-level data to estimate a country-specific 2nd-step regression (i.e., equation (2.3); $\widehat{\beta_{GS}}$ = f(public debt/(3-year tax base), control variables)).

Chapter 3 Global Commodity-Price Shocks and Inflation Targeting in Emerging and Developing Countries

3.1 Introduction

This chapter analyses whether inflation targeting (IT) makes a difference in the output and inflation responses for emerging and developing (EME) countries in the presence of global commodity-price shocks. Trade globalization has amplified the impact of the global commodity cycles on EME economies, deepening their exposure to the external macroeconomic conditions. While a considerable body of literature focuses on the impacts of fuel-price shock in the developed economies, the studies for less developed countries and other commodity-price shocks have been rather scant.[20] This emphasizes the need of research on the effects of commodity price shocks on a wide set of EME countries, which in turn will help guide policy, mitigate the political business cycles, and avoid inconsistent policy discretions.[21] Also, given the exposure to global commodity-price shocks, do the IT emerging countries perform better than their non-IT counterparts do? Although the supportive evidence to IT are quite popular in the literature –i.e., IT helps anchor inflation and reduce inflation volatility without lowering output – the inadequate controlling for the endogeneity of monetary policy in those studies leaves those findings ambiguous. In this context, my study contributes to the literature, not only in revealing the impacts of global fuel, agriculture, and metal prices on a larger sample of less developed economies but also using the structural vector autoregression (SVAR) approach to better control for the endogeneity of multiple macroeconomic variables, and thus providing more robust results.

First, I estimate country-specific SVARs using a global commodity price (agriculture, fuel, or metal price) and a domestic block of five macroeconomic variables including commodity terms-of-trade, GDP growth, inflation, interest rate, and real exchange rate. Then I compare the impulse responses of the domestic variables to the global commodity-price shocks between the IT and the non-IT countries. The sample includes 99 EME countries covering the 1990Q1-

[20] Some example papers on oil shock and developed economies include Hamilton (1983), Cuñado and de Gracia (2003), Jiménez-Rodríguez and Sánchez (2005), Kilian (2009), and Gospodinov and Ng (2013).

[21] Drechsel and Tenreyro (2018) propose a two-sector model for a small net commodity exporter, showing two effects of commodity-price shocks on business cycle: the competitiveness effect and the borrowing cost effect. They also conduct a quantitative analysis for Argentina from 1900 to 2015 and find that commodity-price shocks contribute 38% of the fluctuations of the post-1950 output growth. They argue that the exogenous shocks from international commodity prices are easier to identify and act upon by policymakers, compared to other shocks such as domestic TFP shocks.

2016Q4 period. In the SVAR estimations for small open EME economies, I assume the global commodity-price shocks to be exogenous, i.e., the shocks affect domestic variables contemporaneously but not vice versa, and thus I focus on a smaller group excluding the largest commodity exporters and importers. The sampled countries whose economic performance may affect the global commodity prices include China (top importer of agriculture, fuel, and metal), India (big fuel importer), Russia (big exporter of fuel and metal), Saudi Arabia (top fuel exporter), Venezuela (large fuel exporter), Brazil (top agriculture exporter), and Chile (top metal exporter). The findings are supported by the results for the full sample, a cross-check using interannual data, and an examination of asymmetric effects of positive vis-à-vis negative shocks.

The analysis of symmetric price shocks shows evidence of the transitory inflationary responses, which are quite uniform across the IT and the non-IT countries. However, only the IT group displays long-lasting increases in GDP growth following the shocks. After one year, GDP growth improves by 1.1% in response to a 10% increase in agriculture price, 0.3% to the fuel-price shock, and 0.62% to the metal-price shock. The accumulated increases in GDP growth of the IT countries persist at least for 6 quarters after the shocks. The variance decomposition shows that the shocks account for a bigger variation share of output growth in the IT group (18% - 20%) than in the non-IT group (9% - 13%). Compared to the agriculture-price and metal-price shocks, the fuel-price shock has the largest impact on the variations of inflation in both groups. For the whole sample, the shocks play a modest role in explaining the variations of GDP growth (10% - 14%) and inflation (11% - 16%).

I also find evidence of asymmetric effects, with the domestic variables being more responsive to the negative shocks than the positive shocks. Most of the responses to the positive shocks are negligible. The analysis of asymmetric shocks strengthens my findings that the IT countries are more resilient to the negative price shocks than the non-IT group. More specifically, the negative shocks, i.e., declines in the global commodity prices cause the GDP growth of the IT countries to increase persistently while leaving that of the non-IT countries unchanged. The inflationary impacts are also temporary, similarly for both IT and non-IT groups.

3.2 Literature review

Firstly, this chapter is related to several theoretical and empirical studies that account for the role of global commodity-price shocks in the output variations of small open economies. Mendoza (1995) suggests that the shocks can propagate through many channels, including international capital mobility, the cost of imported inputs, and the overall purchasing power of export.[22] Later on, Kose (2002) extends Mendoza's work and demonstrates three transmission channels for terms-of-trade disturbances, including primary goods, imported capital goods, and intermediate inputs. A common finding stands out in the literature: the commodity price increase is positively associated with output expansion of commodity net exporters but negatively associated with that of commodity net importers. Using data of 28 non-oil exporting developing economies and the G7 countries from 1970 to 1992, Kose (2002) finds that the terms-of-trade disturbances positively affect the output of commodity-exporting countries but negatively impact commodity-importing countries. Similarly, Berument, Ceylan, and Dogan (2010) use a sample of 16 selected countries in the Middle East and North African region (time periods varying across countries) to show that an oil price increase positively affects the output of the oil net exporters but an insignificant effect on the oil net importers.[23] In a theoretical model for a commodity net exporter specifically, Drechsel and Tenreyro (2018) point out two effects linking commodity-price shocks to output fluctuations. First, an increasing commodity price leads to higher commodity trade revenues, significantly pushing up the value-added of the commodity sector, which is strong enough to stunt the suffering impact on the final good sector, and thus increasing total production of the economy (competitiveness effect). Second, a higher commodity price lowers the interest spread between the borrowing country and the world interest rates, bringing more consumption, investment, and hence a hump-shaped increase in the GDP of both commodity sector and final good sector and thus total economy (borrowing cost effect).[24] However, existing evidence on the contribution of global commodity-price and terms-of-trade shocks on output variations differ remarkably across studies, depending upon the employed measurements and assumptions that allow more or fewer disturbances transmitted, amongst other reasons. While Mendoza (1995) shows that terms-of-trade shocks explain 56% of output fluctuations, Kose (2002) suggests a larger number of roughly 88%. Through the lens of SVAR estimations across 38 poor and emerging countries

[22] See Mendoza (1995) for an empirical examination of the relationship between terms-of-trade shocks and business cycles using a three-sector intertemporal equilibrium model and a large panel dataset of G7 countries from 1955 to 1990 and 23 developing countries from 1960 to 1990.
[23] According to Berument et al. (2010), a positive oil price shock increases production cost for oil-importing countries and, therefore, decreases output. Disposable income, consumption, and private investment also decrease along with the fall in capital and labor productivity and the potential output. By contrast, a similar shock for oil-exporting countries increases the energy sector's profit, the countries' currency value, real national income and stimulates investment due to local currency appreciation as well as demand for labor and capital. See Berument et al. (2010) for several related empirical studies in the field.
[24] Higher export earnings may bring a better repayment capacity to the countries depending directly on commodity prices and hence higher collateral value of the economy, which makes creditors lower the required interest rate premium (Drechsel & Tenreyro, 2018).

from 1980 to 2011, Schmitt-Grohé and Uribe (2018) find a smaller role of terms-of-trade shocks, less than 10%, in explaining the output disturbances. This empirical result is defended in an extended calibration for a three-sector model with the macroeconomic variables measured in the same units as in the data on average across countries. Recently, Fernández, Schmitt-Grohé, and Uribe (2017) use a block of three global commodity prices (including agriculture, fuel, and metal prices) and find that 33% of business cycle variations across 138 countries over 1960-2015 are due to the shocks.

Previous studies find that global commodity-price shocks can also propagate into consumer prices. The first-round effect is on the domestic commodity prices. The second-round effect is on production costs or price-setting of firms. The standard sticky-price model predicts a persistent response of consumer prices to all shocks because firms adjust prices slowly. However, in the rational-inattention models, firms pay more attention to volatile shocks, which are more likely detrimental to their profits, leading to fast responses of inflation.[25] Using a sample of 144 countries covering the 2000s, Sekine and Tsuruga (2018) find that commodity-price shocks cause a temporary first-round effect and a weak transitory second-round effect on headline inflation at least in countries with exchange rate flexibility (47% of their sample). The same shocks may have a non-transitory effect on inflation in the countries with exchange rate pegged to the US dollar. Cecchetti and Moessner (2008), however, find no second-round inflationary effect in their sample of 19 developed and emerging economies from 1994 to 2008. Gospodinov and Ng (2013) use the principal component of convenience yield (the linear combination of spot and future commodities prices) to reveal the predictive power of six commodities' prices (including cocoa, orange juice, copper, soybeans, oats, and silver) on inflation of G7 countries during March 1983 to July 2008, but little explanatory power of the International Monetary Fund (IMF) aggregate commodity index. In the research of Chen, Turnovsky, and Zivot (2014) for Australia, Canada, Chile, New Zealand, and South Africa from 1983Q1 to 2010Q3, both aggregate and disaggregate commodity price indexes perform well at predicting consumer price index (CPI) and producers price index inflation; the energy price index is the pivot contributor. Cecchetti and Moessner (2008), however, find a significant role of food prices in predicting future headline inflation. Across countries, the evidence suggests that the inflationary effect of commodity-price shocks is larger and more persistent in

[25] For example, a highly volatile shock such as a technology shock that tends to generate a large loss of firm profits unless a quick price adjustment is made will attract more firms' attention and lead to fast responses of prices (Paciello, 2012).

developing countries than in developed ones (Cecchetti and Moessner (2008), Sekine and Tsuruga (2018)).[26]

Given the exposure to global commodity-price shocks, does adopting an inflation-targeting regime make a difference in output and inflation responses? Early theoretical studies supporting IT argue that an official announcement on an explicit target helps anchor inflation expectations and enhance the central banks' credibility, thereby lowering inflation volatility (Mishkin (1999), Svensson (1997)). The supportive empirical evidence is extensive (see, for example, Vega and Winkelried (2005), Lin and Ye (2009), Capistrán and Ramos-Francia (2010), De Mendonça and e Souza (2012), and Bems, Caselli, Grigoli, Gruss, and Lian (2018)). The counter-arguments point out that the monetary regime is far from a panacea. Mishkin (2000) addresses major disadvantages of IT, including, for instance, the rigidity of inflation targets, the difficulty of controlling inflation, the lag of monetary policy responses, the erosion of the central banks' accountability, as well as the political economy of the increasing temptation for policy discretion and the resultant output instability.

A major branch of the literature focuses on the effects of IT on the output-inflation trade-off and whether or not IT is the optimal policy. The theoretical evidence is largely inconclusive though. Goodfriend and King (2001) found little evidence on the long-run trade-off between inflation and real activity at low inflation rates but a positive long-run association at higher inflation levels. However, this so-called *divine coincidence* hypothesis — the simultaneous price and output stability (thus, no trade-off) — tends to dissipate with real market imperfections such as real wage rigidities (Blanchard & Galí, 2007) or without a zero-inflation condition (Alves, 2014). Erceg, Henderson, and Levin (2000) and Bodenstein, Erceg, and Guerrieri (2008) show in the calibration the existence of the trade-off in stabilizing output gap and inflation. Erceg et al. (2000) find that, under a staggered price setting, a strict price inflation targeting induces relatively large welfare losses and is not the optimal rule. However, Natal (2012) studies a monopolistic competition model focusing on the trade-off in the presence of an oil shock and finds that the IT regime is the optimal rule. For a small commodity-exporting country, Drechsel, McLeay, and Tenreyro (2019) provide a case of an inefficient boom as a result of a positive commodity-price shock with rising inflation as well as output in which IT leads to the outcomes most similar to the optimal policy.

[26] Kilian (2009) is one of the pioneering studies showing that demand and supply shocks have different dynamic effects on the real price of oil and hence the economy. Understanding this is important for policymakers, but isolating the shocks or the purely inflationary effects is not a trivial task.

The empirical findings are more concentrated, tending to support the substantial benefits of IT on inflation and output in emerging markets but weak in advanced countries (Batini and Laxton (2007), Gonçalves and Salles (2008), Ball (2010), Mishkin and Schmidt-Hebbel (2007), and Huang, Yeh, and Wang (2019)). For example, Batini and Laxton (2007) study 13 IT and 29 non-IT emerging economies during 1984-2005, and Ball (2010) studies 20 advanced economies from 1985 to 2007, both using the difference-in-difference approach to show that IT has an insignificant impact on inflation in advanced countries. However, IT is associated with lower inflation, inflation expectations, and inflation volatility while leaving output unchanged in emerging countries. Gonçalves and Salles (2008) use the same approach for 36 emerging economies (13 of which adopted IT) from 1980 to 2005 to reveal a slightly different but still supportive story: IT developing countries tend to experience concrete welfare gains with a larger drop in both inflation and growth volatility. Mishkin and Schmidt-Hebbel (2007) use a panel VAR with lagged IT dummy as an instrumental variable for a sample of 21 industrial and emerging IT countries and 13 emerging non-IT countries to find that IT countries have a smaller inflation response to an oil-price shock and tend to achieve lower inflation in the long-run. More recently, Huang et al. (2019) find a lower output-inflation trade-off across both developed and developing sub-samples thanks to IT adoption. While the findings supportive of IT are quite popular in the empirics, the question of how to solve the endogeneity of monetary policy remains open, which in turn casts some doubt on the existing findings. In this context, my study contributes to the abovementioned literature, using the SVAR approach to better control the endogeneity of multiple macroeconomic variables for a larger sample, and thus providing more robust results.

The remainder of this chapter proceeds as follows. Section 3.3 introduces the methodology for accessing the effects of symmetric shocks in global commodity prices. Section 3.4 reports the benchmark results focusing on the responses of output growth and inflation for the IT and the non-IT countries. I also extend the chapter by including several robustness checks using interannual data, investigating asymmetric shocks, and using a more restricted IT classification. Section 3.5 concludes.

3.3 Methodology

3.3.1 Data

The unbalanced sample covers 99 EME countries over the 1990Q1-2016Q4 period. I use global prices of agriculture, fuel, and metals (including minerals), and country-specific commodity terms-of-trade, GDP growth rate, inflation rate, interest rate, and bilateral real exchange rates for estimations. The data on the three global commodity prices is from World Bank's Pink Sheets, available from 1960 to the present. These commodity prices serve as the common exogenous shocks for all countries in this analysis. The average price index of agriculture is based on the prices of food, beverages, and agricultural raw materials. The fuel index is a weighted average of spot prices of coal, crude oil, and natural gas. The weighted average price index of metal and minerals is computed from the spot prices of aluminum, copper, iron ore, lead, nickel, steel, tin, and zinc. I deflate the nominal series with the US CPI (2015 base index) and average the monthly series to get the quarterly price indices. I use the first difference of the natural logarithm of each series (i.e., quarter-on-quarter growth rate), designating them as agriculture (*a*), fuel (*f*), and metal (*m*), in the estimations.

The first series in the domestic block of macroeconomic variables is the quarter-on-quarter growth rate of commodity terms-of-trade (*tot*). Bertrand and Suhaib (2019) provide both time-variant and -invariant commodity terms-of-trade indices, weighted by commodity trade balance and GDP. In the baseline estimation, I use the rolling index weighted by commodity trade flows.[27] I average the monthly series to get quarterly data. Other series in the domestic block include the quarter-on-quarter growth rate of GDP (*gdp*). I use the real quarterly seasonally-adjusted GDP (in 2015 US dollar) from Oxford Economics/Datastream to calculate GDP quarterly growth rate. I also include the first difference of central bank policy rate (*ir*) in SVAR estimations. The data on interest rate is sourced from the Monetary and Financial Statistics of the IMF and Datastream. I also use the quarter-on-quarter growth rate of real (CPI deflated) bilateral exchange rate of the local currency against the US dollar (*rer*), computed from the nominal exchange rate from the International Financial Statistics (IFS) of the IMF. By construction, an increase in the bilateral exchange rate indicates the local currency's depreciation. The last domestic variable is the first difference of the quarter-on-quarter growth rate of the CPI index (*inf*) sourced from the IFS and Datastream, which went through the

[27] The findings in the baseline estimation remain intact as we use alternative country-specific commodity terms-of-trade indices weighted either by rolling GDP, fixed commodity trade balance, or fixed GDP.

seasonal adjustment using the X-13ARIMA-SEATS program. Only countries with strongly balanced data for at least 30 consecutive quarters are in the final sample.[28]

3.3.2 Differences between inflation-targeting and non-inflation targeting countries

Among 99 EME economies studied, 25 countries explicitly adopted IT as of 2016, and 74 countries did not. Poland is the earliest IT adopter in the sample (in September 1998), followed by Colombia, Brazil, and Chile (all in 1999). The latest IT adopters are Argentina, India, Russia (in December 2015), and Ukraine (in December 2016). I follow Jahan (2012) to classify the IT vis-à-vis the non-IT countries and further collect data on the quarters when they adopted IT officially (i.e., when IT came into force by law). Note that the official date when IT came into force by law in a country may differ from when the central bank implemented the scheme. For example, Chile adopted IT officially in September 1990, but the central bank did not implement the IT scheme until September 1999. Also, no country that has adopted IT has abandoned this scheme, except the countries that subsequently joined the European Union, including Finland and Spain in 1999 and Slovakia in 2009, which are not in my sample.

I first describe some striking differences in GDP growth and inflation performance between the IT group and the non-IT group, as well as the improvement in GDP growth-inflation association in the former group since IT adoption. Figure 3.1 simply depicts the GDP-weighted association between output growth and inflation by country in the IT and the non-IT groups, taking into account all variations during the 1990Q1-2016Q4 period. I use the constant GDP in 1990 (in 2010 $US) by group of countries as the weight to control for the initial development levels. While more non-IT countries display positive correlations (concentrated bars on the right-hand side), most of the IT-countries show negative correlations (concentrated bars on the left-hand side). On average, I find from a t-test for mean differences that the mean correlation coefficient for the IT countries is statistically negative while that for the non-IT countries is not different from zero. To explore further the trend in output-inflation association in the IT countries, I focus on a sub-sample of 15 IT countries having strongly balanced data for a 6-year window: three years before and three years after IT adoption (Figure 3.2). In that figure, the correlation coefficients between output growth and inflation are 12-quarter rolling windowed and GDP-weighted, emphasizing the decrease in average correlation coefficient after IT adoption. Those two figures say that, though the correlation of output and inflation for

[28] See Appendix 3 for detailed data sources and the computation method of the variables.

the IT countries is quite strikingly lower than that for the non-IT countries, the average correlation for the IT countries is not always necessarily negative. Also, the correlation coefficients have little to tell us about the performance of GDP growth and inflation separately. I then add another point in Figure 3.3, depicting the (GDP-weighted) output growth improvement as well as lower and less volatile inflation in the IT countries during three years after IT adoption, compared to the previous 3-year period.[29] I dig deeper into SVAR estimations to deliver more robust results in section 3.4.

3.3.3 Large emerging-economy exporters and importers of commodities

The importance of commodities in the global market has changed substantially over the 1990-2016 period. For example, China's oil import ratio (to total imports) increased from 0.87% in 1990 to 11% in 2016[30], and its global market share of fuel imports increased from 0.34% in 1990 to 11% in 2016. To classify the influential commodity exporters and importers, I rank the median of annual market shares of exports and imports by country-commodity, focusing on three individual commodity groups: agriculture, fuel, and metal, over the 1990-2016 period. For example, the market share of agriculture export of a country is the ratio of their total exported raw materials, food, and beverages to the corresponding global agriculture exports. I use the data on commodity exports and imports from WDI and the United Nations Conference on Trade and Development (UNCTAD), with similar commodity groupings as in World Bank Pink Sheets for the three global commodities.

Among the sampled EME countries, Brazil is the leading exporter in the global market for agricultural products, accounting for 3.4% of global agriculture export, followed by China (2.97%), Thailand (2.1%), and Argentina (2.09%). China and Russia are the top agriculture importers with 4.34% and 2.24% market shares, respectively. The largest exporting share of the global fuel market belongs to Saudi Arabia (10.16%), followed by Russia (9.57%). Other large exporters include Venezuela (3.25%), Iran (3%), Kuwait (2.45%), Algeria (2.44%), and Iraq (2.03%). The two largest importers of fuels are China and India, accounting for 3.68% and 3.08%, respectively. For metals, Chile and Russia are among the largest exporters, accounting for 4.4% and 3.98% of the global markets, followed by South Africa (2.99%), China (2.9%),

[29] In Figures 3.1 – 3.3, we use year-on-year growth rates of GDP and CPI index, in which seasonal components are removed by construction. We weight the correlation coefficients and growth rates by real GDP (in 2010 $US) in 1990 to correct for the possible effects of economy size. Rolling coefficients in 0 are to account for the effects of pre-IT time. Data of real GDP in 1990 is from World Development Indicators (WDI).

[30] Access https://oec.world/ for quick visualization of export and import breakdowns by country.

and Brazil (2.77%). China has the largest importing share (9.5%) of metals, followed by relatively small importers such as Turkey, India, Mexico, Malaysia, and Thailand, each accounting for less than 2% share.

This analysis treats most countries as commodity price-takers, i.e., the domestic economy has no contemporary impact on global commodity prices. However, for large commodity consumers and producers, their domestic business cycles can affect global commodity production and consumption, and thus global commodity prices, rendering the assumption of the exogeneity invalid. To maintain the exogeneity assumption, I exclude the country from estimating if its export share or import share in any of global agriculture, fuel, and metal markets exceed 3%; hence, China, India, Russia, Saudi Arabia, Venezuela, Brazil, and Chile are not in the baseline estimation. This exclusion, however, does not affect the main findings. These seven countries remain in the full sample for a robustness check.

3.3.4 Benchmark specification

The SVAR(4) specification for each country has the stacked form as follows:

$$A_0 Z_t = A_1 Z_{t-1} + \mu_t \tag{3.1}$$

where

$$Z_t = \begin{bmatrix} X_t \\ X_{t-1} \\ X_{t-2} \\ X_{t-3} \end{bmatrix} \text{ and } X_t = \begin{bmatrix} p_t \\ tot_t \\ gdp_t \\ inf_t \\ ir_t \\ rer_t \end{bmatrix} \tag{3.2}$$

The first element in the vector X_t is the quarterly growth rate of a global commodity price (agriculture (a_t), fuel (f_t), or metal (m_t)). The other five variables are country-specific macro series, including the quarterly growth rate of commodity terms-of-trade (tot_t), quarterly real GDP growth rate (gdp_t), first-difference of quarterly inflation rate (inf_t), first-difference of interest rate (ir_t), and quarterly growth rate of the real exchange rate (rer_t).

I estimate the system with four lags. For equation (3.1), the objects A_0 and A_1 are thus 24-by-24 matrices of contemporaneous and lagged structural coefficients, respectively. Specifically, the matrices A_0 and A_1 have the following forms:

$$A_0 = \begin{bmatrix} a_{0,11} & a_{0,12} & a_{0,13} & a_{0,14} \\ a_{0,21} & a_{0,22} & a_{0,23} & a_{0,24} \\ a_{0,31} & a_{0,32} & a_{0,33} & a_{0,34} \\ a_{0,41} & a_{0,42} & a_{0,43} & a_{0,44} \end{bmatrix} \quad \text{and} \quad A_1 = \begin{bmatrix} a_{1,11} & a_{1,12} & a_{1,13} & a_{1,14} \\ a_{1,21} & a_{1,22} & a_{1,23} & a_{1,24} \\ a_{1,31} & a_{1,32} & a_{1,33} & a_{1,34} \\ a_{1,41} & a_{1,42} & a_{1,43} & a_{1,44} \end{bmatrix}$$

, where each block a is a 6-by-6 matrix. The object A_0 is a lower triangular matrix with one on its main diagonal. That is, the blocks $a_{0,11}$, $a_{0,22}$, $a_{0,33}$, and the block $a_{0,44}$ are lower triangular 6-by-6 matrices with one on their main diagonal. Other blocks in the matrix A_0 are all zeros. The last term, μ_t, is a 24-by-1 random vector with zero mean and diagonal variance-covariance matrix Σ. Multiplying the system (1) by A_0^{-1}, I have:

$$Z_t = AZ_{t-1} + \Pi\varepsilon_t \tag{3.3}$$

, where $A \equiv A_0^{-1}A_1$, $\Pi \equiv A_0^{-1}\Sigma^{1/2}$, and $\varepsilon_t \equiv \Sigma^{-1/2}\mu_t$. The term ε_t is a random vector with zero mean and identity variance-covariance matrix.

I treat the global commodity-price shocks and country-specific commodity terms-of-trade as exogenous, and the former can affect the latter but not vice versa. Accordingly, I impose restrictions on the matrix A_1 as follows. For each of the blocks $a_{1,11}$, $a_{1,12}$, $a_{1,13}$, and the block $a_{1,14}$, all elements on their first rows are zeros except the first element, and all elements on their second rows are zeros except the first two elements. The blocks $a_{1,21}$, $a_{1,32}$, and the block $a_{1,43}$ are lower triangular matrices with one on their main diagonals. Other 6-by-6 blocks of the matrix A_1 contain all zeros. The specification allows the global commodity prices to follow a univariate autoregressive process AR(4). The commodity terms-of-trade is determined contemporaneously by the global commodity price. The global commodity prices and commodity terms-of-trade affect the domestic block contemporarily but not vice versa. My main interest is to investigate the domestic variables' responses to the global commodity-price shocks, as long as the commodity terms-of-trade follows the global commodity prices, the order of other domestic variables in the SVAR system is immaterial. In this study, the order of the SVAR variables remains as in equation (3.2). I estimate the matrices A_0, A_1, and the matrix Σ country-by-country by OLS, and then compute the impulse responses and cumulative responses of the domestic variables to a 10% increase in each global commodity price (starting in the 0^{th} quarter) up to 20 quarters following the shocks. The results will focus on comparing the median responses of GDP growth and inflation between the IT and the non-IT countries. I apply the Cholesky decomposition to compute the impulse responses and extract the variance share explained by the global shocks. Following the literature, I then calculate 68% confidence intervals using Monte Carlo bootstrapping approach.

3.4 Results

3.4.1 Preliminary tests

Before the estimation, I check the stationarity of the following variables country-by-country (with 792 series in total): growth rate of global commodity prices (*p*), growth rate of country-specific commodity terms-of-trade (*tot*), growth rate of GDP (*gdp*), first-difference of inflation rate (*inf*), first-difference of interest rate (*ir*), and growth rate of real exchange rate (*rer*) using the Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) unit root tests. At a 5% level of significance, the ADF tests for the level-series including intercept only indicate four non-stationary series including *m*, *gdp*, and *ir* in three countries. As including both intercept and time trend, the ADF tests indicate 9 non-stationary series of *m*, *gdp*, and *ir* across six countries. The PP tests for the level-series including intercept indicate 16 non-stationary series of *gdp* in 16 countries. Including both intercept and time trend in the PP tests results in 32 non-stationary series. Section A3.2 in Appendix 3 lists the non-stationary variables in according countries. I proceed with the ADF results, including the variables above in country-specific SVAR estimations, and make sure the SVAR systems are stable (i.e., all eigenvalues are inside the unit circle).

I check the optimal lag length for each country using a series of criteria including the Akaike Information Criterion (AIC), Hannan-Quinn Criterion (HQ), Schwarz Criterion (SC), and Final Prediction Error Criterion (FPE) with the maximum lags of 10. For the 99-country sample, AIC and FPE suggest the optimal lag length of 10 while HQ and SC suggest the lag length of 1. Table A3.2 in Appendix 3 reports the frequency of optimal lags suggested by those four criteria. Because the range of implied optimal lag lengths is so diverse across those criteria, and that more lags result in a loss of degree of freedom, I proceed with VAR estimations using four lags.[31]

3.4.2 Baseline results

3.4.2.1 Does IT make a difference in the growth and inflation responses?

This section compares the responses of output growth and inflation between the IT group (21 countries) and the non-IT group (71 countries) to the global commodity-price shocks. I find

[31] Detailed results of the ADF and the PP unit root tests for 8 variables and the optimal lag lengths suggested by AIC, HQ, SC, and FPE by country are provided upon request.

that the IT and the non-IT groups have similar inflationary responses, which are also transitory to the shocks while only the IT group shows persistent increases in the GDP growth. I do not detect any differences in the responses of commodity terms-of-trade and interest rates between the two groups.[32]

Following a 10% increase in the global agriculture price, the GDP growth of the IT group rises by 0.43%, but the non-IT group's growth remains unchanged (Figure 3.4, panel A). The inflation rates for both IT and non-IT groups increase by 0.3% - 0.4%. The interest rates of the IT and the non-IT countries are not responsive to the agriculture-price shock. The 10% increases in the global fuel and metal prices are associated with a 0.2% increase in the inflation rate of the IT and the non-IT countries (Figure 3.4, panels B-C). The responses of output growth to these two shocks differ remarkably between the IT and the non-IT groups. The IT group's growth increases by 0.11% (0.17% with a one-quarter delay) following the fuel-price shock. To the global metal-price shock, the GDP growth of the IT countries increases by 0.15% (0.32% with a one-quarter delay). GDP growth of the non-IT group is not responsive to both global fuel- and metal-price shocks. I also find minor and similar responses of the interest rate for the IT and the non-IT groups.

Figure 3.5 reports the cumulative responses from the estimated SVARs. The IT countries show positive and long-lasting cumulative responses of GDP growth following the global agriculture-price shock (for 11 quarters; 1.1% after a year), fuel-price shock (for 6 quarters; 0.3% after a year), and metal-price shock (persistent for years; 0.62% after a year). The non-IT group, however, displays slight adjustment in its GDP growth. The cumulative impacts on inflation and interest rate are very similar across the IT and the non-IT countries. Table 3.1 summarizes the contemporaneous and one-year cumulative effects of the global commodity-price shocks on the domestic variables across the IT and the non-IT countries. I also report the ratios of countries in the IT group and the non-IT group with significant responses of GDP growth and inflation to the global commodity price shocks in Table 3.2. Overall, more IT countries show significant improvements in GDP growth relative to the non-IT countries whereas quite similar ratios of two groups showing significant responses in inflation to the shocks.

[32] The inclusion of China, India, Russia, Saudi Arabia, Venezuela, Brazil, and Chile (thus, 25 IT countries and 74 non-IT countries) does not change our main findings. Country-specific impulse responses and corresponding replication files for impulse responses are available upon request. Additional results weighting the impulse responses by constant GDP in 1990 to control for the country's level of economic development are also consistent with the main findings. The figures for GDP-weighted impulse responses are in Appendix 3.

The different responding patterns in GDP growth and inflation between the IT group and non-IT group could be explained by several reasons. The central banks in the IT countries set either specific inflation target rates, possibly including upper limits, or symmetric bands around a midpoint; all at low single digits. This framework as a combination of "rules" and "discretion" allows the central banks to be flexible as achieving a target rate in the medium term, typically over a two- to three-year horizon and a target band in the short term along with addressing other objectives such as smoothing output, wages, employment level, or exchange rate. As economic shocks happen, the policymakers would not be enforced to do whatever it takes to meet the inflation targets on a period-by-period basis but rather focus on, for example, stabilizing output. Broad empirical evidence also support the effectiveness of inflation targeting framework in delivering low inflation volatility and anchoring inflation expectations. This could explain why there is indiscernible and similar inflation effects observed across the IT countries and non-IT countries whereas the GDP responses are very different between those two groups. Please also note that this chapter is using the explicit IT classification, which is acknowledged as one of my limitations by possibly excluding the implicit IT adopters. However, the data of implicit IT countries are not in hand. Another limitation may belong to the headline inflation data which includes food and energy prices used in this chapter. As the global commodity prices are very volatile and exogenous to most of the economies, the central banks in the IT countries may tend to stabilize core inflation instead of headline inflation. However, a comparable dataset of core inflation is not available for cross-check. In addition, the global commodity price shocks are essentially caused by different shocks such as global demand and global supply shocks which may drive domestic inflation differently. Decomposing the global commodity price shocks could be considered for a future research, possibly applying the approach proposed by Kilian (2009).

3.4.2.2 Variance decomposition

Table 3.2 reports the domestic variables' median variance shares explained by the global commodity-price shocks by country group. On average, the global commodity-price shocks can explain 10% - 14% of variances of GDP growth and 11% - 16% of variances of inflation (panel A). The results are comparable to the estimates in other studies that find a small role of terms-of-trade shocks, less than 10%, to output disturbances (for example, Schmitt-Grohé and Uribe (2018)). Using a block of three global commodity price indices (agriculture, fuel, and metal), however, Fernández et al. (2017) find a much larger role: the block of three global commodity price indices explains about 33% of the output gap's fluctuations. The big difference from my findings could result from their employment of a block of three indices instead of an aggregate index. Across the subsamples, I find the global commodity-price shocks account for 18% - 20% of variances of output and 9% - 18% of variances of inflation for the IT countries, and 9% - 13% of variances of output and 12% - 16% of variances of inflation for the non-IT countries (panel B). The fuel-price shock dominates the agriculture- and metal-price shocks in explaining a greater variance share of inflation: the fuel-price shock accounts for 18% of inflation variations for the IT countries and 16% for the non-IT countries; the agriculture-price shock explains 9% for the IT countries and 14% for the non-IT countries, and the metal-price shock explains 10% for the IT countries and 12% for the non-IT countries.

3.4.3 Robustness checks

3.4.3.1 Interannual data frequency

I re-construct the variables in interannual changes, i.e., the year-on-year (change from the corresponding quarter of the previous year; hence, no seasonal components) and re-estimate SVAR(4) by country. Specifically, p_t is the four-quarter difference (hereafter referred to as *interannual change*) of a real global commodity price index (in the natural log); tot_t is the interannual change of commodity terms-of-trade (in the natural log); gdp_t is the interannual change of real GDP (in the natural log); inf_t is the interannual change of non-seasonally adjusted CPI (in the natural log); ir_t is the interannual change of interest rate; and rer_t is the interannual change of real exchange rate (in the natural log). The first-differences of gdp_t, inf_t, ir_t, and rer_t enter the SVAR estimations. The sample with reconstructed data includes 81 countries but I focus on the impulse responses of 74 countries excluding China, India, Russia, Saudi Arabia, Venezuela, Brazil, and Chile. Detailed impulse responses and variance decomposition for 81 countries are available upon request.

Figure 3.6 provides consistent findings with the baseline results: the IT and the non-IT groups have similar inflation responses but only the IT group has a positive GDP growth response to the global commodity-price shocks. Following a 10% increase in the global agriculture price, both IT and non-IT groups exhibit a resembling increase of 0.53% in inflation with a one-quarter lag while the GDP growth responds positively only for the IT countries, by 0.29% (panel A). Following a 10% increase in the global fuel price, the GDP growth increases on impact by 0.08% for the IT group, peaking at 0.23% in the following quarter whereas the non-IT group's response is negligible (panel B). Although the IT and the non-IT groups have comparable inflationary responses of 0.2% to the fuel-price shock, the impact lasts a-quarter longer for the non-IT countries. In panel C, the 10% increase in the global metal price is associated with a 0.23% increase in inflation, similarly for both groups of countries, and a 0.28% increase in GDP growth (with a one-quarter lag) for the IT countries. In no circumstances the responses of interest rate are visible. That commodity terms-of-trade are negatively associated with the three global commodity-price shocks is also consistent with the baseline results.

3.4.3.2 Asymmetric shocks

I further extend this study by considering the asymmetric effects of positive vis-à-vis negative shocks on output growth and inflation. Previous studies on the asymmetric effects of oil prices on output include, for instance, Cuñado and de Gracia (2003), Jiménez-Rodríguez and Sánchez (2005), and Berument et al. (2010). Cuñado and de Gracia (2003) construct asymmetric variables in different ways: positive year-on-year oil price growth; maximum growth of oil prices compared to one, two, three, and four years; and standardized oil price increases with the standard deviation coming from the GARCH (1,1) specification. Using VAR estimations for 15 European countries over the 1960-1999 period, they find that oil price increases lower growth rate of industrial production index. In a similar approach for G7 plus Norway from 1972Q3 to 2001Q4 using positive and negative oil price changes as separate variables, Jiménez-Rodríguez and Sánchez (2005) also find asymmetric accumulated effects. After 4 quarters, the positive oil-price shocks have a negative impact on GDP growth in most of the countries, except Japan and the UK with a positive association. The impacts of negative oil-price changes are largely insignificant except negative in the US and Canada, and positive in the UK. The positive standardized oil-price shocks, however, increase one-year accumulated GDP growth in Norway but lower in the UK while the impacts of negative standardized oil-price shocks on output remain largely insignificant. Berument et al. (2010) construct asymmetric variables in a similar way and estimate SVARs for 16 selected countries in the Middle East and North African region (time periods varying across countries). They find evidence on asymmetric effects of oil price shocks in Egypt, Iraq, and Tunisia as including positive and negative shocks as separate variables, in Egypt, Iraq, Syria, and Tunisia as including normalized shocks.

In this section, I adjust the specification to satisfy the assumption of exogeneity and allows the estimators to be comparable across groups. Specifically, I distinguish between a positive shock (i.e., a positive percentage change), p_t^+, and a negative shock (i.e. a negative percentage change), p_t^- as follows:

$$p_t^+ = \begin{cases} p_t & if\ p_t > 0 \\ 0 & otherwise \end{cases}$$

$$p_t^- = \begin{cases} p_t & if\ p_t < 0 \\ 0 & otherwise \end{cases}$$

, where p_t is a quarter-on-quarter growth rate of a global commodity price.

Accordingly, vector X_t consists of seven variables, with the positive and negative global commodity-price shocks as exogenous factors and contemporaneously uncorrelated:

$$X_t = \begin{bmatrix} p_t^+ \\ p_t^- \\ tot_t \\ gdp_t \\ inf_t \\ ir_t \\ rer_t \end{bmatrix}$$

This examination leaves us 87 countries. In Figures 3.7 – 3.9, I compare the impulse responses and cumulative responses of GDP growth and inflation between the IT group and the non-IT group excluding China, India, Russia, Saudi Arabia, Venezuela, Brazil, and Chile, to 10% positive and negative changes of the global commodity prices. The results are largely intact as I include those seven large commodity exporters and importers. In sum, I find evidence of asymmetric effects of the global commodity-price shocks. The domestic variables are more responsive to the negative shocks than the positive shocks, especially to the agriculture-price and metal-price shocks whereas most of the responses to the positive shocks are negligible. More importantly, the IT countries appear to be more resilient with long-lasting GDP expansions as responding to the negative price shocks, compared to their non-IT counterparts. Meanwhile, the negative shocks, i.e., declines in the global commodity prices, leave GDP growth of the non-IT countries unchanged. I also find the responding patterns of inflation are temporary, similarly for both IT and non-IT groups, and the responses of interest rates are marginal. I report the detailed results as follows.

Figure 3.7, panel A shows that the negative agriculture-price shock is positively associated with the IT countries' GDP growth, which increases by 0.6%, peaking at 1.2% in the following quarter while the GDP growth's responses of the non-IT group remain unchanged. The responses of inflation and interest rates of both IT and non-IT groups are indifferent across the asymmetric agriculture-price shocks. In panel B, I also find the persistent cumulative expansion in GDP growth of the IT group to the negative agriculture-price shock (2.64% after one year) while that of the non-IT group is negligible. The one-year cumulative responses of inflation to the negative agriculture-price shock and those of output growth and inflation to the positive agriculture-price shock are trivial across both groups.

In Figure 3.8, panel A, the negative fuel-price shock immediately causes a 0.16% expansion in GDP growth and the positive fuel-price shock immediately leads to a 0.22% increase in inflation of the IT countries. The IT countries also display persistent cumulative expansions in GDP growth (0.49% after one year), lasting for five quarters following the negative fuel-price shock (panel B). The one-year accumulated increase in output growth is marginal though. I also find no persistent responses of inflation and interest rates in any group of countries.

As seen in Figure 3.9, panel A, the negative metal-price shock is associated with a 0.2% increase in inflation for both IT and non-IT groups and a 0.32% increase in GDP growth for the IT countries. As in panel B, the cumulative responses of the IT countries' GDP growth to the negative metal-price shock remain positive during the first year (0.7% after a year) and those to the positive metal-price shock became persistently positive after six quarters. However, I find the cumulative responses of the non-IT countries' GDP growth marginal to the shocks. The association between the negative metal-price shock and inflation as well as interest rate are largely transitory and negligible.

3.4.3.3 Restricted IT classification

In this section, I cross-check the findings above (both baseline model and robustness checks) using a more restricted IT classification to exclude the recent IT adopters. As the sample spreads from 1990Q1 up to 2016Q4, I choose 2010Q1 as the threshold: a country is classified as an IT country if their starting date of IT adoption is before 2010Q1 because I strongly believe that seven years would be long enough to see the impact of IT scheme. Although it makes sense to exclude the recent IT countries, the pre-IT time could play a role in achieving GDP growth and stabilising inflation as IT countries may take time to prepare for IT adoption. For this reason, the baseline results is kept as a reference. Using the restricted IT classification leaves 15 IT countries and 77 non-IT countries in the baseline sample, 14 and 60 in the first robustness check using interannual data, and 14 and 66 countries in the second robustness check of asymmetric analysis respectively, excluding the seven large global commodity exporters and importers. The results are shown in Table 3.4, strengthening the findings above that: (1) the IT group displays persistent and discernible GDP growth improvements but not the non-IT group, which is also found true in the symmetric and asymmetric analyses; (2) the inflationary responses are transitory and similar for both groups.[33]

3.5 Conclusion

This chapter analyses the effects of the global commodity-price shocks on 99 EME economies from 1990Q1 to 2016Q4. After symmetric commodity-price shocks, I find that the IT countries display persistent GDP growth improvements, with cumulative increases persistent at least for six following quarters. However, the GDP growth responses of the non-IT countries are marginal. In the analysis of asymmetric shocks, I also find strong evidence indicating that the IT group is more resilient to the negative price shocks with long-lasting improvements in the GDP growth compared to the non-IT group. The impacts of the positive price shocks are generally silent. In any case, inflationary responses are transitory, similarly for both groups. In addition, the variance decomposition shows that the global commodity-price shocks play a modest role in explaining the variations of output and inflation, with the fuel-price shock having largest effects than the agriculture-price and metal-price shocks.

[33] Appendix Section A.3.4 also breaks downs the sample into the commodity net exporters and commodity net importers as examining their responses to the global commodity price shocks. Overall, I find that the commodity net importers are more vulnerable to the global commodity-price shocks than the commodity net exporters in terms of lower GDP growth, as well as higher and persistent inflation. A potential future research will study the impact of inflation targeting framework on the responses of the commodity net exporters and the commodity net importers to the global commodity price shocks.

Table 3.1 Impulse responses to the global commodity-price shocks

	Agriculture-price shock		Fuel-price shock		Metal-price shock	
	IT	Non-IT	IT	Non-IT	IT	Non-IT
			Baseline			
Short-term effect						
tot	-2.06	-2.31	-2.35	-2.24	-1.58	-1.29
gdp	0.43	0.02	0.11	0.02	0.15	0.01
inf	0.29	0.4	0.15	0.18	0.19	0.18
ir	-0.04	-0.02	-0.01	0.02	-0.03	0
rer	-2.45	-1.87	-0.76	-0.24	-1.21	-0.28
Cumulative one-year effect						
tot	-1.13	-2.67	-2.06	-2.61	-2.71	-2.3
gdp	1.1	0.13	0.3	0.07	0.62	0.2
inf	0.23	0.26	0.03	0.07	0.15	0.18
ir	0.78	0.11	0.19	0.06	0.34	0.1
rer	-3.36	-3.76	-1.05	-0.5	-2.47	-0.27
			Robustness check 1: Interannual data frequency			
Short-term effect						
tot	-1.94	-2.38	-2.35	-1.88	-1.36	-1.24
gdp	0.29	0.03	0.08	0.01	0.08	0.03
inf	0.11	0.22	0.18	0.2	0.23	0.14
ir	0	0	-0.03	0.02	0.05	0
rer	-2.4	-1.93	-1.31	-0.35	-1.5	-0.55
			Robustness check 2: Asymmetric shocks			
			Negative shock			
Short-term effect						
tot	-4.69	-3.57	-2.27	-2.3	-2.73	-2.92
gdp	0.6	0.02	0.16	0.02	0.32	0.03
inf	0.36	0.35	0.09	0.14	0.2	0.24
ir	0.01	0.04	-0.07	0.02	-0.02	0
rer	-2.85	-0.97	-0.68	-0.07	-1.51	-0.32
Cumulative one-year effect						
tot	-4.53	-6.84	-2.12	-3.24	-1.91	-3.28
gdp	2.64	0.9	0.49	0.18	0.7	0.19
inf	0.47	0.57	-0.01	0.14	0.19	0.25
ir	0.85	0.25	0.13	0.09	0.5	0.09
rer	-7.79	-3.2	-1.84	-0.35	-0.52	0.42
			Positive shock			
Short-term effect						
tot	0.49	-0.69	-1.66	-2.37	0.17	0.34
gdp	0.09	0.03	0.05	-0.01	0.06	-0.02
inf	0.07	0.19	0.22	0.19	0.11	-0.01
ir	-0.07	-0.05	-0.02	0.03	-0.01	-0.04
rer	-1.87	-2.73	-0.74	-0.63	0.01	-0.16
Cumulative one-year effect						
tot	0.64	-0.1	-2.37	-1.93	-1.4	-1.7
gdp	-0.41	-0.32	-0.18	-0.02	0.54	0.22
inf	0	-0.11	-0.02	-0.11	0.27	0.07
ir	0.17	0.02	0.01	0.03	0.63	0.05
rer	-1.92	-3.61	0.02	-0.38	-4.56	-0.83

Note: The detailed description of variables (*tot*, *gdp*, *inf*, *ir*, *rer*) is provided in Section 3.1 for the baseline estimation and Section 3.3 for the robustness checks.

Table 3.2 Ratios of countries showing significant responses in the IT and Non-IT groups

		Agriculture-price shock		Fuel-price shock		Metal-price shock	
		IT (21)	Non-IT (71)	IT (21)	Non-IT (71)	IT (21)	Non-IT (71)
Baseline							
Short-term effect							
	Significant	16 (76.19%)	32 (45.07%)	14 (66.67%)	34 (47.89%)	11 (52.38%)	39 (54.93%)
gdp	Negative	1 (6.25%)	13 (40.63%)	0 (0%)	11 (32.35%)	0 (0%)	18 (46.15%)
	Positive	15 (93.75%)	19 (59.37%)	14 (100%)	23 (67.65%)	11 (100%)	21 (53.85%)
	Significant	11 (52.38%)	44 (61.97%)	14 (66.67%)	50 (70.42%)	18 (85.71%)	42 (59.15%)
inf	Negative	0 (0%)	5 (11.36%)	0 (0%)	3 (6%)	1 (5.56%)	6 14.29%)
	Positive	11 (100%)	39 (88.64%)	14 (100%)	47 (94%)	17 (94.44%)	36 (85.71%)
Cumulative one-year effect							
	Significant	17 (80.95%)	20 (28.17%)	13 (61.90%)	29 (40.85%)	18 (85.71%)	36 (50.70%)
gdp	Negative	2 (11.76%)	8 (40%)	1 (7.69%)	7 (24.14%)	0 (0%)	8 (22.22%)
	Positive	15 (88.24%)	12 (60%)	12 (92.31%)	22 (75.86%)	18 (100%)	28 (77.78%)
	Significant	9 (42.86%)	28 39.44%)	5 (23.81%)	27 (38.03%)	9 (42.86%)	35 (49.30%)
inf	Negative	2 (22.22%)	5 (17.86%)	1 (20%)	6 (22.22%)	0 (0%)	3 (8.57%)
	Positive	7 (77.78%)	23 (82.14%)	4 (80%)	21 (77.78%)	9 (100%)	32 (91.43%)
Robustness check 1: Interannual data frequency							
Short-term effect							
	Significant	14 (73.68%)	31 (24%)	14 (73.68%)	34 (61.82%)	10 (52.63%)	30 (54.55%)
gdp	Negative	1 (7.14%)	9 (29.03%)	1 (7.14%)	12 (35.29%)	1 (10%)	10 (33.33%)
	Positive	13 (92.86%)	22 (70.97%)	13 (92.86%)	22 (64.71%)	9 (90%)	20 (66.67%)
	Significant	10 (52.63%)	32 (58.18%)	16 (84.21%)	40 (72.73%)	14 (73.68%)	30 (54.55%)
inf	Negative	1 (10%)	6 (18.75%)	2 (12.5%)	5 (12.5%)	1 (7.14%)	3 (10%)
	Positive	9 (90%)	26 (81.25%)	14 (87.5%)	35 (87.5%)	13 (92.86%)	27 (90%)
Robustness check 2: Asymmetric shocks							
Negative shock							
Short-term effect							
	Significant	14 (73.68%)	22 (36.07%)	12 (63.16%)	27 (44.26%)	15 (78.95%)	32 (52.46%)
gdp	Negative	0 (0%)	6 (27.27%)	1 (8.33%)	5 (18.52%)	0 (0%)	9 (28.13%)
	Positive	14 (100%)	16 (72.73%)	11 (91.67%)	22 (81.48%)	15 (100%)	23 (71.87%)
	Significant	9 (47.37%)	32 (52.46%)	10 (52.63%)	34 (55.74%)	12 (63.16%)	37 (60.66%)
inf	Negative	1 (11.11%)	7 (21.88%)	1 (10%)	7 (20.59%)	0 (0%)	5 (13.51%)
	Positive	8 (88.89%)	25 (78.12%)	9 (90%)	27 (79.41%)	12 (100%)	32 (86.49%)
Cumulative one-year effect							
	Significant	19 (100%)	25 (40.98%)	10 (52.63%)	19 (31.15%)	10 (52.63%)	26 (42.62%)
gdp	Negative	0 (0%)	2 (8%)	0 (0%)	2 (10.53%)	0 (0%)	6 (23.08%)
	Positive	19 (100%)	23 (92%)	10 (100%)	17 (89.47%)	10 (100%)	16 (76.92%)
	Significant	7 (36.84%)	27 (44.26%)	4 (21.05%)	23 (37.70%)	5 (26.32%)	21 (34.43%)
inf	Negative	0 (0%)	0 (0%)	1 (25%)	2 (8.70%)	1 (20%)	2 (9.52%)
	Positive	7 (100%)	27 (100%)	3 (75%)	21 (91.30%)	4 (80%)	19 (90.48%)
Positive shock							
Short-term effect							
	Significant	5 (26.32%)	18 (29.51%)	10 (52.63%)	30 (49.18%)	10 (52.63%)	29 (47.54%)
gdp	Negative	2 (40%)	7 (38.89%)	3 (30%)	13 (43.33%)	4 (40%)	21 (72.41%)
	Positive	3 (60%)	11 (61.11%)	7 (70%)	17 (56.67%)	6 (60%)	8 (27.59%)
	Significant	5 (26.32%)	26 (42.62%)	13 (68.42%)	36 (59.02%)	6 (31.58%)	23 (37.70%)
inf	Negative	2 (40%)	8 (30.77%)	1 (7.69%)	8 (22.22%)	2 (33.33%)	12 (52.17%)
	Positive	3 (60%)	18 (69.23%)	12 (92.31%)	28 (77.78%)	4 (66.67%)	11 (47.83%)
Cumulative one-year effect							
	Significant	7 (36.84%)	22 (36.07%)	2 (10.53%)	34 (55.74%)	10 (52.63%)	26 (42.62%)
gdp	Negative	6 (85.71%)	13 (59.09%)	2 (100%)	16 (47.06%)	0 (0%)	6 (23.08%)
	Positive	1 (14.29%)	9 (40.91%)	0 (0%)	18 (52.94%)	10 (100%)	20 (76.92%)
	Significant	5 (26.32%)	19 (31.15%)	3 (15.79%)	20 (32.79%)	5 (26.32%)	21 (34.43%)
inf	Negative	4 (80%)	12 (63.16%)	1 (33.33%)	15 (75%)	1 (20%)	2 (9.52%)
	Positive	1 (20%)	7 (36.84%)	2 (66.67%)	5 (25%)	4 (80%)	19 (90.48%)

Note: Table provides the numbers and ratios of countries with significant impulse responses of GDP growth (*gdp*) and inflation (*inf*) in the IT group and the non-IT group, excluding the 7 large commodity exporters and importers.

Table 3.3 Shares of variance explained by the symmetric global commodity-price shocks

	Number of countries	Share of variance					
		p	tot	gdp	inf	ir	rer
A. Whole sample	92						
Agriculture shock		100	38.31	10.14	13.23	12.10	16.58
Fuel shock		100	68.04	13.71	16.37	12.64	14.39
Metal shock		100	40.61	13.87	11.17	10.39	12.52
B. IT countries versus Non-IT countries							
IT countries	21						
Agriculture shock		100	34.58	17.72	9.47	7.95	19.29
Fuel shock		100	73.01	18.00	17.55	14.53	13.93
Metal shock		100	32.30	20.32	9.53	9.56	13.35
Non-IT countries	71						
Agriculture shock		100	38.36	8.64	14.30	12.27	16.32
Fuel shock		100	62.54	13.21	16.26	12.44	14.57
Metal shock		100	40.78	12.58	11.65	10.71	11.90

Note: Group-specific median shares of variance are computed from country-specific variance decompositions based on country-by-country SVAR estimations in the baseline case.

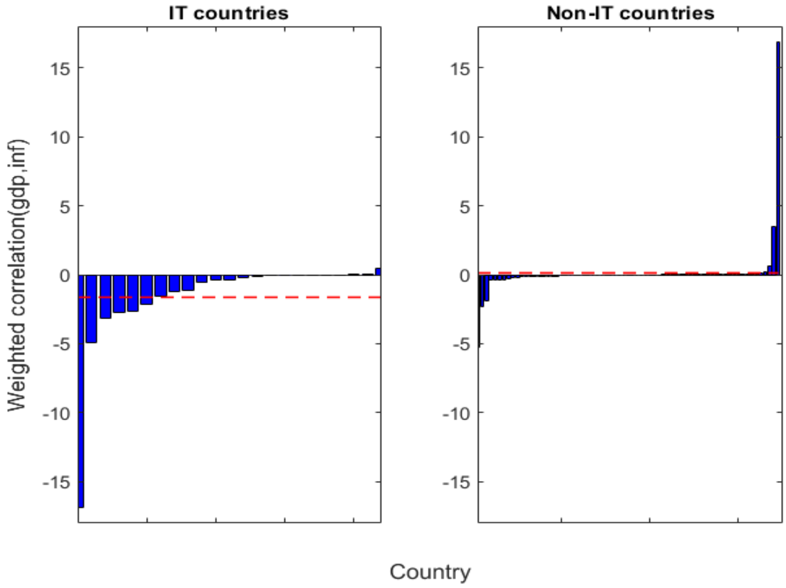
Table 3.4 Impulse responses to the global commodity-price shocks using restricted IT classification

	Agriculture-price shock		Fuel-price shock		Metal-price shock	
	IT	Non-IT	IT	Non-IT	IT	Non-IT
			Baseline			
Short-term effect						
tot	-2.72	-2.13	-2.2	-2.32	-1.53	-1.55
gdp	0.39	0.03	0.11	0.02	0.15	0.01
inf	0.28	0.4	0.12	0.18	0.17	0.18
ir	0.03	-0.02	0.02	0.02	0.04	-0.01
rer	-2.45	-1.92	-0.76	-0.3	-1.21	-0.28
Cumulative one-year effect						
tot	-1.71	-2.49	-2.03	-2.73	-2.37	-2.33
gdp	1.1	0.28	0.3	0.07	0.62	0.27
inf	0.23	0.23	0.03	0.07	0.15	0.18
ir	0.79	0.14	0.19	0.06	0.38	0.1
rer	-3.23	-3.78	-0.97	-0.53	-1.72	-0.41
			Robustness check 1: Interannual data frequency			
Short-term effect						
tot	-2.34	-2.24	-2.23	-1.94	-1.39	-1.27
gdp	0.29	0.04	0.09	0.01	0.08	0.02
inf	0.11	0.23	0.11	0.21	0.21	0.15
ir	0.06	0	-0.01	0.01	0.08	0
rer	-2.37	-2.02	-1.26	-0.38	-1.4	-0.59
			Robustness check 2: Asymmetric shocks			
			Negative shock			
Short-term effect						
tot	-5.27	-3.88	-2.24	-2.33	-2.55	-2.92
gdp	0.62	0.03	0.17	0.02	0.33	0.03
inf	0.25	0.36	0.09	0.14	0.14	0.25
ir	0.02	0.04	-0.05	0.02	0.03	0
rer	-2.66	-1.16	-0.85	-0.1	-1.2	-0.36
Cumulative one-year effect						
tot	-4.62	-6.16	-1.96	-3.26	-1.79	-3.12
gdp	2.62	0.97	0.47	0.21	0.7	0.23
inf	0.43	0.56	-0.03	0.12	0.18	0.24
ir	1.1	0.25	0.15	0.08	0.39	0.09
rer	-7.12	-3.79	-0.95	-0.44	-0.11	0.29
			Positive shock			
Short-term effect						
tot	-0.17	-0.31	-1.6	-2.43	0.2	0.3
gdp	0.05	0.04	0.04	-0.01	0.03	-0.02
inf	0.09	0.18	0.17	0.19	0.11	0.01
ir	-0.07	-0.06	0.06	0.02	0.01	-0.05
rer	-1.91	-2.65	-0.83	-0.63	0.06	-0.19
Cumulative one-year effect						
tot	0.63	-0.02	-1.72	-2.17	-1.5	-1.55
gdp	-0.63	-0.31	-0.06	-0.05	0.36	0.24
inf	0.03	-0.12	-0.03	-0.08	0.25	0.09
ir	0.09	0.03	0.05	0.02	0.62	0.05
rer	-1.11	-3.6	0.99	-0.45	-4.32	-0.92

Note: The detailed description of variables (*tot*, *gdp*, *inf*, *ir*, *rer*) is provided in Section 3.1 for the baseline estimation and Section 3.3 for the robustness checks. This table uses a more restricted IT classification, i.e. a country is classified as an IT country if their starting date of IT adoption is before 2010Q1. Please see the detailed description of variables (tot, gdp, inf, ir, rer) in Section 3.1 for the baseline estimation and Section 3.3 for the robustness checks.

Figure 3.1 GDP growth and inflation correlation, 1990Q1-2016Q4

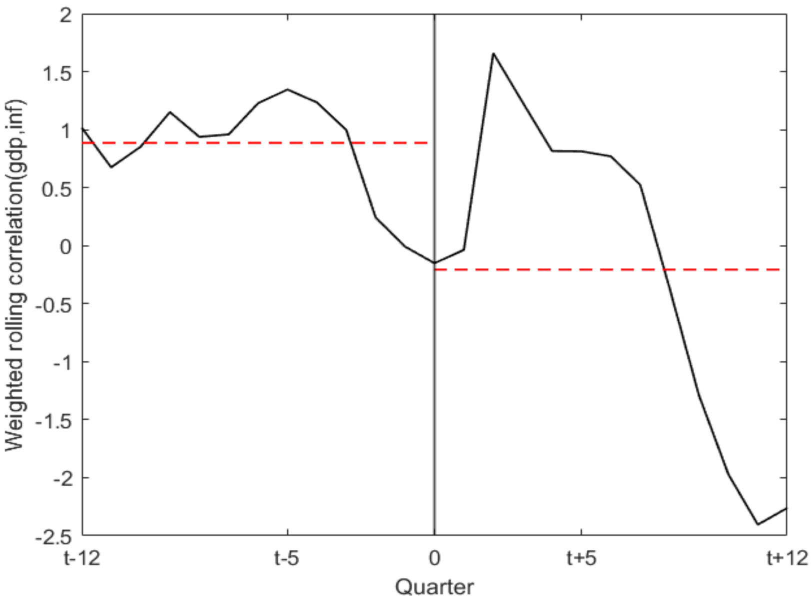
Negative correlation of GDP growth and inflation in the IT countries



Note: Bars: weighted correlation coefficients of year-on-year growth rates of GDP and CPI index by country over the 19990Q1-2016Q4 period; Dash lines: mean correlations by group. The GDP-weighted correlation coefficient of country i in group j is calculated by $weightedcorr_{i|j} = (\frac{GDP1990_i}{\sum_i GDP1990_{i|j}}) * corr_i$, where j = 1, 2 (either IT group or non-IT group), $corr_i$: (unweighted) correlation coefficient of GDP growth and inflation, $GDP1990_i$ in constant 2010 $US. The sample includes 23 IT countries and 69 non-IT countries. Croatia, Hungary, Kuwait, Libya, Moldova, Qatar, São Tomé and Principe are excluded due to missing data of GDP in 1990.

Figure 3.2 GDP growth and inflation correlation in the IT countries, 6-year window

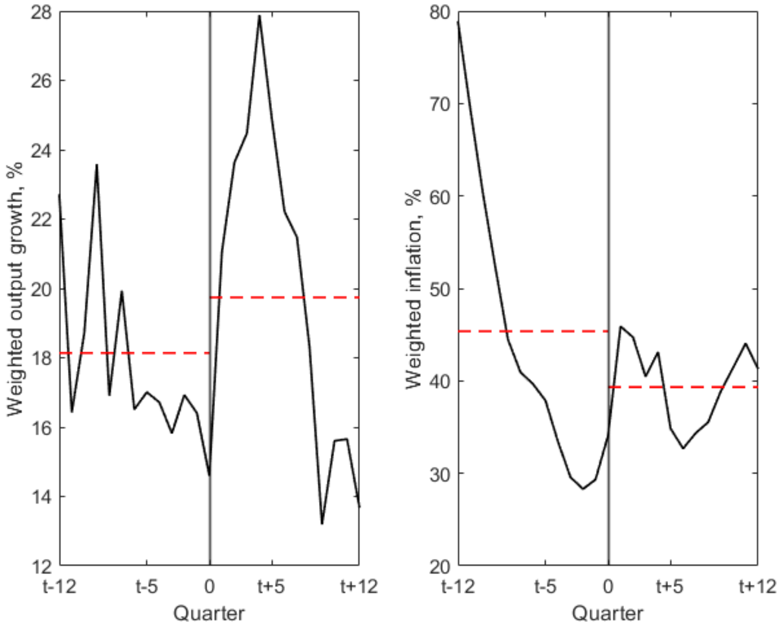
Decreasing correlation of GDP growth and inflation since IT adoption



Note: Solid line: GDP-weighted, 12-quarter rolling correlation coefficients of year-on-year growth rates of GDP and CPI index, averaged across 15 IT countries over three years before and three years after IT adoption; Dash lines: mean correlations. The starting quarter of IT adoption (*t = 0*) varies across countries. The weight is GDP in 1990 (in constant 2010 $US). The sample includes 15 IT countries with non-missing data for three years before and three years after IT adoption (Albania, Armenia, Brazil, Colombia, Ghana, Guatemala, Indonesia, Peru, Philippines, Poland, Romania, South Africa, Thailand, Turkey, and Uruguay).

Figure 3.3 GDP growth and inflation in the IT countries, 6-year window

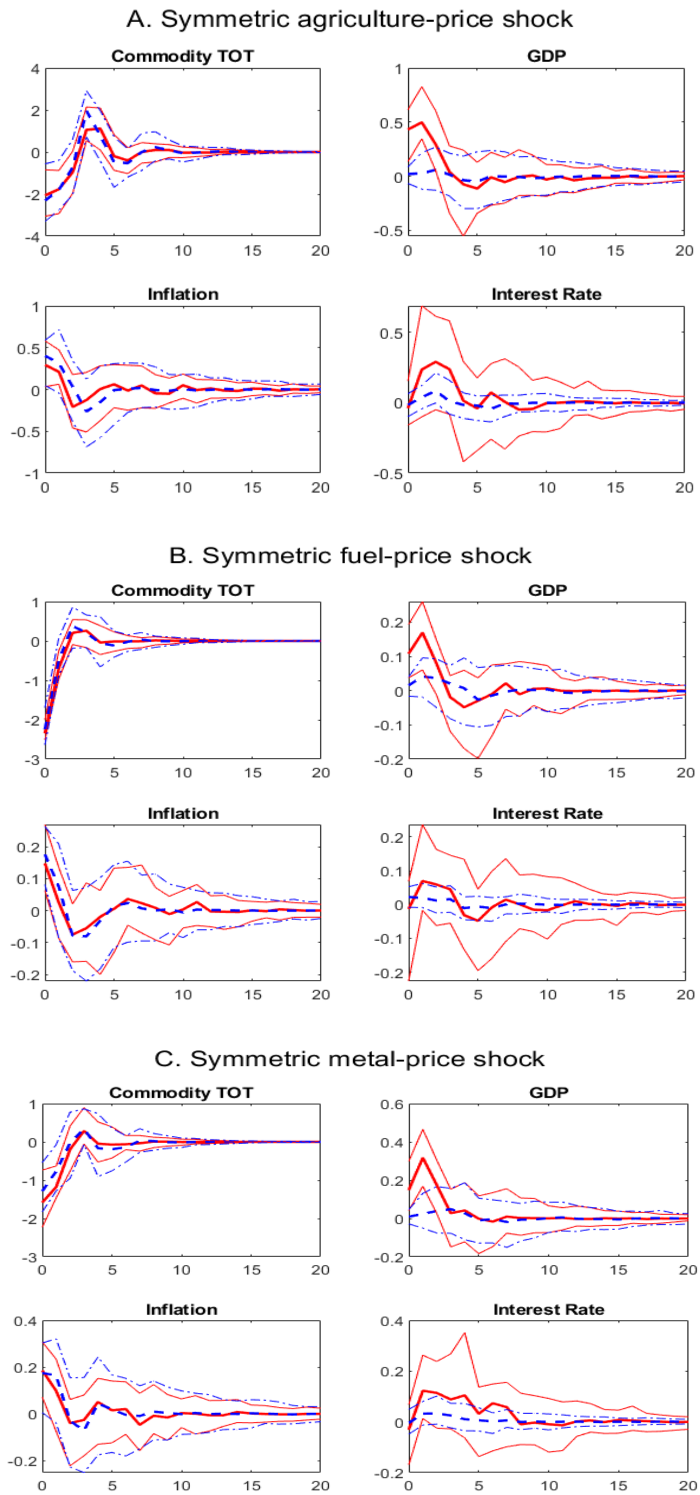
Increasing GDP growth and decreasing inflation since IT adoption



Note: Solid lines: GDP-weighted, year-on-year growth rates of GDP and CPI index, averaged across 18 IT countries, three years before and three years after IT adoption; Dash lines: mean growth rates. The starting quarter of IT adoption ($t = 0$) varies across countries. The weight is GDP in 1990 (in constant 2010 $US). This sample includes 18 IT countries with non-missing data for three years before and three years after IT adoption (Albania, Armenia, Brazil, Colombia, Dominican Republic, Ghana, Guatemala, Indonesia, Paraguay, Peru, Philippines, Poland, Romania, South Africa, Thailand, Turkey, Uganda, and Uruguay).

Figure 3.4 Impulse responses to symmetric commodity-price shocks

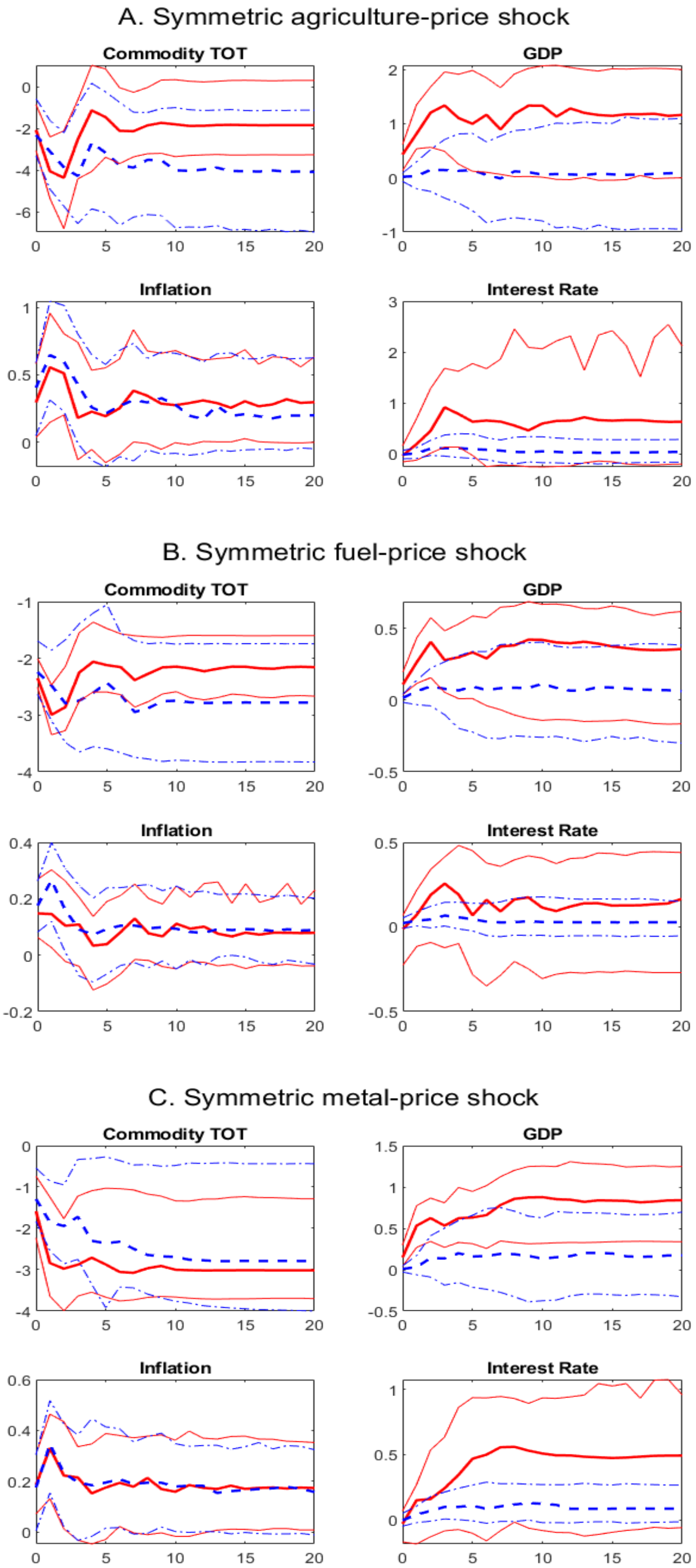
Increasing output growth in the IT countries, comparable inflationary effect across groups



Note: Solid lines: IT countries; Dash lines: Non-IT countries; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure 3.5 Cumulative impulse responses to symmetric commodity-price shocks

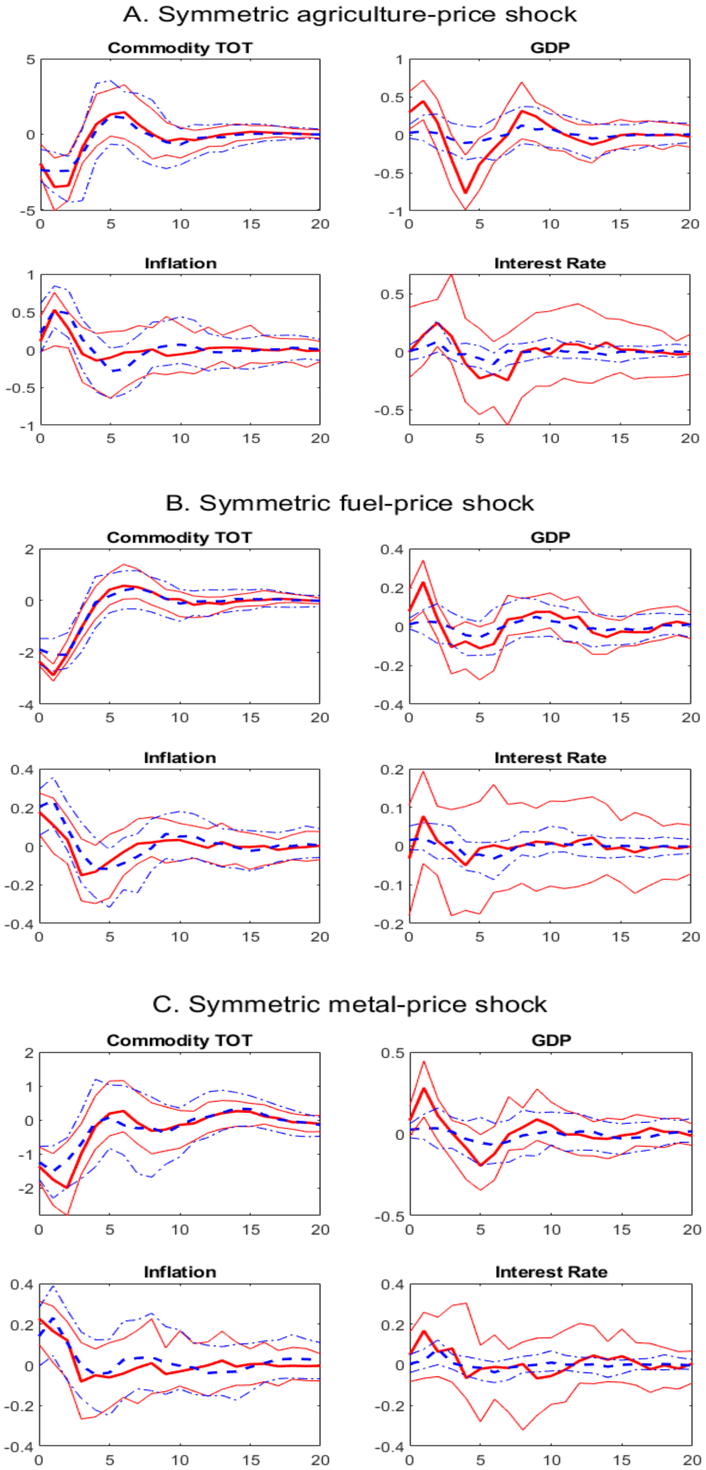
Persistent output growth in the IT countries, transitory inflationary effect for both groups



Note: Solid lines: IT countries; Dash lines: Non-IT countries; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure 3.6 Interannual data frequency: Impulse responses to symmetric commodity-price shocks

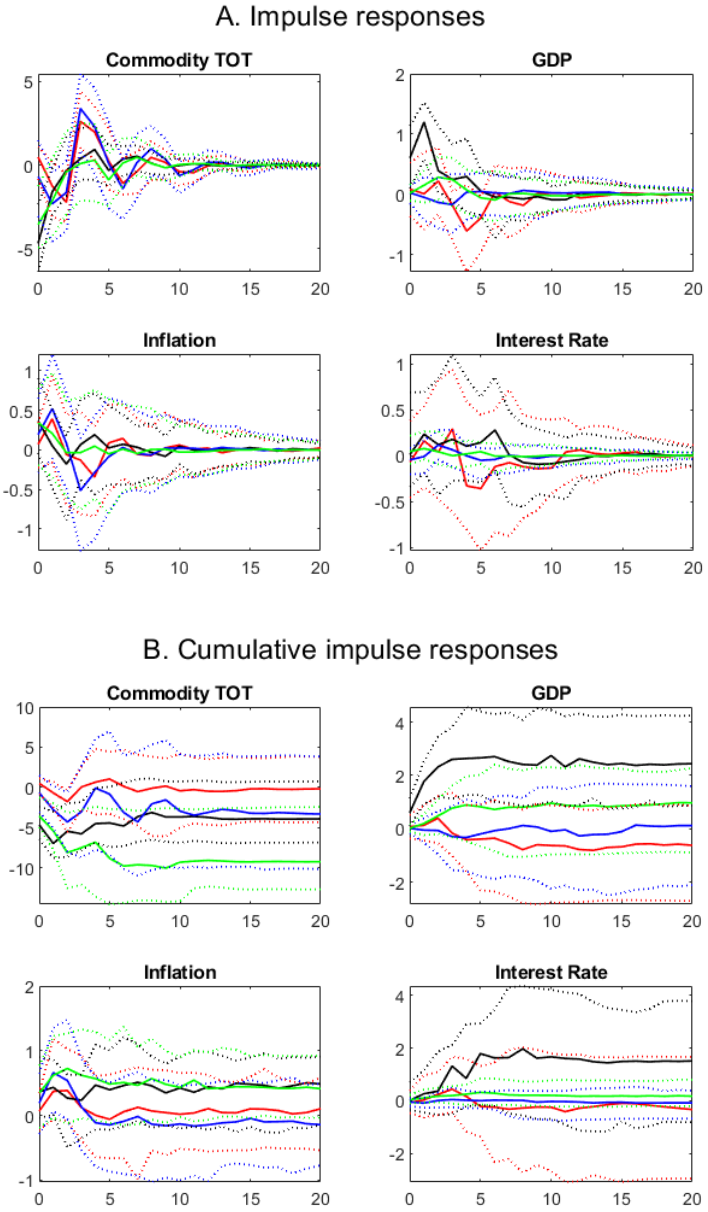
Increasing output growth in the IT countries, comparable inflationary effect for both groups



Note: Solid lines: IT countries; Dash lines: Non-IT countries; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure 3.7 Impulse responses to asymmetric agriculture-price shocks

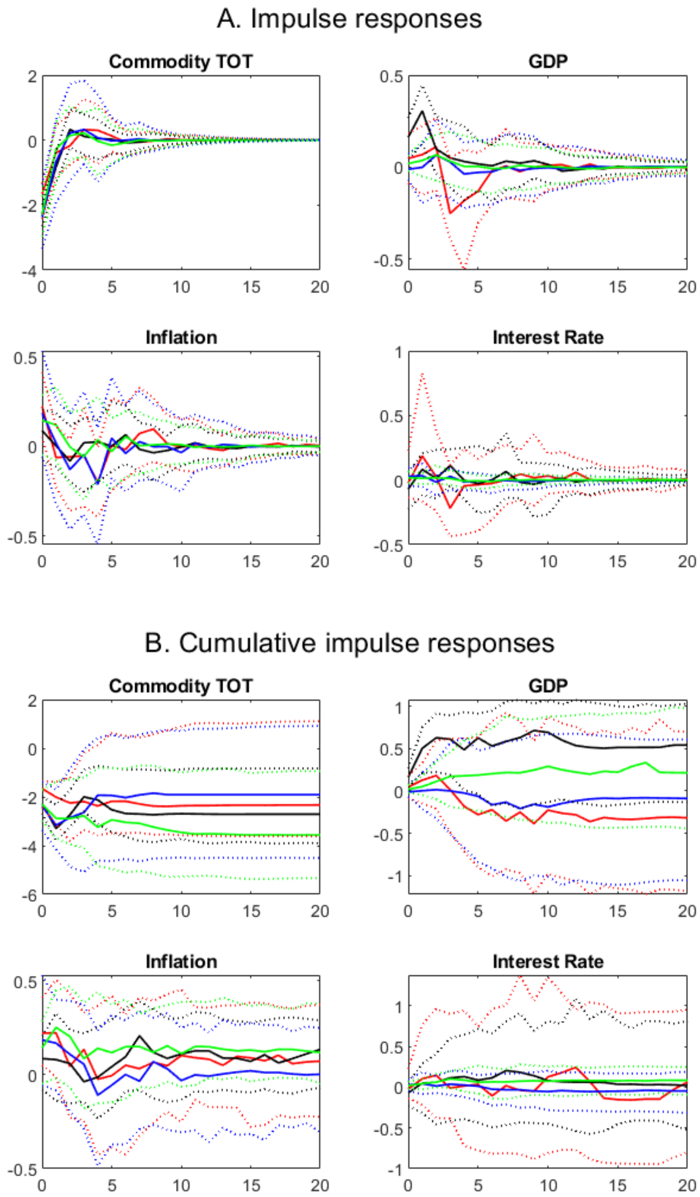
Persistent expansionary growth in the IT countries to the negative agriculture-price shock



Note: Red lines: (+) shock, IT countries; Black lines: (-) shock, IT countries; Blue lines: (+) shock, Non-IT countries; Green lines: (-) shock, Non-IT countries; Solid lines: impulse responses; Dot lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure 3.8 Impulse responses to asymmetric fuel-price shocks

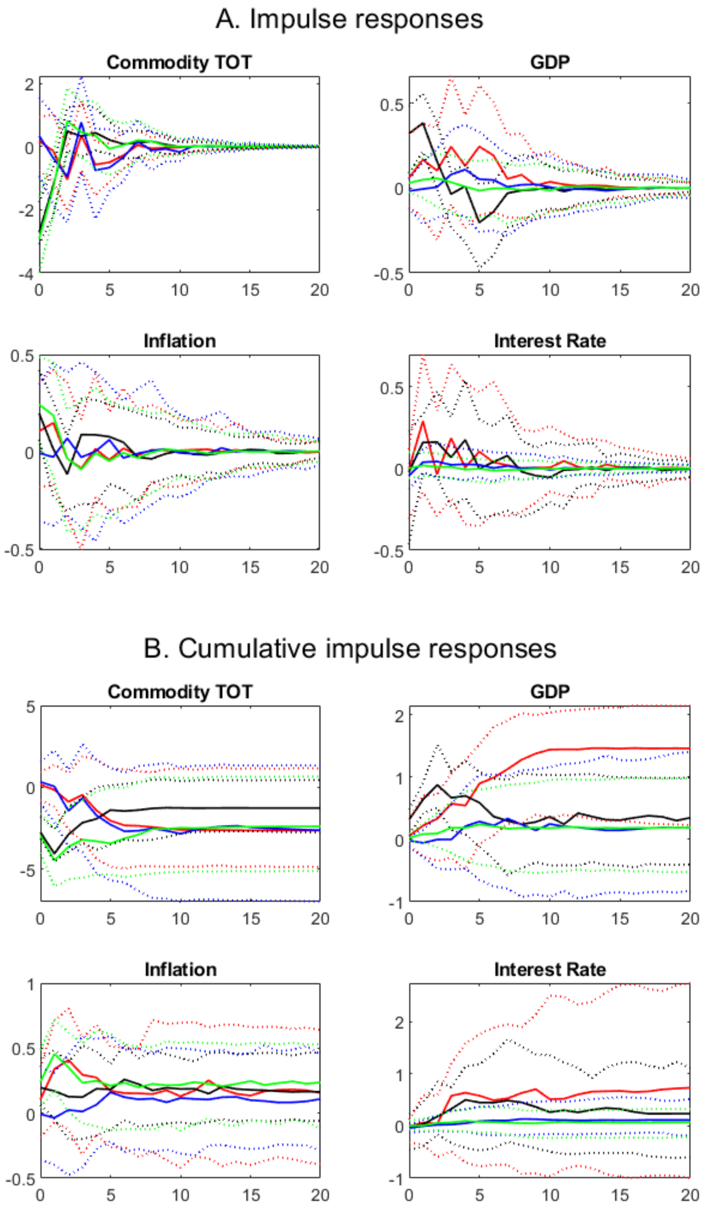
Expansionary growth to the negative shock and inflationary effect to the positive shock in the IT countries



Note: Red lines: (+) shock, IT countries; Black lines: (-) shock, IT countries; Blue lines: (+) shock, Non-IT countries; Green lines: (-) shock, Non-IT countries; Solid lines: impulse responses; Dot lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure 3.9 Impulse responses to asymmetric metal-price shocks

Persistent output expansion to both negative and positive shocks, temporary inflationary effect in the IT countries



Note: Red lines: (+) shock, IT countries; Black lines: (-) shock, IT countries; Blue lines: (+) shock, Non-IT countries; Green lines: (-) shock, Non-IT countries; Solid lines: impulse responses; Dot lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Appendix 3

A3.1 Sample, data source and variable computation

This section details the sampled countries, data source and computation of the domestic variables, focusing on interest rate, CPI, and real bilateral exchange rate.

Table A3.1 Sampled countries

No.	Country	Period	Observation	IT starting date	Commodity exporter/importer	Robustness check		Monetary policy instrument	
						Interannual data	Asymmetric shocks	Instrument	Source
1	Albania	2001q3-2016q4	62	2009q1	CIM/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
2	Algeria	1995q2-2016q4	87		CEX/AIM/FEX/MIM	included	included	Discount rate	MFS
3	Antigua and Barbuda	1998q3-2016q4	74		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
4	Argentina	1990q2-2013q4	95	2015q4	CEX/AEX/FEX/MEX	included	included	Central Bank policy rate	Datastream
5	Armenia	2000q2-2016q4	67	2006q1	CIM/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
6	Azerbaijan	1996q4-2016q4	81		CEX/AIM/FEX/MEX	included	included	Central Bank policy rate	MFS
7	Bahamas, The	1990q2-2016q4	107		CIM/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
8	Bangladesh	1994q1-2016q4	92		CIM/AIM/FIM/MIM	included	included	Central Bank policy rate	MFS
9	Barbados	1990q2-2016q4	107		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
10	Belarus	2000q2-2016q4	67		CIM/AIM/FIM/MIM	included	included	Central Bank policy rate	MFS
11	Benin	2002q1-2016q4	60		CEX/AEX/FIM/MIM		included	Central Bank policy rate	MFS
12	Bolivia	1996q3-2016q4	82		CEX/AEX/FEX/MEX	included		Discount rate	MFS
13	Botswana	2000q2-2016q4	67		CIM/AIM/FIM/MEX	included	included	Discount rate	MFS
14	Brazil	1999q3-2016q4	70	1999q3	CEX/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
15	Bulgaria	2005q2-2016q4	47		CIM/AEX/FIM/MEX	included		Central Bank policy rate	MFS
16	Burkina Faso	2002q1-2016q4	60		CIM/AEX/FIM/MIM	included	included	Central Bank policy rate	MFS
17	Burundi	1990q2-2016q4	107		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
18	Cameroon	1994q4-2016q4	89		CEX/AEX/FEX/MEX	included	included	Discount rate	MFS
19	Central African Republic	1994q4-2015q4	85		CEX/AEX/FIM/MEX	included	included	Discount rate	MFS
20	Chad	1994q4-2016q1	86		CEX/AIM/FEX/MIM	included	included	Discount rate	MFS
21	Chile	1995q4-2016q4	85	1990q3	CEX/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
22	China	1990q2-2016q4	107		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
23	Colombia	1995q3-2016q4	86	1999q4	CEX/AEX/FEX/MIM	included	included	Central Bank policy rate	MFS
24	Comoros	2000q3-2014q3	57		CIM/AIM/FIM/MIM		included	Discount rate	MFS
25	Congo, Rep.	1998q3-2016q4	74		CEX/AIM/FEX/MEX	included	included	Discount rate	MFS
26	Costa Rica	2006q3-2016q4	42		CEX/AEX/FIM/MIM			Central Bank policy rate	MFS
27	Croatia	1992q2-2014q1	88		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
28	Côte d'Ivoire	2002q1-2016q4	60		CEX/AEX/FEX/MIM			Central Bank policy rate	MFS
29	Dominica	1994q2-2016q4	91		CIM/AIM/FIM/MEX	included	included	Discount rate	MFS

30	Dominican Republic	2004q2-2016q4	51	2012q1	CIM/AIM/FIM/MIM	included		Central Bank policy rate	MFS
31	Egypt, Arab Rep.	1990q2-2016q4	107		CIM/AIM/FEX/MIM	included	included	Discount rate	MFS
32	Equatorial Guinea	1994q4-2016q4	89		CEX/AIM/FEX/MIM	included	included	Discount rate	MFS
33	Ethiopia	1990q2-2008q4	75		CIM/AEX/FIM/MIM	included	included	Treasury Bill rate	MFS
34	Gabon	1994q4-2016q4	89		CEX/AEX/FEX/MEX	included	included	Discount rate	MFS
35	Gambia, The	1990q2-2016q4	107		CIM/AIM/FIM/MIM	included	included	Central Bank policy rate	MFS
36	Ghana	1990q2-2016q4	107	2007q2	CEX/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
37	Grenada	1994q2-2016q4	91		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
38	Guatemala	2005q2-2016q4	47	2005q1	CEX/AEX/FIM/MIM	included		Central Bank policy rate	MFS
39	Guinea-Bissau	2002q1-2016q4	60		CEX/AEX/FIM/MIM		included	Central Bank policy rate	MFS
40	Guyana	1994q3-2016q4	90		CEX/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
41	Haiti	1997q1-2016q4	80		CIM/AIM/FIM/MIM	included	included	3-month Central Bank Bill rate	Datastream
42	Hungary	1991q2-2016q4	103	2001q2	CIM/AEX/FIM/MIM	included	included	Discount rate	MFS
43	India	1990q2-2016q4	107	2015q1	CIM/AEX/FIM/MIM	included	included	Discount rate	MFS
44	Indonesia	1990q2-2016q4	107	2005q3	CEX/AEX/FEX/MEX	included	included	Central Bank policy rate	Datastream
45	Iran, Islamic Rep.	2004q2-2016q4	51		CEX/AIM/FEX/MEX			Lending rate	MFS
46	Iraq	2005q3-2016q4	46		CEX/AIM/FEX/MIM			Central Bank policy rate	MFS
47	Jamaica	2002q2-2016q4	59		CIM/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
48	Jordan	2004q3-2016q4	50		CIM/AIM/FIM/MEX		included	Central Bank policy rate	MFS
49	Kenya	2006q4-2016q4	41		CEX/AEX/FIM/MEX			Central Bank policy rate	MFS
50	Kuwait	1992q1-2016q4	100		CEX/AIM/FEX/MIM	included	included	Discount rate	MFS
51	Kyrgyz Republic	2000q2-2016q4	67		CIM/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
52	Lao PDR	1993q2-2011q1	72		CEX/AEX/FIM/MEX	included	included	Discount rate	MFS
53	Lesotho	2000q2-2016q4	67		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
54	Libya	2001q3-2014q1	51		CEX/AIM/FEX/MIM	included	included	Discount rate	MFS
55	Madagascar	2003q1-2016q4	56		CEX/AEX/FIM/MEX	included	included	Treasury Bills rate	MFS
56	Malawi	1990q2-2016q4	107		CEX/AEX/FIM/MIM	included	included	Treasury Bills rate	MFS
57	Malaysia	2004q3-2016q4	50		CEX/AEX/FEX/MIM			Central Bank policy rate	MFS
58	Mali	2002q1-2016q4	60		CIM/AEX/FIM/MIM		included	Central Bank policy rate	MFS
59	Mauritania	1990q2-2004q1	56		CEX/AEX/FIM/MEX		included	Discount rate	MFS
60	Moldova	2000q2-2016q4	67	2012q4	CIM/AEX/FIM/MIM		included	Central Bank policy rate	MFS
61	Morocco	1998q2-2016q4	75		CIM/AIM/FIM/MEX		included	Discount rate	MFS
62	Mozambique	1996q3-2016q4	82		CEX/AIM/FIM/MEX	included	included	Discount rate	MFS
63	Myanmar	1991q2-2016q4	103		CEX/AEX/FEX/MEX	included	included	Discount rate	MFS
64	Namibia	1991q4-2016q4	101		CEX/AEX/FIM/MEX	included	included	Treasury Bills rate	MFS
65	Nepal	1992q1-2016q4	100		CIM/AIM/FIM/MIM	included	included	Central Bank policy rate	MFS
66	Niger	2002q1-2016q4	60		CEX/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
67	Pakistan	1994q2-2016q4	91		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
68	Paraguay	1991q1-2016q4	104	2011q2	CEX/AEX/FEX/MIM	included	included	Discount rate	MFS

69	Peru	2003q4-2016q4	53	2002q1	CEX/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
70	Philippines	2002q2-2015q4	55	2002q1	CIM/AIM/FIM/MIM		included	Discount rate	MFS
71	Poland	1998q2-2016q4	75	1998q3	CIM/AEX/FIM/MEX	included	included	Repurchase agreement rate	MFS
72	Qatar	2003q3-2016q4	54		CEX/AIM/FEX/MIM	included		Central Bank policy rate	MFS
73	Romania	1994q2-2016q4	91	2005q3	CIM/AIM/FIM/MEX	included	included	Discount rate	MFS
74	Russian Federation	1993q4-2016q4	93	2015q1	CEX/AIM/FEX/MEX	included	included	Central Bank policy rate	Datastream
75	Rwanda	1998q1-2016q4	76		CIM/AIM/FIM/MEX	included	included	Repurchase agreement rate	MFS
76	Saudi Arabia	1999q2-2016q4	71		CEX/AIM/FEX/MIM	included	included	Central Bank policy rate	MFS
77	Senegal	2002q1-2016q4	60		CIM/AIM/FIM/MEX	included	included	Central Bank policy rate	MFS
78	Seychelles	1990q2-2008q3	74		CIM/AEX/FIM/MIM	included	included	Discount rate	MFS
79	Solomon Islands	1990q2-2016q4	107		CEX/AEX/FIM/MIM	included	included	Treasury Bills rate	MFS
80	South Africa	1998q3-2016q4	74	2000q1	CEX/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
81	St. Kitts and Nevis	1994q2-2016q4	91		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
82	St. Lucia	1994q2-2016q4	91		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
83	St. Vincent and the Grenadines	1994q2-2016q4	91		CIM/AIM/FIM/MIM	included	included	Discount rate	MFS
84	Swaziland	2000q2-2016q4	67		CEX/AEX/FIM/MIM	included	included	Discount rate	MFS
85	São Tomé and Principe	1997q2-2016q4	79		CIM/AIM/FIM/MIM	included	included	Central Bank policy rate	MFS
86	Tanzania	2007q4-2016q4	37		CIM/AEX/FIM/MEX			Discount rate	MFS
87	Thailand	2000q4-2016q4	65	2000q2	CIM/AEX/FIM/MIM	included	included	Central Bank policy rate	MFS
88	Togo	2002q1-2016q4	60		CIM/AEX/FIM/MEX	included	included	Central Bank policy rate	MFS
89	Tonga	1990q2-2016q4	107		CIM/AIM/FIM/MIM	included	included	Lending rate	MFS
90	Trinidad and Tobago	1990q2-2016q4	107		CEX/AIM/FEX/MIM	included	included	Discount rate	MFS
91	Turkey	1990q2-2016q4	107	2006q1	CIM/AEX/FIM/MIM	included	included	Central Bank policy rate	MFS
92	Uganda	1992q4-2015q4	93	2011q2	CIM/AEX/FIM/MIM	included	included	Discount rate	MFS
93	Ukraine	1994q1-2013q1	77	2016q4	CIM/AEX/FIM/MEX	included	included	Discount rate	Datastream
94	Uruguay	1990q2-2016q4	107	2007q3	CEX/AEX/FIM/MIM	included	included	Discount rate	MFS
95	Vanuatu	1999q2-2016q4	71		CIM/AIM/FIM/MIM		included	Discount rate	MFS
96	Venezuela, RB	1990q2-2016q4	107		CEX/AIM/FEX/MEX	included	included	Discount rate	MFS
97	Vietnam	1996q2-2016q4	83		CEX/AEX/FEX/MIM	included	included	Central Bank policy rate	MFS
98	Yemen, Rep.	2001q3-2014q2	52		CEX/AIM/FEX/MIM		included	Treasury Bills rate	MFS
99	Zambia	2000q2-2016q4	67		CEX/AEX/FIM/MEX	included	included	Discount rate	MFS

Note: Table lists the 99-country sample in the baseline model. Each country is classified to be an inflation targeting country (IT, with according starting date) or non-IT, commodity net exporter (importer) – CEX (CIM), agriculture net exporter (importer) – AEX (AIM), fuel net exporter (importer) – FEX (FIM), metal net exporter (importer) – MEX (MIM). Columns (7) and (8) lists the countries included in each robustness check. The last two columns lists the specific monetary instruments used for each country with the according data source.

A3.1.1 Interest rate

I start with International Financial Statistics Yearbook (2018) to determine the Central Bank policy rate, which is the key underlying financial instrument, and the second-best series for each sampled country. I use the Central Bank policy rate, namely *Monetary policy-related interest rate* in Monetary and Financial Statistics (MFS). Despite named commonly as *Monetary policy-related interest rate*, the key financial instrument varies across countries. For instance, the Central Bank policy rate in Albania is the main policy rate on weekly repurchase agreements while in Azerbaijan, that is the basic six-month rate at which the Central Bank of the Republic of Azerbaijan lends to commercial banks. Readers refer to the Yearbook for more details of various instruments for each country. For the countries whose data of the Central Bank policy rate is not available from MFS, I replace their data by the higher-ordered referenced instrument stated in the Yearbook with data available from MFS. Take Algeria as an example, I proceed with their discount rate as the highest-ordered interest rate in the Yearbook. The MFS report data on Lending rate, Treasury Bills rate, Discount rate, Repurchase Agreement rate, and Government Bond equivalent rate. For the exceptional cases of Argentina, Haiti, Indonesia, Russian Federation, and Ukraine, whose Central Bank policy rates are either missing from MFS or too short and the second-best alternatives are unavailable in MFS, I collect their interest rates from Datastream. The specific interest rates series (quarterly, percent per annum) for 99 countries in the benchmark specification are given in the following table. The first-difference of quarterly interest rate is used in SVAR estimation.

A3.1.2 Consumer Price Index

I collect quarterly Consumer Price Index (CPI) (IFS coded PCPI_IX) for 95 countries from International Financial Statistics (IFS) and monthly CPI for Argentina, Mozambique, and Namibia from Datastream because their data on CPI is missing from IFS. For Venezuela, its shorter CPI series from IFS (2008Q1-2016Q4) is fully replaced by the longer CPI series from Datastream (1990Q1-2016Q4). Data on CPI from both sources are non-seasonally adjusted and 2010-based. The monthly CPI series from Datastream are averaged by quarter to be in quarterly frequency. Note that the X-13ARIMA-SEATS method to remove seasonal components strictly requires the series to be continuous. For the country with discontinuous CPI series, I keep the longer CPI sequence and drop the shorter one. To avoid ending up with no observation after merging the seasonally-adjusted CPI with other domestic variables, I cross-check the data availability for other domestic variables in advance and keep the continuous CPI series for Lesotho since 1998Q2, Libya since 2001Q1, Rwanda since 1997Q3, and Togo since 1995Q3 although their CPI series before those time points are longer. The last steps are to remove the seasonal components in CPI series for each country using the X-13ARIMA-SEATS method, the newest seasonal adjustment software developed by the United States Census Bureau[34], and then calculate the quarter-on-quarter inflation rate (i.e. the percentage change of seasonally-adjusted CPI to the previous quarter). The deseasonalisation is done using the R package *seasonal*.

In the robustness check using interannual data frequency (i.e. year-on-year difference or year-on-year percentage change), the year-on-year inflation rate is computed using quarterly, non-seasonally adjusted, 2010-base year CPI series (i.e. the continuous CPI series before being treated by X-13ARIMA-SEATS). The seasonal components are automatically removed by construction.

A3.1.3 Real bilateral exchange rate

Real bilateral exchange rate against $US (*rer*) of country *i* at quarter *t* is computed as

$$rer_{i,t} = ner_{i,t} * \frac{CPI_{US,t}}{CPI_{i,t}}$$

where *ner* - norminal bilateral exchange rate, CPI series of the US and the emerging countries are mainly sourced from IFS, with the latter two are 2010-based and non-seasonally adjusted, except for the Datastream-sourced data of CPI for Argentina, Mozambique, Namibia, and Venezuela as stated above. The monthly CPI series from Datastream are averaged by quarter to be in quarterly frequency.

A3.2 Unit root tests and optimal lag length

I use the Augmented Dickey-Fuller (ADF) test and the Phillips-Perron (PP) test to check the stationarity of the time-series variables by country. I conduct the tests for each series in their level (including either intercept or both intercept and time trend) and in first-difference (including neither intercept nor time trend in the ADF test and including only intercept in the PP test).

[34] https://www.census.gov/srd/www/x13as/

Table A3.2 Frequency of optimal lag lengths

Lag	AIC	SC	HQ	FPE
1	4	52	25	8
2	3	4	10	9
3	1	.	5	14
4	3	.	.	7
5	2	2	2	2
6	4	4	4	8
7	7	7	7	4
8	6	6	6	6
9	17	10	16	15
10	52	14	24	26
Total	99	99	99	99

Note: AIC: Akaike Information Criterion, HQ: Hannan-Quinn Criterion, SC: Schwarz Criterion, FPE: Final Prediction Error Criterion.

At 5% level of significance, the ADF tests for the level-series including intercept only indicate 4 non-stationary series for Ethiopia (*m*), Jordan (*gdp*, *ir*), and Qatar (*ir*). While the global commodity price indices are common for all countries, that only the growth rate of global metal price is shown to be non-stationary for Ethiopia indicates the imperfection of the ADF test whose result is affected by the sample size. In the ADF tests for level-series including both intercept and time trend indicate 9 non-stationary series for Albania (*ir*), Bulgaria (*gdp*, *ir*), Ethiopia (*m*), Iraq (*ir*, *rer*), Jordan (*gdp*, *ir*), and Qatar (*ir*). The ADF tests for the series in their first-difference imply that all first-differenced series are stationary.

The PP tests for the level-series including intercept only suggest 16 non-stationary series of *gdp* for Antigua and Barbuda, Armenia, Bangladesh, Benin, Burundi, Côte d'Ivoire, Iraq, Jamaica, Jordan, Lesotho, Swaziland, Tanzania, Togo, Trinidad and Tobago, Yemen, Rep., and Zambia. The PP tests for the level-series including both intercept and time trend indicate 32 non-stationary series: Antigua and Barbuda (*gdp*), Armenia (*gdp*), Bahamas (*gdp*), Bangladesh (*gdp*), Barbados (*gdp*), Benin (*gdp*), Burundi (*gdp*), Chad (*gdp*), Côte d'Ivoire, Ethiopia (*a*, *m*), Iraq (*gdp*, *ir*), Jamaica (*gdp*), Jordan (*gdp*), Kenya (*ir*), Lesotho (*gdp*), Mali (*gdp*), Mauritania (*gdp*), Pakistan (*gdp*), Seychelles (*gdp*), Solomond Islands (*gdp*), Swaziland (*gdp*), Tanzania (*gdp*), Togo (*gdp*), Trinidad and Tobago (*gdp*), Yemen (*gdp*), and Zambia (*gdp*) The PP tests for the series in their first-difference suggest that all first-differenced series are stationary.

Favoring the ADF test result and providing that the possible non-stationary series have no impact on the stability of SVAR system, I then proceed with the SVAR estimations. Detailed results of the ADF and the PP unit root test for all eight variables by country are provided upon request.

For the optimal lag lengths, I use a range of criteria such as the Akaike Information Criterion (AIC), Hannan-Quinn Criterion (HQ), Schwarz Criterion (SC), and Final Prediction Error Criterion (FPE), with the maximum lags of 10. A summary of the frequency for the optimal lag lengths are given in Table A3.2. A detailed results for each country in the sample are provided upon request.

A3.3 GDP-weighted impulse responses

Table A3.3 GDP-weighted impulse responses to the symmetric global commodity-price shocks

	Agriculture-price shock		Fuel-price shock		Metal-price shock	
	IT	Non-IT	IT	Non-IT	IT	Non-IT
Short-term effect						
tot	-0.25	-0.56	-0.39	-0.48	-0.24	-0.36
gdp	0.14	0	0.08	0.01	0.05	0
inf	0.09	0.2	0.06	0.12	0.14	0.04
ir	-0.01	0	0	0	-0.01	0
rer	-1	-1.64	-0.18	-0.16	-0.45	-0.09
Cumulative one-year effect						
tot	-0.65	-0.5	-0.42	-0.47	-0.3	-0.43
gdp	0.42	0.07	0.11	0.01	0.35	0.3

inf	0.08	0.13	0	0.02	0.08	0.17
ir	0.19	0.08	0.06	0.03	0.04	0.06
rer	-1.81	-2.97	-0.32	-0.33	-0.91	-0.11

Note: *tot*: quarterly growth rate of commodity terms-of-trade; *gdp*: quarterly growth rate of GDP; *inf*: first difference of quarterly inflation; *ir*: first-difference of central bank policy rate; *rer*: quarterly growth rate of real bilateral exchange rate.

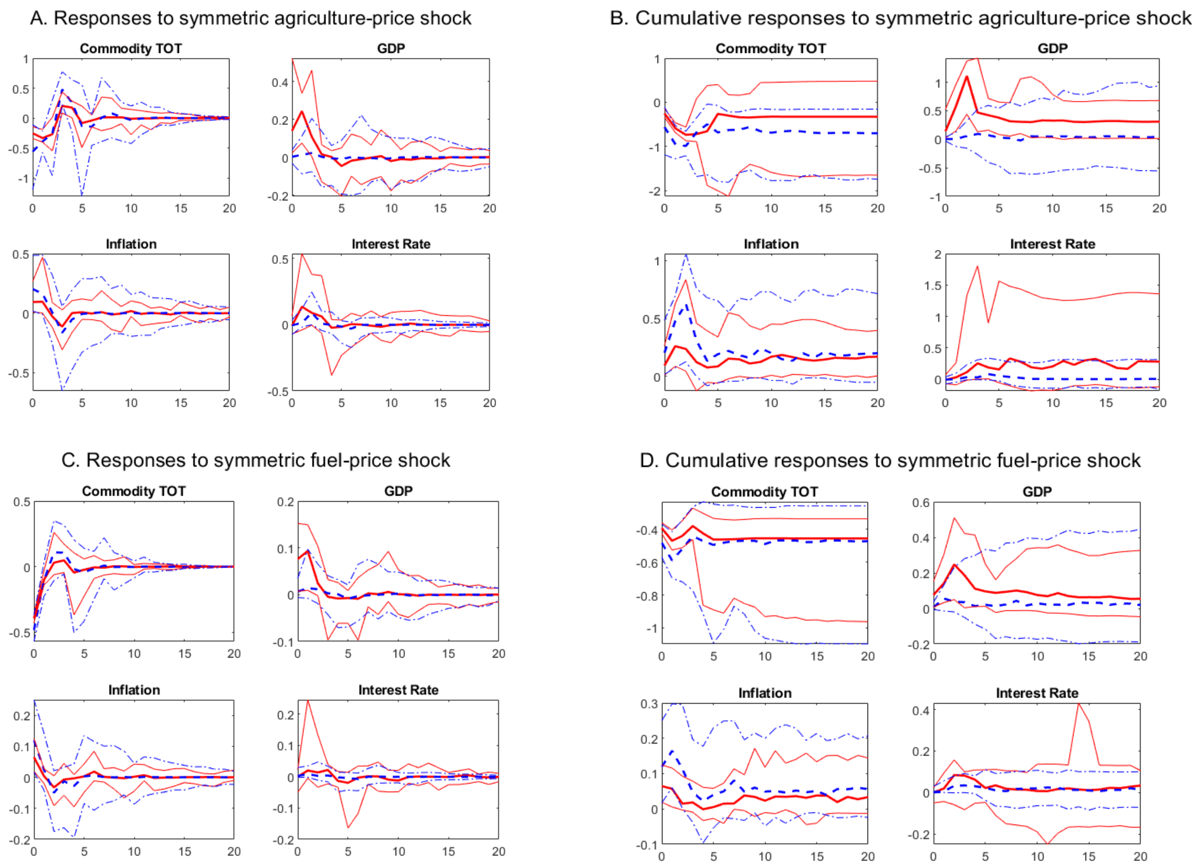
I also cross-check the impulse responses of domestic variables for the IT and non-IT countries to the global commodity-price shocks while controlling for the countries' development levels. I weight the impulse responses in the baseline estimation by country-specific constant GDP of the beginning year of the studied period (1990), which is measured in 2010 $US. Using GDP-weighted impulse responses, I find a consistent result with the benchmark conclusion that the IT group displays persistent improvements in GDP growth than the non-IT group does to the global commodity-price shocks.

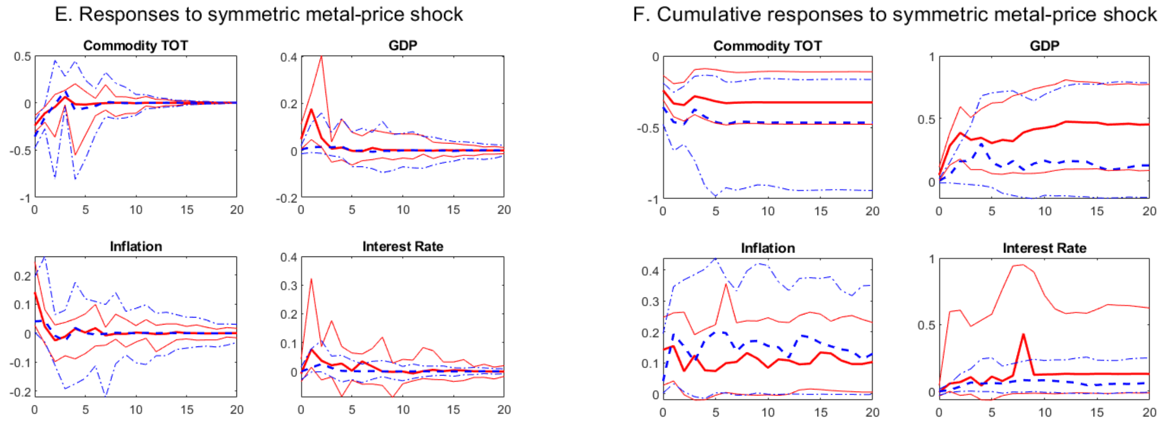
Table A3.3 summarizes the GDP-weighted contemporaneous and cumulative one-year effects of 10% increase in the global commodity-price shocks on the domestic variables across the IT and non-IT countries. The sample includes 85 EM countries excluding seven large commodity exporters and importers (China, India, Russia, Saudi Arabia, Venezuela, Brazil, and Chile) and other seven EME countries without data on constant GDP in 1990 (Croatia, Hungary, Kuwait, Libya, Moldova, Qatar, and São Tomé and Principe).

Figure A3.1 reports the weighted impulse responses in the left panel and the weighted cumulative impulse responses in the right panel.

Figure A3.1 GDP-weighted responses to symmetric global commodity-price shocks

Persistent improvement in GDP growth for the IT countries, transitory effect on inflation for both groups





Note: Solid red lines: IT countries; Dash blue lines: Non-IT countries; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Global agriculture-price shock

As seen in panel A, commodity terms-of-trade of the IT and non-IT groups decline by 0.25% and 0.56%, respectively. Only the former group displays a 0.14% improvement in their GDP growth immediately. Though the inflationary effect is present in both groups, the non-IT countries suffer a greater impact (0.2%) compared to the IT countries (0.09%).

Regarding the cumulative impulse responses shown in panel B, the cumulative fall in commodity terms-of-trade persists in the non-IT group following the global agriculture-price shock. Meanwhile, the negative effect on commodity terms-of-trade for the IT countries ends 2 quarters after the shock. The IT group also displays a persistent expansion of output growth. The effects on inflation are temporary in both groups, ending 2 quarters following the shock.

Global fuel-price shock

As shown in panel C, commodity terms-of-trade falls immediately by 0.39% in the IT countries and 0.48% in the non-IT countries as responding to a 10% improvement in the global fuel price. The IT group also display a marginal increase (0.08%) in GDP growth. The inflationary effect is also higher in the non-IT countries (0.12%) compared to the IT countries (0.06%).

Panel D reveals the cumulative responses to the fuel-price shock in each group: (1) commodity terms-of-trade declines persistently in both groups; (2) the IT countries' GDP growth lasts for 5 quarters after the shock; (3) both groups experience a transitory inflationary effect, lasting up to two quarters following the shock and ending 1-quarter earlier in the IT countries than in the non-IT countries.

Global metal-price shock

Panel E shows that the IT group and non-IT group suffer from a contemporaneous fall in commodity terms-of-trade by 0.24% and 0.36%, respectively. GDP growth in the former group goes up slightly by 0.05%. Their inflation also increases by 0.14% immediately following the metal-price shock while that effect in the non-IT countries is marginal (0.04%).

As seen in panel F, commodity terms-of-trade drops persistently in both groups. The positive response of GDP growth is persistent in the IT countries. The inflationary effect in the non-IT countries ends 2 quarters after the shock, one-quarter later than in their IT counterparts.

A3.4 Responses of commodity net exporters versus importers to global commodity-price shocks

This section details the impulse responses of commodity terms-of-trade, GDP growth, and inflation to the 10% increases in global commodity prices for the following groups: (i) the commodity net exporters vis-à-vis the commodity net importers; (ii) the agriculture net exporters vis-à-vis the agriculture net importers; (iii) the fuel net exporters vis-à-vis the fuel net importers; and (iv) the metal net exporters vis-à-vis the metal net importers. To define a (an) commodity (agriculture/fuel/metal) net exporter (importer), I use the country-specific median of commodity trade balance to GDP over the 1990-2016 period: country is classified as a (an) commodity (agriculture/fuel/metal) net exporter (importer) if its median commodity (agriculture/fuel/metal) trade balance to GDP is positive (negative). Fernández et al. (2017) instead use the dollar amount of commodity trade balance to define a commodity net exporter (importer); I find that the two approaches result in similar country groupings. The data on primary commodities comes from WDI and UNCTAD, consistent with World Bank Pink Sheets. I report here the results from the baseline model for the 92-country sample excluding China, India, Russia, Saudi Arabia, Venezuela, Brazil, and Chile. I also compare the baseline results with those from the robustness checks using interannual data frequency (74-country sample) and examining the asymmetric shocks (80-country sample). Most of the responding patterns described below remain in the full sample (99 countries in the baseline estimation, 81 countries in the robustness check using internannual data frequency, and 87 countries in the robustness check for asymmetric shocks).

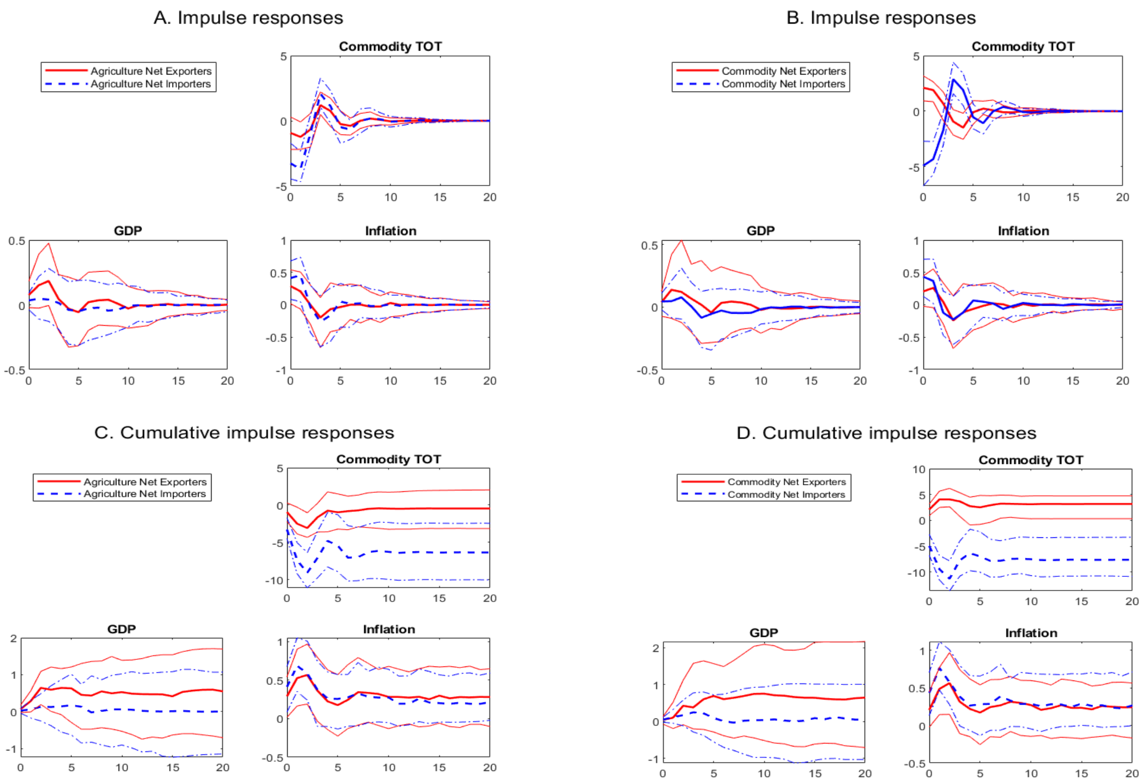
Overall, grouping the EME countries by commodity trade balance to GDP shows a strong association between the global commodity prices and the domestic commodity terms-of-trade. I find that an increase in the global price of agriculture (fuel/metal) leads to an improvement in the commodity terms-of-trade for the agriculture (fuel/metal) net exporters but the opposite for the agriculture (fuel/metal) net importers. Similarly, the global commodity-price shocks increase the commodity terms-of-trade of the commodity net exporters but reduce the commodity terms-of-trade of the commodity net importers. I also find that the commodity (agriculture/fuel/metal) net importers are more vulnerable to the global commodity-price shocks than the commodity (agriculture/fuel/metal) net exporters in terms of lower GDP growth, as well as higher and persistent inflation. The analysis of asymmetric shocks reveals that the domestic variables are more responsive to the negative shocks than to the positive shocks. Following the negative agriculture-price shock, the effects on commodity terms-of-trade and GDP growth are positive and persistent, while inflation tends to be lower for the commodity net exporter than for the net importers.

A3.4.1 Baseline estimation

Global agriculture-price shock

As seen in Figure A3.2, panel A, the 10% increase of the global agriculture price leads to a negligible decrease in commodity terms-of-trade of the agriculture net exporters but a large contraction of 3.26% in the same indicator of the agriculture net importers. GDP growth rate of the former group displays a modestly larger expansion than that of the latter group. Inflation in the agriculture net exporters increases by 0.29%, slightly smaller than that in their importing counterparts (0.41%). In panel B, the responses across the commodity net exporters and the commodity net importers are noticeably contradictory. In the commodity net exporters, commodity terms-of-trade increases by 2.11% whereas GDP growth and inflation remain silent. The commodity net importers, however, experience a 4.9% decline in commodity terms-of-trade, marginal responses of GDP growth, and a 0.43% increase in inflation.

Figure A3.2 Impulse responses to a symmetric agriculture-price shock



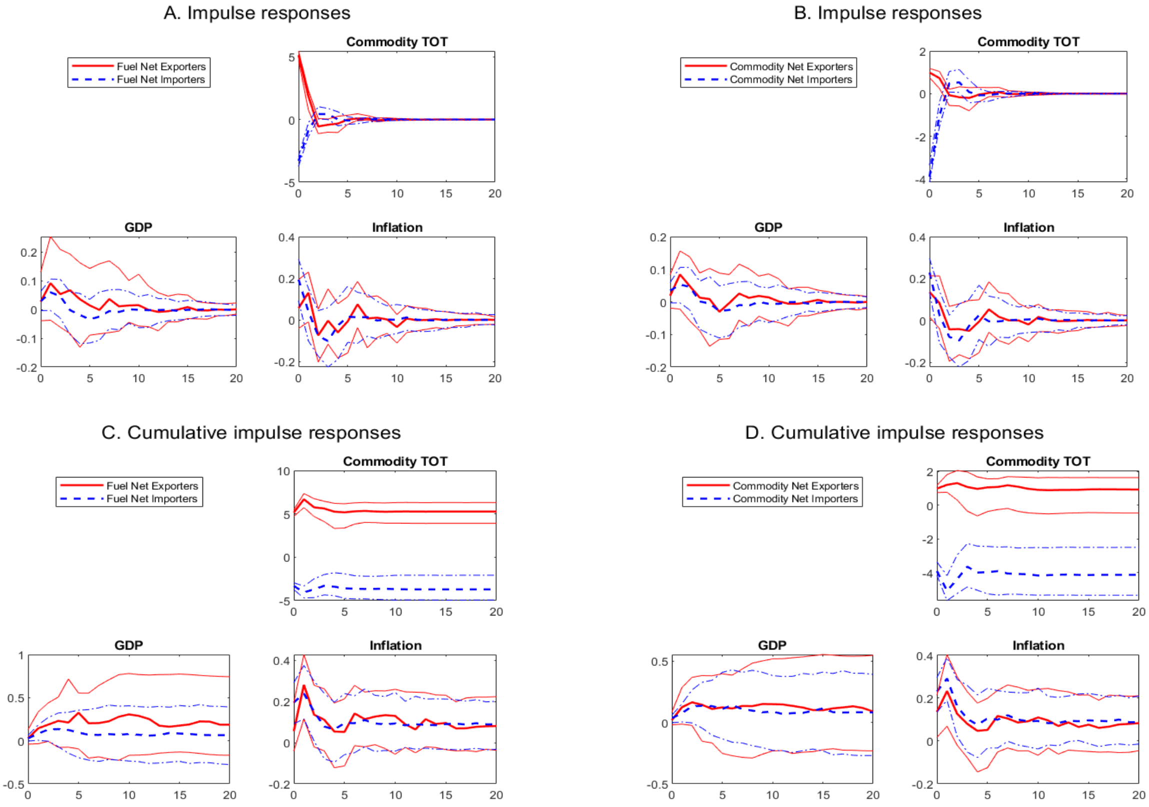
Note: Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure A3.2, panels C-D show the transitory impacts on both GDP growth and inflation across the groups. The accumulated responses of GDP growth in the agriculture net exporters remain positive at 0.6% during quarters 2-3 while those in the agriculture net importers are slight. The cumulative inflationary effects are similar for the agriculture net exporters and the agriculture net importers with a temporary increase of 0.6% ending at quarter 2. The commodity net exporters and the commodity net importers display small cumulative responses of GDP. The positive cumulative inflationary effects remain for quarters 1-2 for the commodity net exporters but remain for the first three quarters for the commodity net importers; both hiting 0.6% at the end of quarter 2 and ending in the following quarter.

Global fuel-price shock

In Figure A3.3, panel A, a 10% improvement in the global fuel price leads to a 5.15% increase in commodity terms-of-trade of the fuel net exporters but a 3.31% decrease in the fuel net importers. The fuel net importers also show an increase of 0.2% in inflation. There is no inflationary effect in the fuel net exporters, however. I also find marginal changes in GDP growth across those two groups. In panel B, the commodity net exporters display a 1% increase in their commodity terms-of-trade when the commodity net importers suffer a 3.91% shrinkage in the same indicator. The latter group also shows a higher inflationary effect (0.23%) than that in the former group (0.14%). I find negligible responses in GDP growth of both groups to the global fuel-price shock.

Figure A3.3 Impulse responses to a symmetric fuel-price shock



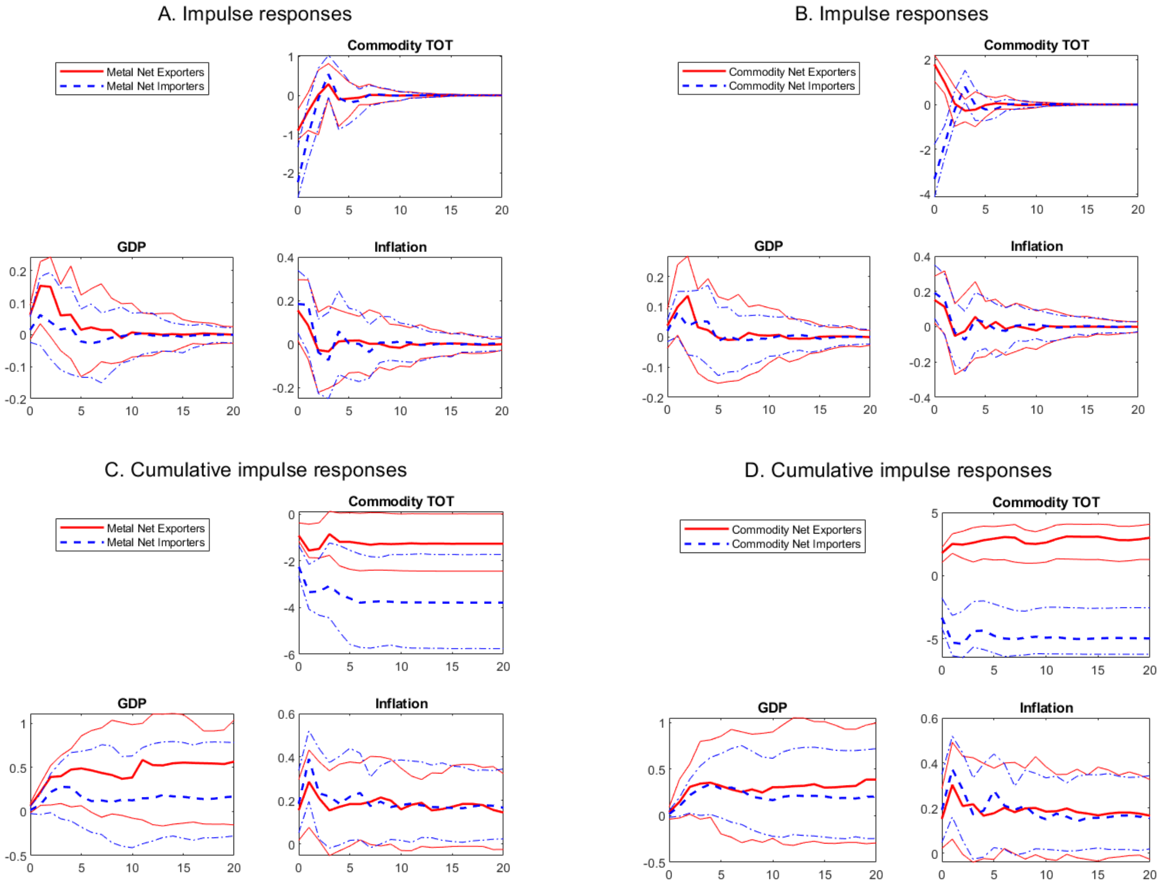
Note: Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Regarding the cumulative responses in Figure A3.3, panels C-D, I do not find any noticeable changes in GDP growth. The inflation responses of the fuel net exporters and the fuel net importers hit their spikes at almost the same level at 0.3% in quarter 1. The cumulative inflationary effect in the fuel net importers lasts until quarter 3. The commodity net exporters experience a transitory accumulated inflation increase, being 0.23% at the end of quarter 1 while the commodity net importers display a comparable but 1-quarter longer-lasting increase in their cumulative inflation response.

Global metal-price shock

As seen in Figure A3.4, panel A, a 10% increase in the global metal price causes commodity terms-of-trade to contract in the metal net exporters and the metal net importers by 0.91% and 2.24%, respectively. The contemporaneous improvement in GDP growth in the former group is marginally larger than that in the latter group. Two groups share the same inflationary effects (0.2%). In panel B, the global metal-price shock leads the commodity terms-of-trade to increase by 1.79% in the commodity net exporters but decrease by 3.33% in the commodity net importers. GDP growth slightly expands by 0.1% equivalently across two groups with a one-quarter delay. The inflation responses are comparable in two groups (0.2%).

Figure A3.4 Impulse responses to a symmetric metal-price shock



Note: Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using bootstrapping method. The x-axis measures quarters after the shock.

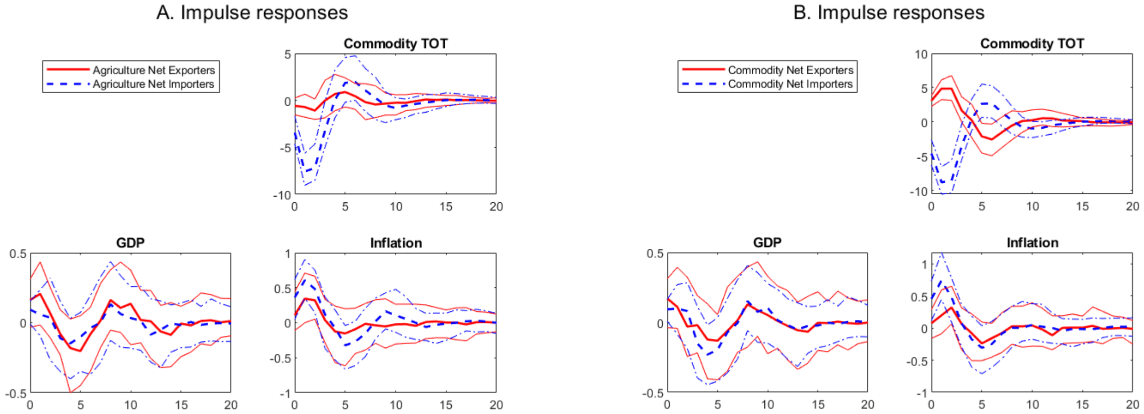
Figure A3.4, panel C shows that the cumulative responses of GDP growth of the metal net exporters are positive for 5 consecutive quarters after the metal-price shock whereas those of the importers are invisible. The one-year cumulative increase in GDP growth of the former group is 0.47%. The accumulated responses of inflation appear similar in two groups, remaining positive before ending up with 0.2% at the end of quarter 2. In Figure A3.4, panel D, the cumulative improvement in GDP growth of the commodity net exporters in quarter 2 is 0.31% while that of the commodity net importers are 0.22% - 0.34% during quarters 2 - 4. The cumulative responses of inflation in the two groups appear similar, both hitting 0.3% - 0.4% in quarter 1. The inflationary effect in the commodity net exporters becomes marginal since quarter 2 but takes the commodity net importers one more quarter to revert to their initial inflation level.

A3.4.2 Interannual data frequency

As seen in Figure A3.5, to the 10% improvement in the global agriculture price, the agriculture net importers experience a 3.43% decrease in commodity terms-of-trade and a 0.36% increase in inflation (panel A). In panel B, the commodity net importers show a 4.57% decrease in commodity terms-of-trade, a slight increase of 0.09% in GDP growth, and a 0.46% increase in inflation. I also find a 3.08% increase in commodity terms-of-trade in the commodity net exporters.

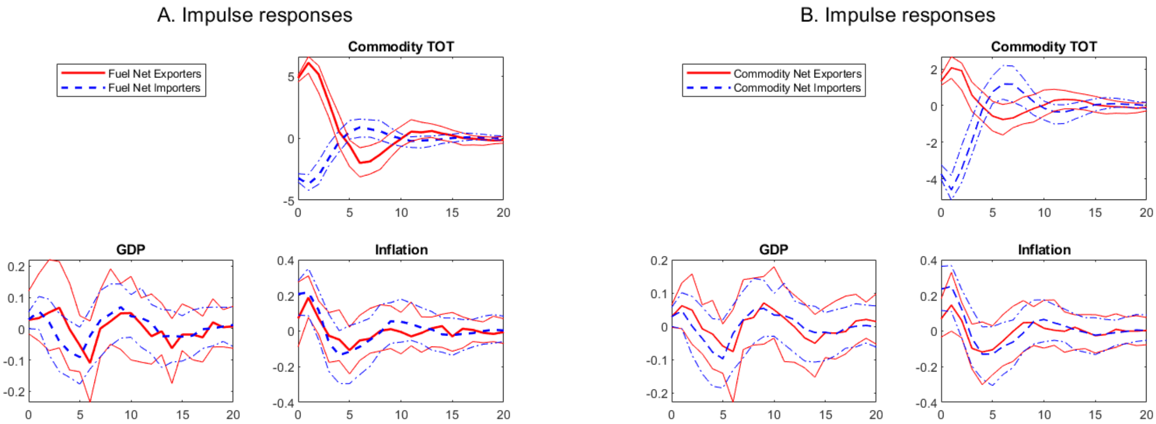
In Figure A3.6, panel A, the 10% increase in the global fuel price leads to a 4.84% contemporaneous improvement in commodity terms-of-trade and a 0.19% increase in inflation, with one-quarter lag, in the fuel net exporters. The fuel net importers, however, suffer from a 3.21% contraction in commodity terms-of-trade and a 0.21% contemporaneous inflationary effect. As shown in panel B, commodity terms-of-trade increases by 1.32% in the commodity net exporters but decreases by 3.73% in the commodity net importers. The latter group also displays an inflationary impact of 0.24%.

Figure A3.5 Interannual data frequency: Impulse responses to a symmetric agriculture-price shock



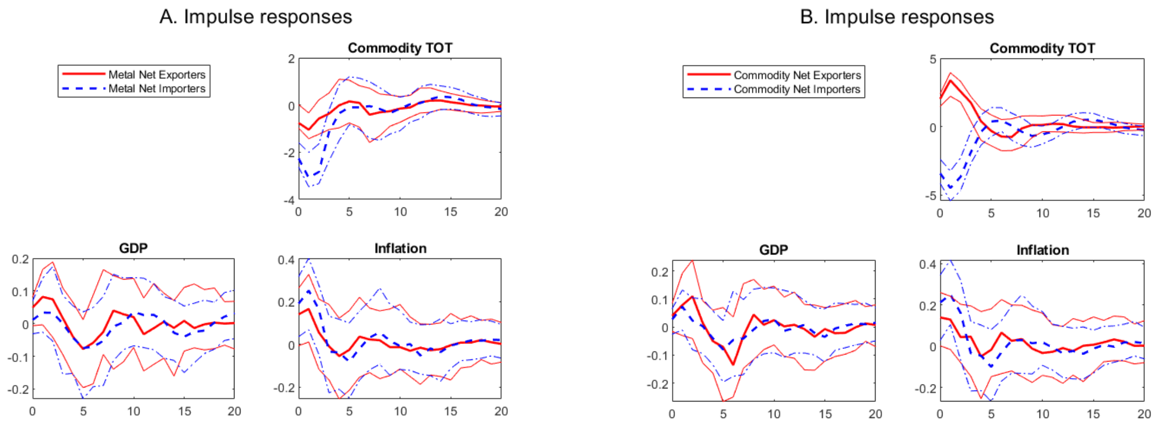
Note: Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure A3.6. Interannual data frequency: Impulse responses to a symmetric fuel-price shock



Note: Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure A3.7 Interannual data frequency: Impulse responses to a symmetric metal-price shock



Note: Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Figure A3.7 shows the responses to a 10% increase in the global metal price. The metal net importers show a 2.26% decrease in commodity terms-of-trade and a 0.19% increase in inflation (panel A). As seen in panel B, commodity terms-of-trade increases by 2.03% in the commodity net exporters but decreases by 3.4% in the commodity net importers. Only the latter group displays an inflationary effect of 0.21%.

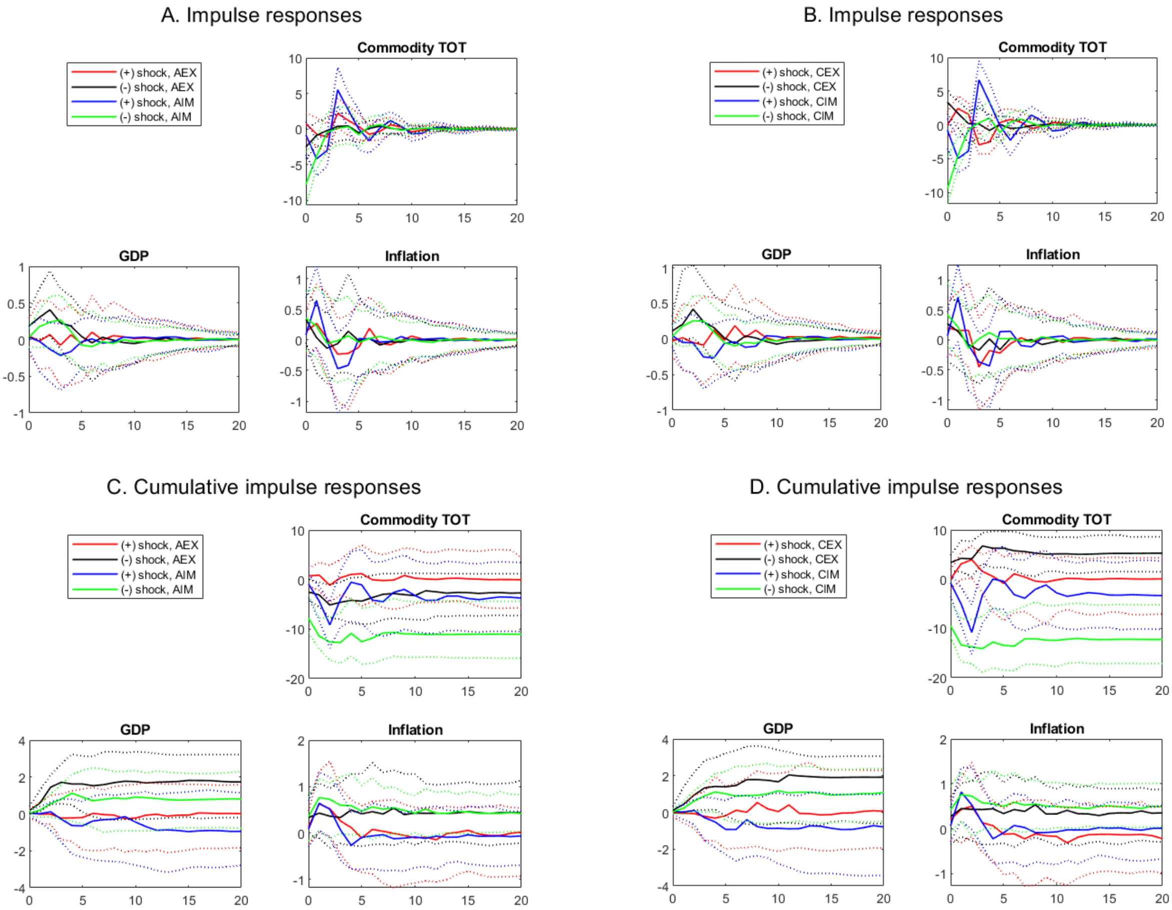
The above results are consistent with the benchmark finding that the global commodity-price shocks are more favourable to the net exporters than the net importers. Particularly, commodity terms-of-trade decreases contemporaneously to the shocks for the net importers but either increases or remains unchanged for the net exporters. The net importing countries also display discernible inflationary effects more often than the net exporting group.

A3.4.3 Asymmetric shocks

Global agriculture-price shock

Figure A3.8, panel A shows the impulse responses of the agriculture net importers and the agriculture net exporters to the asymmetric shocks in the global agriculture price. Only the negative shock leads commodity terms-of-trade to decrease by 2.5% in the agriculture net exporters and 7.8% in the agriculture net importers. That negative shock also brings GDP growth of the former group to increase by 0.18% while the positive shock has no impact on either commodity terms-of-trade or GDP growth. The inflationary effects are negligible in both groups. Similarly, there is marginal effect on the commodity net exporters and the commodity net importers in the case of positive agriculture-price shock (panel B). Meantime, the commodity net exporters benefit from the negative shock with a 3.36% increase in commodity terms-of-trade and a 0.42% expansion in GDP growth, even with a 2-quarter delay. That negative shock, however, causes a 9.49% decrease in commodity terms-of-trade of the commodity net importers.

Figure A3.8. Impulse responses to asymmetric agriculture-price shocks



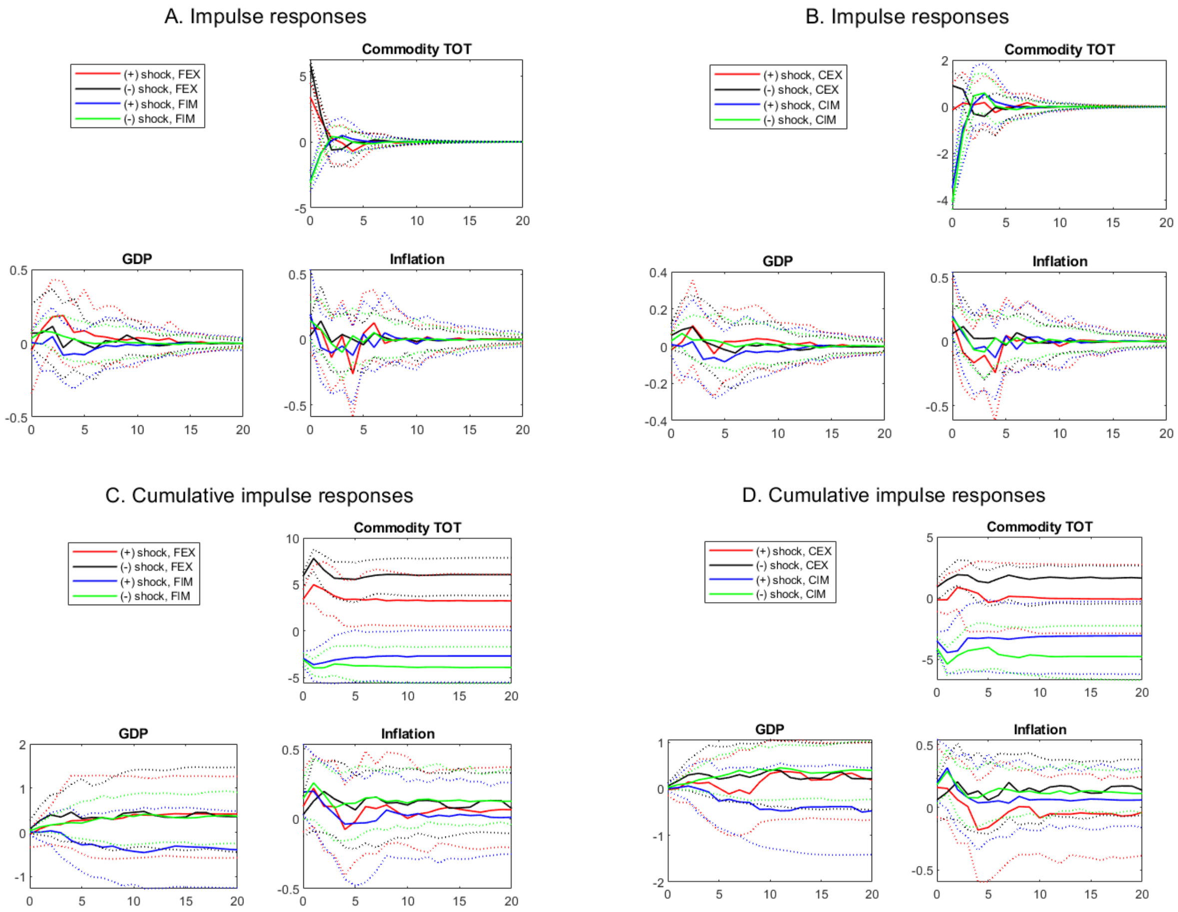
Note: AEX: Agriculture net exporters; AIM: Agriculture net importers; CEX: Commodity net exporters; CIM: Commodity net importers. Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

Looking at the cumulative responses displayed in panels C-D, the agriculture net exporters experience increases in GDP growth, which last for 6 following quarters since the negative shock. One-year accumulated expansion in their GDP growth is 0.5% approximately. The negative agriculture-price shock also leads to cumulative expansions in GDP growth of the commodity net exporters and the commodity net importers, which are discernible for 3 quarters (quarters 2-4) with one-year accumulated increases being 1.41% and 1.11%, respectively. Neither the positive shock nor the negative shock causes an inflationary effect on the commodity net exporters but both shocks lead to a transitory cumulative inflationary effect of 0.8% on the commodity net importers in quarter 1.

Global fuel-price shock

Figure A3.9, panel A shows that the positive and negative fuel-price shocks are favourable to the fuel net exporters, leading their commodity terms-of-trade to increase contemporaneously by 3.39% and 5.88%, respectively. I also find a 0.19% inflationary impact of the positive shock on the fuel net importers but marginal responses of GDP growth. In panel B, those asymmetric fuel-price shocks leave the domestic variables of the commodity net exporters unchanged but lead the commodity terms-of-trade to contract by 4% and inflation to increase by 0.2% for the commodity net importers. GDP growth rate shows slight responses to the shocks.

Figure A3.9 Impulse responses to asymmetric fuel-price shocks



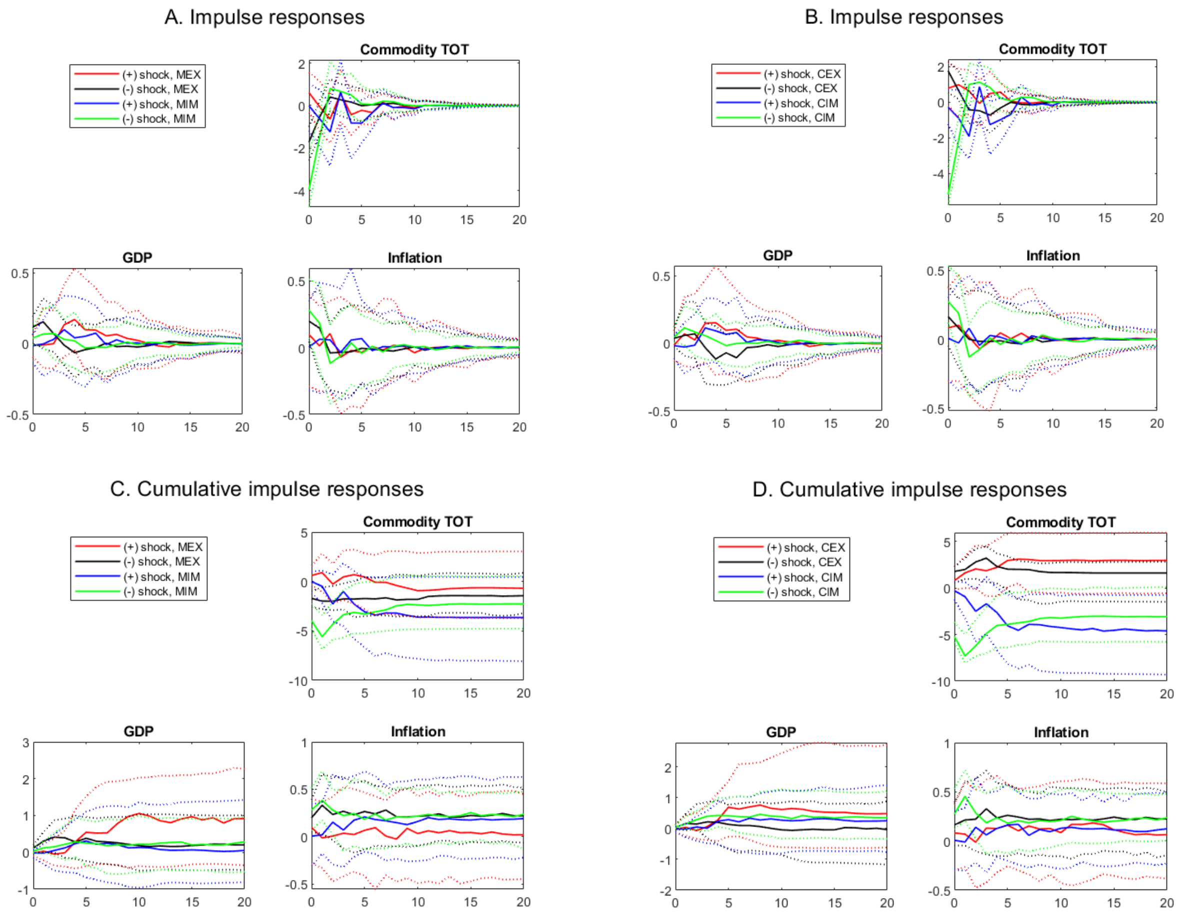
Note: FEX: Fuel net exporters; FIM: Fuel net importers; CEX: Commodity net exporters; CIM: Commodity net importers. Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

In terms of the cumulative responses (panels C-D), I find a temporary inflationary effect of 0.19% on the fuel net importers resulted from the positive fuel-price shock. For the commodity net exporters vis-à-vis the commodity net importers, only the latter group displays a transitory inflationary impact of 0.2%-0.3%. The effects are present for 1 - 2 quarters with the impact of the negative shock lasting one-quarter longer.

Global Metal-price shock

Figure A3.10 reports the impulse responses to the positive and negative shocks in the global metal price. In panel A, to the negative metal-price shock, the metal net exporters and the metal net importers display an increase of 1.71% and 3.94% in commodity terms-of-trade and an inflationary effect of 0.2% and 0.27%, respectively. The positive metal-price shock leads GDP growth of the metal net exporters to rise by 0.12% while leaving other variables unaffected. Regarding the commodity net exporters and the commodity net importers (panel B), the negative metal-price shock brings two contrast effects to their commodity terms-of-trade, with a positive impact of 1.77% on the former group but a negative effect of 5.2% on the latter. The negative shock also leads to a minor increase of 0.05% in GDP growth and a 0.27% increase in inflation of the commodity net importers. No groups show distinguishable responses to the positive metal-price shock.

Figure A3.10. Impulse responses to asymmetric metal-price shocks



Note: MEX: Metal net exporters; MIM: Metal net importers; CEX: Commodity net exporters; CIM: Commodity net importers. Solid lines: median impulse responses; Thinner lines: 68% confidence bands estimated using the bootstrapping method. The x-axis measures quarters after the shock.

As shown in panels C-D for the cumulative impulse responses, the inflationary effects from the negative metal-price shock are transitory, lasting for the first two quarters for the metal net exporters and three quarters for the metal net importers. There is also an increase of 0.16% in GDP growth, which lasts temporarily in quarter 2, and the positive cumulative effects which remain for the first three quarters for the commodity net importers due to the negative metal-price shock.

The robustness check implies the existence of the asymmetric effects of the global commodity-price shocks. The domestic variables are more responsive to the negative shocks than to the positive shocks, especially in the global agriculture and metal prices. Compared to the net importing countries, the net exporting groups experience more favourable effects from the shocks: increasing commodity terms-of-trade, higher GDP growth rate, which also lasts long, and lower inflation.

A3.4.4 Variance decomposition

The variance decomposition in Table A3.4 below, panel A shows that the global commodity-price shocks widely explain 35% - 75% of the variations in commodity terms-of-trade for the commodity net exporters and importers; the global fuel-price shock plays the largest role (60% - 75%). The shocks in global commodity prices also account for 10% - 15% of GDP growth's variations and 10% - 13% of inflation's variations for these groups. Panels B-D show that the fuel-price (metal-price) shock explains a comparable variation share of output growth for both fuel (metal) net exporters and importers. The global agriculture-price shock explains 15% of GDP growth variations for the agriculture net exporters, twice as much as for the agriculture net importers, 8% (panel B).

Table A3.4 Shares of variance explained by the symmetric global commodity-price shocks

	Number of countries	Share of variance					
		p	tot	gdp	inf	ir	rer
A. Commodity net exporters versus commodity net importers							
Commodity net exporters	44						
Agriculture shock		100	34.26	9.57	9.69	13.15	16.17
Fuel shock		100	59.96	12.83	12.92	13.22	13.44
Metal shock		100	40.39	14.75	10.88	13.69	10.82
Commodity net importers	48						
Agriculture shock		100	39.21	10.64	14.79	9.82	17.78
Fuel shock		100	75.46	14.35	18.96	11.81	15.84
Metal shock		100	41.16	13.30	11.66	9.30	12.91
B. Agriculture net exporters versus Agriculture net importers to the agriculture shock							
Agriculture net exporters	44	100	31.69	15.44	12.75	18.07	16.60
Agriculture net importers	48	100	38.97	7.98	14.16	9.82	16.58
C. Fuel net exporters versus Fuel net importers to the fuel shock							
Fuel net exporters	24	100	77.76	13.01	11.16	13.58	11.68
Fuel net importers	68	100	65.96	14.03	18.61	12.29	15.91
D. Metal net exporters versus Metal net importers to the metal shock							
Metal net exporters	39	100	43.09	15.68	11.68	11.19	12.63
Metal net importers	53	100	40.00	12.79	10.45	9.56	12.40

Note: Group-specific median shares of variance are computed from country-specific variance decompositions based on country-by-country SVAR estimations in the baseline case.

Chapter 4 The Effect of Monetary Policy on the New Zealand Dollar: a Bayesian SVAR Approach

4.1 Introduction

As a small open economy, New Zealand is susceptible to fluctuations in the exchange rate in both domestic prices and economic activities. Understanding the responses of exchange rate to monetary policy shocks is important to the monetary policymakers. Along with pursuing price stability as a key objective, the Policy Targets Agreement introduced in December 1999 thus requested the Reserve Bank of New Zealand to "…seek to avoid unnecessary instability in output, interest rates and the exchange rate". Motivated by its importance, this chapter greatly contributes to the literature of monetary policy – exchange rate analyses for New Zealand, which has been rather scant, by using the Bayesian structural vector autoregression (SVAR) approach to re-examine the impact of monetary policy shocks on exchange rate of New Zealand dollar (NZD) against US dollar (USD).

The theory of uncovered interest parity (UIP) is the central block in macroeconomic models connecting the expected changes of exchange rate to the interest rate differentials. As the domestic central bank tightens monetary supply, the UIP theory implies a greater appreciation of domestic currency in the short-run than its long-run level, the so-called overshooting phenomenon. Despite its popularity, the UIP validity has been strongly challenged by empirical evidence, including mine. I estimate an SVAR for five variables of the US and New Zealand money market rates, stock prices, and the bilateral nominal exchange rate, using the Bayesian approach introduced by Baumeister and Hamilton (2015) by explicitly imposing prior information on the structural parameters. By doing so, I am transparent about the influence of prior information on posterior results. The findings show that an unexpected increase in New Zealand's short-run interest rate causes a contemporaneous appreciation of NZD against the US dollar (USD) and even stronger NZD in the long-run than prior to the shock.

The central problem in investigating the interest rate – exchange rate relationship is the endogeneity of the variables. The biggest contribution of this chapter is to employ stock prices and the co-movements between interest rates and stock prices to untangle the unexpected monetary policy shocks from other shocks that simultaneously affect interest rates and exchange rate, including economic news shocks and currency premium shocks. A surprised monetary policy tightening is associated with higher short-term interest rates and lower stock prices whereas a positive economic news shock likely results in increases in both stock prices and interest rates to stabilize economic growth and inflation.

A positive currency risk premium shock, for example, capital flow episodes into New Zealand which lead to NZD appreciation as defined in this chapter, may damage exports and domestic stock prices and encourage the central bank to lower interest rate.

Existing studies have employed various identification approaches to analyse the impacts of monetary policy shocks, largely focusing on unexpected interest rate changes, on exchange rate movements. The first is event-study approach. To isolate the surprise from anticipated monetary policy shocks, a number of papers look at very short windows, for example, in minute, day, or intra-day windows around the central banks' announcement or communication events for the variations of exchange rates. This approach of using high-frequency data is popular in examining the responses of asset prices to monetary policy shocks. For instance, Zettelmeyer (2004) focuses on (immediate but dynamic) responses of exchange rates to the shocks associated with specific policy actions, such as changes in official interest rates or the overnight rate targets, and uses the reactions of market rates as measures of the unanticipated component of the actions. From both OLS and instrumental variable (IV) regressions across the sample of Australia, Canada, and New Zealand during the 1990s, he finds that a 1% increase in the 3-month interest rate appreciates the exchange rate by 2–3%. Later, Kearns and Manners (2006) additionally add the United Kingdom into their 4-country sample and also use the changes of market rates to measure the surprise component of monetary policy reactions. They use instead intraday data – a 70-minute event window – to eliminate the events jointly affecting both interest rate and exchange rate. The studied periods vary across the four countries; the included events for New Zealand occurred during the 17/3/1999 – 10/6/2004 period. The average results for the four countries show that the exchange rate appreciates by around 0.35% to a surprising 25-basis-point increase in the policy interest rates. More recently, Rosa (2011) also uses intraday data with 30- and 60-minute windows for five currencies (the exchange rate of USD against the euro, the Canadian dollar, the British pound, the Swiss franc, and the Japanese yen) and finds a greater impact of the Federal Open Market Committee's monetary policy surprises on exchange rates. On average, the USD exchange value depreciates by 0.5% in response to an unanticipated 25-basis-point cut in the Federal Funds target rate.

Another approach is the identification through heteroscedasticity (ITH), which was first introduced by Rigobon (2003). The idea of this approach is that, to solve the identification problem in simultaneous-equation models, i.e., when the structural estimators must be recovered from the reduced parameters and there are fewer equations than the number of unknown parameters, I need to impose additional information or restrictions. Instead of the exclusion, sign, short-run, and long-run restrictions, which are traditional in the literature, Rigobon (2003) proposed to use the heteroscedasticity of the structural shocks across regimes (or subsamples) contained in the data to add more equations into the system, while keeping other aspects of the structure identical including the assumption of uncorrelated structural shocks. By using the difference in the variance of residuals of the structural equations across regimes, the system is exact-identified. Ehrmann, Fratzscher, and Rigobon (2011), for example, employ the ITH approach to examine the financial linkages between the US and the euro area money markets, bond markets, equity markets, and foreign exchange markets. They estimate a structural system using 2-daily windowed data over 1989-2008 of the US 3-month Treasury bill rate, the US 10-year Treasury bond rate, the S&P 500 index, the 3-month interbank rate (the FIBOR rate before 1999 and the EURIBOR after 1999), the German 10-year government bond, and the S&P Euro index. Other variables are included to control for economic news in the US and the euro area, and oil price changes. This chapter is more related to Ehrmann, Fratzscher, and Rigobon (2005)'s working paper version, estimating the changes instead of levels of the variables. In any case, they compute the rolling windows variances of 20 two-day observations for each variable, i.e., each asset variable is multifactor modeled and a heteroscedasticity regime is identified if at least 16 observations for which the relative variances of at least one asset returns are larger than their average value plus one standard deviation. Then they use the estimated covariance matrices of 7 out of 28 identified heteroscedasticity regimes to generate new data in each bootstrap replication, and choose the estimators to minimize $g'g$ with $g = A' \sum_i A - \Omega_i$, where A is the structural matrix capturing the contemporaneous interactions of the variables, $\sum$ is the variance of the structural shocks, and Ω is the variance-covariance matrix estimated in each regime *i*. As leaving the impact of interest rate on exchange rate unrestricted, they find in the Ehrmann et al. (2005) version that, on average, a 1% increase in the short-term interest rate lead the USD to appreciate by 3.698% against the euro. Their results also stress the existence of international spillover effects within as well as across asset classes and asset prices are more responsive to domestic asset price shocks rather than to international shocks.

Other studies such as Sims (1992), Eichenbaum and Evans (1995), and Karim, Lee, and Gan (2007) rely on the SVAR recursive Cholesky approach to identify monetary shocks and the responses of exchange rates. The exchange rate effects of monetary shocks are ambiguous, however. Sims (1992), for example, finds evidence of exchange rate puzzle with large and persistent domestic currency depreciation for France and Germany following interest rate increases. By contrast, a responding pattern consistent with the theory, i.e., monetary policy contraction raises the value of domestic currency, is found for Japan, the UK, and the US. Similarly, Eichenbaum and Evans (1995) show that a US monetary policy contraction, identified as the shocks to either the Federal Funds rate, the ratio of non-borrowed to total reserves, or the Romer and Romer index of monetary policy, leads to persistent and significant appreciation of the USD. In a study for New Zealand, Karim et al. (2007) also apply the SVAR with Cholesky decomposition method, which imposes a recursive ordering on the structural model, for an 8-variable system including foreign output, non-oil commodity price index, consumer price index, and bank rate which are placed before New Zealand block of output, consumer price index, effective exchange rate, and official cash rate. Using a quarterly data sample covering the 1985Q1-2003Q4 period across four major partners of New Zealand including Australia, Japan, the US, and the UK, they find no evidence of exchange rate puzzles. Both nominal and real effective exchange rates of NZD are found to appreciate immediately and depreciate subsequently to an unexpected monetary policy contraction. The responses in all cases are insignificant, however. Besides, they find a modest role of the monetary policy shocks (from 0.2 to 3%) to explain the variations of exchange rates of NZD.

Cushman and Zha (1997) argue that the recursive approach to monetary policy identification while being plausible for the US studies because the movements in the US interest rates are less likely affected by foreign shocks are less valid for smaller and open economies. The central banks in such economies more likely to adjust interest rates to respond to foreign markets, thus invalidating the assumption of independent interest rates and generating puzzling exchange rate responses to interest rate changes. Cushman and Zha (1997) emphasize the need of using appropriate procedures to identify monetary policy shocks in smaller and open economies than the US. They propose a structural non-recursive approach with zero restrictions for Canada, i.e., a structural VAR model with block exogeneity, allowing monetary policy to react contemporaneously to a range of foreign and domestic variables whose data is available immediately to the policymakers and vice versa. The structural parameters in their system reflecting simultaneous relations will become zero in the recursive Cholesky approach. They identify monetary policy shocks as the changes in the money stock and find evidence consistent with the standard theory: a decline in monetary stock is followed by an immediate and significant Canadian dollar appreciation. Kim and Roubini (2000) also use the non-recursive SVAR approach for monthly data from 7/1974 to 12/1992 for non-US G7 countries whose exchange rates (against the USD) are found to appreciate initially and gradually depreciate after a few months following a monetary contraction. In their study, monetary policy shocks are identified by modeling money supply as a response function of monetary authorities, exchange rate, world oil price, price level, and the US Federal Funds rate.

Another SVAR approach is proposed to impose sign restrictions on structural coefficients (for example, see Kim and Lim (2018)) and/or lagged coefficients such that the signs of impulse responses reflect strong expectations established in the literature. This traditional approach was criticized by Baumeister and Hamilton (2015) for not being agnostic as described, only delivering an identification set that satisfies the imposed sign restrictions, thus limiting the possible distributions. Baumeister and Hamilton (2015) emphasize the need to explicitly acknowledge how the informative priors affect the structural estimation, which makes the Bayesian approach an unambiguous improvement in comparison to the traditional frequentist approach. The Bayesian approach yields a posterior distribution for the structural parameters and other objects of interest such as impulse response functions, which are consistent with the traditional sign-restricted SVAR and handy to check their sensitivity to the imposed priors. In this sense, this chapter serves as a contributor to the scant literature of examining the effect of monetary policy on exchange rate of NZD against USD by using the Bayesian SVAR approach, transparently combining sign restrictions where intuitive and prior modes from existing studies. This study is most close to Grisse (2020), which is the very first paper applying the Bayesian method for the Switzerland case. They find strong evidence to support the UIP theory that the exchange rates of Swiss franc (against euro and USD) overshoot on impact and depreciate in the following weeks after the increases in Swiss short-term interest rates.

The remainder of this chapter proceeds as follows. Section 4.2 introduces the theoretical motivation. Section 4.3 presents the empirical framework and data used to examine the effects of New Zealand monetary policy shocks on exchange rate of NZD against the USD. Section 4.4 reports the results. Section 4.5 extends the study by relaxing several restrictions in the baseline model. Section 4.6 concludes.

4.2 Theoretical motivation

One of the contributions of this chapter is to employ stock prices to disentangle monetary policy shocks from other shocks that jointly drive interest rates, stock prices, and exchange rate including economic news shocks and currency premium shocks. First, I present below the traditional equations linking interest rates, stock prices, and exchange rate as the motivation for my empirical framework. Then I will discuss the related empirical findings, providing useful information for the chosen priors.

Impact of interest rate on exchange rate

Uncovered interest parity is the cornerstone condition for macroeconomic analysis of small open economies. According to the UIP, the basic equilibrium condition of the foreign exchange market is:

$$i_t - i_t^* = E_t(\Delta e_{t+1}) + \rho_t \tag{4.1}$$

, where i_t and i_t^* are the short-term domestic and foreign interest rates, respectively, $E_t(\Delta e_{t+1})$ is the expectation of percentage change in nominal exchange rate of domestic currency against foreign currency ($E_t(\Delta e_{t+1}) = e_{t+1} - e_t$; e_t, e_{t+1} in natural logarithm form; higher e refers to domestic currency appreciation in this chapter), and ρ_t is a risk premium. Equation (4.1) predicts that, when risk premium is very small, a rise in domestic short-term interest rate relative to foreign interest rate should be associated with domestic currency appreciation.

Equation (4.1) can be rewritten as:

$$e_t = e_{t+1} - (i_t - i_t^*) + \rho_t \tag{4.2}$$

Solving equation (4.2) using forward-looking rational expectations (i.e. the law of iterated expectation) after *n* repeated substitutions to get:

$$e_t = E_t(e_{t+n+1}) - \sum_{j=0}^{n} E_t\,(i_{t+j} - i_{t+j}^*) + \sum_{j=0}^{n} E_t(\,\rho_{t+j}) \tag{4.3}$$

Equation (4.3) tells us three possible transmission channels through which monetary policy can influence today's exchange rate: market expectations of future exchange rate (the first term), market expectations of interest rate differentials (the second term), and market expectations of future risk premia (the last term). The first-difference of equation (4.3) says that a tighter domestic monetary policy should be associated with exchange rate appreciation. The predictions from equations (4.1) and (4.3) tell us about the overshooting phenomenon, i.e., the short-run response of exchange rate appreciation is greater than its long-run response when the domestic central bank tightens monetary supply.

Impact of interest rate on stock price

Understanding the effects of monetary policy changes on asset prices is crucially important to monetary policymakers. The most direct and indirect impacts of monetary policy innovations are on financial markets, which in turn affect the macroeconomic volatility, thus understanding these transmission mechanisms will help monetary policymakers react appropriately to achieve the ultimate objectives. In this subsection, I begin with equation (4.4) below to show the theoretical mechanism of interest rate's influence on stock price, and will discuss the related empirical findings on the interactions of monetary policy and asset prices shortly:

$$R_{t+1} \equiv \frac{P_{t+1} + D_{t+1}}{P_t} \tag{4.4}$$

, where R_{t+1} is ex-post stock market return, P_t and P_{t+1} are stock prices at time *t* and *t+1* respectively, and D_{t+1} is the dividend from time *t* to time *t+1*. Taking the logarithm of both sides of equation (4.4), deriving a log-linearization approximation to the logarithm, then solving forward for *n* repeated substitutions, and decomposing ex-post return into excess return and short-term interest rate ($er_t \equiv r_t - i_t$), I get:

$$p_t \approx \sum_{j=0}^{n} \rho^j \kappa + \rho^{n+1}(E_t p_{t+n+1} - E_t d_{t+n+1}) - \sum_{j=0}^{n} \rho^j E_t\big(er_{t+j+1} + i_{t+j+1}\big) + \Big[\sum_{j=0}^{n}(1-\rho)\rho^j E_t d_{t+j+1} + \rho^{n+1} E_t d_{t+n+1}\Big] \quad (4.5)$$

, where p_t, r_t, d_t are logarithms of the stock price, return, and dividend respectively; κ and ρ are parameters with $\kappa > 0$ and $0 < \rho < 1$. In the right-hand side of equation (4.5), the first term is a constant, and the second term, i.e., expected price-dividend ratio, will approach some equilibrium value when n is sufficiently large. The last two terms suggest two channels through which conventional monetary policy shocks can affect stock prices. First, a higher interest rate depresses stock prices by increasing the risk-free components of discount rate and hence a lower present discounted value of dividends (the third term). Second, an increasing interest rate should be associated with a deteriorating growth outlook and thus lower expected dividends (the last term).

Examining the interactions between asset prices and monetary policy has been standing as an attractive topic in the empirical literature. Rigobon & Sack (2003, 2004) examine both sides of the interactions using the heteroscedasticity identification approach. Rigobon and Sack (2003) point out two channels including wealth and the financing cost to businesses through which stock price movements impact the US macroeconomy and thus determine monetary policy decisions (as equity accounts for a large proportion of the US households' total financial wealth and non-financial corporations' assets). Using the daily US data for the 3-month Treasury bill rate and the return on the S&P 500 index from 3/1985 to 12/1999, they find that a 5% unexpected increase in the S&P 500 index increases the Federal Funds rate by about 14 basis points, i.e., 1% increase in the S&P 500 index increases the Federal Funds rate by 0.021%. Rigobon and Sack (2004) study the other side of the relationship, i.e., the impact of monetary policy on asset prices, by implementing the heteroskedasticity identification as IV and generalized-method-of-moments (GMM) regressions using a variety of stock market indices and longer-term interest rates from 03/01/1994 to 26/11/2001. The IV and GMM estimators are very close in magnitude, in particular, a 1% increase in the short-term interest rate causes the S&P 500 index to decline by 6.81% (for the IV estimator) and 7.19% (for the GMM estimator). Motivated by the same question of stock market' responses to monetary policy, Bernanke and Kuttner (2005) use the event-study approach to show that, for the period from 06/1989 to 12/2002, an unexpected 1% easing in the Federal Funds target rate is associated with an approximate 4.68% increase in broad stock indices. They also find that the predominant effects of monetary policy on the stock market come through expected future excess equity returns. The negative impact of monetary tightening on equity prices are in line with Claus, Claus, and Krippner (2018), quantifying the responses of a variety of the US asset price indices to conventional and unconventional monetary policy shocks separately by using a latent factor model with heteroscedasticity identification for monetary and non-monetary policy event days.

4.3 Empirical framework and data

4.3.1 Empirical framework

Based on equations (4.3) and (4.5), I construct the following equations in the linear empirical model, temporarily excluding lagged terms and constant:

$$i_t^* = \alpha_s^* s_t^* + u_t^{i*} \tag{4.6}$$

$$s_t^* = \beta_i^* i_t^* + u_t^{s*} \tag{4.7}$$

$$i_t = \alpha_{i*} i_t^* + \alpha_{s*} s_t^* + \alpha_s s_t + \alpha_e e_t + u_t^i \tag{4.8}$$

$$s_t = \beta_{i*} i_t^* + \beta_{s*} s_t^* + \beta_i i_t + \beta_e e_t + u_t^s \tag{4.9}$$

$$e_t = \gamma_{i*} i_t^* + \gamma_{s*} s_t^* + \gamma_i i_t + \gamma_s s_t + u_t^e \tag{4.10}$$

, where i_t^* and i_t are the US and New Zealand short-term interest rates, respectively; s_t^* and s_t are the US and New Zealand stock price indices; and e_t is nominal exchange rate of NZD against USD, i.e. an increase in exchange rate implies NZD appreciation. As a conventional monetary policy operates by changing short-term interest rate, the structural residuals u_t^{i*} and u_t^i in equations (4.6) and (4.8) are interpreted as the US and New Zealand monetary policy shocks. A positive monetary policy shock, i.e., monetary tightening, is expected to move interest rates up and stock prices down. The last terms in equations (4.7) and (4.9), u_t^{s*} and u_t^s, can be interpreted as the US and New Zealand economic news shocks which are expected to cause interest rates and stock prices to move in the same direction. In equation (4.10), u_t^e is interpreted as currency premium shock, reflecting the shocks to financial risk premia unrelated to monetary policy and economic news shocks. As government bonds provide a hedge against the shocks which make stock investment risky, a positive currency premium shock, which leads to NZD appreciation as defined in this chapter, lowers both stock and bond prices. To cover up, interest rates and stock prices co-vary in the same direction to positive economic news shocks and currency premium shocks but negatively to positive monetary policy shocks. The way I identify the monetary policy shocks by employing the co-movements of interest rates and stock prices in response to monetary policy shocks, economic news shocks, and currency premium shocks are in line with the literature, for example, see Matheson and Stavrev (2014), Cieslak and Schrimpf (2019), and Jarociński and Karadi (2020). These papers, however, use high-frequency co-movements of interest rates and stock prices around the communication events by central banks to isolate the unexpected policy shocks from other shocks contained in the central banks' announcements or communication. The non-monetary policy shocks are defined as economic news shocks by Matheson and Stavrev (2014), as news about economic growth and news affecting financial risk premia by Cieslak and Schrimpf (2019), and as "central bank information shocks", i.e., the way the central banks assess economic outlook, by Jarociński and Karadi (2020).

In equations (4.6) and (4.7), I also assume that the US economic conditions do not immediately respond to those in New Zealand. In equation (4.8), the response of New Zealand interest rate, apart from taking into account the US interest rate, follows the Taylor rule, subject to the growth condition, inflation, and exchange rate. I also describe the initial expectations on the contemporaneous impacts; some of the sign expectations will be relaxed in the baseline model and robustness checks. An increasing domestic stock price may reflect an expected favorable economic growth, which is usually associated with a higher inflation rate and thus a higher interest rate ($\alpha_s, \alpha_s^* > 0$). If the net effect from exchange rate appreciation on exports and imports is negative, I also expect a lower economic growth rate and inflation, and thus a loosening monetary policy ($\alpha_e < 0$). In equations (4.7) and (4.9), higher interest rates should be associated with lower stock prices ($\beta_i, \beta_i^* < 0$). Similarly, for equation (4.9), I expect a negative impact of exchange rate appreciation on exports and domestic stock price ($\beta_e < 0$). Equation (4.10) is motivated by equation (4.3), implicitly assuming that today's interest rate is a linear function of expected interest rates. Equation (4.3) implies that a higher domestic (foreign) interest rate should be associated with domestic currency appreciation (depreciation), i.e. $\gamma_i > 0$ and $\gamma_{i*} < 0$. Similarly, a higher domestic (foreign) stock price indicates an improved economic growth in the domestic (foreign) country and thus domestic currency appreciates or $\gamma_s > 0$ (depreciates or $\gamma_{s*} < 0$). I also assume a co-movement of domestic and foreign interest rates ($\alpha_{i*} > 0$) based on historical data plotted in Figure 4.1 (left panel) and stock prices ($\beta_{s*} > 0$) based on Figure 4.2 (left panel).

I proceed with the SVAR specification as follows

$$Ay_t = b_0 + \sum_{l=1}^{m} B_l y_{t-l} + u_t \tag{4.11}$$

, where y_t is the (*n* x *1*) vector of endogenous variables, the objects A and B_l are (*n* x *n*) matrices of structural and lagged coefficients, b_0 is the (*n* x *1*) vector of constants, u_t is the (*n* x *1*) vector of structural shocks with u_t assumed to be normally distributed u_t ~ N(0, *D*) and the covariance matrix

D being diagonal, and m is the number of lags. Specifically, $y_t = (i_t^*, s_t^*, i_t, s_t, e_t)'$, $u_t = (u_t^{i*}, u_t^{s*}, u_t^i, u_t^s, u_t^e)'$, and

$$A = \begin{bmatrix} 1 & -\alpha_s^* & 0 & 0 & 0 \\ -\beta_i^* & 1 & 0 & 0 & 0 \\ -\alpha_{i*} & -\alpha_{s*} & 1 & -\alpha_s & -\alpha_e \\ -\beta_{i*} & -\beta_{s*} & -\beta_i & 1 & -\beta_e \\ -\gamma_{i*} & -\gamma_{s*} & -\gamma_i & -\gamma_s & 1 \end{bmatrix}$$

In the estimation, i_t^* and i_t are the first-differences of the US and New Zealand short-term interest rates; s_t^* and s_t are the log-differences of the US and New Zealand stock price indices; and e_t is the log-difference of NZD nominal exchange rate against the USD. The very first prior information imposed in the structural matrix A is that New Zealand economic conditions do not affect those in the US in the same week and this is reflected in the upper right block of zero in the matrix A. In the baseline model, I also follow the literature by assuming no international cross-market spillover effects, i.e., the US stock market (interest rate) has no impact on the New Zealand interest rate (stock market), or $\alpha_{s*} = \beta_{i*} = 0$.[35] This assumption will be relaxed later. Without further assumptions, the structural model in equation (4.11) is unidentified. There are 17 parameters to be estimated, including 12 unknown parameters in the matrix A and 5 diagonal elements in the covariance matrix D of the structural shocks while I have only 15 known unique elements in the (*5* x *5*) variance-covariance matrix of the reduced-form residuals. To exactly identify the model, one needs at least two more equality restrictions. In this study, I proceed with the Bayesian approach by following Baumeister and Hamilton (2015), specifying a full prior distribution rather than just sign and zero restrictions for the unknown structural parameters to get a set identification. In particular, the prior information is imposed on the elements in the matrix A, not the inverse matrix (A^{-1}). I will discuss the chosen priors in more detail in the next section.

4.3.2 Data description

The "raw" data includes the US 3-month Treasury bill rate, the S&P 500 index, the New Zealand 3-month Bank bill rate, the NZSE index, and the nominal bilateral exchange rate of NZD against the USD; all in daily frequency. The sample covers the period from 03/01/1999, when data on the NZSE series started being available, to 18/09/2020. I use weekly data based on the final trading day of the week to avoid the different daily timestamps across the markets. I believe that data of higher frequency (such as daily) contain too much noise whereas data of lower frequency (such as monthly

[35] For example, Ehrmann et al. (2011) impose a similar assumption of no international spillover effects across the US and European stock markets and interest rates.

or quarterly) may mute too many variations in stock prices and exchange rates. Appendix 4 Table A4.1 describes data sources in more detail.

For monetary policy reference rates, as the market rates are longer and are more subject to change than the target rates, I choose the US 3-month Treasury bill rate and New Zealand 3-month Bank bill rate instead of the Federal Funds rate and New Zealand Official Cash rate. The left panel in Figure 4.1 shows that the target rates are relatively stable, especially the New Zealand Official Cash rate, compared to the market rates. For example, the Official Cash rate was fixed at 2.5% for almost three years from the week ended on 18/03/2011 to 07/03/2014. As the market rates appear to co-move strongly with the target rates (the Fed Funds rate and the US 3-month Treasury bill rate since 2000 as well as the Official Cash rate and New Zealand 3-month Bank bill rate since 1999), I prefer to estimate the market rates with more variations contained. The left panel in Figure 4.1 also shows the positive correlation between the US and New Zealand market rates, especially clearly from 2004 to 2014. The correlation is not discernible before 2004 or after 2014.

The exchange rate is quoted per NZD, indicating that a higher value of exchange rate reflects NZD appreciation. The right panel in Figure 4.1 implies the positive correlation of New Zealand 3-month Bank bill rate and the nominal exchange rate (in natural logarithm), i.e., higher interest rate is associated with higher NZD value. This positive correlation does not imply a causal relationship between New Zealand monetary policy and exchange rate because the interest rate may be driven by other factors such as foreign interest rates and economic news that may also cause the exchange rate to change. Therefore, the crucial task is to ensure the unexpected interest rate changes to be disentangled from other shocks that jointly drive the movements of both interest rates and exchange rate.

There are various measures of the share market performance in New Zealand. The most popular measurements are the S&P/NZX family of indices. Among the available proxies, I collect data on the following: S&P50NZ, NZSE10, NZSEMC, NZSESC, and NZSE; all in nominal NZD. The S&P50NZ index measures the performance of the 50 largest index-eligible stocks listed on the NZX Main Board by float-adjusted market capitalization. The S&P50NZ data started from 29/12/2000 and is widely considered as New Zealand's pre-eminent benchmark stock price index. The NZSE10 measures the performance of the 10 largest New Zealand listed companies within the S&P50NZ index. The NZSEMC measures the performance of New Zealand's core mid-cap equity market, covering the constituents of the S&P50NZ index but excluding those that are also constituents of the NZSE10 index. The NZSESC index is designed to measure the performance of New Zealand's smaller listed companies that are not covered in the S&P50NZ index. The NZSE index is considered as the total market indicator for the New Zealand equity market, comprising all eligible securities quoted on the NZX Main Board. Apart from those S&P/NZX indices, there is also the MSCI New Zealand index (MSCINZ), which is designed to measure the performance of the large and mid-cap segments of the New Zealand market. The MSCINZ index covers 7 constituents, approximately accounting for 85% of the free float-adjusted market capitalization in New Zealand. The left panel in Figure 4.2 plots the natural logarithms of New Zealand stock prices, implying several common trends in their variations: all indices increase until the global financial crisis and recover afterward before entering another declining phase in early 2020 due to the Covid-19 pandemic. Among those indices, the NZSE series is a composite index based on the prices of stocks excluding dividends, not a total return stock index. Other S&P/NZX series such as S&P50NZ, NZSE10, NZSEMC, and NZSESC as well as the MSCINZ index are size-and-style stock indices by including the stock prices of specific groups of constituents. Despite so-called the benchmark index, the S&P50NZ series is the shortest among the S&P/NZX family indices with available since 2000. While these S&P/NZX indices are strongly correlated, I use the NZSE index for estimation for a longer sample (from 1990 after merging with other variables), which is also in line with the S&P 500 series used to proxy the US stock market performance. In the right panel in Figure 4.2, I plot the S&P 500 and the NZSE indices, all in natural logarithm. I find that the two stock price indices increase over the 1990-2020 period though their growth rates do not resemble all the time.

4.3.3 Unit root tests and optimal lag length

I conduct multiple unit root tests including Augmented Dickey–Fuller (ADF) test, Phillips–Perron (PP) test, Kwiatkowski–Phillips–Schmidt–Shin (KPSS) test, and Zivot & Andrews (ZA) test for stationary testing of the variables. The null hypotheses are different across those tests. In the ADF and PP tests, the null hypothesis of non-stationarity, i.e. series has a unit root, is tested against the alternative hypothesis of stationarity. By contrast, in the KPSS test, the null hypothesis of stationarity is tested against the alternative hypothesis of non-stationarity. Rejection of the null hypotheses in the ADF and the PP tests, and rejection of the alternative hypothesis in the KPSS test indicate the series is stationary. In the ZA test, the null hypothesis is that series has a unit root with a structural break(s) and the alternative hypothesis is that series is stationary with a break(s). Rejection of the null hypothesis in the ZA test indicates the series is stationary with a break(s). However, the ZA test can suggest only one break in one test. I consider all possible cases by including either constant or time trend or both in each test. Table 4.1 briefly reports the unit root test results at 5% significance level for five series (interest rates in percent, stock prices, and exchange rate in natural logarithm), showing whether the tested series is stationary (I(0)) or non-stationary (I(1)). Detailed results are provided in Appendix 4 Table A4.2.

Table 4.1 conclusively suggests at 5% level of significance that, the four series including New Zealand 3-month Bank bill rate, and logs of the S&P 500 index, the NZSE index, and NZD exchange rate are non-stationary in levels and stationary in first-differences. The only inconclusive case is the US 3-month Treasury bill rate, which is suggested to be stationary in the ADF and the PP tests with a time trend included but non-stationary in other tests. Because all four unit root tests suggest that the US 3-month Treasury bill rate is stationary in first-difference, I include the first-differences of the US and New Zealand market rates as well as the log-differences of the US and New Zealand stock price indices and NZD exchange rate in the estimation.

Next, I check the optimal lag lengths based on the Akaike Information criterion (AIC), the Hannan Quinn criterion (HQ), the Schwarz criterion (SC), and the Final Prediction Error criterion (FPE). The AIC and the FPE suggest a similar lag length of 19 while the HQ suggests 4 and the SC suggests 1 as the optimal. Because of the inconsistency of the optimal lags suggested across those criteria, I proceed with 8 lags for the weekly data.

4.3.4 Priors for the structural parameters

I follow Baumeister and Hamilton (2015, 2018) by assigning t-distributions with 3 degrees of freedom as priors for 12 unknown parameters in the structural matrix A. I will specify the prior modes, scales, and sign restrictions where possible and intuitive to the literature. All of the chosen prior modes, except for the effect of exchange rate on New Zealand stock price (β_e), are from Ehrmann et al. (2005). As mentioned in Section 4.1, their paper studies the financial transmission between short-term interest rates, bond yields, and equity returns, and exchange rate within and across the US and the euro area. The reasons I chose that paper as an index of the prior modes are: they use the ITH approach to report the contemporaneous coefficients in the structural matrix A, not the inverse matrix A^{-1}; and they estimate the changes instead of the levels of variables as reported in the Ehrmann et al. (2011) version. The prior modes are the average of their estimated coefficients for the US and the euro markets. For example, the prior mode 0.006 for the effect of stock prices on interest rates (α_s, α_s^*) are average of their reported estimators 0.0113 and 0.001. Column 3 in Table 4.2 provides the prior modes for 12 contemporaneous parameters.

Ehrmann et al. (2005) also report a postive impact of exchange rate on stock price, i.e., a 1% euro appreciation against USD is associated with a 0.5766% increase in the S&P Euro index. The S&P 500 index is irresponsive to exchange rate movements, however. For a small open export-driven economy such as New Zealand, I initially expect instead a negative correlation as exchange rate appreciation more likely damages export and possibly the stock prices. It is, however, inconclusive because the impact also depends on the share of export-oriented constituents in the stock market. Figure 4.3 plots the NZSE index and NZD exchange rate, revealing their positive correlation from 1990 to 2015 though their correlation appears to be reversed since then. To approximately quantify the contemporaneous impact of NZD exchange rate on the NZSE index (β_e), I simply conduct simple OLS estimations which also control for the dynamic effects of both exchange rate and stock price as follows[36]

$$\Delta LNZSE_t = c_0 + c_1 t + c_2 D + \sum_{i=1}^{p} \beta_i \Delta LNZSE_{t-i} + \sum_{j=0}^{q} \gamma_j \Delta LNZDUSD_{t-j} + u_t \quad (4.12)$$

, where Δ*LNZSE* and Δ*LNZDUSD* are the log-differences of the NZSE and NZD exchange rate against the USD. The unit root tests (including the ADF, the PP, the KPSS, and the ZA tests) suggest that the two series *LNZSE* and *LNZDUSD* are non-stationary in levels but stationary in first-differences. I also include *t* for the time trend and *D* as a dummy variable to represent the break dates of *LNZSE* suggested by the ZA test (either *D1* which gets 1 since 19/10/2007 and 0 otherwise, *D2* which gets 1 since 28/12/2007 and 0 otherwise, or *D3* which gets 1 since 23/12/2011 and 0 otherwise). Also, c_0 is a constant and u_t is the error term. As I will assign the estimated coefficient as the prior mode of β_e's t-distribution, at this stage, I ignore other determinants and control variables that could affect the NZD exchange rate and the NZSE index. In the summation terms, p and q are the optimal lag structure, chosen by the AIC. While the Bayesian Information criterion (BIC) suggests the same lag structure of (2,1) for p and q for all cases, the AIC suggests the lag structure of (6,2) if including time trend and a dummy either *D1* or *D2*, and (6,1) if including time trend and *D3*. Although I prefer the BIC as a consistent-model selector, I also do not want to under-fit my model, and thus I proceed with the lag structures chosen by the AIC. The coefficient of interest is γ_0. Apart from the three OLS estimations controlling for different suggested structural breaks, I also conduct an OLS estimation excluding both time trend and dummy variable. The full results are presented in detail in Appendix 4

[36] See Bahmani-Oskooee and Saha (2015) for an extensive review of the studies on the relations between stock prices and exchange rate. Existing studies either use univariate models or control for other determinants of stock prices and exchange rate. In any case of using either linear or non-linear models, most of the studies find no or weak evidence on the long-run equilibrium of the stock prices – exchange rate nexus. Specifically for New Zealand, Obben, Pech, and Shakur (2006) use the weekly data (average of daily data) of the NZSE index and disaggregated New Zealand exchange rates (against the USD, the Australian dollar, the British pound, and the euro) with a cointegrating VAR approach and find ambiguous evidence of the long-run relationship between stock price and exchange rates. In their equation of the USD/NZD exchange rate, the error correction terms, despite being negative, are not significant at 5% significance level, indicating no long-run equilibrium exists between these variables. The short-term coefficients of USD/NZD exchange rate, despite being positive, which implies that the NZD appreciation is associated with New Zealand stock price increases, are not significant either. No contemporaneous coefficients are reported in Obben et al. (2006)'s study. Therefore, we estimate equation (4.12) using OLS, including both contemporaneous for prior mode and the lagged variables to control for the dynamic effects.

Table A4.3. In any case, the contemporaneous coefficients are very close, ranging from 0.21 to 0.218 and all significant at 1% level. As leaving other determinants aside, the estimators implies a positive association between the NZSE index and the exchange rate, i.e. 1% appreciation of NZD is associated with approximately 0.2% increase of the stock price, rather than a causal relationship. I will assign 0.2 as the prior mode of β_e and discuss more this "positive" impact in Section 4.4 after achieving posterior distributions and impulse responses.

Next, I impose sign restrictions on the structural parameters, reflecting the interactions between stock markets, monetary policies, and exchange rate. First, higher stock prices are often associated with economic booms and inflation, and interest rate is expected to increase to stabilize inflation, I assume a positive impact of stock prices on interest rates. This assumption is consistent with the central banks' mandate. Vice versa, I follow the literature to assume a negative effect of unexpected interest rate changes on stock prices. For example, Matheson and Stavrev (2014) impose similar sign restrictions in their bivariate SVAR to examine the US financial market responses following the Federal Reserve's taper talk on 22/05/2013 by disentangling the unexpected monetary shocks from economic news shocks. The intuition is that a positive economic news shock leads to higher stock prices and a higher interest rate to stabilize inflation whereas an unexpected tighter monetary policy leads to a higher interest rate and lower stock prices. Using daily data of 01/2003-06/2014, they find that the immediate rise in the 10-year Treasury bond yields following the May 22 taper talk is mainly driven by monetary policy shocks while the effects of positive news shocks become more prominent during the subsequent months. I also assume a negative effect of exchange rate on monetary policy, i.e., exchange rate appreciation likely lowers interest rate. Although the positive correlation between exchange rate and stock price in New Zealand from the OLS estimations is in line with Ehrmann et al. (2005)' results, the findings on the impact are not conclusive in the literature. On the one hand, exchange rate appreciation can curtail exports, profits, and stock prices of export-oriented companies. On the other hand, appreciation decreases costs of imported inputs, lowers the production costs of non-exporting firms, hence increases their profits and stock prices. For this reason, the effect of NZD exchange rate on New Zealand stock price is left unrestricted. Additionally, despite the positive (negative) prior modes imposed on the effects of New Zealand (the US) interest rate and stock price on NZD exchange rate, I also leave their signs unrestricted for more possible posteriors to be achieved. Although the sign restrictions imposed in traditional SVAR, i.e., sign and exclusion restrictions, are based on the reasonable belief of the researchers on certain impacts, they restrict the set of identification. In addition, I agree with Baumeister and Hamilton (2015, 2018)' criticism on the sign restriction approach which implicitly assumes that the influence of the priors on posterior will vanish asymptotically. Those impacts of key interest are left unrestricted in this chapter to allow more effect scenarios to be obtained. By disclosing the prior information, I am transparent about the effect of the imposed priors on the posterior distributions and impulse responses. Lastly, I impose positive sign restrictions on the effects of the US interest rate (stock market) on New Zealand interest rate (stock market) as those indicators in a small open economy such as New Zealand will tend to follow the US markets.

Once prior modes and degrees of freedom are chosen, the prior scales determine the prior width. I choose the scales reasonably so that they meet the sign restrictions accordingly and more importantly, they are consistent with the previous studies. For instance, the prior for α_s and α_s^* – the effects of stock prices on interest rates – allows a large probability of 70.46% for them to be positive, also covering the estimator of 0.021 found by Rigobon and Sack (2004). For the effect of interest rates on stock prices, I allow a probability of 78.03% for the estimated parameters β_i and β_i^* to be negative, covering other existing estimators of -7.19 and -6.81 found by Rigobon and Sack (2003), and of -4.68 by Bernanke and Kuttner (2005). Wang and Mayes (2012)'s estimators for the effects of New Zealand and Australia monetary policy shocks on stock prices (-3.694 for New Zealand and -1.127 for Australia) using event-study approach are also included in the prior distributions of β_i and β_i^*. For other parameters with unrestricted signs such as β_e, the prior implies a possibility of 32.57% for a negative impact, i.e., NZD exchange rate appreciation drives New Zealand stock price to decrease, and 67.43% for a positive impact. This chosen prior of β_e includes the Ehrmann et al. (2005)' estimator of 0.5766. For γ_i and γ_{i*}, the chosen priors assign a very large probability of 97.3% for γ_i (γ_{i*}) to be positive (negative). The prior distribution of γ_i is in line with existing estimators in the literature, including the coefficients of from 2 to 3 found by Zettelmeyer (2004), of approximately 1.4 by Kearns and Manners (2006), and of 2 by Rosa (2011). By contrast, the priors imposed on γ_s (γ_{s*}) imply an equal probability of about 53.58% for them to be positive (negative).

4.4 Results

Figure 4.4 plots prior distributions (solid red curves) and posterior distributions (blue histograms) for the short-run effects (structural parameters). The key interest is γ_i, i.e., the contemporaneous effect of New Zealand interest rate on NZD exchange rate against the USD, about which the historical data slightly revises my beliefs as the prior and posterior distributions are very similar. Despite being sign-unrestricted, the prior and posterior distributions strongly imply that, following an increase in New Zealand short-term interest rate, the NZD exchange rate appreciates immediately on impact. I also find it less likely to revise my belief about the effect of the domestic stock price on short-term interest rate in the US (α_s^*) but more likely to revise for the New Zealand market (α_s) as the posterior distribution for New Zealand is narrower than the prior distribution. The prior and posterior distributions of γ_s – the contemporaneous effect of New Zealand stock price on NZD exchange rate – also resemble. However, my beliefs about other short-run effects are revised far more strongly when the posterior distributions are typically narrower than the prior distributions. The historical data favors a lower (larger) range for the effect of New Zealand (US) interest rate on stock price. The data also supports a smaller impact of NZD exchange rate on New Zealand short-term interest rate (α_e) compared to the chosen prior. The posterior distribution of foreign interest rate's impact on exchange rate (γ_{i*}), while being unrestricted, is far narrower than the prior, favoring a much smaller effect which is quite close to zero. The data also revises my beliefs about the effect of NZD exchange rate on New Zealand stock price (β_e), with posterior distribution favouring the negative impact, and the effect of the US stock price on NZD exchange rate (γ_{s*}), with posterior distribution favouring the positive impact. For the international spillover effects within the same asset class, the data supports a stronger co-movement of stock prices (β_{s*}) but a smaller for short-term interest rates (α_{i*}).

The median posterior values for the impulse response functions are shown as the solid lines in Figure 4.5, along with the 68% and 95% credibility sets. To a 1% unexpected increase in New Zealand short-term interest rate, I find that the NZD appreciates immediately by 1.51% on impact. The shaded 68% credibility regions exclude zero, strengthening my belief about the contemporaneous effect of monetary policy tightening on exchange rate appreciation. The 95% credibility regions include zero, however. As soon as the interest rate falls back to the initial level, exchange rate gradually depreciates to its original level. The posterior median of the direct impact (1.51) is close to other existing findings for New Zealand, such as 1.8–2 found by Kearns and Manners (2006) but much far from the prior mode (3.698) taken from Ehrmann et al. (2005).

Most of the other contemporaneous effects are as expected including increasing interest rates dampen stock prices, New Zealand short-term interest rate (stock price) co-moves positively with the US interest rate (stock price), a positive US monetary policy shock leads to NZD depreciation, the US short-term interest rate responds positively to the US economic news shock (the evidence for New Zealand is weak as the 68% credibility set includes zero), and the New Zealand interest rate increases in response to a positive currency premium shock. Despite the zero restrictions on the international spillover effects across asset markets, a higher US interest rate does cause New Zealand stock price to decrease for two weeks following the shock. The US stock price has no impact on New Zealand interest rate, however. Interestingly, I find that a positive economic news shock either in Zealand or in the US leads NZD value to increase immediately. Last but not least, the results show a negative response of New Zealand stock price to a positive currency premium shock despite the chosen prior of a positive impact. However, the impact is overall uncertain because the 68% credibility set of the direct response of New Zealand stock price to a currency premium shock also contains zero.

Figure 4.6 plots the median posterior values of cumulative impulse responses. The results show that the NZD exchange rate keeps appreciating persistently in response to a positive monetary policy shock: the 6-month accumulated response to a 1% increase in short-term interest rate is approximately 3.5% and there is no signal of "delay overshooting" over 6 months after the shock. I also expand the horizon up to one year (52 weeks) following the shock and find very similar responses of NZD exchange rate: the one-year accumulated appreciation of NZD exchange rate remains at 3.5%. The results are partly consistent with many existing studies that find contradict evidence to the UIP theory, which predicts subsequent exchange rate depreciations following an initial appreciation after a monetary policy contraction. Again, the findings on the effect of interest rate shocks on exchange rate are largely controversial in the empirical literature. Some studies, such as Sims (1992) show that the exchange rate depreciates after monetary tightening, which is the so-called exchange rate puzzle. Most of the other studies, for example, Cushman and Zha (1997), Kim and Roubini (2000), Kim (2005), as well as Kim and Lim (2018) report the evidence supporting the "delay overshooting" phenomenon with the delay lasting shortly, for example at best 6 months found by Kim and Lim (2018). Scholl and Uhlig (2008), however, document the more prolonged delay from one to three years before exchange rate starts to depreciate. Various explanations for the failure of the UIP theory have been discussed. One of them focuses on the invalidity of the two fundamental behavioral assumptions of the UIP theory in the data: market participants are risk-neutral and they have rational expectations about future exchange rate movements. If market participants are not risk-neutral, they will require a risk premium to hold foreign assets over domestic assets. In a recent paper, Granziera and Sihvonen (2020) relax the second assumption by allowing agents to have sticky expectations about short-term rates and illustrate that the increase in short-term rate forecast with sticky expectation occurs with a lag. Because of sticky expectations, agents have gradually updated their expectations about the short-term rates, the home currency keeps appreciating. This explains the failure of the UIP theory.

In addition, the results of cumulative impulse responses using the 68% credibility regions suggest the persistent impacts of short-term interest rates on stock prices and vice versa, of the US interest rate (stock price) on New Zealand interest rate (stock price), of the US interest rate on NZD exchange rate, of New Zealand stock price on NZD exchange rate, and of NZD exchange rate on New Zealand short-term rate. The persistent appreciation of NZD to a positive economic news shock in the US and the decrease of US stock price to a positive economic news shock in New Zealand, however, are unexplainable.

Table 4.3 reports the US and New Zealand variables' median variance shares, accumulated over 6 months, explained by the monetary policy shocks, economic news shocks, and currency premium shocks. The results show that New Zealand monetary policy shock plays a very modest role in explaining the variance of NZD exchange rate (2.62%), which is very close to Karim et al. (2007)' estimate of 2.92% (for 4-quarter forecast errors). The largest variance share of NZD exchange rate is explained by currency premium shocks (75.84%), followed by New Zealand economics news shocks (11.89%), and the US economic news shocks (9.45%). For other variables for New Zealand, I find that currency premium shocks can explain 7.32% of the variance of the short-term interest rate while the contributions of the US monetary policy and economic news shocks are very small (about 1%). However, the US economic news shocks can explain up to 17.3% of the variations of the New Zealand stock price, followed by New Zealand monetary policy shocks (7.97%) and currency premium shocks (6.75%). As expected, the shocks to New Zealand monetary policy, economic news, and NZD exchange rate attribute very little to the variances of the US variables.

4.5 Robustness check

In this section, I cross-check the baseline results by relaxing several restrictions. One of the assumptions in the benchmark model restricts the international cross-market spillover effects. This restriction, despite being intuitive and similar to Ehrmann et al. (2011) that assume no spillover effects across the US and European stock markets and interest rates, could be relaxed to allow the possible cross-market effects of the US stock price (interest rate) on New Zealand interest rate (stock price). In a sequent check, I also relax the sign restriction on the effect of exchange rate on interest rate (α_e). While the baseline model supports a negative relationship between exchange rate and interest rate - domestic currency appreciation is associated with lower interest rate – this additional check instead allows an opposite scenario to happen when the currency appreciation caused by a positive risk premium may be associated with a higher interest rate, i.e. investors switch to riskier assets rather than government bonds. As stock prices could also increase in that scenario, currency appreciation still leads to a comovement of stock price and interest rate. In any case, the main findings remain – a higher interest rate leads NZD to appreciate immediately and even stay stronger in the long-run. The sub-sections below describe the results in more detail.

4.5.1 International cross-market spillover effects

First, to allow the New Zealand interest rate (stock price) to respond to the US stock price (interest rate), I impose a tight t-distribution with 3 degrees of freedom, prior mode of zero, prior scale of 0.1, and non-restricted sign on α_{s*} - direct effect of the US stock price on the New Zealand interest rate and β_{i*} - direct effect of the US interest rate on the New Zealand stock price. The results are provided in Appendix Figures A4.1 – A4.3 for prior and posterior distributions, impulse responses, and cumulative impulse responses. Figure A4.1 includes the prior and posterior distributions of 14 contemporaneous parameters: the posterior distributions of 12 existing parameters are very similar to the baseline results and those of the two newly added parameters (α_{s*} and β_{i*}) appear very sharply peaked, even more around zero for β_{i*}. In Figure A4.2 for the impulse responses, New Zealand stock price decreases as a response to a higher US interest rate (significantly at 68% credibility set but insignificantly at 95% credibility set). New Zealand interest rate, however, shows insignificant responses to the US stock price changes. This result makes sense as New Zealand monetary policy tends to stabilize the domestic inflation rather than reflects the stock price changes in the US market. The negative responses of the New Zealand stock prices to the US monetary policy, however, are intuitive as reflecting the comovement of stock prices across the two markets. Relaxing the cross-market spillover effects also leaves the responses of other variables unchanged, except a larger appreciation of NZD (3% in the short-run and 5.4% in the long-run approximately) due to a 1% increase in New Zealand interest rate, which is statistically significant at 68% credibility set (Figures A4.2 – A4.3).

4.5.2 Effect of exchange rate on interest rate

This second check relaxes both international cross-market spillover effects and the sign restriction on the impact of exchange rate on interest rate. Specifically, apart from imposing the tight t-distribution for α_{s*} and β_{i*} as in Section 4.5.1, the parameter α_e now has an unrestricted sign. The results are included in Appendix Figures A4.4 – A4.6. Although the posterior distribution of α_e now includes positive values due to the sign relaxation, a large proportion of its posterior distribution falls in negative territory (Figure A4.4), which indicates a tiny impact of the prior distribution on the true effect of exchange rate on interest rate. The posterior distributions of other contemporaneous parameters are unchanged. Consistently, the impulse responses in the short-term and the long-term of all variables are very similar to the benchmark and the above robustness check (Figures A4.5 - A4.6). New Zealand interest rate significantly decreases as a response to exchange rate appreciation, implying the sign restriction imposed in the baseline model is strongly supported by the data. The responding magnitude of exchange rate remains the same as in the above cross-check with a 3% appreciation of NZD in the short-run and 5.4% in the long-run following a tighter monetary policy.

4.6 Conclusion

This chapter revisits an old question in the literature in examining the effects of New Zealand monetary policy shock on NZD exchange rate. By applying the Bayesian SVAR approach, I am transparent about the influence of the chosen priors on posterior distributions and impulse response functions, avoiding being too dogmatic as the traditional SVARs. I estimate a system of five variables including the US and New Zealand short-term interest rates and stock prices, and NZD exchange rate against the USD using the weekly data from 03/01/1999 to 18/09/2020. The contribution of this chapter also belongs to the specification, employing stock prices to disentangle the monetary policy shocks from other shocks that jointly drive interest rates and exchange rate. The monetary policy shocks are identified as unexpected changes in short-term interest rates. The results show that, to an unexpected increase in New Zealand short-term interest rate, the NZD appreciates immediately and keeps appreciating without a sign of "delay overshooting" at least for one year in the estimation. The findings are in line with other empirical studies with significant evidence that contradict the UIP theory prediction. The New Zealand monetary policy shocks, however, contribute very modestly to the variances of NZD exchange rate.

Table 4.1 Summary of the unit root test results

Test type		US 3-month Treasury bill rate		Log(S&P 500)		NZ 3-month Bank bill rate		Log(NZSE)		Log(NZDUSD)	
		Level	1st-difference	Level	1st-difference	Level	1st-difference	Level	1st-difference	Level	1st-difference
ADF	None		**I(0)**		**I(0)**		**I(0)**		**I(0)**		**I(0)**
	Constant	I(1)		I(1)		I(1)		I(1)		I(1)	
	Trend	**I(0)**		I(1)		I(1)		I(1)		I(1)	
PP	Constant	I(1)	**I(0)**	I(1)	**I(0)**	I(1)	**I(0)**	I(1)	**I(0)**	I(1)	**I(0)**
	Trend	**I(0)**		I(1)		I(1)		I(1)		I(1)	
KPSS	Constant	I(1)	**I(0)**	I(1)	**I(0)**	I(1)	**I(0)**	I(1)	**I(0)**	I(1)	**I(0)**
	Constant & trend	I(1)		I(1)		I(1)		I(1)		I(1)	
ZA	Constant	I(1) [15/12/2000]	**I(0)**	I(1) [12/10/2007]	**I(0)**	I(1) [03/10/2008]	**I(0)**	I(1) [19/10/2007]	**I(0)**	I(1) [27/09/2002]	**I(0)**
	Trend	I(1) [16/05/2014]		I(1) [02/02/1996]		I(1) [21/12/1990]		I(1) [23/12/2011]		I(1) [13/09/2013]	
	Constant & trend	I(1) [02/11/2007]		I(1) [01/09/2000]		I(1) [03/10/2008]		I(1) [28/12/2007]		I(1) [09/08/2002]	

Note: The variables include the US 3-month Treasury bill rate; log of the S&P 500 index; New Zealand 3-month Bank bill rate; log of New Zealand NZSE index; log of NZD nominal exchange rate against USD. Results are at 5% significance level. Suggested break dates are in square brackets. Detailed test-statistics and critical values are provided in Appendix 4 Table A4.2.

Table 4.2 Priors for the structural coefficients

Parameter	Meaning	Prior mode	Prior scale	Sign restriction
α_s, α_s^*	Effect of stock price on interest rate	0.006	0.01	+
α_e	Effect of exchange rate on interest rate	- 0.048	0.04	–
β_i, β_i^*	Effect of interest rate on stock price	- 1.423	1.6	–
β_e	Effect of exchange rate on stock price	0.2	0.4	none
$\gamma_i, -\gamma_{i*}$	Effect of interest rate on exchange rate	3.698	1.2	none
$\gamma_s, -\gamma_{s*}$	Effect of stock price on exchange rate	0.039	0.4	none
α_{i*}	Effect of foreign interest rate on domestic interest rate	0.256	0.4	+
β_{s*}	Effect of foreign stock price on domestic stock price	0.308	0.4	+

Table 4.3 Decomposition of variance of 6-month-ahead forecast errors

	US monetary policy shock	US economic news shock	NZ monetary policy shock	NZ economic news shock	Currency premium shock
US 3-month Treasury bill rate	**0.18** [98.45%] (0.16, 0.19)	**0** [0.73%] (0, 0.01)	**0** [0.24%] (0, 0)	**0** [0.31%] (0, 0)	**0** [0.27%] (0, 0)
US stock price	**0.03** [0.79%] (0.01, 0.21)	**3.39** [97.76%] (3.11, 3.66)	**0.02** [0.45%] (0, 0.04)	**0.02** [0.48%] (0, 0.05)	**0.02** [0.51%] (0, 0.05)
NZ 3-month Bank bill rate	**0** [0.6%] (0, 0)	**0** [1.06%] (0, 0)	**0.02** [89.24%] (0.01, 0.02)	**0** [1.78%] (0, 0)	**0** [7.32%] (0, 0)
NZ stock price	**0.02** [0.1%] (0.01, 0.06)	**0.5** [17.3%] (0.39, 0.63)	**0.2** [7.97%] (0.08, 0.43)	**1.54** [67.88%] (0.69, 1.87)	**0.17** [6.75%] (0.01, 1.13)
NZD exchange rate	**0.01** [0.2%] (0, 0.04)	**0.18** [9.45%] (0.13, 0.25)	**0.07** [2.62%] (0.01, 0.26)	**0.22** [11.89%] (0.01, 1.03)	**1.26** [75.84%] (0.53, 1.58)

Note: Estimated contribution of each structural shock to the 6-month-ahead median squared forecast error of each variance in bold, and expressed as a percent of total MSE in brackets. Parentheses indicate 95% credibility intervals.

Figure 4.1 The US and New Zealand interest rates and exchange rate

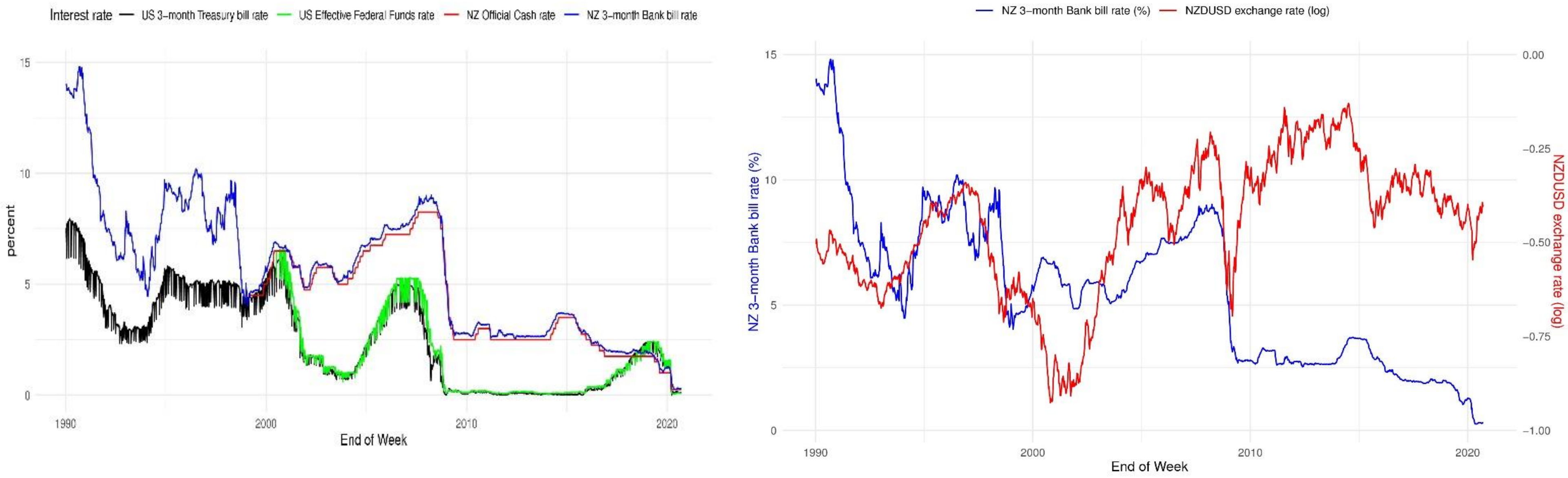


Figure 4.2 The US and New Zealand stock price indices

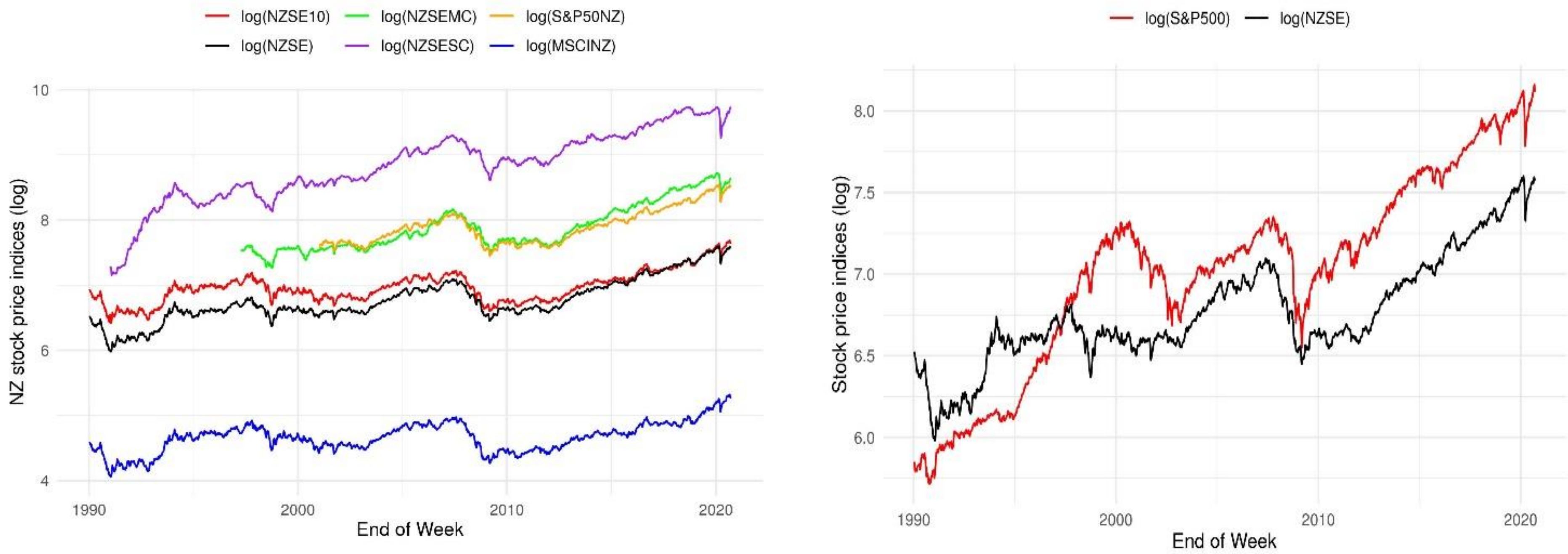


Figure 4.3 NZSE index and New Zealand exchange rate

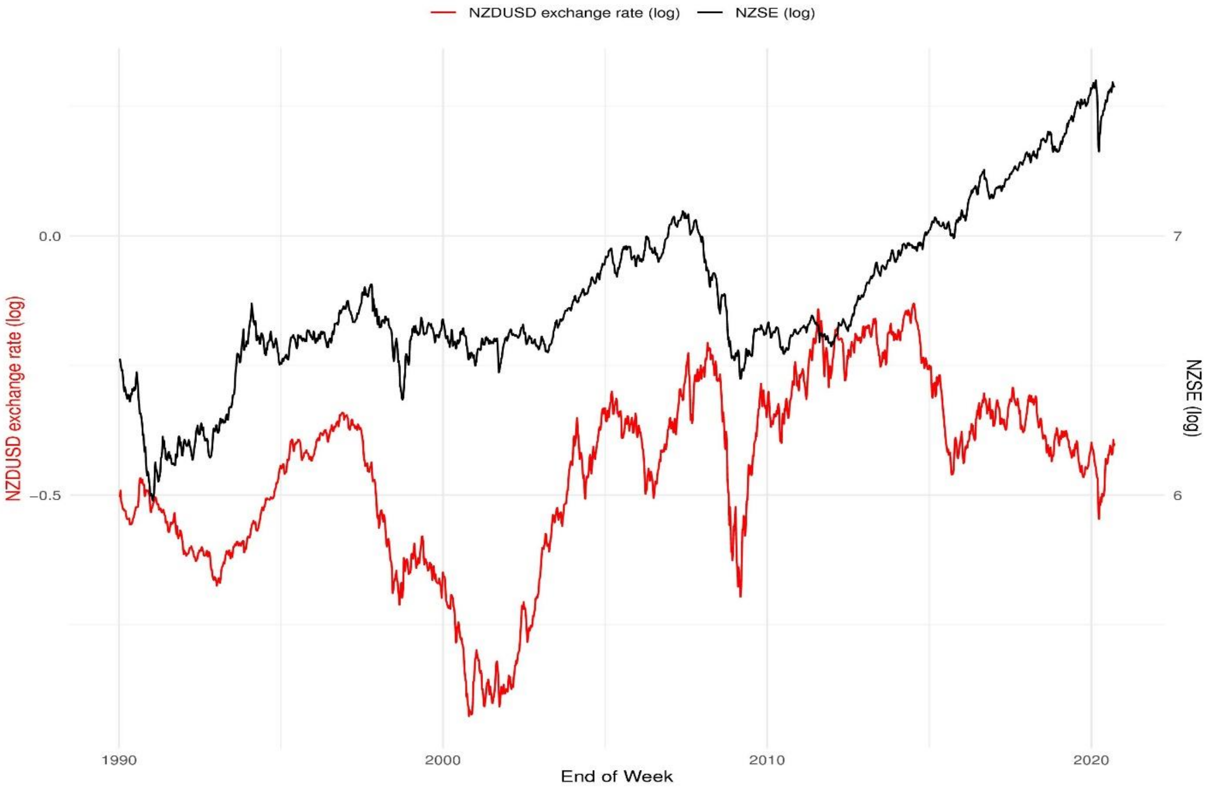
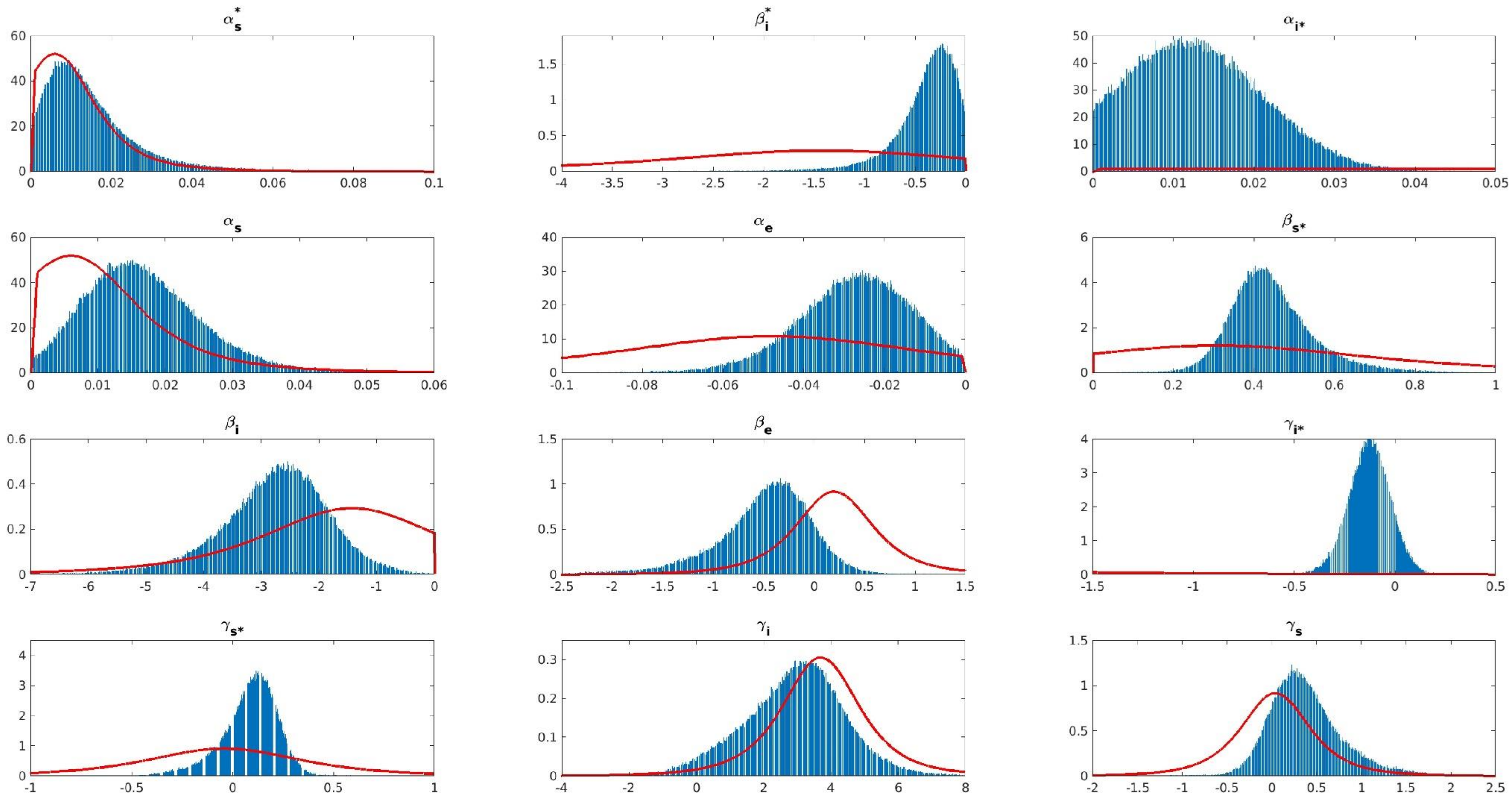
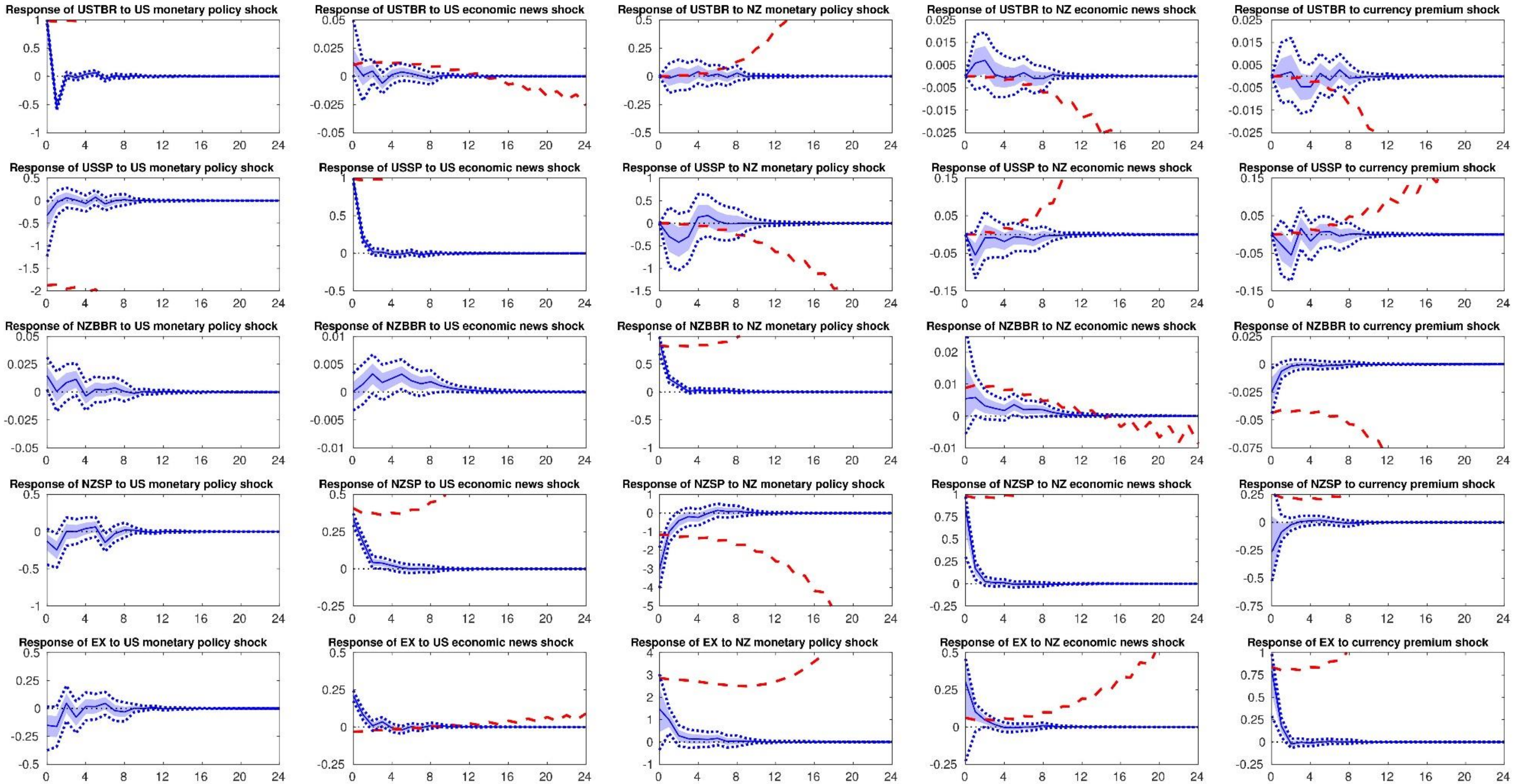


Figure 4.4 Prior and posterior distributions



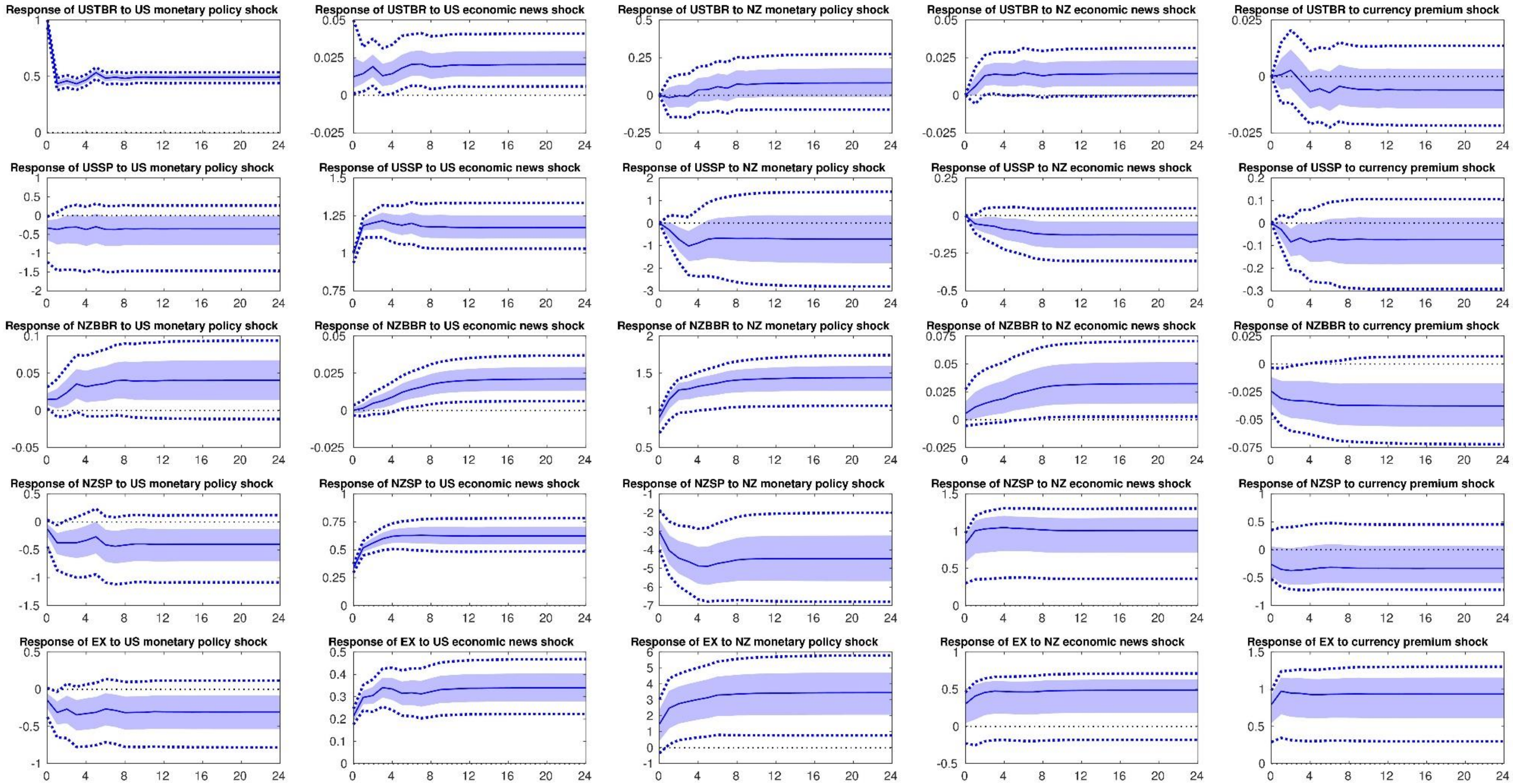
Note: Prior distribution (red lines) and posterior distributions (blue histogram) for contemporaneous coefficients.

Figure 4.5 Structural impulse-response functions and prior median



Note: Structural impulse-response functions. Solid blue lines: posterior median. Shaded region: 68% posterior credibility set. Dotted blue lines: 95% posterior credibility set. Dashed red lines: prior median. USTBR: US 3-month Treasury bill rate, USSP: US stock price, NZBBR: New Zealand 3-month Bank bill rate, NZSP: New Zealand stock price, EX: exchange rate of NZD against USD.

Figure 4.6 Cumulative impulse-response functions



Note: Cumulative impulse-response functions. Solid blue lines: posterior median. Shaded region: 68% posterior credibility set. Dotted blue lines: 95% posterior credibility set. USTBR: US 3-month Treasury bill rate, USSP: US stock price, NZBBR: New Zealand 3-month Bank bill rate, NZSP: New Zealand stock price, EX: exchange rate of NZD against USD.

Appendix 4

I present below briefly how the algorithm is implemented in this study. The fundamental prior assumptions follow Baumeister and Hamilton (2015).

The structural model has the following form:

$$\mathbf{A}\mathbf{y_t} = \mathbf{B}\mathbf{x_{t-1}} + \mathbf{u_t} \tag{A4.1}$$

for $\mathbf{y_t}$ an (*n* x 1) vector of observed variables, **A** an (*n* x *n*) matrix of the contemporaneous structural relations, $\mathbf{x_{t-1}}$ a (*k* x 1) vector (with *k* = *mn* + *1*) consisting of a constant and *m* lags of **y** ($\mathbf{x'_{t-1}} = (\mathbf{y'_{t-1}}, \mathbf{y'_{t-2}}, \ldots, \mathbf{y'_{t-m}}, \mathbf{1})'$), and $\mathbf{u_t}$ an (*n* x 1) vector of structural residuals assumed to be independent and identically distributed (i.i.d.) *N*(**0, D**) and mutually uncorrelated (**D** diagonal). In this study, I set *m* to 8 lags.

The prior information about the contemporaneous structural coefficients is represented in the form of an arbitrary prior distribution *p*(**A**), incorporating exclusion restrictions, sign restrictions, and informative prior beliefs about elements of **A**. I also assume that there are no restrictions on the lag coefficients in **B** other than the prior beliefs about the distribution *p*(**B|D, A**). Then the overall prior distribution of the parameters of the matrices **A**, **B**, and **D** is:

$$p(\mathbf{A, D, B}) = p(\mathbf{A})\, p(\mathbf{D}|\mathbf{A})\, p(\mathbf{B}|\mathbf{D, A})$$

Prior for *p*(A): I assign the prior *p*(**A**) as truncated t-distributions with 3 degrees of freedom for the unknown elements of **A**. The prior modes, scales, and sign restrictions in the baseline model are presented in Table 4.1.

Prior for *p*(D|A): The prior beliefs about structural variance reflect in part the scale of the underlying data. Let d_{ii} denote the (*i*, *i*) diagonal elements of the structural variance matrix **D** and be independent across equations, so $p(\boldsymbol{D}|\boldsymbol{A}) = \prod_{i=1}^{n} p(d_{ii}|\boldsymbol{A})$ with the reciprocals of diagonal elements of **D** following Gamma distribution $d_{ii}^{-1}|\boldsymbol{A} \sim \Gamma(\kappa, \tau_i(\boldsymbol{A}))$, where κ is the parameter and $\tau_i(\boldsymbol{A}) = \kappa a_i' \hat{S} a_i$. Here, a_i' denotes the *i*th row of **A** and $\hat{S}$ is the variance-covariance matrix of residuals from estimating the univariate AR(*m*) models for the endogenous variables. I follow Baumeister and Hamilton (2015) by setting κ = 2, which in turn puts a weight on the prior equivalent to $2\kappa = 4$ observations of data. The prior mean for d_{ii}^{-1} is κ/τ_i, which is chosen to equal the reciprocal of the *i*th diagonal element of $\boldsymbol{A}\hat{\boldsymbol{S}}\boldsymbol{A}'$. The prior variance of d_{ii}^{-1} is κ/τ_i^2.

Prior for *p*(B|D, A): Let b_i' denote the *i*th row of the matrix **B** and be independent across equations, so $p(\mathbf{B}|\mathbf{D}, \mathbf{A}) = \prod_{i=1}^{n} p(b_i|\mathbf{D}, \mathbf{A})$. Assuming that $b_i|\mathbf{D}, \mathbf{A} \sim N(m_i, d_{ii}M_i)$, so m_i denotes the prior mean for b_i and $d_{ii}M_i$ denotes the prior variance associated with this prior. I allow m_i and M_i to be functions of **A** but not of **D**. I set the prior mean $m_i = \eta' a_i$ with $\eta_{n\,\mathrm{x}\,k} = (0_{n\,\mathrm{x}\,1}\; 0_{n\,\mathrm{x}\,nm})$ because I expect all variables (in first-differences) to have zero persistence; i.e., I impose no long-run restrictions for all equations. I also put more confidence in my prior beliefs that coefficients on higher-order lags are zero, represented by smaller diagonal elements for M_i associated with higher lags. Let $v'_{1(1\,\mathrm{x}\,m)} = (\frac{1}{1^{2\lambda_1}}, \frac{1}{2^{2\lambda_1}}, \ldots, \frac{1}{m^{2\lambda_1}})$ and $v'_{2(1\,\mathrm{x}\,n)} = (s_{11}^{-1}, s_{22}^{-1}, \ldots, s_{nn}^{-1})'$ where $\sqrt{s_{ii}}$ denotes the estimated standard deviation of a univariate eighth-order autoregression fit to variable i. Then I form $v_3 = \lambda_0^2 \begin{bmatrix} v_1 \otimes v_2 \\ \lambda_3^2 \end{bmatrix}$ and the covariance matrix M_i is taken as a diagonal matrix whose row *r* column *r* element is the *r*th element of v_3: $M_{i,rr} = v_{3r}$. I set the hyperparameters following Baumeister and Hamilton (2015): $\lambda_1 = 1$ (which governs how quickly the prior for lagged coefficients tightens to zero as the lag *m* increases), $\lambda_3 = 100$ (which makes the prior on the constant term essentially irrelevant), and $\lambda_0 = 0.1$ (which summarizes the overall confidence in the prior).

I follow Baumeister and Hamilton (2015) to calculate the target function which is built based on the join distribution of **A**, **B**, and **D** conditional on the data, and use a random walk Metropolis-Hastings algorithm to generate draws of **A**, **B**, **D** from the posterior distribution, with the scalar tuning parameter set to get an acceptance ratio of 30%. In this chapter, I set the tuning parameter to 0.45 and the acceptance ratio is 29.14%. The results are based on 2,000,000 draws with 1,000,000 burn-in draws.

Table A4.1 Data sources

Variable	Source	Starting daily data
US 3-month Treasury bill rate, % per annum	https://fred.stlouisfed.org	4/01/1954 -
US Effective Fed Funds rate, % per annum	https://fred.stlouisfed.org	3/07/2000 -
S&P 500 index, nominal USD (S&P500): The stock market index that measures the stock performance of 500 large companies listed on stock exchanges in the United States.	Datastream (code: S&PCOMP)	4/03/1957 -
New Zealand 3-month Bank bill rate, % per annum	Reserve Bank of New Zealand	4/01/1985 -
New Zealand Official Cash rate, % per annum	Reserve Bank of New Zealand	17/03/1999 -
S&P/NZX 10 index (NZSE10), nominal NZD: The index measures the performance of 10 of the largest New Zealand listed companies within the S&P/NZX 50 index.	Datastream (code: NZ10CAP)	30/06/1988 -

S&P/NZX All index (NZSE), nominal NZD: The index is considered the total market indicator for the New Zealand equity market. It comprises all eligible securities quoted on the NZX Main Board. Constituents are not screened for liquidity.	Datastream (code: NZSEALL)	3/01/1990 -
S&P/NZX MidCap index (NZSEMC), nominal NZD: The index is designed to measure the performance of New Zealand's core mid-cap equity market. The eligible universe is drawn from the constituents of the S&P/NZX 50 Index, excluding those that are also constituents of the S&P/NZX 10 index.	Datastream (code: NZMCAPC)	7/04/1997 -
S&P/NZX SmallCap index (NZSESC), nominal NZD: The index is designed to measure the performance of New Zealand's smaller listed companies that sit outside of the S&P/NZX 50 index.	Datastream (code: NZSMCIC)	28/12/1990 -
S&P/NZX 50 Portfolio index (S&P50NZ), nominal NZD: The index comprises the same constituents as the S&P/NZX 50 index, but with a 5% cap on the float-adjusted market capitalization weights of the constituents.	Datastream (code: NZ50CAP)	29/12/2000 -
MSCI New Zealand index (MSCINZ), nominal NZD: The MSCINZ is designed to measure the performance of the large and mid cap segments of the New Zealand market. With 7 constituents, the index covers approximately 85% of the free float-adjusted market capitalization in New Zealand.	Datastream (code: MSNZEAL)	31/12/1981 -
Nominal exchange rate of New Zealand dollar against US dollar	Reserve Bank of New Zealand	1/06/1973 -

Table A4.2 Detailed unit root test results

Test	Type	Variable	Test statistic	1%	5%	10%	Result	Break
ADF	level, constant	USTBR	-2.69	-3.43	-2.86	-2.57	I(1)	
ADF	level, constant	log(S&P500)	-0.91	-3.43	-2.86	-2.57	I(1)	
ADF	level, constant	NZBBR	-2.33	-3.43	-2.86	-2.57	I(1)	
ADF	level, constant	log(NZSE)	0.09	-3.43	-2.86	-2.57	I(1)	
ADF	level, constant	log(NZDUSD)	-1.77	-3.43	-2.86	-2.57	I(1)	
ADF	level, trend	USTBR	-3.56	-3.96	-3.41	-3.12	I(0)	
ADF	level, trend	log(S&P500)	-1.79	-3.96	-3.41	-3.12	I(1)	
ADF	level, trend	NZBBR	-2.62	-3.96	-3.41	-3.12	I(1)	
ADF	level, trend	log(NZSE)	-1.78	-3.96	-3.41	-3.12	I(1)	
ADF	level, trend	log(NZDUSD)	-2.16	-3.96	-3.41	-3.12	I(1)	
ADF	1st-difference, none	USTBR	-44.85	-2.58	-1.95	-1.62	I(0)	
ADF	1st-difference, none	log(S&P500)	-26.24	-2.58	-1.95	-1.62	I(0)	
ADF	1st-difference, none	NZBBR	-21.77	-2.58	-1.95	-1.62	I(0)	
ADF	1st-difference, none	log(NZSE)	-24.25	-2.58	-1.95	-1.62	I(0)	
ADF	1st-difference, none	log(NZDUSD)	-26.35	-2.58	-1.95	-1.62	I(0)	
PP	level, constant	USTBR	-2.64	-3.44	-2.864	-2.57	I(1)	
PP	level, constant	log(S&P500)	-0.81	-3.44	-2.864	-2.57	I(1)	
PP	level, constant	NZBBR	-2.41	-3.44	-2.864	-2.57	I(1)	
PP	level, constant	log(NZSE)	0.16	-3.44	-2.864	-2.57	I(1)	
PP	level, constant	log(NZDUSD)	-1.68	-3.44	-2.864	-2.57	I(1)	
PP	level, trend	USTBR	-3.76	-3.97	-3.42	-3.13	I(0)	
PP	level, trend	log(S&P500)	-1.7	-3.97	-3.42	-3.13	I(1)	
PP	level, trend	NZBBR	-2.78	-3.97	-3.42	-3.13	I(1)	
PP	level, trend	log(NZSE)	-1.76	-3.97	-3.42	-3.13	I(1)	
PP	level, trend	log(NZDUSD)	-2.02	-3.97	-3.42	-3.13	I(1)	
PP	1st-difference, constant	USTBR	-95.58	-3.44	-2.86	-2.57	I(0)	
PP	1st-difference, constant	log(S&P500)	-34.38	-3.44	-2.86	-2.57	I(0)	
PP	1st-difference, constant	NZBBR	-31.65	-3.44	-2.86	-2.57	I(0)	
PP	1st-difference, constant	log(NZSE)	-30.9	-3.44	-2.86	-2.57	I(0)	
PP	1st-difference, constant	log(NZDUSD)	-31.72	-3.44	-2.86	-2.57	I(0)	
KPSS	level, constant	USTBR	12.37	0.74	0.46	0.35	I(1)	
KPSS	level, constant	log(S&P500)	15.55	0.74	0.46	0.35	I(1)	
KPSS	level, constant	NZBBR	12.4	0.74	0.46	0.35	I(1)	
KPSS	level, constant	log(NZSE)	12.56	0.74	0.46	0.35	I(1)	
KPSS	level, constant	log(NZDUSD)	8.01	0.74	0.46	0.35	I(1)	
KPSS	level, constant and trend	USTBR	0.7	0.22	0.15	0.12	I(1)	
KPSS	level, constant and trend	log(S&P500)	1.98	0.22	0.15	0.12	I(1)	

KPSS	level, constant and trend	NZBBR	0.55	0.22	0.15	0.12	I(1)	
KPSS	level, constant and trend	log(NZSE)	1.18	0.22	0.15	0.12	I(1)	
KPSS	level, constant and trend	log(NZDUSD)	1.21	0.22	0.15	0.12	I(1)	
KPSS	1st-difference, constant	USTBR	.11	0.74	0.46	0.35	I(0)	
KPSS	1st-difference, constant	log(S&P500)	.12	0.74	0.46	0.35	I(0)	
KPSS	1st-difference, constant	NZBBR	.17	0.74	0.46	0.35	I(0)	
KPSS	1st-difference, constant	log(NZSE)	.22	0.74	0.46	0.35	I(0)	
KPSS	1st-difference, constant	log(NZDUSD)	.08	0.74	0.46	0.35	I(0)	
ZA	level, constant	USTBR	-4.27	-5.34	-4.8	-4.58	I(1)	15/12/2000
ZA	level, constant	log(S&P500)	-2.77	-5.34	-4.8	-4.58	I(1)	12/10/2007
ZA	level, constant	NZBBR	-3.9	-5.34	-4.8	-4.58	I(1)	03/10/2008
ZA	level, constant	log(NZSE)	-3.3	-5.34	-4.8	-4.58	I(1)	19/10/2007
ZA	level, constant	log(NZDUSD)	-3.28	-5.34	-4.8	-4.58	I(1)	27/09/2002
ZA	level, trend	USTBR	-4.03	-4.93	-4.42	-4.11	I(1)	16/05/2014
ZA	level, trend	log(S&P500)	-2.18	-4.93	-4.42	-4.11	I(1)	02/02/1996
ZA	level, trend	NZBBR	-3.13	-4.93	-4.42	-4.11	I(1)	21/12/1990
ZA	level, trend	log(NZSE)	-2.63	-4.93	-4.42	-4.11	I(1)	23/12/2011
ZA	level, trend	log(NZDUSD)	-2.43	-4.93	-4.42	-4.11	I(1)	13/09/2013
ZA	level, constant and trend	USTBR	-4.67	-5.57	-5.08	-4.82	I(1)	02/11/2007
ZA	level, constant and trend	log(S&P500)	-3.36	-5.57	-5.08	-4.82	I(1)	01/09/2000
ZA	level, constant and trend	NZBBR	-3.89	-5.57	-5.08	-4.82	I(1)	03/10/2008
ZA	level, constant and trend	log(NZSE)	-4.86	-5.57	-5.08	-4.82	I(1)	28/12/2007
ZA	level, constant and trend	log(NZDUSD)	-3.36	-5.57	-5.08	-4.82	I(1)	09/08/2002
ZA	1st-difference, constant	USTBR	-44.96	-5.34	-4.8	-4.58	I(0)	
ZA	1st-difference, constant	log(S&P500)	-26.81	-5.34	-4.8	-4.58	I(0)	
ZA	1st-difference, constant	NZBBR	-22.17	-5.34	-4.8	-4.58	I(0)	
ZA	1st-difference, constant	log(NZSE)	-24.76	-5.34	-4.8	-4.58	I(0)	
ZA	1st-difference, constant	log(NZDUSD)	-26.57	-5.34	-4.8	-4.58	I(0)	

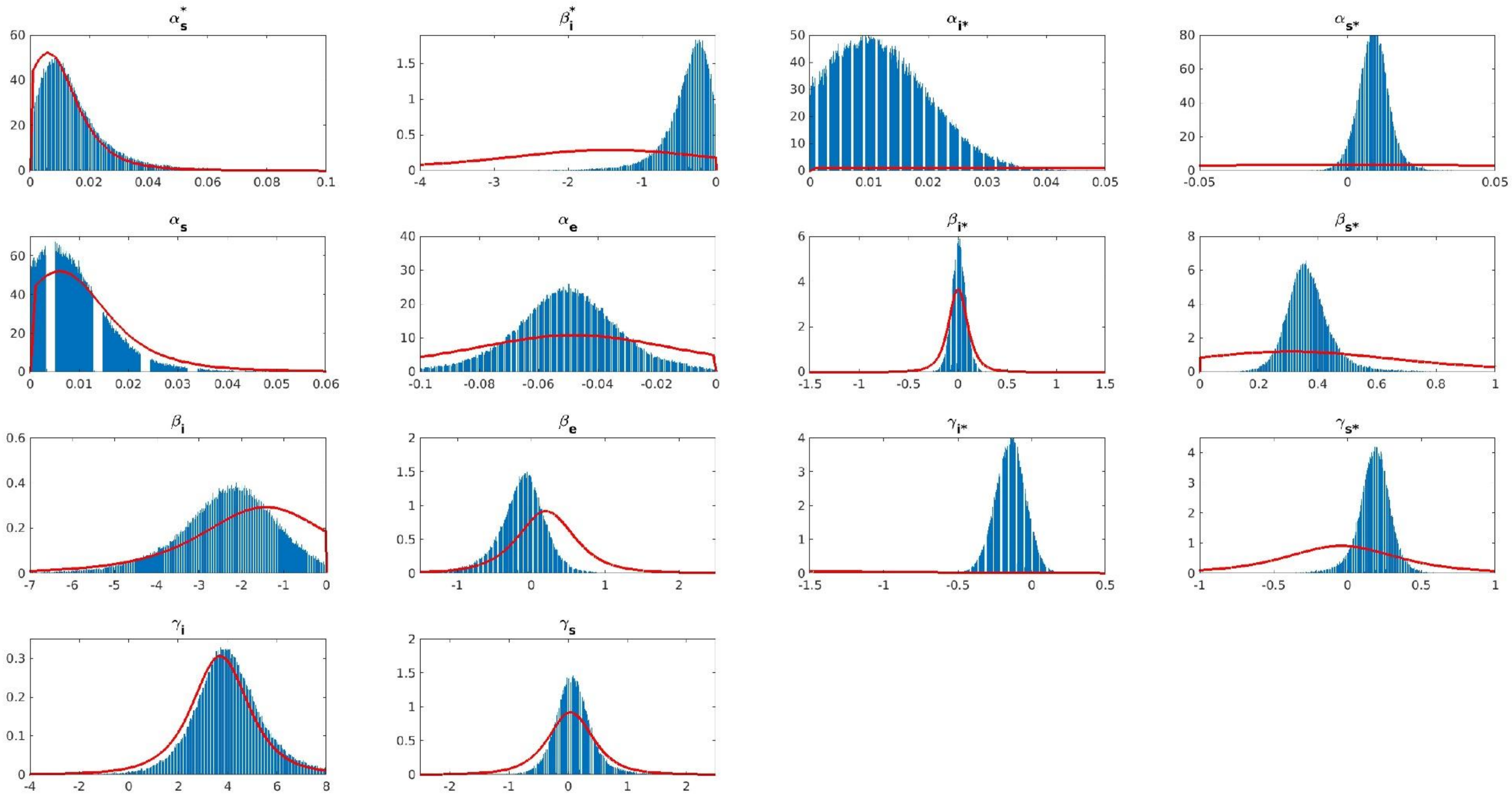
Note: Table reports the test statistics and critical values at 1%, 5%, and 10% levels of significance in the ADF, the PP, the KPSS, and the ZA unit root tests for five series, including US 3-month Treasury bill rate (USTBR), natural log of the S&P500 index (log(S&P500)), New Zealand 3-month Bank bill rate (NZBBR), natural log of New Zealand stock price index (log(NZSE)), and natural log of nominal bilateral exchange rate of NZD against USD (log(NZDUSD)). The sampled period is 12/01/1990 – 27/12/2019 using all available data. Column Result indicates the series is non-stationarity (I(1)) or stationary (I(0)) at 5% level of significance. The last column reports the suggested break dates in the ZA test. (1) ADF test: Null hypothesis: Series is non-stationary; Alternative hypothesis: Series is stationary. Reject the null hypothesis if test-statistic is larger than the critical values at 5% significance in absolute term. (2) PP test: Null hypothesis: Series is non-stationary; Alternative hypothesis: Series is stationary. Reject the null hypothesis if test-statistic is larger than the critical values at 5% significance in absolute term. (3) KPSS test: Null hypothesis: Series is stationary; Alternative hypothesis: Series is non-stationary. Reject the null hypothesis if test-statistic is larger than the critical values at 5% significance in absolute term. (4) ZA test: Null hypothesis: the series has a unit root with structural break(s); Alternative hypothesis: Series are stationary with break(s). Reject the null hypothesis if t-value statistic is lower than the critical value at 5% significance (left tailed test).

Table A4.3 Short-run impact of NZD exchange rate on New Zealand stock price: Simple OLS estimations

Dependent variable: ΔLNZSE	(1)	(2)	(3)	(4)
L.ΔLNZSE	0.251***	0.249***	0.250***	0.241***
	(0.040)	(0.040)	(0.040)	(0.040)
L2.ΔLNZSE	-0.026	-0.028	-0.027	-0.029
	(0.037)	(0.037)	(0.037)	(0.037)
L3.ΔLNZSE	0.026	0.024	0.024	0.022
	(0.033)	(0.033)	(0.033)	(0.033)
L4.ΔLNZSE	0.017	0.015	0.015	0.014
	(0.032)	(0.032)	(0.032)	(0.032)
L5.ΔLNZSE	-0.056**	-0.058**	-0.057**	-0.057**
	(0.028)	(0.028)	(0.028)	(0.028)
L.ΔLNZDUSD	0.052*	0.052*	0.052*	
	(0.032)	(0.031)	(0.032)	
ΔLNZDUSD	**0.218*** **	**0.217*** **	**0.217*** **	**0.210*** **
	(0.042)	(0.042)	(0.042)	(0.040)
D1		-0.002		
		(0.001)		
D2			-0.002	
			(0.001)	
D3				0.003**
				(0.001)
Time trend		0.000	0.000	-0.000
		(0.000)	(0.000)	(0.000)
Observations	1596	1596	1596	1596
R-squared	0.100	0.102	0.101	0.101

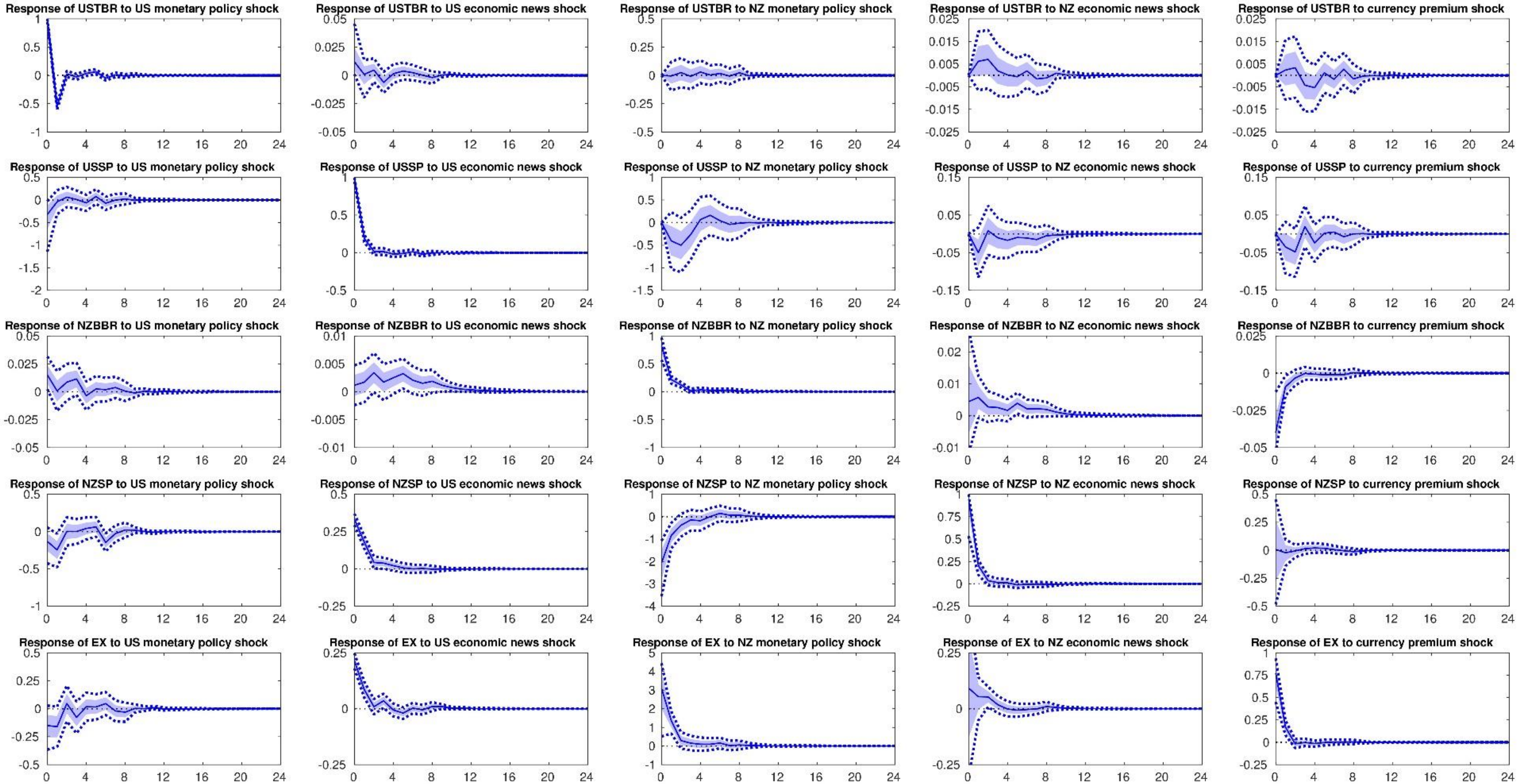
Note: Δ*LNZSE* and Δ*LNZDUSD* are log-differences of the New Zealand stock price and exchange rate of NZD against USD; the dummy variables: *D1* gets 1 since 19/10/2007 and 0 otherwise, *D2* gets 1 since 28/12/2007 and 0 otherwise, or *D3* gets 1 since 23/12/2011 and 0 otherwise. The variables *L.ΔLNZSE, L2.ΔLNZSE*, etc. are the lags of dependent variables. Standard errors are in parentheses. ***, **, *: 1%, 5%, 10% significance.

Figure A4.1 Allowing international cross-market effects: Prior and posterior distributions



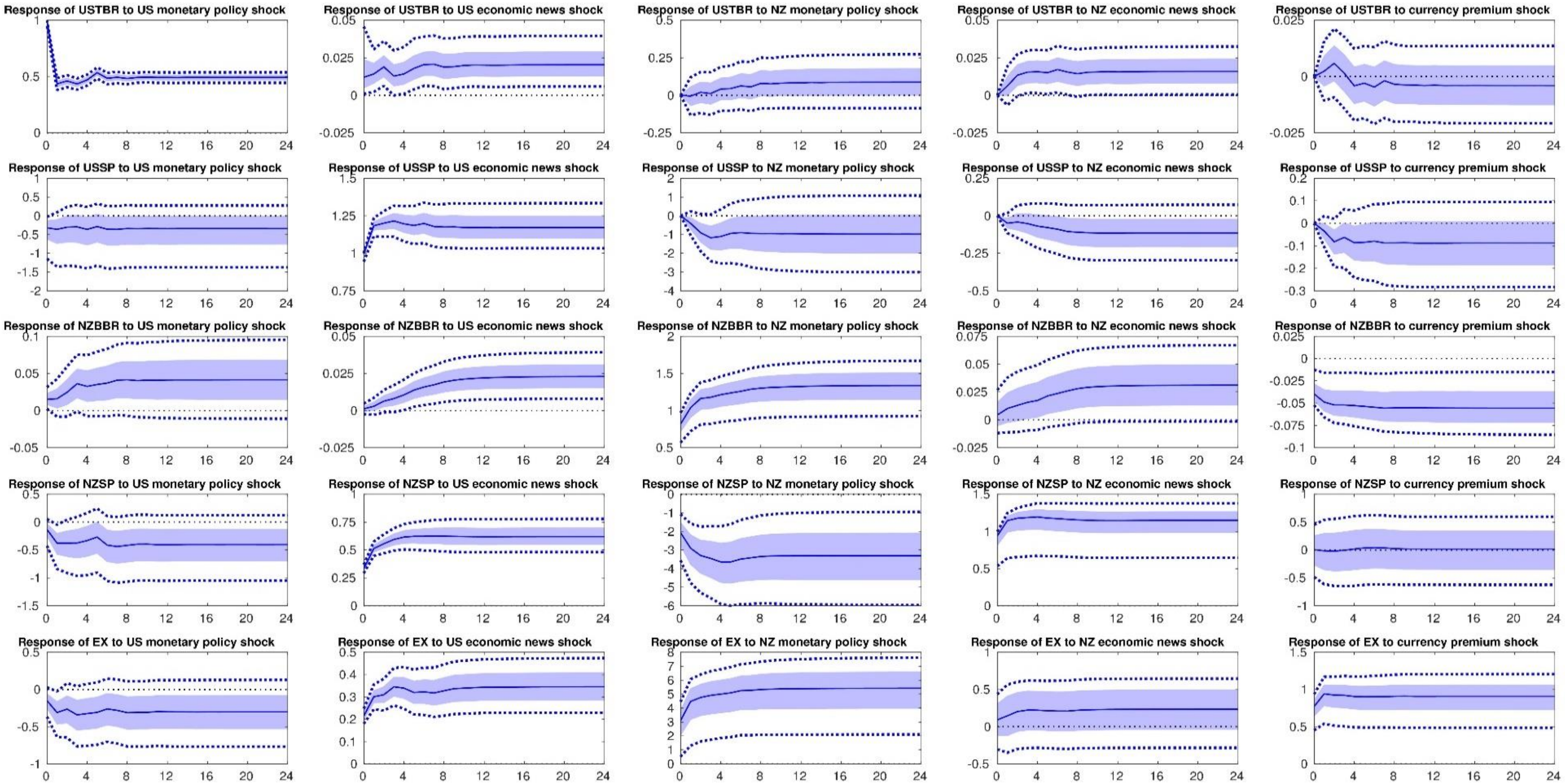
Note: Prior distribution (red lines) and posterior distributions (blue histogram) for contemporaneous coefficients.

Figure A4.2 Allowing international cross-market effects: Structural impulse-response functions and prior median



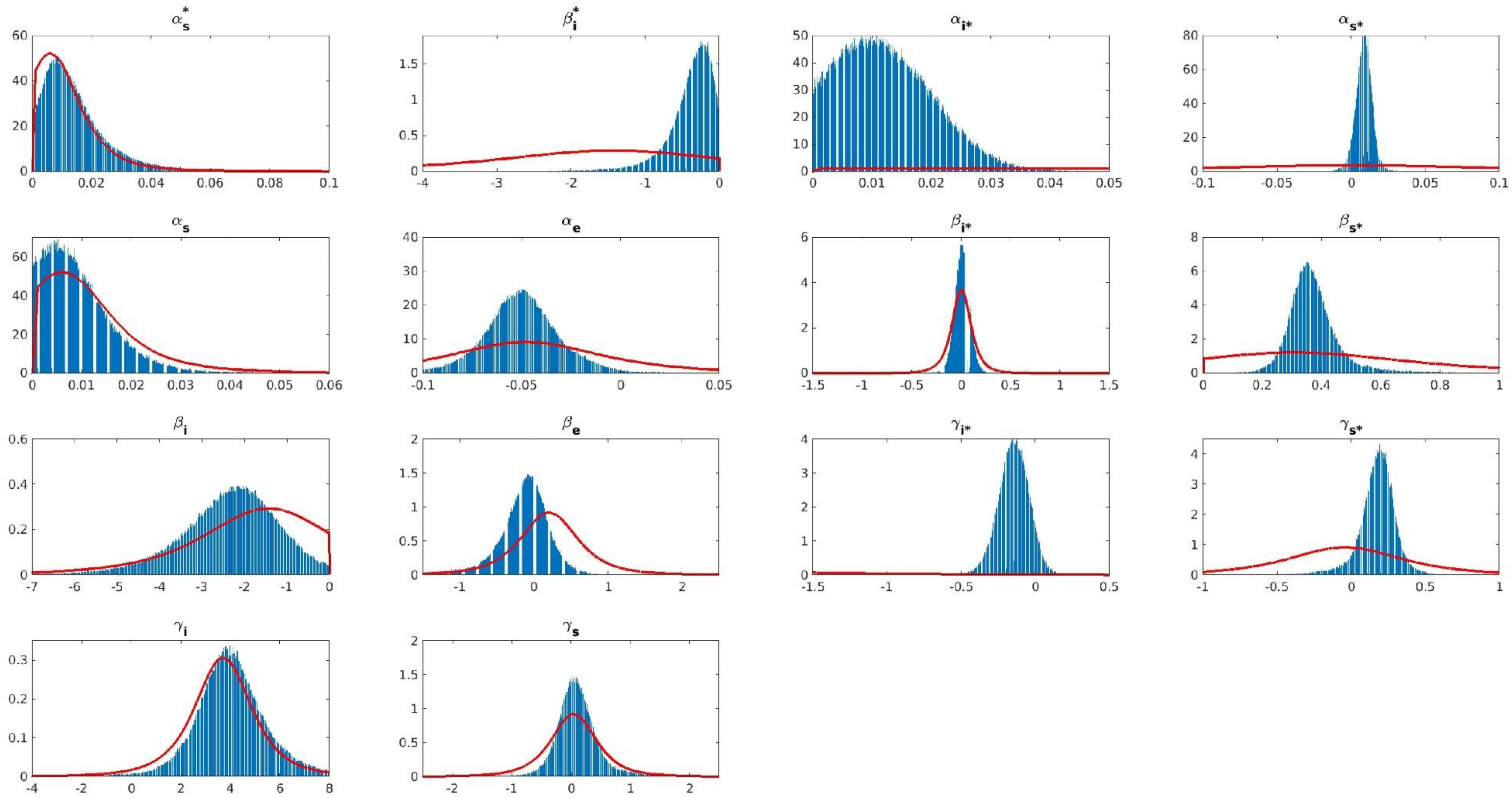
Note: Structural impulse-response functions. Solid blue lines: posterior median. Shaded region: 68% posterior credibility set. Dotted blue lines: 95% posterior credibility set. USTBR: US 3-month Treasury bill rate, USSP: US stock price, NZBBR: New Zealand 3-month Bank bill rate, NZSP: New Zealand stock price, EX: exchange rate of NZD against USD.

Figure A4.3 Allowing international cross-market effects: Cumulative impulse-response functions



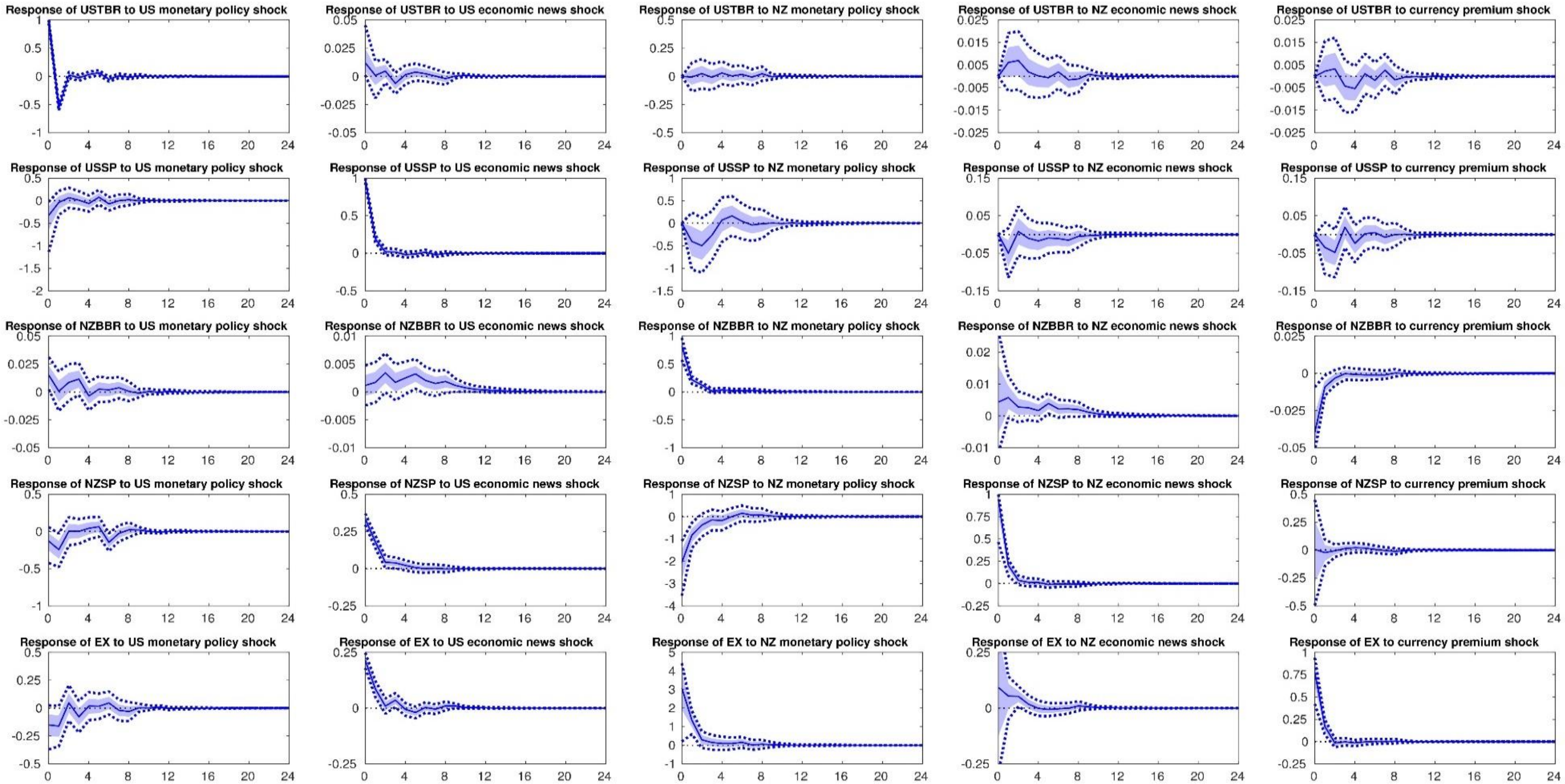
Note: Cumulative impulse-response functions. Solid blue lines: posterior median. Shaded region: 68% posterior credibility set. Dotted blue lines: 95% posterior credibility set. USTBR: US 3-month Treasury bill rate, USSP: US stock price, NZBBR: New Zealand 3-month Bank bill rate, NZSP: New Zealand stock price, EX: exchange rate of NZD against USD.

Figure A4.4 Allowing international cross-market effects and relaxing effect of exchange rate on interest rate: Prior and posterior distributions



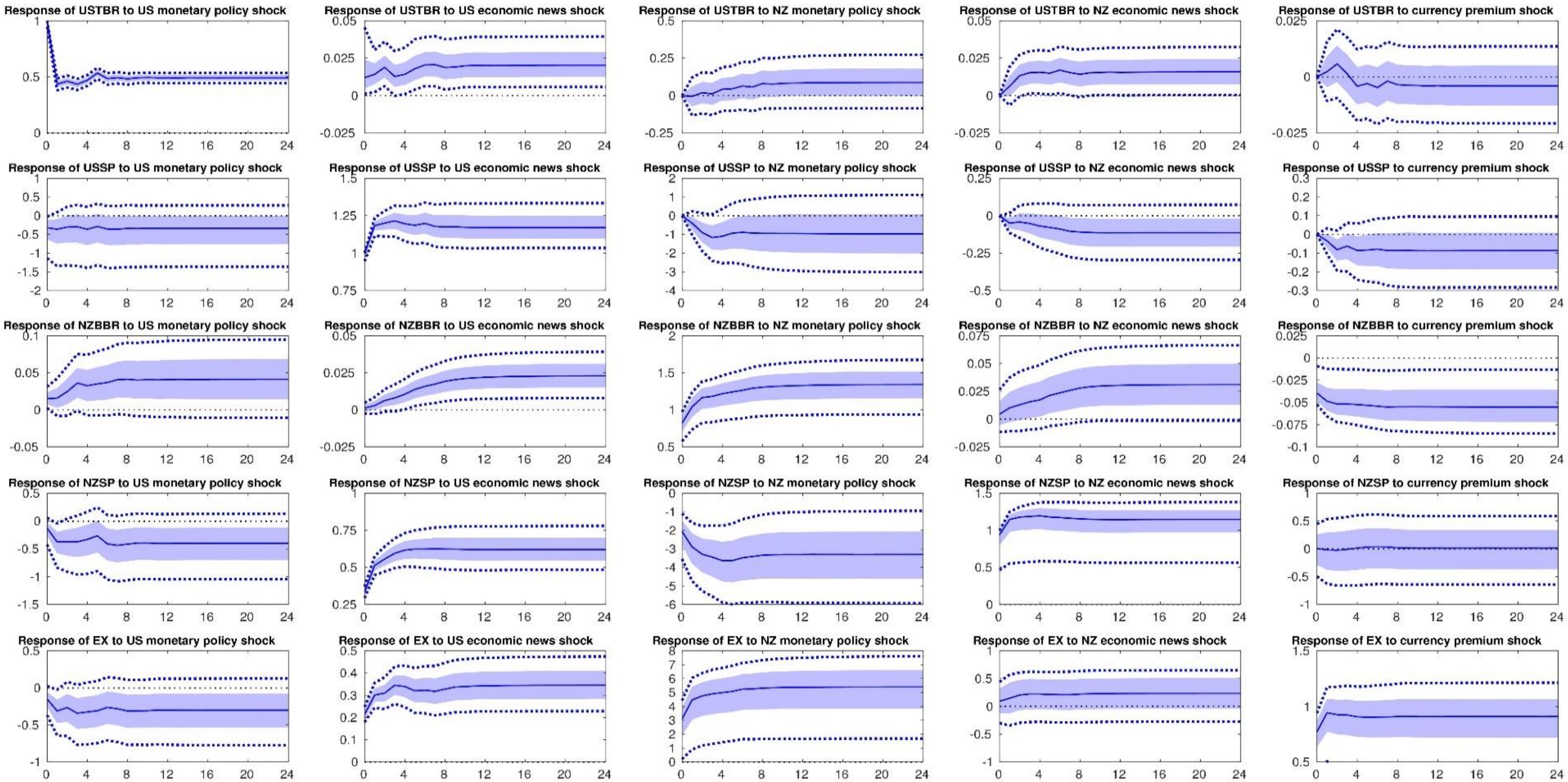
Note: Prior distribution (red lines) and posterior distributions (blue histogram) for contemporaneous coefficients.

Figure A4.5 Allowing international cross-market effects and relaxing effect of exchange rate on interest rate: Structural impulse-response functions and prior median



Note: Structural impulse-response functions. Solid blue lines: posterior median. Shaded region: 68% posterior credibility set. Dotted blue lines: 95% posterior credibility set. USTBR: US 3-month Treasury bill rate, USSP: US stock price, NZBBR: New Zealand 3-month Bank bill rate, NZSP: New Zealand stock price, EX: exchange rate of NZD against USD.

Figure A4.6 Allowing international cross-market effects and relaxing effect of exchange rate on interest rate: Cumulative impulse-response functions



Note: Cumulative impulse-response functions. Solid blue lines: posterior median. Shaded region: 68% posterior credibility set. Dotted blue lines: 95% posterior credibility set. USTBR: US 3-month Treasury bill rate, USSP: US stock price, NZBBR: New Zealand 3-month Bank bill rate, NZSP: New Zealand stock price, EX: exchange rate of NZD against USD.

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