Essays on Fiscal Policy

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Abstract

This dissertation examines how the effects of fiscal shocks vary across countries and economic conditions. The three essays analyze fiscal policy transmission in developing, resource-dependent, and advanced economies. The first chapter focuses on the time-varying effects of government expenditure shocks in six WAEMU countries using a Bayesian TVP-SVAR model with stochastic volatility. The findings show that government spending generally supports output, but the strength and persistence of this effect differ across countries. This indicates that even within a monetary union, fiscal policy effectiveness depends importantly on country-specific structural conditions, making fiscal design especially important where monetary autonomy is limited.

The second chapter investigates the effects of oil revenue shocks in four African oil-exporting economies using a hierarchical Bayesian panel VAR and country-specific structural VAR models. The findings show that oil revenue windfalls generate short-run gains in government expenditure and output, but these effects fade within a few years and are accompanied by inflationary pressures and real exchange rate appreciation. This indicates that oil-driven fiscal expansions provide only temporary macroeconomic support and underscores the importance of institutions that smooth revenue volatility and strengthen resilience to external shocks.

The third chapter studies whether the effects of government spending shocks on business investment depend on macroeconomic uncertainty in seven OECD countries using a threshold panel VAR. The results suggest that fiscal expansions crowd out business investment in low-uncertainty periods but have a more muted effect when uncertainty is high, indicating that private sector responses to fiscal policy may be state-dependent. More broadly, the findings are consistent with the view that the investment effects of fiscal policy depend not only on the policy shock itself, but also on the prevailing economic environment.

Overall, this thesis shows that the effectiveness of fiscal policy cannot be reduced to

a single multiplier or uniform transmission mechanism. Instead, the fiscal multiplier varies with structural characteristics and prevailing economic conditions. Taken together, the three essays provide evidence that institutional arrangements, resource dependence, and economic uncertainty influence the transmission of fiscal policy. Collectively, the findings suggest that the effects of fiscal policy are state-dependent, underscoring the importance of accounting for economic and institutional conditions when designing fiscal interventions rather than relying on uniform policy prescriptions.

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Declaration of Authorship

I hereby certify that I am the author of this thesis. The first chapter was co-authored with my supervisor, Dr. Yazid Dissou.

This entire work was done during the PhD program, and I have not submitted any part of this thesis for any other degree. Where the published work of other authors has been consulted, I have made the necessary citations.

Dedicated to my precious daughters, Dayspring and Charis.

You are the future and that future is bright!

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Chapter 1

Government Spending and Output Dynamics in WAEMU Countries: A Time-Varying Structural VAR Approach

Abstract

In this paper, we investigate the time-varying impact of government expenditure on output in six member countries of the West African Economic and Monetary Union (WAEMU), which share a currency and rely heavily on fiscal measures for macroeconomic stabilization. Using a Bayesian Time-Varying Parameter Structural VAR (TVP-SVAR) model with stochastic volatility on annual data, we estimate both impulse responses and fiscal multipliers associated with government consumption shocks. Unlike advanced economies, WAEMU countries face structural constraints, including fixed exchange rates, limited monetary autonomy, and weak fiscal capacity. Our study finds that fiscal shocks produce positive effects on GDP, with their magnitudes and persistence varying across countries. Senegal displays stable and relatively large fiscal multipliers, whereas countries like Togo and Benin exhibit smaller and more volatile responses. We also find that price level responses to fiscal shocks are short-lived, likely constrained by the WAEMU's fixed exchange rate regime. Our study highlights the importance of designing fiscal policy that responds to context,

especially in developing economies within monetary unions.

1.1 Introduction

A common view in the literature is that fiscal policy plays a fundamental role in macroeconomic stabilization and economic growth, especially in developing countries where structural constraints limit the effectiveness of intervention by monetary policy. This is particularly true for the West African Economic and Monetary Union (WAEMU), in which eight countries—Benin, Burkina Faso, Côte d'Ivoire, Guinea-Bissau, Mali, Niger, Senegal, and Togo—share a common monetary policy and operate under a fixed exchange rate regime. Under these circumstances, fiscal policy remains the most important instrument for managing economic fluctuations and stimulating growth. In this context, the key question is whether the impact of government spending on output has remained constant over time or has varied in response to changing economic conditions, policy frameworks, and external factors. While this question is important for understanding the historical evolution of fiscal policy effectiveness in the WAEMU region, the analysis here is intended as a retrospective assessment rather than a tool for forecasting or designing future policy interventions. The focus is on interpreting how past fiscal shocks have propagated through the economy, providing insights into the dynamics of fiscal transmission over time.

Despite an extensive literature, there is no clear consensus on how fiscal policy instruments affect macroeconomic aggregates. Theoretical views on fiscal multipliers—the change in output resulting from a one-unit change in government spending or taxation—differ significantly. Keynesian models emphasize the stabilizing role of government intervention, particularly during recessions, when private demand is insufficient to maintain full employment and output. In these frameworks, fiscal expansion can offset declines in private spending and smooth business cycle fluctuations (Keynes, 1936; Blinder, 1987; Auerbach and Gorodnichenko, 2012b; Christiano et al., 2011b). In contrast, neoclassical models suggest that government spending crowds out private consumption by absorbing resources that would otherwise be available to the private sector (Barro, 1974; Aiyagari et al., 1992; Baxter and King, 1993; Blanchard and Perotti, 2002). The development of New Keynesian models aimed to reconcile these differences (Rotemberg and Woodford, 1997; Clarida et al., 1999; Woodford, 2003), yet empirical findings remain mixed (Ramey, 2011a; Coenen et al., 2012; Leeper et al., 2017). To better assess the impact of fiscal policy, empirical studies increasingly rely on structural methodologies, such as Structural Vector Autoregressions (SVARs), which incorporate institutional information to identify fiscal shocks (Blanchard and Perotti, 2002).

Most of the models used to estimate fiscal multipliers assume a stable relationship between government spending and output, which neglects how institutional and macroeconomic conditions may influence the effectiveness of fiscal policy. The results of a growing number of studies indicate that fiscal multipliers are highly state-dependent. For example, Auerbach and Gorodnichenko (2012b) and Ramey and Zubairy (2018) point out that these multipliers fluctuate in response to many factors such as business cycles, monetary policy stances, debt sustainability and external shocks. This issue is particularly relevant in WAEMU economies, where the lack of independent monetary policy due to the common currency arrangement places a greater emphasis on fiscal measures for economic stabilization.

Unlike advanced economies, where central banks can implement countercyclical policies to complement fiscal efforts, policymakers in the WAEMU region depend mainly on government spending to manage demand and stimulate growth. Understanding whether fiscal multipliers remain stable or evolve over time provides insights into the effectiveness of fiscal interventions in the region. Adding to the complexity of the fiscal landscape is the growing concern over rising public debt. In recent years, several governments in the region have implemented expansionary fiscal strategies, which have led to a sustained increase in debt levels. Moreover, as noted in IMF (2022), other external factors — such as fluctuating commodity prices, global financial conditions and economic disruptions such as the COVID-19 pandemic — can significantly impact fiscal policy by modifying its transmission mechanisms. All these factors call into question the conventional assumption of constant fiscal multipliers and emphasize the need for a time-varying approach to their estimation using Structural Vector Autoregression (SVAR) models.

Most of the existing empirical studies focus on economies characterized by strong fiscal institutions, deep financial markets, and flexible monetary policies. However, such assumptions are not always valid in the context of developing countries, where factors such as weak fiscal capacity, external financing constraints, informality and monetary rigidities affect the transmission of fiscal policy (Gavin and Perotti, 1997; Ilzetzki et al., 2013). As a result, fiscal multipliers in these economies may be smaller and more volatile than those observed in advanced economies. This shows the need for a region-specific empirical analysis (Eden and Kraay, 2014; Ramey, 2019).

Recent advances in econometric methods, such as structural vector autoregressions with time-varying parameters (TVP-SVARs), offer new possibilities for examining how fiscal multipliers evolve over time. While this approach has been increasingly applied in monetary economics (Cogley and Sargent, 2005; Primiceri, 2005), its use in fiscal policy analysis is limited, particularly in low-income economies. Existing studies, such as Pereira and Lopes, 2014 and Muraraşu et al., 2023, have explored time-varying fiscal multipliers in the context of the United States and Europe, but there is little research on developing economies. In this study, we contribute to the literature by providing the first empirical analysis of time-varying fiscal multipliers in WAEMU countries.

Although the WAEMU comprises eight member states, our empirical analysis focuses on six: Benin, Burkina Faso, Mali, Niger, Senegal, and Togo. Two countries–Guinea-Bissau and Côte d'Ivoire–were excluded for methodological reasons. Guinea-Bissau was omitted due to limited data availability and its relatively recent accession to the Union, which constrains comparability. Côte d'Ivoire was excluded from the main estimation because the unit root tests failed to reject the null hypothesis of a unit root in the first-differenced GDP series, indicating that the series may be integrated of order two (I(2)), whereas the remaining variables appear to be integrated of order one (I(1)). To preserve a consistent econometric specification across the main sample, Côte d'Ivoire was therefore excluded from the baseline estimation. Nevertheless, given Côte d'Ivoire's economic importance within the WAEMU, the estimation results are reported in Appendix A.7 for reference.

This paper contributes to the literature on fiscal policy in several ways. First, it provides new evidence on the time-varying nature of fiscal multipliers in the WAEMU region, where empirical research remains scarce despite the central role of fiscal policy in the absence of independent monetary policy. Second, it applies a Bayesian time-varying parameter structural VAR (TVP-SVAR) with stochastic volatility to a set of low-income economies, thereby extending a methodology that has been used predominantly in advanced economies to a developing-country context. Third, the paper documents substantial cross-country heterogeneity in both the magnitude and persistence of fiscal effects within a monetary union, highlighting the importance of structural differences even under a common policy framework. Finally, by estimating a time-varying framework, the analysis assesses whether fiscal transmission has evolved over time and finds relatively limited evidence of systematic temporal variation in fiscal multipliers. Taken together, these findings suggest that fiscal multipliers in the WAEMU are relatively stable over time, despite changes in economic conditions, policy environments, and external shocks, while differing substantially across member countries.

The remainder of the paper is structured as follows: Sections 2 and 3 review the empirical and theoretical literature, Section 4 provides a brief economic summary of the countries in the study, Section 5 outlines the empirical strategy, Section 6 presents the results, Section 7 deals with robustness checks, and Section 8 concludes.

1.2 Literature Review

Structural VAR (SVAR) models are widely employed in estimating fiscal multipliers. Unlike standard Vector Autoregressive (VAR) models, SVARs incorporate theoretical or institutional restrictions to identify fiscal shocks more precisely. For instance, Blanchard and Perotti (2002) use features of the fiscal framework governing taxes and public transfers to estimate automatic stabilizer elasticities, which allows them to isolate fiscal policy shocks. However, as noted by Caggiano et al. (2015), a key limitation of traditional SVAR models is their assumption of linearity, which prevents them from capturing state-dependent multipliers, such as variations in fiscal multipliers across different phases of the business cycle.

Since fiscal multipliers tend to be larger during recessions than expansions, traditional linear models may not fully capture their dynamics. To account for this state-dependent nature, some authors have developed nonlinear VAR models such as the smooth transition VAR (STVAR) model (Auerbach and Gorodnichenko, 2012b). Caggiano et al. (2015) refine the STVAR model further by differentiating business cycle phases into mild recessions, deep recessions, weak expansions, and strong expansions. Their findings reveal that fiscal multipliers exhibit the strongest countercyclical behavior when comparing deep recessions with strong expansions.

The TVP-VAR model accounts for structural changes by allowing coefficients and error variances to evolve gradually over time. Pioneered by Cogley and Sargent (2001) and Primiceri (2005) in evaluating monetary policy, TVP-VAR models have shown evidence of time variation in inflation and unemployment dynamics. Applications to fiscal policy include Pereira and Lopes (2014) who employ TVP-VAR to analyze the U.S. economy. Their results point to a weakening influence of fiscal policy over time, particularly for tax shocks. More recently, Muraraşu et al. (2023) applied this framework to assess fiscal responses to COVID-19 in Central and Eastern European countries.

An alternative approach to examining changes in fiscal transmission is to estimate separate fixed-coefficient SVARs over different sub-samples (such as before and after the 1994 CFA franc devaluation) and compare the resulting impulse responses. While such an approach is computationally simpler, it requires the timing of structural breaks to be specified ex ante and assumes that changes in the transmission mechanism occur discretely. In contrast, the Bayesian TVP-SVAR allows the coefficients and shock variances to evolve gradually over time without imposing predetermined break dates. This flexibility is particularly relevant in the WAEMU context, where institutional reforms, fiscal frameworks, and macroeconomic conditions are likely to have evolved progressively rather than instantaneously. Moreover, by exploiting information from the full sample, the TVP-SVAR avoids estimating separate models on relatively short sub-samples and produces a continuous time path of impulse responses and fiscal multipliers, thereby providing a richer characterization of the evolution of fiscal policy effectiveness.

Several studies have explored factors influencing fiscal multipliers, including monetary accommodation, exchange rate regimes, development levels, and indebtedness. Ilzetzki et al. (2013) analyze high-income and developing countries separately and find significant differences in their responses to a 1% increase in government spending. High-income countries exhibit a 0.08% positive response in output, but developing countries show a negative 0.01% response. In high-income countries, fiscal shocks have minimal exchange rate effects, whereas in developing economies, fiscal policy significantly depreciates the exchange rate. They estimate a fiscal multiplier of 0.39 for high-income countries and -0.03 for developing economies, with cumulative long-run multipliers of 0.66 and -0.63, respectively.

Exchange rate regimes also significantly affect fiscal multiplier estimates. Born et al. (2013) find that in response to fiscal shocks, countries with fixed exchange rate regimes experience a 1.5% increase in output, whereas those with flexible exchange rates see a more modest 0.5% rise. Additionally, under fixed exchange rates, the real exchange rate tends to appreciate, while it depreciates in economies with flexible exchange rates. Similarly, Ilzetzki et al. (2013) find that the effects of fiscal policy are stronger under predetermined exchange rates, with multipliers starting at 0.15 and rising to 1.4 in the long run, while remaining negative under flexible exchange rate regimes. These findings are consistent with the predictions of the Mundell-Fleming model, which suggests that fiscal policy is more effective in economies with fixed exchange rates and limited capital mobility.

Countries within a currency union typically operate under fixed exchange rate regimes, which, in theory, should lead to larger fiscal multipliers. Several studies have examined optimal fiscal policy responses in such settings. For instance, Gali and Monacelli (2008) argue that local government spending plays a stabilizing role by helping to maintain union-wide price stability during localized demand shocks, while Cook and Devereux (2019) find that fiscal policy should be directed toward the most severely affected regions to maximize its effectiveness. However, research on developing country currency unions remains limited, with most studies focused on the Eurozone (Afonso and Leal, 2019) or states in the U.S. (Nakamura and Steinsson, 2014). Our study adds to this body of research by analyzing fiscal multipliers within the West African Economic and Monetary Union (WAEMU), a developing-country currency union characterized by fixed exchange rates and diverse economic structures.

1.3 Economic Overview of Countries in the Study

This section provides a brief economic overview of the six West African countries under study: Benin, Burkina Faso, Mali, Niger, Senegal, and Togo. These countries share several economic characteristics, including reliance on agriculture, commodity exports, and membership in the West African Economic and Monetary Union (WAEMU). The following description highlights the general patterns observed in the deflator, GDP, and government expenditure series over the period 1970–2020, providing background for the econometric analysis that follows.

Figure 1.1 shows substantial common movements in the GDP deflator across the WAEMU countries over the sample period. Following periods of rising prices during the late 1970s and late 1980s to early 1990s, all countries experienced sustained declines in the GDP deflator during the mid-1990s, indicating periods of pronounced disinflation and, in several countries, declining price levels. Thereafter, price levels evolved more gradually before increasing again from the mid-2000s onward. Mali exhibits the largest fluctuations, whereas Senegal displays a more gradual evolution, with its GDP deflator in 2020 remaining close to its early-1990s level. The common

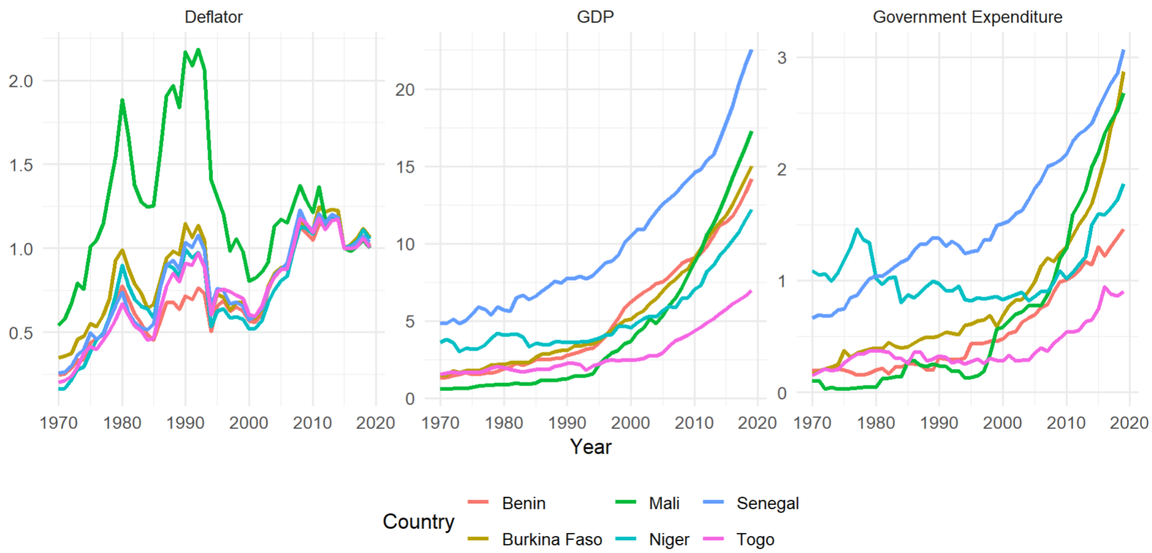


Figure 1.1: GDP and Government Expenditure are expressed in billions of U.S. dollars at constant 2015 prices; Price Level is measured by the GDP deflator (base year = 2015). All series are plotted in levels.

timing of these movements across countries suggests the influence of common regional and external factors rather than purely country-specific developments. This pattern is consistent with the shared monetary and exchange-rate framework of the WAEMU, together with common exposure to external shocks such as commodity-price and terms-of-trade movements, although identifying the precise drivers of these price dynamics is beyond the scope of the present study.

GDP values increase steadily across all countries, though at different scales. Senegal records the highest GDP levels throughout the period, while Togo remains the lowest. Mali, Burkina Faso, Benin, and Niger occupy intermediate positions, with gradual and continuous growth visible from the mid-1990s onward. The upward trend in all countries becomes more pronounced after 2000, with Senegal showing the steepest rise, followed by Mali and Burkina Faso. Despite differences in magnitude, the general pattern across the group is one of persistent output growth over the study period.

Government expenditure follows an overall increasing trend for all six countries. Senegal shows the largest and most sustained rise, while Togo and Benin remain relatively lower throughout most of the period. Mali, Burkina Faso, and Niger display gradual increases that become more noticeable after 2000. In the earlier decades, fluctuations are more visible in some countries, but from the 2000s onward, spending rises consistently across the region. The general pattern suggests a steady expansion of government expenditure in line with growing economic activity over time.

Taken together, the plots indicate that while WAEMU countries share broadly similar long-term movements in prices, output, and government spending, there are clear differences in magnitude and volatility across individual economies. These differences highlight the coexistence of common regional patterns and country-specific characteristics. Overall, the data suggest that despite operating under a shared monetary framework, each country's macroeconomic evolution has followed its own path. Therefore, we analyze the impact of fiscal policy shocks on economic growth separately for each country.

1.4 Data

Building on the approaches of Primiceri (2005) and other related studies (e.g., Rotemberg and Woodford, 1992; Cogley and Sargent, 2001; Cogley and Sargent, 2005; Stock and Watson, 2001), we adopt a compact system of equations to minimize over-parameterization. Our empirical model focuses on three key macroeconomic variables: real government consumption expenditure (in billions of U.S. dollars), real Gross Domestic Product (GDP, also in billions of U.S. dollars), and price level, proxied by the GDP deflator. We specifically use government consumption—rather than total expenditure or public investment, because of data availability and the need for consistency across countries and over time.

To ensure stationarity, the endogenous variables are expressed in first differences rather than logarithmic differences. This specification facilitates the interpretation of impulse responses in monetary units and allows fiscal multipliers to be computed directly from cumulative changes in government expenditure and output. As a robustness check, the model was also estimated using logarithmic differences. The resulting impulse responses exhibit similar dynamic patterns, with differences arising only in scale, indicating that the main findings are robust to the choice of transformation (see Appendix A.6).

The choice of variables reflects both theoretical and practical considerations. GDP captures overall economic activity and is the standard measure of output in fiscal multiplier studies. Price level, included via the GDP deflator, accounts for the price-level effects of fiscal shocks and allows us to examine the interaction between fiscal policy and nominal variables, which is a relevant consideration in low-inflation environments like WAEMU. Government consumption is used as the fiscal policy instrument because it is more consistently and reliably reported across WAEMU countries, and it captures the direct impact of recurrent public spending on the economy, such as wages, purchases of goods and services, and subsidies. This choice aligns with empirical work by Perotti et al. (2007) and Ilzetzki et al. (2013). They argue that government consumption is a more exogenous and analytically manageable component for identifying fiscal shocks in VAR models, in contrast to broader expenditure categories that often include endogenous or cyclical elements.

Due to data limitations, tax variables and other macroeconomic indicators such as monetary policy rates and commodity prices are excluded. The dataset is annual and relatively short, which constrains the number of parameters that can be reliably estimated in the VAR without overfitting. Including additional variables would substantially increase the parameter space, reducing precision and stability. Moreover, consistent pre-1994 monetary data and representative commodity price series are unavailable for all countries. A compact three-variable system (government expenditure, GDP deflator, and real GDP) is therefore used to maintain tractability and ensure credible inference.

All data are sourced from the United Nations Economic Commission for Africa (UNECA), which provides comprehensive macroeconomic statistics for African countries.

Table 1.1: Summary Statistics by Variable across Countries

Variable	**Country**	**Obs**	**Mean**	**Std. Dev.**	**Min**	**Max**
GDP	Benin	49	0.26	0.24	-0.11	0.91
	Burkina Faso	49	0.28	0.25	-0.09	0.88
	Mali	49	0.34	0.38	-0.32	1.17
	Niger	49	0.18	0.32	-0.67	0.90
	Senegal	49	0.36	0.40	-0.36	1.40
	Togo	49	0.11	0.16	-0.36	0.44
Government spending	Benin	49	0.03	0.05	-0.08	0.16
	Burkina Faso	49	0.06	0.08	-0.10	0.32
	Mali	49	0.05	0.08	-0.07	0.31
	Niger	49	0.02	0.09	-0.32	0.28
	Senegal	49	0.05	0.05	-0.07	0.21
	Togo	49	0.02	0.05	-0.08	0.19
Price level	Benin	49	0.02	0.08	-0.23	0.16
	Burkina Faso	49	0.02	0.11	-0.43	0.23
	Mali	49	0.01	0.18	-0.66	0.34
	Niger	49	0.02	0.10	-0.35	0.19
	Senegal	49	0.02	0.10	-0.35	0.22
	Togo	49	0.02	0.09	-0.28	0.17

Source: Author's computations based on [UNECA data (first difference)].

The sample employed spans from 1970 to 2019 for all countries in the study.

Before estimating the model, unit root tests were conducted to assess the statistical properties of the data. Stationarity of the time series variables was examined using the Augmented Dickey-Fuller (ADF) test. The results indicated that all variables in their level form were non-stationary, as the null hypothesis of a unit root could not be rejected at 5% level of significance. However, after taking the first differences of the variables, the ADF test results showed stationarity, confirming that the series were integrated of order one, I(1). Given these findings, the model was estimated using the level-differenced variables to ensure stationarity[1].

Côte d'Ivoire, which accounts for approximately 40% of WAEMU GDP, was excluded from the baseline specification following mixed evidence on the time-series properties of real GDP. The ADF test fails to reject the null hypothesis of a unit root, whereas the KPSS test fails to reject the null hypothesis of stationarity at the conventional 5% significance level. The conflicting outcomes reflect the well-known differences in the hypotheses and finite-sample properties of the two tests, suggesting that the integration properties of the series should be interpreted with caution. Given this ambiguity, Côte d'Ivoire is excluded from the baseline estimation results but is included in the Appendix A.7 for reference. The findings should be interpreted with caution, but they are broadly consistent with the baseline conclusions.

Since the model is estimated using first-differenced data, the impulse response functions (IRFs) reflect how changes in one variable influence changes in another variable over time. Specifically, a shock to government expenditure represents an unexpected increase in the change of government spending — for example, a larger-than-usual rise in expenditure compared to the previous year. The response of GDP then shows how its own change from one year to the next reacts to that shock. These responses capture short-run dynamics in the rate of change rather than permanent effects on

[1]As a robustness check, the model was re-estimated using log-differences rather than the level differences employed in the baseline specification. The impulse responses were qualitatively similar (see Appendix A.6, with the main difference being their magnitude, reflecting the change in the scaling of the variables.

levels. Summary statistics of the variables in the model can be found in Table 1.1

1.5 Empirical Strategy

As discussed in the introduction, this study employs a Time-Varying Parameter Bayesian SVAR (TVP-SVAR). Given the distinct economic structures, policy frameworks, and institutional characteristics of the selected West African countries, we opt for estimating the TVP-VAR model separately for each country rather than employing a Panel TVP-VAR with hierarchical structure. This approach allows us to capture the specific dynamics and responses to fiscal policy shocks in each country without imposing a common structure. It is well known that a hierarchical structure could potentially provide some benefits in terms of information sharing across countries, but the potential loss of country-specific information, the risk of misspecification due to restrictive assumptions and the challenges in the estimation outweigh these benefits in our context.

A key advantage of the Bayesian approach is its ability to address potential overparameterization in VAR models. Unlike frequentist estimation, which relies solely on sample information, the Bayesian framework combines the likelihood with prior distributions for the coefficients and covariance matrices, ensuring well-defined posteriors even in high-dimensional settings or with relatively short samples. This approach helps stabilize estimation, improve inference reliability, and prevent the imprecision that often arises when the model includes many parameters relative to the available observations [Doan et al., 1984; Litterman, 1986; Koop et al., 2010].

1.5.1 Time Varying Parameter Bayesian VAR

The fixed-coefficient assumption may not be valid when dealing with data that exhibit structural breaks, policy changes, or time-varying economic relationships. As a result, the model may perform poorly when relationships between variables change over time. Time-varying parameter VARs (TVP-VARs), also referred to as time-varying parameter VARs (TVP-VARs) differ from standard fixed-coefficient VARs by allowing the coefficients in the linear VAR model to change over time according to a specified law of motion. Furthermore, TVP-VARs are frequently specified with stochastic volatility, allowing the variances of the shocks affecting the VAR system to evolve over time [Lubik and Matthes, 2015]. These models are particularly useful for capturing nonlinearities in macroeconomic time series (see Auerbach and Gorodnichenko, 2012b; Caggiano et al., 2015; Granger and Teräsvirta, 1993).

Model Specification

Consider the model:

$$y_t = \alpha_t + A_{1,t}y_{t-1} + A_{2,t}y_{t-2} + \cdots + A_{p,t}y_{t-p} + \varepsilon_t, \tag{1.1}$$

where

$$y_t = \begin{bmatrix} \Delta GOV_t \\ \Delta DEF_t \\ \Delta GDP_t \end{bmatrix},$$

with ΔGOV_t, ΔDEF_t, and ΔGDP_t denoting the first differences of real government expenditure, the GDP deflator, and real GDP, respectively. The vector α_t contains the time-varying intercepts, $A_{1,t}, \ldots, A_{p,t}$ are 3×3 matrices of time-varying autoregressive coefficients, and ε_t is a 3×1 vector of reduced-form shocks.

The model allows both the coefficients of the VAR system and the covariance matrix of the error terms to evolve over time. The residuals are distributed according to:

$$\varepsilon_t \sim N(0, \Sigma_t). \tag{1.2}$$

Note that the residual covariance matrix is allowed to be period-specific. The model can be expressed in compact form for each period as:

$$y_t = X_t'\beta_t + \varepsilon_t, \tag{1.3}$$

where $X_t' = I_n \otimes (1, y_{t-1}', \ldots, y_{t-p}')$ is a vector of lagged endogenous variables (including intercept) and $\beta_t = (\text{vec}([\alpha_t, A_{1,t}, \ldots, A_{p,t}]'))$ is the vectorized form of coefficients.

The time-varying coefficients follow a random walk process:

$$\beta_t = \beta_{t-1} + \nu_t, \tag{1.4}$$

where: $\nu_t \sim N(0, Q)$ and Q is the covariance matrix of the shocks, assumed to be endogenously determined by the model.

Stochastic Volatility

Our TVP VAR specification follows Dieppe et al. (2016), who introduced stochastic volatility through the dynamic process for Σ_t. They assume that Σ_t can be decomposed as:

$$\Sigma_t = F \Lambda_t F, \tag{1.5}$$

with F denoting a lower-triangular matrix having ones along its diagonal:

$$F = \begin{bmatrix} 1 & 0 & \ldots & 0 \\ f_{21} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ f_{n1} & \ldots & f_{n(n-1)} & 1 \end{bmatrix},$$

and $\Lambda_t = \text{diag}(\bar{s}_1 \exp(\lambda_{1,t}), \ldots, \bar{s}_n \exp(\lambda_{n,t}))$, with known scaling constants $\bar{s}_i$ and dynamic volatility terms $\lambda_{i,t}$ evolving as:

$$\lambda_{i,t} = \gamma \lambda_{i,t-1} + \upsilon_{i,t}, \quad \upsilon_{i,t} \sim \mathcal{N}(0, \phi_i), \tag{1.6}$$

where γ is the autoregressive coefficient and ϕ_i is the variance of the shocks.

The parameters to be estimated include the time-varying VAR coefficients, $\beta = \beta_t : t = 1, \ldots, T$; the covariance matrix of the shocks in the dynamic process, Q; the elements $f^{-1} = fi^{-1} : i = 2, \ldots, n$ associated with the F matrix; the dynamic coefficients $\lambda = \lambda i, t : i = 1, \ldots, n;\ t = 1, \ldots, T$; and the heteroskedasticity parameters $\phi = \phi_i : i = 1, \ldots, n$.

Estimating the Bayesian TVP-VAR

Estimating a time-varying parameter vector autoregressive (TVP-VAR) model involves additional complexities relative to fixed-parameter VARs, due to both time variation in coefficients and stochastic volatility in the innovations. To handle these challenges, we employ Bayesian estimation techniques that efficiently incorporate prior beliefs and update them using observed data. This framework is particularly well-suited to the high-dimensional and nonlinear nature of the TVP-VAR model.

We use the BEAR (Bayesian Estimation, Analysis and Regression) toolbox in `MATLAB` to implement the estimation, following the methodology described in Primiceri, 2005, Cogley and Sargent, 2005, Pereira and Lopes, 2014, and Dieppe et al., 2016. Estimation is conducted via a blocked Gibbs sampling algorithm, iteratively sampling each parameter from its full conditional distribution. The key steps of the algorithm are as follows:

1. Draw time-varying coefficients $\beta_{1:T}$ using a Forward Filtering Backward Sampling (FFBS) algorithm. The forward step applies the Kalman filter to compute the distribution of β_t given the data up to time t, and the backward step samples the entire trajectory $\beta_{1:T}$ recursively.

2. Draw time-varying covariance matrices $\Sigma_{1:T}$: identify F and Λ_t such that $\Sigma_t = F\Lambda_t F'$, with F a lower triangular matrix that has ones on the main diagonal, and Λ_t a diagonal matrix.

3. Draw the innovation variance matrix Q for the evolution of β_t, using the ob-

served innovations $\Delta\beta_t = \beta_t - \beta_{t-1}$. Each term ω_i in Q is sampled from an inverse-Gamma distribution.

Following the BEAR specification, the covariance matrix Q governing the evolution of the state coefficients and the covariance matrix of the volatility innovations are assumed to be diagonal, as in Chan and Jeliazkov (2009) and Murarașu et al. (2023). This simplification, which departs from Primiceri (2005), is adopted for parsimony and identification, ensuring stable posterior sampling with limited annual data. Although this assumption implies independent evolution of coefficients and precludes the possibility of correlated structural shifts, it is employed in applied Bayesian VAR studies for reasons of numerical stability and tractability. Future research could relax this restriction to allow for correlated coefficient dynamics.

The lag length for the time-varying Bayesian VAR models is set to one, as determined by the deviance information criterion (DIC) for the fixed-coefficient Bayesian VAR model (see Appendix A.4). By choosing a small lag length for the time-varying VAR model, we mitigate the risk of overparameterization. Simulations are conducted using 10,000 iterations of the Gibbs sampler, with the first 2,000 iterations discarded (the burn-in sample) to ensure convergence. Following Pereira and Lopes (2014), we perform robustness checks to examine whether the results under two lags differ largely. Details of the prior and conditional posterior distributions are provided in Appendix A.6.1.

Identification and Structural Interpretation

The identification strategy follows a recursive, or Cholesky, scheme to extract structural shocks from the time-varying parameter structural VAR (TVP-SVAR) models. This approach imposes the minimum number of restrictions required for exact identification while leaving the likelihood function unchanged. Consistent with Primiceri (2005), the reduced-form VAR is first estimated, and the structural shocks are subsequently recovered from the estimated variance–covariance matrices.

The recursive ordering used in this study places government spending first, followed by the GDP deflator and real GDP. This ordering assumes that government final consumption expenditure can contemporaneously affect both prices and output but does not respond to either variable within the same period. The GDP deflator is allowed to respond contemporaneously to fiscal shocks but is predetermined with respect to output, while real GDP responds contemporaneously to both fiscal and price shocks. This assumption is supported by the nature of government final consumption expenditure, which largely reflects expenditures on compensation of employees and the purchase of goods and services required for the provision of public services. Such expenditures are typically determined through the budget process and not readily adjusted within a fiscal year in response to contemporaneous economic shocks. The recursive identification strategy is widely used in the fiscal VAR literature (Blanchard and Perotti, 2002; Fatás and Mihov, 2001; Perotti, 2005; Woldu and Szakálné Kanó, 2023) and remains the most common and computationally straightforward identification approach in Bayesian TVP-VAR models. Nevertheless, its application to annual data requires some caution. Unlike quarterly data, annual observations provide greater scope for governments to observe and respond to economic developments within the same year, making the assumption that government spending does not react contemporaneously to output or price shocks more restrictive. The results should therefore be interpreted in light of this identifying assumption.

1.6 Results

1.6.1 Diagnostic Checks

In the Bayesian estimation framework used in this study, the stationarity of the posterior distributions and the convergence of the Markov chains are crucial for reliable inference. We employ four standard diagnostic tests: trace plots, Geweke's diagnostic, the autocorrelation function (ACF), and the effective sample size (ESS). The trace plots of the posterior draws for estimated parameters in the GDP equation exhibit a stable and well-mixed pattern without visible trends or drifts. This suggests that the Markov chains have reached a stationary distribution, providing no evidence of non-convergence. Geweke's test compares the mean of the first and last part of a Markov chain. For all parameters checked, most of the computed Geweke Z-scores fall within the conventional acceptance range ($|Z| < 1.96$). This indicates that the two means are equal, thereby supporting convergence. The Effective Sample Size (ESS) measures how many independent draws the MCMC chain effectively represents, ensuring robust inference. With 10000 MCMC draws, all ESS values are greater than 7000, indicating good mixing and low autocorrelation. The Autocorrelation Function (ACF) assesses the correlation of successive MCMC draws for each parameter. The ACF plots indicate that autocorrelations are close to zero at all lags, suggesting that the draws are approximately independent and that the chains exhibit good mixing and convergence. We therefore conclude that all four diagnostic tests provide evidence of convergence. Outputs from all four diagnostic tests can be found in the Appendix A.5.

1.6.2 Time-varying responses

This study focuses on the impact of government expenditure, with the main analysis presenting the impulse responses of GDP to a change in government spending. We use the TVP-SVAR model to estimate annual impulse responses, capturing the effects of fiscal policy shocks on the economy over the next 15 periods. The results are presented as the median response, which is preferred over the mean due to the minimal influence of extreme responses that may arise from unstable draws. To examine time variation in responses, we display the impulse responses to shocks for the mid-sample year and selected years.

Both government expenditure and GDP are measured in billions of U.S. dollars and are expressed in first differences. Consequently, a 1-unit shock to government expenditure corresponds to a one-billion-dollar increase in real government expenditure relative to the previous year, and the resulting responses indicate the year-on-year change in real GDP, also measured in billions of U.S. dollars. The responses of the GDP deflator represent the year-on-year change in the price-level index resulting from that one-billion-dollar increase in government spending.

Response of GDP to a shock in Government Expenditure

Figure 1.2 shows the impulse responses of GDP to a one-unit shock in government expenditure in the mid-sample year (1995). The responses are positive across all six countries and persist for several periods. The immediate effect ranges from 0.001 for Benin to 0.1 for Senegal, indicating modest but positive short-run impacts. Since the variables are in first differences, the IRFs capture temporary changes in GDP growth rather than permanent effects on output levels. Except for Benin and Togo, the responses are statistically significant at the 68% credibility level.[2]

Overall, the results suggest that fiscal spending accelerations yield short-lived and heterogeneous growth effects, consistent with findings for developing economies (Huidrom et al., 2018; Born et al., 2013). The modest impact responses reflect structural and institutional constraints that often limit fiscal transmission in low-income countries, unlike the stronger and more persistent effects observed in advanced economies (Auerbach and Gorodnichenko, 2012b).

Figure 1.3 further shows cumulative responses, which highlight cross-country differences in fiscal effectiveness. In Burkina Faso and Mali, the cumulative effects build gradually, suggesting more persistent responses to fiscal shocks. By contrast, Benin, Niger, Senegal, and Togo display flatter cumulative profiles, pointing to short-lived impacts consistent with the "fiscal fade-out" pattern commonly observed in aid-dependent economies (Ilzetzki et al., 2013; Honda et al., 2020).

Figure 1.4 compares the impulse responses for 1972, 1995, and 2019. While the direction and magnitude of short-run effects are broadly stable—especially in Benin and Senegal—the cumulative responses (Figure 1.5) show moderate time variation. However, once 68% credibility interval are considered (Figure 1.6), the differences across years are not statistically significant, implying that the overall fiscal transmission mechanism remains stable over time.

[2] In Bayesian VAR analysis, 68% interval is typically reported to convey uncertainty without cluttering figures.

Across countries, Niger, Senegal, and Togo exhibit faster adjustments, with GDP stabilizing within six periods of a shock—likely reflecting more flexible economic structures. In contrast, Benin, Burkina Faso, and Mali take longer, often exceeding fifteen periods to return to steady state, suggesting more persistent fiscal effects due to slower adjustment or structural rigidities. Reliance on volatile exports such as cotton in Benin and Togo may further prolong recovery following fiscal shocks.

Response of the Price Level to a shock in Government Expenditure

Figure 1.7 presents the estimated impulse response functions of the change in the price level—as measured by the GDP deflator—to a shock in the change of government expenditure for the mid-sample year (1995). The responses vary considerably across countries: Benin, Mali, and Senegal exhibit positive initial responses, whereas Burkina Faso, Niger, and Togo display negative impact responses. All responses are presented with 68% Bayesian credibility interval.

A positive response of the price level to government spending shocks is broadly consistent with standard demand-side mechanisms, whereby fiscal expansions increase aggregate demand and place upward pressure on prices. In the context of government consumption—particularly public wages or recurrent transfers—this effect can manifest quickly, especially in economies with limited spare capacity or inelastic supply in key sectors. By contrast, negative price level responses, as observed in Burkina Faso, Niger, and Togo, may reflect fiscal stimulus occurring during periods of weak demand, where excess capacity absorbs the demand increase without generating higher price levels.

The impulse responses are non-monotonic in several countries: initial increases in the price level are followed by reductions, or vice versa. These oscillating patterns could reflect the short-lived nature of fiscal demand boosts. Since government consumption typically does not contribute to lasting productivity gains, any inflationary effects are likely to fade quickly. Indeed, across all countries, price level responses exhibit a rapid return to baseline, consistent with the WAEMU region's institutional architecture.

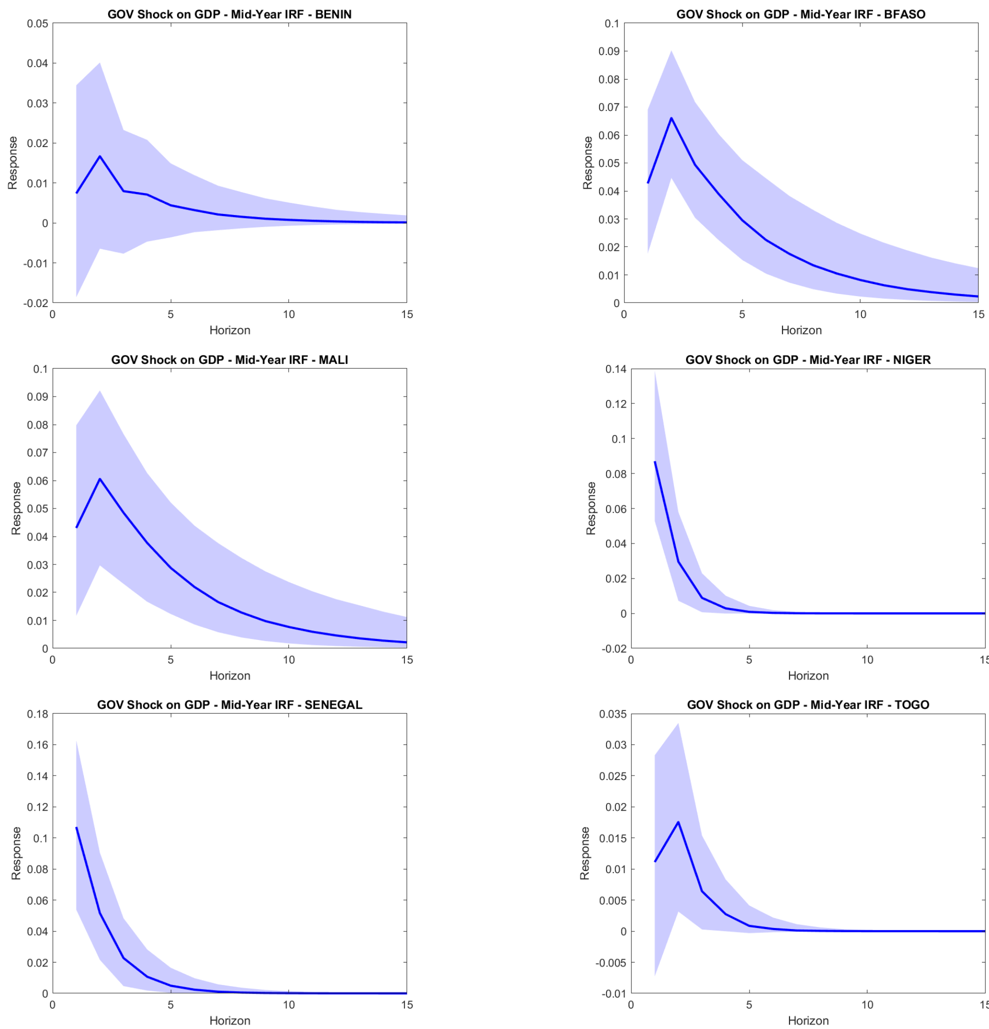


Figure 1.2: Mid-sample year impulse responses of GDP to a 1-unit shock in Government Expenditure with 68% credibility interval

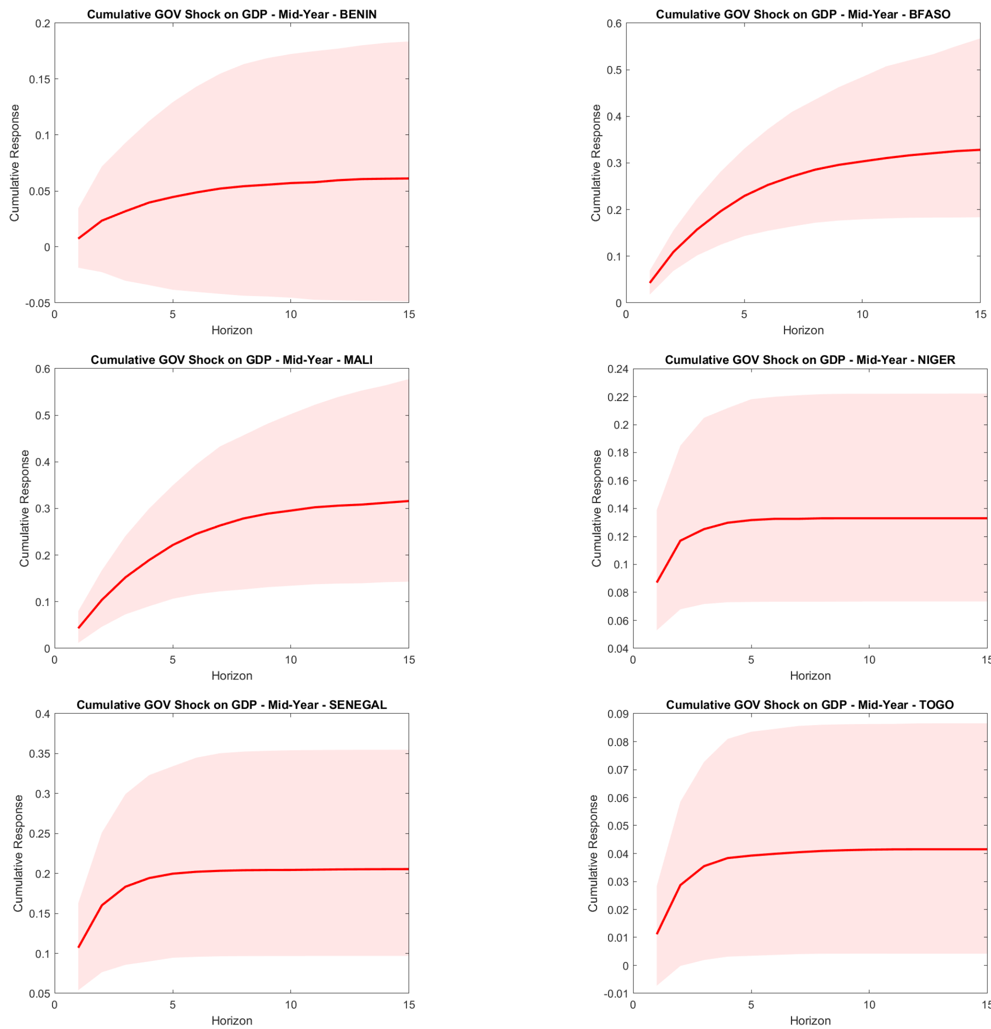


Figure 1.3: Cumulative impulse responses of GDP to a 1-unit shock in Government Expenditure, with 68% credibility interval

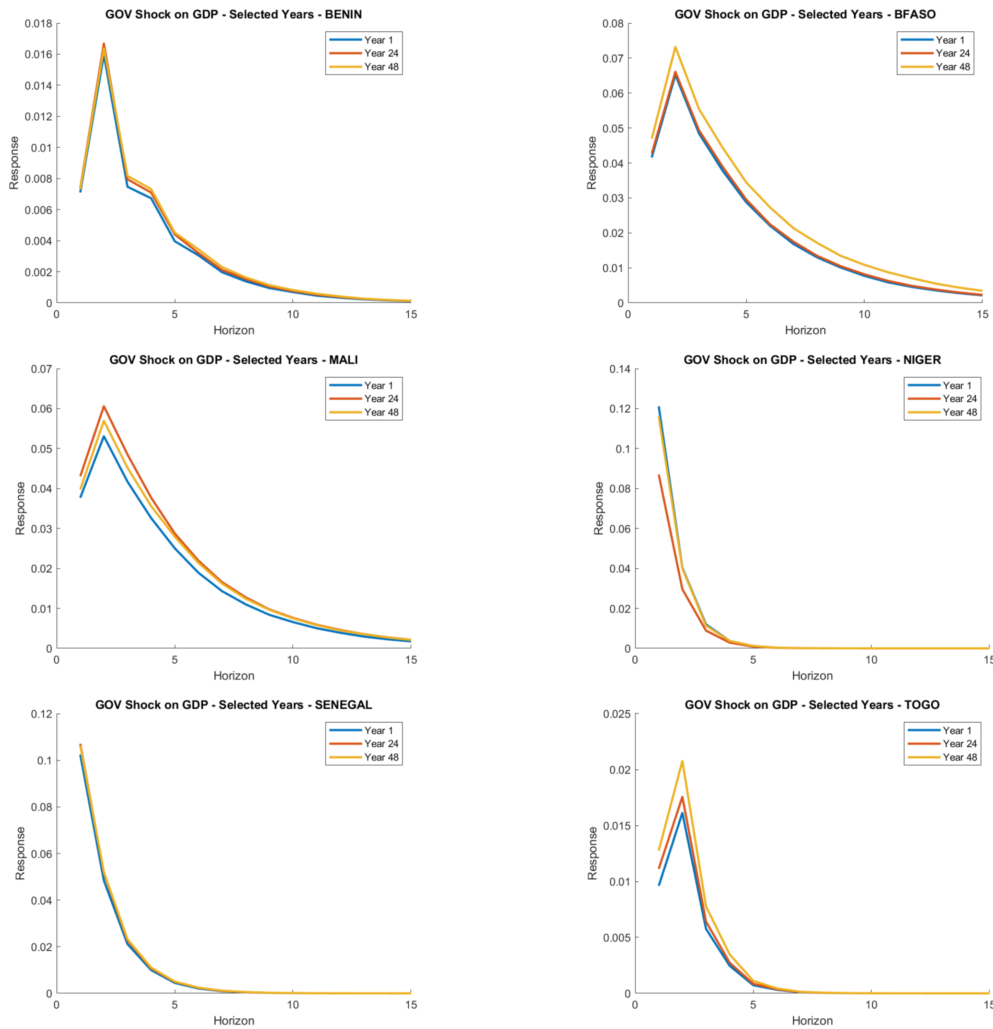


Figure 1.4: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years

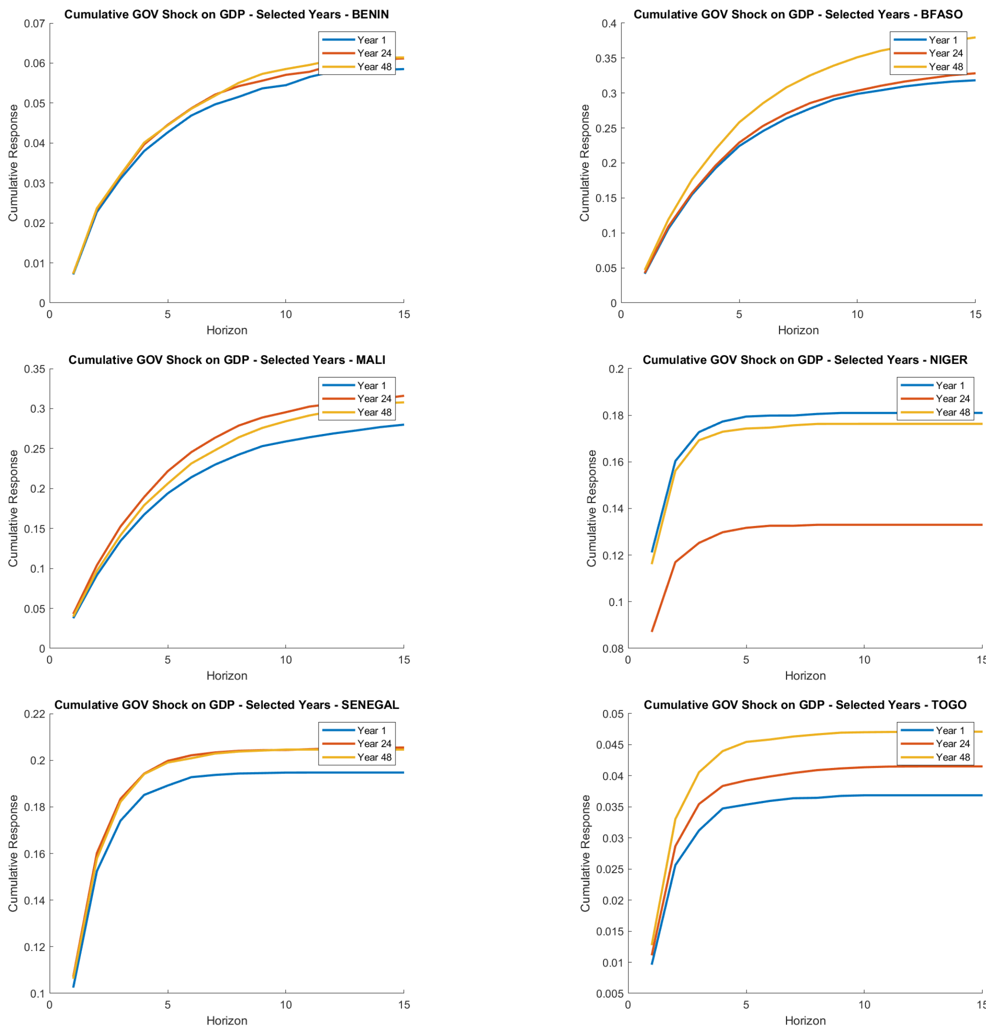


Figure 1.5: Cumulative impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years

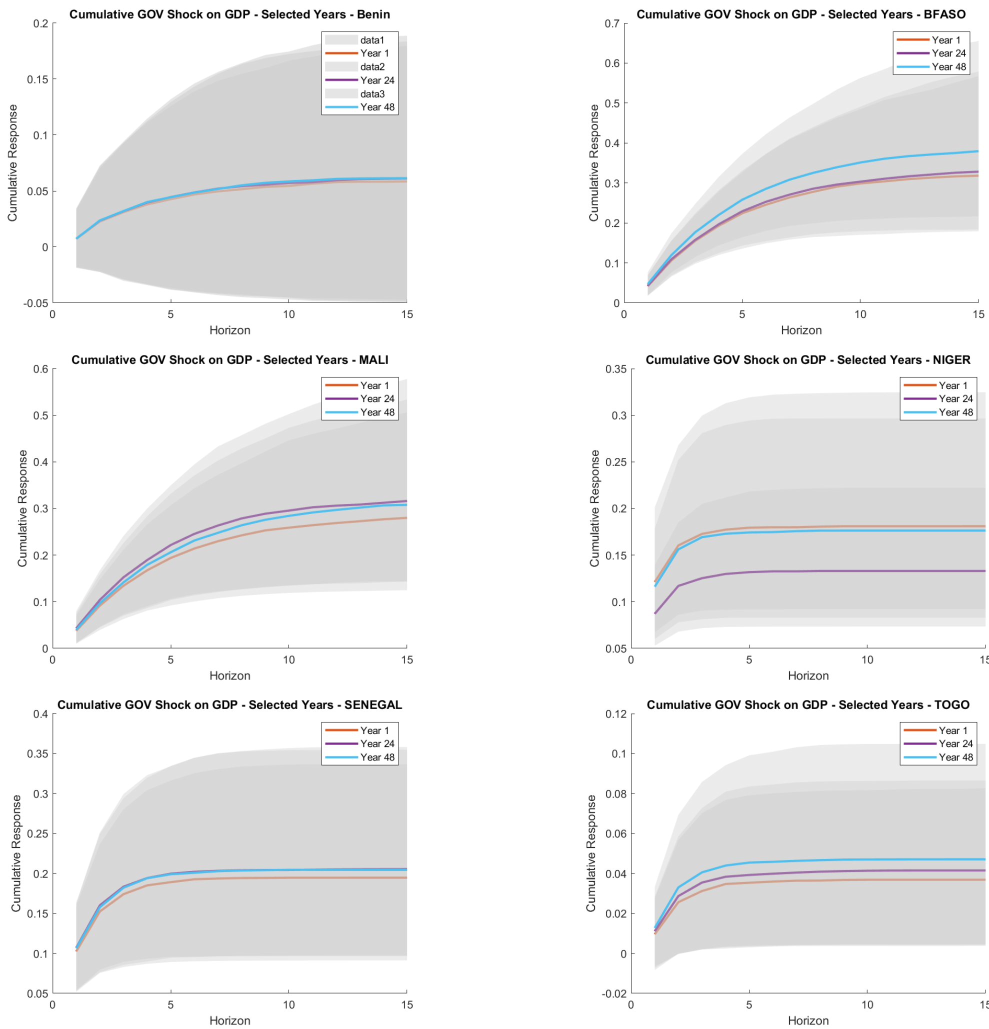


Figure 1.6: Cumulative impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years with 68% credibility interval

The presence of a common monetary policy under the BCEAO, alongside a fixed exchange rate regime anchored to the euro, provides a strong nominal anchor that curtails sustained price increases. This macroeconomic arrangement likely explains why inflationary effects from fiscal shocks appear transitory and contained, despite variations in short-run impact.

The results suggest that government consumption shocks in WAEMU countries can temporarily affect price levels, but the effects dissipate rapidly—reflecting both the type of expenditure and the constraints imposed by the monetary union. The differences across countries emphasize the role of national fiscal structures, expenditure composition, and domestic market conditions in influencing short-run inflation dynamics in a monetary union structure.

Overall, the TVP-SVAR results indicate that fiscal policy impacts in these West African countries exhibit some degree of time-varying characteristics in the first fifteen years following a fiscal shock. For most countries, the period-by-period impulse response of GDP changes to an exogenous government spending shock exhibits little time variation, whereas the cumulative responses display more noticeable changes over time.

1.6.3 Time-varying Fiscal Multipliers

This section analyzes the short-run and long-run fiscal multipliers (SRFM and LRFM) over time, based on impulse response functions derived from a time-varying parameter structural VAR model using annual data in first differences. The fiscal multipliers quantify the cumulative output response relative to the cumulative government consumption response following a structural shock to the change in government spending.

The short-run fiscal multiplier at time t is defined as:

$$SRFM_t = \frac{\sum_{j=0}^{3} IRF_{\Delta GDP}(t,j)}{\sum_{j=0}^{3} IRF_{\Delta GOV}(t,j)}$$

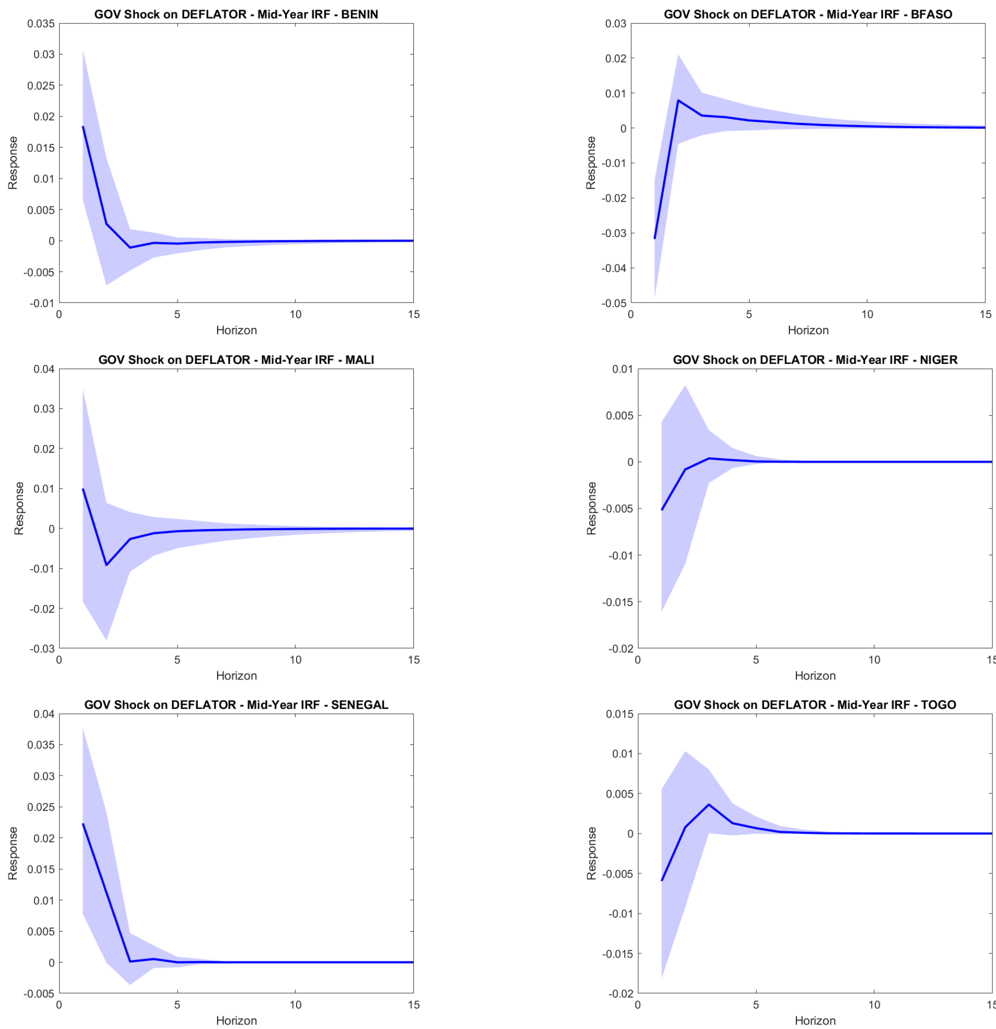


Figure 1.7: Mid-sample year impulse responses of Price Level to a 1-unit shock in Government Expenditure with 68% credibility interval

where:

- $\sum_{j=0}^{3} IRF_{\Delta GDP}(t, j)$ is the cumulative change in GDP over the first 3 years in response to a shock to the change in government spending.
- $\sum_{j=0}^{3} IRF_{\Delta GOV}(t, j)$ is the cumulative change in government spending over the first 3 years.

Similarly, the long-run fiscal multiplier (LRFM) is calculated as:

$$LRFM_t = \frac{\sum_{j=0}^{15} IRF_{\Delta GDP}(t, j)}{\sum_{j=0}^{15} IRF_{\Delta GOV}(t, j)}$$

Figures 1.8 and 1.9 display the short-run and long-run fiscal multipliers, respectively, with 68% Bayesian credibility interval.

The estimated short-run fiscal multipliers (SRFMs) and long-run fiscal multipliers (LRFMs) vary across countries, ranging from about 0.8–4.4 in the short run and 0.9–6.3 in the long run. Senegal consistently records the highest multipliers, while Togo shows the lowest. This pattern reflects differences in economic size and capacity: larger and more diversified economies, such as Senegal, tend to amplify fiscal shocks more effectively than smaller economies like Togo, where limited fiscal space and weaker transmission channels constrain impact.

The unusually high multipliers across most WAEMU countries partly reflect the macroeconomic environment following the 1994 CFA franc devaluation, when large inflows of external aid and debt relief—equivalent to roughly 30% of the zone's GDP—supported public investment and social spending. These consumption-oriented programs, combined with underutilized labor and infrastructure, likely produced stronger short-run effects typical of developing economies.

High multipliers, particularly in Senegal, also stem from the composition of expenditure. Because the measure of government spending used here reflects current consumption, increases translate directly into higher public wages and government

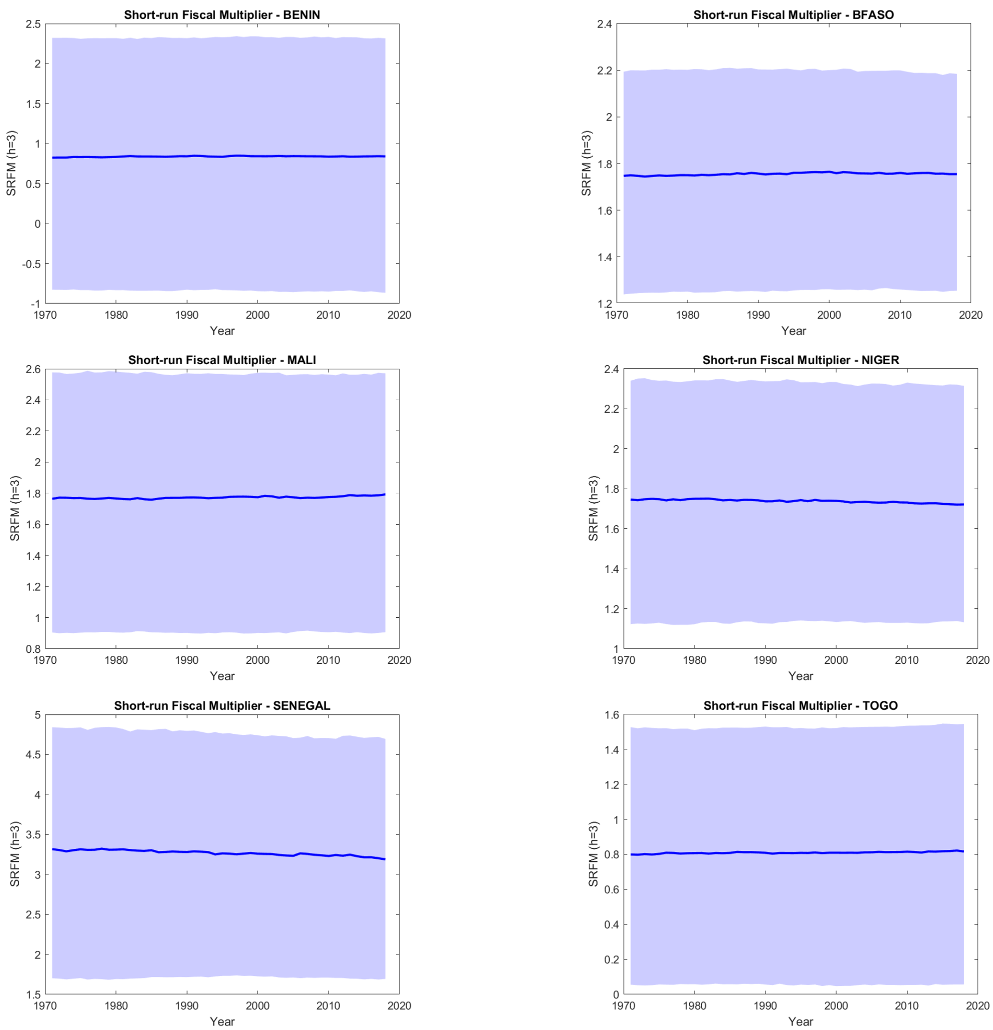


Figure 1.8: Short run fiscal multiplier, with 68% credibility interval

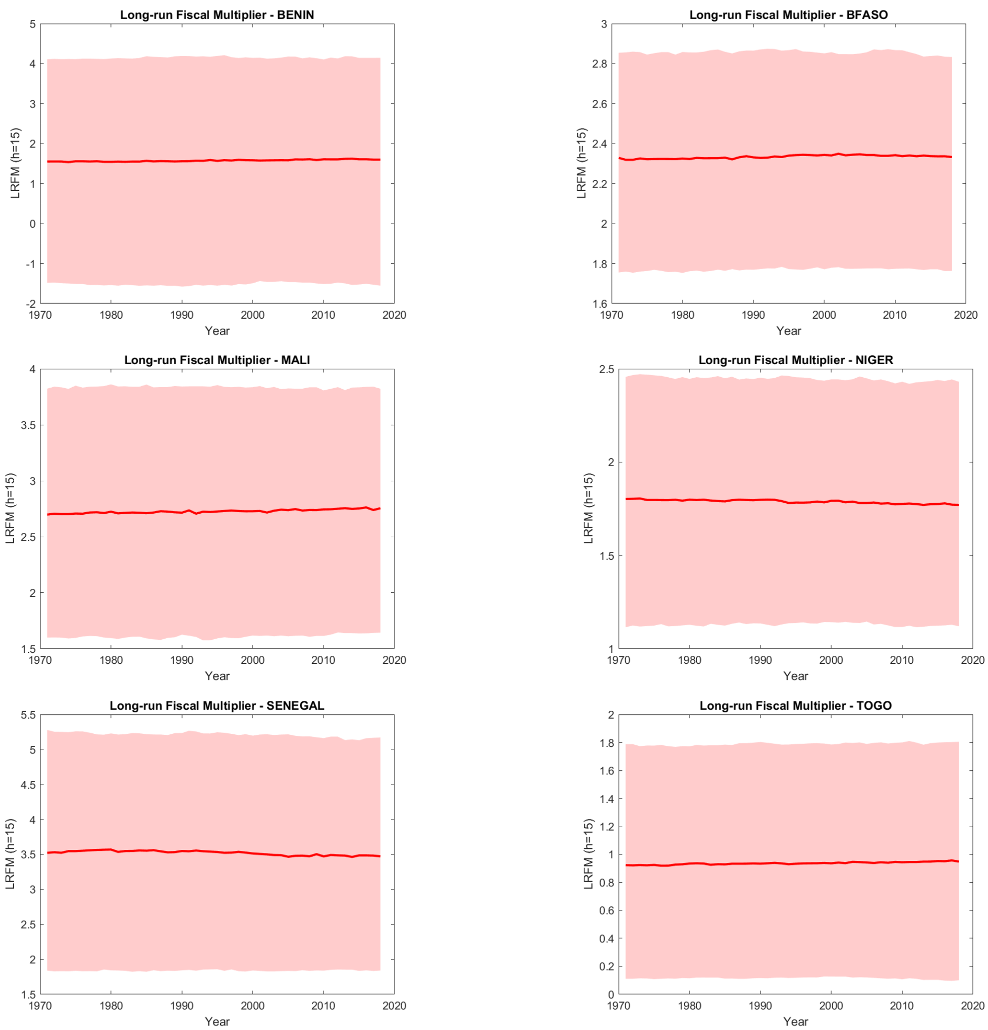


Figure 1.9: Long run fiscal multiplier, with 68% credibility interval

purchases, with immediate spillovers to household income and private demand. In economies with high propensities to consume and weak automatic stabilizers, such shocks can produce rapid and sizable output responses. In Senegal's case, part of the large multiplier may also arise mechanically from a small cumulative government spending response relative to GDP, inflating the ratio by construction.

Multipliers above unity for all countries except Benin align with New Keynesian models that allow for highly expansionary fiscal effects under nominal rigidities, liquidity constraints, and limited monetary offset—conditions characteristic of monetary unions like the WAEMU. When monetary policy is constrained and many households consume out of current income, fiscal expansions can raise demand more than proportionately.

Finally, both short-run and long-run multipliers exhibit limited time variation. Although GDP and expenditure responses shift over time, they tend to move proportionately, leaving the ratio between them relatively stable. This implies that while fiscal shocks evolve in intensity, the overall responsiveness of output to government spending remains broadly consistent over time.

The relatively large magnitude of the estimated fiscal multipliers should, however, be interpreted with caution, as several measurement and specification features may contribute to their size. First, since the multiplier is computed as a ratio of cumulative responses, a relatively small or volatile response of government spending—the denominator—can mechanically inflate the multiplier. Second, the variables are measured in levels expressed in billions of U.S. dollars, and cross-country differences in economic size and scaling may affect the magnitude and comparability of the estimated responses. Third, because the model is estimated in first differences, the impulse responses capture changes in growth rates rather than level effects, which may amplify cumulative responses relative to standard level-based multiplier definitions. Taken together, these considerations imply that the estimated multipliers should be interpreted primarily as indicators of relative fiscal effectiveness across countries and over time, rather than as precise quantitative measures of output gains.

Table 1.2: Fiscal Multipliers and Persistence Ratios

Country	**SRFM**	**LRFM**	**PR**
Benin	0.84	1.59	1.90
Burkina Faso	1.75	2.33	1.33
Mali	1.77	2.72	1.54
Niger	1.74	1.78	1.02
Senegal	3.25	3.54	1.09
Togo	0.81	0.93	1.15

Source: Author's computations

To assess the persistence of fiscal policy effects, we compute a persistence ratio (PR_t) defined as the ratio of the long-run fiscal multiplier to the short-run fiscal multiplier.

$$PR_t = \frac{LRFM_t}{SRFM_t}$$

Figure 1.10 presents the persistence ratios with 68% credibility interval. This measure indicates how the effects of Government Expenditure shocks on GDP evolve over time: values above one imply that a large share of the total impact occurs beyond the short run, reflecting a more persistent response.

All WAEMU countries exhibit persistence ratios greater than one, with particularly high values for Benin, Burkina Faso, and Mali. This suggests that fiscal shocks in these economies build gradually, consistent with delayed transmission in developing countries where weak infrastructure, implementation lags, limited financial depth, and supply-side rigidities slow the propagation of fiscal policy effects. The ratios are also relatively stable across time, indicating that structural and institutional features influencing fiscal transmission remain persistent over the sample period.

The 68% credibility interval around both the short-run and long-run fiscal multipliers (Figures 1.8 and 1.9) are relatively wide because each multiplier is a ratio of two estimated quantities, each with its own posterior uncertainty. When the cumulative response of government expenditure is small or volatile, even modest uncertainty in GDP responses can lead to large variation in the multiplier. A similar compounding effect applies to the persistence ratio (Figure 1.10), which is a ratio of two multipliers. This nested structure amplifies uncertainty in both components, resulting in wider interval and requiring cautious interpretation of the central estimates.

1.7 Robustness Check/Sensitivity Analysis

To ensure the robustness of the results in this study, we undertake some sensitivity analyses and present the details in this section. First, we conduct sensitivity analyses by modifying the autoregressive (AR) coefficient on the residual variance from the default value of 0.85 to 0.7. This hyperparameter governs the assumed persistence in the volatility of shocks across time. A higher value, closer to 1.0, implies that volatility is highly persistent, with shocks to variance having a long-lasting effect (near-unit-root process) in volatility. Lowering the AR coefficient to 0.7 assumes less persistence, allowing volatility to revert more quickly to its mean and enabling the model to adapt more flexibly to evolving shock dynamics. Adjusting this parameter can influence the evolution of uncertainty over time and, consequently, the magnitude and smoothness of impulse responses.

We find that adjusting the AR coefficient to 0.7 does not change the sign or magnitude of the impulse responses for most countries. For instance, with Burkina Faso's data, the impact response of GDP to a positive shock in government expenditure remains within the range of 0.04 to 0.05. However, for Benin, using an AR coefficient of 0.7 produces impulse responses with more visible time variation. This may reflect that the data for this country already signal underlying shifts in volatility. Weakening the prior on smoothness in volatility by lowering the AR coefficient gives the stochastic volatility component more flexibility, allowing it to capture these changes more prominently, resulting in greater variation in the estimated IRFs. When γ is set to 1, reflecting a random walk, the results for all countries are similar to those

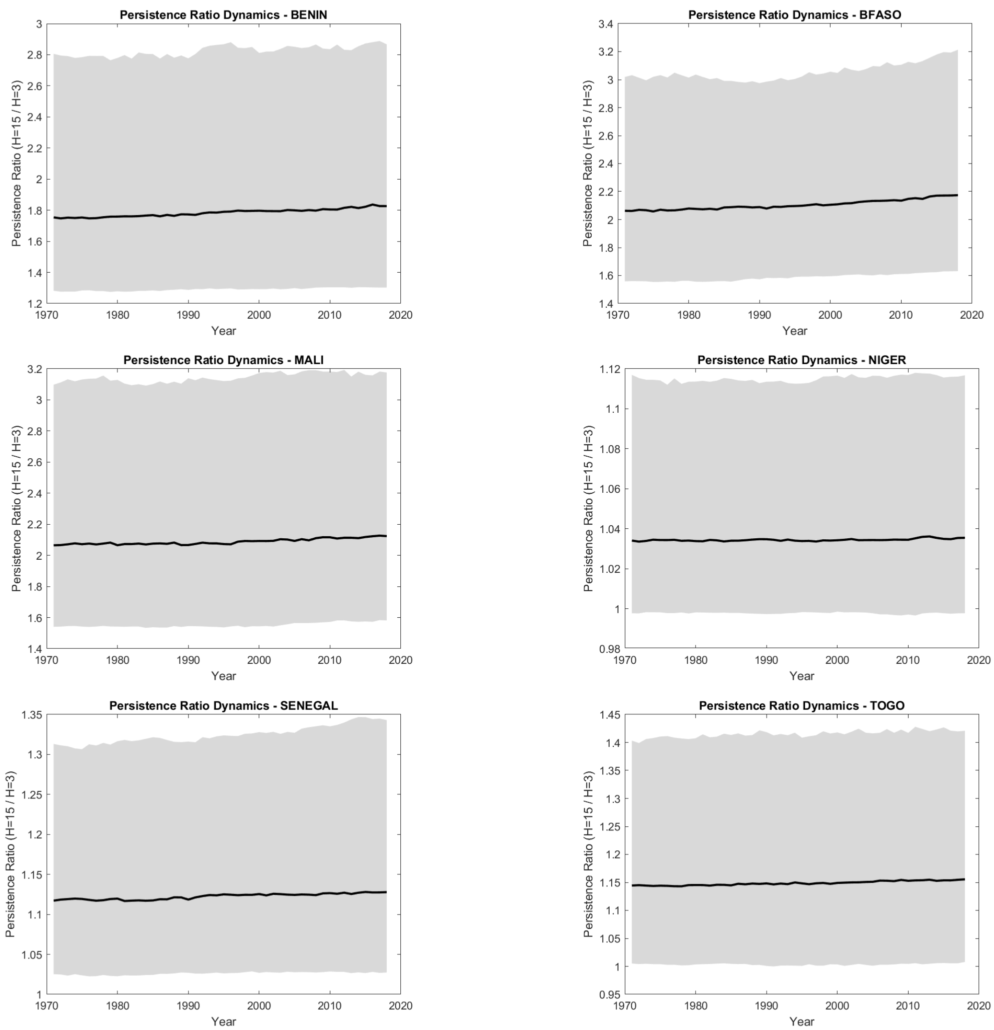


Figure 1.10: Persistence Ratio, with 68% credibility interval

obtained in the main specification.

In their implementation of the TVP-SVAR model, Dieppe et al. (2016) assume an inverse gamma prior for each heteroskedasticity variance parameter, ϕ_i to be inverse gamma with shape $\frac{\alpha_0}{2}$ and scale $\frac{\delta_0}{2}$. They implement a loose prior which is weakly informative, by using low hyperparameter values: $\alpha_0 = \delta_0 = 0.001$. This choice prevents biasing the results toward small or specific variance values, ensuring that the posterior estimates are primarily driven by the data. However, a potential drawback of this approach is its sensitivity to data quality: with noisy data or small sample sizes, the posterior estimates may become unstable due to insufficient regularization. While this loose prior works well with large datasets, we considered whether tightening the prior would enhance stability, particularly since our dataset is relatively small compared to others in the literature. Consequently, we adjusted the hyperparameter values to $\alpha_0 = \delta_0 = 0.01$. We find that this change has little effect on most countries, with their impulse responses remaining largely stable. However, for Benin, the GDP response begins to exhibit some time variation.

Additionally, following Pereira and Lopes (2014), we adjusted the lag length in the TVP-SVAR specification from 1 to 2. We adopted a lag length of 1 in the original specification following the DIC recommendation as well as to mitigate the risk of overparameterization and manage the computational demands of the simulation procedures. Increasing the lag length to 2 had minimal impact on the results: the sign and direction of the responses remained consistent with only slight changes in time-variation and their magnitudes. Also, variables generally took a longer time to reach their steady-state levels.

Finally, the ordering of the variables was reversed so that government spending was placed last rather than first in the Cholesky decomposition. This alternative specification preserves the qualitative features of the impulse responses: the signs and directions of the effects remain the same, indicating that the main conclusions are not sensitive to the baseline ordering assumption. Only the magnitudes of the responses exhibit some variation, which is expected when altering the recursive structure. The robustness checks confirm that our key results are stable across various model specifications.

1.8 Conclusion

This study provides new evidence on fiscal policy dynamics in the WAEMU by examining time-varying fiscal multipliers within a Bayesian Time-Varying Parameter Structural VAR (TVP-SVAR) framework. The analysis addresses how the effects of government spending in low-income countries operating under a fixed exchange rate and centralized monetary regime have evolved over time.

Unlike most studies focused on advanced economies, WAEMU countries face rigid macroeconomic constraints, making fiscal policy their primary stabilization tool. The results show that government spending shocks generally boosted GDP, although the magnitude and persistence of these effects differed across countries. While the estimation framework allows fiscal multipliers to change over time, the results indicate that they remained broadly stable over the sample period. Instead, the main differences were observed across countries, suggesting that fiscal multipliers are not uniform even within a monetary union.

Methodologically, incorporating stochastic volatility within a TVP-SVAR framework allows the model to capture gradual shifts in fiscal transmission that standard fixed-parameter approaches cannot detect. This is crucial in a region shaped by policy reforms, commodity price swings, and shocks such as the 1994 devaluation. The findings also indicate that countries with stronger fiscal management, such as Senegal, exhibited higher and more stable multipliers, while those with weaker institutions showed lower and more volatile effects. Inflationary responses were short-lived and varied mainly due to fiscal, not monetary, differences. While the estimated fiscal multipliers are relatively large in some cases, they should be interpreted with caution. Their magnitude partly reflects the ratio-based construction of the multipliers, potential scaling differences across countries, and the use of first-differenced data, which captures changes in growth rates rather than level effects.

From a policy perspective, these findings suggest that uniform fiscal policy prescriptions within the WAEMU may lead to suboptimal outcomes. Instead, fiscal interventions should account for country-specific structural conditions, including institutional capacity, economic diversification, and exposure to external shocks

Robustness checks confirm the stability of the results, though the inability to separate anticipated from unanticipated shocks remains a limitation. Future work could distinguish between current and capital spending, consider financing sources, and explore fiscal–monetary interactions using higher-frequency data.

Chapter 2

Oil Revenue Shocks, Fiscal Policy and Macroeconomic Dynamics in African Oil Exporters: Evidence from a Bayesian Panel VAR

Abstract

This paper examines the impact of oil revenue shocks on fiscal and macroeconomic outcomes in four African oil-exporting economies—Algeria, Angola, Egypt, and Nigeria. Unlike studies that focus solely on oil price shocks, this analysis employs oil revenue to capture the combined effects of price and production, offering a more comprehensive measure of fiscal exposure to oil market fluctuations. Using a hierarchical Bayesian panel VAR alongside country-specific structural VAR models, the results reveal that oil revenue shocks generate a positive and statistically significant short-run responses in government expenditure, GDP and inflation, with effects fading after three to four years. The magnitude of these responses varies across countries, with Algeria and Egypt exhibiting moderate effects compared to the larger responses of Angola and Nigeria. The appreciation of the real effective exchange rate following oil revenue windfalls is also transitory. Robustness checks using alternative oil revenue definitions, exchange rate measures, and sample compositions confirm the stability of the main results. The findings highlight the temporary

nature of oil windfalls and the importance of fiscal frameworks that smooth revenue volatility and enhance macroeconomic resilience.

2.1 Introduction

The economic performance of oil-exporting economies remains closely tied to the volatility of global energy markets. In these countries, oil is not merely an export commodity but the primary driver of government revenue, foreign currency earnings, and overall macroeconomic balance. Consequently, these countries have frequently experienced cycles of rapid growth followed by downturns, driven by fiscal policies that amplify these trends, unpredictable government spending, and symptoms consistent with Dutch disease. Therefore, understanding how shocks originating in the oil sector transmit to domestic macroeconomic conditions is crucial for both policymakers and researchers—particularly in Africa, where dependence on oil revenues remains high and fiscal buffers are often limited.

A large macroeconomic literature has examined how oil market fluctuations affect output, inflation, and external balances. Seminal contributions by Hamilton (1983) and Kilian (2009) show that oil price movements can have significant macroeconomic effects, depending on whether shocks originate from supply, global demand, or oil-specific factors. Subsequent studies emphasize that the transmission of oil shocks depends on structural characteristics such as monetary policy regimes, labor market flexibility, and trade exposure to oil (Bernanke et al., 1997; Blanchard and Gali, 2007; Nasir et al., 2018). More recent work, including Baumeister and Hamilton (2019) and Caldara et al. (2019), refines this framework by jointly modeling oil prices and production in order to identify distinct oil market shocks.

While this literature provides important insights into the macroeconomic effects of oil market disturbances, it is primarily concerned with identifying the sources of oil price and production fluctuations rather than with the fiscal transmission mechanisms that are central to oil-exporting economies. In particular, although oil production is often included as a variable, the focus is typically on oil market dynamics rather than on total oil revenues and their direct implications for government finances. For oil-dependent countries, fluctuations in oil revenues—determined jointly by prices and production—constitute the primary channel through which external oil shocks affect fiscal outcomes and broader macroeconomic performance. This motivates our focus on oil revenue shocks, defined as unanticipated changes in total income or government receipts derived from the oil sector. By directly modeling oil revenues, we adopt a measure that is more closely aligned with the fiscal and external constraints faced by oil-exporting economies and therefore more informative for policy analysis (Cunado and De Gracia, 2005; Dreger and Rahmani, 2016; Adedokun, 2018).

Despite some progress in understanding oil market shocks, important gaps remain. Much of the existing research is confined to case studies of individual countries, often employs recursive identification strategies, and fails to account for cross-country variation in the transmission of fiscal responses. Moreover, little attention has been paid to how institutional factors, exchange rate systems, or fiscal governance frameworks shape responses to oil revenue fluctuations. These issues are particularly relevant for African oil-exporting nations, where fiscal policy tends to be procyclical and reliant on volatile external income. Addressing these gaps is crucial, at a time when oil revenue volatility has intensified following the COVID-19 pandemic, geopolitical tensions, and the global energy transition. A clearer understanding of how oil revenue shocks influence government budgets, economic growth, inflation dynamics, and currency valuations provides a basis for formulating effective, countercyclical fiscal strategies.

Our study contributes to the literature in three main ways. First, we adopt a sign-restriction identification strategy that aligns with theoretical predictions, allowing us to distinguish oil revenue shocks from aggregate demand and supply disturbances (Uhlig, 2005; Rubio-Ramirez et al., 2010; Baumeister and Hamilton, 2019). Second, we integrate country-specific Structural VARs (SVARs) with a Bayesian Hierarchical Panel VAR (HPVAR) approach. This allows us to capture both the common patterns across countries and the heterogeneity in their individual responses (Canova and Ciccarelli, 2013; Jarociński, 2010). By combining these two frameworks, we achieve efficient estimation in short samples while gaining richer comparative insights into how fiscal and macroeconomic dynamics evolve across Algeria, Angola, Egypt, and Nigeria. Third, the analysis highlights the central role of fiscal policy as the main transmission channel of oil revenue shocks for countries in the study, showing that oil windfalls operate primarily through government expenditure dynamics rather than purely through external or supply-side mechanisms.

Our findings indicate that the positive effects of oil revenue shocks on government expenditures, output, and the real exchange rate are largely transitory, typically fading within three to four years. Importantly, the results show that these effects vary significantly across countries, reflecting differences in institutional capacity and fiscal frameworks. By examining how fiscal capacity shocks interact with broader macroeconomic adjustments under varying institutional frameworks, we provide fresh insights that can guide the development of stabilization policies, fiscal governance frameworks, and sovereign wealth strategies in Africa's resource-rich nations.

The remainder of the paper is organized as follows. Section 2 reviews the related literature on oil shocks and their macroeconomic implications. Section 3 outlines the theoretical framework guiding the analysis. Section 4 presents an overview of the economies under study by examining the historical evolution of key variables through graphical analysis. Section 5 describes the empirical strategy, including the data and methodology employed. Section 6 discusses the empirical results, while Section 7 provides robustness checks to assess the stability of the findings. Section 8 concludes the paper.

2.2 Literature Review

2.2.1 Oil Shocks and Macroeconomic responses

Early foundational work by Hamilton (1983) established a strong historical relationship between oil price spikes and subsequent economic downturns, showing that oil price increases preceded seven out of eight U.S. recessions since World War II and arguing that oil price shocks cause substantial declines in output and employment. However, subsequent research has shown that these severe macroeconomic disruptions were largely associated with the structural conditions of the 1970s. Studies such as Hooker (1996), Blanchard and Gali (2007), and Baumeister and Peersman (2013) document that the transmission of oil shocks to the real economy in advanced oil-importing economies has become considerably weaker over time. Blanchard and Gali (2007) attribute this moderation to reduced energy intensity in production, more flexible labor markets that limit wage–price spirals, and more credible monetary policy frameworks.

A related strand of the literature emphasizes that the macroeconomic effects of oil shocks depend critically on their underlying sources and on country-specific structural characteristics. Kilian (2009) shows that supply-driven oil price increases tend to have relatively mild effects, aggregate demand shocks can be expansionary, and oil-specific demand shocks are strongly contractionary. The impact of oil shocks also varies across countries depending on structural and policy characteristics such as the prevailing monetary policy stance, labor market flexibility, external conditions, and whether the country is an oil exporter or importer (Bernanke et al., 1997; Blanchard and Gali, 2007; Nasir et al., 2018; Yang et al., 2023). Further advances refine the identification of global oil shocks by distinguishing between demand-driven and risk-based movements in oil markets (Baumeister and Hamilton, 2019; Caldara et al., 2019), while Aastveit et al. (2015) show that global demand shocks are important in explaining fluctuations in both oil prices and output. These findings suggest that the macroeconomic effects of oil shocks are complex, nonlinear, and highly context-dependent.

While these studies have deepened our understanding of oil price fluctuations, a growing body of work argues that oil price movements alone do not adequately capture the shocks relevant for oil-dependent economies. For these countries, the true disturbance to fiscal and external balances stems not from price changes in isolation, but from the variation in total oil income—the product of price and production volume. This perspective introduces the concept of an oil revenue shock, defined as an unexpected change in total oil income or government receipts from the sector (Cunado and De Gracia, 2005; Adedokun, 2018; Bouyacoub, 2025). Because it reflects both price and quantity effects, an oil revenue shock directly captures the wealth transfer and fiscal impulse experienced by oil-exporting countries.

Empirical evidence strongly supports this view. Adedokun (2018) finds for Nigeria that oil revenue shocks are stronger predictors of government expenditure than oil price shocks, while Bouyacoub (2025) documents long-run relationships between oil prices, oil revenues, and fiscal outcomes in Algeria. Similarly, Dreger and Rahmani (2016) show that oil revenue fluctuations exert persistent effects on growth and fiscal stability across MENA countries. These studies confirm that oil revenue shocks are a more comprehensive and fiscally relevant measure of the macroeconomic impulse in oil-exporting economies. However, they remain largely single-country analyses and typically rely on recursive identification, which may confound the simultaneity between fiscal and output responses.

2.2.2 Definition and Measurement of Oil Revenue Shocks

The literature defines oil revenue shocks in several ways. A common measure is the product of oil prices and production volumes (P×Q), representing total oil income (Ahmad and Mottu, 2002; Bouyacoub, 2025). However, implementing this measure is often constrained by data limitations, especially for developing countries where consistent time series on production and benchmark prices are unavailable for long historical periods. For instance, global benchmarks such as Brent prices only begin in 1987. To ensure broader comparability across countries and over time, we follow multiple studies by using oil rents (% of GDP) as a proxy. Oil rents, calculated as the value of crude oil production at world prices minus production costs, capture both price and quantity dynamics while expressing potential resource income relative to the size of the economy. Although they do not directly represent fiscal revenues, they are widely used in the resource-curse literature (Mehlum et al., 2006; Majumder et al., 2020; Farzanegan and Thum, 2020; Mignamissi and Kuete, 2021; Sweidan and Elbargathi, 2022) and provide a reliable indicator of the resource-driven capacity that shapes government spending and macroeconomic conditions.

In developing economies, oil revenue or price shocks tend to have amplified and more persistent effects due to limited diversification, structural rigidities, and weak fiscal institutions. Collier and Goderis (2009) observe that while commodity booms raise growth in the short run, they tend to lower long-run growth in countries with weak governance and high resource dependence. El Anshasy and Bradley (2012) similarly find that oil-exporting developing countries often exhibit procyclical fiscal behavior, with government spending and GDP moving in tandem with oil price movements, contributing to macroeconomic instability. These findings highlight the importance of institutional quality and fiscal frameworks in mediating the effects of oil shocks.

Oil revenue is measured using oil rents as a percentage of GDP, which captures changes in both international oil prices and production volumes. Three of the four countries in the sample (Algeria, Angola, and Nigeria) were subject to OPEC production agreements during part or all of the sample period, which tends to impose a ceiling on oil production through production quotas. Production volumes are also influenced by other factors such as investment decisions and production capacity, implying that the quantity component of oil revenue is less clearly exogenous than the price component. Although using an international oil price would provide a cleaner source of exogenous variation, it would not capture changes in government revenue which is the focus of this study. The oil rents measure is therefore retained as the most appropriate indicator of government oil revenue, while recognizing that the estimated oil revenue shocks reflect the combined effects of price and quantity movements and should be interpreted accordingly.

2.2.3 Evidence from African Oil-Exporting Economies

In the African context, several studies have analyzed the macroeconomic effects of oil price shocks using time-series or panel methods. Omojolaibi and Egwaikhide (2014) find that oil price volatility affects GDP, fiscal balances, and investment across

African oil exporters. Rotimi and Ngalawa (2017) show that positive oil price shocks raise GDP and appreciate the real exchange rate, often followed by delayed inflationary effects. While these panel approaches provide valuable regional insights, they may obscure cross-country heterogeneity related to fiscal institutions, exchange-rate regimes, or policy frameworks.

For Nigeria, an extensive body of research has examined the transmission of oil shocks. Studies such as Oyelami and Olomola (2016), Olayeni et al. (2020), Iwayemi and Fowowe (2011), and Ayadi (2005) find that oil price shocks affect GDP and the exchange rate, with evidence of asymmetry—negative shocks tend to generate stronger and less predictable depreciations. However, peer-reviewed studies on other major African oil exporters remain scarce. Algeria's experience, for instance, illustrates how fiscal buffers such as the Revenue Regulation Fund (FRR) and a managed exchange rate can moderate oil shocks, while Angola's liberalized regime and post-conflict vulnerabilities amplify volatility. Egypt's partial diversification and large energy subsidies create distinct transmission channels through budgetary and price mechanisms. Libya, though a major oil exporter, is often excluded from empirical analyses due to data discontinuities and its unique institutional setting.

Existing studies rarely examine these countries jointly, nor do they explicitly distinguish between oil price and oil revenue shocks. This leaves unresolved questions about the comparative fiscal transmission mechanisms across Africa's resource-dependent economies and how structural differences shape the propagation of shocks. The present study addresses this gap by jointly analyzing Algeria, Angola, Egypt, and Nigeria using a consistent framework that allows for both shared dynamics and country-specific heterogeneity.

2.2.4 Methodological Developments: From SVAR to Hierarchical Panel VAR

Methodologically, the study of oil shocks has evolved considerably. Early empirical work relied on reduced-form time-series regressions that could not distinguish between exogenous and endogenous sources of variation. The introduction of Structural Vector Autoregressions (SVARs) by Sims (1980) and Bernanke et al. (1997) allowed for the identification of structural shocks using zero restrictions through recursive Cholesky decompositions. Kilian (2009) later applied this framework to oil markets. Subsequent studies introduced sign restrictions to impose theoretically consistent constraints without enforcing rigid contemporaneous exclusions (Uhlig, 2005; Rubio-Ramirez et al., 2010). This innovation greatly improved the credibility of identification schemes and is now widely applied in the analysis of oil market disturbances (Chudik and Fidora, 2011; Kilian and Murphy, 2012; Kilian and Murphy, 2014;Baumeister and Hamilton, 2019; Kilian, 2022). In this study, we combine zero and sign restrictions to identify structural shocks. Given the annual frequency of our data, relying solely on contemporaneous zero restrictions would be overly restrictive, so we complement them with sign constraints that reflect theory-consistent responses.

Recognizing that shocks may transmit differently across economies, recent research increasingly employs Hierarchical Panel VAR (HPVAR) models to combine efficiency gains from data pooling with flexibility in capturing country-specific dynamics. The HPVAR approach (Canova and Ciccarelli, 2013; Gelman, 2006) treats country coefficients as random draws from a common distribution, thus exploiting cross-sectional information without imposing full homogeneity. Jarociński (2010) demonstrates its usefulness for identifying both common patterns and heterogeneity in monetary policy transmission across Europe. Building on these insights, recent developments in hierarchical Bayesian modeling (Gefang et al., 2023; Ter Steege, 2024) have further advanced the analysis of large VAR systems and improved computational efficiency.

In this study, we build on this framework by combining country-level SVARs with a Bayesian HPVAR framework. The dual approach enables the identification of both the average effect of oil revenue shocks across African exporters and the country-specific responses that reflect differences in fiscal institutions, macroeconomic structures, and adjustment mechanisms. This integrated framework represents a significant step forward compared with existing single-country or homogenous-panel studies, allowing for more credible inference on how oil revenue shocks shape fiscal and macroeconomic outcomes in resource-dependent African economies.

Overall, the literature demonstrates that oil shocks—whether in prices or revenues—play a central role in shaping macroeconomic outcomes in oil-exporting economies. Yet despite extensive research, we identify some gaps that motivate this study. Existing studies often rely on oil prices rather than total revenues and focus on single-country settings. Very few have examined the heterogeneous fiscal transmission of oil revenue shocks across African economies using a unified framework. This paper addresses these gaps by integrating a sign-restricted SVAR identification strategy within a hierarchical Bayesian panel framework, offering new comparative insights into how fiscal and macroeconomic responses to oil revenue shocks differ across countries and over time.

2.3 Theoretical Framework

The macroeconomic consequences of oil shocks have been extensively examined in the literature, beginning with the seminal contribution of Hamilton (1983), who demonstrated that oil price increases are systematically followed by economic downturns. His findings challenged the view that oil price movements merely coincided with recessions driven by other factors. For oil-importing economies, higher oil prices increase production costs and consumer prices, generating stagflationary pressures. In contrast, for oil-exporting economies, a positive oil price shock raises export receipts and national income, stimulating aggregate demand. This global income transfer from consumers to producers, emphasized by Barsky and Kilian (2002), represents a key adjustment mechanism whose effects depend on the relative marginal propensities to consume and import of both groups.

Subsequent theoretical work has identified several transmission channels through which oil shocks affect macroeconomic outcomes. Given the low short-run price elasticity of oil demand, firms and households cannot immediately adjust consumption and production patterns (Kliesen, 2008), leading to costly sectoral reallocations. At the firm level, monopolistic producers may raise markups following cost shocks (Rotemberg and Woodford, 1996), while reduced energy use can lower capital utilization and output (Finn, 2000). These mechanisms are reinforced by uncertainty effects: as Bernanke (1983) argued, oil price volatility can delay investment and hiring decisions, amplifying output fluctuations. DePratto et al. (2009) incorporated these mechanisms into open-economy New Keynesian models, showing that oil shocks influence output and inflation through multiple channels—supply-side (cost-push), demand-side (income and uncertainty), and real activity effects.

Another important transmission mechanism operates through the real exchange rate. The Dutch Disease hypothesis (Corden, 1984) argues that commodity booms can cause the real exchange rate to appreciate, which in turn undermines the competitiveness of non-oil tradable sectors and limits structural diversification. This channel is particularly relevant for African oil exporters, where production structures remain concentrated in the primary sector. The Dutch Disease mechanism connects directly with the broader resource-curse literature, which argues that dependence on natural resources often results in slower long-run growth and greater macroeconomic volatility (Sachs and Warner, 1995; Auty, 2002; Collier and Goderis, 2012). These adverse outcomes are often amplified by fiscal mismanagement, rent-seeking, and underinvestment in human and physical capital (Sachs and Warner, 2001; Frankel, 2010). In weak institutional environments, positive revenue shocks can trigger the voracity effect, whereby competing interest groups drive excessive public spending during booms, outpacing revenue growth and worsening macroeconomic instability (Tornell and Lane, 1999).

Building on these theoretical insights, we do not focus not on oil price shocks per se in this study, but on oil revenue shocks—unexpected variations in total income derived from the oil sector. For oil-exporting economies, the fiscal sector serves as the primary transmission channel through which these shocks affect the broader economy. When oil revenues rise, government expenditure typically expands, stimulating aggregate demand and output; conversely, negative shocks force contractionary fiscal adjustments that depress growth. This fiscal channel captures both the direct spending response and its multiplier effects on private activity. At the same time, higher oil revenues appreciate the real exchange rate by increasing foreign exchange inflows and domestic absorption, giving rise to the Dutch Disease channel, which erodes external competitiveness and constrains diversification.

Accordingly, the theoretical framework underlying this study emphasizes two principal mechanisms: (i) The fiscal transmission channel, through which oil revenue shocks alter government spending, aggregate demand, and output dynamics; and (ii) The real exchange rate (Dutch Disease) channel, through which these shocks influence external competitiveness and relative price adjustments. By examining these channels jointly, the analysis captures both the short-run demand effects of fiscal expansions and the medium-run structural implications of real exchange rate movements. These mechanisms provide the theoretical basis for the empirical strategy that follows, where Structural and Hierarchical Panel VAR models are used to quantify the macroeconomic and fiscal responses to oil revenue shocks in African oil-exporting economies.

2.4 Economic Overview of Countries in the Study

Macroeconomic series for the four countries display both common patterns and country-specific dynamics, with Nigeria standing out for its pronounced volatility. Nigeria's profile is dominated by sharp swings in oil revenue, notably the boom of the 2000s. The surge in oil income during that decade was closely mirrored by government expenditure, indicating a strongly procyclical fiscal stance. The inflation series for Nigeria is particularly striking, showing an extended period of very high inflation during the 1980s that far exceeded the levels observed in the other countries. Despite these large fluctuations in oil revenue and spending, Nigeria's GDP has followed an upward trajectory, accelerating markedly since the 2000s.

Algeria also exhibits substantial fiscal dependence on oil revenues, with government expenditure rising steadily from the 1990s onward. However, spending has fluctuated less than oil revenue, suggesting some degree of expenditure smoothing. Inflation was high during the 1980s and early 1990s but subsequently declined and remained relatively contained. Algeria's GDP displays a remarkably consistent and robust growth pattern throughout the period, making it one of the most stable economies in the sample in terms of real output performance.

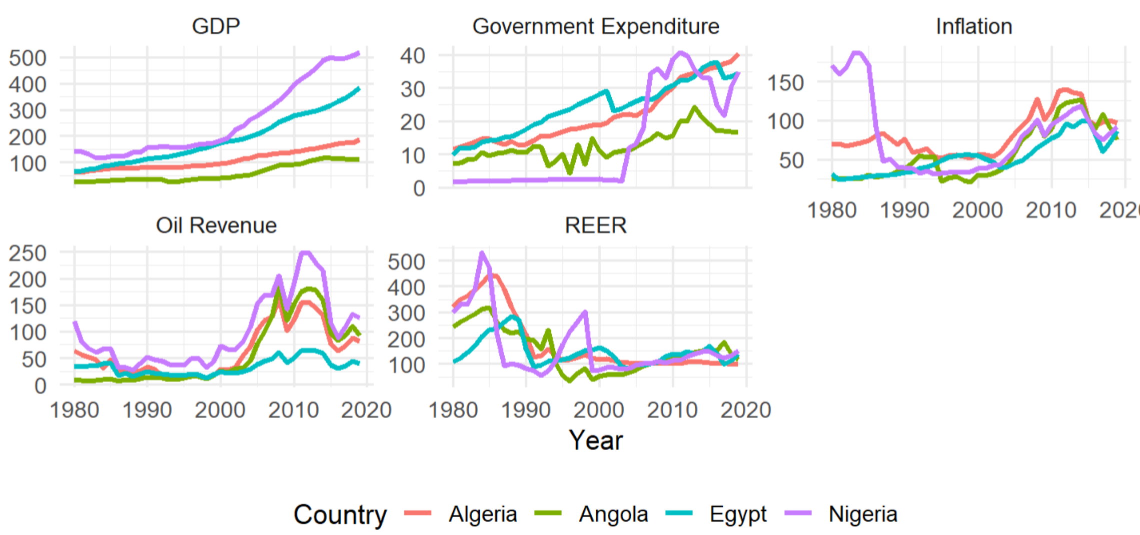


Figure 2.1: GDP and Government Expenditure are expressed in billions of U.S. dollars at constant 2015 prices; Inflation is measured by the GDP deflator (base year = 2015); Oil revenue (World Bank Brent Price × EIA oil production volume) is in billions of U.S dollars; REER is CPI -based. All series are plotted in levels.

Angola's macroeconomic trends illustrate a strong dependence on oil revenues. Both GDP and government expenditure expanded rapidly during the 2000s oil boom, reflecting the impact of rising export earnings. Oil revenues surged between 2004 and 2012 before declining sharply after 2014. Inflation, initially moderate, rose considerably in the 2000s, consistent with the regional trend of higher prices during oil booms. These patterns point to the prevalence of procyclical fiscal policy and highlight the macroeconomic volatility characteristic of oil-dependent economies.

In contrast, Egypt's economy appears far less dependent on oil revenues. The oil revenue series remains consistently low and stable throughout the period. Correspondingly, government expenditure follows a smooth and persistent upward trend, without the pronounced procyclicality observed elsewhere. Egypt's GDP exhibits the most steady and sustained growth trajectory in the sample, highlighting its relatively diversified production base and lower exposure to oil-market fluctuations.

Overall, the plots of the macroeconomic indicators reveal a blend of shared and idiosyncratic dynamics across the four countries. Periods of synchronized movements in oil revenues—such as the shared oil boom and inflationary pressures of the 2000s—suggest the presence of common shocks operating at the regional or global level. Conversely, episodes such as Nigeria's extreme inflation in the 1980s or Algeria's comparatively smoother fiscal adjustments highlight country-specific dynamics shaped by domestic institutions and policy frameworks.

These observations provide a strong empirical rationale for the dual methodological approach adopted in this study. The Hierarchical Panel VAR (HPVAR) framework captures the common regional responses to oil revenue shocks and the degree of cross-country variation, while the country-specific SVARs identify the mechanisms driving national differences in transmission. The combination of both methods therefore offers a coherent analytical strategy: the HPVAR reveals the systematic effects of oil revenue shocks across African oil exporters, whereas the SVARs uncover the heterogeneity that reflects each country's unique macroeconomic and fiscal context. The observed data patterns thus provide both the visual and empirical foundation for this integrated econometric design.

2.5 Empirical Analysis

2.5.1 Data

Given the annual dataset adopted in this study, we adopt a compact system of equations to minimize over-parameterization. Our empirical models focus on four key macroeconomic variables in their log-difference forms: oil revenue (proxied by oil rents), government consumption expenditure, Gross Domestic Product (GDP), inflation (proxied by the GDP deflator). We also adopt the real effective exchange rate (REER) in place of inflation to capture the Dutch Disease channel, a key transmission mechanism discussed in the theoretical section. Given the limited sample size (39 observations), including both inflation and REER in the same model would risk overparameterization in the single-country SVAR models. The study period is from 1980 to 2019 (before log-difference transformation).

We use government consumption as our measure of government spending due to data availability and consistency across countries over time. This approach aligns with prior studies, such as Perotti (2007) and Ilzetzki et al. (2013), which argue that government consumption offers a cleaner and more direct measure for identifying fiscal shocks in VAR models. It is preferred over broader expenditure aggregates because it is less influenced by cyclical fluctuations and other automatic stabilizers, making it less endogenous to current economic conditions. To ensure consistency across countries, we use the Bruegel CPI-based REER series constructed with 65 trading partners, as it offers longer coverage for our sample economies[1]. We verify robustness by comparing results obtained with the Bruegel REER series to those for Algeria and Nigeria, using the World Bank data. The findings remain consistent, suggesting that our results are not sensitive to the choice of REER source.

Consistent with standard practice in the macroeconomics and fiscal policy literature, all variables were initially expressed in logarithmic form. This transformation helps stabilize variance, reduces skewness, and allows the interpretation of results in terms of growth rates and elasticities, which are economically meaningful. Stationarity was examined using the Augmented Dickey-Fuller (ADF) test on the log-transformed series. The results indicated that several variables in log levels were non-stationary,

[1] The REER is measured as the nominal effective exchange rate (NEER) adjusted for relative consumer prices between the home country and its trading partners. Trade weights are derived from gross trade values following Bayoumi et al. (2006), as in the conventional approach used by major institutions such as the IMF and BIS. An increase in the REER implies a real appreciation—domestic goods become relatively more expensive—indicating a possible loss of external competitiveness. In this study, a positive REER response to an oil shock is interpreted as evidence of the Dutch disease effect.

Table 2.1: Data Definitions and Sources

Variables	Definition of variables	Data source
Oil revenue	Oil rents (% of GDP)	World Development Indicators (WDI), World Bank
Government expenditure	General government final consumption expenditure (billions US$, constant 2015)	UNECA
GDP	Gross Domestic Product (billions US$, constant 2015)	UNECA
Inflation	GDP deflator (index; 2015 = 100)	UNECA
REER	CPI-based real effective exchange rate (65 trading partners)	Bruegel REER Database

Notes: All variables are transformed into log-differences prior to estimation.

as the null hypothesis of a unit root could not be rejected at the 5% significance level. While some variables appeared stationary in their log-transformed levels for certain countries, to maintain consistency across all countries and variables, the first differences of the log-transformed series were taken, after which the ADF test confirmed stationarity, implying that the series are integrated of order one, I(1). Accordingly, the model was estimated using the log-difference series[2]. Results for the stationarity test can be found in Appendix B.1.

Unit root test results for the REER (in log-levels) are mixed across countries, with the null of a unit root rejected for Algeria and Egypt but not for Nigeria and Angola. Given this heterogeneity, a common log-difference specification is adopted for all countries to maintain a consistent panel VAR specification and ensure that impulse responses are directly comparable across the countries. Applying different transformations across countries would complicate both estimation and interpretation. Nevertheless, because differencing may alter the persistence of the responses,

[2]Variables are expressed in first differences of logarithms to account for the heterogeneity of Algeria, Angola, Egypt, and Nigeria. Unlike the relatively homogeneous WAEMU countries, level differences here could produce highly volatile series that obscure co-movements, whereas log-differencing stabilizes variance and allows meaningful cross-country comparisons.

Table 2.2: Summary Statistics by Variable across Countries

Variable	**Country**	**Obs**	**Mean**	**Std. Dev.**	**Min**	**Max**
GDP	Algeria	39	0.03	0.02	-0.02	0.07
	Angola	39	0.04	0.07	-0.27	0.14
	Egypt	39	0.05	0.02	0.01	0.09
	Nigeria	39	0.03	0.05	-0.12	0.14
Government Spending	Algeria	39	0.03	0.04	-0.09	0.11
	Angola	39	0.02	0.32	-0.84	1.10
	Egypt	39	0.03	0.06	-0.24	0.19
	Nigeria	39	0.08	0.34	-0.29	1.90
Oil Revenue	Algeria	39	-0.02	0.27	-0.89	0.45
	Angola	39	0.01	0.42	-0.93	1.01
	Egypt	39	-0.05	0.32	-0.84	0.56
	Nigeria	39	-0.03	0.57	-1.84	1.08
Inflation	Algeria	39	0.01	0.11	-0.29	0.21
	Angola	39	0.03	0.21	-0.88	0.36
	Egypt	39	0.03	0.11	-0.31	0.19
	Nigeria	39	-0.02	0.18	-0.65	0.23
REER	Algeria	39	-0.03	0.12	-0.52	0.17
	Angola	39	-0.02	0.28	-0.80	0.68
	Egypt	39	0.01	0.19	-0.60	0.21
	Nigeria	39	-0.02	0.33	-1.39	0.46

Source: Author's computations (variables in log-differences).

an alternative specification is also estimated in which oil revenue and the REER are included in levels, while the remaining variables are retained in growth rates (see Appendix B.15 for the impulse responses of government revenue and REER to an oil revenue shock).

Since the model is estimated using log-differenced data, the impulse response functions (IRFs) reflect how growth in one variable responds to growth in another over time. Specifically, a shock to government expenditure represents an unexpected increase in the growth rate of spending — for example, a sharper-than-anticipated rise relative to the previous year. The response of GDP then shows how its growth rate reacts to that shock. These responses therefore capture short-run dynamics in growth rates rather than permanent effects on levels.

2.5.2 Estimation Strategy

As mentioned earlier, this study employs two VAR models. Given the distinct economic structures, policy frameworks, and institutional characteristics of the selected African countries, we estimate a fixed-coefficient SVAR model separately for each country. We also recognize that the countries in the sample share certain similar characteristics– developing African nations, strong reliance on oil exports and political instability in certain periods– and as such, adopt a Panel SVAR with hierarchical structure as a second methodology. This approach allows us to capture the specific dynamics and responses to fiscal policy shocks in each country, while also allowing for information sharing across countries through a common structure. If the results from the SVARs and HPVAR converge, this strengthens confidence in the findings. If they diverge, the differences are themselves informative, as they highlight country-specific structural or institutional features that may be obscured in pooled estimation.

This study uses a Bayesian VAR framework to analyze the data. For the individual country SVAR model, the Independent Normal–Wishart (INW) prior is used. This prior helps to stabilize the parameter estimates by pulling them towards a prior mean, a process known as Bayesian shrinkage, which reduces estimation variance and improves stability. To handle the panel structure of the data, we adopt a hierarchical prior specification. Under this framework, country-specific coefficients are assumed to be drawn from a common cross-country distribution, which induces partial pooling across countries. Even though all countries are observed over the same time period, this structure improves estimation precision by shrinking country-specific parameters toward a common mean in a short-sample setting with many parameters. This shrinkage reduces estimation noise and guards against overfitting, while still allowing for meaningful cross-country heterogeneity. These features are particularly important given the use of annual data and enhance the robustness of the resulting estimates.

2.5.3 Fixed Coefficients Bayesian VAR

In this section, we discuss the fixed-coefficient SVAR model. The benchmark VAR specification is given by

$$y_t = \alpha + A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + \varepsilon_t, \tag{2.1}$$

where

$$y_t = \begin{bmatrix} \Delta OIL_t \\ \Delta GOV_t \\ \Delta GDP_t \\ \Delta INF_t \end{bmatrix},$$

for the baseline model, with ΔOIL_t, ΔGOV_t, ΔGDP_t, and ΔINF_t denoting the growth rates of oil revenue, real government expenditure, real GDP, and inflation, respectively. In the alternative specification, ΔINF_t is replaced by $\Delta REER_t$, the growth rate of the real effective exchange rate. The vector α contains the intercepts, $A_1, \ldots, A_p$ are 4×4 coefficient matrices, and ε_t is a 4×1 vector of reduced-form shocks. Shocks are assumed to be distributed according to the normal distribution $\varepsilon_t \sim N(0, \Sigma)$. The model can be rewritten equivalently as:

$$y_t = X_t' \beta + \varepsilon_t, \tag{2.2}$$

where $X_t' = I_n \otimes [1, y_{t-1}', \ldots, y_{t-p}']'$ and β is a vector that incorporates the coefficients in α and $A_1, \ldots, A_p$.

Structural VARs allow us to obtain responses of variables to orthogonal shocks. Equation (2.2) above can be referred to as the reduced form VAR. The alternative specification of this model is its structural VAR form:

$$y_t = X_t'\beta + D\eta_t, \tag{2.3}$$

where we assume $\eta_t \sim N(0, I)$ for simplicity, so that the structural shocks are mutually orthogonal and have unit variance. This implies that the structural shocks are uncorrelated and arise independently. This assumption is necessary to obtain meaningful impulse response functions that are not affected by correlated shocks.

$$\varepsilon_t = D\eta_t. \tag{2.4}$$

Bayesian VAR estimation

In the fixed-coefficient Bayesian VAR (BVAR) framework, the key parameters are the coefficient vector β and the residual covariance matrix Σ. This study adopts the Independent Normal-Wishart (INW) prior following Dieppe et al. (2016) and Murarașu et al. (2023), which specify a multivariate normal prior for β with mean β_0 and covariance matrix Ω_0:

$$\beta \sim N(\beta_0, \Omega_0). \tag{2.5}$$

The prior for Σ follows an inverse Wishart distribution with scale matrix S_0 and degrees of freedom α_0:

$$\Sigma \sim IW(S_0, \alpha_0). \tag{2.6}$$

Compared to the standard Normal-Wishart prior, the INW prior provides greater flexibility in how uncertainty is modeled. Instead of forcing a fixed relationship between the variability of the model's coefficients and the variability of the residuals, it treats them as separate. In simpler terms, the INW prior allows the uncertainty in the estimated relationships among variables to differ from the uncertainty in the model's shocks, making the estimation more realistic and less restrictive (Pacifico, 2021).

A trade-off is that the INW prior does not yield a closed-form posterior. Nevertheless, its conditional posterior distributions have tractable forms, enabling estimation via numerical methods such as the Gibbs sampler. The likelihood function, prior specifications, and kernel distribution are presented in Appendix (B.5.1), and the Gibbs sampling algorithm is implemented following Dieppe et al. (2016).

2.5.4 Panel VAR Model with Hierarchical Priors

We employ a Hierarchical Panel VAR (HPVAR) to capture heterogeneity and commonalities across oil-exporting economies. Unlike fixed-coefficient SVARs suited for single-country analysis or traditional panel VARs that impose identical dynamics, the HPVAR accommodates cross-country heterogeneity while avoiding possible overparameterization with separate VARs. Using a Bayesian framework, the model allows coefficients to vary across units while assuming they are generated from a common distribution characterized by a shared mean and variance (Dieppe et al., 2016, Canova and Ciccarelli, 2013). The panel data framework allows the model to pool information across countries and improve precision, especially with short samples. This makes the HPVAR well suited to studying oil revenue shocks in African exporters, as it captures both country-specific dynamics and shared responses. The hierarchical prior follows Dieppe et al. (2016) in line with Jarociński (2010). It improves on Zellner and Hong (1989) by estimating not only the country-level coefficient vectors β_i but also their residual covariance matrices Σ_i and the common mean and covariance of the VAR coefficients b and Σ_b.

In this context, the hierarchical structure allows the model to capture both shared and idiosyncratic components of fiscal transmission. Common patterns across countries reflect the global nature of oil revenue shocks, while country-specific coefficients capture differences in fiscal policy responses, absorption capacity, and macroeconomic adjustment mechanisms. This distinction is particularly important in the context of African oil-exporting economies, where institutional and structural heterogeneity plays a central role in driving the persistence and magnitude of macroeconomic responses to external shocks.

Consider individual unit i, where each country responds only to its own lagged variables. The country-specific VAR is given by

$$y_{i,t} = C_i + A_i^1 y_{i,t-1} + \cdots + A_i^p y_{i,t-p} + \varepsilon_{i,t}, \tag{2.7}$$

where

$$y_{i,t} = \begin{bmatrix} \Delta OIL_{i,t} \\ \Delta GOV_{i,t} \\ \Delta GDP_{i,t} \\ \Delta INF_{i,t} \end{bmatrix},$$

for the baseline model, with $\Delta OIL_{i,t}$, $\Delta GOV_{i,t}$, $\Delta GDP_{i,t}$, and $\Delta INF_{i,t}$ denoting the growth rates of oil revenue, real government expenditure, real GDP, and inflation, respectively. In the alternative specification, $\Delta INF_{i,t}$ is replaced by $\Delta REER_{i,t}$, the growth rate of the real effective exchange rate. The vector C_i contains the country-specific intercepts, $A_i^1, \ldots, A_i^p$ are 4×4 coefficient matrices, and $\varepsilon_{i,t}$ is a 4×1 vector of reduced-form shocks.

The reduced-form innovations are assumed to satisfy

$$\varepsilon_{i,t} \sim \mathcal{N}(0, \Sigma_i). \tag{2.8}$$

In compact form,

$$y_i = \bar{X}_i \beta_i + \varepsilon_i, \tag{2.9}$$

where $y_i = \text{vec}(y'_{i,t})$ is the stacked vector of endogenous variables for country i over T periods, $\bar{X}_i = I_n \otimes [1, y'_{i,t-1}, \ldots, y'_{i,t-p}]$ is the block-diagonal regressor matrix, $\beta_i = \text{vec}(C_i, A_i^1, \ldots, A_i^p)$ is the vector of coefficients, and $\varepsilon_i = \text{vec}(\varepsilon'_{i,t})$ is the stacked vector of error terms.

The lag length is set to one for all VAR specifications, as selected by the Deviance Information Criterion (DIC), which is particularly suitable for Bayesian models and helps avoid over-parameterization in small samples.

Details on the likelihood, priors, and posterior are provided in Appendix (B.5.2). Since the parameters are interdependent, analytical marginal distributions cannot be derived. Instead, conditional distributions are obtained and sampled via Gibbs sampling. The Gibbs sampler iteratively draws from these conditionals to approximate the joint posterior, implemented using the BEAR toolbox in MATLAB (Dieppe et al., 2016). Impulse response functions are computed from the posterior draws to examine how shocks propagate through the system over time.

2.5.5 Identification Strategy

The identification strategy is twofold, combining zero and sign restrictions within the same framework to ensure robustness. Zero restrictions, pioneered by Sims (1980) and applied to oil shocks by Bernanke et al. (1997), assume that certain variables do not respond to specific shocks in the same period. In this context, we assume that oil revenue is not affected contemporaneously by GDP, consistent with the view that revenue from oil exports is largely exogenous to domestic macroeconomic fluctuations in the short run.

Only selective zero restrictions are imposed here, complemented by sign restrictions to enhance theoretical plausibility. The latter impose economically consistent directions of response, as advocated by Uhlig (2005) and applied to the oil market by Kilian and Murphy (2012). For instance, an oil revenue shock is required to generate a contemporaneous increase in government spending, GDP and inflation. This restriction follows the fiscal-resource literature, which highlights the expansionary effect of resource windfalls on output and prices through higher public spending and aggregate demand.

The sign restrictions are motivated by standard theoretical predictions associated with positive oil revenue shocks in oil-dependent economies. A positive oil revenue shock is expected to raise government revenues contemporaneously, reflecting the direct fiscal impact of higher oil prices or production. Increased fiscal space and higher national income generate demand pressures, leading to higher inflation and higher output in the short run. Consistent with the Dutch disease hypothesis, higher oil revenues also induce real exchange rate appreciation, reflecting increased domestic prices and capital inflows. Government spending is assumed to increase following oil revenue windfalls due to procyclical fiscal behavior commonly observed in resource-rich economies. These restrictions are imposed only on impact to identify the structural shocks.

Economically, the restriction pattern is intended to distinguish exogenous oil revenue innovations from other structural disturbances. An exogenous increase in oil revenues, arising from changes in world oil prices or oil production, is expected to raise government fiscal resources, leading to higher government expenditure, stronger aggregate demand, upward pressure on domestic prices, and an appreciation of the real exchange rate through the Dutch disease mechanism. By contrast, domestic demand, fiscal, monetary, or productivity shocks may generate similar movements in one or more macroeconomic variables, but they would not simultaneously satisfy the full set of restrictions. In particular, the zero restriction prevents domestic macroeconomic conditions from contemporaneously driving oil revenues, while the joint sign restrictions require that the identified shock produces the complete macroeconomic response expected following an exogenous oil revenue windfall. Consequently, the admissible structural shocks are those that are consistent with both the institutional exogeneity of oil revenues and the theoretical transmission mechanism linking oil revenue windfalls to domestic economic activity.

The implementation in the BEAR toolbox follows the methodology of Arias et al.

(2018), which allows for the joint imposition of zero, sign, and magnitude restrictions within a Bayesian VAR framework.

2.6 Results

2.6.1 Diagnostic Checks

In the Bayesian estimation framework employed in this study, model stationarity is essential for reliable inference. Posterior distribution plots of the VAR coefficients provide a preliminary visual check for convergence. Smooth, unimodal posterior distributions suggest well-sampled densities. Convergence was further assessed using the Effective Sample Size (ESS) and the Autocorrelation Function (ACF). ESS quantifies the number of effectively independent draws, ensuring robust MCMC inference; with 5,000 MCMC draws, ESS values ranged from 3,500 to 5,000, indicating good mixing and low autocorrelation. The ACF measures correlations across successive lags, helping to evaluate chain independence. The ACF plots show rapid decay to zero, followed by random fluctuations, confirming efficient mixing and convergence. For all parameters examined, most Geweke Z-scores fell within the conventional acceptance range ($|Z| < 1.96$), indicating no significant difference between the means of the early and late segments of the chain and thus supporting convergence. We therefore conclude that all four diagnostic tests provide evidence of convergence. Outputs from all four diagnostic tests can be found in the Appendix B.2

2.6.2 Results and Discussion

This section presents the estimated impulse responses from both the hierarchical panel VAR (HPVAR) and the country-specific structural VAR (SVAR) models (with 68% credibility intervals). The HPVAR captures common dynamics across the panel of oil-exporting economies, while the SVARs allow for country-level heterogeneity in the transmission of shocks. For clarity and coherence, the discussion is organized around three key response channels. Each subsection first summarizes the HPVAR results and then contrasts them with the individual-country SVAR outcomes.

To guide the interpretation of these results, we emphasize three main transmission mechanisms. First, oil revenue shocks operate primarily through a fiscal channel, whereby increases in government revenues translate into higher public spending, thereby stimulating aggregate demand and output in the short run. Second, this demand expansion generates inflationary pressures, particularly in economies with limited productive capacity or weaker monetary anchors. Third, higher oil revenues induce real exchange rate appreciation through a spending effect, as increased domestic absorption raises the relative price of non-tradable goods, consistent with the Dutch disease hypothesis. These mechanisms provide a coherent framework for interpreting both the common patterns identified by the HPVAR and the cross-country heterogeneity revealed by the SVAR models.

In this analysis, the contemporaneous responses of government expenditure, GDP, inflation and the REER to an oil revenue shock should be interpreted as identifying assumptions rather than empirical findings, as they are imposed through the sign restrictions used to identify the structural oil revenue shock.

Responses of Fiscal Policy to Oil Revenue Shocks

This subsection examines how government expenditure reacts to innovations in oil revenue. Figure 2.2 displays the impulse responses derived from the HPVAR and the country-specific SVAR models. Under the HPVAR framework, the impact response of government expenditure growth to an oil revenue shock is around 0.02 percent for Algeria and Egypt, and approximately 0.15 percent for Angola and Nigeria, with a 1 percent increase in the growth rate of oil revenue. As imposed, these responses suggest that fiscal spending tends to rise following positive oil revenue shocks, consistent with a procyclical fiscal stance. The country-specific SVAR results reveal variation in the magnitude of these effects- the contemporaneous impact is estimated at 0.02 percent for Algeria, 0.03 percent for Egypt, 0.04 percent for Angola, and 0.14 percent for Nigeria, given a 1 percent unexpected increase in oil revenue growth across the sample. The responses generally persist for about four periods before going back to the steady state.

The relatively stronger fiscal responses observed for Nigeria and Angola likely reflect their greater fiscal dependence on oil revenues and more limited mechanisms for insulating budgets from commodity price swings. In both countries, oil income constitutes a dominant share of public revenue, which may make expenditure more responsive to oil price fluctuations. Previous studies have shown that fiscal policy in highly oil-dependent economies often exhibits a procyclical pattern, with spending rising during oil booms and contracting in downturns (El Anshasy and Bradley, 2012; Lopez-Murphy and Villafuerte, 2010). In contrast, the more muted responses for Algeria and Egypt may be associated with stronger fiscal stabilization practices or a more diversified revenue base. Algeria's stabilization fund, for instance, may have helped smooth fiscal outcomes, while Egypt's lower reliance on oil receipts may have moderated its sensitivity to oil price movements.

Overall, comparison between the HPVAR and SVAR estimates suggests that while all four economies share exposure to oil shocks, the scale of fiscal adjustment depends on domestic institutions and revenue composition. More fundamentally, these results indicate that oil revenue shocks operate primarily through a fiscal transmission channel, whereby increases in public revenues translate into higher government spending and short-run macroeconomic expansion. This finding is consistent with broader evidence on cross-country heterogeneity in fiscal responses among resource-exporting countries.

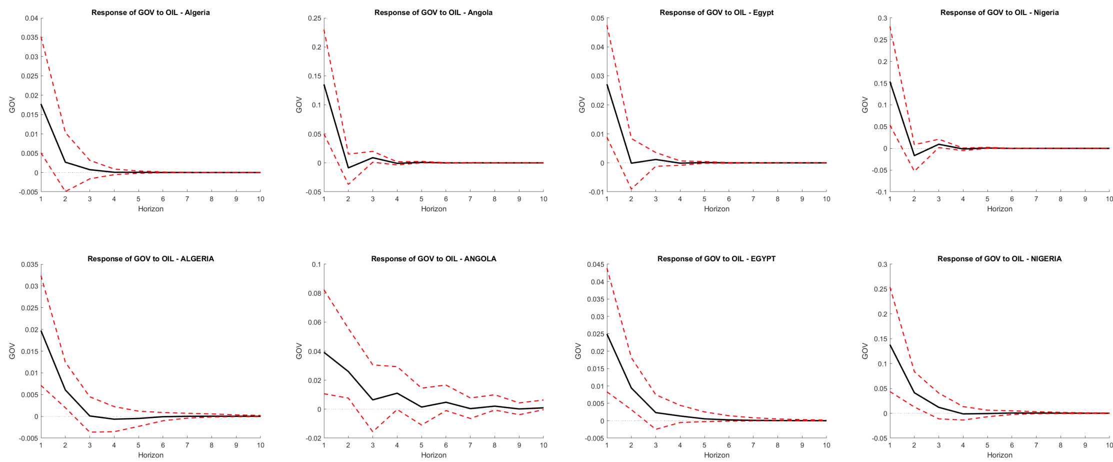


Figure 2.2: Impulse responses of government expenditure growth to a shock in oil revenue growth for Algeria, Angola, Egypt, and Nigeria. Top row: HPVAR model; Bottom row: SVAR model.

Responses of Macroeconomic Conditions to Oil Revenue Shocks

This subsection analyzes how key macroeconomic indicators—real GDP growth and inflation—respond to oil revenue shocks. Figure 2.3 displays the impulse responses of GDP growth, while Figure 2.4 presents the corresponding responses for inflation. Together, these figures illustrate how fluctuations in oil revenues influence output dynamics and price stability, as well as the persistence of these effects over time.

In the HPVAR, the response of GDP growth is immediately positive as imposed, and persists for roughly three periods. Impact effects range from 0.007 percent in Egypt to 0.03 percent in Angola, implying that a 1 percent increase in oil revenue growth is associated with a 0.007–0.03 percent contemporaneous rise in GDP growth. The country-specific SVAR results reveal a similar pattern but with greater persistence, lasting up to five periods, and impact magnitudes ranging from 0.008 percent (Egypt) to 0.04 percent (Angola).

These positive responses over the first four periods are consistent with the notion that oil revenue windfalls provide additional fiscal resources for government expenditure, supporting domestic demand and short-term economic expansion. Differences in magnitude and persistence across countries likely reflect variation in oil dependence, policy frameworks, and institutional quality. Angola and Nigeria exhibit the largest responses, reflecting their heavier reliance on oil revenues and more procyclical fiscal behavior. Egypt's smaller effects correspond to its more diversified economy and limited exposure to oil-sector fluctuations, while Algeria's moderate reaction may stem from high oil dependence combined with stronger fiscal management and slower transmission of oil windfalls to domestic spending.

The positive impulse responses of inflation to oil revenue shocks beyond the impact period are consistent with the presence of both demand-pull and cost-push inflationary pressures. Angola records the largest inflationary effect: a 1 percent increase in oil revenue growth raises inflation by about 0.09 percent in the subsequent period. Nigeria follows closely, with a 0.08 percent increase, while Algeria and Egypt experience more moderate responses of 0.04–0.05 percent. The pronounced inflationary effects in Angola and Nigeria likely reflect their strong fiscal procyclicality and limited monetary anchors, which amplify demand pressures during oil booms. In contrast, Algeria and Egypt's more diversified economic bases and relatively disciplined fiscal policies appear to dampen the transmission of oil revenue shocks to domestic prices.

Overall, the results suggest that oil revenue windfalls stimulate output in the short run but also generate inflationary pressures, especially in economies with weaker fiscal stabilization mechanisms. To some extent, these findings reinforce the role of countercyclical macroeconomic policies to mitigate overheating risks and sustain economic stability in oil-dependent economies.

Examining the Dutch Disease Effect

This subsection examines whether the estimated REER dynamics following an identified oil revenue shock are consistent with the Dutch disease hypothesis. Theoretically, a positive and persistent appreciation of the real effective exchange rate (REER)

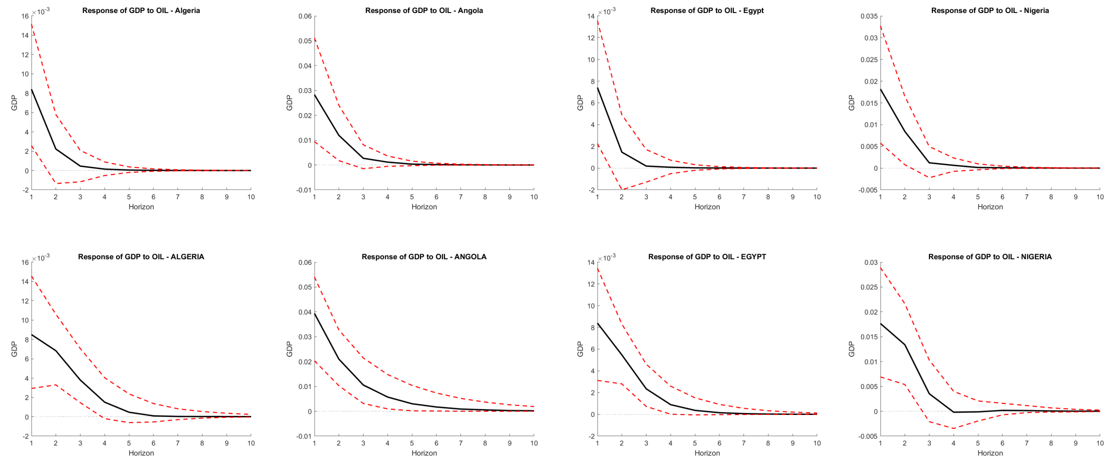


Figure 2.3: Impulse responses of GDP growth to a shock in oil revenue growth for Algeria, Angola, Egypt, and Nigeria. Top row: HPVAR model; Bottom row: SVAR model.

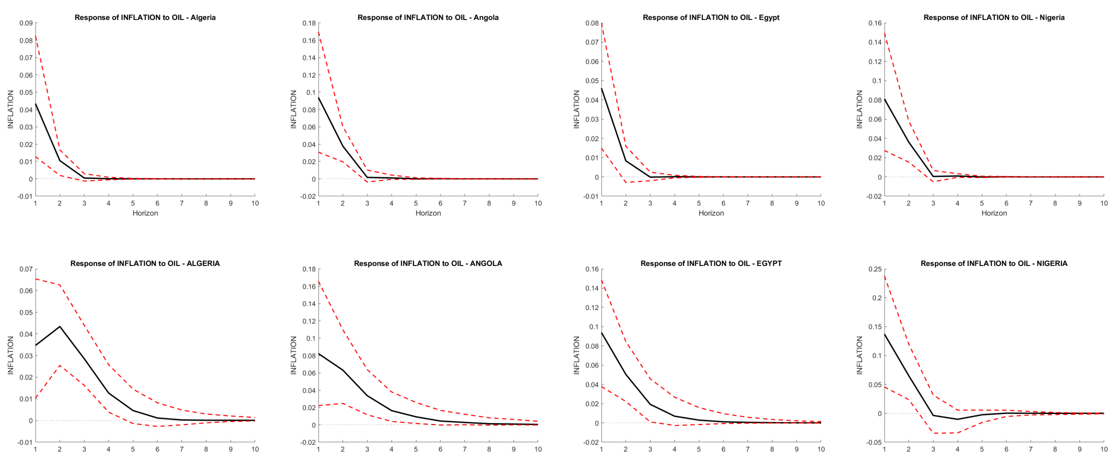


Figure 2.4: Impulse responses of Inflation to a shock in oil revenue growth for Algeria, Angola, Egypt, and Nigeria. Top row: HPVAR model; Bottom row: SVAR model.

following an oil windfall would suggest that oil inflows place upward pressure on domestic prices, potentially eroding the competitiveness of the tradable sector.

Figure 2.5 presents the impulse responses of REER growth to an oil revenue shock estimated by the HPVAR and the country-specific SVAR models. Across both models, the estimated REER responses remain positive over the short run, typically peaking within the first two to three periods before gradually dissipating toward zero. This short-lived appreciation may reflect temporary fiscal expansions or monetary interventions, such as foreign exchange sterilization, that mitigate the full impact of oil inflows on domestic prices and the exchange rate.

The HPVAR results reveal larger and more heterogeneous short-run appreciations, particularly for Nigeria, highlighting the model's ability to capture both shared and country-specific dynamics. The stronger response in Nigeria likely reflects its heavy dependence on oil revenues and limited export diversification, which amplify the effect of oil windfalls on domestic absorption and prices. In contrast, Algeria exhibits smaller responses, consistent with its relatively structured fiscal framework and stabilization mechanisms that help moderate exchange-rate pressures.

Quantitatively, the impact responses from the HPVAR are approximately 0.05 for Algeria, 0.09 for Angola, 0.08 for Egypt, and 0.13 for Nigeria, implying that a 1 percent increase in oil revenue growth leads to about a 0.05–0.13 percent appreciation in the REER. The SVAR estimates produce similar magnitudes, reinforcing the robustness of the findings across model specifications.

In Egypt, the relatively pronounced REER response may be partly explained by its exchange-rate regime transitions. The 1991 unification of Egypt's multiple exchange-rate systems and the subsequent liberalization in 2016 likely increased short-run exchange-rate sensitivity to external shocks, even though the broader economy remains more diversified than those of other oil exporters.

In Nigeria, the SVAR results show a small depreciation in the third and fourth periods (though not statistically significant), suggesting that the medium-term effects are

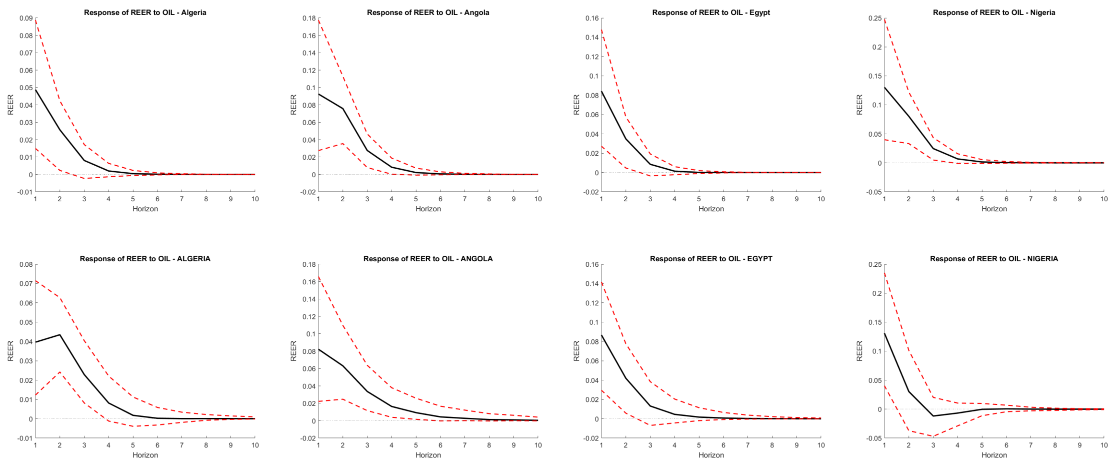


Figure 2.5: Impulse responses of REER to a shock in oil revenue for Algeria, Angola, Egypt, and Nigeria. Top row: HPVAR model; Bottom row: SVAR model.

weak or short-lived. This pattern aligns with previous studies (Ismail, 2010; Sala-i Martin and Subramanian, 2013), which argue that the degree of REER appreciation depends on how oil revenues are spent—with higher spending on imported goods or foreign-denominated assets generating little appreciation pressure. Overall, the estimated REER responses exhibit considerable cross-country heterogeneity in both magnitude and persistence, reflecting differences in fiscal institutions, expenditure patterns, and exchange-rate management.

2.7 Robustness Check

To verify that the baseline results are not driven by measurement choices or sample composition, three robustness exercises were conducted. These tests assess the stability of the estimated impulse responses under alternative definitions of key variables, data sources, and country coverage. Specifically, We: (i) replace the oil-rent variable with a quantity-based proxy of oil revenue constructed as the product of crude oil production and Brent crude prices; (ii) re-estimate the model using the World Bank real effective exchange rate (REER) data—available for Nigeria and Algeria—instead of the Bruegel dataset; and (iii) re-estimate the hierarchical panel

HPVAR model excluding Nigeria, to test whether its large size and potential data volatility unduly influence the pooled results. The impulse response functions from these estimations are presented in Appendix B.3.

Alternative Oil Revenue Definition

In the first robustness check, an alternative measure of oil revenue is constructed by multiplying each country's crude oil production (in million barrels per day), obtained from the U.S. Energy Information Administration, by the corresponding annual average Brent crude oil price from the World Bank Pink Sheet. Brent price data prior to 1987 are reconstructed based on historical estimates and were therefore excluded from the main baseline specification. Besides providing an alternative measure of oil revenue, this exercise addresses the concern that oil rents expressed as a share of GDP incorporate domestic output in the denominator, making the contemporaneous exogeneity assumption less transparent. The alternative proxy removes this direct dependence on GDP while continuing to capture variation in oil revenues.

When this quantity-based proxy replaces the oil-rents variable, the estimated impulse responses of government expenditure, REER, and GDP growth to oil revenue shocks remain largely unchanged in both sign and shape relative to the baseline. The primary difference is that convergence to the steady state occurs after approximately six periods, compared to four periods in the baseline. Minor variations in response magnitudes are observed for government expenditure and the REER, but the overall patterns of adjustment and persistence remain consistent. These results confirm the robustness of the findings to the specific definition of oil revenue.

Alternative REER Data Source

The second robustness exercise evaluates whether the results are sensitive to the choice of REER data source. While the baseline analysis uses the Bruegel REER dataset for its consistent cross-country coverage, the World Bank REER series is employed here for Nigeria and Algeria. After transforming the World Bank REER data into logarithmic differences and re-estimating the models, the resulting impulse responses exhibit the same qualitative patterns as those obtained from the baseline specification. The direction, timing, and magnitude of REER responses to oil revenue shocks remain highly similar, confirming that the results are not affected by dataset-specific measurement differences.

Excluding Nigeria from the Panel

Given Nigeria's large economic size and relatively volatile data, the third robustness check re-estimates the HPVAR model after excluding Nigeria from the sample. The resulting impulse responses preserve the same shapes and signs as in the baseline, with only minor differences in magnitude. This indicates that the pooled results are not disproportionately influenced by any single country and that the HPVAR estimates reflect common structural dynamics across the panel rather than being driven by Nigeria's economy alone.

Taken together, the three robustness checks—alternative oil revenue definitions, REER data sources, and country composition—consistently validate the baseline findings. Across all exercises, the impulse responses remain stable in sign, shape, and general magnitude. The persistence and direction of fiscal, macroeconomic, and exchange-rate adjustments to oil revenue shocks are therefore robust to plausible changes in model specification and data construction. These results reinforce confidence in the empirical analysis and support the conclusion that the observed effects of oil revenue shocks reflect genuine underlying structural relationships shared by the oil-exporting economies in the sample.

2.8 Conclusion

This study examined the effects of oil revenue shocks on fiscal and macroeconomic dynamics in four African oil-exporting economies. Unlike much of the existing literature that focuses on oil-price shocks, this analysis uses oil revenue as the external impulse—capturing both price and production effects and thus providing a more accurate measure of fiscal exposure in resource-dependent economies. Structural identification was achieved through zero and sign restrictions, while the empirical strategy combined a Hierarchical Panel VAR (HPVAR) and country-specific SVARs to address short-sample constraints and reveal cross-country heterogeneity.

This paper contributes to the literature in three ways. First, it offers a more comprehensive framework for analyzing external shocks in oil exporters by focusing on oil revenue rather than oil prices—thereby aligning the analysis with the fiscal realities that shape macroeconomic responses. Second, it advances empirical methodology by applying a hierarchical Bayesian VAR to a panel of African economies, an approach that leverages common information while preserving country-specific dynamics. Third, it provides comparative evidence on the magnitude and persistence of fiscal and macroeconomic responses to oil revenue shocks across African oil-exporting economies, highlighting similarities and differences.

The results show that the positive short-run impact of oil revenue shocks on government expenditure and output growth is transitory, typically dissipating within three to four periods. Inflation responses vary across countries—large in Nigeria and Angola but modest in Algeria and Egypt—reflecting differences in fiscal discipline, monetary management, and economic diversification. Also, the real exchange rate appreciates temporarily in response to oil windfalls.

Overall, the findings provide evidence that oil windfalls yield temporary and uneven benefits, shaped by fiscal institutions and policy frameworks. Strengthening fiscal rules, stabilization funds, and revenue diversification mechanisms remains essential to mitigate volatility and enhance macroeconomic resilience.

Future research could explore alternative identification strategies, including external-instrument approaches and market-based oil shock measures. It could also use higher-frequency or longer time-series data to better capture short-run adjustment dynamics and examine how the composition of fiscal spending, particularly the balance between capital and current expenditures, influences long-term growth and structural transformation.

Chapter 3

Fiscal Policy and Business Investment under Uncertainty: Evidence from a Threshold Panel VAR

Abstract

This paper examines the state-dependent effects of government spending shocks on business investment across seven OECD countries. Using a threshold panel vector autoregression (TVAR) framework, we allow the transmission of fiscal policy to vary between high- and low-uncertainty regimes. The results indicate differences in the responses of business investment across uncertainty regimes. During low-uncertainty periods, government spending shocks are associated with a decline in business investment, consistent with crowding-out effects. In contrast, during high-uncertainty episodes, the investment response is muted and statistically insignificant, suggesting that elevated uncertainty weakens firms' responsiveness to fiscal stimulus. Forecast error variance decompositions further indicate that while investment is largely driven by its own innovations in stable environments, macroeconomic fundamentals—particularly output and interest rates—play a greater role under high uncertainty. Overall, the findings suggest that uncertainty may influence the effectiveness of fiscal policy in shaping private capital formation.

3.1 Introduction

Fiscal policy remains a primary tool for stabilizing economic activity and supporting growth, particularly during periods of economic stress. Advanced OECD economies have relied heavily on government spending to counteract recessions, stimulate demand, and restore confidence in periods of severe economic shocks, like the global financial crisis and the COVID-19 pandemic—episodes characterized by both large fiscal expansions and heightened economic uncertainty. Yet, despite its widespread use, the effect of government spending on private investment remains theoretically ambiguous and empirically contested. This question is especially important because business investment plays a central role in long-run economic performance, and elevated uncertainty may alter how firms respond to fiscal shocks, potentially reshaping the transmission of policy during turbulent periods.

A substantial body of literature examines the impact of government spending on private investment, with mixed findings. Neoclassical models emphasize crowding-out effects, arguing that deficit-financed public spending raises real interest rates and the user cost of capital, thereby reducing private investment (Barro, 1974; Barro, 1981; Baxter and King, 1993). In contrast, Keynesian frameworks allow for crowding-in effects when higher public expenditure stimulates aggregate demand and improves firms' expectations (Woodford, 2011; Christiano et al., 2011a). Empirical evidence reflects this ambiguity: some studies document crowding out (Blanchard and Perotti, 2002; Alesina et al., 2002; Galí et al., 2007; Ramey, 2011b), while others find positive, non-monotonic, or negligible effects (Aschauer, 1989; Burnside et al., 2004; Perotti et al., 2007; Mountford and Uhlig, 2009; Arin et al., 2015; Ramey and Zubairy, 2018).

However, most studies rely on aggregate measures of private investment that combine residential and non-residential components, potentially obscuring the behavior of firms' capital expenditure decisions (Aschauer, 1989; Blanchard and Perotti, 2002; Galí et al., 2007; Arin et al., 2015). Since business investment is closely linked to productivity, capacity expansion, and long-run growth (Dixit and Pindyck, 1994;

Greenwood et al., 1997), isolating its response to fiscal shocks is particularly important—particularly during episodes of increased uncertainty, when firms may delay irreversible capital expenditures despite expansionary policy (Bloom, 2009; Bloom et al., 2018). Only a limited number of studies focus specifically on business investment. For OECD countries, Alesina et al. (2002) find that fiscal expansions reduce business investment not only through higher interest rates but also through rising public wages that increase real wages and compress firm profitability, while Perotti et al. (2007) documents heterogeneous responses across countries, with identical fiscal shocks generating negative investment effects in the United States and positive effects in Australia. Although these studies advance our understanding of how business investment responds to fiscal shocks, they rely on linear specifications that assume constant transmission across macroeconomic conditions.

An expanding body of research highlights nonlinear fiscal effects, showing that multipliers vary across business cycle phases, financial stress conditions, or monetary regimes (Auerbach and Gorodnichenko, 2012a; Baum et al., 2012; Ilzetzki et al., 2013; Arin et al., 2015; Caggiano et al., 2015; Riera-Crichton et al., 2015). However, this evidence largely focuses on output rather than investment. A growing literature also emphasizes the role of uncertainty in shaping investment dynamics. Real options theory shows that irreversibility increases the option value of waiting under heightened uncertainty (Bernanke, 1983; Pindyck, 1990), and empirical studies document that heightened uncertainty leads to lower levels of investment and hiring (Bloom et al., 2007; Bloom, 2009). Measures of economic policy uncertainty are similarly associated with declines in corporate investment (Bloom, 2009; Julio and Yook, 2012; Baker et al., 2016). Together, these findings highlight uncertainty as a key determinant of investment behavior, particularly in volatile policy environments.

These considerations motivate the central question of this paper: Does economic uncertainty alter the effect of government spending shocks on business investment? We examine whether investment responses differ between low- and high-uncertainty regimes using a threshold panel VAR (TVAR) estimated on quarterly data for seven OECD countries. The threshold framework allows fiscal transmission to vary systematically across regimes by permitting the VAR coefficients to shift once uncertainty crosses a critical level. This approach is appropriate for the analysis, as investment behaviour may change sharply when uncertainty rises enough to increase the option value of waiting and discourage immediate capital expenditure.

This paper contributes to the state-dependent fiscal policy literature by being the first to examine how government spending affects business investment across high- and low-uncertainty regimes. While prior nonlinear studies focus mainly on output or aggregate private investment, firms' capital expenditure decisions under varying uncertainty have received limited attention. The paper also provides the first cross-country evidence using a panel of advanced OECD economies.

The results suggest differences in investment responses across uncertainty regimes. In low-uncertainty periods, government spending shocks are associated with negative cumulative responses of business investment, consistent with crowding-out effects. In high-uncertainty episodes, the response is muted and statistically insignificant, suggesting that elevated uncertainty may weaken fiscal transmission rather than reverse it. The FEVD supports this pattern: investment is largely self-driven in stable environments, whereas output and interest rate shocks play a larger role under high uncertainty.

The remainder of the paper is organized as follows. Section 2 reviews the related literature. Section 3 outlines the theoretical framework guiding the analysis. Section 4 discusses the variables adopted in the study and analyzes the data. Section 5 describes the empirical strategy. Section 6 discusses the empirical results, while Section 7 provides robustness checks to assess the stability of the findings. Section 8 concludes the paper.

3.2 Literature Review

3.2.1 Fiscal Policy and State Dependence

A large empirical literature studies the macroeconomic effects of government spending shocks, particularly in advanced economies. Early SVAR contributions focus on identifying exogenous fiscal shocks and estimating their effects on output and consumption. One strand exploits institutional timing assumptions to isolate discretionary policy shocks (Blanchard and Perotti, 2002; Perotti, 2005; Pereira and Lopes, 2014; Caldara and Kamps, 2008), while others rely on sign and timing restrictions grounded in theory (Mountford and Uhlig, 2009; Arias et al., 2018; Canova and Pappa, 2011). A third approach adopts narrative identification strategies based on historical records (Romer and Romer, 2010; Ramey, 2011b; Ramey and Shapiro, 1998). Across methodologies, the common finding is that government spending raises output in the short run. However, these studies typically rely on linear models that impose constant transmission over time, implying multipliers that average across states and potentially mask nonlinearities.

Recognizing the limitations of linear models, a growing literature introduces state dependence into fiscal analysis. Using cross-country data, Ilzetzki et al. (2013) show that multipliers vary with structural characteristics such as exchange rate regimes, trade openness, and debt levels, while Ramey (2019) emphasize the role of identification, expectations, and macroeconomic conditions in shaping estimated effects. Within nonlinear frameworks, Auerbach and Gorodnichenko (2012b) find that U.S. government spending multipliers are larger during downturns, and Arin et al. (2015) report stronger spending effects in "bad times." Related cross-country evidence further documents that fiscal transmission differs across macroeconomic states and policy environments (Baum et al., 2012; Caggiano et al., 2015; Riera-Crichton et al., 2015). Collectively, these studies establish that fiscal transmission is state-dependent.

Despite these advances, the investment channel remains insufficiently examined.

State-dependent fiscal analyses predominantly center on output, leaving limited systematic evidence on how business investment responds to fiscal shocks across macroeconomic states, particularly in cross-country contexts. Even when investment is included in VAR systems, it is rarely the primary object of analysis and is typically measured as aggregate private investment rather than business (non-residential) capital formation. This distinction matters because theory suggests that investment may respond differently from output across economic states, given the roles of accelerator effects, adjustment costs, and uncertainty-related frictions.

3.2.2 Fiscal Policy, Investment and Uncertainty

As noted earlier, private investment has received comparatively less attention, despite its central importance for capital accumulation and long-run growth, and the existing empirical evidence remains mixed. For the United States, Blanchard and Perotti (2002) document crowding out of private investment following spending shocks, whereas Furceri and Sousa (2011) show that public investment can crowd in private capital when it complements infrastructure. Cross-country evidence reinforces this heterogeneity: Perotti (2005) finds that investment responses vary across OECD countries and over time, and Afonso and Jalles (2015) show that the composition of public spending affects outcomes: social security expenditures tend to reduce private investment, whereas health-related spending may have positive effects.

Importantly, most of this evidence relies on aggregate private investment, combining residential and non-residential components with distinct financing structures and sensitivities. Business investment, however, is especially responsive to expectations and plays a central role in productivity (Dixit and Pindyck, 1994; Greenwood et al., 1997). Studies focusing specifically on business investment remain relatively scarce. Alesina et al. (2002) show that fiscal expansions can reduce business investment through higher wages and compressed profits, while Perotti et al. (2007) and Mountford and Uhlig (2009) document heterogeneous and non-monotonic responses across countries. Tax-based analyses document a negative relationship between corporate taxes and business investment (Djankov et al., 2010; Mertens and Ravn, 2013).

A further limitation of this literature is its reliance on linear specifications that assume constant transmission across macroeconomic states. Yet theory suggests that investment responses may depend critically on the prevailing environment. Theory predicts that greater uncertainty increases the option value of waiting, leading firms to postpone hiring decisions and irreversible investment (Bernanke, 1983; Dixit and Pindyck, 1994). Consistent with this prediction, Bloom (2009) empirically shows that uncertainty shocks lead to sharp declines in output, employment, and investment, followed by rapid rebounds. Bloom et al. (2018) extended this work with the Economic Policy Uncertainty (EPU) index, documenting significant declines in investment and output following EPU shocks. Ludvigson et al. (2021) further distinguish financial from macroeconomic uncertainty, showing that economic fluctuations are primarily driven by the former.

While financial frictions may also influence investment dynamics, the focus on Economic Policy Uncertainty (EPU) is motivated by both theoretical and empirical considerations. The central question of this study is whether the transmission from fiscal policy to business investment depends on uncertainty surrounding the policy and macroeconomic environment. Real-options theory predicts a higher value of waiting with increased uncertainty, causing firms to delay irreversible investment even in the presence of demand stimulus. EPU directly captures this expectations-based channel. By contrast, financial frictions operate primarily through credit supply constraints and balance-sheet effects, representing a distinct mechanism. Focusing on EPU as the threshold variable therefore isolates the role of policy-related uncertainty in shaping regime-dependent investment responses.

Although nonlinear fiscal models have become more common, only a few studies treat uncertainty explicitly as a regime-defining variable (Alloza, 2017; Arčabić and Cover, 2016). Using a threshold VAR for the United States, Arčabić and Cover (2016) show that private investment responses to government spending differ across uncertainty regimes; however, they do not isolate business investment from broader measures of private domestic investment, limiting conclusions about firms' capital expenditure decisions. Likewise, Berg (2019) analyzes fiscal effectiveness under varying levels of business uncertainty using German firm-level survey data, but their focus remains on multiplier effects rather than on the macroeconomic transmission to business investment. Both studies rely on single-country evidence, which limits the external validity of their findings. Although advanced OECD economies share broadly similar institutional and macroeconomic structures, the timing and duration of low- and high-uncertainty regimes may differ across countries because of differences in policy frameworks, economic conditions, and exposure to uncertainty shocks. Evidence from a single country therefore cannot establish whether regime-dependent fiscal effects represent a general pattern across advanced economies. A panel framework addresses this limitation by allowing each country to transition between uncertainty regimes at different points in time while pooling information across countries to estimate common regime-specific impulse responses.

3.2.3 Threshold Panel VAR

A prominent approach in state-dependent fiscal analysis employs regime-switching models. Auerbach and Gorodnichenko (2012b) use a smooth transition VAR (STVAR) and show that U.S. government spending multipliers are larger in recessions than in expansions (see also Granger et al., 1993; Arin et al., 2015; Caggiano et al., 2015). Local projection methods (Jordà, 2005; Stock and Watson, 2007) have likewise been used to estimate state-dependent multipliers while relaxing the dynamic restrictions imposed by standard SVARs (Auerbach and Gorodnichenko, 2012a; Auerbach and Gorodnichenko, 2013; Riera-Crichton et al., 2015). A distinct but related framework is the Threshold VAR (TVAR), which allows regime changes to occur once an observable variable crosses a critical value, generating discrete shifts in system dynamics rather than smooth transitions. This structure is particularly appropriate when fiscal transmission may change abruptly across economic conditions (Baum et al., 2012; Fazzari et al., 2015; Woldu and Szakálné Kanó, 2023).

Multivariate TVARs have been widely applied to study macroeconomic and financial asymmetries (Balke and Fomby, 1997; Calza and Sousa, 2005; Li and St-Amant, 2010). Extending this framework to panel data, the Panel Threshold VAR (PTVAR) permits regime dependence to be analyzed in multi-country settings (Hansen, 1999; Fazzari et al., 2015; Soave, 2023). Recent studies apply panel threshold models to fiscal and macro-financial questions, showing that fiscal transmission varies across financial stress regimes (Afonso et al., 2018) and broader macroeconomic environments (Woldu and Szakálné Kanó, 2023; Davidson and Moccero, 2024). These applications demonstrate the usefulness of threshold methods for identifying regime-dependent policy effects in cross-country data.

An alternative way to model state dependence is the interacted VAR framework, in which impulse responses vary continuously with the level of a state variable, such as economic policy uncertainty. In this approach, VAR coefficients are interacted with the state variable so that responses change smoothly as uncertainty rises. This method has been implemented primarily in single-country settings with long time series, such as Berg (2019) for Germany, where sufficient observations allow reliable estimation of the additional interaction parameters. Extending interacted VARs to a panel context is more challenging. One option is to impose homogeneous interaction effects across countries, which risks masking cross-country heterogeneity in fiscal transmission. Alternatively, estimating separate interacted VARs for each country reduces efficiency and limits comparability, especially when individual time series are relatively short. These trade-offs are particularly relevant in panels of moderate length, such as the present sample.

Given these considerations, this study adopts a Panel Threshold VAR framework. The PTVAR allows fiscal transmission to differ discretely across low- and high-uncertainty regimes while pooling information across seven OECD countries to improve estimation precision. With only 91 quarterly observations per country, estimating a separate nonlinear model for each country would not be reliable. By exploiting the cross-sectional dimension, the model provides a tractable and transparent framework for identifying regime-dependent impulse responses across advanced economies.

3.3 Theoretical Framework

The effect of fiscal policy on private business investment remains contested. Neoclassical models emphasize crowding out: debt-financed government spending raises real interest rates or future tax burdens, lowering the expected after-tax return on capital and reducing private investment (Barro, 1974; Baxter and King, 1993). Empirical evidence often supports this view. For example, Blanchard and Perotti (2002) show that U.S. government spending shocks increase output and consumption but reduce private investment.

In contrast, Keynesian models allow for crowding-in effects when the economy operates below potential and prices are sticky. In such settings, government spending stimulates aggregate demand, raises firms' expected sales, and may increase investment through accelerator effects (Woodford, 2011; Christiano et al., 2011a). Complementarities between public and private capital can further enhance investment incentives, particularly when infrastructure raises the marginal productivity of private inputs (Aschauer, 1989; Bhatta and Drennan, 2003). However, Keynesian theory does not yield a uniform prediction: investment rises if demand effects dominate, but may fall if higher interest rates or financing costs offset them (Blanchard and Perotti, 2002; Alesina et al., 2002; Galí et al., 2007; Ramey, 2011b). Thus, theoretical predictions are inherently ambiguous.

Both frameworks highlight the importance of expectations and uncertainty. When investment is partially irreversible, high uncertainty raises the option value of waiting, causing a delay in capital expenditures (Bernanke, 1983; Dixit and Pindyck, 1994). Empirical evidence shows that uncertainty shocks depress investment and weaken the effectiveness of expansionary policies (Bloom et al., 2007; Bloom, 2009). Under elevated uncertainty, firms may perceive fiscal expansions as temporary or insufficient to compensate for heightened risk, while rising risk premia can further dampen responsiveness to demand stimuli. Conversely, credible fiscal interventions could reduce uncertainty and amplify investment. Whether fiscal transmission strengthens or weakens in high-uncertainty environments is therefore ultimately an empirical question.

The theoretical justification for regime dependence arises from real options models of irreversible investment. In these frameworks, firms face non-convex adjustment costs and "inaction regions," investing only when expected returns exceed a critical threshold (Bernanke, 1983; Dixit and Pindyck, 1994). As uncertainty rises, the option value of waiting increases nonlinearly, widening the inaction region and raising the hurdle rate for investment. Once uncertainty exceeds a critical level, firms' responses to demand or policy shocks may change discretely rather than smoothly. Fiscal expansions that stimulate investment in low-uncertainty environments may have limited effects when uncertainty is sufficiently high. This discontinuity provides a theoretical foundation for modeling fiscal transmission within a threshold framework that distinguishes between low- and high-uncertainty regimes.

3.4 Data Analysis

3.4.1 Variable Selection and Data Transformation

The empirical analysis incorporates a set of macroeconomic and policy-relevant variables: business investment, real GDP, government spending, inflation, a short-term interest rate, and economic policy uncertainty (EPU). Table 3.1 summarizes the definition, data source, and transformation applied to each series.

Business investment, proxied by private non-residential gross fixed capital formation, is the primary outcome variable, as it captures firms' capital accumulation decisions[1]. Government spending is included to identify fiscal shocks following standard VAR-based approaches. We adopt government consumption expenditure, which captures discretionary fiscal interventions and limits feedback from automatic stabilizers like

[1]Under the System of National Accounts, Gross fixed capital formation (GFCF) records investment in fixed assets, including buildings and machinery and Private non-residential GFCF excludes residential investment. Focusing on this measure isolates firms' fixed capital expenditure decisions, which are expected to respond more directly to fiscal policy and uncertainty than broader measures of aggregate investment.

Table 3.1: Variables, Data Sources, and Transformations

Variable	Definition / Construction	Source	Transformation
Business Investment	Private non-residential GFCF	OECD Economic Outlook	Real terms, quarterly growth rates
GDP	Quarterly national accounts GDP (expenditure approach)	OECD Economic Outlook	Real terms, quarterly growth rates
Government Spending	Final consumption expenditure of general government	OECD Economic Outlook	Real terms, quarterly growth rates
Inflation	GDP deflator	OECD Quarterly GDP and Components	Quarterly growth rates
Short-term Interest Rate	Short-term nominal policy rate	OECD Key Short-Term Economic Indicators	First-difference
Economic Policy Uncertainty (EPU)	Log EPU index, cyclical component	Baker et al. (2016) obtained from policyuncertainty.com	Lagged one quarter, cyclical component of log index

income taxes and transfer programs (Blanchard and Perotti, 2002; Ilzetzki et al., 2013).

GDP captures aggregate demand conditions and the accelerator mechanism linking output fluctuations to investment. Inflation is included to account for price dynamics, while short-term interest rates are included to account for the monetary policy stance and the financing-cost channel. Economic policy uncertainty (EPU) is proxied by the cyclical component of the log EPU index (Baker et al., 2016), lagged by one quarter.

The short-term interest rate is used as the baseline measure of monetary conditions because it is consistently available across all countries over the sample period. Although alternative indicators, such as long-term government bond yields, corporate bond spreads, or shadow policy rates, may better capture firms' financing conditions in some circumstances, comparable quarterly data for these measures are not available across all seven countries for the full sample. As a result, alternative financial indicators could not be used without substantially reducing the sample size.

The sample consists of seven advanced OECD economies: Australia, Canada, France, Germany, Korea, the United Kingdom, and the United States. These countries were selected based on the availability of consistent quarterly data on economic policy uncertainty (EPU) and private non-residential gross fixed capital formation over a sufficiently long time span. Many OECD members do not provide quarterly business investment series disaggregated from residential and public components, or lack a comparable EPU index, which would prevent the construction of a balanced panel. Focusing on advanced OECD economies ensures institutional comparability, similar fiscal and monetary frameworks, and well-developed financial systems, which supports the assumption of broadly comparable transmission mechanisms across countries.

3.4.2 Data Inspection

Figure 3.1 presents the macroeconomic series for the 7 OECD countries in our sample, revealing notable patterns. The EPU panel shows that the cyclical component of policy uncertainty peaks precisely when economic growth rates are lowest. These deviations from the trend are highly volatile, with the most extreme synchronized spikes coinciding with 2002-2003 geopolitical uncertainty and the 2008–2009 Global Financial Crisis.

Business Investment exhibits the most pronounced volatility among the real variables. The growth rate of investment swings between 0.1% to -0.2%, significantly exceeding the range of GDP growth. The deepest contractions in investment (e.g., 2009) align closely with spikes in the lagged EPU cycle, suggesting a potentially negative relationship in which heightened policy uncertainty follows economic downturns and drags on capital accumulation.

The Government Spending panel reveals cross-country variation in observed fiscal growth rates, reflecting the discretionary and country-specific implementation of fiscal policy. Empirically, the uncertainty series exhibits episodic spikes rather than gradual drift, and these spikes coincide with sharp contractions in investment. This

Figure 3.1: Real GDP, Government Consumption, and Business Investment are expressed as quarterly growth rates (log-differences); Economic Policy Uncertainty (EPU) is measured as the cyclical component of the log-index, lagged by one quarter.

motivates treating high-uncertainty periods as a distinct state. A threshold VAR provides a parsimonious representation of such state dependence by allowing the propagation of shocks to differ across low- and high-uncertainty regimes.

3.4.3 Pre-estimation Tests

Unit Root Tests

To assess the time-series properties of the variables, we first conducted country-specific Augmented Dickey–Fuller (ADF) tests including an intercept and four lags. The null hypothesis of a unit root was rejected for all variables across all countries at conventional significance levels. These results indicate that output growth, government spending growth, investment growth, inflation, interest rates (in first-difference), and the lagged economic policy uncertainty index are stationary in levels for each country. The consistency of these findings across all seven economies provides strong evidence that the series do not exhibit stochastic trends over the sample period.

To complement the individual tests, we implemented panel unit-root tests using the Im–Pesaran–Shin (IPS) and Fisher-type ADF procedures. Both approaches strongly reject the null hypothesis that all panels contain unit roots, with p-values close to zero for all variables. The panel results confirm the country-level evidence and further strengthen the conclusion that the variables are stationary in their transformed form. Taken together, the individual and panel unit-root tests provide a robust basis for proceeding with the estimation of the panel VAR models. Tables containing results of the unit root tests can be found in Appendix C.1.1

Lag Order Selection

Table C.3 reports the panel VAR lag selection criteria based on the modified Bayesian (MBIC), Akaike (MAIC), and Hannan–Quinn (MQIC) information criteria (Andrews and Lu, 2001; Abrigo and Love, 2016). All three criteria unanimously support a first-order VAR specification. Accordingly, we adopt a one-lag specification for the panel VAR. This lag length is sufficient to capture the short-run interactions among business investment, output, government spending, inflation, and interest rates, while avoiding the risk of over-parameterization that arises in higher-order panel VARs with a relatively small cross-section.

Table 3.2: Summary Statistics

Variable	**Obs**	**Mean**	**Std. Dev.**	**Min**	**Max**
Spending	637	0.005974	0.008480	-0.028677	0.037766
Output	637	0.006033	0.007888	-0.069502	0.042468
Investment	637	0.007165	0.035779	-0.315202	0.382603
Inflation	637	0.004514	0.008386	-0.033523	0.045120
Interest rates	637	-0.053839	0.600703	-6.650000	7.120000
EPU_lag	637	-0.000630	0.306765	-0.950043	1.026894

Note: Summary statistics are computed for the full panel of 7 countries across the sample period (1997Q2–2019Q4).

3.5 Empirical Strategy

To examine the dynamic relationship between economic uncertainty and macroeconomic aggregates, we employ a Panel Vector Autoregression (PVAR) framework following Abrigo and Love (2016). The PVAR assumes all the variables are endogenous, while controlling for unobserved country-specific heterogeneity. The analysis proceeds in two steps: we first estimate a linear PVAR to establish baseline dynamics over the full sample (1997Q2–2019Q4), and then estimate a threshold PVAR to assess whether these dynamics differ across low- and high-uncertainty regimes.

3.5.1 Linear PVAR

We begin by estimating a linear panel VAR model on the full sample. The reduced-form system is given by

$$y_{i,t} = B(L)y_{i,t-1} + \omega_i + u_{i,t}, \tag{3.1}$$

where $y_{i,t}$ is a vector of endogenous variables for country i at time t, comprising *Government Spending*, *Output*, *Investment*, *Inflation*, and *Interest Rates*. The matrix $B(L)$ is a lag polynomial that includes lagged values of $y_{i,t}$ up to the selected order. The term ω_i captures time-invariant country-specific effects, while $u_{i,t}$ denotes a vector of reduced-form innovations.

GMM Estimation of the Panel VAR

Estimation follows the generalized method of moments (GMM) framework for panel vector autoregressions developed by Holtz-Eakin et al. (1988). Their contribution extends dynamic panel estimation to systems of endogenous variables, allowing for the estimation of VAR models in panel settings while accounting for unobserved heterogeneity.

The inclusion of lagged dependent variables renders standard pooled OLS and fixed-effects estimators inconsistent, as the regressors are mechanically correlated with the fixed effects (Nickell, 1981). Holtz-Eakin et al. (1988) address this problem by first eliminating fixed effects through a linear transformation and then exploiting internal instruments derived from deeper lags of the endogenous variables. Assuming that the idiosyncratic errors are not serially correlated, deeper lags of the endogenous variables are orthogonal to contemporaneous shocks, making them suitable internal instruments. The GMM estimator imposes these orthogonality conditions to identify the parameters, yielding consistent estimates in dynamic panel models.

Estimation is conducted in Stata using the panel VAR implementation of Abrigo and Love (2016), based on the GMM framework developed by Holtz-Eakin et al. (1988). Country-specific fixed effects are removed using forward orthogonal deviations (FOD) rather than first differencing. FOD preserves orthogonality between transformed errors and lagged instruments while avoiding the loss of initial observations, improving efficiency in panels with moderate time dimensions. All variables in the VAR are treated as endogenous, allowing consistent estimation of dynamic interdependencies across macroeconomic variables.

3.5.2 Panel Threshold VAR

To allow for state-dependent dynamics in the transmission of fiscal shocks, we extend the linear panel VAR to a Panel Threshold VAR (PTVAR) framework following the specifications in Balke and Fomby (1997), Calza and Sousa (2005), Afonso et al. (2018), Woldu and Szakálné Kanó (2023) and Davidson and Moccero (2024). We partition the sample based on an exogenous threshold variable—Economic Policy Uncertainty (EPU). The PTVAR model is specified in reduced form as:

$$y_{i,t} = B_1(L)y_{i,t-1} + B_2(L)y_{i,t-1} \cdot I(\text{EPU}_{i,t-1} > \gamma) + \omega_i + u_{i,t}, \tag{3.2}$$

where $y_{i,t} = \left[\text{Government Spending}_{i,t}, \text{Output}_{i,t}, \text{Investment}_{i,t}, \text{Inflation}_{i,t}, \text{Interest Rate}_{i,t}\right]'$ is a five-dimensional vector of endogenous variables for country i at time t. The lag polynomial $B_i(L)$ captures the dynamic effects of past values of the endogenous variables. The indicator function $I(\text{EPU}_{i,t-1} > \gamma)$ defines the high-uncertainty regime: it takes the value 1 if the lagged cyclical component of the Economic Policy Uncertainty (EPU) index exceeds the threshold γ, and 0 otherwise. This specification allows the dynamic relationships among the variables to differ depending on the level of policy uncertainty. The term ω_i captures time-invariant country-specific effects, while $u_{i,t}$ represents the vector of reduced-form innovations with covariance matrix Σ_u.

3.5.3 Identification Strategy

Structural shocks are identified with a recursive identification scheme using the Cholesky factor of the variance–covariance matrix. This approach imposes zero contemporaneous restrictions, assuming that variables ordered earlier can affect those ordered later within the same quarter, but not vice versa.

The endogenous variables are ordered as

$$Y_{it} = [\text{Government Spending}, \text{Output}, \text{Investment}, \text{Inflation}, \text{Interest Rate}].$$

Following Blanchard and Perotti (2002) and Perotti et al. (2007), government spending is placed first, reflecting implementation lags that prevent contemporaneous responses to economic shocks. Output and investment are ordered next, allowing real activity to respond immediately to fiscal shocks but only gradually to nominal disturbances. Inflation follows, consistent with price rigidities, and then the short-term interest rate, permitting the central bank to react contemporaneously to all other variables (Christiano et al., 2005).

3.5.4 Economic Policy Uncertainty and Threshold Determination

Economic policy uncertainty is measured using the country-specific Economic Policy Uncertainty (EPU) indices developed by Baker et al. (2016). The index is constructed from newspaper coverage frequency, identifying articles that reference the economy, uncertainty, and policy. The EPU index is constructed at the monthly frequency and normalized to have a mean of 100 over a base period. For consistency with the macroeconomic variables, we aggregate the monthly series to quarterly frequency by taking simple averages within each quarter. In the empirical analysis, the quarterly EPU series is first expressed in logarithms, and the Hodrick–Prescott filter is then applied to the logged series to extract its cyclical component. This specification is consistent with prior empirical work that applies logarithmic transformation and HP filtering to uncertainty measures (Bloom, 2009; Caggiano et al., 2014; Sengupta et al., 2025; Liu et al., 2023). The resulting cyclical log EPU measure is lagged by one quarter to account for information transmission delays and mitigate potential endogeneity.

The EPU index is HP-filtered to remove its low-frequency trend so that the threshold variable reflects deviations from the underlying level of policy uncertainty rather than persistent movements in the index. This distinction is important because long-run increases in the EPU series may partly reflect changes in the construction or measurement of the index, such as evolving media coverage, rather than sustained increases in economically meaningful uncertainty. By focusing on the cyclical component, the analysis identifies periods of unusually high or low policy uncertainty relative to the prevailing trend, thereby reducing the likelihood that the estimated regimes are driven by gradual shifts in the level of the index. Because periods of elevated policy uncertainty often coincide with recessions, financial crises, or other adverse macroeconomic conditions, the estimated regime differences should be interpreted as reflecting fiscal transmission under periods characterized by relatively high or low cyclical policy uncertainty, rather than isolating the effect of uncertainty from all other concurrent economic developments.

This study employs the Economic Policy Uncertainty (EPU) index rather than a fiscal policy uncertainty (FPU) measure. Although FPU indices exist (Hollmayr and Matthes, 2015; Popiel, 2020; Anzuini et al., 2020; Wen et al., 2022), they are typically country-specific, constructed using different methodologies, and not consistently available across OECD economies. Empirical results are often sensitive to the chosen FPU measure, and constructing a harmonized cross-country index would introduce substantial methodological challenges beyond the scope of this study. Moreover, state-dependent fiscal analyses generally rely on broad measures of economic or policy uncertainty rather than fiscal-specific indicators (Alloza, 2017; Berg, 2019). The EPU index provides a transparent, widely used, and consistently available proxy for uncertainty across advanced economies.

Rather than adopting the endogenous threshold estimator of Hansen (1999) as implemented by Wang (2015), we define regimes using country-specific percentiles of the EPU distribution. The Hansen–Wang approach imposes a common global cutoff across countries and can yield wide confidence intervals in finite samples, potentially producing economically implausible classifications. A single threshold assumes identical uncertainty "tipping points" across heterogeneous economies, despite substantial cross-country differences in EPU levels and volatility. We therefore classify high-uncertainty periods as quarters above the 85th percentile of each country's cyclical EPU distribution, allowing crisis states to be defined relative to domestic norms. As shown in Table C.4, the resulting cutoffs vary across countries, reflecting heterogeneity in uncertainty dynamics.

Using percentile-based thresholds is consistent with established practice in nonlinear macroeconomic analysis when the objective is to isolate extreme states. For example, Bloom (2009) and Caggiano et al. (2017) identify high-uncertainty episodes using a dummy that equals one when detrended volatility exceeds 1.65 standard deviations above its mean, thereby focusing on large exogenous jumps. Similarly, Caggiano et al. (2015) calibrate their nonlinear fiscal model so that roughly 15% of observations are classified as recessions, effectively defining regimes using tail realizations of the transition variable. Consistent with this approach, defining high-uncertainty periods as the top 15% of each country's EPU distribution ensures that the regime captures economically meaningful spikes—such as the 2001–2002 global instability, the 2008–2009 Global Financial Crisis, and the Eurozone Sovereign Debt Crisis—rather than normal cyclical fluctuations. The IRFs presented in Appendix C.3 using the 80th percentile confirm that the findings are not sensitive to the choice of cutoff.

3.6 Results

3.6.1 Impulse Responses of Business Investment to Government Spending

Figure 3.2 reports the impulse response function (IRF) of business investment to an identified government spending shock obtained from the linear panel VAR model, the high-uncertainty and low-uncertainty states using the PTVAR model.

The IRF from the linear model reveals a non-monotonic short-run response of investment to a government spending shock, though not statistically significant. On impact, a 1% increase in government spending growth leads to a decline in the growth rate of business investment by 0.002% up to the first period after the shock. This negative initial response is consistent with crowding-out, with evidence that in the short run, public sector expansion competes directly with the private sector for scarce resources. If the economy is near capacity, higher public spending can bid up wages and input prices, raising production costs and further compressing expected returns on private investment. Under these conditions, the private sector reacts by scaling back investment growth even though overall demand is stronger. At horizon two, the response turns positive, before gradually converging back toward zero at longer horizons. The non-monotonic investment responses observed are consistent with evidence from VAR-based studies showing that private investment responses to government spending shocks sometimes display irregular patterns or oscillatory dynamics rather than smooth adjustment paths (Ramey, 2011b; Perotti, 2005; Ilzetzki et al., 2013).

The cumulative impulse response remains consistently negative across all horizons, implying that, when investment responses are aggregated over time, government spending shocks lead to a small but persistent decline in business investment. The cumulative response stabilizes relatively quickly, suggesting that most of the transmission occurs in the short term. This evidence of crowding-out aligns with the findings of other studies such as Blanchard and Perotti (2002), Alesina et al. (2002) and Afonso and Jalles (2015).

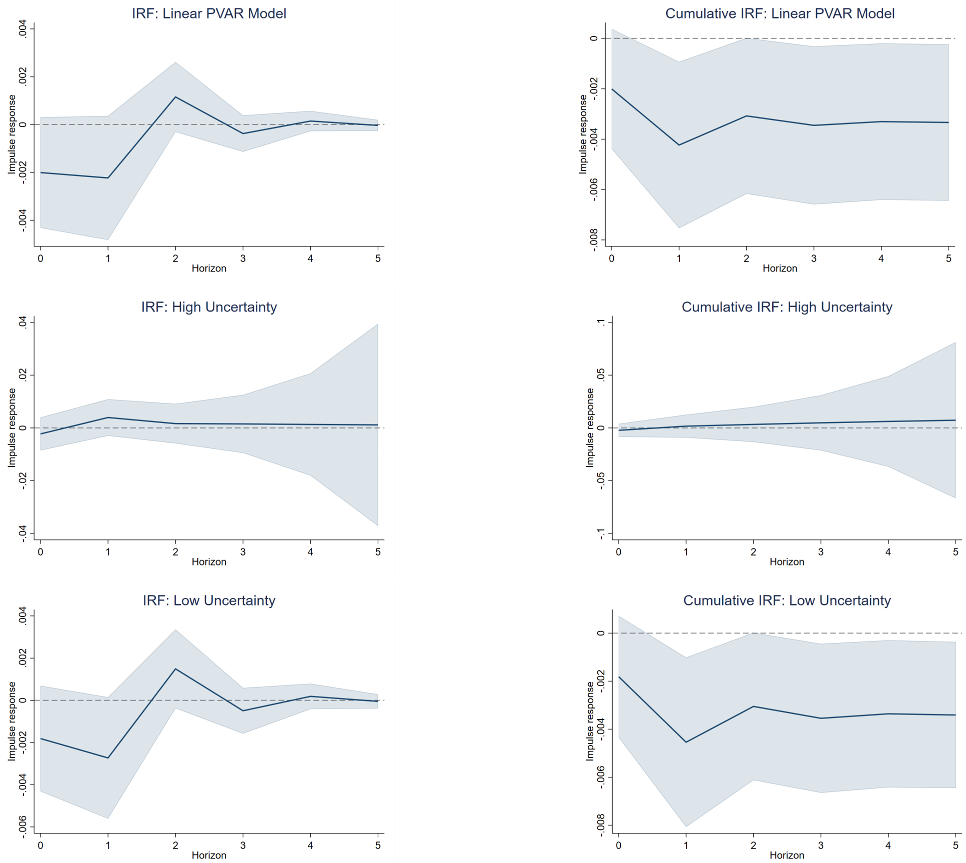


Figure 3.2: IRFs and Cumulative IRFs of business investment growth to a 1% shock in Government Spending growth, using linear and threshold VAR models at 95% confidence interval

When divided into high and low uncertainty regimes, we find some evidence of state-dependence in the response of business investment to a government spending shock. The IRF and cumulative IRF under low uncertainty closely resemble those from the linear PVAR model. This suggests that during normal, low-uncertainty periods, firms behave in line with standard classical predictions: an increase in government spending leads to a decline in private capital accumulation. When fiscal expansion is financed through debt, the increased sovereign demand for credit competes with private sector borrowing. This shift in the loanable funds market exerts upward pressure on real interest rates, thereby raising the cost of capital for firms. As borrowing costs rise, the net present value of marginal investment projects declines, prompting firms to postpone or cancel planned expenditures. At the same time, increased public sector demand may absorb scarce productive resources and drive up input prices, raising production costs for private firms. Private investment is therefore displaced both financially–through higher borrowing costs–and physically, through the reallocation of real resources toward the public sector.

The impulse response functions under high uncertainty reveal a different pattern from the low-uncertainty case. The standard IRF indicates a negative impact response of business investment (-0.002%) at horizon 0, suggesting an initial crowding-out effect. However, this negative response is short-lived. By horizon 1, investment turns slightly positive, before gradually converging back toward zero over subsequent periods. The confidence bands are relatively wide and encompass zero throughout the forecast window, indicating that the response is statistically imprecise. The cumulative IRF reinforces this interpretation. Investment displays a gradual but very modest positive accumulation over time. Nevertheless, the confidence intervals widen substantially as the horizon increases and remain centered around zero, implying that the cumulative effect is economically small and statistically indistinguishable from no effect.

The findings imply that high uncertainty may dampen the responsiveness of business investment to government spending shocks. Unlike in low-uncertainty periods—where classical crowding-out through higher interest rates and resource reallocation is more apparent—the high-uncertainty regime is characterized by muted and statistically weak responses. This pattern is consistent with the findings of Bernanke (1983), Dixit and Pindyck (1994), and Bloom (2009), who argue that heightened uncertainty can lead firms to postpone irreversible investment decisions. However, these results should be interpreted with appropriate caution because the high-uncertainty regime is estimated from a relatively small number of observations. Since only the upper 15% of each country's EPU distribution is classified as high uncertainty, the resulting estimates are less precise, particularly at longer horizons, as reflected in the wider confidence intervals. Table C.5 shows that observations coinciding with the

Global Financial Crisis account for only 16.5% of the high-uncertainty regime. The remaining observations are distributed across other periods of elevated uncertainty, including the early-2000s global slowdown, the Eurozone sovereign debt crisis, and country-specific policy uncertainty episodes.

3.6.2 IRFs to Shocks in GDP and Interest rates

Figures 3.3 and 3.4 present the impulse response functions of business investment to shocks in GDP and interest rates. Although the primary focus of this study is the effect of fiscal shocks, examining these responses provides insight into how uncertainty shapes investment dynamics through demand and financial channels.

Figures 3.3 show that the response of business investment to a positive GDP shock differs across uncertainty regimes both in magnitude and persistence. Under high uncertainty, the growth rate of business investment rises by approximately 0.017% on impact, remains positive at 0.015% in the first quarter, and gradually declines to about 0.008% by horizon five. The cumulative response increases steadily from roughly 0.017% on impact to 0.073% by the fifth quarter, indicating a sustained demand-driven expansion in capital expenditure. However, the confidence bands widen substantially after the first two quarters, and statistical significance fades beyond the short run, reflecting increased volatility in investment behaviour. In contrast, during low-uncertainty periods, the response is smaller and short-lived: investment increases by about 0.009% on impact and 0.004% in the first quarter, but converges to zero from horizon two onward. The cumulative effect stabilizes around 0.012% – 0.013% and remains relatively flat thereafter, with tighter confidence intervals. These results suggest that the demand channel operates in both regimes, but under high uncertainty, the response is larger in point estimates yet more dispersed and less precisely estimated over time.

The response of business investment to a monetary tightening also differs across uncertainty regimes. Under high uncertainty, investment falls by about –0.010% on impact and –0.0117% in the first quarter, with confidence bands at horizon 1 entirely

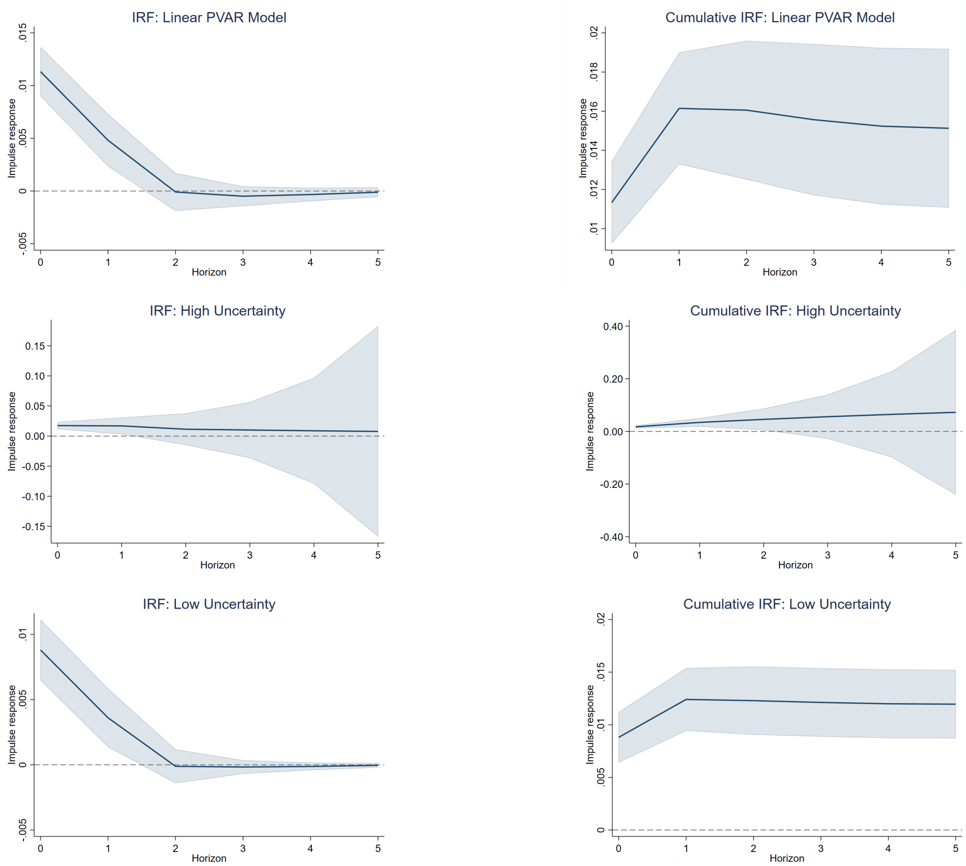


Figure 3.3: IRFs and Cumulative IRFs of business investment growth to a one–percentage-point innovation in GDP growth, using linear and threshold VAR models at 95% confidence interval

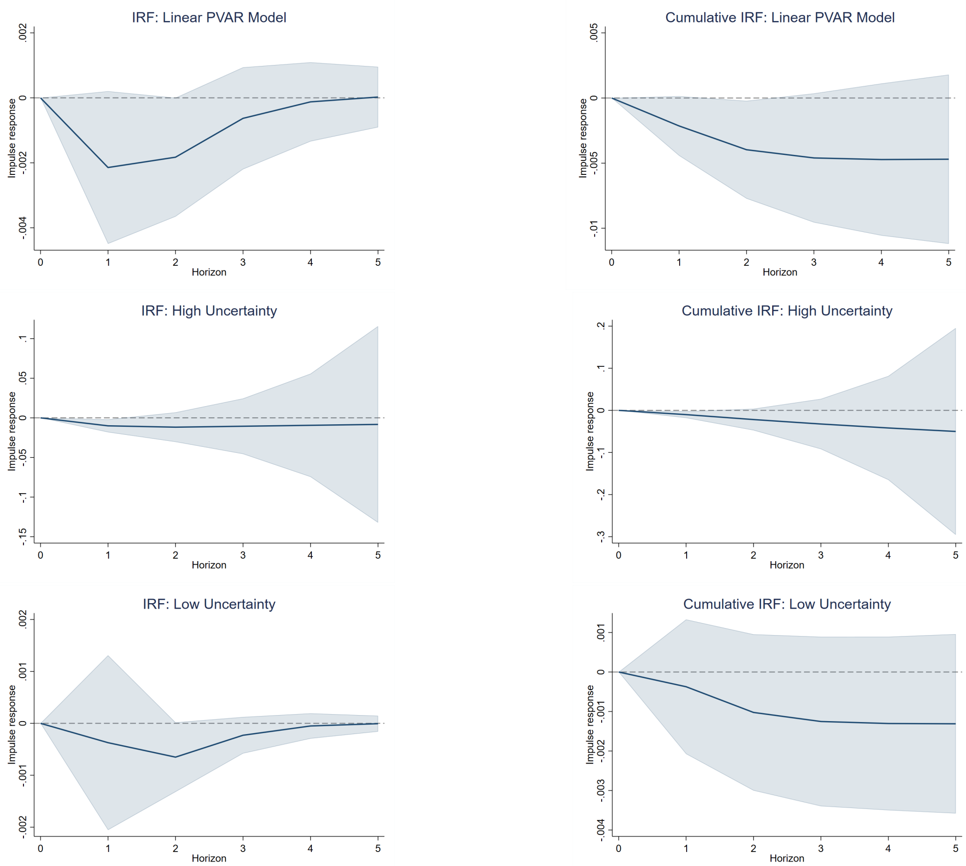


Figure 3.4: IRFs and Cumulative IRFs of business investment growth to one–percentage-point increase in the change in the policy rate, using linear and threshold VAR models at 95% confidence interval

below zero (–0.018% to –0.002%), indicating short-run statistical significance. The cumulative decline deepens steadily to approximately –0.050% by horizon five, pointing to a persistent contraction in capital expenditure, although significance dissipates as confidence intervals widen. In contrast, during low-uncertainty periods, the effect is negligible: investment declines by only –0.00037% on impact, remains near zero thereafter, and the cumulative response stabilizes around –0.0013%, with confidence intervals including zero at all horizons. These results indicate that the financial channel becomes substantially stronger when uncertainty is high, while interest rate shocks have minimal influence on investment in stable environments.

3.6.3 Forecast Error Variance Decomposition (FEVD)

Figure 3.5 presents the Forecast Error Variance Decomposition (FEVD) of business investment across the three estimated specifications: the linear benchmark, the low-uncertainty regime, and the high-uncertainty regime. While the impulse response functions presented in the previous sections describe the direction and magnitude of investment responses to structural shocks, the FEVD provides complementary information by quantifying the proportion of the forecast error variance in business investment attributable to each identified shock over the forecast horizon. Comparing the low- and high-uncertainty regimes reveals a marked change in the relative importance of the different shocks in explaining fluctuations in business investment.

In the low-uncertainty regime, the forecast error variance of business investment is dominated by its own structural innovation, which accounts for close to 90% of the variance over the forecast horizon. Output, government spending, inflation, and interest rate shocks each explain only a small proportion of the variation in investment forecast errors. This indicates that, within the model, fluctuations in business investment during periods of low uncertainty are explained primarily by the investment-specific structural shock rather than by the identified macroeconomic shocks. It should be noted, however, that this residual innovation may capture a range of factors not explicitly modelled, including firm-specific disturbances, expectations, financial frictions, technology shocks, or measurement error.

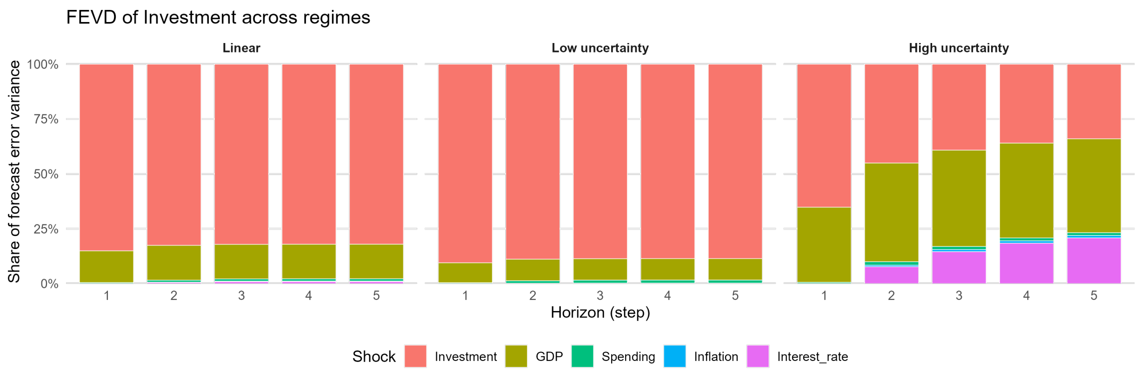


Figure 3.5: FEVD of Investment with the linear model, high-uncertainty regime and low-uncertainty regime, for Horizons 1-5.

In the high-uncertainty regime, the composition of the forecast error variance changes substantially. The contribution of the investment-specific shock declines markedly—from approximately 65% at the one-quarter horizon to around 34% by the fifth quarter—while the contributions of macroeconomic shocks increase. In particular, output shocks account for an increasingly large share of the forecast error variance over time, and the contribution of interest rate shocks also rises relative to the low-uncertainty regime. The FEVD further shows that the importance of output and interest rate shocks increases with the forecast horizon in the high-uncertainty regime. This finding complements the impulse response analysis by indicating that these shocks account for an increasing share of the uncertainty surrounding future investment as the horizon lengthens.

3.7 Robustness Checks

3.7.1 Adopting the Global EPU Index

As an additional robustness exercise, the threshold variable is replaced with the hp-filtered global EPU index of Baker et al. (2016), so that all countries are assigned to the same uncertainty regime in a given quarter. This alternative specification addresses the concern that country-specific thresholds may complicate the interpretation of a pooled threshold panel VAR by allowing countries to occupy different regimes within the same calendar quarter. The results (see impulse response functions of business investment to a government spending shock in Appendix C.1) are qualitatively similar to those obtained using country-specific EPU indices. Under high uncertainty, the impulse response of private investment to a government spending shock is negative on impact, oscillates before converging to zero by the third period, and the cumulative response remains close to zero throughout the horizon. Under low uncertainty, the impulse response is likewise negative on impact and oscillates before converging by approximately the fifth period, while the cumulative response remains negative. These findings indicate that the main results are robust to both a country-specific regime classification, which captures domestic episodes of uncertainty, and a common global regime classification, which imposes a more restrictive but econometrically cleaner pooling structure.

3.7.2 Using the Raw EPU Index

As a robustness exercise, the threshold panel VAR is re-estimated using the unfiltered EPU index to define the uncertainty regimes. This exercise responds to concerns regarding the use of the HP filter and assesses whether the main findings depend on isolating the cyclical component of economic policy uncertainty. Inspection of the resulting regime classification shows that observations in the later years of the sample are disproportionately assigned to the high-uncertainty regime across all countries, reflecting the upward trend in the EPU index rather than pronounced deviations from trend. Consequently, periods characterized by relatively favorable macroeconomic conditions are frequently classified as episodes of high uncertainty. This pattern suggests that the unfiltered index captures persistent changes in the level of measured policy uncertainty in addition to transitory uncertainty shocks.

Consistent with this alternative regime classification, the estimated impulse responses (see Appendix C.2) differ from those of the baseline specification. Under the low-uncertainty regime, both the impulse responses and cumulative responses of business investment are similar in direction and magnitude to the baseline results. In contrast, the responses under the high-uncertainty regime resemble those estimated for the low-uncertainty regime more closely: business investment declines on impact and in the first period following a government spending shock, while the cumulative response remains negative throughout the forecast horizon. The reduced distinction between the two regimes suggests that defining high uncertainty using the unfiltered EPU index weakens the model's ability to distinguish periods of unusually elevated policy uncertainty from periods characterized by a persistently higher level of measured uncertainty. These findings support the use of the HP-filtered EPU index in the baseline specification, as it identifies episodes of unusually high policy uncertainty relative to each country's underlying trend rather than classifying observations primarily on the basis of the long-run increase in measured policy uncertainty.

3.8 Conclusion

This study examines the state-dependent effects of government spending shocks on business investment across seven OECD countries using a threshold panel VAR framework that distinguishes between high- and low-uncertainty regimes. The results suggest that the response of business investment differs across uncertainty regimes. During low-uncertainty periods, government spending shocks are associated with a decline in business investment, with cumulative impulse responses remaining negative, consistent with crowding-out effects. This pattern is consistent with classical explanations in which deficit-financed fiscal expansions increase borrowing costs and reallocate resources toward the public sector.

In contrast, during high-uncertainty episodes, evidence of crowding out becomes weaker and statistically insignificant. Investment responds only modestly, with cumulative effects remaining close to zero. While these muted responses are consistent with theories that heightened uncertainty delays irreversible investment, the lack of statistical significance warrants cautious interpretation. The FEVD results provide additional insight by showing that, under high uncertainty, investment becomes more sensitive to output and interest rate shocks than in low-uncertainty periods, suggesting that the factors influencing investment dynamics differ across uncertainty regimes.

While the analysis provides new cross-country evidence on state-dependent fiscal transmission, it is subject to certain data constraints. The availability of the Economic Policy Uncertainty (EPU) index for most OECD countries begins in 1997, limiting the time dimension of the sample. In addition, consistent data on private non-residential gross fixed capital formation are not available for all OECD economies, restricting the analysis to a subset of advanced countries. Future research could extend this framework by examining government investment rather than government consumption to assess whether public capital spending generates different state-dependent responses in business investment.

General Conclusion

This dissertation examined how the effects of fiscal shocks vary across countries and macroeconomic environments. Using evidence from developing, resource-dependent, and advanced economies, the findings show that fiscal policy transmission is highly context-specific: rather than being captured by a single multiplier or stable transmission mechanism, the effects depend on structural characteristics, institutional capacity, and prevailing macroeconomic conditions. The three essays jointly illustrate this variation.

In the WAEMU, fiscal effects differ across member states despite a common monetary framework, highlighting the role of country-specific structural constraints. In African oil-exporting economies, oil windfalls yield short-term fiscal gains, but these effects are temporary and uneven due to reliance on external markets. In advanced OECD countries, government spending impacts business investment differently depending on macroeconomic uncertainty, with crowding out occurring in stable periods but weakening when firms delay investment under heightened uncertainty.

The central contribution of this dissertation is to show that the fiscal multiplier is not fixed, but varies with structural characteristics and prevailing economic conditions. More broadly, the dissertation advances the view that fiscal policy should be assessed in context, rather than through uniform assumptions about its effects. This contributes to the literature by integrating evidence across distinct economic settings and showing that differences in institutional capacity, resource dependence, and uncertainty can fundamentally alter fiscal outcomes.

From a policy perspective, the findings suggest that effective fiscal policy design requires more than the size of public expenditure alone. Fiscal interventions must be adapted to the institutional and macroeconomic context in which they are implemented. In practice, this means strengthening fiscal institutions in constrained monetary unions, improving stabilization and diversification mechanisms in resource-dependent economies, and complementing fiscal stimulus with measures that reduce uncertainty and support private-sector confidence, when investment responses are weak.

Future research could build on this work by examining how the composition and financing of fiscal shocks shape their macroeconomic effects. In particular, distinguishing between current and capital expenditure, as well as between spending- and revenue-based shocks, would help clarify whether different fiscal instruments produce different transmission patterns across countries and over time. Further research could also investigate how fiscal policy interacts with monetary conditions and macroeconomic uncertainty, especially in environments where private sector responses may vary across states of the economy. Such extensions would deepen understanding of how fiscal policy can be better designed to support both short-run stabilization and long-run growth.

Appendix A

Appendix to Chapter One

A.1 Robustness Checks

A.1.1 Residuals AR Coefficient

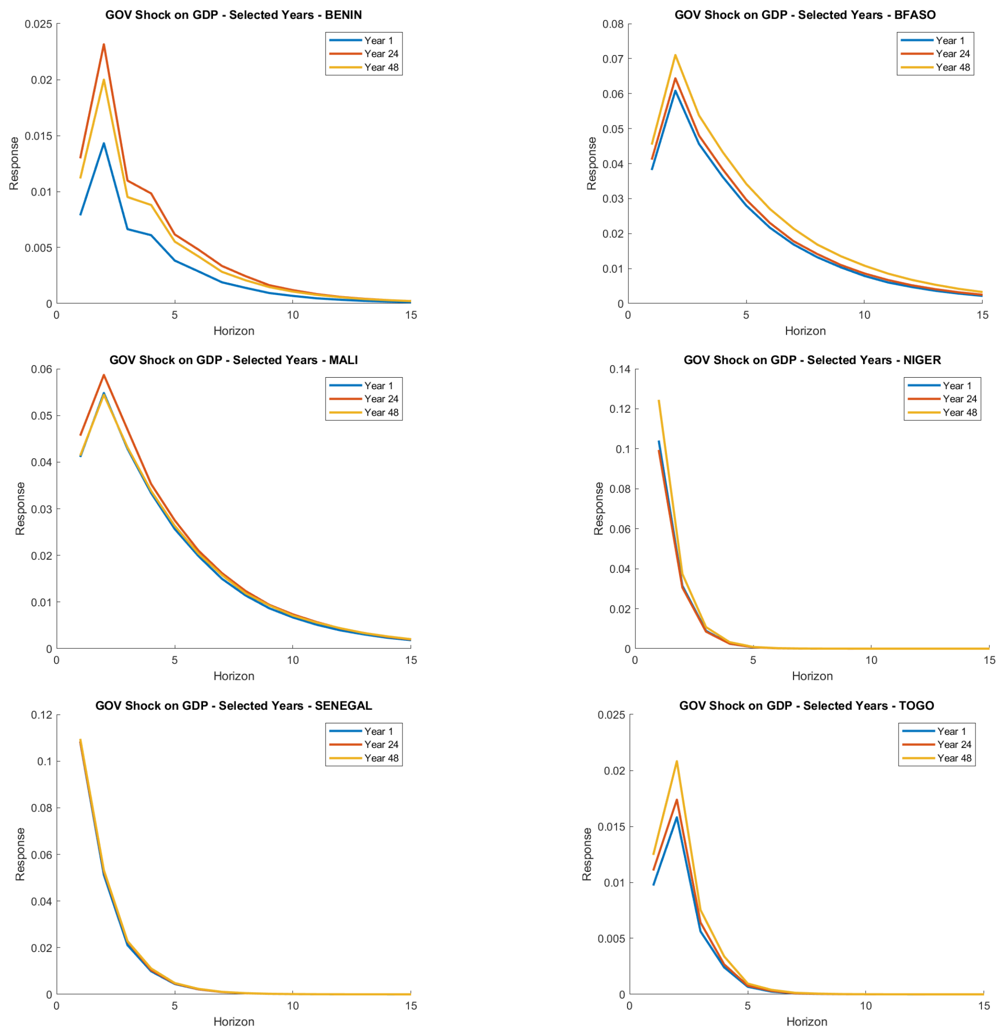


Figure A.1: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years using $\gamma = 0.7$

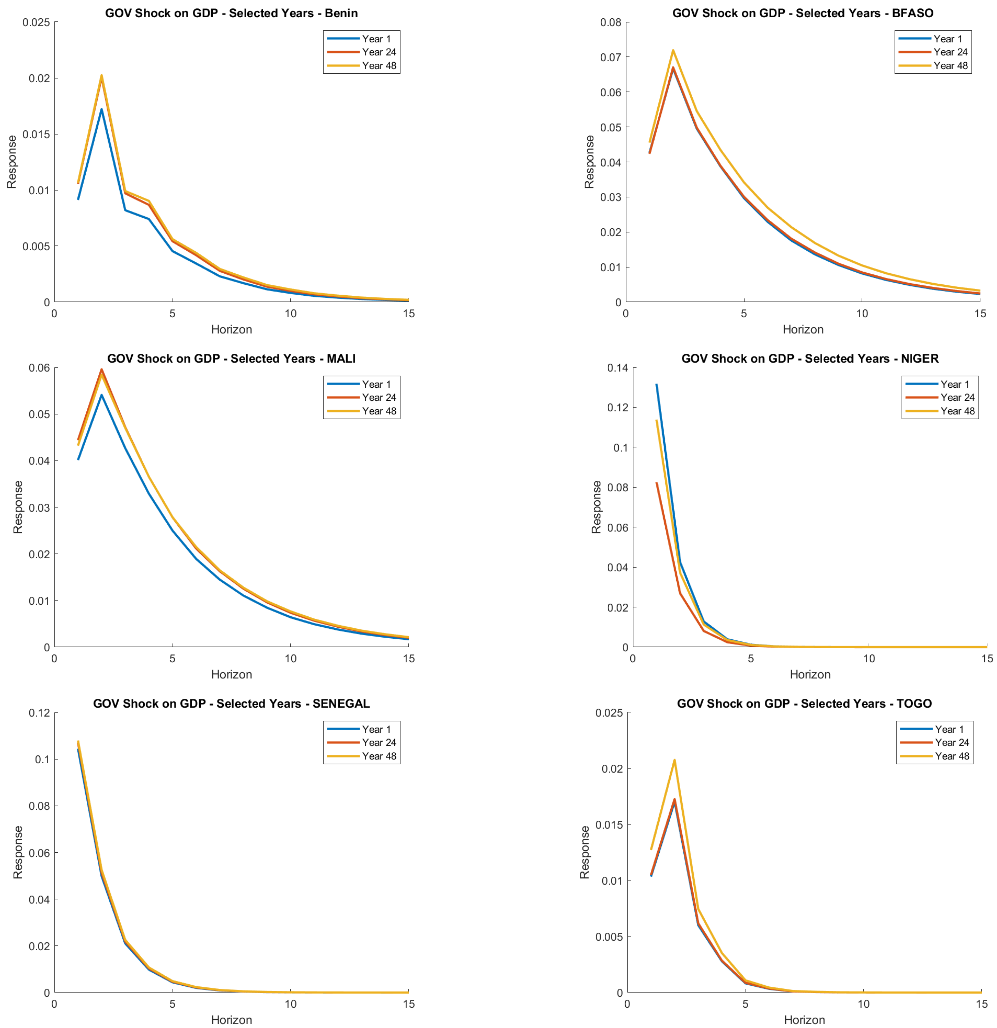


Figure A.2: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years using $\gamma = 1$

A.1.2 Two Lags

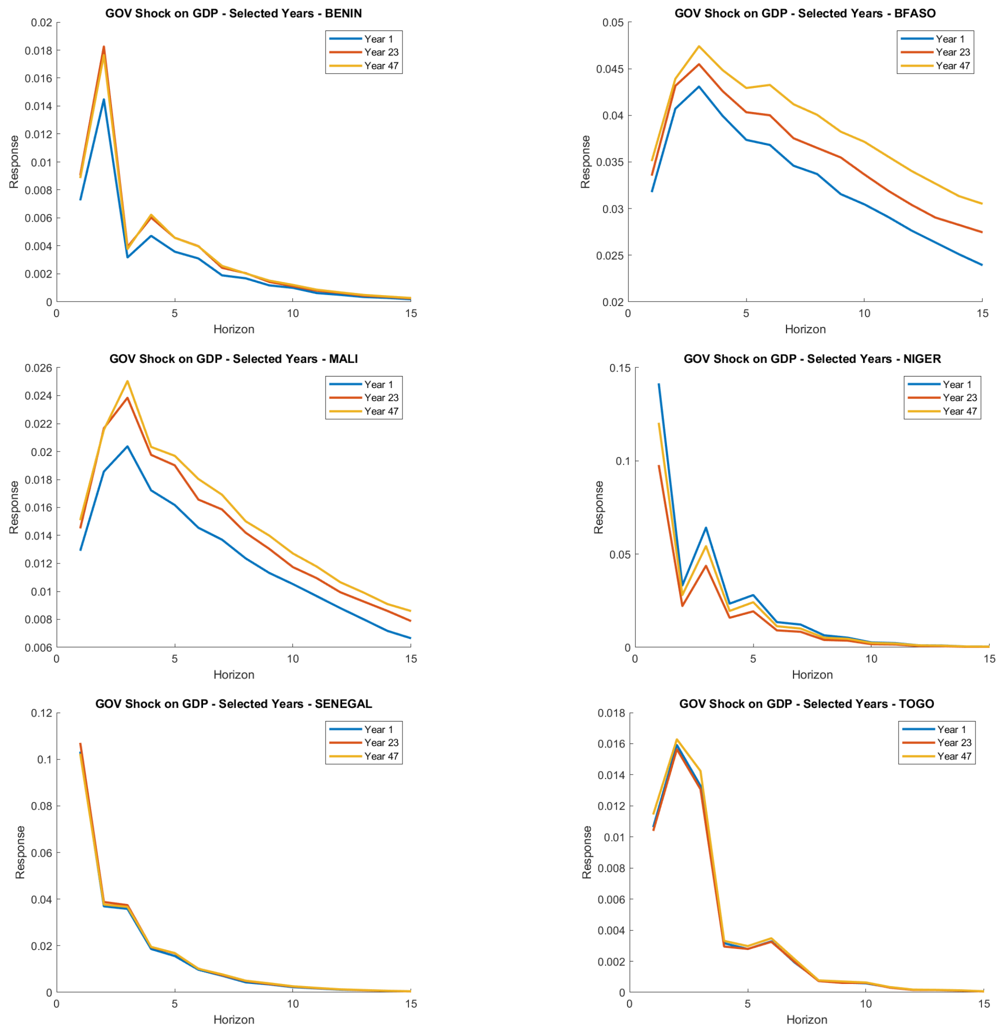


Figure A.3: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years using two lags

A.1.3 Prior Tightness

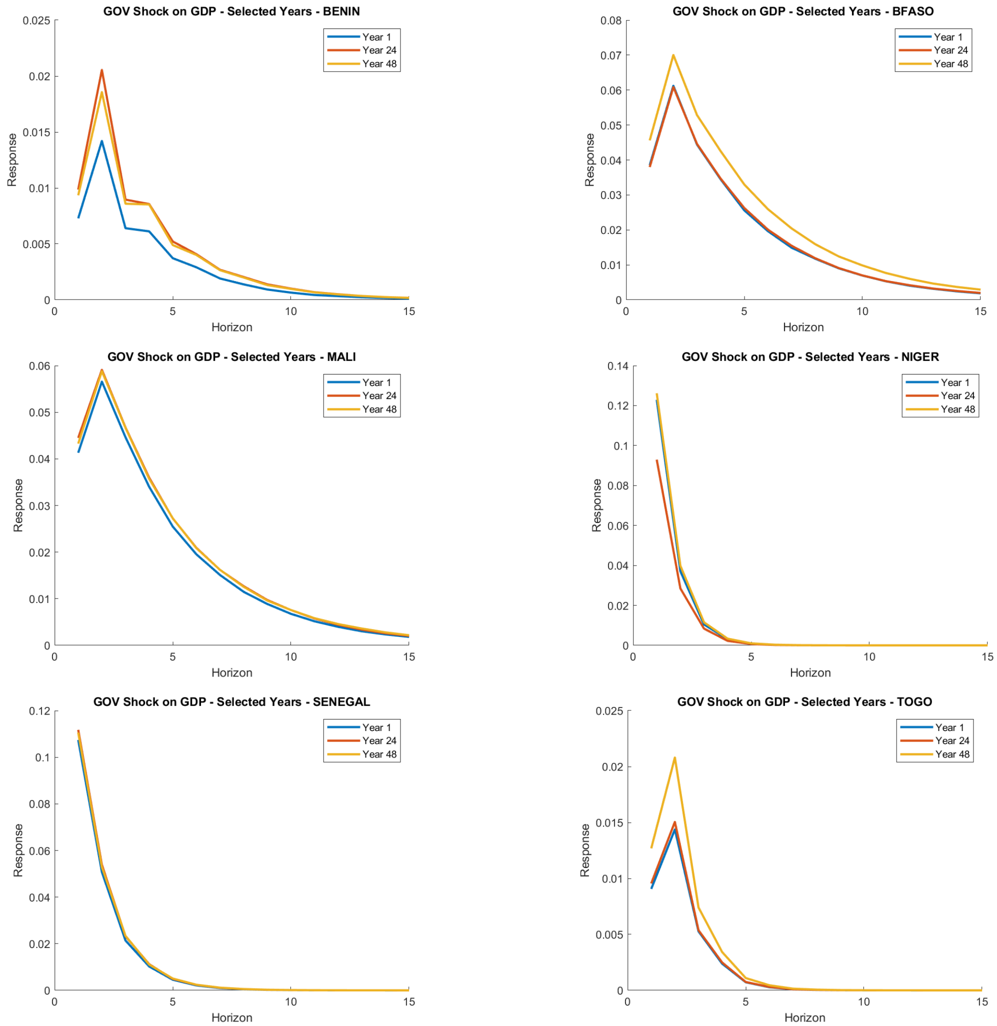


Figure A.4: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years using $\alpha_0 = \delta_0 \leq 0.01$

A.1.4 Ordering Government Expenditure Last

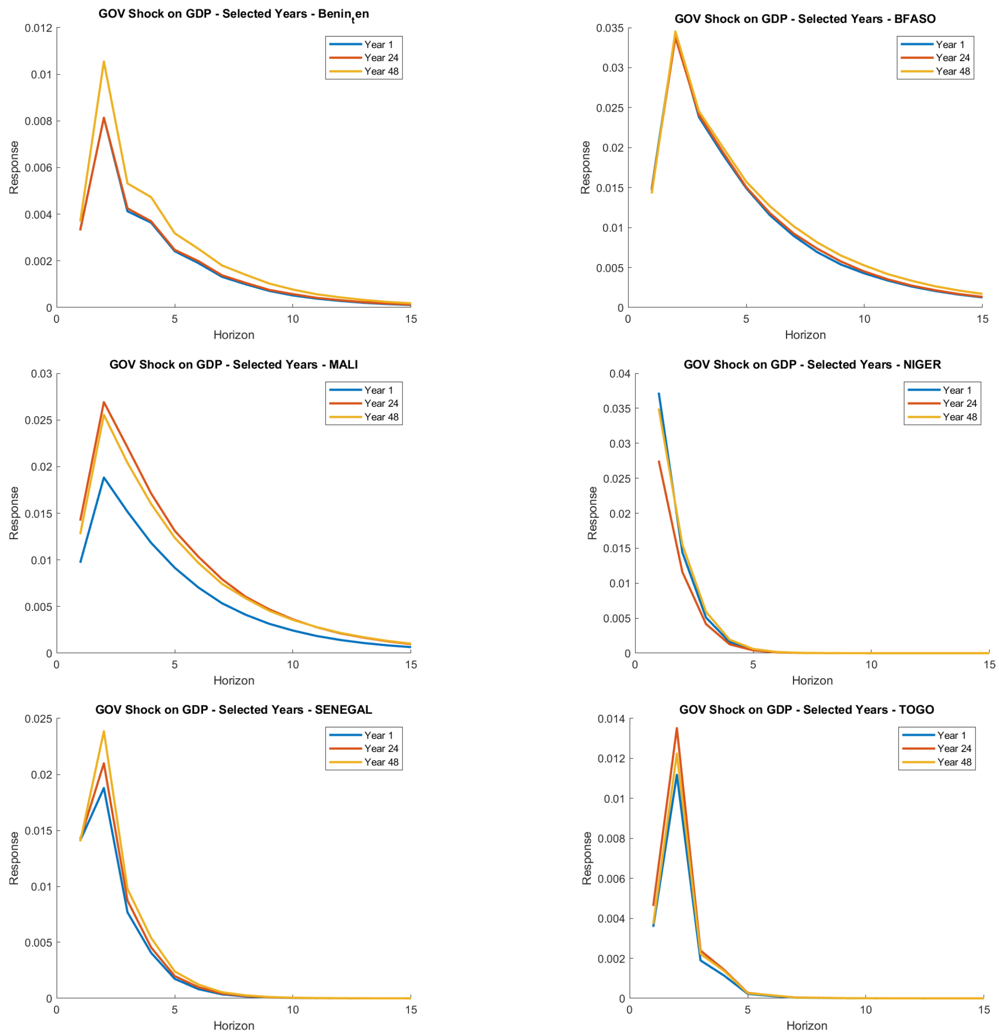


Figure A.5: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years with government expenditure ordered last

A.1.5 Variables in Log Differences

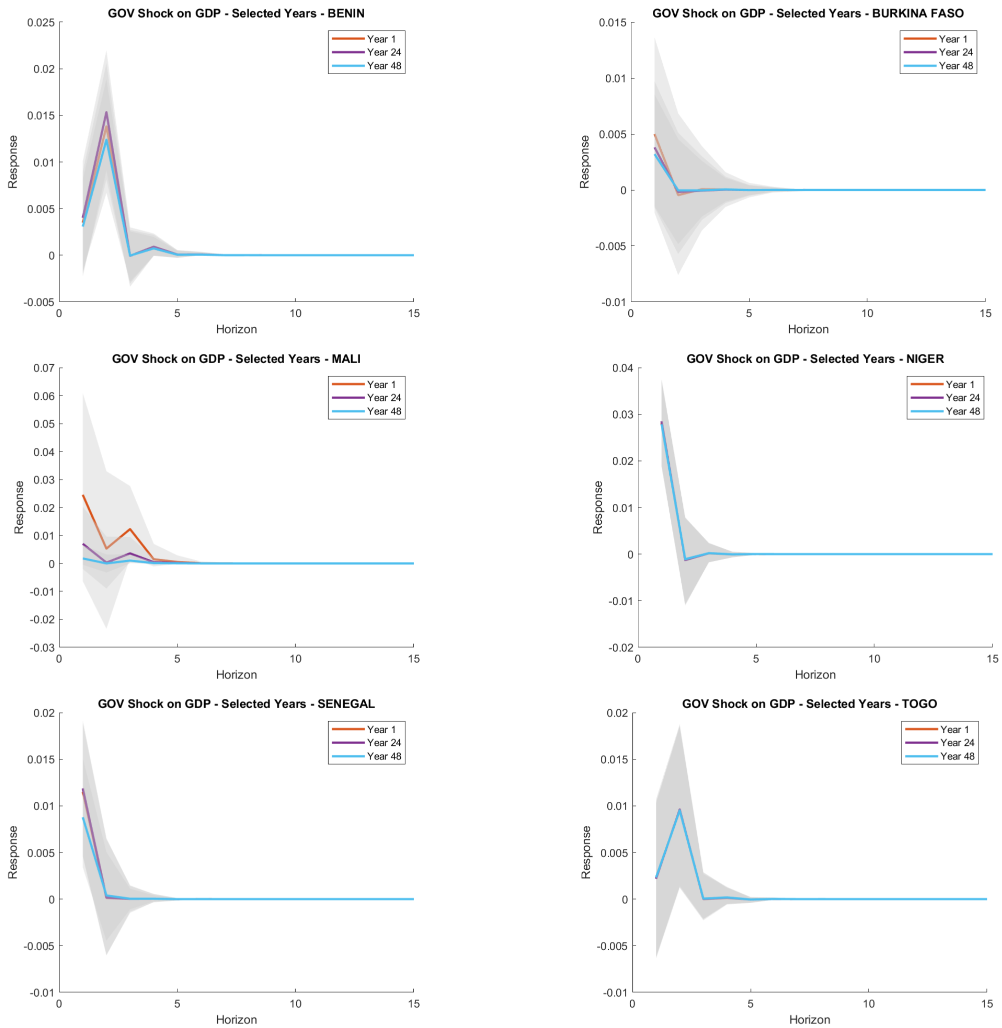


Figure A.6: Impulse responses of GDP to a 1-unit shock in Government Expenditure for selected years (68% credible intervals) with variables in log differences

A.1.6 Côte d'Ivoire

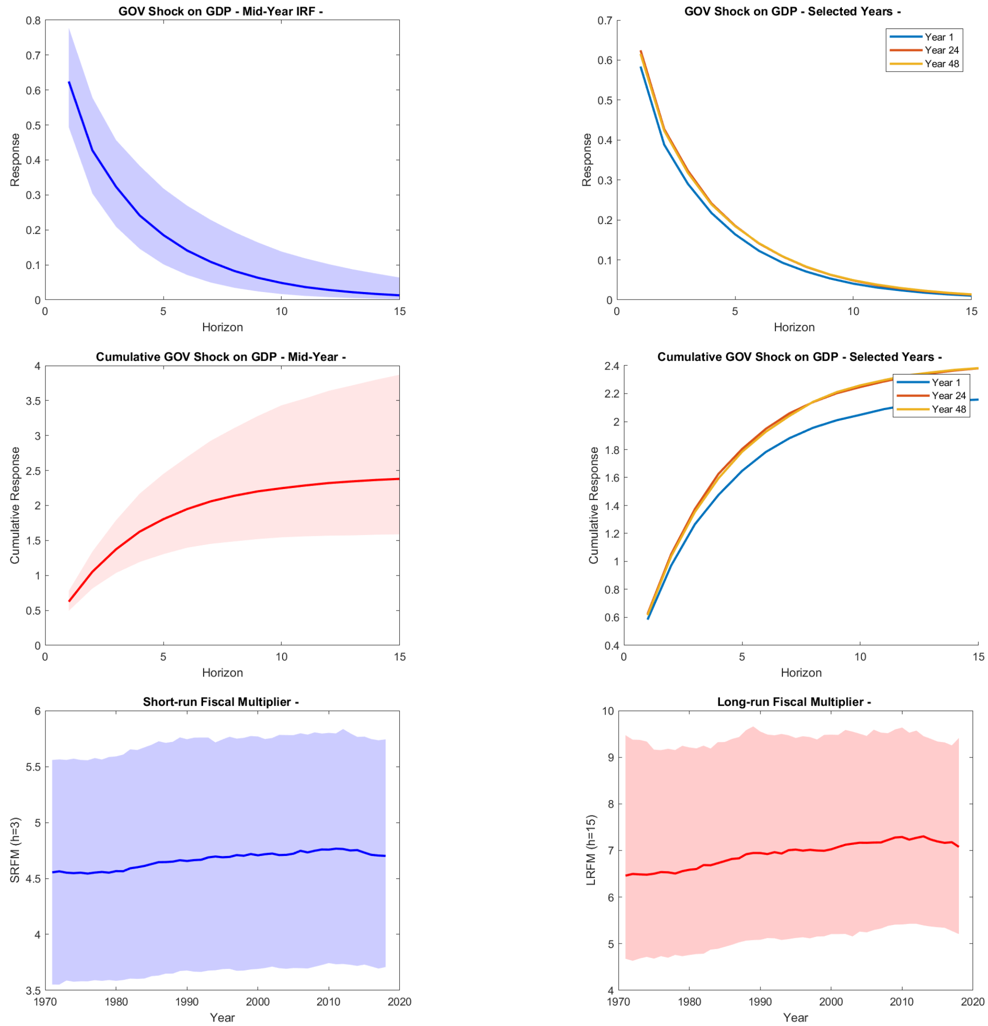
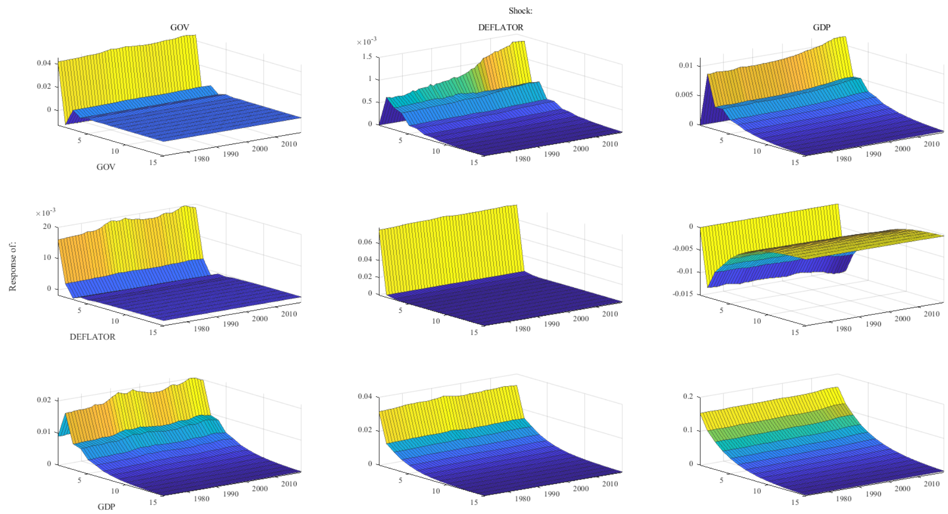
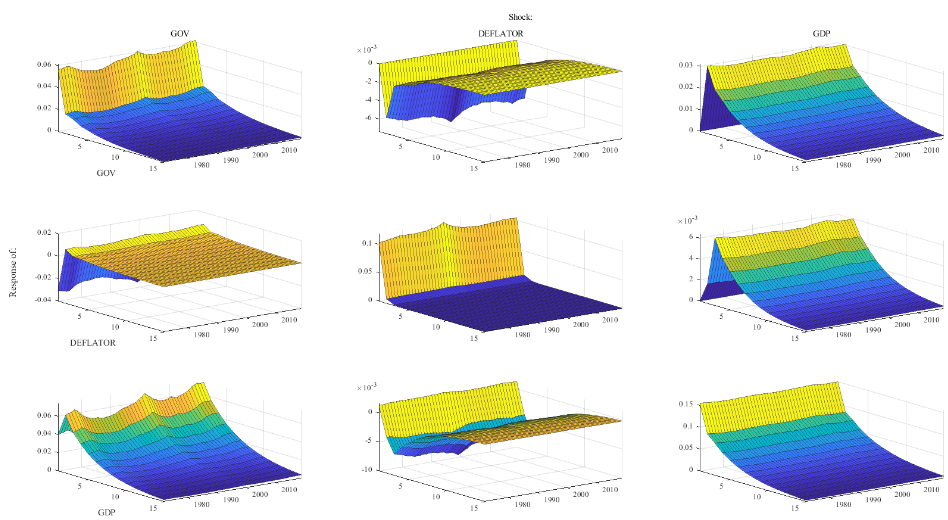


Figure A.7: Impulse responses of GDP to a 1-unit shock in Government Expenditure and Fiscal Multipliers for Côte d'Ivoire

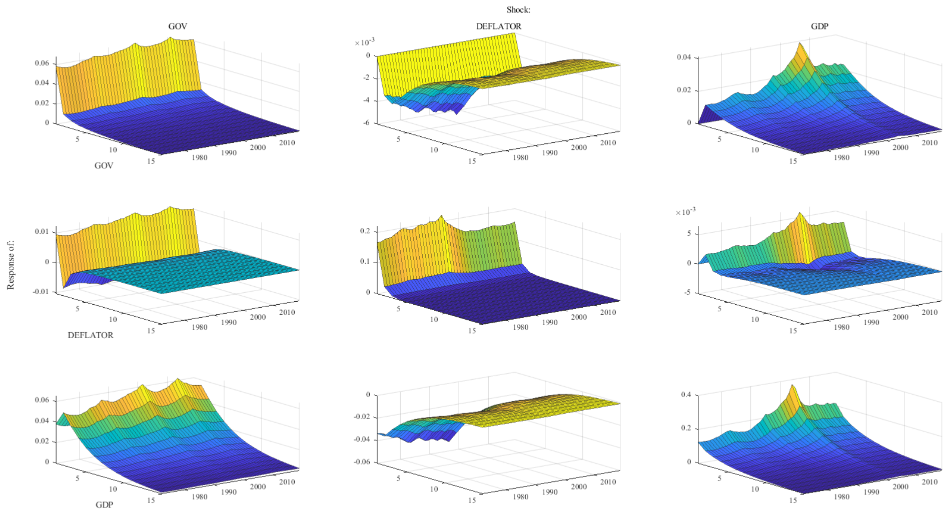
A.2 Impulse Responses in 3D Graphs



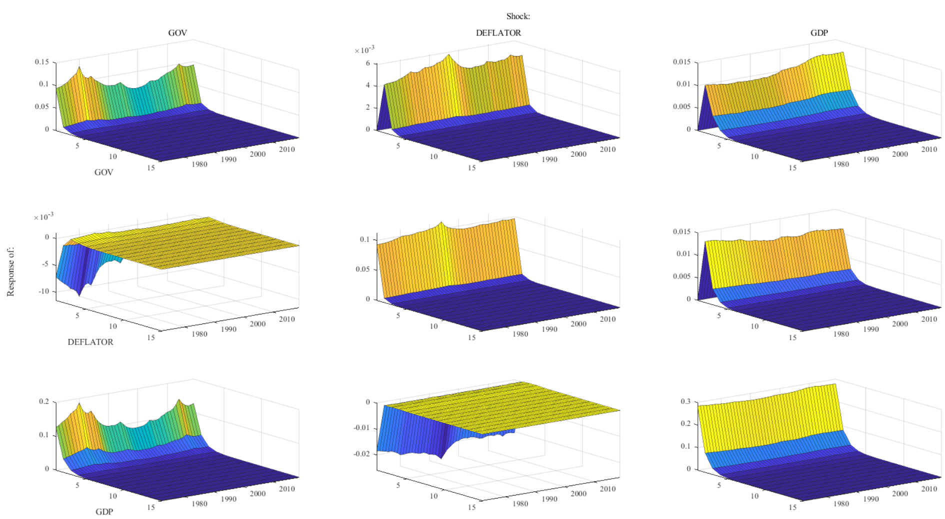
Benin



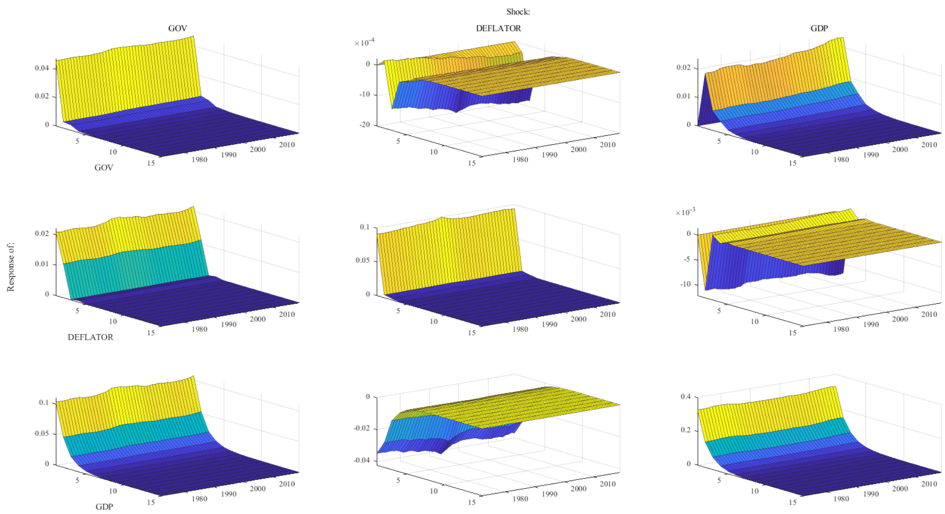
Burkina Faso


Shock:
GOV
DEFLATOR
GDP
Response of:
GOV
DEFLATOR
GDP

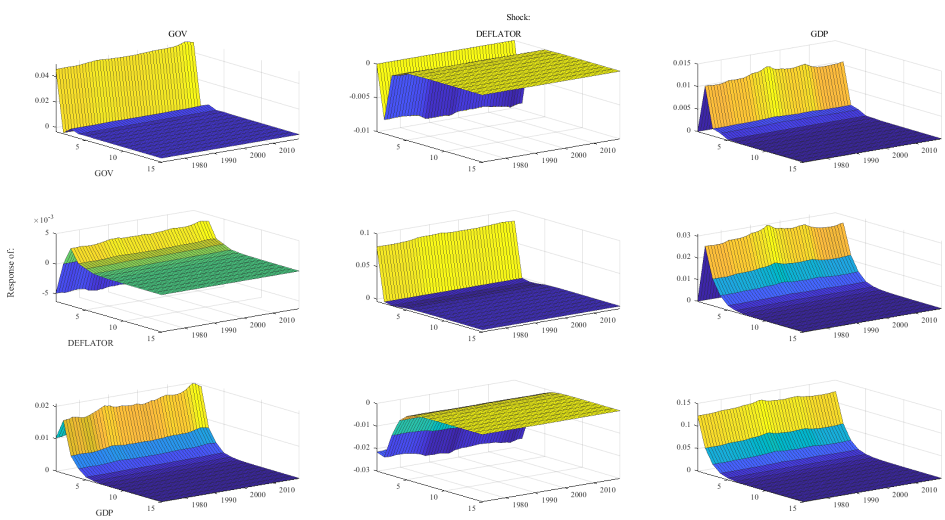
Mali


Shock:
GOV
DEFLATOR
GDP
Response of:
GOV
DEFLATOR
GDP

Niger



Senegal



Togo

A.3 Unit root tests

Table A.1: Augmented Dickey–Fuller (ADF) Unit Root Test Results

Country	**Variable**	**Levels**		**First Differences**	
		t-Statistic	**p-value**	**t-Statistic**	**p-value**
Benin	GDP	2.816	1.000	-3.534	0.047
	Government Spending	-0.405	0.985	-10.030	0.000
	Deflator	-2.264	0.444	-6.270	0.000
Burkina Faso	GDP	4.133	1.000	-5.364	0.000
	Government Spending	5.638	1.000	-4.177	0.010
	Deflator	-2.221	0.468	-5.187	0.000
Mali	GDP	4.362	1.000	-5.184	0.001
	Government Spending	1.009	1.000	-5.484	0.000
	Deflator	-2.090	0.538	-5.506	0.000
Niger	GDP	1.909	1.000	-7.1889	0.000
	Government Spending	0.245	0.998	-5.914	0.000
	Deflator	-2.071	0.549	-6.087	0.000
Senegal	GDP	2.756	1.000	-5.524	0.000
	Government Spending	1.711	1.000	-5.231	0.001
	Deflator	2.173	0.493	-6.192	0.000
Togo	GDP	1.754	1.000	-5.161	0.001
	Government Spending	-0.349	0.987	-5.784	0.000
	Deflator	-2.267	0.443	-6.191	0.000

Note: The null hypothesis is the presence of a unit root. Tests include an intercept; p-values are MacKinnon (1996) one-sided values.

A.4 Lag Length Selection using DIC

Table A.2: Deviance Information Criterion

Country	**Lag 1**	**Lag 2**	**Lag 3**	**Lag 4**
Benin	-260.70*	-252.07	-243.90	-236.27
Burkina Faso	-213.59*	-211.00	-210.16	-201.96
Mali	-133.62*	-132.18	-125.78	-121.65
Niger	-129.59*	-125.59	-124.16	-121.21
Senegal	-173.98*	-169.06	-165.06	-159.03
Togo	-273.68*	-264.29	-256.15	-248.40

A.5 Diagnostic checks

A.5.1 Trace plots

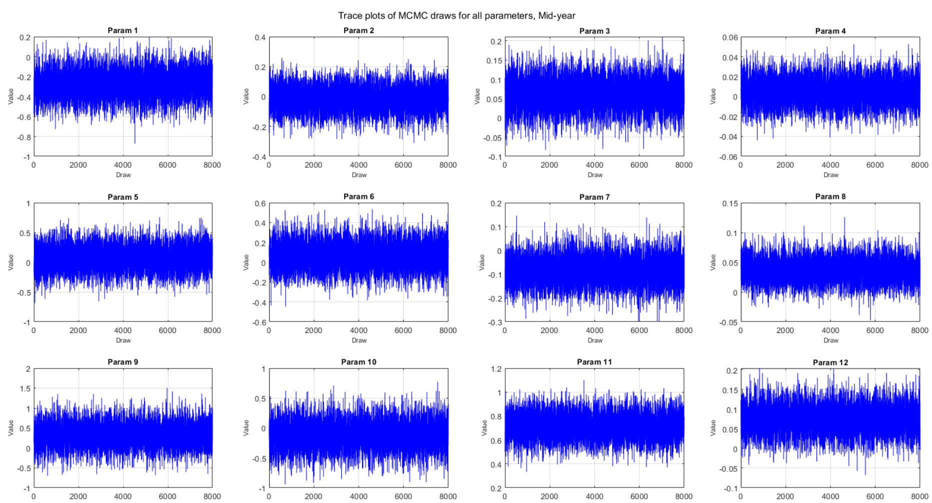


Figure A.8: Benin

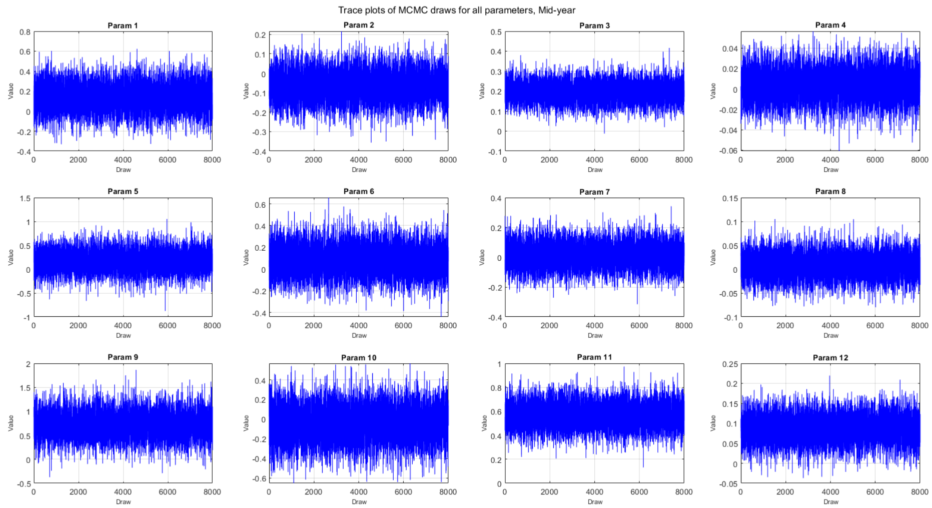


Figure A.9: Burkina Faso

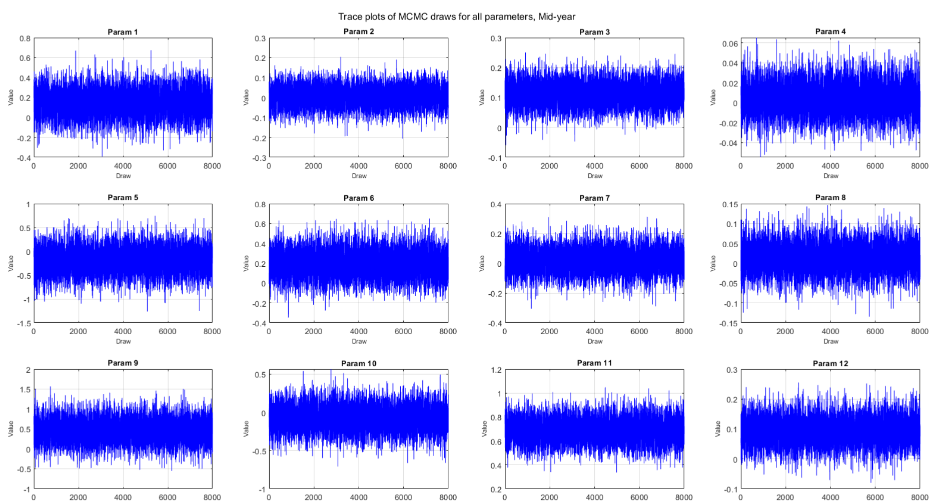


Figure A.10: Mali

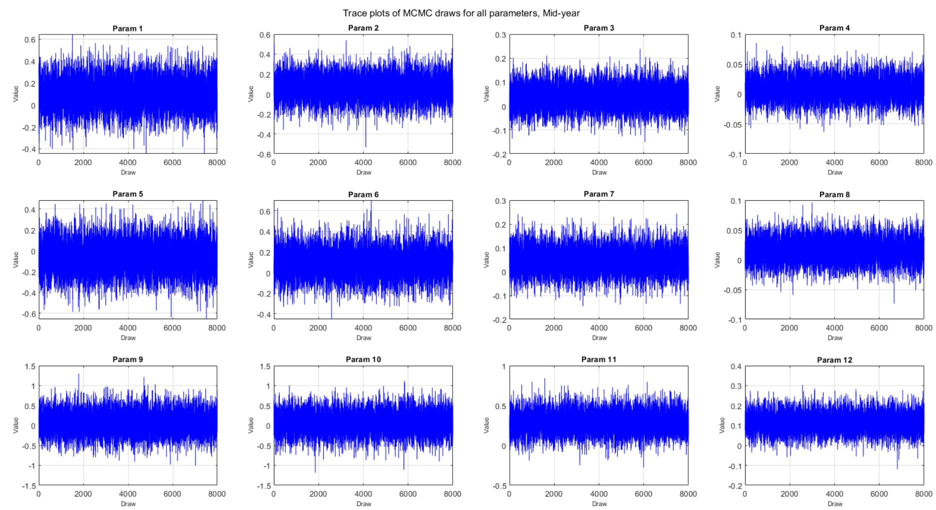


Figure A.11: Niger

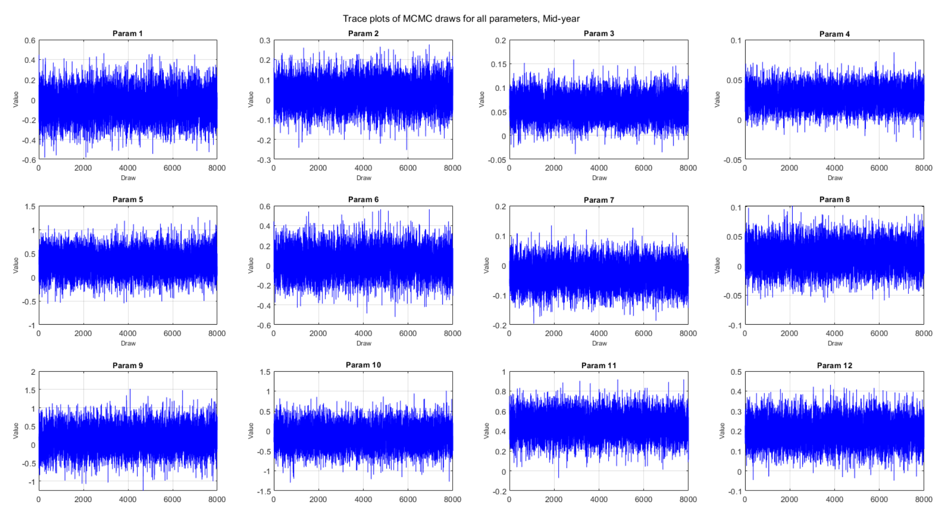


Figure A.12: Senegal

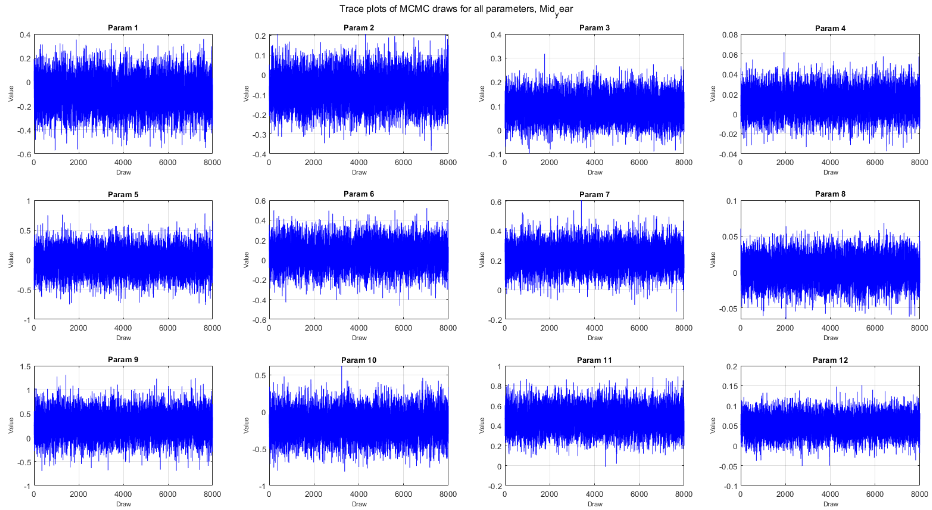


Figure A.13: Togo

A.5.2 Autocorrelation Function (ACF)

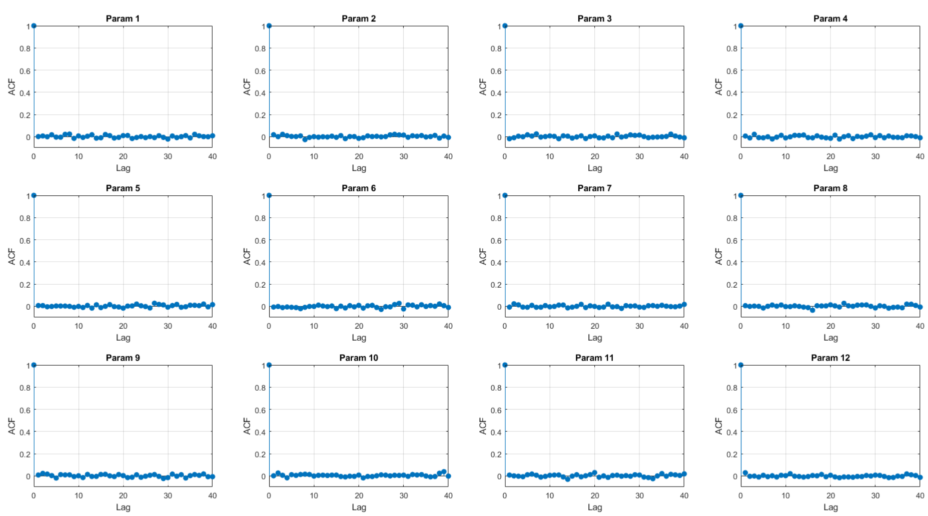


Figure A.14: Benin

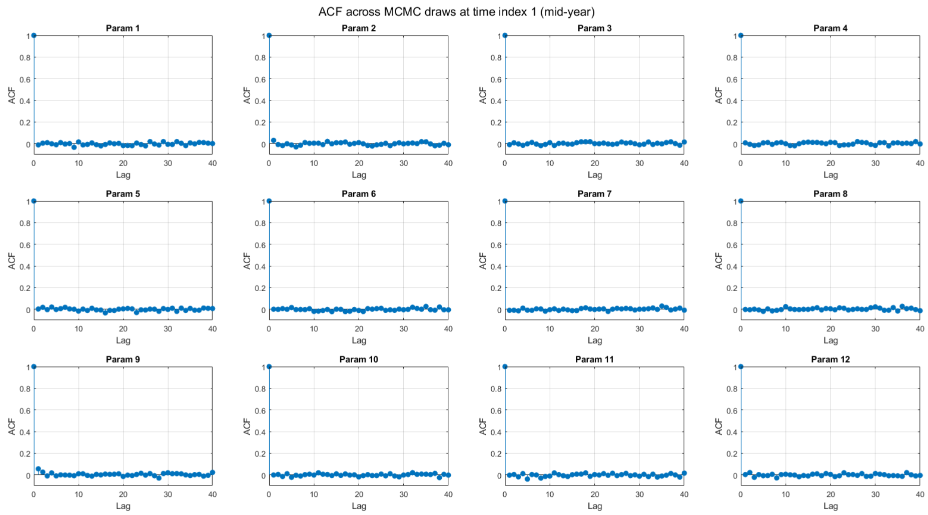


Figure A.15: Burkina Faso

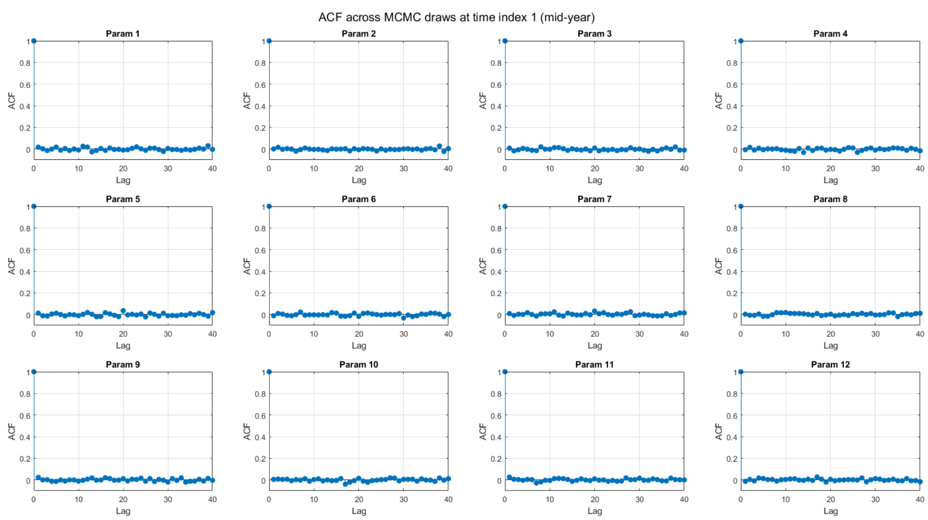


Figure A.16: Mali

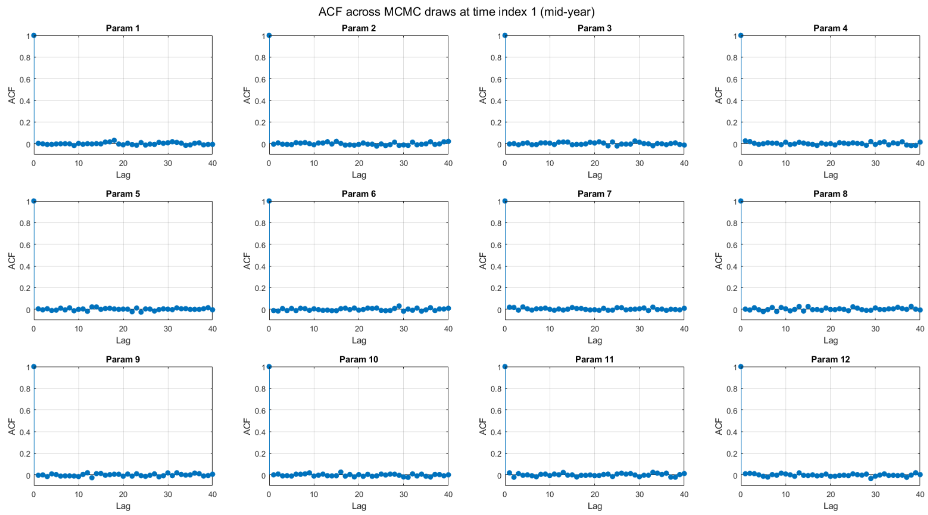


Figure A.17: Niger

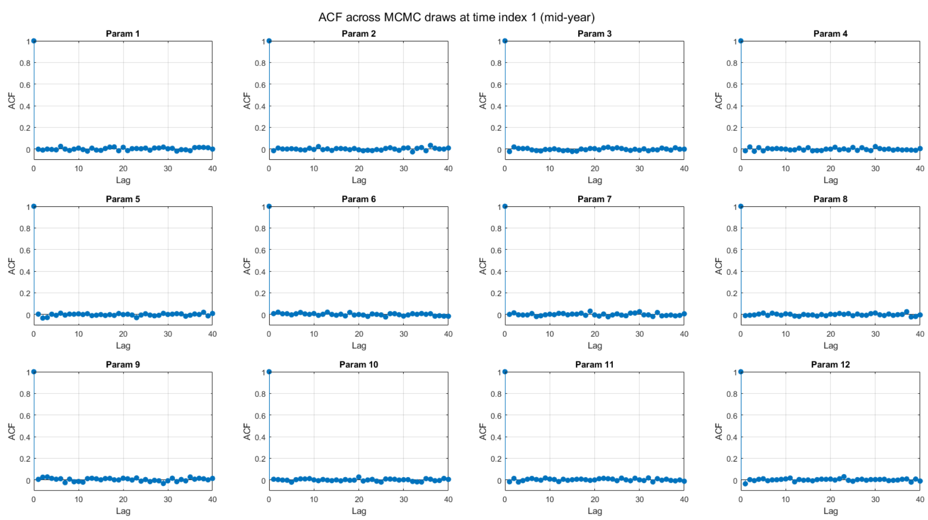


Figure A.18: Senegal

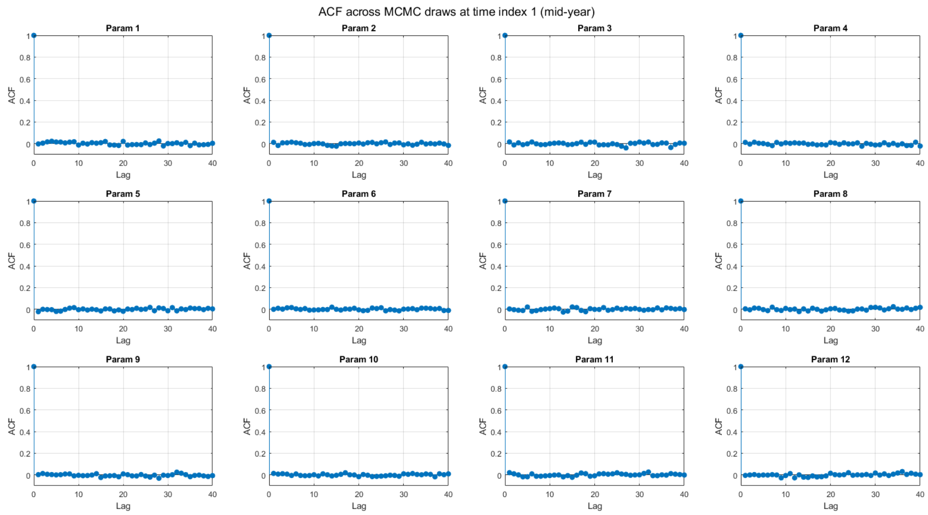


Figure A.19: Togo

A.5.3 Effective Sample Size (ESS)

Table A.3: Effective Sample Size (ESS) Values by Country

Parameter	Benin	Burkina Faso	Mali	Niger	Senegal	Togo
1	7881.0	8000.0	7695.2	7947.5	8000.0	8000.0
2	7157.2	7537.9	7734.0	8000.0	8000.0	7805.0
3	8000.0	8000.0	7862.5	8000.0	8000.0	7748.3
4	7906.4	7871.6	8000.0	7280.4	8000.0	7808.8
5	7806.5	7678.6	7804.8	7933.0	7946.9	8000.0
6	8000.0	7988.6	8000.0	8000.0	7364.6	7326.2
7	8000.0	8000.0	7860.0	7462.5	7747.1	7949.9
8	7914.1	8000.0	7955.0	7971.4	8000.0	7932.4
9	7287.2	6869.2	7601.3	8000.0	6759.5	7627.9
10	7509.3	8000.0	7624.0	7862.4	7849.3	7414.4
11	7876.4	8000.0	7466.1	7707.7	8000.0	7533.1
12	7569.2	7618.1	8000.0	7445.4	8000.0	8000.0

A.5.4 Geweke Z-Scores

Table A.4: Geweke Z Scores

Parameter	Benin	Burkina Faso	Mali	Niger	Senegal	Togo
1	-3.848	+0.619	-0.528	-1.243	+0.048	-0.289
2	+0.785	-1.727	-0.501	+1.438	+0.308	+1.401
3	+0.223	+0.628	-0.592	-0.422	-0.521	+0.334
4	+0.319	-1.558	+1.624	-1.140	+0.506	+0.138
5	-3.644	-1.504	-1.500	+1.529	-0.878	+0.819
6	-0.414	+0.497	+0.284	+1.225	-0.441	-0.797
7	-0.701	+0.950	+1.466	+0.931	+0.384	+0.026
8	+0.368	+0.347	+0.032	+0.123	-0.228	+1.635
9	-2.262	+0.378	-3.363	+2.293	-1.571	-0.554
10	-0.952	-1.342	+0.969	-1.301	+1.041	+0.954
11	-0.873	+0.519	+2.258	-2.623	+0.036	-0.711
12	+0.990	-0.280	-0.867	+1.518	+0.161	+0.004

A.6 Bayesian Estimation

A.6.1 TVP-SVAR

The likelihood is given as:

$$f(y|\beta, f^{-1}, \lambda) \propto \prod_{t=1}^{T} |F\Lambda_t F'|^{-1/2} \exp\left(-\frac{1}{2}(y_t - \bar{X}_t\beta_t)'(F\Lambda_t F')(y_t - \bar{X}_t\beta_t)\right) \quad \text{(A.1)}$$

The prior formulation for $\pi(\beta|Q)$ is given by:

$$\pi(\beta|Q) \sim N(0, Q_0) \quad \text{(A.2)}$$

The prior for each term ω_i in Q follows an inverse Gamma distribution:

$$\pi(\omega_i) \sim IG(\frac{\chi_0}{2}, \frac{\psi_0}{2}) \tag{A.3}$$

with hyperparameters $\chi_0 = \psi_0 = 0.001$ to reflect a weakly informative prior.

f_i^{-1} is denoted as elements of the ith row in F^{-1} that are not zero or one. The prior is given by:

$$f^{-1} \sim N(f_{i0}^{-1}, \Upsilon_{i0}) \tag{A.4}$$

The conditional prior for the general case of $t = 2, \ldots, T$, that is, any sample period but the first one:

$$\pi(\lambda_{i,t}|\lambda_{i,t-1}, \phi_i) \sim \mathcal{N}(\gamma\lambda_{i,t-1}, \phi_i) \tag{A.5}$$

For the first sample period:

$$\pi(\lambda_{i,1}|\phi_i) \sim \mathcal{N}(0, \phi_i\omega) \tag{A.6}$$

Thus, the joint density is given as:

$$\pi(\lambda_i|\phi_i) = \pi(\lambda_{i,1}|\phi_i) \prod_{t=2}^{T} \pi(\lambda_{i,t}|\lambda_{i,t-1}, \phi_i) \tag{A.7}$$

The prior for each heteroskedasticity variance parameter ϕ_i is given by:

$$\pi(\phi_i) \sim IG(\frac{\alpha_0}{2}, \frac{\delta_0}{2}) \tag{A.8}$$

Given the likelihood function in Equation A.1 and the prior distribution formulations, the conditional posterior distributions for the parameters of interest (β, ω_i, f_i^{-1}, $\lambda_{i,t}$, ϕ_i) are:

$$\pi(\beta \mid y, Q, \Sigma) \sim N(\bar{B}, \bar{Q}) \tag{A.9}$$

$$\pi(\omega_i \mid y, \beta, \omega_{-i}, \Sigma) \propto \omega_i^{-\bar{\chi}-1} \exp\left(-\frac{\bar{\psi}}{\omega_i}\right) \tag{A.10}$$

$$\pi(f_i^{-1} \mid y, \beta, f_{-i}^{-1}, \lambda, \phi) \sim N(\bar{f}_i^{-1}, \bar{\Upsilon}_i) \tag{A.11}$$

$$\begin{aligned}\pi(\lambda_{i,t} \mid y, \beta, f^{-1}, \lambda_{-i,-t}, \phi) \propto \exp&\left[-\frac{1}{2}\bar{s}_i^{-1} e^{-\lambda_{i,t}} (\varepsilon_{i,t} + (f_i^{-1})' \varepsilon_{-i,t})^2 + \lambda_{i,t}\right] \\ &\times \exp\left[-\frac{1}{2} \frac{(\lambda_{i,t} - \bar{\lambda}_i)^2}{\bar{\phi}_i}\right]\end{aligned} \tag{A.12}$$

$$\pi(\phi_i \mid y, \beta, f^{-1}, \lambda_i, \phi_{-i}) \sim IG\left(\frac{\bar{\alpha}}{2}, \frac{\bar{\delta}}{2}\right) \tag{A.13}$$

Sampling the log-volatility $\lambda_{i,t}$

As described in Dieppe et al. (2016), sampling the log-volatility term $\lambda_{i,t}$ requires a Metropolis–Hastings step because its conditional posterior distribution in Equation A.12 is non-standard. The acceptance function for the Metropolis-Hastings step is given as:

$$\alpha\big(\lambda_{i,t}^{(n-1)}, \lambda_{i,t}^{(n)}\big) = \exp\Bigg[-\frac{1}{2}\Big(e^{-\lambda_{i,t}^{(n)}} - e^{-\lambda_{i,t}^{(n-1)}}\Big)\bar{s}_i^{-1}\big(\varepsilon_{i,t} + (f_i^{-1})'\varepsilon_{-i,t}\big)^2\Bigg] \times \exp\Bigg[-\frac{1}{2}\Big(\lambda_{i,t}^{(n)} - \lambda_{i,t}^{(n-1)}\Big)\Bigg] \tag{A.14}$$

For the first period ($t = 1$), draw a candidate from $\mathcal{N} \sim (\bar{\lambda}, \bar{\phi})$, with:

$$\bar{\lambda} = \frac{\gamma\lambda_{i,2}}{\omega^{-1} + \gamma^2}, \qquad \bar{\phi} = \frac{\phi_i}{\omega^{-1} + \gamma^2} \tag{A.15}$$

The acceptance function is given by:

$$\alpha\big(\lambda_{i,1}^{(n-1)}, \lambda_{i,1}^{(n)}\big) = \exp\Bigg[-\frac{1}{2}\Big(e^{-\lambda_{i,1}^{(n)}} - e^{-\lambda_{i,1}^{(n-1)}}\Big)\bar{s}_i^{-1}\big(\varepsilon_{i,1} + (f_i^{-1})'\varepsilon_{-i,1}\big)^2\Bigg] \times \exp\Bigg[-\frac{1}{2}\Big(\lambda_{i,1}^{(n)} - \lambda_{i,1}^{(n-1)}\Big)\Bigg] \tag{A.16}$$

For the last period ($t = T$), draw a candidate from $\mathcal{N} \sim (\bar{\lambda}, \bar{\phi})$, with:

$$\bar{\lambda} = \gamma\lambda_{i,T-1}, \qquad \bar{\phi} = \phi_i \tag{A.17}$$

Appendix B

Appendix to Chapter Two

B.1 Unit root tests

Table B.1: Augmented Dickey–Fuller (ADF) Unit Root Test Results

Country	**Variable**	**Levels**		**First Differences**	
		t-Statistic	**p-value**	**t-Statistic**	**p-value**
Algeria	OIL	-2.45	0.136	-4.86	0.000
	GDP	0.83	0.993	-3.77	0.007
	GOV	0.66	0.989	-5.03	0.000
	INFLATION	-1.06	0.723	-5.78	0.000
	REER	-8.49	0.000	-2.44	0.021
Angola	OIL	-2.60	0.104	-7.34	0.000
	GDP	-0.62	0.854	-3.77	0.007
	GOV	-1.45	0.549	-13.64	0.000
	INFLATION	-1.23	0.651	-5.96	0.000
	REER	-1.70	0.423	-5.94	0.000
Egypt	OIL	-2.32	0.130	-5.87	0.000
	GDP	-0.63	0.852	-4.43	0.001
	GOV	-1.78	0.385	-6.64	0.000
	INFLATION	-1.12	0.697	-4.92	0.000
	REER	-3.17	0.030	-4.75	0.000
Nigeria	OIL	-2.69	0.085	-7.72	0.000
	GDP	0.017	0.954	-3.41	0.017
	GOV	-0.51	0.877	-6.40	0.000
	INFLATION	-1.53	0.51	-3.97	0.004
	REER	-2.26	0.189	-4.76	0.000

Note: Null hypothesis is the presence of a unit root. Tests include an intercept; p-values are MacKinnon (1996) one-sided values.

B.2 Diagnostic checks

B.2.1 Trace plots

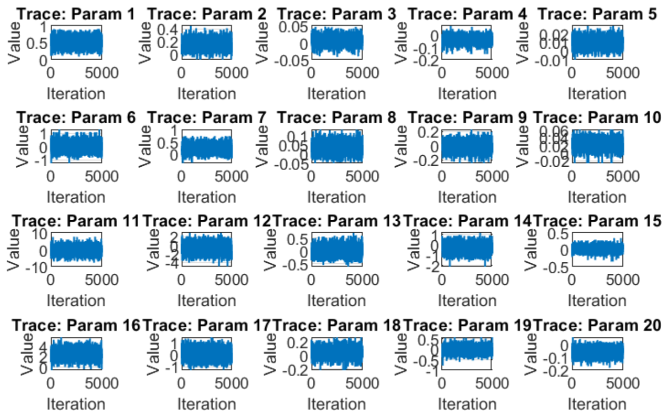


Figure B.1: Trace plots: Algeria

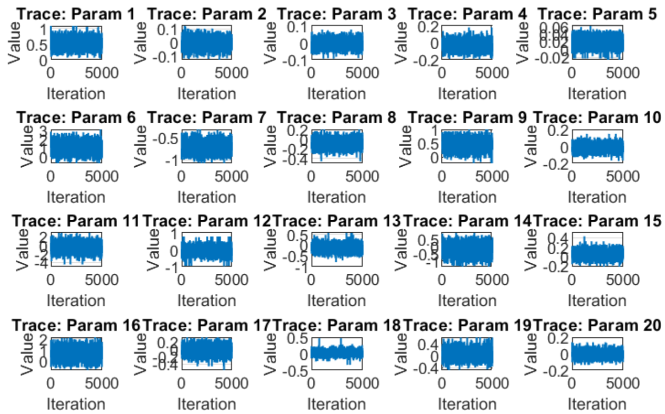


Figure B.2: Trace plots: Angola

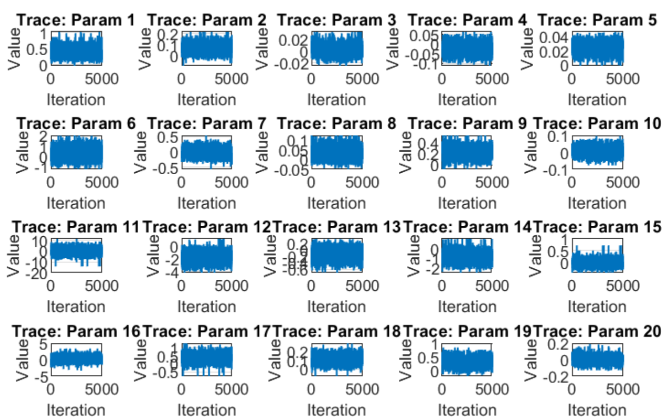


Figure B.3: Trace plots: Egypt

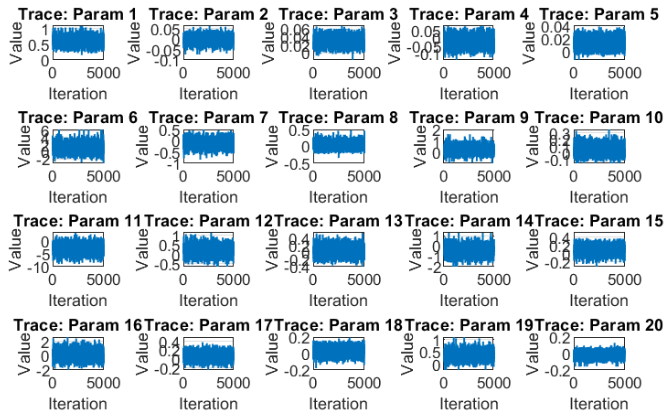


Figure B.4: Trace plots: Nigeria

B.2.2 Autocorrelation Function (ACF)

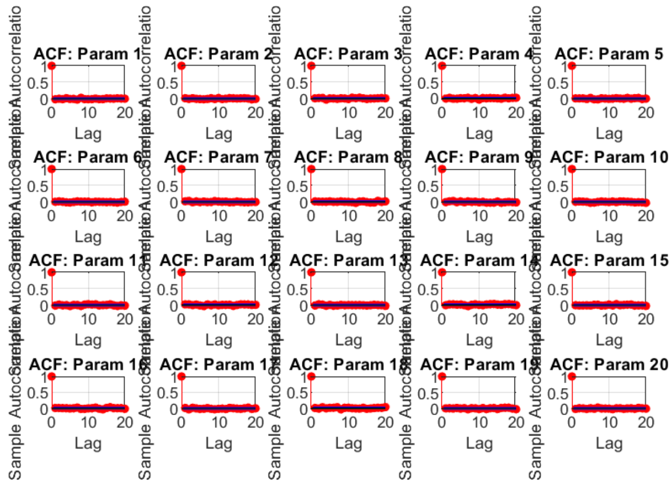


Figure B.5: Autocorrelation Function: Algeria

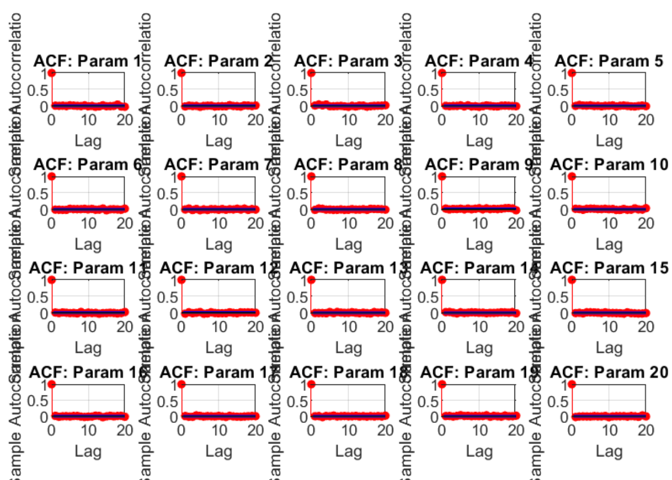


Figure B.6: Autocorrelation Function: Angola

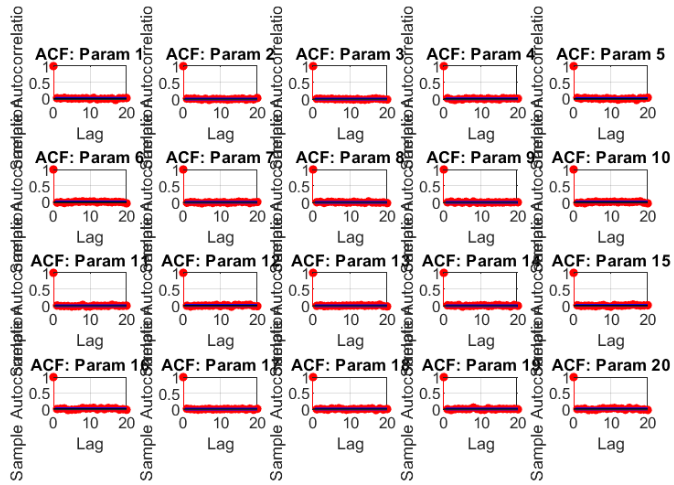


Figure B.7: Autocorrelation Function: Egypt

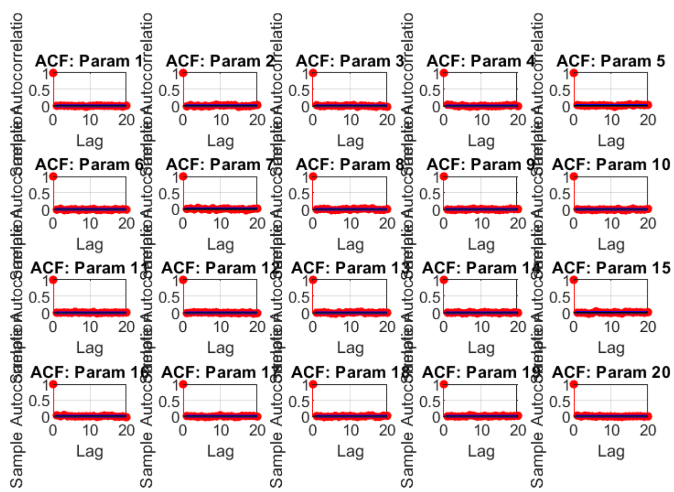
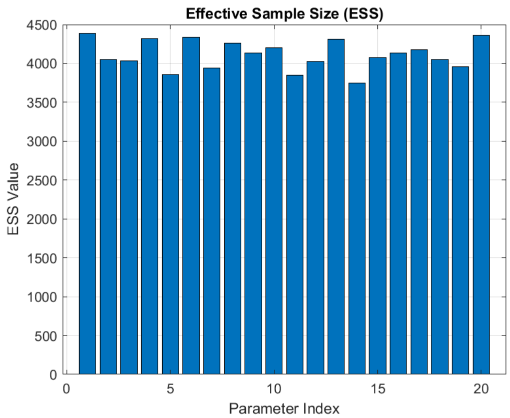
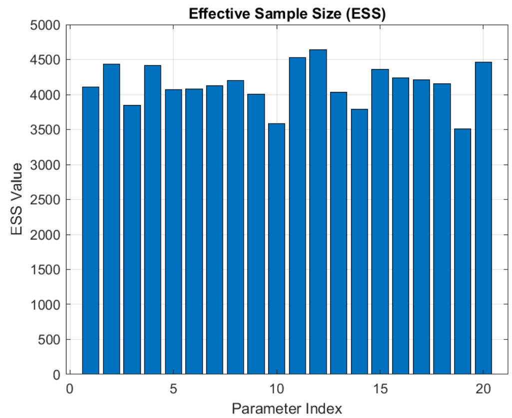


Figure B.8: Autocorrelation Function: Nigeria

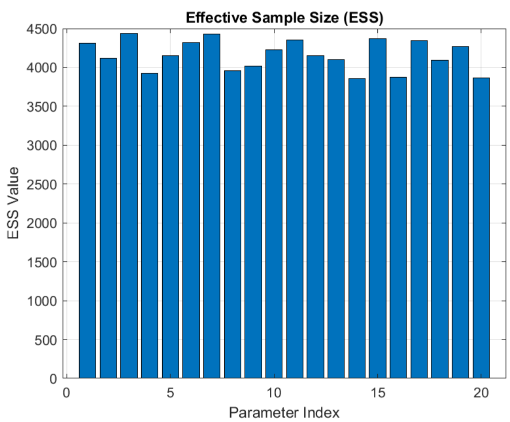
B.2.3 Effective Sample Size (ESS)



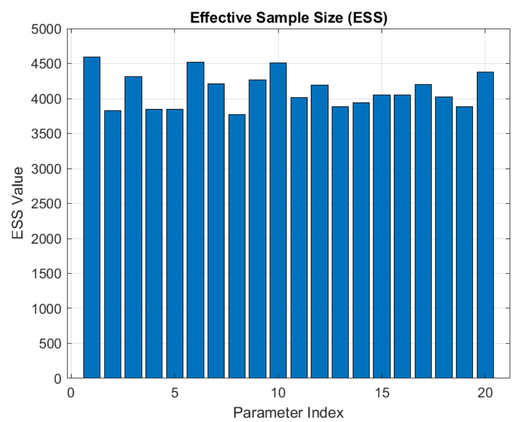
(a) Algeria



(b) Angola

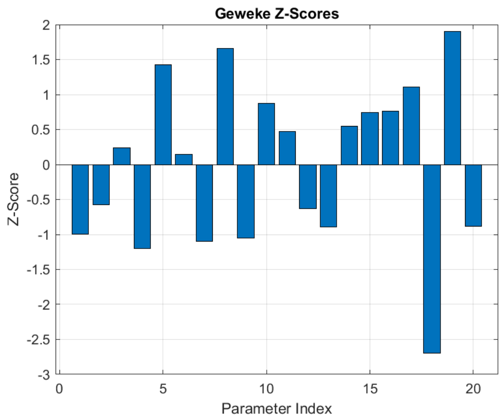


(c) Egypt

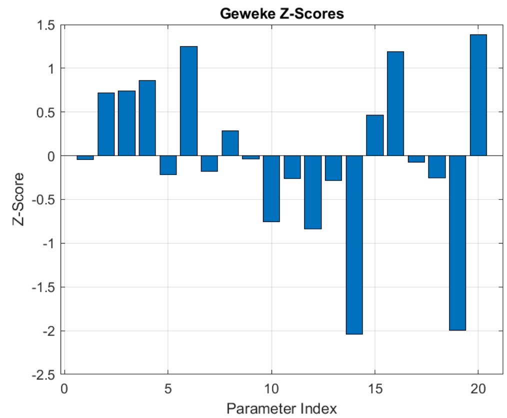


(d) Nigeria

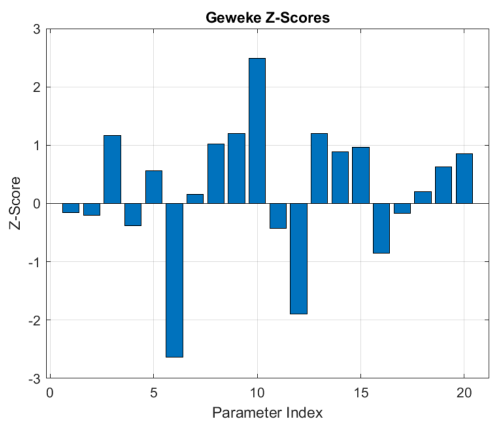
B.2.4 Geweke ZScores



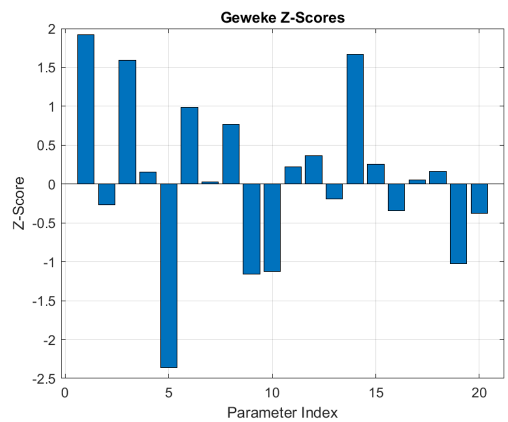
(a) Algeria



(b) Angola



(c) Egypt



(d) Nigeria

B.3 Robustness Check

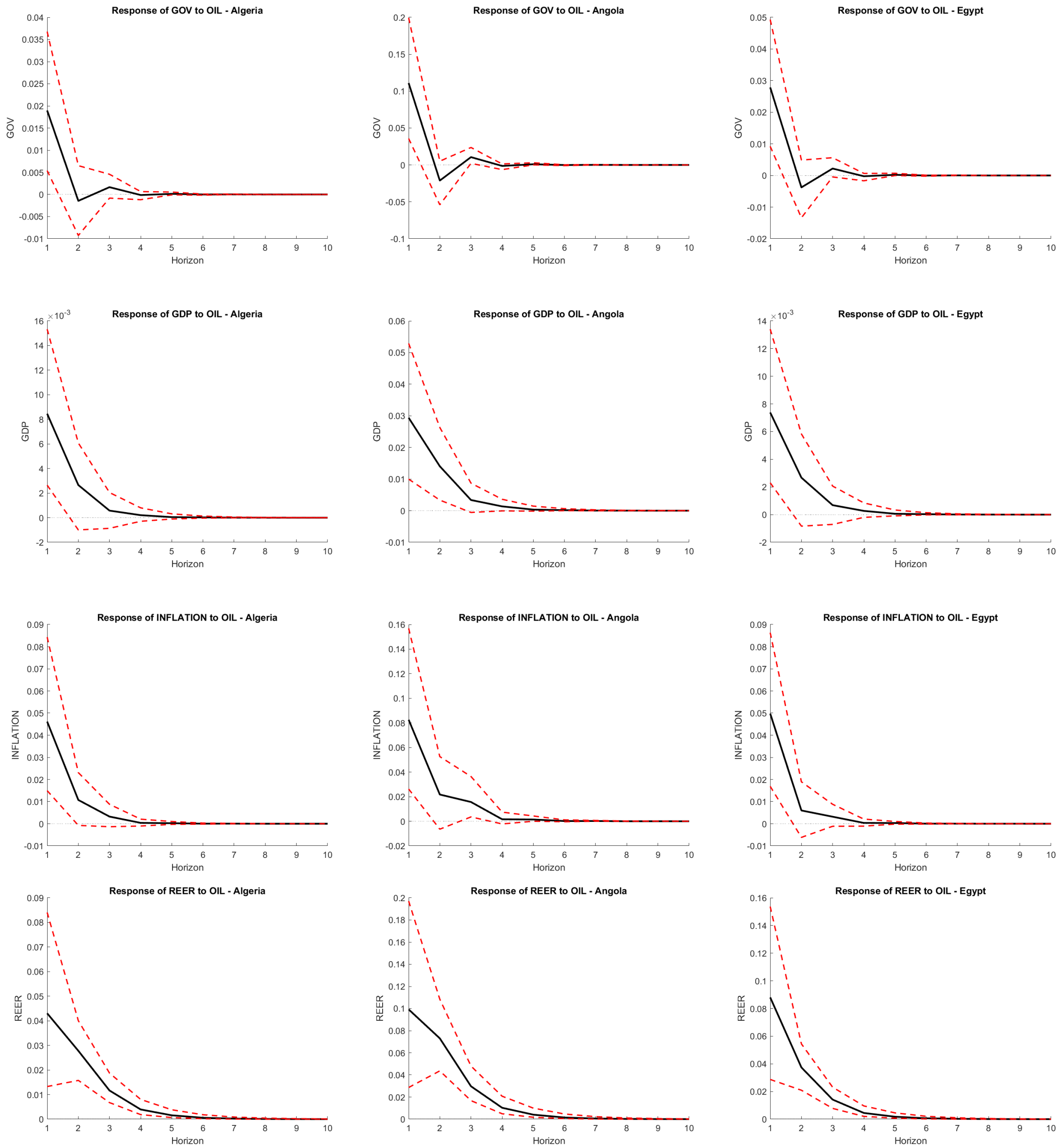


Figure B.11: Impulse responses of government expenditure, GDP, Inflation and REER to a shock in oil revenue growth for Algeria, Angola and Egypt (excluding Nigeria) using the HPVAR model

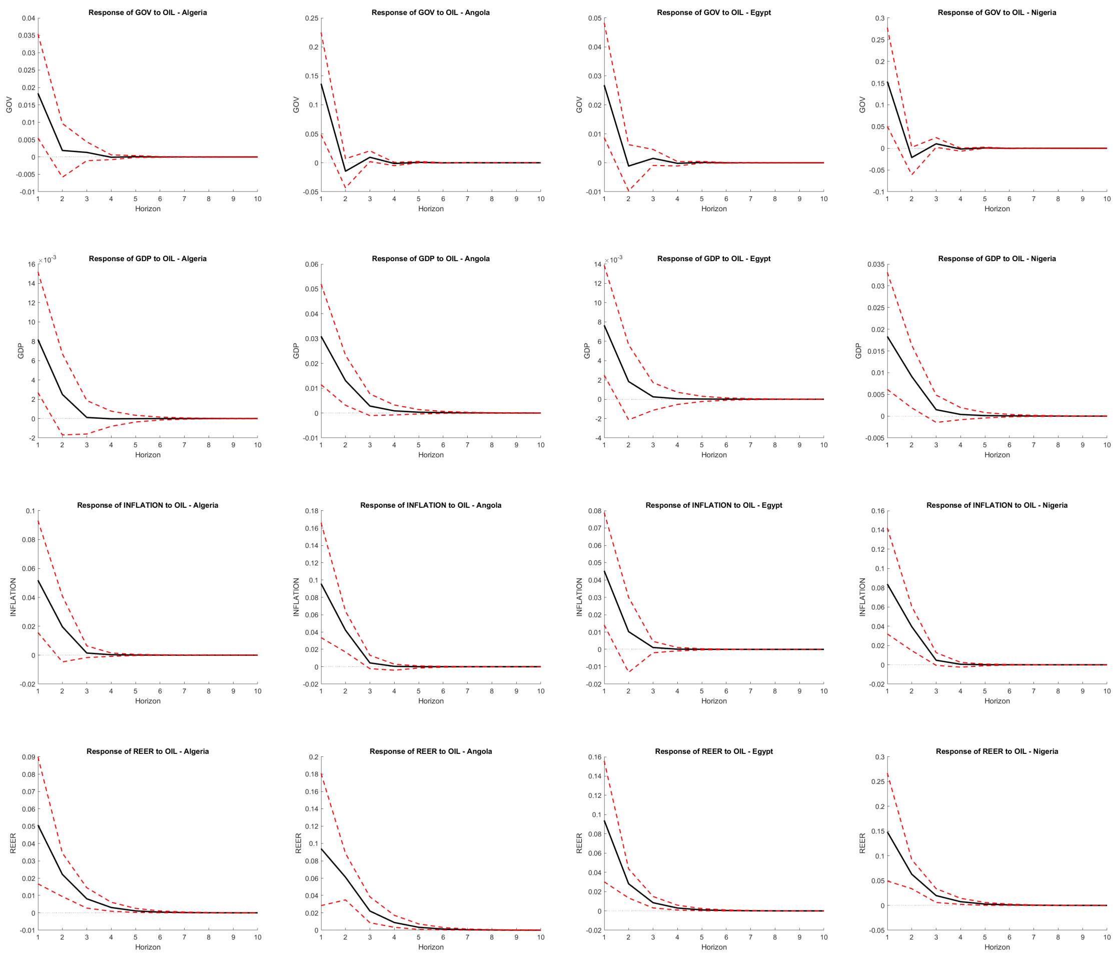


Figure B.12: Impulse responses of government expenditure, GDP, Inflation and REER to a shock in oil revenue (P $\times$ Q) growth for Algeria, Angola, Egypt and Nigeria using the HPVAR model

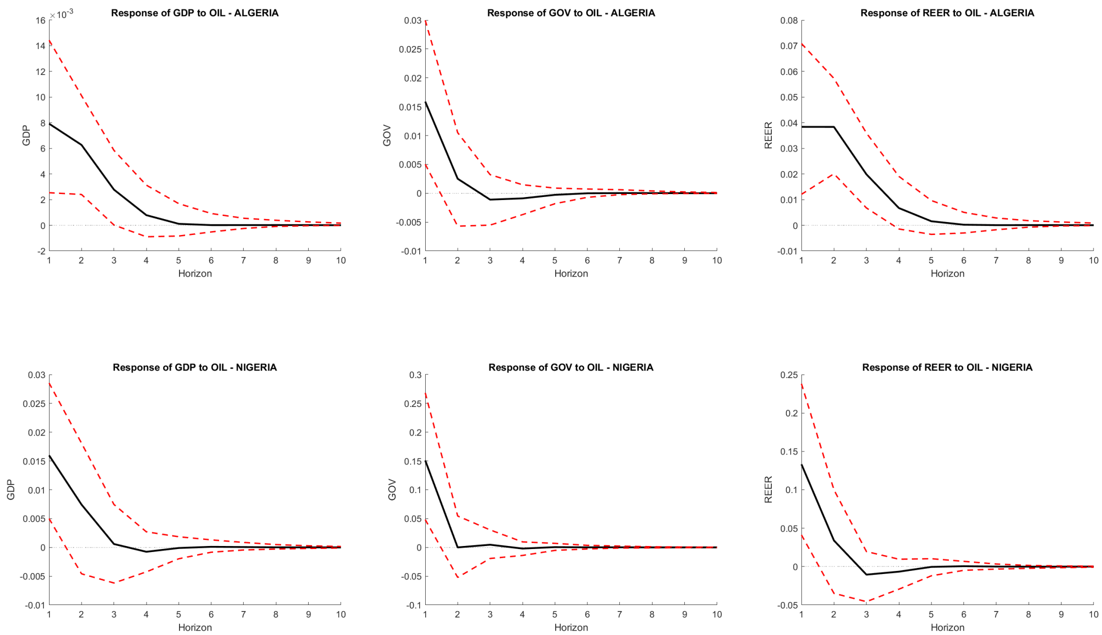


Figure B.13: Impulse responses of government expenditure, GDP, and REER to a shock in oil revenue for Algeria and Nigeria (with World Bank REER data) using the SVAR model

B.4 IRFs of Government Expenditure and the REER to Oil Revenue Shock with an Alternative Specification

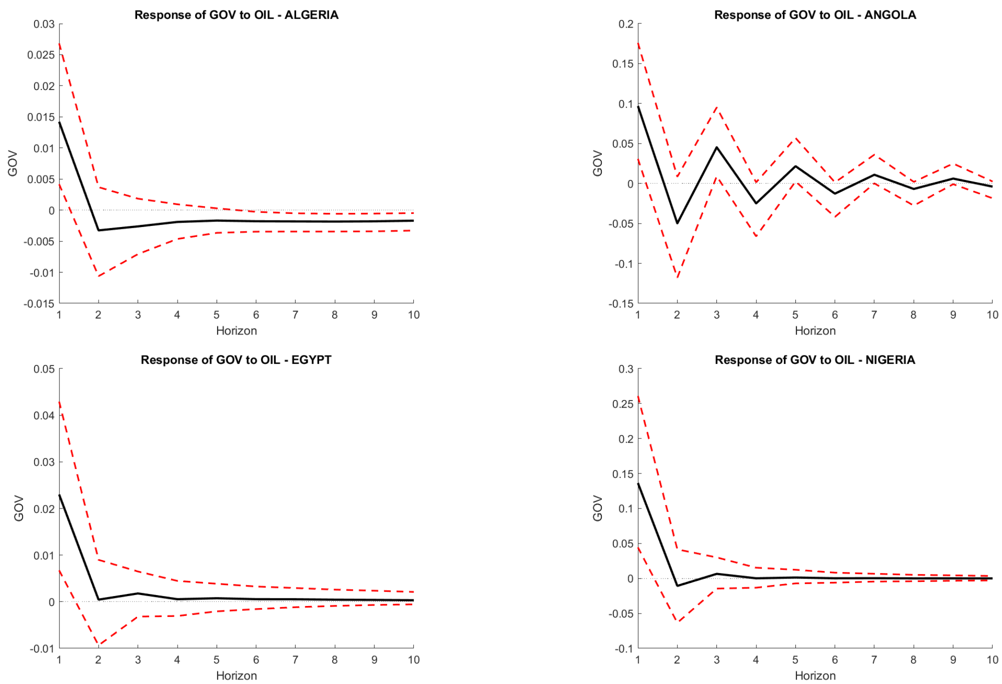


Figure B.14: Impulse responses of government expenditure to a shock in oil revenue, with Oil Rents (% of GDP) and the REER measured in levels

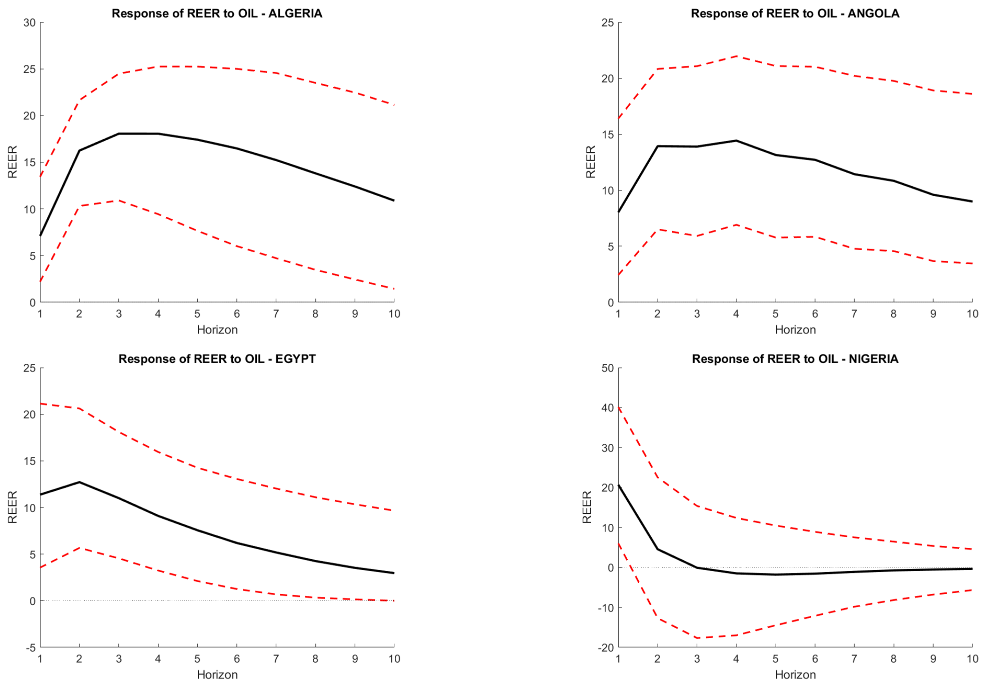


Figure B.15: Impulse responses of the real effective exchange rate to a shock in oil revenue, with Oil Rents (% of GDP) and the REER measured in levels

B.5 Bayesian Estimation

B.5.1 Fixed-coefficient SVAR

The likelihood is given as:

$$f(y|\beta,\Sigma) \propto |\det(\Sigma)|^{-\frac{T}{2}} \times \exp[-\frac{1}{2}(\beta-\hat{\beta})'(\Sigma\otimes(X'X)^{-1})^{-1}(\beta-\hat{\beta})] \\ \times exp[-\frac{1}{2}tr\{\Sigma^{-1}(Y-X\hat{B})'(Y-X\hat{B})\}] \tag{B.1}$$

The prior distribution for β is assumed to follow a multivariate normal distribution with mean β_0 and an arbitrary covariance matrix Ω_0. So we have,

$$\beta \sim N(\beta_0, \Omega_0) \tag{B.2}$$

Thus, the prior density for β is given by:

$$\pi(\beta) \propto exp[-\frac{1}{2}(\beta-\beta_0)'\Omega_0^{-1}(\beta-\beta_0)] \tag{B.3}$$

The prior distribution for Σ is an inverse Wishart distribution, with scale matrix S_0 and degrees of freedom α_0:

$$\sum \sim IW(S_0, \alpha_0) \tag{B.4}$$

Consistently, the prior density of Σ is given by:

$$\pi(\Sigma) \propto \det(\Sigma)^{-\frac{(\alpha_0+n+1)}{2}} \times exp\{-\frac{1}{2}tr[\Sigma^{-1}\S_0]\} \tag{B.5}$$

The kernel of a multivariate distribution with mean β and variance-covariance matrix Ω is given by:

$$\pi(\beta|\Sigma, y) \propto \exp[-\frac{1}{2}(\beta - \hat{\beta})'\bar{\Omega}^{-1}(\beta - \bar{\beta})] \tag{B.6}$$

where:

$$\bar{\Omega} = [\Omega_0^{-1} + \Sigma^{-1} \otimes X'X]^{-1} \tag{B.7}$$

$$\bar{\beta} = \bar{\Omega}[\Omega_0^{-1}\beta_0 + (\Sigma^{-1} \otimes X')y] \tag{B.8}$$

Hence:

$$\pi(\beta|\Sigma, y) \sim N(\bar{\beta}, \bar{\Omega}) \tag{B.9}$$

The conditional distribution for $\pi(\Sigma)$ is then determined to be:

$$\begin{aligned} \pi(y|\beta, \Sigma) \propto & \,|\det(\Sigma)|^{-\frac{T}{2}} \\ & \times exp\left\{-\frac{1}{2}tr\left[\Sigma^{-1}[(Y - X\hat{B})'(Y - X\hat{B}) + S_0]\right]\right\} \end{aligned} \tag{B.10}$$

And the kernel of an inverse Wishart distribution is:

$$\pi(\Sigma|\beta, y) \propto IW(\hat{S}, \hat{\alpha}) \tag{B.11}$$

With equation (B.12) as the scale matrix and equation (B.13) as the degrees of freedom:

$$\hat{S} = (Y - XB)'(Y - XB) + S_0 \tag{B.12}$$

$$\hat{\alpha} = T + \alpha_0 \tag{B.13}$$

Obtaining these conditional distributions allow for the use of Gibbs sampling to obtain random draws from the unconditional posterior distributions of the parameters of interest.

B.5.2 Hierarchical Panel VAR

The likelihood is given as:

$$f(y|\beta,\Sigma) \propto \prod_{i=1}^{N} |\bar{\Sigma}_i|^{1/2} \exp\left(-\frac{1}{2}\big(y_i - \bar{X}_i\beta_i\big)'\big(\bar{\Sigma}_i\big)^{-1}\big(y_i - \bar{X}_i\beta_i\big)\right) \tag{B.14}$$

The vectors of coefficients β_i follow a normal distribution, with a common mean b and common variance Σ_b: $\beta_i \sim \mathcal{N}(b, \Sigma_b)$

The prior density for β is given by:

$$\pi(\beta|b,\Sigma_b) \propto \prod_{i=1}^{N} |\Sigma_b|^{-1/2} \exp\left(-\frac{1}{2}(\beta_i - b)'(\Sigma_b)^{-1}(\beta_i - b)\right) \tag{B.15}$$

The prior distribution for Σ_i is simply the classical diffuse prior given by: $\pi(\Sigma_i) \propto |\Sigma_i|^{(n+1)/2}$

Thus the prior full density for Σ is given by:

$$\pi(\Sigma_i) \propto \prod_{i-1}^{N} |\Sigma_i|^{(n+1)/2} \tag{B.16}$$

For the hyperparameters b and Σ_b, the assumed hyperpriors are the following. For b, the selected form is simply a diffuse (improper) prior: $\pi(b) \propto 1$

For Σ_b, the adopted functional form is designed to replicate the VAR coefficient covariance matrix of the Minnesota prior. It relies on a covariance matrix Ω_b. The full covariance matrix is then defined as:

$$\Sigma_b = (\lambda_1 \otimes I_q)\Omega_b \tag{B.17}$$

$(\lambda_1 \otimes I_q)$ is a diagonal matrix with all its diagonal entries being equal to λ_1. Considering Ω_b (the coefficient covariance matrix) as fixed and known, but treating λ_1 as a random variable conveniently reduces the determination of the full prior for Σ_b to the determination of the prior for the single parameter λ_1.

In-between values for λ_1 imply some degree of information sharing between units, and ideally λ_1 should be designed to provide a good balance between individual and pooled estimates. To achieve this result, a we use a prior distribution of λ_1 that is an inverse Gamma distribution with shape $s_0/2$ and scale $\upsilon_0/2$: $\lambda_1 \sim IG(s_0/2, \upsilon_0/2)$

Thus:

$$\pi(\lambda_1|s_0/2, \upsilon_0/2) \propto \lambda^{-\frac{s_0}{2}-1} \exp\left(-\frac{\upsilon_0}{2\lambda_1}\right) \tag{B.18}$$

To avoid having the results being sensitive to the choice of values for $s_0/2$ and $\upsilon_0/2$ as shown in Jarocinski (2010b) and Gelman (2006), we make the priors weakly informative ($s_0, \upsilon_0 = 0.001$).

The full posterior distribution is equal to the product of the data likelihood function in (B.14) with the conditional prior distribution for β in (B.15), the prior for Σ in (B.16) and the hyperpriors $\pi(b)$ and $\pi(\Sigma_b)$

$$\pi(\beta, b, \Sigma_b, \Sigma|y) \propto \pi(y|\beta, \Sigma)\pi(\beta|b, \Sigma_b)\pi(b)\pi(\Sigma_b)\pi(\Sigma) \tag{B.19}$$

The conditional distributions can be obtained by considering the joint posterior distribution for all parameters and retaining only terms involving β, Σ, b and Σ_b. Dieppe et al. (2016) relegate terms that do not involve these parameters to the proportionality constant, as they do not contain information about their distribution.

The conditional posterior for β_i becomes:

$$\pi(\beta_i|\beta_{-i}, y, b, \Sigma_b, \Sigma) \propto \pi(y|\beta_i, \Sigma)\pi(\beta_i|b, \Sigma_b) \tag{B.20}$$

where β_{-i} denotes the set of all β coefficients less β_i. The distribution is conditionally multivariate normal:

$$\pi(\beta_i|\beta_{-i}, y, b, \Sigma_b, \Sigma) \sim \mathcal{N}(\bar{\beta}_i, \bar{\Omega}_i) \tag{B.21}$$

as derived in Dieppe et al. (2016).

The conditional distribution for b is given as:

$$\pi(b|y, \beta, \Sigma_b, \Sigma) \propto \exp\left(-\frac{1}{2}(b - \beta_m)'(N^{-1}\Sigma_b)^{-1}(b - \beta_m)\right) \tag{B.22}$$

where $\beta_m = N^{-1}\Sigma_{i=1}^{N}\beta_i$ is the arithmetic mean over the β_i vectors. It follows a multivariate normal distribution:

$$\pi(b|y, \beta, \Sigma_b, \Sigma) \sim \mathcal{N}(\beta_m, N^{-1}\Sigma_b) \tag{B.23}$$

The conditional posterior for Σ_b:

$$\pi(\Sigma_b|y, \beta, b, \Sigma) \propto \lambda_1^{-\frac{\bar{s}}{2}-1} exp\left(-\frac{\bar{\upsilon}}{2}\frac{1}{\lambda_1}\right) \tag{B.24}$$

This is the kernel of a inverse Gamma distribution with shape $\frac{\bar{s}}{2}$ and scale $\frac{\bar{\upsilon}}{2}$:

$$\pi(\Sigma_b|y, \beta, b, \Sigma) \sim IG\left(\frac{\bar{s}}{2}, \frac{\bar{\upsilon}}{2}\right) \tag{B.25}$$

The conditional posterior distribution for the set of residual covariance matrices $\Sigma = \Sigma_1, \Sigma_2, \ldots, \Sigma_N$:

$$\pi(\Sigma_i|\Sigma_{-i}, y, \beta, b, \Sigma_b) \propto |\Sigma_i|^{-(T+n+1)/2} \exp\left(-\frac{1}{2} tr[\Sigma_i^{-1}\tilde{S}_i]\right) \tag{B.26}$$

This is the kernel of an inverse Wishart distribution with scale $\tilde{S}_i$ and degrees of freedom T:

$$\pi(\Sigma_i|\Sigma_{-i}, y, \beta, b, \Sigma_b) \sim IW\big(\tilde{S}_i, T\big) \tag{B.27}$$

B.5.3 Implementation of Sign and Zero Restrictions

Identification of structural shocks in both the SVAR and HPVAR models is conducted using sign and zero restrictions within a Bayesian VAR framework, following the methodology of Arias et al. (2018) as implemented in the BEAR toolbox.

Estimation proceeds by drawing the reduced-form VAR coefficients and the residual covariance matrix Σ from their joint posterior distribution. Conditional on these draws, reduced-form impulse response functions are computed. Structural impulse response functions are then obtained through an orthogonalization step that maps reduced-form innovations into mutually orthogonal structural shocks with unit variance.

Specifically, for each posterior draw of Σ, a preliminary factorization $h(\Sigma)$ satisfying

$$h(\Sigma)h(\Sigma)' = \Sigma$$

is obtained. Structural representations are generated by post-multiplying this factor by an orthogonal rotation matrix Q, such that the resulting structural impact matrix is given by

$$D = h(\Sigma)Q.$$

The orthogonal matrix Q is drawn to be uniformly distributed over the space of orthonormal matrices. In practice, this is achieved by drawing a random matrix with independent standard normal entries and applying a QR decomposition to obtain an orthogonal rotation.

Given D, structural impulse response functions are computed at the horizons specified by the identification scheme. Sign and zero restrictions are imposed directly on these impulse responses using selection matrices that isolate the restricted variables, shocks, and horizons. For each candidate draw, the restrictions are checked jointly across all specified horizons. Draws that do not satisfy all restrictions are discarded, while admissible draws are retained. This accept–reject procedure is repeated until a sufficient number of posterior draws satisfying the identifying restrictions is obtained.

When zero restrictions are included, the BEAR implementation follows the sequential construction of the orthogonal matrix Q described in Arias et al. (2018), ensuring that zero restrictions are satisfied exactly by construction prior to checking the remaining sign restrictions. This approach guarantees that posterior inference is conducted over the correct conditional distribution of the structural parameters.

The same identification strategy is applied in both the single-country SVAR and hierarchical panel VAR settings. In the HPVAR case, posterior draws additionally reflect the hierarchical prior structure across countries, but the construction of admissible orthogonal rotations and the enforcement of sign and zero restrictions follow the same algorithmic steps.

Full technical details of the algorithm are provided in Arias et al. (2018) and in the BEAR technical documentation. As the present paper relies on the standard BEAR implementation without modification, these details are not reproduced further here.

Appendix C

Appendix to Chapter Three

C.1 Pre-estimation Tests

C.1.1 Unit-root Tests

Table C.1: Augmented Dickey–Fuller Tests with Drift (lags = 4)

	Output		Spending		Investment		Inflation		Interest Rate		EPU_lag	
Country	t	p	t	p	t	p	t	p	t	p	t	p
Australia	-3.884	0.0001	-3.626	0.0003	-3.049	0.0016	-4.119	0.0000	-4.689	0.0000	-3.550	0.0003
Canada	-3.214	0.0009	-2.931	0.0022	-3.917	0.0001	-4.410	0.0000	-4.356	0.0000	-3.978	0.0001
France	-3.374	0.0006	-5.380	0.0000	-4.089	0.0001	-3.132	0.0012	-4.752	0.0000	-3.558	0.0003
Germany	-4.346	0.0000	-2.734	0.0039	-4.804	0.0000	-4.349	0.0000	-4.838	0.0000	-3.826	0.0001
Korea	-5.746	0.0000	-3.909	0.0001	-7.149	0.0000	-5.130	0.0000	-6.010	0.0000	-4.199	0.0000
United States	-3.874	0.0001	-2.953	0.0021	-4.716	0.0000	-5.495	0.0000	-4.680	0.0000	-3.752	0.0002
United Kingdom	-3.406	0.0005	-2.899	0.0024	-3.435	0.0005	-3.128	0.0012	-3.345	0.0006	-3.420	0.0005

Note: Results from country-by-country Augmented Dickey–Fuller tests with an intercept (drift) and four lags. The null hypothesis is a unit root (random walk with drift). All p-values are MacKinnon approximate p-values.

Table C.2: Panel Unit-Root Tests (IPS and Fisher-ADF, 4 lags)

	IPS Test		Fisher-ADF Test	
Variable	W-t-bar	p-value	Z-statistic	p-value
Output	-10.8198	0.0000	-7.7688	0.0000
Spending	-5.8349	0.0000	-6.2678	0.0000
Investment	-8.5856	0.0000	-9.0788	0.0000
Inflation	-9.9663	0.0000	-8.5810	0.0000
Interest Rate	-8.2704	0.0000	-9.8070	0.0000
EPU_lag	-9.3919	0.0000	-7.1530	0.0000

Notes: The table reports Im–Pesaran–Shin (IPS) and Fisher-type panel unit-root tests based on augmented Dickey–Fuller regressions with four lags. The null hypothesis in both tests is that all panels contain unit roots. Panel means are included and cross-sectional means are removed. All reported p-values indicate rejection of the unit-root null at the 1% significance level.

C.1.2 Lag Order Selection

Table C.3: Panel VAR Lag Selection Criteria

Lag	MBIC	MAIC	MQIC
1	-396.3541	-69.9069	-197.2372
2	-283.4472	-65.8158	-150.7027
3	-144.9922	-36.1765	-78.6199

Note: MBIC = Modified Bayesian Information Criterion; MAIC = Modified Akaike Information Criterion; MQIC = Modified Hannan–Quinn Criterion. All criteria favor a parsimonious specification with one lag, which balances model fit and parameter parsimony while adequately capturing short-run dynamics.

C.2 Threshold Cutoffs

Table C.4: Country-Specific Percentile Cutoffs of the EPU Cycle

Country	80th Percentile	85 Percentile
Australia	0.3300	0.4261
Canada	0.1983	0.3381
France	0.2512	0.3396
Germany	0.2069	0.2815
Korea	0.2678	0.3554
United Kingdom	0.2790	0.3494
United States	0.2215	0.2620

Note: The table reports country-specific 80th and 85th percentile thresholds of the lagged cyclical component of the Economic Policy Uncertainty (EPU) index. High-uncertainty regimes in robustness specifications are defined as quarters in which EPU_{t-1}^{cyc} exceeds the respective country-specific percentile cutoff.

Table C.5: High-Uncertainty Quarters Coinciding with the Global Financial Crisis

Country	High-Uncertainty Quarters	During GFC	Share (%)
Australia	13	3	23.1
Canada	13	2	15.4
France	13	2	15.4
Germany	13	1	7.7
Korea	13	1	7.7
United Kingdom	13	2	15.4
United States	13	4	30.8

C.3 Robustness Checks

C.3.1 Global EPU Index

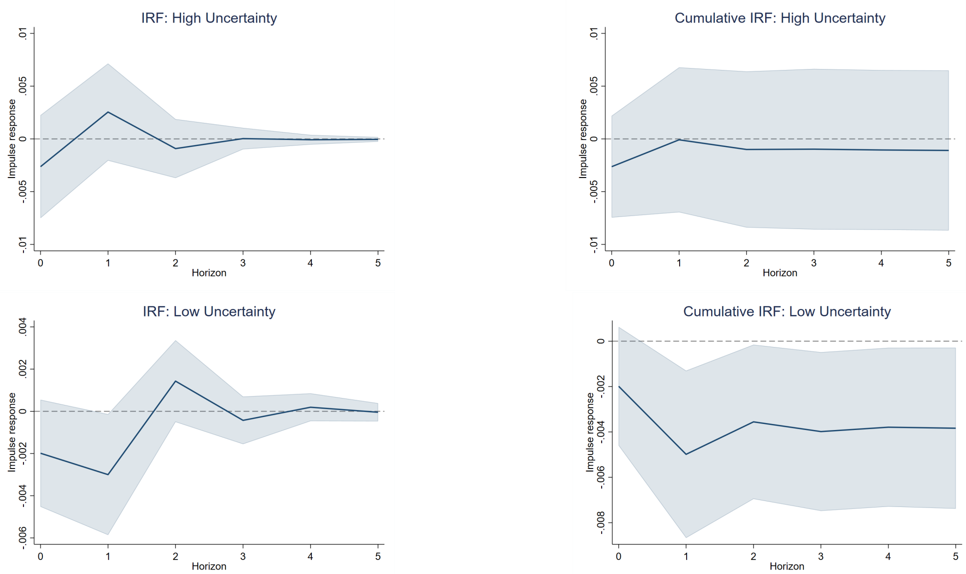


Figure C.1: IRFs and Cumulative IRFs of business investment growth to a 1-unit shock in Government Expenditure growth, using the Global EPU Index

C.3.2 Unfiltered EPU Index

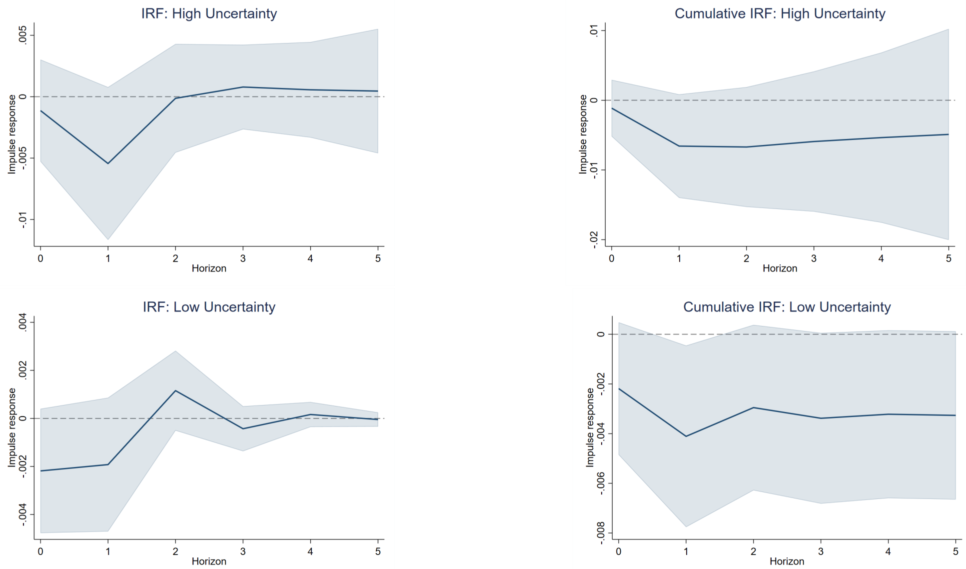


Figure C.2: IRFs and Cumulative IRFs of business investment growth to a 1-unit shock in Government Expenditure growth, using the EPU Index in levels

C.3.3 Country-specific Thresholds using 80th Percentile

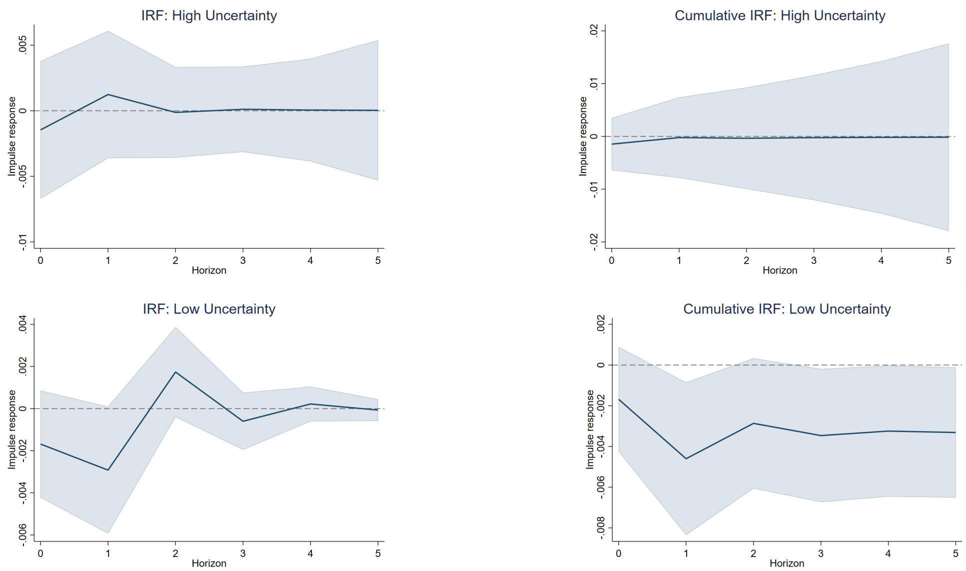


Figure C.3: IRFs and Cumulative IRFs of business investment growth to a 1-unit shock in Government Spending growth, using the country-specific 80th percentile uncertainty threshold

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