

Essays on fiscal deficit, debt and monetary policy: A nonlinear approach

by

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B.S.S., Shah Jalal University of Science and Technology, Bangladesh, 2002

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AN ABSTRACT OF A DISSERTATION

submitted in partial fulfillment of the  
requirements for the degree

DOCTOR OF PHILOSOPHY

Department of Economics  
College of Arts and Sciences

KANSAS STATE UNIVERSITY  
Manhattan, Kansas

2017

# Abstract

This essay empirically investigates the dynamics between government debt and budget deficits in the United States during a recession as opposed to an expansion. We use four different budget deficits definitions to develop a more comprehensive insight. We estimate a threshold VAR model on quarterly data from 1947: Q1 to 2016: Q3 on debt to GDP and budget deficits to GDP ratio for the United States. Specification test using LR test rejects the null for a linear VAR against nonlinear VAR. The nonlinear impulse responses indicate, with an increase to budget deficits to GDP ratio, government debt to GDP ratio rise faster during a recession as opposed to an expansion, and tend to move in a counter-cyclical manner with an increase in the output gap. We can thus infer that governments chose economic stability over fiscal balance during recessions. With an increase in government debt to GDP ratio, nonlinear impulse response show budget deficits to GDP ratio grow faster during an expansion as opposed to a recession and exhibit counter-cyclicity with an increase in the output gap. All four budget deficits definitions depict similar pattern. Robustness check, using cyclically adjusted primary budget deficit published by the congressional Budget Office, also confirm the above findings.

In this essay, we explore the presence of a long run relationship between the monetary base and the government debt using monthly data from 1942:1 to 2015:12. We apply formal statistical methods including cointegration and threshold cointegration tests to investigate the presence of a long-run relationship and estimate a threshold vector error-correction model (TVECM henceforth) to analyze the short-run dynamics. We find the presence of a threshold cointegration between the monetary base and government debt. As for the short-run dynamics, TVECM estimates show that the speed of adjustment is significant for the growth in debt equation in both regimes with the signs indicating government adjusting the debt in the

short-run. But the U.S. Fed does not change the monetary base, hence we do not find any evidence of debt monetization in the U.S. We evaluate our findings over two sub-samples: 1946 to 2015 and 1946 to 2007 for robustness purposes. Findings from both sub-samples conform to our findings from the full sample.

In this essay, we investigate the impacts of growth in the budget deficit and money supply on real interest rate are integral to contemporary macroeconomic policy. We employ threshold VAR and nonlinear impulse responses using quarterly data from 1959 to 2015. We find that growth in money supply and budget deficits have an asymmetric impact on inflation, short-term interest rate, and real interest rates. Growth in money supply and budget deficits tend to make the real interest rate negative in a bad state. In a good state, on the other hand, growth in money supply tend to increase the real interest rate but growth in budget deficits tend to decrease the real interest rate over the forecast horizon.

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# Acknowledgments

First of all, blessed be the name of Allah, thy Lord, full of majesty, bounty, and honor.

I would like to extend my sincere gratitude to all my faculty members at the Department of Economics at Kansas State University. The formal academic training I received from our faculty members has enriched my understanding of economics and enhanced my research capabilities to accomplish this dissertation research. I am particularly grateful to Dr. Steven P. Cassou for guiding me personally and professionally. All the chapters in this dissertation are the result of your thorough and meticulous guidance over the past three years. I would like to thank Dr. Lance Bachmeier for his guidance with estimation and coding help. I would also like to thank the faculty members who gave me input during the graduate student research workshops.

I would like to thank my spouse Jamil Sharmin for being patient with me and supporting me in achieving this degree. You are a blessing in my life; without your gracious support, I would not be able to reach this far. You are my courage and my strength.

I would like to thank my parents Mr. Kaosar Ahmed Haydory and Mrs. Rokeya Haydory, my parents-in-law Mr. Ruhul Amin, and Mrs. Ishrat Naaz, and my children Ahraz Ahmed Haydory and Zayna Ayesha Haydory. You people are my inspiration.

I would like to thank my sisters and their family, my friends and all the well wishers' for your continuous support and blessing. Thanks a lot, everyone.

# Dedication

To my Family.

# Chapter 1

## Dynamics between government debt and budget deficits in the United States: a threshold VAR analysis

### 1.1 Introduction

Budget deficits, resulting either from a tax cut or an increase in government spending, will impact government debt. Rising government debt may subsequently shape future budget balance through policies aimed at deficit reduction. Government authorities justify austerity and deficit reduction by arguing for sustainable fiscal policy, which will prevent further debt accumulation, and therefore put less pressure on monetary policy ([Uctum and Wickens, 2000](#)). In recent times, the expansionary fiscal policy was used to curb the depressing consequences of the Great Recession resulting in increased budget deficits, subsequently raising government debt in the United States. The post-WWII period has seen several recessions in the United States, for which the expansionary fiscal policy tools were adopted. During which tax cuts, and/or, spending increase were pursued to stabilize the economy. Political conflicts such as the Korean War, the Cold War, and post- 9/11 era are marked by rising defense spending. We also observe an increasing trend in non-discretionary government spending

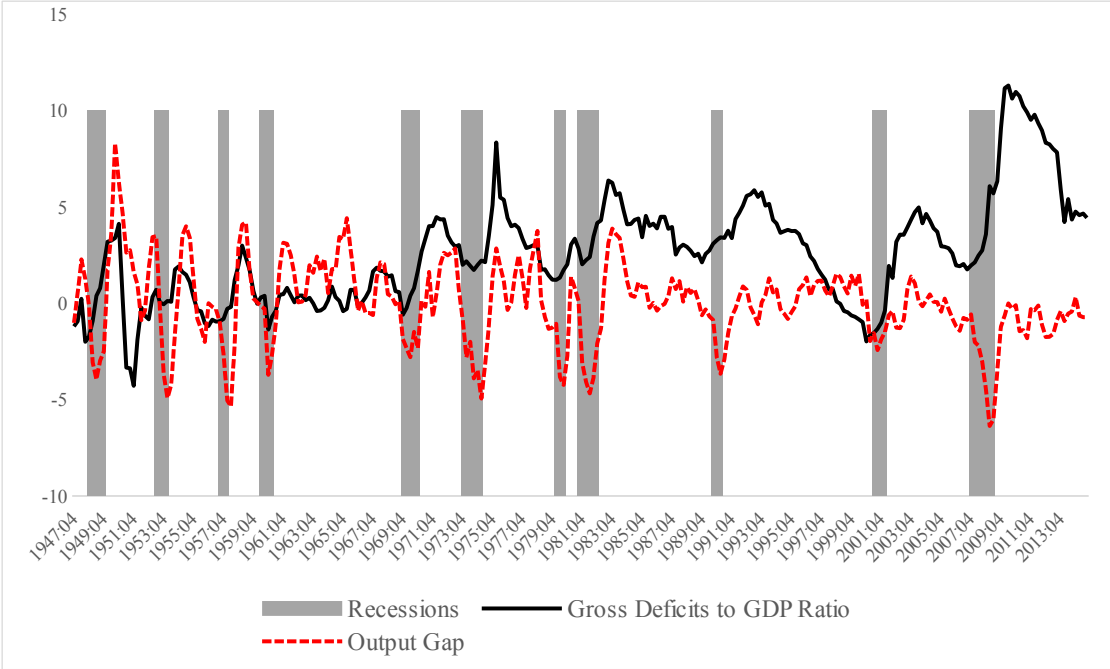
since the late 1980s. All these factors jointly contributed towards increases in the budget deficit. Subsequently, government debt in the United States has grown over time along with budget deficits ([Thornton, 2012](#)). Economists have investigated the sustainability of the fiscal policy with rising debt ([Davig, 2005](#); [Quintos, 1995](#)). [Bohn \(1998, 2007\)](#) uses a policy reaction function interpretation to investigate government debt sustainability in the United States. He argues that historically corrective measures were implemented to respond to government debt whenever the debt to GDP ratio is increasing. Sustainability of debt and budget deficits is well researched in the existing literature. A political economy approach explains the budget deficits and debt as a consequence of political competition between the incumbent and successor, where the incumbent prefers debt and budget deficits to limit government spending choices for the successor ([Alesina and Tabellini, 1990](#); [Persson and Svensson, 1989](#)). However, is the relationship between the government debt and the budget deficits asymmetric in nature? Do we observe cyclicalities? [Cassou et al. \(2015\)](#) investigate asymmetry between the government debt and the primary budget deficit using a structure analogous to [Bohn \(1998, 2007\)](#), and find asymmetry for sustainability as well as cyclicalities of the budget deficit in the United States.

An expansionary fiscal policy is prescribed during recessions to stabilize the economy. Policy makers generally are more concerned to stabilize the economy during economic recessions. As such we observe expansionary fiscal policies are designed and implemented, with either an increase in the government spending or tax cuts, during recessions. Thus we observe an increase in the budget deficit leading to more government debt during recessions. We do not observe the use of expansionary fiscal policy during economic expansions. Hence, we can argue that government's policy response during economic recessions and expansions are nonlinear in nature. Figure 1.1 below presents primary budget deficit to GDP ratio and the output gap for the United States from 1947 to 2016. The shaded areas represent recessions using the National Bureau of Economic Research business cycle dates. We observe the primary budget deficit to GDP ratio to rise during recessions. At the same time, there are several periods where the deficit to GDP ratio increases even though the economy was



expanding. The business cycle fluctuations motivate the government’s policy choice between economic stabilization versus balancing the budget. Analyzing the nature of the dynamics of the budget deficit and the government debt along with business cycle fluctuations is thus imperative. In recent times during the Great Recession, we observed European governments with large debt burdens tend to go for austerity measures at times of economic stress, despite the usual policy prescription that economic stabilization needs fiscal expansion. The nature of the variations in policy choice adopted during a recession as opposed to an expansion, either pro-cyclical or counter-cyclical, can help us infer about the government’s policy choices.

Figure 1.1: Budget Deficit to GDP ration and the Output Gap



In the existing literature, it remains unanswered how the budget deficits impact the government debt during recessions as opposed to expansions, and vice-versa. We investigate how the budget deficit impacts the dynamics of the government debt during recessions as well as during expansions, and vice-versa. This essay augments the existing literature in a number of ways. First, we use four different measure of budget deficits, which provide us a more comprehensive picture of the budget balances. Second, we analyze the dynamics of

the government debt and the budget deficits during recessions and expansions in a more objective manner. As such we use the regime-switching threshold vector auto-regression (henceforth TVAR) model to derive the impulse response, which enables us to understand the dynamics during a recession as opposed to an expansion. Finally, our sample covers the whole of the post-WWII period from 1947: Q1 to 2016: Q3, which includes the Great Recession. The non-linear impulse responses indicate, with an increase in budget deficits to GDP ratio, the government debt tends to rise faster during the recessions vis-à-vis the expansions, and tend to grow in a counter-cyclical manner with an increase in the output gap. We can thus infer that governments chose economic stability over fiscal balance during recessions. However, with an increase to government debt to GDP ratio, budget deficits tend to grow faster during the expansions vis-à-vis the recessions, and exhibit counter-cyclicity with an increase in the output gap.

The rest of the paper is organized as follows: Section 1.2 discusses the relevant literature, in Section 1.3 we discuss the data and methodology, and Section 1.4 presents the results, and Section 1.5 discusses the conclusion.

## 1.2 Literature Review

Research investigations on the government debt and the budget deficits vary in their approach. However, the literature on the dynamics of the government debt and the budget deficits focusing on asymmetry and cyclicity are sparse. The issue of debt sustainability is well researched; we can find single country studies using the time-series econometric techniques as well as cross-country evidence based on the panel data models. There are papers that use the theoretical structural models to analyze numerous aspects along with sustainability related to the budget deficit and debt. The question of debt sustainability is a key aspect of the existing literature. Empirical studies based on time-series techniques focus on stationarity and cointegration between the budget deficit and the debt. This approach

implies that the deficit and the debt are sustainable as they satisfy certain statistical properties. The theoretical papers investigate how larger debt could potentially impact future tax structure or economic growth. The economic question of debt sustainability is often loosely defined or remain vague.

There are two approaches that dominate the literature analyzing the sustainability of government debt and budget deficits in the literature using the time-series techniques. The first consists of testing the stationarity of the debt and/or the budget deficits. Other studies look for a cointegrating relationship linking the primary budget deficits, the stock of outstanding debt and interest payments for the United States. Results vary with the specification of the budget constraint used in the research. [Barro \(1986\)](#) investigate the relationship between debt and budget deficits while addressing cyclicity in the econometric specification. He reports a strong counter-cyclical behavior of budget deficits in the United States. [Hamilton and Flavin \(1985\)](#) reject the non-stationarity of constant-dollar undiscounted government debt in the United States under the assumption of constant real interest rates. [Wilcox \(1989\)](#) allows for stochastic interest rates and trends, finding discounted government debt in the United States is non-stationary. [Hamilton and Flavin \(1985\)](#), [Trehan and Walsh \(1988\)](#), [Quintos \(1995\)](#), and [Davig \(2005\)](#) find evidence in support of sustainability. In contrast, [Kremers \(1989\)](#) and [Hakkio and Rush \(1991\)](#) show that in recent years fiscal policy violates the inter-temporal budget constraint. In particular, [Kremers \(1989\)](#) showed the empirical finding of [Hamilton and Flavin \(1985\)](#) is reversed once their ADF regression for the budget balance, which suffers from autocorrelation, is extended to include the second lagged difference of the budget balance as an additional regressor in order to eliminate the autocorrelation problem. [Ahmed and Rogers \(1995\)](#), using historical data that goes back to 1700s, find strong evidence favoring the sustainability of the fiscal policy in the United States and some support for the sustainability of the United Kingdom. [Arestis et al. \(2004\)](#) find that large budget deficits in the United States are sustainable in the long run as policy makers will intervene to reduce per capita deficit when it reaches a certain threshold. They use threshold autoregression on data spanning from 1947 to 2002, where the regression equa-

tion stems from the intertemporal budget constraint of the government.

[Bohn \(2007\)](#) argues that using stationarity and cointegration methodology in analyzing the sustainability of budget deficit and government debt is not the valid approach. [Cassou et al. \(2015\)](#) investigate the sustainability and cyclical asymmetry in the structure analogous to [Bohn \(1998\)](#). They use three different model specifications over a short and a long sample (this includes the Great Recession) to investigate asymmetry between primary budget deficits and debt in the United States. They find fiscal policy is asymmetric for the long sample, as the response of primary budget deficits to lagged government debt is asymmetric. But the asymmetry disappears for the short sample. They find evidence in support of fiscal sustainability for the short sample but this disappears for the long sample.

Fiscal rules, such as the stability and growth pact (SGP), aim at constraining government spending. Governments revert to creative accounting to circumvent such rules and in the process hide budget deficits. [Von Hagen and Wolff \(2006\)](#) provide empirical evidence of creative accounting in the European Union countries. Their two-stage IV regression in a dynamic panel setup find the tendency to substitute budget deficits with creative accounting is especially strong for cyclical component of the deficit. In times of recession, interest in reducing the budget deficit is particularly large. [Hatzinikolaou and Simos \(2013\)](#) suggest a new test for sustainability of budget deficits. In particular, they define sustainability by requiring formally that both the intertemporal budget constraint is satisfied (i.e. the discounted debt converges to zero) and the undiscounted debt is bounded. According to this more restrictive definition, [Hatzinikolaou and Simos \(2013\)](#) found empirical evidence against the sustainability of budget deficits in the United States. [Balassone et al. \(2010\)](#) uses a panel for 14 European Union countries and finds cyclical asymmetry between budget deficits and government debt. [Mahdavi \(2014\)](#) uses a panel data model on data comprised from 48 U.S. states to investigate the state fiscal sustainability using the framework proposed by [Bohn \(1998\)](#). He finds the positive response of primary budget surpluses (negative primary budget

deficits) to government debt is the sufficient condition for sustainability.

There are theoretical papers that also address the linkage between debt and budget deficits, explore their impacts on growth, as well as potential changes in the tax structure of future budgets. [Barro \(1979, 1987\)](#), [Lucas and Stokey \(1983\)](#) explain the role of budget deficits and debt accumulation to explain redistribution of income over time and across generations, as well as to explain the how these two minimize the deadweight losses of taxation for the provision of public goods. These papers, perhaps, explain the accumulation of government debt in peace times. [Uctum and Wickens \(2000\)](#) derived conditions suitable for determining the sustainability of fiscal policy in the long run, in the medium term, and in the presence of debt and budget deficits ceilings for the E.U. and the United States using data ranging from 1965-1994. Using infinite horizon models for long run they do not find evidence for sustainability for many countries. However, there is some evidence that the government discounted net debt is mean-reverting for a few countries, implying that their fiscal policies are sustainable. The evidence in favor of sustainability is strengthened for most countries when data are extended to incorporate future fiscal consolidation plans. They argue that in practice governments also need a medium-term framework for fiscal policy. [Uctum and Wickens \(2000\)](#) have shown that in the absence of ceilings most countries have sustainable fiscal policies in the medium term. This is in contrast to the result of the infinite horizon analysis. Furthermore, they find that imposing deficit or debt limits in the medium term throws most government budgets onto an inter-temporally inconsistent path. They further argue that such arbitrary ceilings confound the government debt and fiscal sustainability as they do not take account of cyclical factors.

[Bräuninger \(2005\)](#) analyze the impact of public debt on endogenous growth using an overlapping generations model setup. He assumes that government fixes the budget deficits ratio. When the budget deficits ratio is below a critical level there are two potential steady states - capital, output, and government debt all grow at a constant rate, and, increase in deficits ratio reduces the growth rate. As and when the budget deficits ratio exceeds the

critical level, there is no steady state with capital growth declining continuously. [Marcet and Scott \(2009\)](#) investigate the relationship between optimal tax and stochastic behavior of debt for the United States. Under complete market, they find, debt is the same or less persistent and it declines in response to a shock that causes deficits to rise. Under incomplete market, debt is persistent and increases in response to a shock that causes higher deficits. Authors argue that the U.S. data is more akin to incomplete market. These papers, although differ in approach and research questions, indicate that investigating dynamics between government debt and budget deficits in a linear framework may not be sufficient to develop an in-depth insight.

Budget deficits and debt are explained as strategic tools used by competing political governments in the political economy literature. These models are intended to explain the growth in budget deficits and debt in peace times when economic stabilization is unwarranted. [Alesina and Tabellini \(1990\)](#) find a bias toward budget deficits by the incumbent in a model of political competition between political actors- incumbent and future successor. The incumbent uses debt accumulation strategically to influence the choices of its successors. Disagreement among the political parties and uncertainty about who will get elected next period prevent the incumbent from internalizing the cost of leaving debt to its successors. The resulting equilibrium stock of public debt tends to be higher than socially optimum choice. Their results also show that government debt tends to get larger with public consumption expenditure. [Persson and Svensson \(1989\)](#) consider a model where two policy-makers have different views about the level of public expenditure. Their results show that a conservative incumbent policy-maker (one who likes less public expenditure) has a bias towards deficits to force the liberal successor to spend less. Conversely, the liberal incumbent may choose to leave a surplus to its conservative successor.

The extent of an increase in government spending or the tax cuts intended to invigorate aggregate demand during recessions in recent years in the United States render the above arguments less attractive as most of these literature do not address the state of the macroe-

economy and subsequent response from the policy makers. In Europe, on the other hand, we observe the case of austerity measures domination the policy space in recent years. The Great Recession era presents us with a divergent picture on how budget deficits and government debt interact in designing macroeconomic stabilization policy. The aforementioned literature does not fully explain the governments rationale for macroeconomic policies in contemporary times. Even though, we find a large volume of empirical literature that acknowledge the importance and significance of cyclicalilty while analyzing budget deficit and government debt. In particular, the literature thus far does not explore the dynamics between government debt and budget deficit during recessions juxtaposed against expansions. The definition of budget deficits also focuses largely on primary budget deficits. In this paper, we will explore the dynamics between government debt and budget deficits using the nonlinear time-series econometric specification that allows us to incorporate the business cycle aspects. We use four different definitions for budget deficits in the United States. Thus we expand the discussion on dynamics between government debt and budget deficits filling a gap hitherto unaddressed.

### 1.3 Data and Methodology

In this paper, we use quarterly data from 1947: Q1 to 2016: Q3 collected from the National Income and Product Account (henceforth NIPA) tables and Federal Reserve Bank of Dallas for the United States. We collected real GDP, the GDP deflator, and government spending and receipt data from the NIPA tables available in the United States Bureau of Economic Analysis website. Data on government debt are collected from Federal Reserve Bank of Dallas website.

We use the following definitions to measure budget deficits<sup>1</sup> in the United States. Primary deficits and primary current deficits show by how much government receipts fall short of the

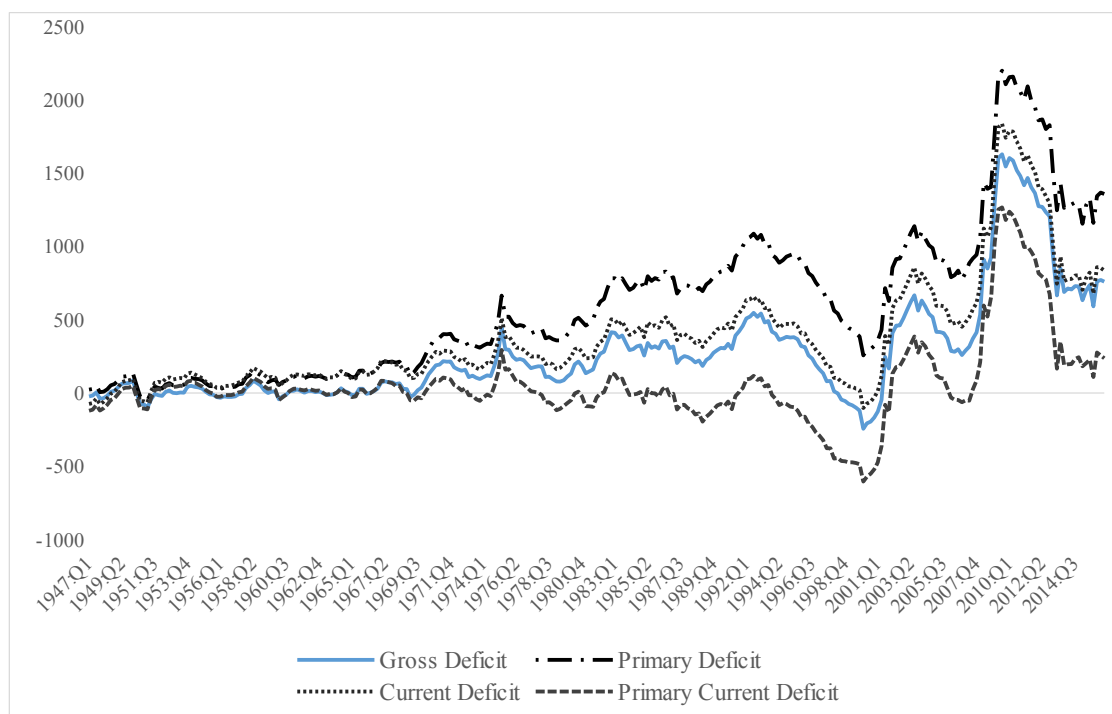
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<sup>1</sup>We use the budget deficits definition as in Macroeconomics, 9th Ed., authored by Andrew B. Abel, Ben S. Bernanke and Dean Croushore published by Pearson.

spending outlays each period. On the other hand, gross deficits and current deficits show how much the government needs to borrow each period.

- Gross Budget Deficit = Government Outlays - Government Receipts = (Government Purchase + Transfer and Subsidies + Net Interest Payment) - Government Receipts
- Primary Budget Deficit = Gross Deficit - Net Interest Payment
- Current Budget Deficit = (Gross Deficit + Government Investment Expenditure) - Government Receipts
- Primary Current Budget Deficit = Current Deficit - Net Interest Payment

Figure 1.2: Budget Deficits in the United States, 1947:Q1-2016Q3



Note: 1. Author's own calculation from the Bureau of Economic Analysis website.  
 2. All the measures are in real terms.

Figure 1.2 show the budget deficits in the United States <sup>2</sup>, as per the four definitions we are using for this paper. We observe the budget deficits show relatively less variability until the early 1970s. The mid-1970s visibly depict a decline in the budget deficits and a

<sup>2</sup>We use the GDP deflator to convert the nominal values into real terms.



similar pattern emerges during the presidency of Bill Clinton. The 1980s depict a large rise in budget deficits and also during the Great Recession years. Subsequently, we observe a continuous decline in the budget deficits during the last few years in our sample.

All the variables used are in real terms in the subsequent analysis. We use the following variables: debt to GDP ratio, budget deficits to GDP ratio<sup>3</sup> and the output gap. There are different ways to measure the output gap in the existing literature. For example, one can use the Hodrick-Prescott (HP) filter (Hodrick and Prescott, 1997). However, Mise et al. (2005) argue that such filters are not adequate to fully capture the output gap as the HP filter produces series with spurious dynamic relations that have no basis in the underlying data-generating process. Following Cassou et al. (2015) the output gap is defined as the deviation of the observed annual output growth rate from its long-term average. For the output gap, we computed the difference between the observed annual growth rate and the average growth rate over the sample period of 1947 to 2016. In particular, we computed the growth rate in percentage terms by multiplying 100 times the log difference between the current value of real GDP and the value four quarters earlier. In the next step these growth rates were averaged and then the average was subtracted from the annual growth rate series. The resulting series has positive values when the current growth rate exceeds the average and negative values when the growth rate is below the average.

In this paper, the endogenous threshold VAR is estimated following the method proposed by Tsay (1998), and, Lo and Zivot (2001) using the output gap as our threshold variable. They generalize the univariate and single equation estimation by Tong (1990), Chan (1993) and Hansen (1996) for a multivariate VAR. We will estimate an endogenous threshold VAR (TVAR) model and compute the nonlinear impulse responses to explore the dynamics between government debt and budget deficits. Why endogenous threshold VAR? Afonso et al. (2011) note that the endogenous TVAR model has a number of interesting features. First,

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<sup>3</sup>We have four different budget deficits to GDP ratio for our four different budget deficit definitions. We use each for a separate estimation.

it is a relatively simple way to capture possible nonlinearities such as asymmetric reactions to shocks or the existence of multiple equilibria. As the effects of the shocks are allowed to depend on the size and the sign of the shock, and also on the initial conditions, the impulse response functions are no longer linear, and it is possible to distinguish, for instance, between the effects of fiscal developments under different economic states. Second, another advantage of the endogenous TVAR methodology is that the variable by which different regimes are defined itself is an endogenous variable included in the VAR. Therefore, this makes it possible that regime switches may occur after the shock to each variable. In particular, fiscal expansion through a budget deficits shock might either boost the output or increase the debt that harms the prospects of economic growth, and the overall effect to the economy of a fiscal expansion through increased budget deficits might become negative.

The endogenous threshold VAR (TVAR) can be specified in the following manner, where the lag order is based on the Akaike information criterion (AIC):

$$\begin{aligned}
YG_t &= I_t \left[ \alpha_L + \sum_{i=1}^p \alpha_i DT_{t-i} + \sum_{j=1}^p \alpha_j DF_{t-j} + \sum_{k=1}^p \alpha_k YG_{t-k} \right] + \\
(1 - I_t) &\left[ \alpha_H + \sum_{i=1}^p \alpha_i DT_{t-i} + \sum_{j=1}^p \alpha_j DF_{t-j} + \sum_{k=1}^p \alpha_k YG_{t-k} \right] + \epsilon_{t,YG}
\end{aligned} \tag{1.1}$$

$$\begin{aligned}
DT_t &= I_t \left[ \beta_L + \sum_{i=1}^p \beta_i DT_{t-i} + \sum_{j=1}^p \beta_j DF_{t-j} + \sum_{k=1}^p \beta_k YG_{t-k} \right] + \\
(1 - I_t) &\left[ \beta_H + \sum_{i=1}^p \beta_i DT_{t-i} + \sum_{j=1}^p \beta_j DF_{t-j} + \sum_{k=1}^p \beta_k YG_{t-k} \right] + \epsilon_{t,DT}
\end{aligned} \tag{1.2}$$

$$\begin{aligned}
DF_t &= I_t \left[ \gamma_L + \sum_{i=1}^p \gamma_i DT_{t-i} + \sum_{j=1}^p \gamma_j DF_{t-j} + \sum_{k=1}^p \gamma_k YG_{t-k} \right] + \\
(1 - I_t) &\left[ \gamma_H + \sum_{i=1}^p \gamma_i DT_{t-i} + \sum_{j=1}^p \gamma_j DF_{t-j} + \sum_{k=1}^p \gamma_k YG_{t-k} \right] + \epsilon_{t,DF}
\end{aligned} \tag{1.3}$$

where  $YG_t$  is the measure of the output gap at time  $t$ ,  $DT_t$  is debt to GDP ratio at time  $t$ , and  $DF_t$  is deficit to GDP ratio at time  $t$ .  $\alpha$ ,  $\beta$ , and  $\gamma$  are the estimated parameters of the

model, and  $p$  represent the lag length selected using the information criterion. The output gap variable is our endogenous threshold variable such that:

$$I_t = \begin{cases} 1 & \text{if } YG_{t-d} \leq \tau \\ 0 & \text{if } YG_{t-d} > \tau \end{cases} \quad (1.4)$$

$I[\cdot]$  is an indicator function that takes the value of 1 when the output gap  $YG_{t-d}$  is less than or equal to the threshold  $\tau$ , and 0 otherwise. The time lag for the output gap is set to 1 (that is  $d=1$ ). We arrange the output gap variable in an ascending order and trim the top 15% and the bottom 15% in order to avoid over-fitting. We compute a VAR model for each of the output gap values and obtain the sum-squared of the residuals. We chose the output gap corresponding to the lowest sum-squared of the residuals as our threshold and the corresponding model as our threshold VAR model. We interpret the model below the threshold as bad state, which is assumed to be the state of recession. The model above the threshold is interpreted as good state, which we assume to be the state of expansion.

Since we are estimating a threshold VAR model it is imperative that we address the idea of a specification test. We carry out a likelihood ratio test with the null of linear VAR model against the nonlinear alternative. Since we are estimating a threshold VAR model and the threshold value is not known a priori, the testing procedure involves non-standard inference, because  $\tau$  is not identified under the null hypothesis of no threshold. Therefore, first, the TVAR model is estimated for all possible values of  $\tau$  (to avoid over-fitting, the possible values were set so that at least 15% of the observations fall under each regime). This test is the multivariate extension proposed by [Lo and Zivot \(2001\)](#) of the linearity test of [Hansen \(1996\)](#). Instead of an F-test comparing the SSR for the univariate case, a Likelihood Ratio (LR) test comparing the covariance matrix of each model is computed.

$$LR_{ij} = T \left[ \ln(\det \hat{\Sigma}_i) - \ln(\det \hat{\Sigma}_j) \right] \quad (1.5)$$

where  $\hat{\Sigma}_i$  is the estimated covariance matrix of the model with  $i$  regimes (e.g.  $i=1$  imply an one threshold regime). In this paper, we test a linear model against 2 regime threshold alternative.

In a linear model, the impulse responses can be derived directly from the estimated coefficients and the estimated responses are symmetric both in terms of the sign and the size of the structural shocks. Furthermore, these impulse responses are constant over time as the covariance structure does not change. These convenient properties do not hold within the class of nonlinear models as shown in [Koop et al. \(1996\)](#). As per [Enders \(2010\)](#), interpretation of the impulse responses is history dependent. The impact of a shock on the time path of the system depends on the magnitude of the current and subsequent shocks. The sign, as well as size of the shocks, are important. For example, the impact of a negative shock on the time-path of the system in a contractionary regime will be different than an expansionary regime. The moving average representation of the TVAR is nonlinear in the structural disturbances. Because some shocks may lead to switches between regimes, and thus their Wold decomposition does not exist ([Afonso et al., 2011](#)). Consequently, in contrast to linear models, we cannot construct the impulse responses as the paths the variables follow after an initial shock, assuming that no other shock hits the system. To cope with these issues [Koop et al. \(1996\)](#) proposed nonlinear impulse response functions defined as the difference between the forecasted paths of variables with and without a shock to a variable of interest. Formally, the nonlinear impulse responses functions (NIRF) are defined as:

$$NIRF_Y(k, \epsilon_t, \Omega_{t-1}) = E(Y_{t+k} | \epsilon_t, \Omega_{t-1}) - E(Y_{t+k} | \Omega_{t-1}) \quad (1.6)$$

where  $Y_{t+k}$  is a vector of forecasts at horizon  $k$ ,  $\Omega_{t-1}$  is the available information set, and  $\epsilon_t$  is the shock at time  $t$ . The following discussion draws on [Afonso et al. \(2011\)](#). This formulation implies that the impulse response functions depend on the initial conditions and that there is no restriction regarding the symmetry of the shocks. Therefore, in order to get the complete information about the dynamics of the model, the impulse responses have to be simulated

for various sizes and for the signs of the shocks. The algorithm proceeds as follows. First, the shocks for the periods from  $\theta$  to  $q$  are drawn from the residuals of the estimated VAR model. Then, for each initial value, that is for each point of our sample, this sequence of shocks is fed through the model to produce forecasts conditional on initial conditions. These steps are repeated for the same initial condition and the same set of residuals except for the shock to the variable of interest, which is set to (+/-) 1 standard error and (+/-) 2 standard errors at time 0. Second, we can calculate the forecasts conditional on the shocks and on the initial conditions with and without an additional shock at  $t = 0$ , and the difference between these two is the impulse response function. This procedure is replicated 500-times for each initial condition, and then we compute averages over the initial conditions from each regime to get the average impulse responses. We compute nonlinear impulse responses for two regimes defined by the threshold VAR above.

## 1.4 Results and Analysis

In this section, we present the results and analysis. We begin our analysis comparing the model specification. Table 1.2 below present the likelihood ratio based specification tests described earlier. We estimate four threshold VAR models with a lag-length of 2 (as per the AIC), where each model is estimated with a separate budget deficits definition. In all cases, we reject the null of linear VAR against the alternative of 2 regimes one threshold VAR<sup>4</sup>. Thus, estimating a threshold model is justified<sup>5</sup>. The endogenously estimated threshold defines the two regimes. For example, when output gap falls below 1.64 the corresponding model is our bad state or the recessionary state. When the output gap is above threshold

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<sup>4</sup> Model with Primary Budget Deficit to GDP ratio has 81.8% observation in the regime above the threshold and 18.2% are below threshold, model with Gross Budget Deficit to GDP ratio has 81.8% observation in the regime above the threshold and 18.2% are below threshold, model with Current Budget Deficit to GDP ratio has 80.7% observation in the regime above the threshold and 19.3% are below threshold, and model with Primary Current Budget Deficit to GDP ratio has 80.7% observation in the regime above the threshold and 19.3% are below threshold.

<sup>5</sup>The homoskedastic bootstrap distribution of the critical values are based on the residuals from the  $H_0$  model.

value of 1.64 the corresponding model is our good state or the expansionary state.

Table 1.1: Specification Test: All TVAR Models

Model	Threshold $\tau$	$H_0$ : Linear VAR
Gross Budget Deficit to GDP ratio	1.64	44.03* (0.06)
Primary Budget Deficit to GDP ratio	1.64	49.51** (0.02)
Current Budget Deficit to GDP ratio	1.57	63.30*** (0.00)
Primary Current Budget Deficit to GDP ratio	1.54	64.31*** (0.00)

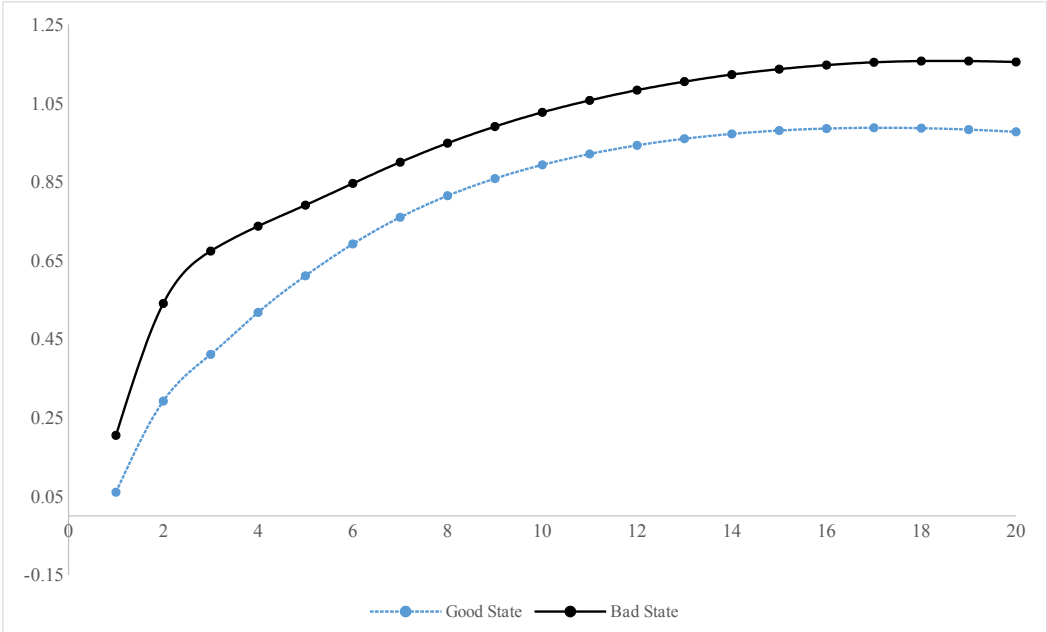
Note:  $p$ -value in the parenthesis

We proceed to nonlinear impulse response analyses below. We estimate two sets of impulse response- a good state and a bad state as per our threshold VAR model. We present the nonlinear impulse responses for each variable to a positive 1 standard deviation increase in the other. Figures below present the good state (expansion) and bad state (recession) comparisons. These impulse responses present us a visual comparison of how one variable is behaving over the forecast horizon with respect to an increase in the other during good state and bad state. We estimate four different models for each of the budget deficit definitions. For the purpose of brevity, we will present and discuss the nonlinear impulse responses for the models estimated with primary budget deficit to GDP and primary current deficit to GDP ratios<sup>6</sup>. By definition, primary budget deficit and primary current deficit measures depict by how much government receipts fall short of the government spending outlays each period. Figures 1.3 and 1.4 below present the nonlinear impulse responses of the debt to GDP ratio. A +1 standard deviation increase in the primary budget deficit to GDP ratio increases the debt to GDP ratio over the forecast horizon. We also observe the debt to GDP ratio grows faster during a bad state (recessions) as opposed to a good state (expansions).

<sup>6</sup>Please see Appendix A for the remaining nonlinear impulse responses.

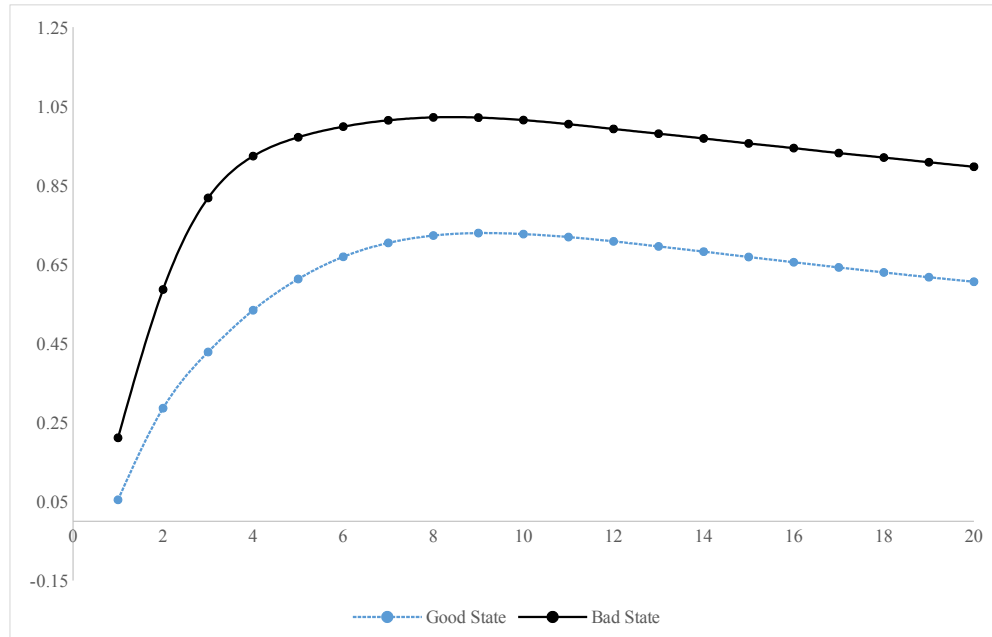
For example, The debt to GDP ratio rises to about 0.7 after 4 quarters in a bad state, whereas in the good state it rises to about 0.5 after about 4 quarters with respect to an increase in the primary budget deficit to GDP ratio. The impulse responses indicate the debt to GDP ratio shows an asymmetric response. The debt to GDP ratio rises much faster during a bad state and remains at a higher level than the responses during a good state. We find the same pattern of response in debt to GDP ratio for all the budget deficit definitions used in this essay (the nonlinear impulse responses for the rest are in the appendix A). Arguably, the government prioritizes economic stabilization during a recession instead of maintaining fiscal balance. Thus adopting an expansionary fiscal policy that raises budget deficits, henceforth increasing the debt. These findings are consistent with the history of United States as there were several examples of tax cuts, and/or, increases in governments discretionary spending (at times both) by the government to stimulate the economy during recessions. Budget deficits, resulting from the fiscal stimulus, always increase the government debt in the United States. [Cassou et al. \(2015\)](#) also find asymmetry, and the impulses responses conform to theoretical predictions in [Marcet and Scott \(2009\)](#).

Figure 1.3: Response of Debt to GDP ratio



Note: +1 standard deviation increase in primary budget deficit to GDP ratio

Figure 1.4: Response of Debt to GDP ratio



Note: +1 standard deviation increase in primary current deficit to GDP ratio

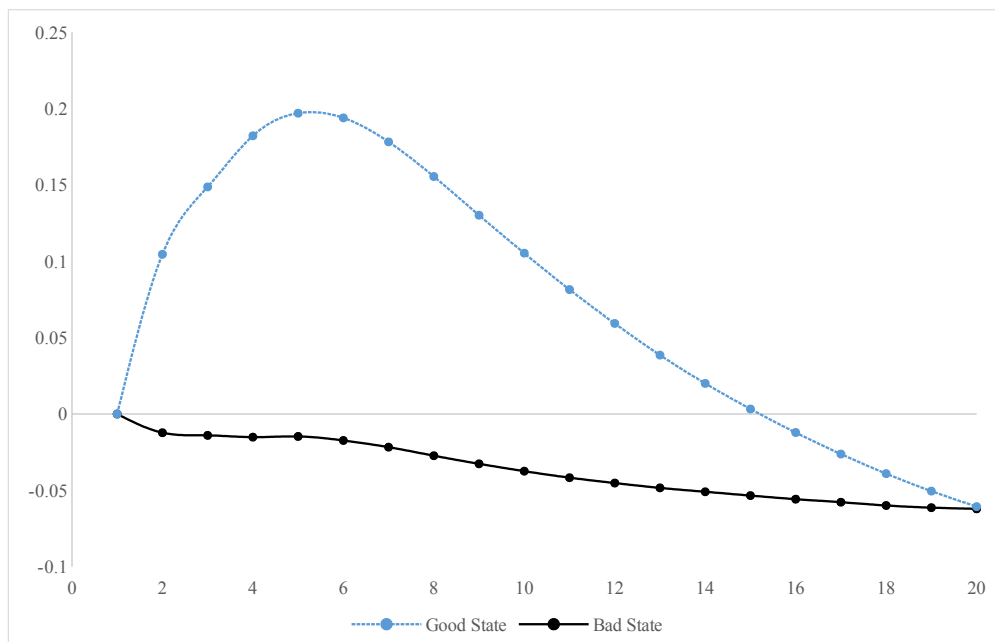
Figures 1.5 and 1.6 present the good state (expansions) and bad state (recessions) comparisons of the nonlinear impulse response of budget deficits to GDP ratio to a +1 standard deviation increase in the debt to GDP ratio. For both the budget deficit definitions, we observe a positive 1 standard deviation increase in the debt to GDP ratio does not increase budget deficit to GDP ratio in a bad state. During a good state, We observe the budget deficit to GDP ratio rises very fast after a quarter before it starts to become negative after about 5 quarters. In the bad state, on the other hand, the budget deficit to GDP ratio responds very slowly and rises after about 3 quarters over the forecast horizon (for the primary current budget deficit to GDP ratio). Budget deficits are related to macroeconomic policy, and, contingent upon the state of the economy. During recessions, the government has the incentive use an expansionary fiscal policy. But a higher debt level may erase the motivation in favor of austerity<sup>7</sup> During an economic expansion, on the other hand, the necessity to stabilize the economy is not binding for the government. We observe the budget deficits to GDP ratio to increase in a good state for few quarters and subsequently they fall. In the U.S., the

<sup>7</sup>The economic situation across some of the European economies such as Greece, U.K., and Portugal during the Great Recession are examples of such a scenario.



sharp rise in the budget deficits to GDP ratio during a good state as opposed to a bad state could stem from the interest rate differential in the U.S. We know the interest rates usually remain at a higher level during economic expansions and are usually lower during recessions. As such, the rise in debt may increase the budget deficit in the good state due to higher interest rate. Also, the components of the budget as well. The non-discretionary component of the budget is dictated by certain statutes and require Congressional deliberation for any changes. As per the impulse responses, the deficit does not remain high, rather they decline after about 4 quarters. During the Regan era, we observe the budget deficit to rise very fast due to the tax cuts. These results conform to predictions in [Uctum and Wickens \(2000\)](#). We observe the deficit to GDP ratio gradually fall over the forecast horizon. Despite an initial surge, the ensuing decline in budget deficits over the forecast horizon indicate governments prefer to reduce deficits over time when debt to GDP ratio rises conforming to findings in [Arestis et al. \(2004\)](#).

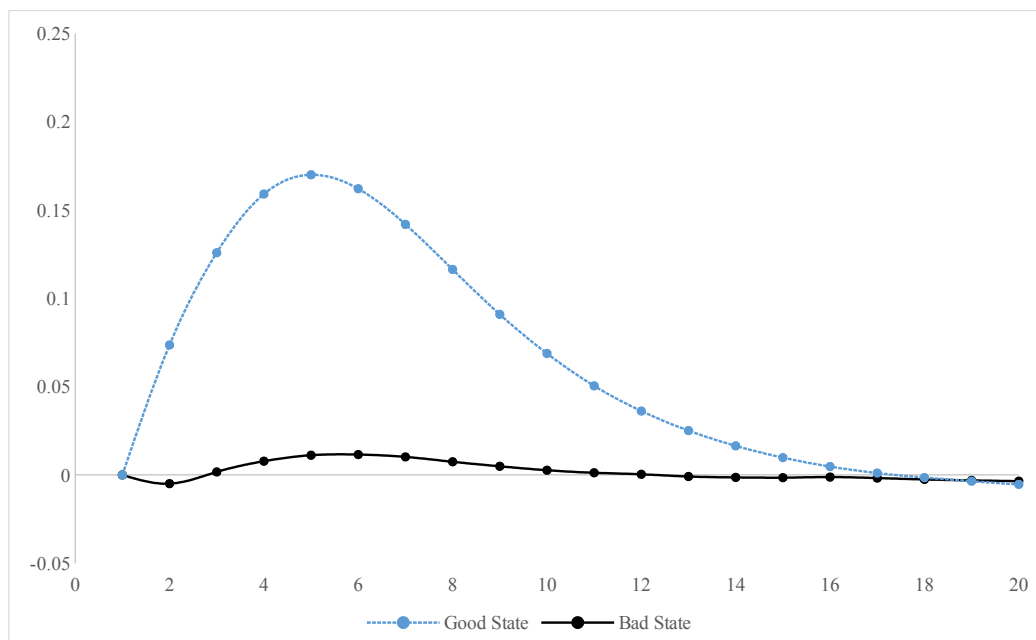
Figure 1.5: Response of Primary Budget Deficit to GDP ratio



Note: +1 standard deviation increase in debt to GDP ratio

In the figures 1.7 and 1.8, we present the nonlinear impulse responses of the debt to

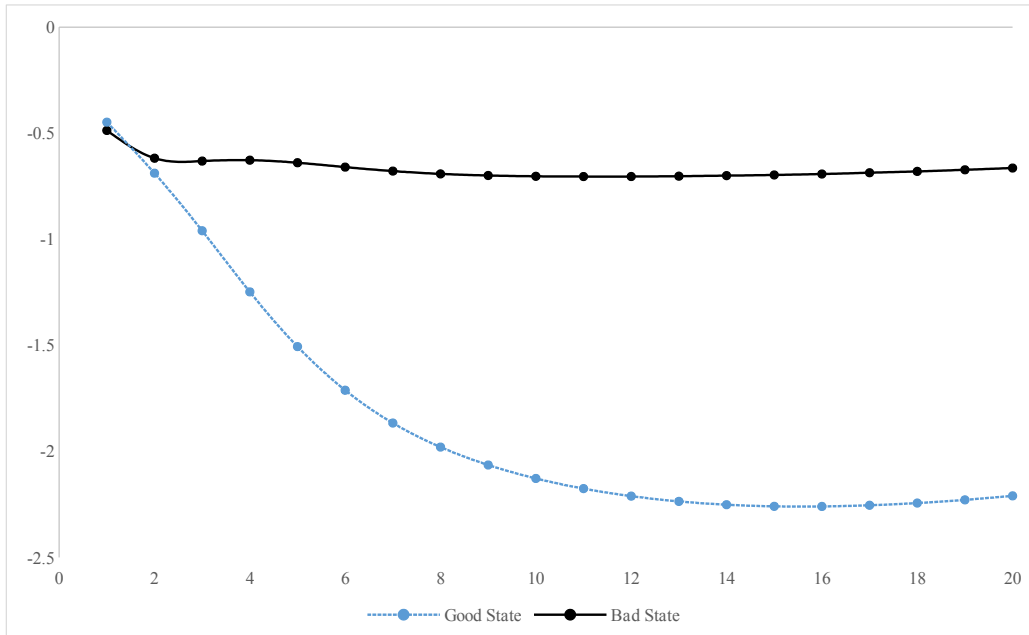
Figure 1.6: Response of Primary Current Budget Deficit to GDP ratio



Note: +1 standard deviation increase in debt to GDP ratio

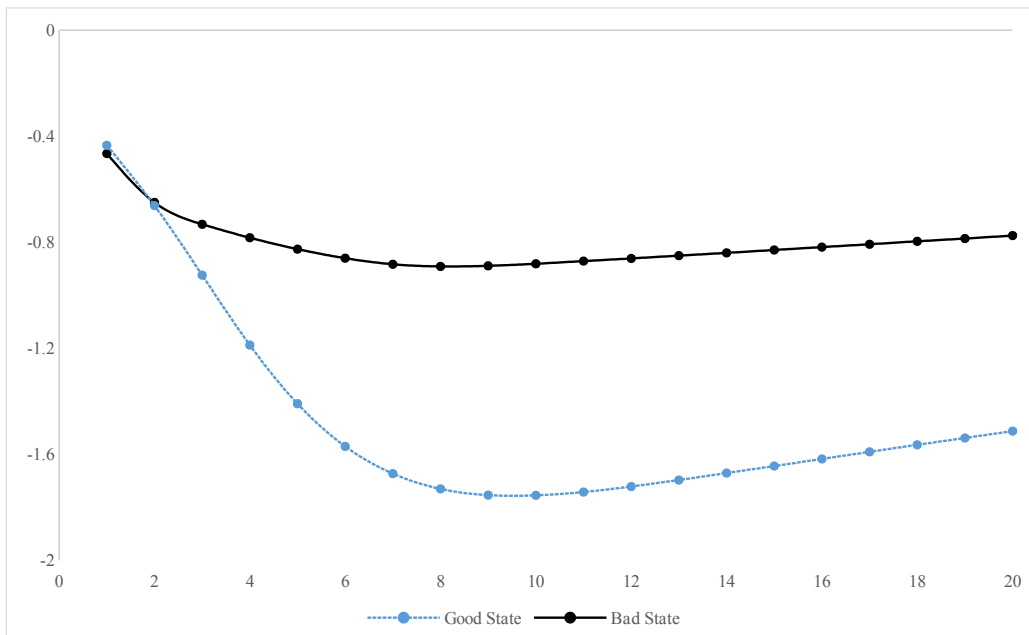
GDP ratio with respect to a +1 standard deviation increase in the output gap. For intuitive purposes let's assume that these positive 1 standard deviation increase in the output gap are positive technology shocks that boost the economy's productivity, thus output. A positive 1 standard deviation increase in output gap implies the economy is expanding. We can argue that when the economy starts to recover or expand, components of the aggregate demand increase and unemployment rate starts to go down. Figures 1.7 and 1.8 show the nonlinear impulse responses for each state. Here the debt to GDP ratio falls immediately with respect to a +1 increase in the output gap. We observe the debt to GDP ratio is falling in both states, although the rate is faster in a good state as opposed to a bad state. This shows that debt to GDP postulates some counter-cyclicality in bad state conforming to findings of [Cassou et al. \(2015\)](#). The decline in the debt to GDP ratio could potentially emerge from the fact that the increase in the output raises the government's tax revenues through the tax code along with reduction in the welfare payments (this reduces government spending), thus reducing budget deficit, and at the same time ability for subsequent debt repayment.

Figure 1.7: Response of Debt to GDP ratio



Note: +1 standard deviation increase in the Output Gap

Figure 1.8: Response of Debt to GDP ratio

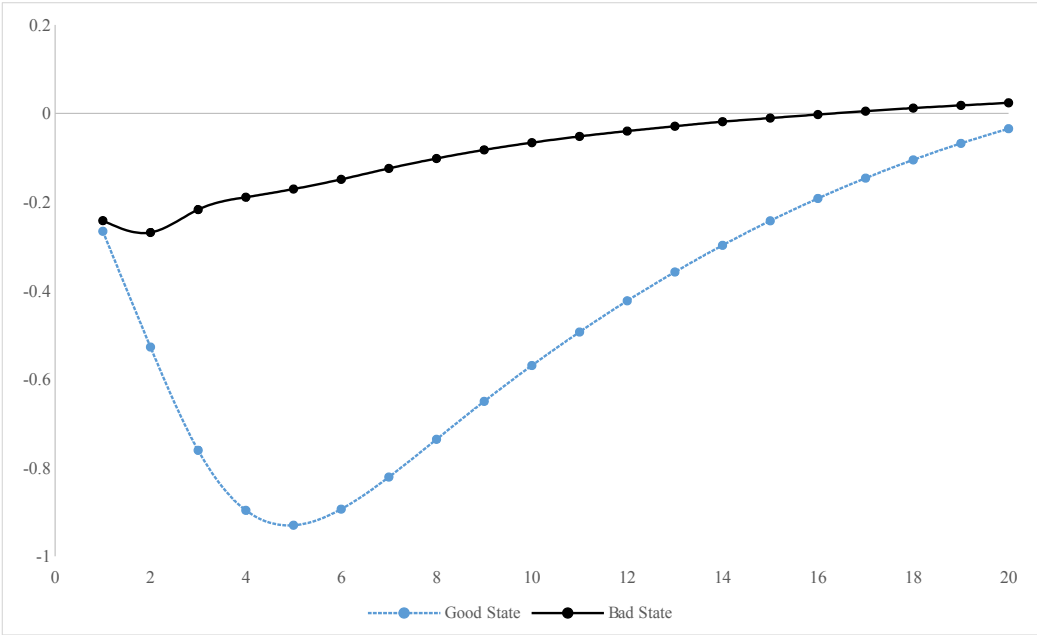


Note: +1 standard deviation increase in the Output Gap

Figures 1.9 and 1.10 below, we present the nonlinear impulse responses of budget deficits to GDP ratio with respect to a positive 1 standard deviation increase in the output gap. The impulse response below indicates an increase in the output gap reduces budget deficits

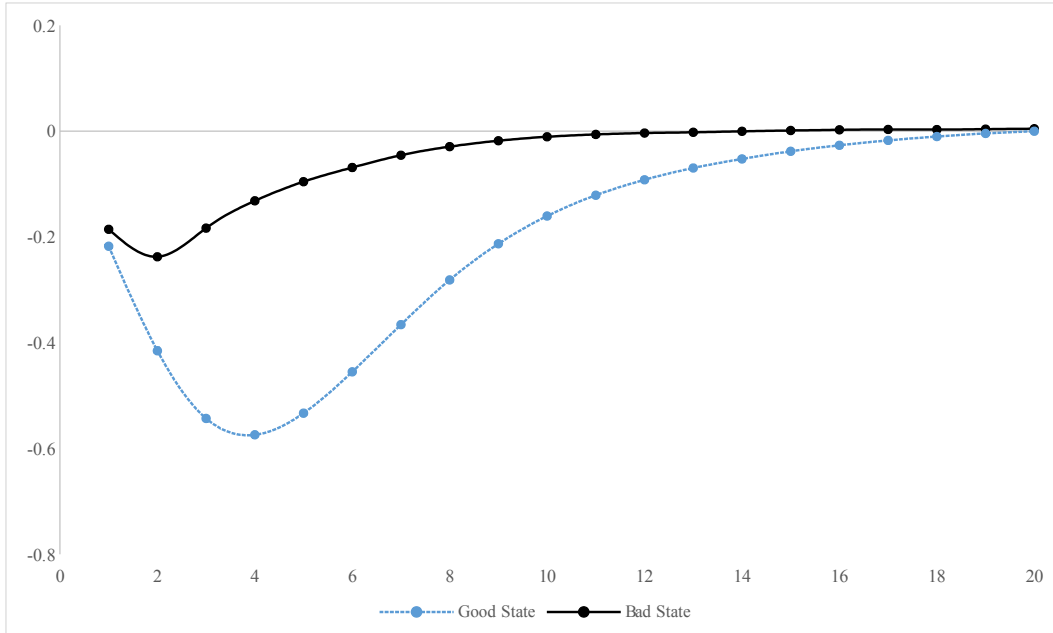
on impact. A positive increase in the output gap may reduce budget deficits through increases in tax revenues as well as a reduction in some of the components of discretionary government spending. We observe the budget deficit to GDP ratio, during the bad state, do not portray a steady decline as opposed to the good state. On impact, the deficit to GDP ratio falls but during the bad state, it worsens relative to the good state. However, the budget deficits to GDP ratio do not show a steady reduction over the forecast horizon for both states. We observe the asymmetric response of budget deficits to GDP ratio over the forecast horizon and degree of counter-cyclicality is different in a bad state relative to a good state. Arguably, during a bad state (recession) the government policy makers opt for economic stabilization adopting expansionary fiscal policy tools. This corroborates to our earlier findings. The relatively slow and non-decreasing nature of the deficit to GDP ratio in a bad state also indicates to the government’s policy choice.

Figure 1.9: Response of Primary Budget Deficit to GDP ratio



Note: +1 standard deviation increase in the Output Gap

Figure 1.10: Response of Primary Current Budget Deficit to GDP ratio



Note: +1 standard deviation increase in the Output Gap

### 1.4.1 Evaluation using Cyclically Adjusted Data

Thus far, our analysis of the dynamics between budget deficits and debt is conducted using data from the NIPA sources. One potential criticism may arise due to cyclical adjustments. As we know, there are automatic stabilizers designed within the fiscal system, especially the tax code and some of the spending programs (especially transfers) to counter business cycle fluctuations. These automatic stabilizers influence the government revenues as well as the expenditures during expansions and contractions of the business cycle. Thus the budget deficit measures may reflect this cyclical nature. [Golinelli and Momigliano \(2009\)](#) use the cyclically adjusted data, [Cassou et al. \(2015\)](#) note that using cyclically adjusted data may be advantageous because they reflect the true reaction function of the policy makers. The cyclically adjusted data eliminate the built-in fiscal measures, thus showing actual contemporaneous budgetary decisions. However, there are economists who argue that automatic stabilizers are also policy decisions, hence using the observed data is sufficient. Regardless of the debate, it is useful to investigate whether the results described above are robust using the cyclically adjusted budget data. For our analysis, we use cyclically adjusted data computed

by the Congressional Budget Office of the United States (CBO).

In the process, we face two problems. First, one problem with this data is a period of availability is 1963: Q3 to 2013: Q3, whereas our analysis uses data from 1947: Q1 to 2015: Q3. Although there are several CBO sources of data, we chose to use the data from a single source that was put out in 2014 in a document called The Budget and Economic Outlook: 2014 to 2024. A second problem that we face is the other definitions of deficits. We are only able to compute the primary budget deficits, but could not compute the other definitions used in this paper. Thus, we confine our analysis to the cyclically adjusted primary budget deficit to GDP ratio. As in our previous analysis, we estimate an endogenous threshold VAR using the output gap as our endogenous threshold variable and compute the nonlinear impulse responses. We estimate a TVAR model with 2 lags (as per AIC). We begin our analysis with specification test presented in Table 1.2 below. We reject the null of linear VAR model against the alternative of two regimes one threshold VAR model<sup>8</sup>. The endogenously estimated threshold defines our two regimes. For example, when output gap falls below -1.84 the corresponding model is our bad state or recessionary state. When the output gap is above threshold value of -1.84 the corresponding model is our good state or expansionary state.

Estimated nonlinear impulse responses are presented below. We present the impulse response for debt to GDP ratio to positive 1 standard deviation increase in budget deficits to GDP ratio (see Figure 1.11). We again find the asymmetry presented earlier and we observe the government debt to GDP ratio grow faster during a bad state (recession) as opposed to a good state (expansion). With budget deficits to GDP ratio, we can again find conformity to our findings presented earlier. Cyclically adjusted budget deficit to GDP ratio rise faster during a good state (expansion) as oppose to a bad state (recession) with a +1 standard deviation increase to the debt to GDP ratio. With respect to an increase in the output gap, we can observe the counter-cyclical behavior for the debt to GDP ratio as well as for the

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<sup>8</sup>84.9% observation in the regime above the threshold and 15.1% are below the threshold.

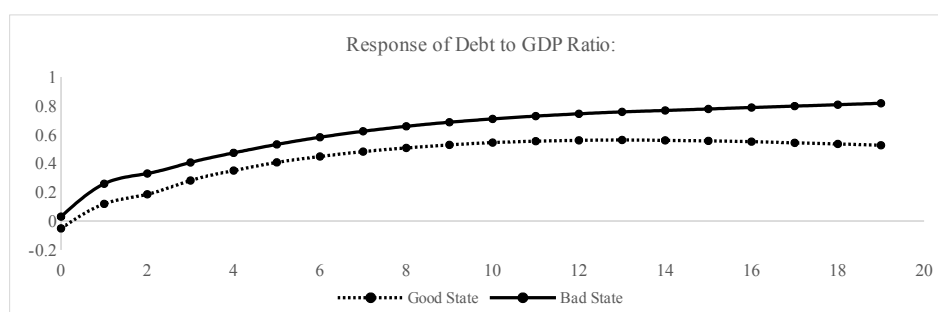
budget deficits to GDP ratio.

Table 1.2: Specification Test: Cyclically Adjusted Data

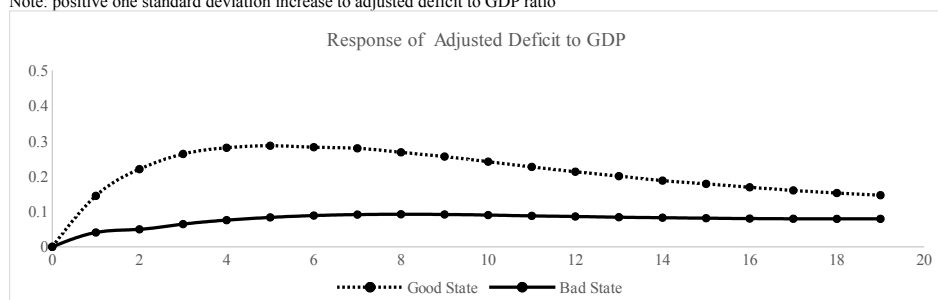
Model	Threshold $\tau$	$H_0$ : Linear VAR
Gross Budget Deficit to GDP ratio	-1.84	53.44** (0.01)

*p*-value in the parenthesis

Figure 1.11: Response of Debt to GDP ratio



Note: positive one standard deviation increase to adjusted deficit to GDP ratio

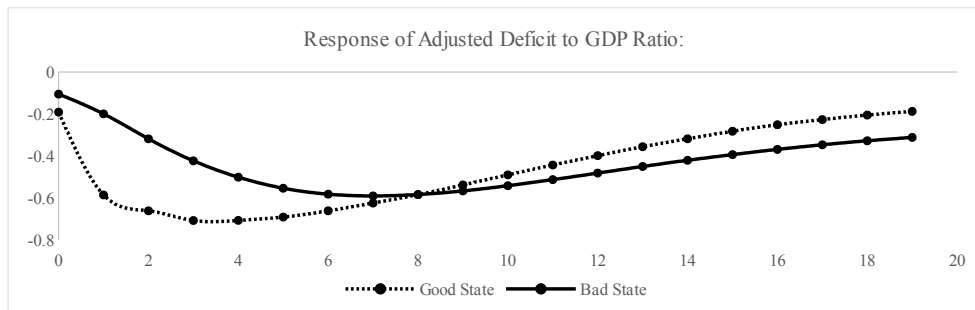


Note: positive one standard deviation increase to debt to GDP ratio

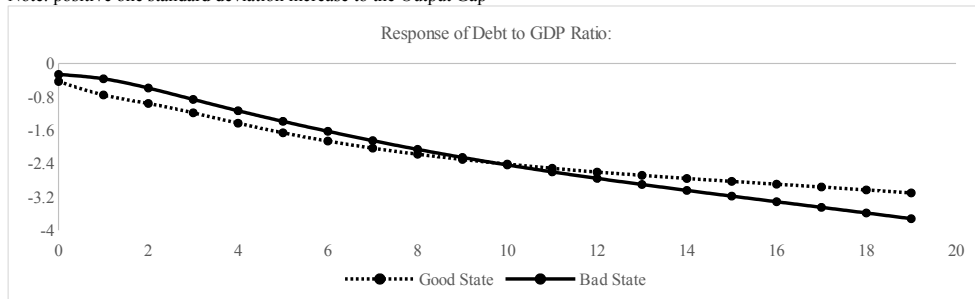
When we explore the nonlinear impulse responses of the debt to GDP and cyclically adjusted budget deficit to GDP ratios (see Figure 1.12) we again observe conformity to our findings presented earlier. Both these variables indicate presence of counter-cyclical behaviour with asymmetry.

These robustness check conform to our findings presented in the previous section. The observed asymmetry and counter-cyclical behaviour conform to our earlier finding that government

Figure 1.12: Response of Debt to GDP ratio



Note: positive one standard deviation increase to the Output Gap



Note: positive one standard deviation increase to the Output Gap

prefers economic stabilization over fiscal balances and debts during recessions.



## 1.5 Conclusion

In this paper, we investigate the dynamics between government debt and budget deficits in the United States during recessions as well as expansions using quarterly data for the post-World War II period. We employ the threshold VAR model and compute the nonlinear impulse responses. The econometric specification test confirm the endogenous threshold VAR is the correct specification against the linear VAR model. The estimated nonlinear impulse responses during a bad state and a good state allow us to explore the dynamics between the two in a more objective manner.

Nonlinear impulse response analysis for the good state and bad state comparisons show that response of the government debt to GDP ratio to an increase in the budget deficit to GDP ratio is asymmetric, and counter-cyclical to an increase in the output gap. The debt to GDP ratio rises faster during a bad state (recession) as opposed to a good state (expansion). These results imply policy makers chose economic stabilization over balancing the budget during recessions. However, the response of budget deficits to GDP to an increase to government debt to ratio is asymmetric, and counter-cyclical to an increase in the output gap. We observe the budget deficits to GDP ratio to rise more during a good state (expansion) as opposed to a bad state (recession). We check the robustness of our findings using a time-series of primary budget deficits cyclically adjusted for variations arising from automatic stabilizers, which conform to our earlier findings. The findings above provide important empirical insight on governments choice of economic stabilization over fiscal balance during recessions in the United States in the post- World War II period.

The adoption of expansionary fiscal policy during recessions is imperative for economic stabilization. However, such a policy increases the national debt. According to our analysis, we argue for deficit reduction during economic expansions as a policy recommendation.

# Chapter 2

## Relationship between the monetary base and the fiscal debt in the U.S.: evidence from threshold cointegration

### 2.1 Introduction

Does fiscal policy influence the Fed's choice of monetary policy strategy? Increasing government debt could potentially influence discretionary monetary policy choices by the central bank (Blake and Kirsanova, 2012; Díaz-Giménez et al., 2008). Corsetti and Dedola (2014) argue that unconventional monetary policy- targeting monetary aggregates instead of inflation- can act as a backstop for an economy facing default risk with sovereign debt crisis. Taylor (1993) is considered as the point of reference in defining the Fed's monetary policy strategy, and a plethora of research has examined forward-looking as well as backward-looking versions of the rule. There are empirical studies that explore the presence of regimes in the U.S. monetary policy by investigating how the Taylor rule evolves with the Fed Chairmans tenure (Bae et al., 2012). Asymmetry in monetary policy is another aspect investigated extensively in recent times (Cassou et al., 2012). Despite recurrent scrutiny from a variety of dimensions, there is a lack of comprehensive consensus on how the Fed's monetary policy strategy evolve.

Fiscal debt, arising from budget deficits, is a likely contender to influence monetary policy. The Fed's open market operations involve purchase and sale of government issued securities, which in turn affect the money supply through the high-powered money or the monetary base. Empirical investigations grounded on the [Taylor \(1993\)](#) in analyzing monetary policy strategy pay little or no attention to the possible nexus between the government debt and monetary aggregates such as the monetary base. We can develop more insight on the Fed's monetary policy strategy by analyzing the relationship between monetary base and fiscal debt.

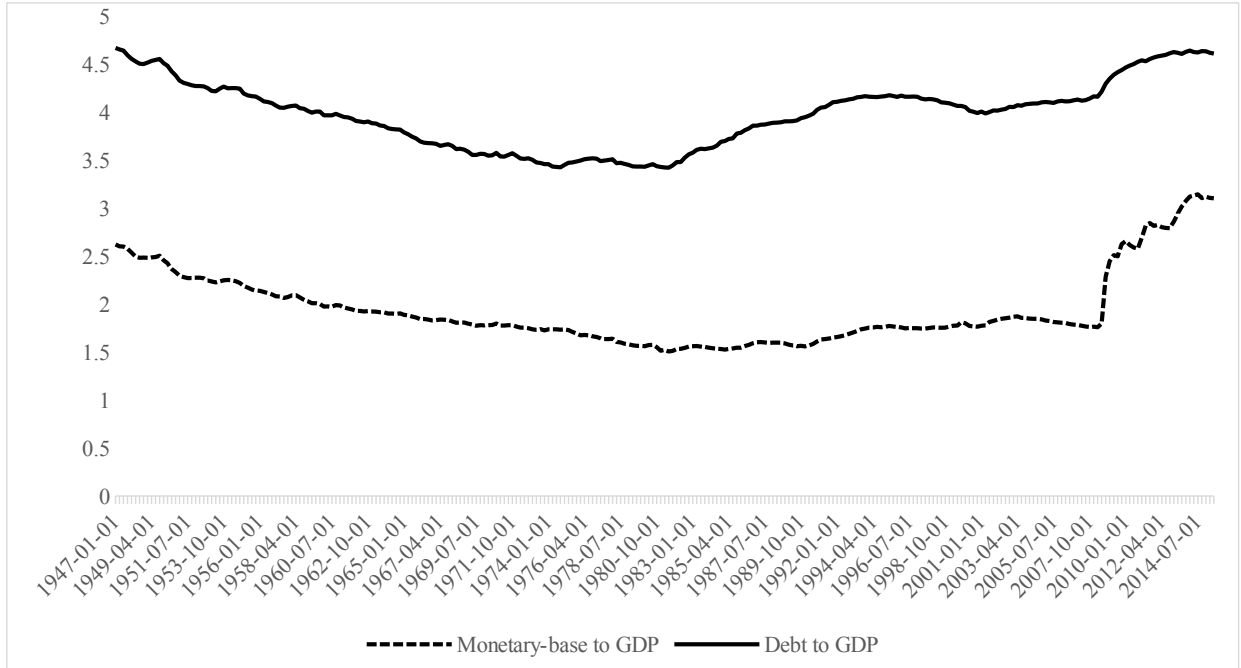
[Sims \(2013\)](#) argues in favor of a potential linkage between monetary and fiscal policy, but [Mankiw \(2016\)](#) argues that there is a dearth of empirical evidence. A fiscal expansion, through a tax cut or increased spending (or both), may lead to budget deficits that sway government debt. In the Taylor rule, the real federal funds rate reacts to deviations in contemporaneous inflation from a target, and, deviations in real output from its long-run potential level. But the rule does not specify any responsiveness to fiscal policy. Whereas the Fed's open market operations involve buying and selling of government issued securities. Government debt could bind the central bank to pursue a monetary policy in an accommodative manner. Monetization of fiscal debt is one area that has been under public scrutiny from the policy makers and economists alike ([Thornton, 2010](#)). Fiscal vulnerabilities arising from high government debt is likely to create new and complex interactions between public debt management and monetary policy ([Blake and Kirsanova, 2012](#); [Blommestein and Turner, 2011](#)).

Empirical evidence of a long-run relationship or co-movement between fiscal debt and monetary base can provide insight on how government debt becomes relevant to the Fed's monetary policy strategy. Figure 2.1 below illustrates a similar pattern in the log of U.S. government debt to GDP and the log of monetary base to GDP ratios since 1947. We observe both the ratios decline until the mid-1970s, and then start to rise again. Both these variables indicate a pattern showing a potential long-run relationship or co-movement between

the two. In this paper, we investigate formally using statistical methods if there is a long-run relationship between government debt and monetary base in the U.S. Since there is empirical evidence that infers the presence of regimes in monetary policy, we explore the presence of a long-run relationship while considering regimes or threshold (Bae et al., 2012; Cassou et al., 2012; Sims and Zha, 2006). We use the threshold cointegration technique to investigate the presence of a long-run relationship and estimate a threshold vector error-correction model (TVECM henceforth) to understand the short-run dynamics. The presence of a long-run relationship, established through formal statistical procedures, imply we can obtain the dis-equilibriums or movements deviating away from the long-run relationship over time. We can use these dis-equilibriums in a vector error-correction model to analyze the short-run movements. Our contribution to the literature takes two forms. First, using monthly data from 1942 to 2015 we find evidence of a threshold long-run relationship between government debt and monetary base and threshold cointegration in the U.S. Typical cointegration tests do not detect the presence of this long-run relationship. Second, the short-run dynamics from the TVECM shows no evidence of debt monetization by the Fed, and the fiscal authorities adjust debt in the short-run. This implies that the federal government debt does not influence the Fed's monetary policy choices in the U.S. Thus, we find no evidence of monetary accommodation by the Fed implying independence or autonomy of the Fed.

We organize the rest of the paper as follows: in Section 2.2 we discuss the literature; Section 2.3 we outline the data and methodology; Section 4.4 discusses the results, and Section 2.5 presents the conclusion.

Figure 2.1: Log of Monetary base to GDP and Government debt to GDP, 1947 to 2015



## 2.2 Literature Review

Recently, there is renewed interest in the interaction between fiscal authority and the central bank in the literature. There is a rich volume of theoretical papers discussing these interactions as well as empirical papers investigating for evidence.

[Sims \(2013\)](#) argues that fiscal and monetary policy are intertwined, and interactions between the fiscal authority and the central bank are key in determining the price level for an economy. Theoretical papers such as [Leeper and Leith \(2016\)](#) develop the theory of price-level determination using jointly optimal monetary and fiscal policy. They argue that, from a theoretical perspective when the fiscal authorities adopt an active fiscal rule, the central banks ability to control inflation depends on the maturity structure of the outstanding debt and the nature of its policy response. A central bank, pursuing active monetary policy targeting inflation, may find the inflation rate permanently deviating from the inflation target in the face of a fiscal shock.

Chen et al. (2015) allow both a monetary and fiscal policy described by rules and/or optimal policy which are subject to switches over time. Building on models by Bianchi and Ilut (2017), these authors use a Bayesian estimation method that finds that monetary and fiscal policy are often in conflict. They compare different permutations of rule-based policy to a time-consistent set of optimal policy regimes and argue that time-consistent optimal policy rule offers a data-preferred description of fiscal and monetary policy relative to the rule-based approach. In particular, the fiscal authority is a Stackelberg leader and the monetary authority implements a time-consistent monetary policy with switches in their degree of inflation conservatism. Corsetti and Dedola (2014) argue that the central bank can provide a reinforcement which helps rule out a sovereign debt crisis with the former's purchase of government debt as unconventional monetary policy, instead of the conventional monetary policy regarding the choice on inflation. Using a monetary model similar to that of Calvo (1988), two potential states of the economy (high output state and low output state), the authors<sup>1</sup> argue that a monetary backstop is a feasible option in situations with large government debt and such a monetary backstop prevents high inflation.

Blommestein and Turner (2011) argue for the significance of a fiscal and monetary coordination in the post-financial crisis era. Albeit there could be potential conflicts or tensions between the debt managers and monetary policy makers. Blake and Kirsanova (2012) investigate the stabilization bias that arises in a model of monetary and fiscal policy stabilization using linear-quadratic rational expectation model, where the monetary authority put higher weight on inflation stabilization than the rest of the society. These authors find that if the steady-state level of debt is high, then the monetary authority has to take an active part in debt stabilization. Díaz-Giménez et al. (2008) using a cash-in-advance production economy analyze the implications for the optimal sequential design of monetary policy with nominal and indexed public debt. These authors argue that a Calvo (1988) model is by design unable to show how debt, either nominal or indexed, can impact the choice of monetary

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<sup>1</sup>In this paper authors considers both the monetary authority and fiscal authority are independent of each other. Both authorities are benevolent maximizing the same objective function, consolidate their budget constraints, but take their optimal decisions independently treating each others instruments as constant.

policy. In the cash-in-advance production economy, the rational expectations equilibria for an initial given level of outstanding debt, nominal debt is a burden on optimal monetary policy. [Lambertini and Rovelli \(2003\)](#) examine the relations between monetary and fiscal policies in the process of macroeconomic stabilization. These authors argue, based on their alternative game situations, both fiscal and monetary authority prefer a Stackelberg to that of a Nash game, and the fiscal authority becoming the Stackelberg leader.

In recent times, the theoretical macroeconomic literature discusses avenues for fiscal and monetary policy coordination. But the question of whether there is empirical evidence of fiscal policy influencing monetary policy strategy? The empirical literature lacks a consensus. [Allen and Smith \(1983\)](#) find a positive and significant impact of total Treasury borrowing upon the growth of the monetary base, and they conclude there is evidence of monetary accommodations in the U.S. These authors also argue for regimes in the monetary authority's policy preference. [Cebula \(1988\)](#) finds the federal budget deficit does, in fact, exercise a positive and significant impact upon the nominal Moody's Aaa-rated corporate bond rate, thus report positive empirical evidence supporting a link between fiscal policy and monetary policy. [Darrat \(1990\)](#) argues that the findings in [Cebula \(1988\)](#) are subject to methodological error, and using Engle-Granger procedure of cointegration analysis refuted the above finding. These studies again do not provide any conclusive evidence.

[Devereux \(2010\)](#) while analyzing the effectiveness of fiscal and monetary policy in fighting recessions argues that due to the zero lower bound on interest rates, governments used deficit-financed spending increase, tax cuts and also directly increase the monetary base. Given the liquidity trap at the zero lower bound government spending financed by deficits may be far more expansionary than that financed by tax increases in such an environment. Monetary policies aimed at directly increasing monetary aggregates may be effective, even if interest rates are unchanged. This study is indicative of coordination between monetary and fiscal policy, however, it does not establish an empirical long run relationship. [Glenn and Samad \(2012\)](#) explore a fiscal and monetary policy inter-relation and inflation over the long-

run for the U.S. The authors do not find evidence of fiscal policy dominating monetary policy and vice-versa. [Choi and Holmes \(2014\)](#) using a Markov regime-switching model in investigating the relevance of the Ricardian Equivalence Theorem for the relationship between the budget deficit and the real interest rate. They find that the U.S. economy switches between a Ricardian Equivalence regime characterized by an insignificant relationship between the adjusted primary budget deficit and real long-term interest rate and a regime characterized by a positive relationship. However, these studies do not provide empirical evidence of a central bank accommodating fiscal debt using monetary aggregates such as the monetary base.

[Taylor \(1993\)](#) specified a policy reaction function for the Fed where the real federal funds rate varies to changes in the difference between inflation and the inflation target, and the output gap. [Bae et al. \(2012\)](#) note that there are studies that investigate the backward-looking as well as forward-looking versions of the Taylor rule, however, these studies lack consensus regarding the nature, evolution, or even existence of monetary policy regime. [Clarida et al. \(2000\)](#) find evidence that there are significant differences in the manner in which monetary policy is conducted in pre- and post- 1979. [Orphanides \(2004\)](#) argues that results in [Clarida et al. \(2000\)](#) should be considered with caution. [Orphanides \(2004\)](#) argues that the Fed uses real-time data while determining the suitable monetary policy and do not use the ex-post data. The author shows that the differences found in [Clarida et al. \(2000\)](#) disappears when using real-time data. [Primiceri \(2005\)](#) finds evidence of time-variation in monetary policy; however, finds little evidence of interest rate responding to high inflation and unemployment episodes. [Sims and Zha \(2006\)](#) argue that the Taylor rule (1993) could potentially be misleading in explaining regimes in monetary policy as there could be episodes where monetary policymakers were focusing on monetary aggregates instead. They use an SVAR model that explicitly allows for changes in the policy regime, and report that the data is best explained by a model with no changes at all in coefficients, either of the policy rule or of the private sector block of the model. They find variance of the structural disturbances change across regimes. [Kim and Nelson \(2006\)](#) find three separate regimes in US monetary



policy in a time-varying model for forward-looking monetary policy rule. [Bae et al. \(2012\)](#) employ a two-step MLE procedure to estimate a Markov-switching model to investigate a forward-looking Taylor rule for monetary policy in which policy changes are non-recurrent and monotonic but probabilistic. They find the presence of five regimes in the U.S. monetary policy. There are studies which investigate asymmetry in monetary policy preference. [Cassou and Vázquez \(2014\)](#) consider a time-varying parameter extension of the ([Ruge-Murcia, 2003, 2004](#)) model to explore asymmetric preference of the Fed. [Cassou et al. \(2012\)](#) also report similar asymmetry in the Fed's preference using the Taylor rule (1993).

The theoretical literature discussed above indicates a fiscal and monetary coordination, and/or accommodation of fiscal policy by the monetary authority, especially when fiscal authorities face higher debt burden. There is a plethora of empirical research on the issue of regimes and asymmetry in monetary policy strategy by the Fed. These studies use the Taylor (1993) rule as their benchmark. The Taylor (1993) rule, given its specification, refers to the Fed's target. The research, using the Taylor (1993) rule, however, produce a varied set of evidence. We do not find empirical studies that investigate the potential relationship between the government debt and the monetary base while considering regimes in the policy regime. Also, there is a dearth of empirical evidence if the Fed is accommodating government debt through debt monetization in the short-run. This is one area that merits investigation to fill-up the gap in the literature.

## 2.3 Data and Methodology

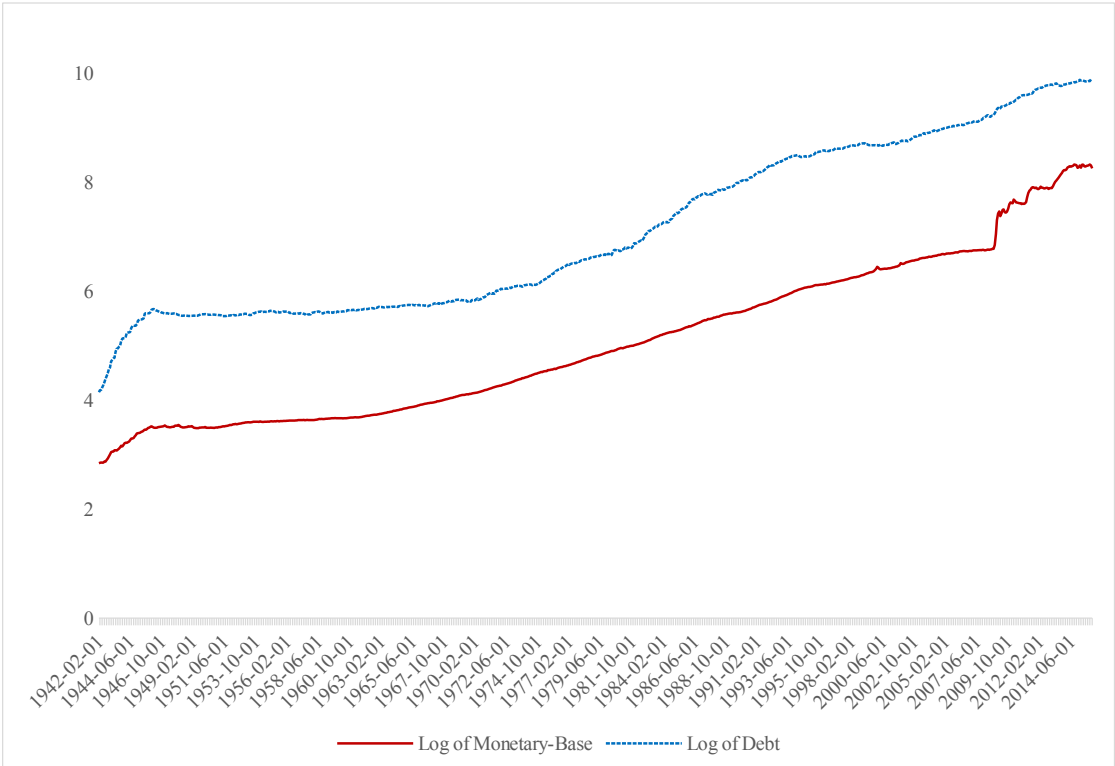
In this section, we discuss the data and methodology use in this essay. We are using monthly data from 1942: 1 to 2015: 12 on government debt and monetary base of the U.S. The government debt data series is obtained from the Federal Reserve Bank of Dallas<sup>2</sup> and the monetary base data is obtained from the Federal Reserve Bank of St. Louis website. The GDP de-

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<sup>2</sup>Federal Reserve Bank of Texas compiled the data from various issues of the US Treasury Bulletins.

flator series are obtained from the National Income and Product Account (NIPA) tables. We use the GDP deflator to convert the government debt and monetary base data into real terms. The variables used in this essay are in their natural log, where  $LB_t$  denote the natural log of monetary base and  $LD_t$  denote the natural log government debt. Figure 2.2 below plots the data from 1942 to 2015. The plot shows that both the series move together over time. Both figures 2.1 and 2.2 indicate the presence of a long-run relationship between the two.

Figure 2.2: Natural log of Monetary Base and Government Debt, 1942 to 2015



There are three potential cases for cointegration or a long-run relationship between government debt and monetary base. First, there is no cointegration or no long-run relationship. Second, there is cointegration or a long-run relationship. Finally, there could be threshold cointegration or a threshold long-run relationship. In our empirical methodology we investigate all three cases. As a pre-requisite to testing for cointegration between the two, we begin our analysis by investigating non-stationarity or the presence of a unit root in the data. We begin our analysis with simple unit root tests, namely augmented [Dickey and Fuller \(1981\)](#)

and Phillips and Ouliaris (1990) tests. In the presence of a structural break various Dickey and Fuller (1981) test statistics are biased toward non-rejection of a unit root and Phillips and Ouliaris (1990) procedure assumes the date of the structural break is known (Enders, 2010). Plots of the data in Figure 2.2 indicate that both these series may have structural breaks at unknown dates. Hence, we use the Lee and Strazicich (2003) and Lumsdaine and Papell (1997) tests to investigate the stationarity property of the data. The Lee and Strazicich (2003) test has two versions: the crash model investigates an abrupt change in the level, and the break model allows for simultaneous changes in level and trend.

We then proceed to investigate if the two data series have cointegration or a long-run relationship. Before describing the threshold model, we first discuss the simple non-threshold cointegration model. The simple cointegrating relationship is given by:

$$LD_t = \beta LB_t + \mu_t \quad (2.1)$$

We use the Johansen (1991) procedure as well as the Phillips and Ouliaris (1990) tests to investigate the presence of cointegration between the two variables. These procedures are extensively described in other empirical papers, thus we do not discuss them in detail. In general, the tests above assume the long-run relationship is invariant to time. We proceed to test the presence of cointegration in a threshold model. The cointegration methodology suggested by Engle and Granger (1987) involve estimation of equation (1) and testing if the residual has a unit root or not. Since the augmented Dickey and Fuller (1981) test on the error-correction term  $\hat{\mu}_t$  from equation 1 do not consider potential nonlinearity in the process, we run the following regression and conduct the RESET test.

$$\Delta\mu_t = \beta_0 + \beta_1\hat{\mu}_{t-1} + \beta_2\Delta\hat{\mu}_{t-1} + \beta_3\Delta\hat{\mu}_{t-1}^2 + \beta_4\Delta\hat{\mu}_{t-1}^3 + \epsilon_t \quad (2.2)$$

In order to inspect the case for potential non-linearity, we perform the following test  $H_0 : \beta_3 = \beta_4 = 0$ . In addition, Pippenger and Goering (1993) and Balke and Fomby (1997)

showed that tests of unit root have lower power in the presence of asymmetric adjustment. As such, we consider a TAR structure for the error-correction term as follows:

$$\Delta\mu_t = \rho_1 D(\hat{\mu}_{t-1} - \tau) + \rho_2(1 - D)(\hat{\mu}_{t-1} - \tau) + \sum_{i=1}^l \gamma_{T,i} \Delta\hat{\mu}_{t-1} + v_{T,t} \quad (2.3)$$

In the above threshold regression,  $\rho_i$  and  $\gamma_{T,i}$  are the parameters of the model for  $i = 1, \dots, l$  and  $v_{T,t}$  is the error term with an added subscript to distinguish it from the non-threshold error term.  $D$  is the dummy such that:

$$D_t = \begin{cases} 1, & \text{if } \Delta\hat{\mu}_{t-1} \geq \tau \\ 0, & \text{if } \Delta\hat{\mu}_{t-1} < \tau \end{cases} \quad (2.4)$$

In this case  $\tau$  represents the endogenously chosen threshold following the method by [Chan \(1993\)](#). We arrange the endogenously chosen threshold variable in ascending order and trim 15% at the top and 15% at the bottom to avoid over-fitting. We estimate the model for each value of the threshold variable and save the sum squared of residuals. We chose the model with the lowest sum squared of residuals as our chosen model and the corresponding value of the endogenous threshold variable as our threshold. As in the [Engle and Granger \(1987\)](#) the lag length ( $l$ ) typically is chosen by some type of information criterion so that the model is well specified and the results in the  $\gamma_{T,i}$  being white noise. We will explore two important aspects in the above structure. First, we will investigate the presence of threshold cointegration. Second, we will test if the lag length is greater than 1 in the threshold model. Testing for a threshold cointegration is explored by [Enders and Siklos \(2001\)](#). [Enders and Siklos \(2001\)](#) describe two possible test statistics:  $H_0 : \rho_1 = \rho_2 = 0$  and  $H_0 : \rho_i = 0, i = 1, 2$ . The authors refer the first one as  $\Phi^*$  statistics and the second one as  $t - max$  statistics respectively, and also note that the former has more power than the  $t - max$  statistics. The null hypothesis  $H_0 : \rho_1 = \rho_2 = 0$  implies no cointegration between the variables, and alternative implies the presence of threshold cointegration. As with the [Engle and Granger \(1987\)](#) method, the test statistics do not have standard distributions for the [Enders and Siklos \(2001\)](#) tests. They describe the methods for generating the proper critical values.

Since the exact nature of non-linearity may be unknown, we use an M-TAR (momentum TAR) with the threshold defined on  $\hat{\mu}_{t-1}$  as well (Enders, 2010). M-TAR adjustments can be especially useful when the policy makers are viewed as attempting to smooth out any large changes in the series Enders and Siklos (2001). In addition, we will also use the Hansen and Seo (2002) method to explore threshold cointegration. Hansen and Seo (2002) propose two heteroscedasticity-consistent LM tests for the null of linear cointegration (i.e. there is no threshold effect), against the alternative of threshold cointegration. The first test can be used when the true cointegrating vector is known a priori, and is denoted by:

$$\sup LM^0 = \sup_{\gamma^L \leq \gamma \leq \gamma^U} LM(\beta_0, \gamma) \quad (2.5)$$

where  $\beta_0$  is the known cointegrating vector. When the true cointegrating vector is unknown we use the following, where  $\tilde{\beta}$  is the null hypothesis estimate  $\beta$ .

$$\sup LM^0 = \sup_{\gamma^L \leq \gamma \leq \gamma^U} LM(\tilde{\beta}_0, \gamma) \quad (2.6)$$

In both tests,  $[\gamma^L, \gamma^U]$  is the search region set such that  $\gamma^L$  is the  $\pi_0$  percentile of  $\hat{\mu}_{t-1}$  and  $\gamma^U$  is the  $(1 - \pi_0)$  percentile of  $\hat{\mu}_{t-1}$ . Andrews (1994) suggest setting  $\pi_0$  between 0.05 and 0.15, we use 0.15 in this paper. We proceed with the second test where the cointegrating vector is unknown, and employ parametric residual bootstrap procedures to approximate the null distribution of the sup LM test and calculate asymptotic critical values and p-values.

Testing for whether the lag length is greater than 1 is important because this reveals information relevant to the proper error-correction structure in the threshold vector error-correction model (VECM) we will use this to explore the short-run dynamics. As noted by Krishnakumar et al. (2009), if the lag length is only equal to 1, then the threshold structure appears in the error-correction term only in the threshold VECM. While if the lag length is greater than 1 the threshold structure extends to all the lagged dependent variables including the error-correction term. One way to test this is to check whether a second lag improves

the fit in the model described by equation 3.

Once the presence of threshold cointegration relationship is confirmed, we proceed to estimate the threshold VECM to analyze the short-run dynamics using either  $\hat{\mu}_{t-1}$  or  $\Delta\hat{\mu}_{t-1}$  as the threshold variable. There are two alternative ways of estimating a threshold vector error correction model (TVECM). We can estimate the TVECM either by considering the threshold effect in the error correction term or we can estimate a model with a threshold effect in all the dependent variables. We use the typical Akaike information criterion (AIC) to select the appropriate lag lengths for the models described below.

The model below is the case where we consider threshold effect in all the lagged dependent variables:

$$\begin{aligned} \Delta LD_t &= D_t[\alpha_{d,1} + \rho_{d,1}\hat{\mu}_{t-1} + \sum_{l=1}^p \beta_{d,1,d,l}\Delta LD_{t-p} + \sum_{l=1}^p \beta_{d,1,d,l}\Delta LB_{t-p}] + \\ (1 - D_t)[\alpha_{d,0} + \rho_{d,0}\hat{\mu}_{t-1} + \sum_{l=1}^p \beta_{d,0,d,l}\Delta LD_{t-p} + \sum_{l=1}^p \beta_{d,0,d,l}\Delta LB_{t-p}] + \varepsilon_{d,t} \end{aligned} \quad (2.7)$$

$$\begin{aligned} \Delta LB_t &= D_t[\alpha_{b,1} + \rho_{b,1}\hat{\mu}_{t-1} + \sum_{l=1}^p \beta_{b,1,d,l}\Delta LD_{t-p} + \sum_{l=1}^p \beta_{b,1,d,l}\Delta LB_{t-p}] + \\ (1 - D_t)[\alpha_{b,0} + \rho_{b,0}\hat{\mu}_{t-1} + \sum_{l=1}^p \beta_{b,0,d,l}\Delta LD_{t-p} + \sum_{l=1}^p \beta_{b,0,d,l}\Delta LB_{t-p}] + \varepsilon_{b,t} \end{aligned} \quad (2.8)$$

$$D_t = \begin{cases} 1, & \text{if } \Delta\hat{\mu}_{t-1} \geq \tau \\ 0, & \text{if } \Delta\hat{\mu}_{t-1} < \tau \end{cases} \quad (2.9)$$

where  $\alpha_{j,0}$ ,  $\alpha_{j,1}$ ,  $\rho_{j,1}$ ,  $\rho_{j,0}$ ,  $\beta_{j,d,i}$  and  $\beta_{j,b,i}$  for  $j=d,b$  are the parameters of the model,  $l = 1, \dots, p$  refer to the lag length and  $\varepsilon_{j,t}$  are the error terms. In the specification above, the subscripts make use of the following mnemonics. The first subscript indicates which equation the parameter or error term is from, the second subscript in the  $\rho_{j,1}$  and  $\rho_{j,0}$  refer to the dummy representing the state described earlier. The third and fourth subscripts attached

to the lagged difference term correspond to the type of variable that is differenced and the lag length respectively.

We can alternatively compute the following model instead, where threshold effect is observed only in the error correction term only and not on the other parameters:

$$\Delta LD_t = \alpha_d + \rho_{d,1}D_t\hat{\mu}_{t-1} + \rho_{d,0}(1 - D_t)\hat{\mu}_{t-1} + \sum_{l=1}^p \beta_{d,1,d,l}\Delta LD_{t-p} + \sum_{l=1}^p \beta_{d,1,d,l}\Delta LB_{t-p} + \varepsilon_{d,t} \quad (2.10)$$

$$\Delta LB_t = \alpha_b + \rho_{b,1}D_t\hat{\mu}_{t-1} + \rho_{b,0}(1 - D_t)\hat{\mu}_{t-1} + \sum_{l=1}^p \beta_{b,1,d,l}\Delta LD_{t-p} + \sum_{l=1}^p \beta_{b,1,d,l}\Delta LB_{t-p} + \varepsilon_{d,t} \quad (2.11)$$

$$D_t = \begin{cases} 1, & \text{if } \Delta\hat{\mu}_{t-1} \geq \tau \\ 0, & \text{if } \Delta\hat{\mu}_{t-1} < \tau \end{cases} \quad (2.12)$$

## 2.4 Results and Analysis

The following sub-sections summarize the estimation results starting with stationarity tests, followed by the cointegration and the threshold cointegration tests. Finally, we present the threshold VECM estimates.

### 2.4.1 Unit Root Test

We begin our analysis with stationarity tests. Figure 2.2 indicate the presence of potential structural breaks in the data. We will pursue stationarity tests for known and unknown structural break points in the data. Table 2.1 below furnishes the stationarity test results using augmented [Dickey and Fuller \(1981\)](#) and [Perron \(1989\)](#) tests on the level data. These

results indicate both variables are non-stationary at levels as we can not reject the null of unit root in the data <sup>3</sup>. We note that the augmented Dickey-Fuller test statistics as well as the Perron test statistics in the Table 2.1 indicate that intercept and trend are important in the procedure. The trend and the intercept are prone to structural breaks. The power of the tests presented below in Table 2.1 are susceptible to the presence of structural break, and the Perron (1989) tests assume the date of the structural breaks are known Enders (2010). Table 2.2 furnishes Lumsdaine and Papell (1997) and Lee and Strazicich (2003) tests which allow us to test for stationarity with unknown structural breaks. These results also confirm<sup>4</sup> that the variables are non-stationary or have a unit root at level.

Table 2.1: Unit Root Test: Level

Test	$LB_t$	$LD_t$
ADF Test: no Trend and Intercept	5.12	3.02
ADF Test: with Intercept	1.89	0.96
ADF Test: with Trend and Intercept	-0.63	-2.60
Phillips-Perron Test: with Intercept	2.21	-0.63
Phillips-Perron Test: with Intercept and Trend	-0.06	-1.36

Table 2.2: Unit Root Test: Unknown Structural Break

Test	$LB_t$	$LD_t$
Lumsdaine-Papell Test: Break in Intercept	-2.92	-5.02
Lumsdaine-Papell Test: Break in Trend	-5.59	-5.55
Lumsdaine-Papell Test: Break in both	-5.60	-6.47
Lee-Strazicich Test: Crash Model	-1.63	-3.18
Lee-Strazicich Test: Break Model	-3.08	-3.27

Tables 2.3 below summarize the stationarity tests for both the variables in their first differences. We use the variables in their natural log, thus we can interpret the first difference as

<sup>3</sup>ADF test 5% critical value are for no intercept and trend model -1.94, intercept only model -2.87, and both intercept and trend model -3.42. Phillip Perron test 5% critical value are for intercept only model -2.87, and both intercept and trend model -3.42.

<sup>4</sup>Lumsdaine-Papell Test 5% critical values are -6.16 for breaks in intercept, -6.62 for breaks in trend, and -6.75 for break in both intercept and trend.



their respective growth rates. We investigated the plot of the growth rates for both variables, which do not indicate the presence of a trend in the first differences. Hence, we will not discuss the case for a trend in our analysis of stationarity presented below in Table 2.3. Results for all the tests indicate the variables are stationary in their first difference<sup>5</sup> i.e. they are  $I(1)$ .

Table 2.3: Unit Root Test: First Difference

Test	$LB_t$	$LD_t$
ADF Test: no Trend and Intercept	-5.67**	-4.93**
Phillips-Perron Test: with Intercept	-15.20**	-27.17**
Lumsdaine-Papell Test: Break in Intercept	-8.81**	-6.33**
Lee-Strazizich Test: Crash Model	-3.81	-4.18

## 2.4.2 Cointegration and Threshold VECM

In order to investigate the presence of a long-run relationship or cointegration between the two variables, we need them to be non-stationary at the level and be first order difference stationary. Since our findings discussed above confirm that our variables are  $I(1)$ , we can proceed to test for cointegration or a long-run relationship between the two. The following furnishes the results from the cointegration tests between the natural logs of debt and the natural log monetary-base. Results in Table 2.4 display the [Johansen \(1991\)](#) and the [Phillips and Ouliaris \(1990\)](#) procedures, which indicate that we cannot find evidence in support of cointegration or a long-run relationship between the two<sup>6</sup>. We can not reject the null of no cointegration for these test at 10%, 5% and 1% significance levels for both these tests.

<sup>5</sup>ADF test 5% critical value is -1.93, Phillips Perron Test is -2.86, Lumsdaine-Papell test is -6.16.

<sup>6</sup> Critical values are 14.90, and 8.18 for the Eigenvalue Test, and 17.95, and 8.18 for the Trace Test at 5%, Critical values is 33.71 for  $P_U$ , and s 55.22  $P_Z$  tests at 5%.

Table 2.4: Cointegration Test

Johansen-Juselius (1990) Procedure		
	$\lambda_{max}$	$\lambda_{trace}$
$r \leq 1$	0.81	0.81
$r = 0$	8.15	8.96
Phillips-Ouliaris (1990) Procedure		
Test type: $P_U$ :	6.67	
Test type: $P_Z$ :	3.29	

Since, the tests above do not confirm evidence of cointegration we proceed to investigate for the presence of threshold cointegration. Following the [Engle and Granger \(1987\)](#) procedure, in our first step, we compute the following long-run relationship as per equation 2.1 described earlier:

$$LD_t = 1.37LB_t + \mu_t \quad (2.13)$$

In the second step, we obtain the residuals  $\hat{\mu}_t$  and conduct unit root test on the residuals. The residual have a unit root as per the [Dickey and Fuller \(1981\)](#) tests. This implies that there is no evidence of cointegration between the two. The linear cointegration test may fail to reject the null of no cointegration due to presence of possible nonlinearity in the adjustment. [Siklos and Granger \(1997\)](#) and [Balke and Fomby \(1997\)](#) find evidence in favor of asymmetric adjustment. We first conduct a RESET test (described in equation 2.2) and cannot reject the null  $H_0 : \beta_3 = \beta_4 = 0$  for the linearity. This linearity test enables us to check for nonlinear alternatives. Following the [Enders and Granger \(1998\)](#) threshold unit root we conduct the threshold unit root test on the obtained residuals. In this test, we consider both TAR and momentum-TAR (M-TAR) adjustments for the threshold variable. Results below, Table 2.5, indicate that we can reject the case for unit root for momentum-TAR. The momentum-TAR model with a lag length of 9<sup>7</sup> implies we should apply threshold effect in all the right hand side variables in our threshold vector error-correction model. This result imply that the residual is stationary with a threshold effect. We now proceed with the formal threshold

<sup>7</sup> 5% critical values are 6.12 for TAR model, and 5.54 for M-TAR model

cointegration tests described earlier. Table 2.6 below presents the results for the Hansen and Seo (2002) and Enders and Siklos (2001) tests. The Enders and Siklos (2001) test, as well as the Hansen and Seo (2002) test, confirm threshold cointegration between the two<sup>8</sup>. The corresponding threshold values for both these tests have the same sign and similar magnitude.

Table 2.5: Enders-Granger Threshold Unit Root test

Model	Threshold Unit Root	Lags
TAR Model	2.54	9
Momentum TAR Model	5.58**	9

Table 2.6: Threshold Cointegration Test

Test	Test Statistic	Threshold
Enders and Siklos Test (2002)	12.45**	-0.0156
Hansen and Seo Test (2002)	32.45**	-0.0659

The results presented above confirm the presence of threshold cointegration or a threshold long-run relationship between the natural log of government debt and the natural log of monetary-base in the U.S.

As discussed earlier, the evidence of a long-run relationship confirmed through formal statistical procedure enables us to explore the short-run dynamics between the two. The residuals from equation 2.1 provide us the disequilibriums or deviation away from the long-run relationship. These residuals define our error-correction term. We use these residuals in our threshold vector error-correction model (TVECM) to analyze the short-run dynamics between the two. We use the first difference of the error-correction term as our endogenous threshold variables to define the dummy for the TVECM model below.

<sup>8</sup> 5% critical value for the Enders and Siklos is 6.63, and Hansen and Seo is 31.60.

$$D_t = \begin{cases} 1, & \text{if } \Delta\hat{\mu}_{t-1} \geq -0.0152 \\ 0, & \text{if } \Delta\hat{\mu}_{t-1} < -0.0152 \end{cases} \quad (2.14)$$

As discussed earlier, we estimate equations 2.7 and 2.8, where the threshold effect is applied to all the right-hand side variables. The [Enders and Granger \(1998\)](#) threshold Unit root test uses a lag length of 9, and as per [Krishnakumar et al. \(2009\)](#), the above specification is our preferred one. We will present and discuss the estimation output from the equations 2.10 and 2.11 as well to check the robustness of our findings. We will use a lag length of 3 as per AIC for the estimation using the full sample.

The results in Table 2.7 and 2.8 below present the lagged error-correction terms for both growth in government debt ( $\Delta LD_t$ ) and growth in monetary-base ( $\Delta LB_t$ ) equations. Table 2.7 present the case where threshold effect is applied to all the right-hand side variables, and Table 2.8 present the case where threshold effect is applied to the error-correction term only. For the purpose of brevity as well as for relevance we will present and discuss part of the estimation output with the error-correction terms<sup>9</sup>. The results indicate that government adjusts their debt whenever the error correction term deviates from the long-run equilibrium and the Fed does not accommodate debt. We denote the periods where the difference in the error correction terms are above the threshold as regime 1 ( $\Delta ECT_{t-1} = \Delta LD_{t-1} - 1.3745\Delta LB_{t-1} - 0.0152$ ). Regime 2 refer to the periods when the difference in the error correction term are below the threshold ( $\Delta ECT_{t-1} = \Delta LD_{t-1} - 1.3745\Delta LB_{t-1} < -0.0152$ ). In regime 1, when government debt growth is larger than the growth in monetary-base, we observe that the government reduce debt growth as the speed of adjustment has a value of -0.002 (the coefficient of the lagged error-correction term). It is statistically significant as well. The Fed, on the other hand, reduces growth in monetary-base. The speed of adjustment parameter has a value of -0.0005, but it is not statistically significant. Though it is not significant, the sign implies that the Fed tightens the base instead of accommodating

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<sup>9</sup>The complete estimation output is presented at the appendix.

the debt growth. We find similar results when we consider threshold effect in the error-correction term only presented in Table 2.8. In regime 2, when government debt growth is smaller than the growth in monetary-base, we observe that the government increase debt growth as the speed of adjustment has a value of 0.007 (the coefficient of the lagged error-correction term). It is statistically significant as well. The Fed, on the other hand, increase growth in monetary-base. The speed of adjustment parameter has a value of 0.005, but it is not statistically significant. We find similar results when we consider threshold effect in the error-correction term only presented in Table 2.8.

Table 2.7: TVECM: Error Correction Term (Threshold effect in All)

$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.37\Delta LB_{t-1}) \geq -0.0152$		
	$\Delta LD_t$	$\Delta LD_t$
$ECT_{t-1}^H$	<b>-0.002**</b> (0.003)	-0.0005 (0.001)
$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.37\Delta LB_{t-1}) < -0.0152$		
$ECT_{t-1}^L$	<b>0.007**</b> (0.002)	0.005 (0.004)

Notes: 1. Eicker-White heteroskedasticity consistent standard error  
2. Approximately 84.72% observation in higher regime, and 15.28% in lower regime.

Table 2.8: TVECM: Error Correction Term (Threshold effect in the ECT only)

$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.37\Delta LB_{t-1}) \geq -0.0152$		
	$\Delta LD_t$	$\Delta LD_t$
$ECT_{t-1}^H$	<b>-0.002**</b> (0.0008)	-0.002 (0.001)
$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.37\Delta LB_{t-1}) < -0.0152$		
$ECT_{t-1}^L$	<b>0.007**</b> (0.002)	0.001 (0.003)

Notes: 1. Eicker-White heteroskedasticity consistent standard error  
2. Approximately 84.72% observation in higher regime, and 15.28% in lower regime.

Figures 2.3 and 2.4 plots the first difference in the error correction term and the threshold,

where we have two regimes: above and below the threshold. The shaded areas represent the National Bureau of Economic Research recession dates. For the purpose of visual inspection, we split the figure into two time periods: <sup>10</sup> 1942:1 to 1976:12 and 1977:1 to 2015:12. We observe the frequency at which regime 2 (below the threshold) appear is relatively fewer times until 1976. With rising debt to GDP starts in the subsequent period, we observe the frequency of regime 2 increases substantially. The magnitude of the variation in the difference in the error-correction terms is visibly more during this period. But, we do not find a systemic pattern with the recessions. As per figures 2.3 and 2.4, we observe a lot of volatility and regime-switching appearing during the World War II years and post-financial crisis years. The period from post-World War II to the early part of the 1970s, we observe regime 1 more prevalent, where the scale of the first difference in the error correction term indicates relatively lesser volatility. As in figure 2.1, we observe that during this period debt to GDP ratio in the U.S. has been steadily falling. Starting from the early 1980s we start observing the scale in the first difference in the error correction term show relatively more volatility. During this era, we also observe the frequency at which regime 2 appears increase relative more. As in figure 2.1, we also observe the debt to GDP ratio in the U.S. start to rise again after about four decades. [Arestis et al. \(2004\)](#) also find the presence of regimes with the federal budget deficits and report similar findings for the U.S.

Based on the results discussed above, we find no evidence of the Fed accommodating government debt or debt financing. These findings remain consistent with both the specifications in the threshold structure. Figure 2.3 and 2.4, we observe there are a lot of volatility in during the World War II periods and post-financial crisis years. The prevalence of regime 2 during the period from 1946 to 1966 are relatively few and mostly coincide with recessions. The frequency at which regime 2 appear is much higher since the early 1970s. They also appear during the first part of Volckers tenure as the Fed Chair. We also observe the regime 2 to appear in periods on or after recessions.

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<sup>10</sup>1977 onwards we observe the debt to GDP ratio in the U.S. starts to rise again after almost three decades of continuous decline.

Figure 2.3: Difference in Error Correction Term, Recessions, and Threshold, 1942-1976

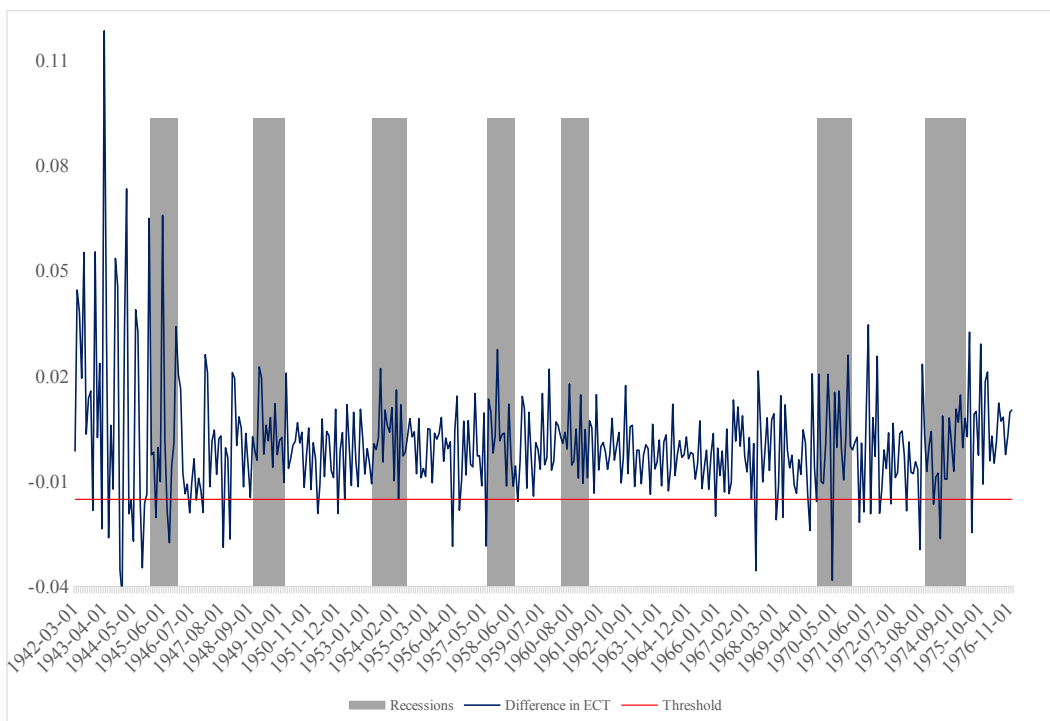
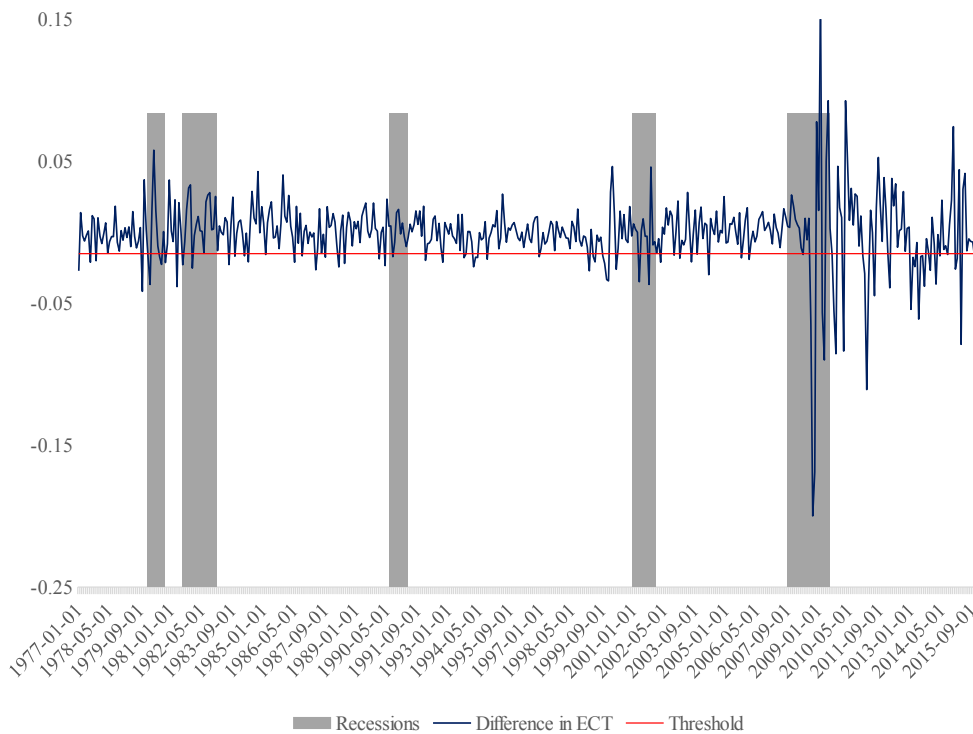


Figure 2.4: Difference in Error Correction Term, Recessions, and Threshold, 1977-2015



### 2.4.3 Sensitivity Analysis

In this section, we proceed to check the validity of our findings over two sub-samples. The full sample ranges from 1942 to 2015, which include World War II years and the financial crisis years. The findings presented earlier are likely to have influence from these two events. We observed a lot of volatility in the first difference of the error-correction term (see Figures 2.3 and 2.4). Government budget deficits and subsequent rise in debt were necessitated due to some special circumstances. Hence, it is imperative to investigate the sensitivity of our results and findings over sub-samples. As such we create two sub-samples: sub-sample 1 comprises of data from 1946 to 2015, and sub-sample 2 comprises of data from 1946 to 2007. The results below present and discuss the threshold cointegration and threshold VECM results for these two sub-samples.

Table 9 below present the Ender and Granger Threshold Unit Root Test. We can reject the presence of a unit root for the sub-sample 1 but not for the sub-sample 2<sup>11</sup>. On the other hand, the formal threshold cointegration tests in Table 10, indicate there is evidence of threshold cointegration between the natural log of government debt and the natural log of monetary-base in both the sub-samples<sup>12</sup>.

The estimation output of the threshold vector error correction model for sub-sample 1 has the same findings as the full sample. We observe the government to adjust debt growth but no evidence of monetary authority adjusting monetary-base. However, estimation output of the threshold vector error correction model for sub-sample 2 gives findings which differ from the full sample as well as sub-sample 1.

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<sup>11</sup>5% critical values are 6.12 for TAR model, and 5.54 for M-TAR model.

<sup>12</sup>The Enders and Siklos test 5% critical value is 6.63, and the Hansen and Seo critical values are 31.32 and 32.37.



Table 2.9: Threshold Unit Root Test: Sub-samples

Sub-sample: 1946 to 2015		
Model	Threshold Unit Root Test	Lags
TAR	2.88	15
Momentum TAR	8.13**	15
Sub-sample: 1946 to 2007		
TAR	2.80	15
Momentum TAR	1.95	15

Table 2.10: Threshold Cointegration Test: Sub-samples

Threshold Cointegration Test	Test Statistic	Threshold
Sub-sample: 1946 to 2015		
Enders and Siklos Test (2002)	14.10**	-0.0146
Hansen and Seo Test (2002)	59.57**	-0.0116
Sub-sample: 1946 to 2007		
Enders and Siklos Test (2002)	9.27**	0.0108
Hansen and Seo Test (2002)	62.37**	-0.2528

The results in Table 2.11 and 2.12 below present the lagged error-correction terms for both growth in government debt ( $\Delta LD_t$ ) and growth in monetary-base ( $\Delta LB_t$ ) equations. Table 2.11 present the case where threshold effect is applied to all the right-hand side variables, and Table 2.12 present the case where threshold effect is applied to the error-correction term only. For the purpose of brevity as well as for relevance we will present and discuss part of the estimation output with the error-correction terms<sup>13</sup>. The results indicate that government adjusts their debt whenever the error correction term deviates from the long-run equilibrium and the Fed does not accommodate debt. We denote the periods where the difference in the error correction terms are above the threshold as regime 1 ( $\Delta ECT_{t-1} = \Delta LD_{t-1} - 1.42\Delta LB_{t-1} - 0.0106$ ). Regime 2 refer to the periods when the difference in the error correction term are below the threshold ( $\Delta ECT_{t-1} = \Delta LD_{t-1} - 1.42\Delta LB_{t-1} < -0.0106$ ).

<sup>13</sup>The complete estimation output is presented in Appendix C.

Table 2.11: TVECM: Error Correction Term (Threshold effect in All)

	$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.42\Delta LB_{t-1}) \geq -0.0106$	
	$\Delta LD_t$	$\Delta LD_t$
$ECT_{t-1}^H$	<b>-0.003**</b> (0.001)	<b>-0.002**</b> (0.0009)
	$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.42\Delta LB_{t-1}) < -0.0106$	
$ECT_{t-1}^L$	<b>0.009**</b> (0.003)	0.001 (0.002)

Notes: 1. Eicker-White heteroskedasticity consistent standard error  
 2. Approximately 77.78% observation in higher regime, and 22.22% in lower regime.

Table 2.12: TVECM: Error Correction Term (Threshold effect in the ECT only)

	$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.42\Delta LB_{t-1}) \geq -0.0152$	
	$\Delta LD_t$	$\Delta LD_t$
$ECT_{t-1}^H$	-0.002 (0.001)	<b>-0.001*(0.0009)</b>
	$\Delta ECT_{t-1} = (\Delta LD_{t-1} - 1.42\Delta LB_{t-1}) < -0.0152$	
$ECT_{t-1}^L$	<b>0.007**</b> (0.004)	0.001 (0.003)

Notes: 1. Eicker-White heteroskedasticity consistent standard error,  
 2. Approximately 77.78% observation in higher regime, and 22.22% in lower regime.

In regime 1, when government debt growth is larger than the growth in monetary-base, we observe that the government reduce debt growth as the speed of adjustment has a value of -0.003 (the coefficient of the lagged error-correction term). It is statistically significant as well. The Fed, on the other hand, reduces growth in monetary-base. The speed of adjustment parameter has a value of -0.002, which is statistically significant. The sign implies that the Fed tightens the monetary base instead of accommodating the debt growth. We find similar results when we consider threshold effect in the error-correction term only presented in Table 2.8. In regime 2, when government debt growth is smaller than the growth in monetary-base, we observe that the government increase debt growth as the speed of adjustment has a value

of 0.009 (the coefficient of the lagged error-correction term). It is statistically significant as well. The Fed, on the other hand, increase growth in monetary-base. The speed of adjustment parameter has a value of 0.001, but it is not statistically significant. We find similar results when we consider threshold effect in the error-correction term only presented in Table 2.12.

The sensitivity analysis using two sub-samples again conform to our findings for the full sample. We find formal statistical evidence supporting the presence of a threshold long-run relationship between the natural log of government debt and the natural log of monetary base. Threshold VECM models show that in the short-run government adjust the debt and there is no evidence of accommodation by the Fed or debt monetization.

## 2.5 Conclusion

Empirical evidence of fiscal policy influencing monetary policy is scarce, even though theoretical papers in the literature in recent times consider such cases. In this paper, we investigate to find formal statistical evidence of a long-run relationship between the monetary-base and government debt using monthly data from 1942:1 to 2015:12 for the U.S. in an attempt to find empirical evidence of fiscal policy influencing monetary policy.

The linear cointegration procedures fail to detect evidence of a long-run relationship. However, threshold cointegration procedures developed by [Enders and Siklos \(2001\)](#) and [Hansen and Seo \(2002\)](#) find evidence in favor of a threshold long-run relationship between the government debt and the monetary-base in the U.S. In addition to the full sample, we use two sub-sample to investigate if the evidence for the threshold long-run relationship or threshold cointegration are sensitive to sample periods, excluding the World War II and financial crisis years. As one can argue that the long-run relationship is a more recent phenomenon. We find evidence of a long-run relationship in both cases. As for the short-run dynamics using the threshold vector error correction model postulate that the monetary-base

do not adjust to changes in the government debt in any of the regimes. Rather the fiscal authorities adjust debt level in both regimes. Thus showing no evidence of regime-based monetary accommodation to fiscal debt in the U.S. These findings remain consistent over the sub-samples. The sample excluding the World War II and financial crisis era indicate at times of rapid growth in the government debt, the Fed tightens the monetary-base growth. We do not find evidence of debt monetization in the U.S. Thus, the findings above reinforcing the argument in favor of independent or autonomous Fed.

# Chapter 3

## Response of short-run interest rate and inflation to budget deficits and money supply: A nonlinear approach

### 3.1 Introduction

In theory, an increase in budget deficits leads to a rise in interest rate as the supply of loanable funds fall given demand. A monetary expansion, on the hand, leads to a fall in the interest rate. During recessions, we observe budget deficit to increases due to expansionary fiscal policy, and central banks adopt an expansionary monetary policy. Movement of real interest rate with budget deficits and monetary policy is integral to contemporary macroeconomic discourse. Theory relate these two in explaining the goods market as well as the money market outcomes, and also we use them to predict outcomes of policy intervention. Business cycle necessitates macroeconomic policy interventions, and, both monetary and fiscal policies impact inflation and interest rate. In assessing success (or failure) of a policy intervention, economists often refer to the behavior of economic agents. Economic agents decision on labor supply and consumption, and, their expectation formation subsequently impact their resource allocation decisions. In theory, these decisions are linked to inflation

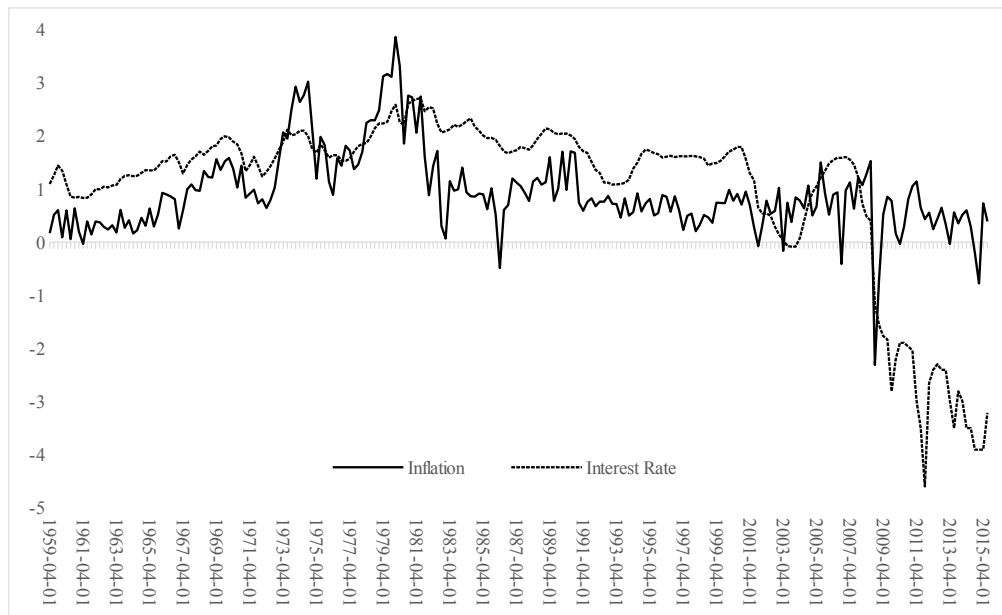
and interest rate for optimizing economic agents.

Macroeconomic impact of the budget deficits is an area of concern for academics as well as for policy makers. Budget deficits, arising either from tax cuts or increased government purchases, vary with the business cycle, and the Great Recession made the budget deficits the core of a heated political debate. Analyzing the likely impact of the budget deficits on inflation and the short-run interest rate is thus imperative. Similarly, theory relates money supply directly to inflation and short-run interest rates, and it is very common to observe monetary policy interventions precede fiscal policy. Analyzing the potential impact of the budget deficits on inflation and interest without considering money supply will lead to erroneous conclusions ([Laubach, 2009](#)).

U.S. data on inflation and short-run interest rates exhibit variability with the business cycle, thus are likely to respond in an asymmetric manner to fiscal deficit and money supply (see Figures 1 and 2). We may also discover cyclicity (or counter-cyclical behavior) in their behavior. Policy makers use both expansionary fiscal policy and monetary policy to stabilize the economy during a recession. We also observe the use of contractionary monetary policy during economic expansions. As such, the use of macroeconomic stabilization policy has a nonlinear dimension in their implementation corresponding to the state of the economy. We also observe a simultaneous nature in the use of such policies during recessions, usually, monetary policy precedes fiscal policy. Use of expansionary fiscal policy often leads to larger budget deficits. We also observe variations in the growth of money supply during recessions as part of the expansionary monetary policy. This paper intends to explore the dynamics between inflation and short-term interest rate with respect to a change in budget deficits and money supply during a recession as opposed to an expansion. The principal contributions of this paper are as follows: (1) in this paper, we investigate the impact of growth in budget deficits and money supply on inflation and short-term interest rate in the presence of business cycles. We will use a threshold VAR model to incorporate business cycle into our econometric specification. (2) The movement in inflation and short-term interest rate allow

us to infer potential movement of real interest in an asymmetric framework, and (3) the sample period in our analysis includes the financial crisis, thus giving us a more up-to-date analysis. Our results indicate both inflation and short-term interest rate exhibit asymmetry as well as cyclicality. Time path for real interest (inferred from the movements of inflation and short-term interest rates) is also indicative of asymmetry.

Figure 3.1: Inflation and Short-term Interest Rate in the U.S.



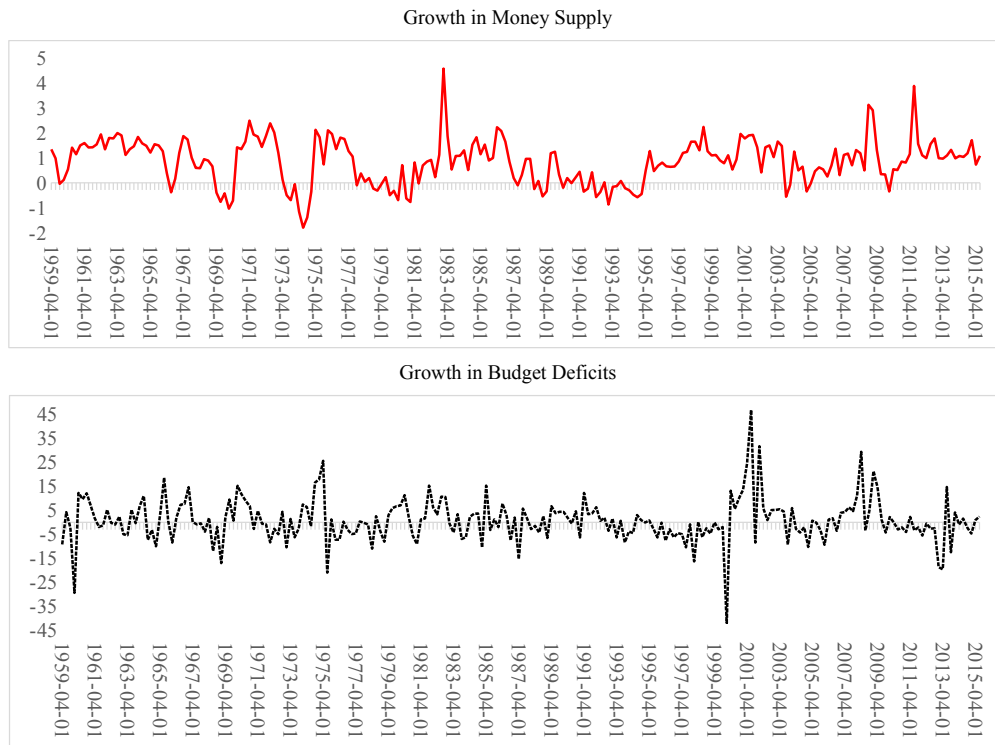
Note: 1. Quarterly data, 1959:Q1 to 2015:Q3  
 2. We use CPI to measure inflation and 3 month Treasury bill rate as a measure of short-term interest rate

The rest of the paper is organized as follows: Section 3.2 discusses the relevant literature, Section 3.3 elaborates the data and methodology, Section 3.4 presents the results and analysis, and finally Section 3.5 present the conclusion.

## 3.2 Literature Review

We can find a plethora of empirical literature investigating the impact the budget deficits on the interest rate. Most of the studies use a measure of real interest rate or use the Treasury bill rates in their time series econometric specification to explore the relationship, and the

Figure 3.2: Growth in Money Supply and Budget Deficit in the U.S.



findings are often contradictory in nature. There are studies that explore the impact of monetary policy on real interest rates. It is worth noting that there is a gap in empirical literature where the impact on real interest rate is explored while considering the budget deficits and money supply together in the econometric specification.

In a simple Keynesian framework, an increase in government spending increases interest rates, leading to a reduction in private spending through the crowding-out effect. Contrary to the conventional Keynesian view, the expansionary fiscal contraction hypothesis suggests that fiscal contractions can, through their impact on expectations, lead to growth in consumption and investment. In this hypothesis, a large or persistent fiscal contraction follows a previous expansionary fiscal stance, and signal the governments adjustment that has been delayed [Barry and Devereux \(1995\)](#); [Perotti \(1999\)](#); [Sutherland \(1997\)](#). [Alesina and Perotti \(1997\)](#) using cross-country data find evidence of such episodic fiscal contractions are more likely to happen in the economies that need a fiscal adjustment. In a neoclassical framework,



the effects of government spending stem mainly from its crowding-out effect and a wealth effect. The wealth effect arises because increases in government spending today imply increases in future taxes to balance the budget or reduce the deficit. The resulting fall in wealth reduces consumer demand, increases labor supply, and lowers interest rates [Barry and Devereux \(1995\)](#); [Devereux and Love \(1995\)](#), and the increase in labor supply, in turn, increases the marginal productivity of capital and spur investment [Burnside et al. \(2004\)](#); [Ramey and Shapiro \(1998\)](#). The size of the wealth effect depends on whether the change in government spending has purely transitory or persistent effects. A common theme that arises from the aforementioned literature- budget deficits are a significant aspect.

[Makin \(1983\)](#) uses quarterly data from 1959 to 1981 for 3 month Treasury bill finds limited or no evidence of an increase in the interest rate due to the budget deficits. [Wachtel and Young \(1987\)](#) report the existence of an empirical link between interest rates and future budget deficits. [Evans \(1987\)](#) on the other hand, finds no evidence of budget deficits causing interest rates to rise using three statistical techniques and argue that Ricardian equivalence may explain as to why larger budget deficits are not associated with higher interest rates. [Mankiw \(1987\)](#) argue that an increase in government spending depresses the real interest rate because wealth effect reduces private consumption and increases the marginal utility of consumption. This, in turn, lowers the marginal rate of substitution and thus marginal productivity of capital. [Engen and Hubbard \(2004\)](#) explore the impact of the government debt and the budget deficit on 10 year Treasury bill rate using U.S. data from 1953 to 2004 using a structural VAR framework. Their impulse responses indicate debt and budget deficit increases the long-term interest rate measured with 10 year Treasury bill rate. [Terzi \(2007\)](#) finds no evidence of any linkage between the budget deficit and interest rate for the U.S. and Europe. [Laubach \(2009\)](#) finds interest rate to rise with government debt and budget deficits. This study shows that statistically significant and economically plausible estimates of the effects of budget deficits and debt on interest rates can be obtained by focusing on long-horizon forward rates and projections of the budget deficits or debt.

Choi (1999) finds that interest rate response to money supply shock varies with the state of the monetary policy stance. In this study author defines three states in monetary policy- tight, loose and neutral, based on an estimated policy stance index. Threshold VAR estimates suggest liquidity and expected inflation effect vary across states, where liquidity effect is dominant in the loose and neutral states but a reversed liquidity effect in a tight regime. Choi and Devereux (2006) explore the asymmetric effect of government spending and as to how the real interest rate is relevant. They find the effect of fiscal policy depend on the level of real interest rates since Ricardian Equivalence effect is smaller at lower financing costs of the fiscal policy. Choi and Holmes (2014) investigate the relevance of Ricardian equivalence in explaining the relationship between the budget deficit and real interest rate using two centuries of annual U.S. data. They estimate a Markov regime-switching model that shows that the U.S. economy switches between Ricardian equivalence regime, characterized by an insignificant relationship between adjusted primary budget deficit and real long-term interest rates, and a regime characterized by the traditional view of a positive relationship.

There are panel data studies which investigate the impact of budget deficits on inflation in a cross-country setup. Dwyer (1982) report a lack of evidence that higher current or past budget deficits lead to inflation. Recent analyses of cross-country data suggest that the positive association between fiscal deficits and inflation is strong among high-inflation and developing countries but not among low-inflation and industrial economies (Catao and Terrones, 2005; Fischer et al., 2002).

Lack of consensus could potentially arise from governments asymmetric policy preference. Choi (1999) and Choi and Holmes (2014) address the non-linearity by using threshold VAR and Markov regime-switching model. But a key gap arises as none of the earlier studies consider the budget deficits and money supply together in their empirical framework. In addition, the financial crisis era is characterized by substantial changes in both fiscal and monetary policy actions. We observed a rapid rise in the budget deficits on the fiscal side. Short-term Interest rate stayed at the zero lower bound and the Fed also adopted quantita-

tive easing. We employ a regime switching threshold VAR to investigate as to how money supply and budget deficit influence short-run interest rate and inflation.

### 3.3 Data and Methodology

In this paper, we use U.S. quarterly data from 1959: Q2 to 2015: Q3 on inflation, interest rate, the percentage change in the budget deficit and percentage change in money supply. Real GDP, deflator, and government spending and receipt data are collected from the National Income and Product Account (NIPA) tables. The money supply and the Consumer Price Index (CPI) data are collected from Federal Reserve Bank of St. Louis FRED database. All the variables used in this paper are in their natural log.

We use the following definitions for measuring deficit ([Abel et al., 2014](#)). Primary deficit and primary current deficit shows by how much government receipts fall short of the spending outlays.

- Gross Deficit = Government Outlays - Government Receipts = (Government Purchase + Transfer and Subsidies + Net Interest Payment) - Government Receipts
- Current Deficit = (Gross Deficit + Government Investment Expenditure) - Government Receipts
- Primary Current Deficit = Current Deficit - Net Interest Payment

We will estimate a 5 variable endogenous threshold vector autoregression (TVAR) model with the output gap as our endogenous threshold variable. There are different ways to measure the output gap in the existing literature. For example, one can use the Hodrick-Prescott (HP) filter ([Hodrick and Prescott, 1997](#)). However, [Mise et al. \(2005\)](#) argue that such filters are not adequate to fully capture the output gap as the HP filter produces series with spurious dynamic relations that have no basis in the underlying data-generating process.

Following [Cassou et al. \(2015\)](#) output gap is defined as the deviation of the observed annual output growth rate from its long-term average. For the output gap, we computed the difference between the observed annual growth rate and the average growth rate over the sample period of 1959 to 2015. In particular, we computed the growth rate in percentage terms by multiplying 100 times the log difference between the current value of real GDP and the value four quarters earlier. Next, these growth rates were averaged and then the average was subtracted from the annual growth rate series to give a series that has positive values when the current growth rate exceeds the average and negative values when the growth rate is below the average.

In this paper, the endogenous threshold VAR is estimated following the method proposed by [Tsay \(1998\)](#) and [Lo and Zivot \(2001\)](#). Authors' generalize the univariate and single equation estimations by [Tong \(1990\)](#), [Chan \(1993\)](#) and [Hansen \(1996\)](#) for a multivariate VAR. We will use an endogenous threshold VAR (TVAR) model and compute the nonlinear impulse responses to explore the dynamics between debt and deficit. Why endogenous threshold VAR? [Afonso et al. \(2011\)](#) note that the endogenous TVAR model has a number of interesting features. First, it is a relatively simple way to capture possible nonlinearities such as asymmetric reactions to shocks or the existence of multiple equilibria. As the effects of the shocks are allowed to depend on the size and the sign of the shock, and also on the initial conditions, the impulse response functions are no longer linear, and it is possible to distinguish, for instance, between the effects of fiscal developments under different economic states. Second, another advantage of the TVAR methodology is that the variable by which different regimes are defined itself is an endogenous variable included in the VAR. Therefore, this makes it possible that regime switches may occur after the shock to each variable. In particular, the fiscal policy shock might either boost the output or increase the budget deficit that harms the prospects of economic growth, and the overall effect to the economy of a fiscal expansion might become negative. We can write an endogenous threshold VAR (of lag order  $p$ ) specified in the following manner. We will use the Akaike Information Criteria (AIC) to select the appropriate lag lengths.

$$\begin{aligned}
Inf_t &= I_t \left[ \alpha_L + \sum_{i=1}^p \alpha_i Int_{t-i} + \sum_{j=1}^p \alpha_j \Delta MS_{t-j} + \sum_{k=1}^p \alpha_k \Delta Def_{t-k} + \sum_{l=1}^p \alpha_l YG_{t-l} \right] + \\
(1 - I_t) &\left[ \alpha_H + \sum_{i=1}^p \alpha_i Int_{t-i} + \sum_{j=1}^p \alpha_j \Delta MS_{t-j} + \sum_{k=1}^p \alpha_k \Delta Def_{t-k} + \sum_{l=1}^p \alpha_l YG_{t-l} \right] + \epsilon_{t,Inf}
\end{aligned} \tag{3.1}$$

$$\begin{aligned}
Int_t &= I_t \left[ \beta_L + \sum_{i=1}^p \beta_i Inf_{t-i} + \sum_{j=1}^p \beta_j \Delta MS_{t-j} + \sum_{k=1}^p \beta_k \Delta Def_{t-k} + \sum_{l=1}^p \beta_l YG_{t-l} \right] + \\
(1 - I_t) &\left[ \beta_H + \sum_{i=1}^p \beta_i Inf_{t-i} + \sum_{j=1}^p \beta_j \Delta MS_{t-j} + \sum_{k=1}^p \beta_k \Delta Def_{t-k} + \sum_{l=1}^p \beta_l YG_{t-l} \right] + \epsilon_{t,Int}
\end{aligned} \tag{3.2}$$

$$\begin{aligned}
\Delta MS_t &= I_t \left[ \gamma_L + \sum_{i=1}^p \gamma_i Inf_{t-i} + \sum_{j=1}^p \gamma_j Int_{t-j} + \sum_{k=1}^p \gamma_k \Delta Def_{t-k} + \sum_{l=1}^p \gamma_l YG_{t-l} \right] + \\
(1 - I_t) &\left[ \gamma_H + \sum_{i=1}^p \gamma_i Inf_{t-i} + \sum_{j=1}^p \gamma_j Int_{t-j} + \sum_{k=1}^p \gamma_k \Delta Def_{t-k} + \sum_{l=1}^p \gamma_l YG_{t-l} \right] + \epsilon_{t,MS}
\end{aligned} \tag{3.3}$$

$$\begin{aligned}
\Delta Def_t &= I_t \left[ \rho_L + \sum_{i=1}^p \rho_i Inf_{t-i} + \sum_{j=1}^p \rho_j Int_{t-j} + \sum_{k=1}^p \rho_k \Delta MS_{t-k} + \sum_{l=1}^p \rho_l YG_{t-l} \right] + \\
(1 - I_t) &\left[ \rho_H + \sum_{i=1}^p \rho_i Inf_{t-i} + \sum_{j=1}^p \rho_j Int_{t-j} + \sum_{k=1}^p \rho_k \Delta MS_{t-k} + \sum_{l=1}^p \rho_l YG_{t-l} \right] + \epsilon_{t,Def}
\end{aligned} \tag{3.4}$$

$$\begin{aligned}
YG_t &= I_t \left[ \theta_L + \sum_{i=1}^p \theta_i Inf_{t-i} + \sum_{j=1}^p \theta_j Int_{t-j} + \sum_{k=1}^p \theta_k \Delta MS_{t-k} + \sum_{l=1}^p \theta_l \Delta Def_{t-l} \right] + \\
(1 - I_t) &\left[ \theta_H + \sum_{i=1}^p \theta_i Inf_{t-i} + \sum_{j=1}^p \theta_j Int_{t-j} + \sum_{k=1}^p \theta_k \Delta Def_{t-k} + \sum_{l=1}^p \theta_l \Delta Def_{t-l} \right] + \epsilon_{t,YG}
\end{aligned} \tag{3.5}$$

where  $Inf_t$  is inflation rate at time  $t$ ,  $Int_t$  is interest rate at time  $t$ ,  $\Delta Def_t$  is growth in budget deficit at time  $t$ , and  $\Delta MS_t$  is growth in money supply at time  $t$ , and  $YG_t$  is the measure of the output gap at time  $t$ .  $\alpha_j, \beta_j, \gamma_j, \delta_j$  and  $\theta_j$  are the parameters of the model, and  $p$  represent the lag length selected using the information criterion. The output gap variable is our endogenous threshold variable such that:

$$I_t = \begin{cases} 1, & \text{if } YG_{t-d} \leq \tau \\ 0, & \text{if } YG_{t-d} > \tau \end{cases} \quad (3.6)$$

$I_t$  is an indicator function that takes the value of 1 when output gap  $YG_{t-d}$  is less than or equal to the threshold, and 0 otherwise. The time lag for the output gap is set to 1 ( $d=1$ ). We arrange the output gap variable in an ascending order and trim the top 15% and the bottom 15% in order to avoid over-fitting. We compute a VAR model for each of the output gap values and obtain the sum-squared of the residuals. We chose the output gap corresponding to the lowest sum-squared of the residuals as our threshold and the corresponding model as our threshold VAR model. We interpret the model below the threshold as bad state, which is assumed to be the state of recession. The model above the threshold is interpreted as good state, which we assume to be the state of expansion.

We conduct specification test of linear VAR model against nonlinear alternative. The threshold value is not known a priori and the testing procedure involves non-standard inference, because  $\tau$  is not identified under the null hypothesis of no threshold. Therefore, first, the TVAR model is estimated for all possible values of  $\tau$ , (to avoid over-fitting, the possible values were set so that at least 15% of the observations). This test is the multivariate extension proposed by [Lo and Zivot \(2001\)](#) of the linearity test of [Hansen \(1996\)](#). Instead of a F-test comparing the SSR for the univariate case, a Likelihood Ratio (LR) test comparing the covariance matrix of each model is computed.

$$LR_{ij} = T \left[ \ln(\det \hat{\Sigma}_i) - \ln(\det \hat{\Sigma}_j) \right] \quad (3.7)$$

where  $\hat{\Sigma}_i$  is the estimated covariance matrix of the model with  $i$  regimes (e.g.  $i=1$  imply an one threshold regime). We test a linear model against 1 threshold alternative.

In a linear model, the impulse responses can be derived directly from the estimated coefficients and the estimated responses are symmetric both in terms of the sign and of the size of the structural shocks. Furthermore, these impulse responses are constant over time as the covariance structure does not change. However, these convenient properties do not hold within the class of nonlinear models as shown by [Koop et al. \(1996\)](#). The moving average representation of the TVAR is nonlinear in the structural disturbances, because some shocks may lead to switches between regimes, and thus their Wold decomposition does not exist. Consequently, in contrast to linear models, we cannot construct the impulse responses as the paths the variables follow after an initial shock, assuming that no other shock hits the system. To cope with these issues, [Koop et al. \(1996\)](#) proposed nonlinear impulse response functions defined as the difference between the forecasted paths of variables with and without a shock to a variable of interest. Formally, the nonlinear impulse responses functions (NIRF) are defined as:

$$NIRF_Y(k, \epsilon_t, \Omega_{t-1}) = E(Y_{t+k} | \epsilon_t, \Omega_{t-1}) - E(Y_{t+k} | \Omega_{t-1}) \quad (3.8)$$

where  $Y_{t+k}$  is a vector of forecasts at horizon  $k$ ,  $\Omega_{t-1}$  is the available information set, and  $\epsilon_t$  is the shock at time  $t$ . The following discussion draws on from [Afonso et al. \(2011\)](#). This formulation implies that the impulse response functions depend on the initial conditions and that there is no restriction regarding the symmetry of the shocks. Therefore, in order to get the complete information about the dynamics of the model, the impulse responses have to be simulated for various sizes and for the signs of the shocks. The algorithm proceeds as follows. First, the shocks for the periods from  $0$  to  $q$  are drawn from the residuals of the estimated VAR model. Then, for each initial value that is, for each point of our sample, this sequence of shocks is fed through the model to produce forecasts conditional on initial conditions. These steps are repeated for the same initial condition and the same set of residuals except for the

shock to the variable of interest, which is set to (+/-) 1 standard error and (+/-) 2 standard errors at time 0. Second, we can calculate the forecasts conditional on the shocks and on the initial conditions with and without an additional shock at  $t = 0$ , and the difference between these two is the impulse response function. This procedure is replicated 500-times for each initial condition, and then we compute averages over the initial conditions from each regime to get the average impulse responses for both regimes. Using these forecasts, we compute the time path for real interest rate following the Fishers equation:

$$r_t^j = i_t^j - \pi_{t+1}^j \quad (3.9)$$

where  $r, i$  and  $\pi$  represent real interest rate, short-term interest rate and inflation rate respectively. The subscript  $j = \textit{good}$  or  $\textit{badstate}$ , superscript  $t$  represent forecast horizon.

### 3.4 Results and Analysis

We compute a threshold VAR with 1 lag, where the optimal lag length is selected based on AIC. We proceed with the specification test based on likelihood ratio as per [Lo and Zivot \(2001\)](#) described earlier. Table 3.1 below presents the likelihood ratio test where we reject the null of the linear model against threshold VAR. Based on the threshold VAR, with the threshold is defined at  $-2.073^1$ . The endogenously estimated threshold defines our two regimes. For example, when the output gap falls below  $-2.073$  the corresponding model is our bad state or recessionary state. When the output gap is above the threshold value of  $-2.073$  the corresponding model is our good state or expansionary state. Using the TVAR and the threshold, we proceed to compute the generalized impulse responses. We compute nonlinear impulse responses for  $+1, +2, -1$  and  $-2$  standard deviation change in the variables. For brevity, we will present the impulse responses for  $+1$  standard deviation increase for the good and the bad states. We derive the real interest rate for the two states of the economy

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<sup>1</sup>We have 12.4% observations in the regime below the threshold, and 87.6% observations above the threshold.



using equation 3.9 discussed earlier.

Table 3.1: Specification Test

Model	Threshold $\tau$	$H_0$ : Linear VAR
Threshold VAR	-2.07	250.88*** (0.00)

Note: 1. p-value in the parenthesis

2. Boot strap critical values are 118.50 for 90%, 129.83 for 95%, and 138.90 for 99%.

We estimate nonlinear impulse response for two states- a good state and a bad state as per out threshold VAR models. These impulse responses present us a visual comparison of how one variable is behaving over the forecast horizon with respect to an increase in the other during good state and bad state. Figure 3.3 below shows the 15 quarter ahead nonlinear impulse responses of inflation and short-term interest rate with a +1 standard deviation increase to growth in money supply. We observe both inflation and 3-month Treasury bill interest rate responds after one-quarter and show asymmetry. Inflation rises relatively faster in the good state as opposed to the bad state. The low aggregate demand in the bad state perhaps plays an important role. The 3-month Treasury bill rates keep rising in the good state. However, in the bad state, it reduces initially but after 5 quarters it starts to rise again. When we combine the inflation and short-run interest rates to explore the dynamics in real interest rate, we observe real interest rate also behave in an asymmetric manner (see figure 3.4 below). The growth in money supply lower real interest rate in the bad state, which remains negative for about 8 quarters. Though it becomes positive afterward but remains very low. During the good state, we observe the real interest to remain negative for about 4 quarters. But it keeps rising eventually becoming positive after 4 quarters. In analyzing the behavior of the real interest rate during a bad state (recession) and a good state (expansion) we will refer to the liquidity effect, price-level effect and expected inflation effects.<sup>2</sup> In the

<sup>2</sup>Please refer to Chapter 5, Mishkin, F.S. (2014). Economics of Money, Banking and Financial Markets, 11th Ed., Pearson for discussion on income, price-level and expected-inflation effect

good state, the liquidity effect is dominant for about a year. But the expected inflation and the price-level effects become relatively more dominant causing real interest rate to rise faster afterward. In the bad state, on the other hand, the liquidity effect is dominant for about 5 quarters pushing the interest rate lower. But beyond that the expected inflation and price-level effects start to dominate more than liquidity effect, thus, the real interest rate starts rising.

Figure 3.3: Nonlinear Impulse Response with shock to growth in Money Supply

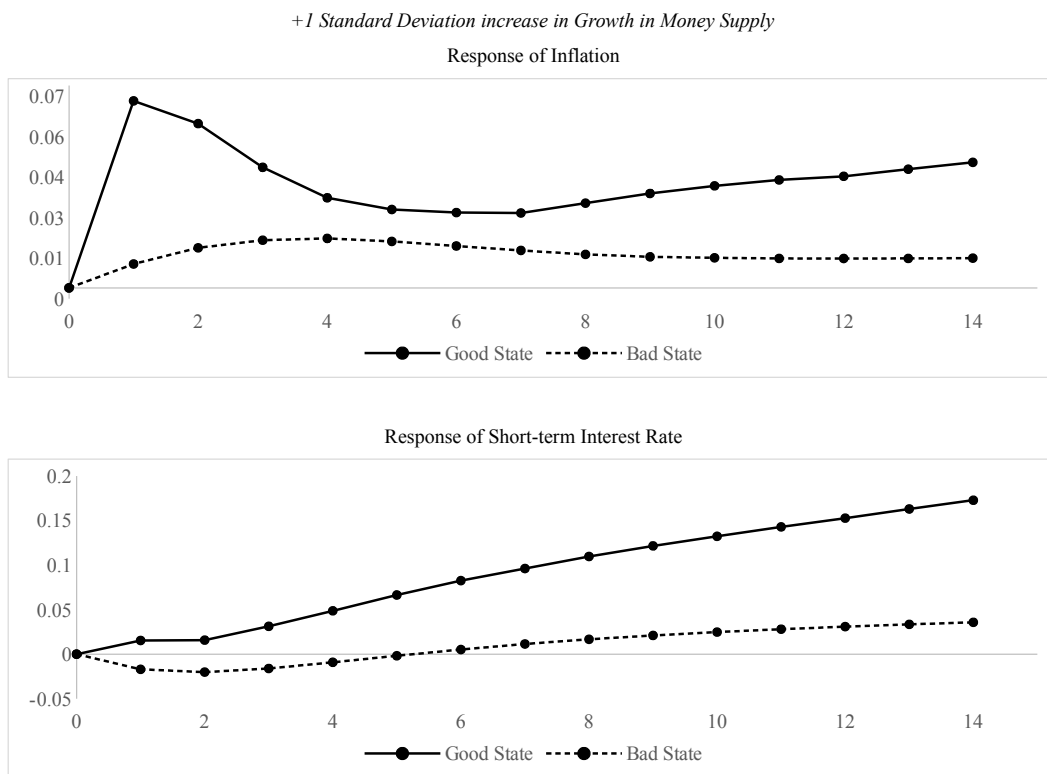
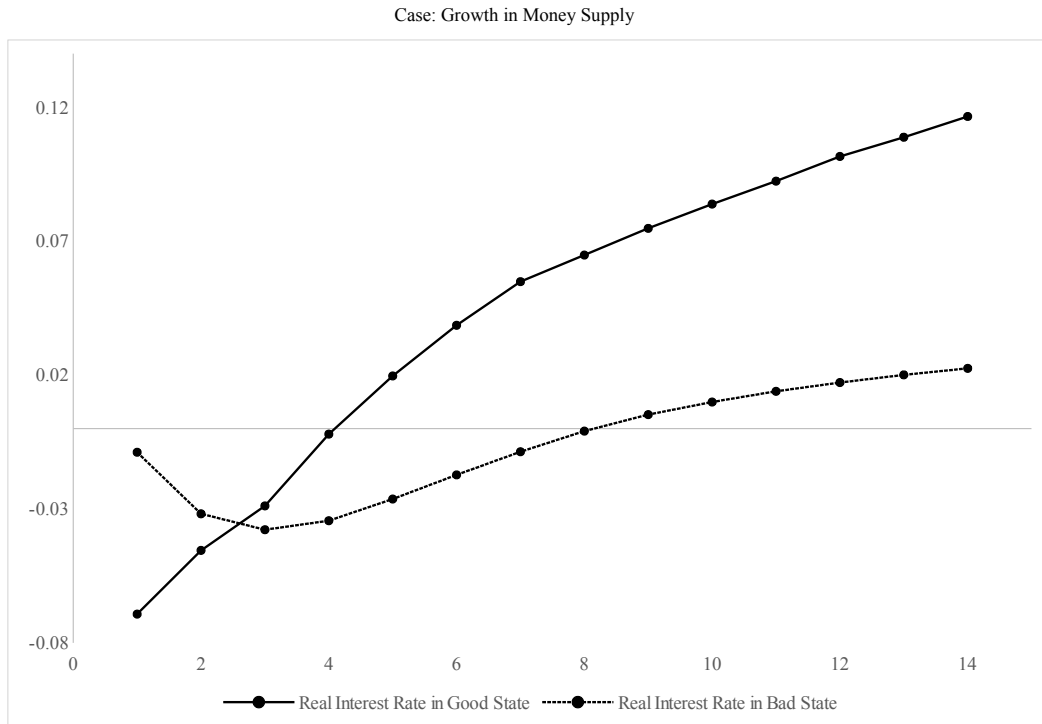


Figure 3.5 below shows the 15 quarter ahead nonlinear impulse responses for inflation and short-term interest rate with +1 standard deviation increase in the growth in the budget deficit. We can find that both inflation and 3-month Treasury bill interest rate responds after one-quarter and display asymmetry. We observe that the inflation rate rises faster in the good state, whereas we observe deflation in the bad state. Conceivably increased deficit causes a larger decline in aggregate demand through a decline in consumption (a negative wealth effect) in the bad state. Although, after 10 quarters the increase in budget deficit

Figure 3.4: Time path for Real Interest Rate



cause deflation in the good state. The short-run interest rate measured by the 3-month Treasury bill rates falls throughout the forecast horizon in the bad state. However, in the good state, it rises initially but after 2 quarters it starts to fall throughout the forecast horizon. When we combine the inflation and short-run interest rates to explore the dynamics of the real interest rate, we observe real interest rate also behave in an asymmetric manner. An increase in the budget deficits in the good states keeps the real interest rate negative. During the bad state, the real interest rate remains positive for about 5 quarter, then gradually goes down over the forecast horizon. [Mankiw \(1987\)](#) argue that government spending depresses the real interest rate due to a wealth effect. Economic agents reduce private consumption in anticipation of an increase in tax in the future. This increases the marginal utility of consumption lowering the marginal rate of substitution, hence marginal productivity of capital also declines. These results conform to [Laubach \(2009\)](#) and [Choi and Holmes \(2014\)](#). The increase in real interest rate on impact (and remain positive for about a year) in the bad state (recession) bear evidence for potential crowding out effect (see figure 3.6).

Figure 3.5: Nonlinear Impulse Response with shock to growth in Budget Deficit

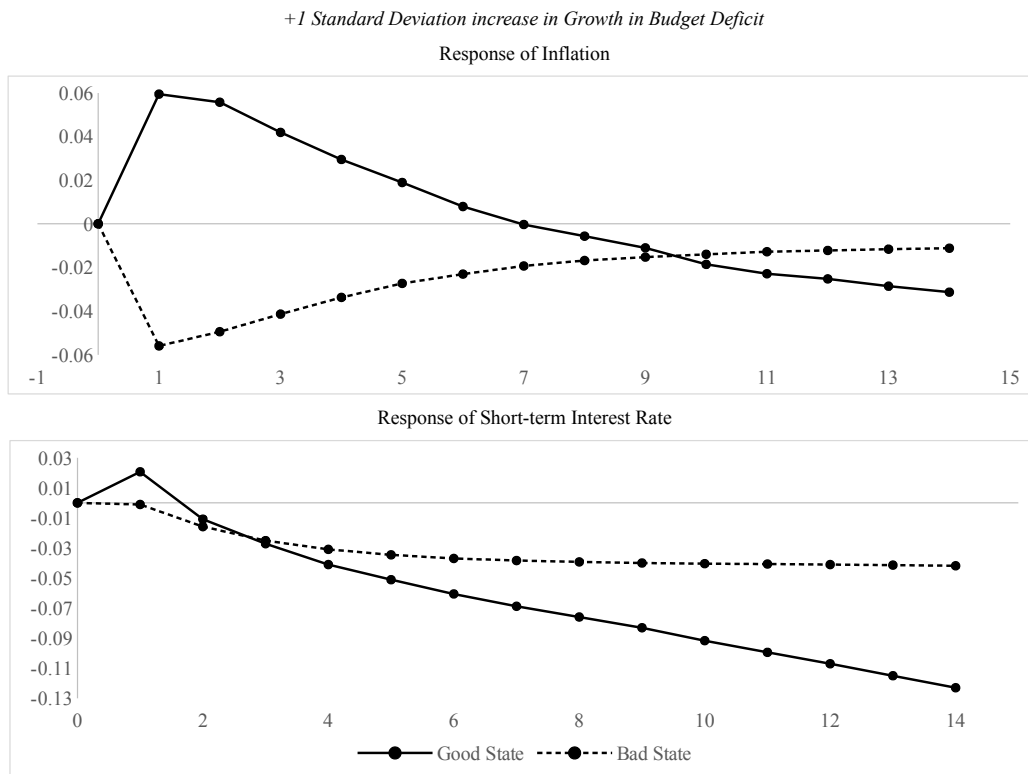


Figure 3.6: Time path for Real Interest Rate

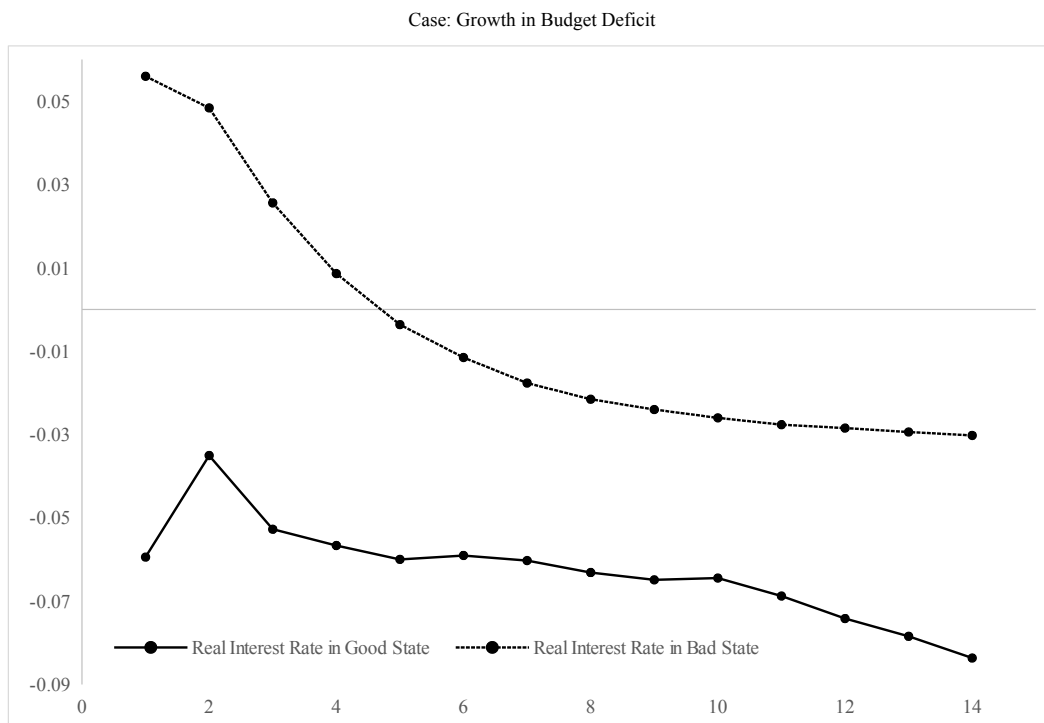
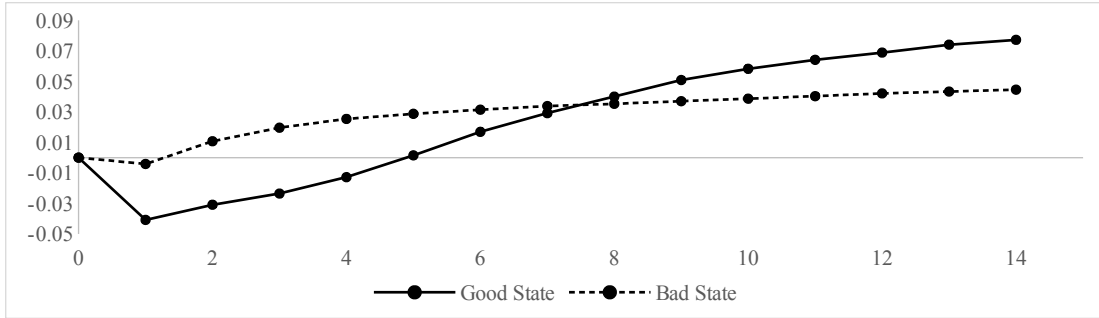


Figure 3.7 below shows the 15 quarter ahead nonlinear impulse responses for inflation and interest rate with +1 standard deviation increase in the output gap. For intuitive purposes, we can interpret this as a case of a positive supply shock to the economy. An increase in the output gap implies positive response of the economy, where output in the economy is increasing. During economic expansions we observe the demand to grow, unemployment to go down, investment and consumption increase over time. Similar to cases of growth in the money supply and the growth in the budget deficit, we can find that both inflation and short-term 3-month Treasury bill interest rate responds after one-quarter and display asymmetry. We observe initial deflation, although the deflationary phenomena are short-lived in the bad state as opposed to the good state. Thus we observe some evidence of counter-cyclicity, where the counter-cyclicity is more prevalent in the good state in contrast to the bad state. Short-term 3-month Treasury bill rate keeps rising in both states although the increase is larger in the good state as opposed to the bad state. When we combine the inflation and the short-run interest rate we find that in the bad state the real interest rate rises initially but then fall again with respect to a positive increase in the output gap (see figure 3.8). However, during the good state, we observe the real interest rate to remain low for up to 4 quarters and then start rising. The output gap also finds an asymmetric response to inflation and short-term interest rate. The real interest rate is pro-cyclical, arguably the expanding economy increases demand and supply.

Figure 3.7: Nonlinear Impulse Response with shock to Output Gap

+1 Standard Deviation increase in the Output Gap

Response of Inflation



Response of Short-term Interest Rate

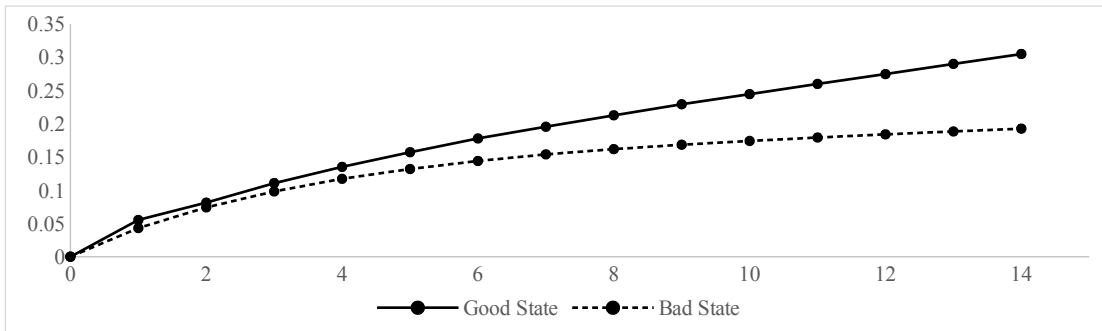
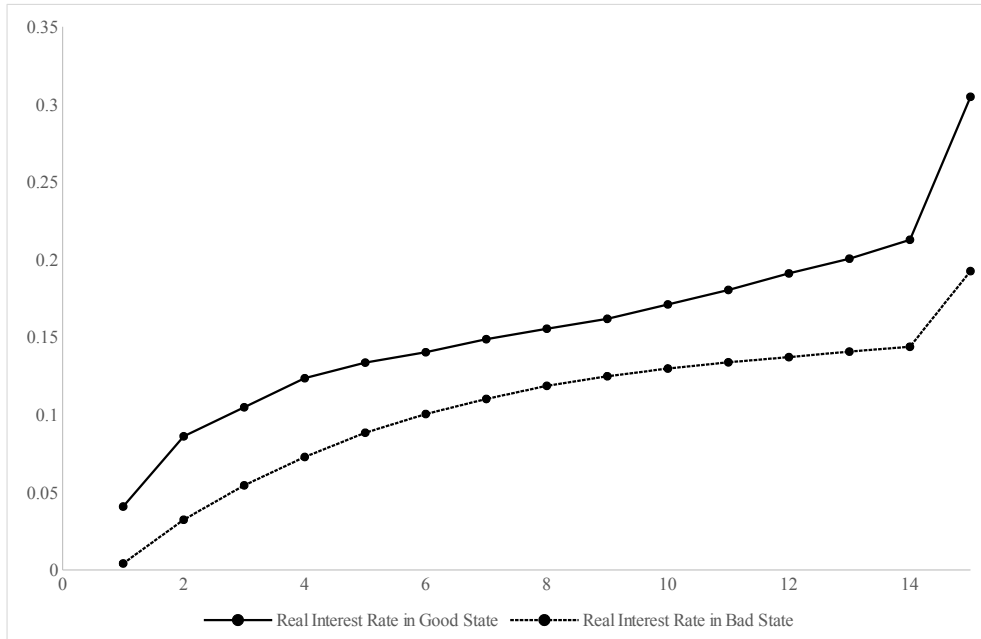


Figure 3.8: Time path for Real Interest Rate

Case: Increase in the Output Gap



## 3.5 Conclusion

In this essay, we analyze the response of inflation, short-term interest rate, and real interest rate during a good state (expansion) and a bad state (recession). It is imperative for us to understand as to how the economy will respond to increases in the fiscal deficit and the money supply during a recession as opposed to economic expansion. Economic agents resource allocation decision is closely related to these variables. In the presence of business cycle, these variables are likely to respond in a differential manner. We employ a threshold VAR model and nonlinear impulse response functions to investigate the dynamic behavior of inflation, short-term interest rate, and real interest rate.

In our analysis, we find specification test of linear VAR against nonlinear VAR rejects the null. Thus we estimate a TVAR and proceed with the nonlinear impulse response analysis. In this paper using quarterly U.S. data, we find that both inflation and short-run interest rates behave asymmetrically with increases in the growth in money supply and the budget deficits. When we combine the future trajectory of inflation and short-run interest rates to analyze the behavior of real interest rate, we observe asymmetry and counter-cyclical. We find the liquidity effect is dominant in the bad state when there is a growth in money supply, but the expected inflation and the income effects dominate the liquidity effect eventually. An increase in the growth in budget deficits tends to be deflationary in the bad state, and short-term interest rate declines. Time path for real interest rate indicates that the wealth effect is relatively more dominant in the good state. Time path for projected real interest rate indicates asymmetry. The output gap also finds an asymmetric response to inflation and short-term interest rate. Real interest rate is pro-cyclical.

Thus, existing opposing or contradictory conclusions regarding the effects of deficit and money supply on interest could potentially arise from asymmetry. As a policy recommendation, we argue for deficit reduction during economic expansions.

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# Appendix A

## Nonlinear Impulse Response

### Computation Algorithm for Chapters

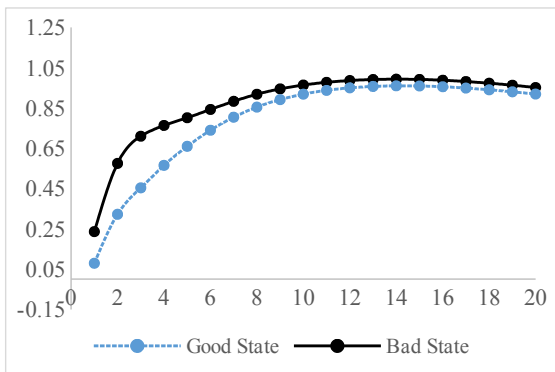
### 1 and 3

Algorithm to compute nonlinear impulse responses:

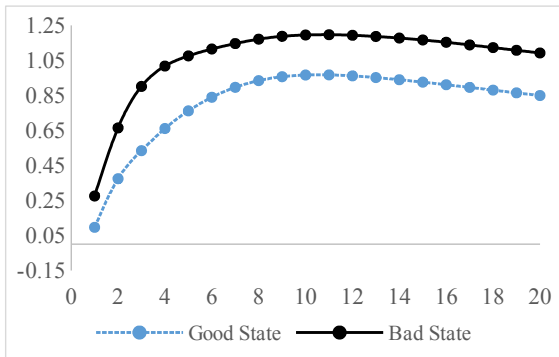
1. The shocks for the periods from 0 to  $q$  are drawn from the residuals of the estimated VAR model
2. For each initial value this sequence of shocks is fed through the model to produce forecasts conditional on initial conditions.
3. Repeat step 2) with the initial shock into one variable equal to  $\pm 1$  or 2 SD to get forecasts if there was an initial shock.
4. The difference between the forecasts from step 2 and 3 is the impulse response function. Repeat this 500-times and derive an average impulse response for this particular initial condition
5. Repeat steps 2-4 for each initial conditions. Final impulse responses are average impulse responses over initial conditions of each regime.

Figure A.1: Nonlinear Impulse Response

Figure: Response of Debt to GDP Ratio

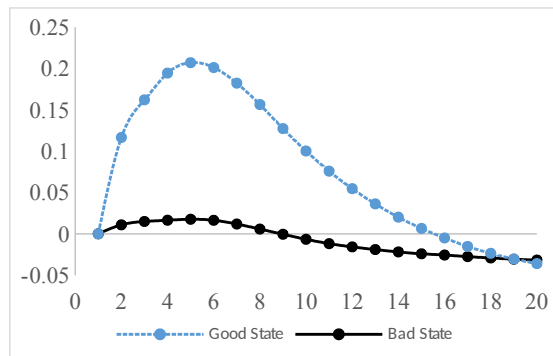


Note: +1 S.D. increase to Gross Deficit to GDP ratio



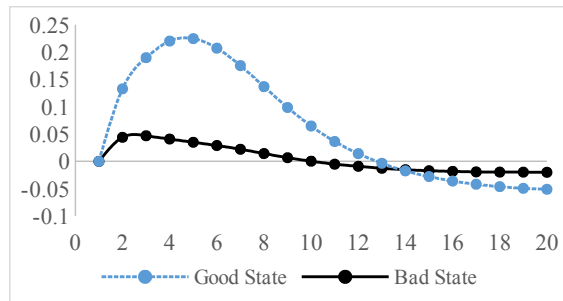
Note: +1 S.D. increase to Current Deficit to GDP ratio

Figure: Response of Gross Deficit to GDP Ratio



Note: +1 S.D. increase to Debt to GDP ratio

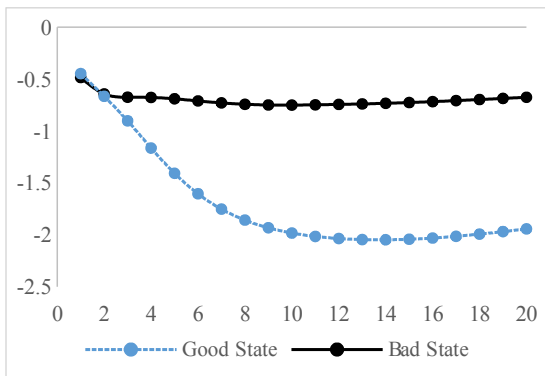
Figure: Response of Current Deficit to GDP Ratio



Note: +1 S.D. increase to Debt to GDP ratio

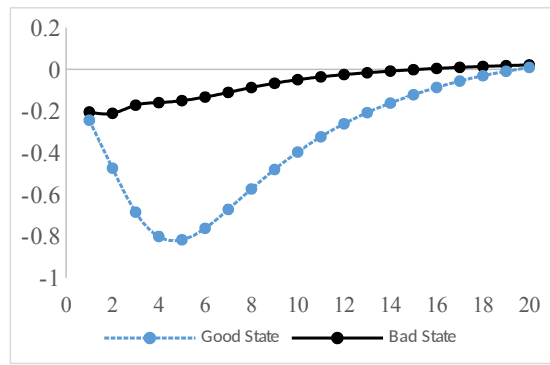
Figure A.2: Nonlinear Impulse Response

Figure: Response of Debt to GDP Ratio



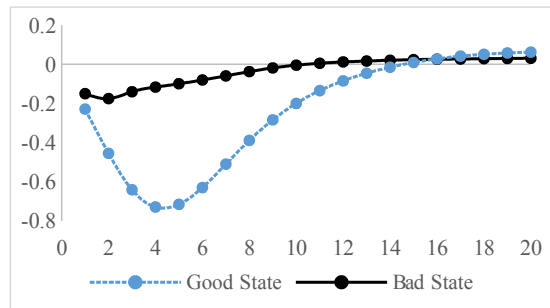
Note: +1 S.D. increase to the Output Gap

Figure: Response of Gross Deficit to GDP Ratio

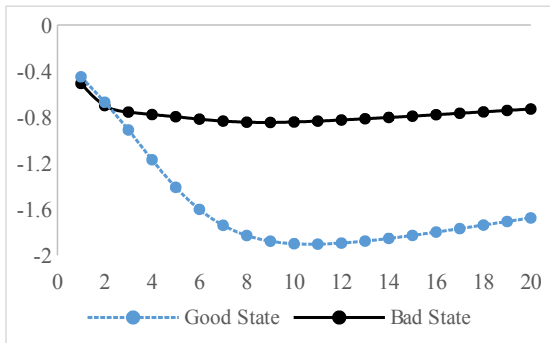


Note: +1 S.D. increase to the Output Gap

Figure: Response of Current Deficit to GDP Ratio



Note: +1 S.D. increase to the Output Gap



Note: +1 S.D. increase to the Output Gap

# Appendix B

## TVECM Tables for Chapter 2

The following tables show the full model with threshold effect on all the right hand side variables.

Table B.1: Threshold Vector Error Correction Model  
 Full Sample, Threshold in All RHS variables

$\Delta ECT_{t-1} = ( \Delta LD_{t-1} - 1.37\Delta Lb_{t-1}  \geq -0.0152)$		
	$\Delta LD_t$	$\Delta LD_t$
Constant	0.001*** (0.0006)	0.003*** (0.0006)
$ECT_{t-1}^H$	-0.002** (0.003)	-0.0005 (0.001)
$\Delta LD_{t-1}^H$	0.25*** (0.04)	0.003 (0.03)
$\Delta LD_{t-2}^H$	0.0006 (0.03)	0.08*** (0.02)
$\Delta LD_{t-3}^H$	0.16*** (0.04)	0.07*** (0.02)
$\Delta LB_{t-1}^H$	0.13* (0.08)	0.16 (0.16)
$\Delta LB_{t-2}^H$	0.08 (0.05)	-0.17*** (0.02)
$\Delta LB_{t-3}^H$	0.11** (0.05)	0.13 (0.09)
$\Delta ECT_{t-1} = ( \Delta LD_{t-1} - 1.37\Delta Lb_{t-1}  < -0.0152)$		
Constant	0.008*** (0.002)	0.003 (0.005)
$ECT_{t-1}^L$	0.007** (0.002)	0.005 (0.004)
$\Delta LD_{t-1}^L$	0.37* (0.22)	0.37 (0.29)
$\Delta LD_{t-2}^L$	0.36** (0.19)	0.02 (0.14)
$\Delta LD_{t-3}^L$	-0.04 (0.09)	-0.08 (0.09)
$\Delta LB_{t-1}^L$	0.03 (0.08)	0.89*** (0.24)
$\Delta LB_{t-2}^L$	-0.07 (0.12)	-0.32 (0.16)
$\Delta LB_{t-3}^L$	0.02 (0.07)	-0.05 (0.15)

Note: 1. Eicker-White Heteroskedasticity Consistent Standard Error,  
 2. Approximately 84.72% observation in higher regime and 15.28% in lower regime.

Table B.2: Threshold Vector Error Correction Model  
Sub-sample: 1946 to 2007, Threshold in All RHS variables

$\Delta ECT_{t-1} = ( \Delta LD_{t-1} - 1.42\Delta Lb_{t-1}  \geq -0.0106)$		
	$\Delta LD_t$	$\Delta LD_t$
Constant	0.001 (0.0008)	0.002*** (0.0004)
$ECT_{t-1}^H$	<b>-0.003**</b> <b>(0.001)</b>	<b>-0.002**</b> <b>(0.0009)</b>
$\Delta LD_{t-1}^H$	0.26*** (0.05)	-0.02 (0.02)
$\Delta LD_{t-2}^H$	-0.005 (0.041)	0.07** (0.02)
$\Delta LD_{t-3}^H$	0.18*** (0.04)	0.07*** (0.01)
$\Delta LB_{t-1}^H$	0.08 (0.12)	0.27** (0.08)
$\Delta LB_{t-2}^H$	0.19 (0.12)	-0.04 (0.07)
$\Delta LB_{t-3}^H$	-0.02 (0.11)	0.11** (0.05)
$\Delta ECT_{t-1} = ( \Delta LD_{t-1} - 1.42\Delta Lb_{t-1}  < -0.0106)$		
Constant	-0.003 (0.003)	0.00001 (0.001)
$ECT_{t-1}^L$	0.009** (0.003)	0.001 (0.002)
$\Delta LD_{t-1}^L$	0.23 (0.20)	-0.04 (0.07)
$\Delta LD_{t-2}^L$	0.33** (0.13)	0.08* (0.04)
$\Delta LD_{t-3}^L$	-0.03 (0.09)	0.13*** (0.03)
$\Delta LB_{t-1}^L$	0.42 (0.39)	-0.008 (0.22)
$\Delta LB_{t-2}^L$	0.64 (0.42)	0.43** (0.14)
$\Delta LB_{t-3}^L$	0.57** (0.25)	0.26** (0.11)

Note: 1. Eicker-White Heteroskedasticity Consistent Standard Error,  
2. Approximately 77.78% observation in higher regime and 22.22% in lower regime.

# Appendix C

## Sensitivity Analysis TVECM Tables for Chapter 2

The following tables show the full model with threshold effect on the error correction term only.

Table C.1: Threshold Vector Error Correction Model  
Full Sample, Threshold in ECT Only

	$\Delta LD_t$	$\Delta LD_t$
Constant	0.003*** (0.0006)	0.002*** (0.0008)
$ECT_{t-1}^H$	<b>-0.002***</b> <b>(0.0008)</b>	-0.002 (0.001)
$ECT_{t-1}^L$	<b>0.007***</b> <b>(0.002)</b>	0.001 (0.003)
$\Delta LD_{t-1}^H$	0.24*** (0.03)	0.04 (0.04)
$\Delta LD_{t-2}^H$	0.05 (0.04)	0.07** (0.03)
$\Delta LD_{t-3}^H$	0.14*** (0.03)	0.008 (0.03)
$\Delta LB_{t-1}^H$	0.11** (0.04)	0.66*** (0.18)
$\Delta LB_{t-2}^H$	-0.02 (0.06)	-0.22** (0.09)
$\Delta LB_{t-3}^H$	0.04 (0.03)	0.03 (0.13)

Note: 1. Eicker-White Heteroskedasticity Consistent Standard Error,  
2. Approximately 84.72% observation in higher regime and 15.28% in lower regime.



Table C.2: Threshold Vector Error Correction Model  
Full Sample, Threshold in ECT Only

	$\Delta LD_t$	$\Delta LD_t$
Constant	0.000006 (0.001)	0.002*** (0.0004)
$ECT_{t-1}^H$	-0.002 (0.001)	<b>-0.001*</b> <b>(0.0009)</b>
$ECT_{t-1}^L$	<b>0.007*</b> <b>(0.004)</b>	-0.0008 (0.002)
$\Delta LD_{t-1}^H$	0.16*** (0.04)	-0.01 (0.02)
$\Delta LD_{t-2}^H$	0.04 (0.04)	0.07*** (0.02)
$\Delta LD_{t-3}^H$	0.12*** (0.04)	0.08*** (0.01)
$\Delta LB_{t-1}^H$	0.38*** (0.14)	0.22** (0.08)
$\Delta LB_{t-2}^H$	0.31** (0.12)	0.01 (0.07)
$\Delta LB_{t-3}^H$	0.14 (0.11)	0.13*** (0.04)

Note: 1. Eicker-White Heteroskedasticity Consistent Standard Error,  
2. Approximately 77.78% observation in higher regime and 22.22% in lower regime.