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Cyclical Fiscal Multipliers:

Policy Mix and Financial Friction Puzzle

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Cyclical Fiscal Multipliers: Policy Mix and Financial Friction Puzzle
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ABSTRACT: This paper investigates dynamic relationships between U.S. government expenditure multipliers and the economy's cyclical position from 1949 to 2018 using a Time-Varying Parameter Vector Autoregression (TVP-VAR) model. We challenge the existing literature, which predominantly relies on predefined economic regimes and assumes a stable relationship between fiscal multipliers and business cycles. Our findings identify two distinct periods: fiscal multipliers were counter-cyclical from 1949 to the late 1980s, followed by a significant decline in their effectiveness during recessions thereafter. These variations are attributed to the prevailing fiscal-monetary policy mix; with higher fiscal multipliers during earlier recessions resulting from sharp shifts toward a fiscally led policy stance, followed by a decline after the Dot-com recession due to a transition toward a monetary-led policy mix. We find particularly low multipliers during the global financial crisis, which provides new insights into the evolving role of financial frictions in the transmission of fiscal policy.

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Contents

1 Introduction

The primary argument supporting discretionary fiscal spending during recessionary phases, over other conventional strategies, hinges on the size and direction of fiscal multipliers. This approach becomes particularly crucial when other policy measures fall short in mitigating recessionary pressures.[1] In this context, empirical literature has predominantly discussed two contrasting perspectives. Auerbach and Gorodnichenko (2012), for example, contend that public expenditure multipliers are counter-cyclical. Conversely, Ramey and Zubairy (2018), to name but one, challenge this view by demonstrating the absence of significant differences in multiplier values across various phases of business cycles.

The above studies implicitly assume a stable relationship between the size of fiscal multipliers and phases of business cycles over time. While their modeling approaches simplify the analysis and enhance intuition, they carry the risk of misleading conclusions, as this relationship is likely to change dynamically over time. Our paper addresses these concerns by investigating the following key question: How do fiscal multipliers evolve during a recession, particularly when considering the potential influence of higher-order non-linearities that are often overlooked by aggregate time series models?[2]

To answer this question, we extend the Time-Varying Parameter Vector Autoregression (TVP-VAR) model developed by Belmonte et al. (2014) to estimate a series of time-varying fiscal multipliers. The model incorporates a decision mechanism that adjusts the degree of time variation in its parameters. Then, following Eisenstat et al. (2016), a Tobit prior is applied to regulate the extent of time variation in the VAR parameters, while a Lasso prior is imposed on their variances. This approach enhances the efficiency of stochastic model specification search in large TVP-VAR models while simultaneously leading the model toward a more stable, time-invariant VAR. Furthermore, adopting the TVP-VAR framework allows us to extend the VAR lag polynomial to four lags, in line with the reasoning of Blanchard and Perotti (2002).[3] Including all four lags in a fiscal VAR enables the capture of delayed effects and helps mitigate potential endogeneity issues, considering that the fiscal year spans four quarters.

[1] For instance, in the wake of the Covid-19 pandemic, the economic landscape has cast doubt on the efficacy of monetary policies, see Bernanke (2022), Bhar and Malliaris (2021) and Levin and Sinha (2020).

[2] Even though regime-switching models are often used to capture such non-linearities, they are limited by the number of regimes to focus, see for instance Callegari et al. (2012). In addition, higher order non-linearity, if not explicitly estimated by the researchers, can produce conflicting conclusions regarding the size or direction of fiscal multipliers. For example, Ascari et al. (2023) show that the qualitative differences in results between Ramey (2011b) and Blanchard and Perotti (2002) disappear once the econometrician conditions on the prevailing policy regime. Similarly, Ramey and Zubairy (2018) focus on the average effect of a recession on the output response to a discretionary government expenditure shock. If the fiscal multiplier decreases during certain recessions and increases during others, depending on some underlying realization of a regime, their study's framework can not reveal any statistically significant differences between recessions and expansions.

[3] Existing contributions estimate TVP-VAR models of order two. Iwata and IIboshi (2023) extends the lag order to four and applies identification similar to this paper. However, they detrend variables prior to estimation, assuming linear and quadratic trends, and do not focus on the state-dependent nature of the multiplier. Detailed definitions are provided in Appendix B.

The identification scheme of this paper integrates sign restrictions with short-term zero restrictions, following Rubio-Ramirez et al. (2010). Moreover, the use of sign restrictions circumvents a key limitation of Cholesky factorization in VAR models that include tax revenues as an endogenous variable.[4] To ensure accuracy, we estimate the model on stationary series in de-trended levels using Hamilton (2018)'s linear projection method which allows us to avoid the re-scaling bias pointed out in Ramey and Zubairy (2018). We employ data from the Bureau of Economic Analysis and the Federal Reserve Economic Database from Q1 1948 to Q2 2018.

Our findings outline that the relationship between multipliers and the business cycle stage is not constant over time. The results divide the post-World War II (WWII) period into two subperiods. The first period, spanning from 1949 to the late 1980s, exhibits counter-cyclical fiscal expenditure multipliers. In contrast, the subsequent period is characterized by pro-cyclical fiscal multipliers. This shift raises the possibility of a monetary explanation, particularly in light of the change in the dynamics of the fiscal multipliers during the Federal Reserve's prolonged disinflationary effort under Paul Volcker. These two phases are explained as follows.

First, our study reveals that the elevated fiscal multipliers during the recessions of the late 1960s, 1970s, and early 1980s were likely driven by specific policy shifts rather than being an inherent feature of the recessions. This challenges the arguments by Auerbach and Gorodnichenko (2012), who suggest that multipliers are consistently higher during recessions. We present evidence that a shift towards more aggressive fiscal policies, such as unfunded increases in government spending or tax cuts during these periods, could have enhanced the economic impact of fiscal interventions.

In the second phase since 1980s, we attribute the reduction in the average multiplier size to the shift toward a monetary-led policy mix. We independently reach the conclusion that multipliers tend to be larger during periods characterized by a fiscally-led policy mix, where fiscal policy is more active and monetary policy takes a passive role. This is consistent with findings from the literature for example Leeper et al. (2017); Bianchi and Ilut (2017); Bianchi et al. (2023). Moreover, unlike Ascari et al. (2023), we provide evidence that expenditure multiplier are lower (and occasionally negative) even for unanticipated fiscal shocks.

Our analysis uncovers a surprising pattern in the post-2000s period: multipliers in some recent recessions have decreased, contradicting the widely accepted view that fiscal multipliers generally rise during downturns. The particularly low government spending multiplier values during the global financial crisis is especially striking.[5] Examining the relationship between multipliers and financial frictions across different subsamples suggests that the role of financial frictions in the U.S. economy may have shifted in recent years, affecting the potency of fiscal stimulus. Our results propose that heightened financial constraints during the global financial crisis, coupled with rising household debt levels, could have dampened the impact of government

[4] See Appendix C

[5] Valery Ramey's Mundell-Fleming Lecture at the IMF's Jacques Polak Conference presented evidence of a similar outcome regarding the multipliers associated with government transfers.

spending on economic activity under a monetary-led policy mix.

To ensure the robustness of our results, we incorporate professional forecasts into our TVP-VAR model, following Berg (2015). We focus on removing the anticipated component from structural shocks by controlling only for the previous quarter's forecast. This approach balances model integrity while addressing policy anticipation. The robustness findings confirm that accounting for policy anticipation does not alter our initial results.

Literature overview

By extending the lag polynomial to the fourth order, detrending data using Hamilton (2018)'s linear projection method, not discarding unstable draws of the lag polynomial in Gibb's sampling, and estimating the model in levels, to the best of our knowledge, this paper is the first to present state-dependent fiscal multipliers in a TVP-VAR framework.[6] In this regard, this study fills two important gaps in the modelling literature. First, by estimating government expenditure multipliers that depend on the business cycle stage using Bayesian time-varying parameter model. Second, our approach allows us to investigate whether the state-dependent nature of the public expenditure multiplier remains the same throughout the post-WWII period in the US. Therefore, we evaluate the correctness of the prevailing set of modelling strategies, analysing the state-dependence of fiscal multipliers.

The theoretical and empirical literature on fiscal multipliers is extensive. Earlier papers focused on the average multiplier size e.g. Ramey and Shapiro (1998), Blanchard and Perotti (2002), Mountford and Uhlig (2005), Ramey (2011b), Barro and Redlick (2011) and largely avoided the topic of state-dependence.[7] Within the empirical realm, the more recent literature that investigated state-dependent nature of the fiscal policy can be broadly categorized into two model classes. The first category estimates state-dependent multipliers using models that rely on explicit regimes or states. The second class of models that can capture such non-linearity relies on time-varying parameter (TVP) models.

The first category, e.g. Auerbach et al. (2020), Shaheen and Turner (2020), Hwang and Kim (2021) and Alloza (2022) to name but a few, is pioneered by the smooth transition VAR approach of Auerbach and Gorodnichenko (2012) which estimates a set of multipliers for various types of spending, allowing the output response to depend on the business cycle stage.[8] Some of the contributions adopt modifications of Jorda (2005)'s local projection method for calculating multipliers that depend on an economic state. For instance, Ramey and Zubairy (2018), using a local

[6]The prevailing belief has been that models of this type are incapable of generating multipliers dependent on the business cycle phase Ramey (2011a); Auerbach and Gorodnichenko (2012). The findings of Kirchner et al. (2010), Pereira and Lopes (2014), and Berg (2015) have lent support to this perspective. Our results, however, demonstrate that the inability to discern such a relationship is largely attributable to their modeling choices.

[7]For a more comprehensive overview of the theoretical contributions on this topic, see Ramey (2011a).

[8]Auerbach and Gorodnichenko (2012) show the U.S. government expenditure multipliers to be as low as -0.33 and as high as 2.24 during expansions and recessions, respectively. Using a non-linear VAR setup, Bachmann and Sims (2012) support higher multipliers during the US economic recessions.

projection instrumental variable (LP-IV) model, conclude that dependence of the government expenditure multiplier on a business cycle stage does not hold for various alternative specifications of models, shocks and states. Bernardini and Peersman (2018) extend the previous model to accommodate more than two states simultaneously. Similar to Ramey and Zubairy (2018), the estimation results offer mixed evidence. Both studies produce state-dependent multipliers in some model specifications.

The aforementioned studies share a common feature: they assume that the relationship between the multiplier's value and the stage of the business cycle remains fixed. These models are primarily designed to calculate the average difference in multiplier values between periods of recession and economic expansion. They often overlook the possibility that this relationship may arise due to underlying policy shifts and reaction of market agents to them, thus, allowing it to evolve over time.

In our approach, contrary to the previous studies, we explicitly account for the potential variability in this relationship. This flexibility in our modeling framework is of significant importance. To illustrate the importance of this aspect in the estimation of multipliers, one needs to consider the fluctuation in the share of unfunded government expenditure, as estimated by Bianchi et al. (2023). This metric exhibited substantial oscillations in the post-WWII period. It displayed distinct peaks during recessions before the 1980s, followed by a less evident pattern in the 1980-1990s period, and small troughs after the dot-com crisis and during the Great Recession.

Building upon the insights from their model, these variations in the pattern of unfunded government expenditure have the potential to amplify multipliers during earlier recessions and reduce them during later ones. This dynamic underscores the significance of our approach, which acknowledges the evolving nature of the multiplier relationship with the business cycle. This offers a more nuanced and accurate understanding of economic dynamics over time.

The second category of models adepts at capturing non-linearity, centered on the utilization of time-varying parameter (TVP) models to calculate time-varying government expenditure multipliers. Kirchner et al. (2010) who to the best of our knowledge are the first to introduce the Bayesian TVP-VAR analysis into the fiscal debate, estimate the model for the European Union (EU).[9] The study does not analyze the dependence of the multiplier value on the stage of the business cycle. Pereira and Lopes (2014), in addition, focus on computing time-varying effects of the fiscal policy for the U.S.[10] They conclude a small degree of time variation in the output response to a discretionary government expenditure shock. Furthermore, they focus on estimating elasticities and not extending the analysis to incorporate multiplier calculation.

On the other hand, papers such as Berg (2015), Afonso et al. (2018) and Iwata and IIboshi (2023), to name but a few, investigate the changing dynamics of government spending multipliers

[9] This study concludes that short-run fiscal multipliers increased until the late 1980s, reaching values above unity, and subsequently decreased to 0.5 by the end of the sample.

[10] Unlike our approach, the study removes the simple quadratic trend and allows for two lags only.

in the United States by employing different types of a time-varying parameter VAR models. Eisenstat et al. (2016) estimate a a time-varying parameter (TVP) vector error correction model (VECM) following Blanchard and Perotti (2002), but their contribution mainly lies in modifying the model of Belmonte et al. (2014) with alternative prior definition. The above studies do not reveal a relationship between government expenditure multipliers and the stage of business cycles. Unlike these studies, our findings raise questions about whether the inability to detect the state-dependent nature of the expenditure multiplier stems from the modeling choices commonly used in the literature.

In addition, our findings challenge the prevailing academic view that financial frictions enhance fiscal multipliers by promoting growth and lowering financial intermediation costs, see for example, Carrillo and Poilly (2013) and Canzoneri et al. (2016). While this relationship is evident in the earlier part of our sample, it appears to reverse in the 2000s. These findings suggest that shifts in the financial landscape, characterized by greater global financial integration, constrained monetary policy under near-zero interest rates, and rising household debt, may have reshaped the traditional dynamics of fiscal policy. Moreover, our empirical results align with the findings of structural models, such as Ghiaie and Rouillard (2022), which outline that financial frictions contribute to lowering tax multipliers. This outcome is consistent with our findings on government spending multipliers, further reinforcing the role of financial constraints in shaping fiscal policy effectiveness.

This paper proceeds as follows. Section two presents the methodology, and section three discusses the identification strategy and data. Section four presents the time-varying government expenditure multipliers and discusses dependence on the stage of business cycles. Section five estimates multipliers in a framework, which acknowledges that policy actions can be anticipated by market agents. Section six delves into the significant variation observed in the estimated multiplier series and outlines the possible drivers behind this variation. Conclusion follows in section seven.

2 Methodology

This study employs the TVP-VAR model to compute a series of time-varying fiscal multipliers, drawing inspiration from Frühwirth-Schnatter and Wagner (2010) and Belmonte et al. (2014). The model incorporates a decision mechanism that regulates the extent of time variation in the model parameters. Following Eisenstat et al. (2016), we use a Tobit prior to regulate the extent of time variation in the VAR parameters and apply a Lasso prior to their variances.[11] Notably, the latter feature proves valuable as it allows for the successful expansion of our model to encompass all four quarterly lags of the fiscal variables.

[11] This approach enhances the efficiency of the stochastic model specification search in large TVP-VAR models while simultaneously guiding the model toward greater stability by favoring a time-invariant VAR structure.

Although the methodology accommodates both a time-varying lag polynomial and a time-varying variance-covariance matrix, it discourages excessive parameterization. Allowing the model parameters to follow random walks introduces a significant degree of non-linearity in their transitions over time.[12] The following system of equations describes the econometric model:

$$Y_t = X_t\beta_t + \Sigma_t u_t, \qquad u_t \sim N(0, I_n) \tag{2.1}$$

$$\beta_t = \beta_{t-1} + v_t \qquad v_t \sim N(0, \Omega) \tag{2.2}$$

$$log(\sigma_t) = log(\sigma_{t-1}) + \theta_t \qquad \theta_t \sim N(0, W) \tag{2.3}$$

where Y_t is an $n \times 1$ vector of observed endogenous variables, X_t is an $n \times m$ matrix of observations on explanatory variables (both endogenous and exogenous variable vectors, their lags, and some contemporaneous elements of Y_t), u_t denotes structural shocks, and Σ_t is a diagonal matrix containing standard deviations of the structural shocks. β_t is a vector containing all coefficients of X_t, and $log(\sigma_t)$ is a vector containing logs of all the non-zero elements of a diagonal matrix Σ_t. Ω and W are the variance-covariance matrices of the disturbances from the parameter laws of motion.

At this point, it is crucial to acknowledge that $\Omega = \tilde{\Omega}^{\frac{1}{2}}\Phi\Phi'\tilde{\Omega}^{\frac{1}{2}\prime}$, where $\tilde{\Omega}^{\frac{1}{2}} = diag(\omega_1, \ldots, \omega_m)$ contains the indicators used to access the degree of time-variance and Φ is a lower-unitriangular matrix. This setup is standard, for instance see Eisenstat et al. (2016). More details are explained in Appendix E.

The set of equations 2.1 to 2.3 pins down the problem under analysis, which is solved using Bayesian techniques.[13] As the joint posterior density is unknown, parameters of interest are sampled iteratively from conditional densities using the Gibb's sampler. Each set of the presented results is based on 150,000 iterations of the Gibb's sampler; the procedure is sensitive to initial values and is subject to the autocorrelation of the sampled draws. Therefore, in each case, a burn-in period of 100,000 is eliminated; subsequently, every 25th draw of the remainder is used to approximate the posterior density function. As in Primiceri (2005), var($[u_t \quad v_t \quad \theta_t]'$) matrix is assumed to be diagonal.

The Bayesian estimation also requires a set of prior distributions. Considering data limitations and to preserve comparability between different subsample estimations, we deviate from the commonly accepted routine of constructing priors based on the pre-sample estimation. Conversely, we assume uninformative priors, as in Eisenstat et al. (2016).[14]

[12] The VAR's time-varying parameters adhere to random walk laws of motion. This modelling choice can be considered a strong assumption. Nonetheless, estimating a time-invariant version of the model on a rolling sample of 100 observations proves it to be a reasonable assumption. All the resulting autoregressive coefficients follow either a random walk or an AR(1) process with a coefficient close to 1.

[13] The Bayesian treatment of the problem, as argued in Cogley and Sargent (2005), allows the treatment of coefficients as random variables and alienates the method from the Lucas Jr (1976) critique.

[14] This choice comes at the cost of broader confidence bands. Nonetheless, as can be seen in section 4, we conclude that the multiplier series are subject to two chronologically ordered regimes. By using uninformative priors, we avoid imposing a prior distribution constructed using data from the first regime on the posterior

The estimation provides a set of VAR parameters for each period t under analysis, subsequently utilized to compute impulse response functions (IRFs). As the calculated posterior distributions are ergodic, the IRFs, being functions of the estimated parameters, are computed individually for each parameter draw. Consequently, the collective set of IRFs constitutes the IRF distribution. In a similar fashion, posterior distributions of the multipliers are derived, illustrating a straightforward method for significance assessment. The analysis produces two multiplier types:

First, the cumulative multiplier:

$$K_t^{sum} = \frac{\Sigma_{h=0}^{H} f_{t+h}^{y}}{\Sigma_{h=0}^{H} f_{t+h}^{g}} \tag{2.4}$$

Second, the impact multiplier:

$$K_t^{imp} = f_t^y \tag{2.5}$$

where f_{t+h}^{y} refers to the output response at horizon h to a fiscal shock taking place at time t. f_{t+h}^{g} refers to the fiscal variable counterpart.

The cumulative multiplier 2.4 is assumed to be equal to the integral of the output response divided by the integral of the government expenditure response; they comprise responses to a discretionary government expenditure shock. As argued in Ramey and Zubairy (2018), this definition of the multiplier tends to provide lower multiplier values than alternative definitions.

3 Identification strategy and data

This paper proposes an identification scheme that combines sign restrictions with short-term zero restrictions following Rubio-Ramirez et al. (2010).[15] Additionally, the implementation of sign restrictions evades an important pitfall of Cholesky factorisation in VAR models including tax revenues as an endogenous variable. Appendix C provides an elaborate explanation of how such a pitfall can arise in a fiscal VAR framework. The proposed identification strategy can be summed up by:

$$Z_t = \begin{array}{c} \\ G_0 \\ NT_0 \\ Y_0 \end{array} \begin{array}{c} \begin{array}{ccc} \varepsilon_t^G & \varepsilon_t^T & \varepsilon_t^Y \end{array} \\ \begin{pmatrix} + & 0 & 0 \\ \times & + & \times \\ \times & - & + \end{pmatrix} \end{array} \tag{3.1}$$

distribution of the model coefficients in the second one.

[15] The actual implementation of the routine follows the strategy outlined by Binning (2013). Short-term zero restrictions are based on the reasoning of Blanchard and Perotti (2002)

where ε_t^G refers to the structural government expenditure shock, ε_t^T to the structural tax shock, and ε_t^Y to the structural output shock.

Identification scheme 3.1 combines the assumption of the lagged discretionary fiscal policy response, as in Blanchard and Perotti (2002), and assigns a minimum amount of sign restrictions necessary for identification. Under this modeling choice, the output shock is allowed to have an immediate effect on net taxes, and a net taxes shock lowers the output contemporaneously. In other words, the latter assumption draws an equivalence between an increase in both net taxes as well as marginal tax rates, which only holds if the fiscal policy stance remains at the uphill of the Laffer curve.[16] Appendix D presents an elaborate explanation of the identification strategy.

The TVP-VAR model incorporates key variables: government expenditure (G_t), tax revenue net of transfer payments (NT_t), and GDP (Y_t), all expressed in real per capita terms to account for population growth and the nominal effects of inflation.[17]

GDP, government expenditure, and net taxes are non-stationary time series, and therefore, estimations in levels may lead to spurious results. On the other hand, detrending data may strip away valuable information contained within the series and impact our results. Consequently, the primary goal in our data preparation is to devise a stationary transformation that retains the variation we wish to analyze.

This study introduces the linear projection method by Hamilton (2018) as a tool to obtain such a stationary series for the fiscal VAR analysis. Hamilton (2018)'s procedure is standard in the literature and generates stationary cyclical components of macroeconomic series centered around zero. Similar to the Hodrick-Prescott filter, it renders stationary series in levels. Unlike the Hodrick-Prescott filter, it avoids creating artificial correlations absent from the actual data generation process. Since these resulting stationary series are incorporated into the model in levels, the estimated multipliers remain unaffected by the rescaling bias identified by Ramey and Zubairy (2018). Thus, the TVP-VAR model, estimated on cyclical components using Hamilton (2018)'s linear projection method, directly produces multipliers rather than elasticities. Figure 1 presents these stationary series alongside the recession dates identified by the National Bureau

[16] The literature contains evidence in favour of the validity of this assumption, at least, for the 1995–2007 period. For example, Trabant and Uhlig (2011) find that, on average, both labour and capital tax rates for the US were below levels prompting maximum tax revenues. Blinder (1981) also supports the argument that the US tax burden, in a broad sense, is unlikely to be on the downhill of the Laffer curve.

[17] In the fiscal VAR literature, two primary approaches to data preparation prevail. The first approach, followed by Mountford and Uhlig (2005), Kirchner et al. (2010), Auerbach and Gorodnichenko (2012), and Bachman and Sims (2012), utilizes data in levels to preserve the long-term relationships between government expenditure, taxes, and output. Auerbach and Gorodnichenko (2012) do acknowledge that, in the presence of non-stationarity, a properly constructed VECM may outperform their modeling strategy. Nevertheless, specifying a specific cointegration vector can be challenging, and conducting the analysis in first differences does not align with economic theory or the empirical consensus on the topic. The second approach, adopted by Blanchard and Perotti (2002), Pereira and Lopes (2014), and Berg (2015), underscores the importance of estimating the multivariate model on stationary data. Therefore, this approach detrends the data before estimation or explicitly accounts for the trend within the model.

of Economic Research (NBER).

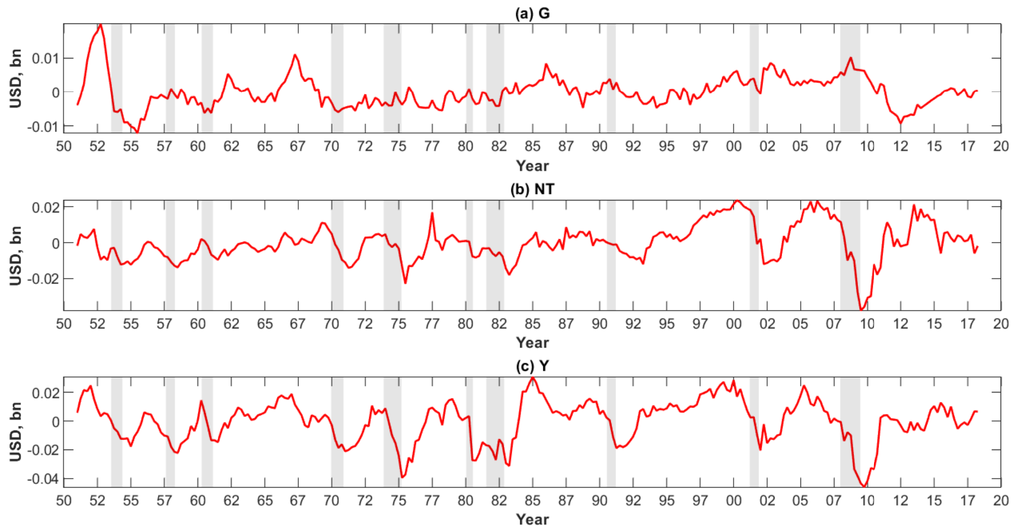

Figure 1: Stationary transformations of the real per capita government expenditure (A), taxes net of transfers (B), GDP (C) over NBER recession dates: obtained via Hamilton's linear projection method

We run the baseline TVP-VAR model on the 1948Q1–2018Q2 sample, assuming four lags and no intercept terms. Data are obtained from the Bureau of Economic Analysis and the Federal Reserve Economic Database. Fiscal variables and GDP are taken from the latest release of the national income and product accounts' tables. Net taxes follow the definition of Blanchard and Perotti (2002) (detailed definitions are presented in Appendix B). The time series enter the model in levels of real per capita terms, detrended via the Hamilton (2018) linear projection technique.

The VAR model of Blanchard and Perotti (2002) serves as our point of comparison with the existing literature, particularly with earlier time-varying fiscal VAR studies such as Pereira and Lopes (2014) and Berg (2015). It is important to acknowledge at this point that the model's relatively small set of endogenous variables may raise concerns that the identified shocks are non-fundamental, as they are conditioned on a more limited information set than that available to market agents; see discussion in Chung and Leeper (2009), Leeper et al. (2013), and Mertens and Ravn (2010). To address this, we introduce an alternative specification that conditions the VAR on professional forecasters' projections. Conditioning the VAR on such forecasts helps align the information set with that of real market agents, ensuring the correct identification of shocks. However, even the baseline model specification may already be resistant to this issue.

Ascari et al. (2023) argue that the incorrect identification of shocks stems from the misspecification of linear VARs estimated on data from two distinct regimes, each requiring a separate VAR structure. If this hypothesis holds, the TVP-VAR methodology should be immune to this issue, as it accommodates shifts in both the lag polynomial and the variance-covariance matrix over time. However, omitting key state variables, such as public debt, presents significant limitations. The estimated IRFs may fail to fully capture the role these variables play in the transmission of discretionary fiscal shocks.

4 Government expenditure during economic slack

The estimation of the baseline model leads to several noteworthy outcomes. First, the TVP-VAR framework provides estimates of state-dependent IRFs. Second, several years into Paul Volcker's tenure as Chairman of the Federal Reserve, coinciding with Ronald Reagan's presidency, the effectiveness of discretionary government expenditure as a tool for stimulating economic activity persistently declined. Third, the relationship with the stage of the business cycle undergoes a structural break in the 1980s: NBER recessions are characterized by local peaks in the multiplier values before the late 1980s, and the relationship inverts during the last two recessions of our sample period. Furthermore, we demonstrate that, until 1990, interest rate spreads effectively predict future shifts in the multiplier value. These findings underscore the importance of a careful consideration of the relationship between fiscal policy and the business cycle, while also pointing to a potential monetary explanation.

Figure 2 shows that the time-varying responses of output to a government expenditure shock exhibit state-dependent behaviour. In order to maintain clarity, this section focuses on responses of output and resulting multipliers. Similar to Pereira and Lopes (2014), we focus on the median IRFs to minimise the effect of occasionally unstable draws. The remaining IRFs can be found in Appendix A. Our approach provides a substantial degree of time variation in the post-WWII timeline in the US. Unlike Pereira and Lopes (2014), our estimates depend on the stage of business cycles. We also observe a significant change in both the shapes and magnitudes of output responses before and after the late 1980s: negative output responses to discretionary government expenditure shocks emerge during and after the global financial crisis.

The majority of heterogeneity in the IRFs emerges at distant horizons. The output response on impact remains approximately close to unity throughout the entire sample of years. Regarding the immediate impact, government expenditure shocks exhibit a straightforward Keynesian accounting effect: an additional US$1 of government expenditure increases aggregate demand by roughly the same amount in the same quarter. Various crowding-in and crowding-out effects emerge over the medium to long term. With this in mind, we seek to determine whether these shifts in long-term effects depend on the stage of the business cycle.

Examining the multiplier measures, it becomes evident that they indeed vary with the stage of the business cycle. Table 1 presents the average, minimum, and maximum values for each type of multiplier estimated using the baseline model. The highest multiplier values consistently occur during recessions, as defined by the NBER, but only in those preceding the late 1980s. In contrast, the 4-year cumulative multiplier reaches its lowest point during the global finan-

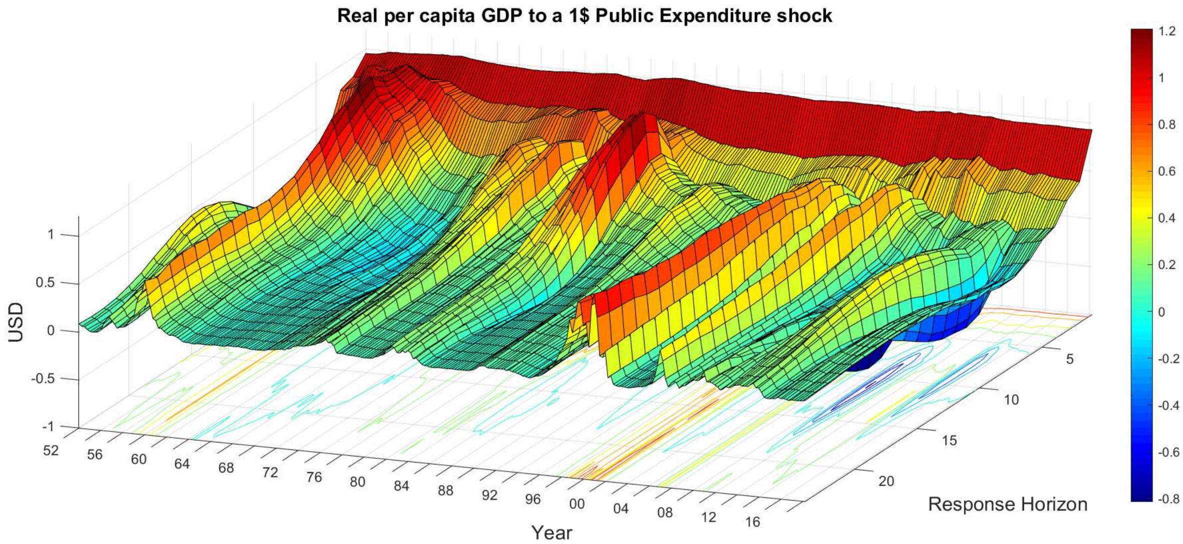

Figure 2: Median output response to an US$1 government expenditure shock as function of time: obtained via a mixture of signs and zero restrictions. Output response is measured in real US dollars.

Table 1: Descriptive statistics for the estimated multiplier series.

Multiplier	Average	Min		Max	
		date	value	date	value
Impact	0.97***	2017Q4	0.93***	1960Q4	1.03***
Sum (1-year)	0.89**	2011Q1	0.65*	1958Q1	1.38**
Sum (2-year)	0.83	2014Q4	0.05	1982Q4	2.23*
Sum (4-year)	0.78	2009Q2	-0.64	1958Q1	2.29*
Sum (5-year)	0.77	2014Q4	-0.33	1958Q1	2.35*

* $p < 0.32$, ** $p < 0.1$, *** $p < 0.05$

cial crisis. Plotting the estimated multipliers over time provides further insights into how this relationship has evolved.

Figure 3 presents the estimated path of the impact and cumulative government expenditure multipliers. The choice of the confidence bands follows the general pattern in the TVP-VAR literature. All three multiplier series demonstrate a pronounced decrease that begins in the 1980s. The 1-year cumulative multiplier almost exclusively falls below the impact value after 1985. Focusing on the 2-year multiplier, it is evident that more long-term crowding out takes place at the same time; discretionary government expenditure seems to be especially ineffectual in the aftermath of the global financial crisis. In line with Kirchner et al. (2010), Pereira and Lopes (2014) and Berg (2015), we conclude that the potential of government expenditure in stimulating output fell sharply after the 1980s.

The baseline model successfully estimates the multiplier series that are higher during recessions; however, the estimated relationship has certain limitations. First, the multiplier is

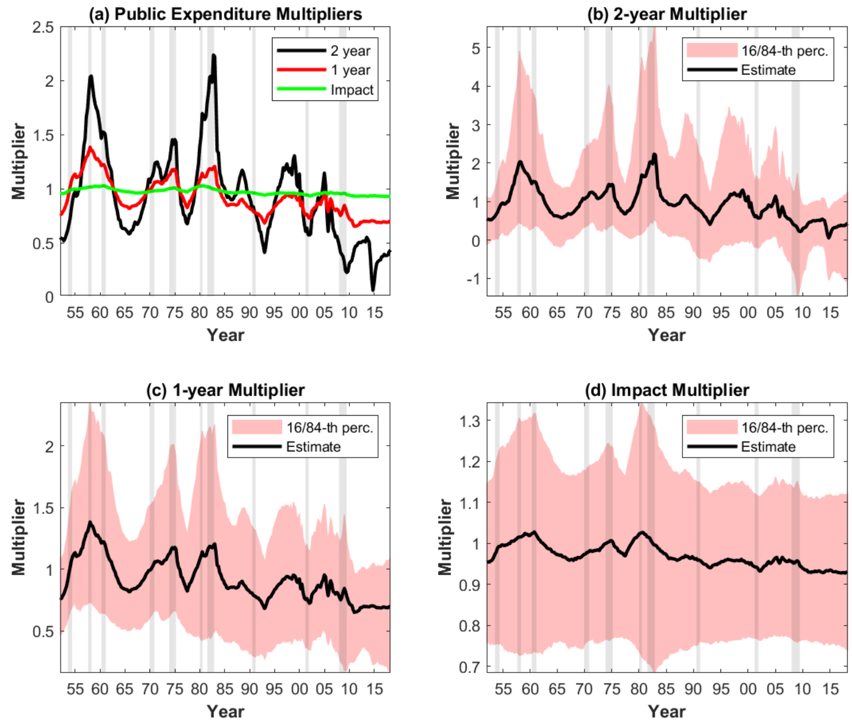

Figure 3: Public expenditure multipliers over NBER recession dates. Median multiplier values are presented in figure (a). The rest of the figures contain multipliers along with respective confidence bands, for 2-year (b) and 1-year (c) cumulative multipliers as well as the impact multiplier (d). Confidence bands are in red, calculated as 16th and 84th percentiles of the posterior multiplier distributions.

only higher during recessions before the late 1980s. The 2-year cumulative multiplier series demonstrate spikes around the recessions of 1957–58, 1974–75, and 1981–82. Second, at some point after the 1980s, the relationship changes its direction. During the two most recent US recessions, the 2-year cumulative government expenditure multiplier witnesses a decline. This evidence supports our hypothesis that the mixed results presented by Ramey and Zubairy (2018) may not be attributed to the fact that fiscal policy does not robustly depend on the stage of the business cycle but to the fact that such a relationship may not be constant over time.

We conclude that the post-WWII period in the US can be divided into two parts. The first part stretches from the beginning of our sample to the late 1980s, while the second proceeds until the modern-day. The relationship between the stage of the business cycle and the fiscal multiplier changes between these two periods.

We further confirm that the estimated multiplier values depend on the stage of the business cycle before the late 1990s, using a forward-looking indicator of economic slack. The slope of the yield curve (or the term spread) has been widely considered an early-stage predictor of the US recessions. Such spread represents a mix of market sentiments regarding the future path of the economy and the monetary policy playbook. We seek to determine if the spreads between the federal funds rate and yields on bonds of different maturities can predict future multiplier values. We obtain five interest rate spreads from the federal reserve economic database: (i) 10-year treasury constant maturity minus federal funds rate, (ii) 5-year treasury constant maturity minus federal funds rate, iii) 1-year treasury constant maturity minus federal funds rate, (iv) 6-month treasury bill minus federal funds rate, and (v) 3-month treasury bill minus federal funds rate.

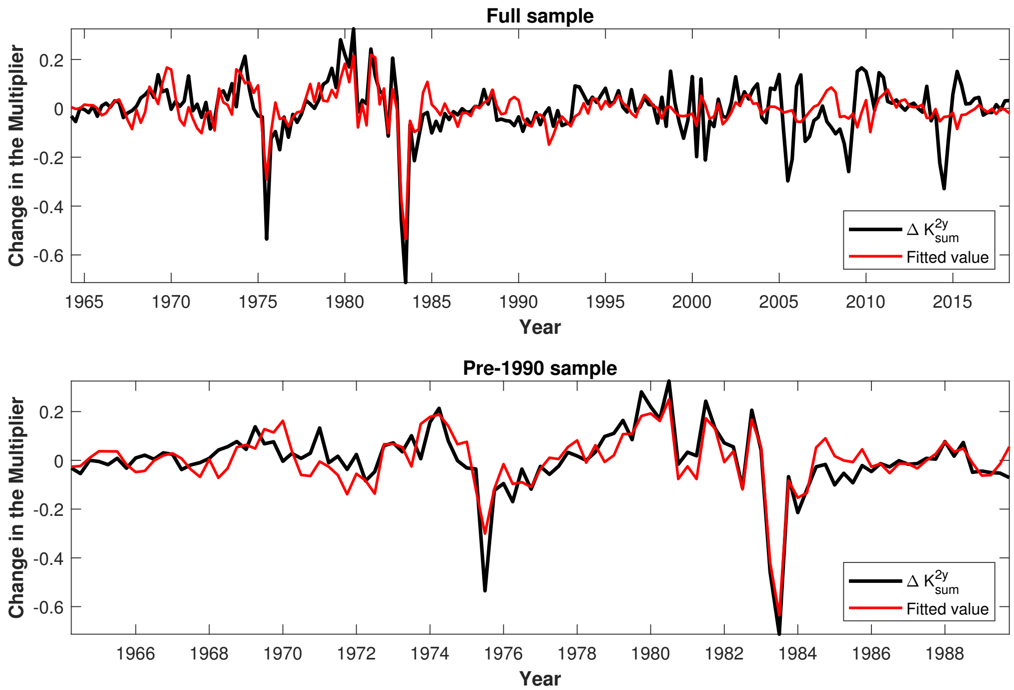

Figure 4: Predictability of changes in the 2-year cumulative multiplier value by the interest rate spreads.

We project the multiplier value estimated in our baseline estimation on the eight lags of these spreads. Given that the TVP-VAR coefficients are modelled as random walks, we estimate the linear regression in first differences. A modelling choice distinguishes our study from both Kirchner et al. (2010) and Berg (2015) that perform a similar ex post estimation in levels. We can see in Figure 4 that past yield spreads perform well in explaining the change in the 2-year cumulative multiplier before 1990. Nonetheless, this relationship seems negligible in the subsequent period. The fit improves if we limit estimation to observations before 1990.

This result is robust to the case of the multiplier calculated using unanticipated discretionary shocks considered in the next section. This result highlights three key insights. First, it confirms that the expenditure multiplier was counter-cyclical prior to the late 1980s. Second, there appears to be a structural break in the relationship between the multiplier and the business cycle during the 1980s. Third, the differing oscillations of the term spread across business cycle stages in the two periods (see also, Figure 10) suggest that a possible reason for this structural break is the shift in the monetary policy framework, along with the resulting changes in how economic agents form expectations.

5 Structural shocks and policy anticipation

The results from the previous section should be interpreted with caution. As widely discussed in the literature, econometricians may have access to a more limited information set than a representative market agent. Consequently, the structural shocks estimated by a TVP-VAR model based on such restricted information sets are likely to be anticipated by market agents in advance.[18] Structural shocks identified by such models represent a linear combination of the true past and future structural shocks, leading to biased IRFs (Mertens and Ravn, 2010).

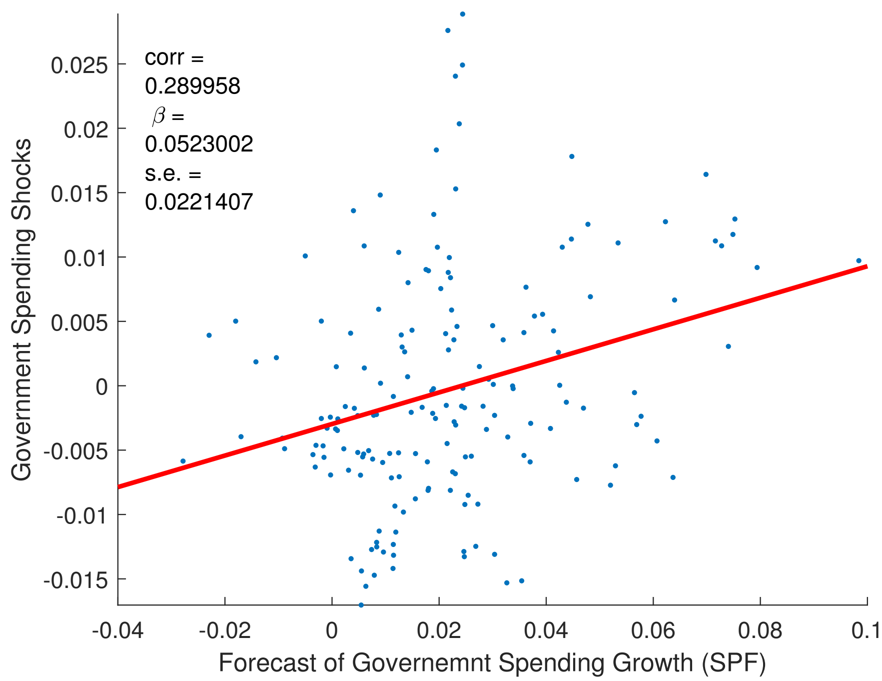

Figure 5: Forecastability of VAR shocks to government spending.

Figure 5 displays a scatter plot that juxtaposes the structural shocks identified through our baseline methodology with government expenditure growth forecasts derived from the Survey of Professional Forecasters. Our analysis of the estimated slope coefficient suggests that, at least in the previous quarter, the shocks identified by our model align with the expectations of professional forecasters. However, it is important to highlight that the R^2 value, at 0.0408, indicates that forecasters can predict only a small fraction of the variance in our structural

[18] Ramey (2011a) presents evidence suggesting that shocks identified using conventional VAR methods fail to account for surprise discretionary expenditures. The author concludes that professional forecasts of government expenditure and war dates Granger-cause VAR structural shocks.

shocks.[19] This observation motivates a closer examination of the timing issue and its potential impact on the series of estimated multipliers.

Two primary methodologies address the issue of shock anticipation in the analysis of fiscal multipliers. The first method involves utilizing narrative shock series, which are inherently unanticipated. Notable instances in the government expenditure multiplier literature encompass events such as the military build-up dates by Ramey and Shapiro (1998), along with military-related news highlighted by Ramey (2011b) and Ramey and Zubairy (2018). On the other hand, the second approach leverages professional forecasts to filter out the anticipated element from the shock. This approach has been extended in various ways, including the utilization of Auerbach and Gorodnichenko (2012)'s ST-VAR model, the expectations-augmented TVP-VAR framework proposed by Berg (2015), and the development of unanticipated shock measures by Auerbach and Gorodnichenko (2013), Abiad et al. (2016), Tsangarides et al. (2017), Furceri and Li (2017) and Furceri et al. (2018).

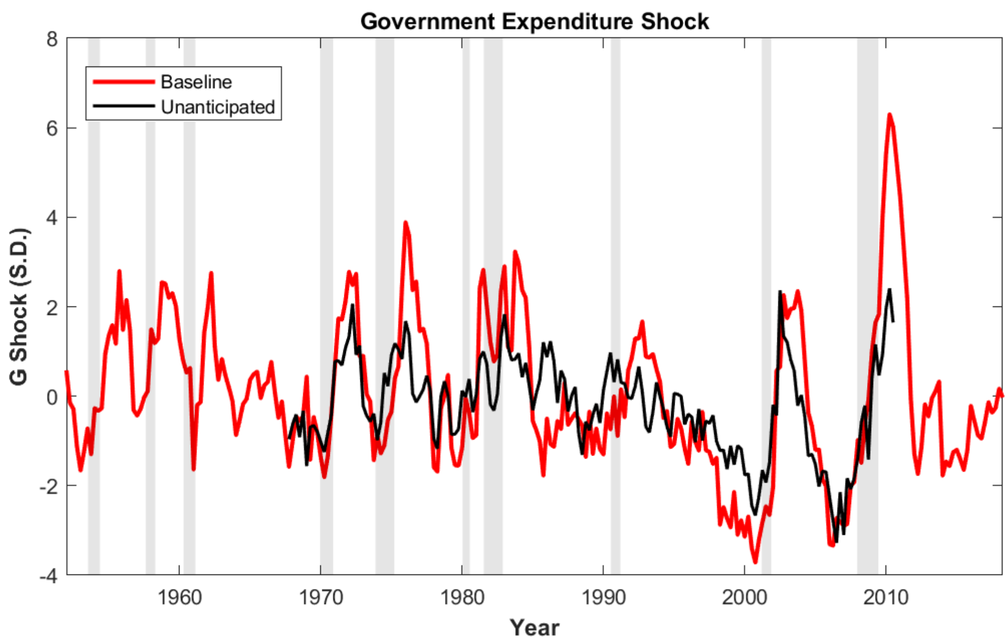

Figure 6: Government expenditure shocks identified by the baseline model (red) and the extension that controls for the professional forecasts (black), over NBER recession dates.

The first method, employing narrative shock series, offers distinct advantages but comes with drawbacks. Constructing an appropriate narrative shock measure can be a challenging task, as these shocks often pertain to specific, unique events. For instance, Ramey and Shapiro (1998)'s dates or the military news events analyzed by Ramey and Zubairy (2018) focus on

[19] Mertens and Ravn (2010) have argued that the challenges in estimating conventional VARs related to the degree of anticipation tend to decrease as the market agents' ability to foresee future shocks diminishes. This insight provides additional context for our examination of the dynamics of these shocks and their impact on fiscal multipliers.

military spending shocks. While they are valuable for understanding the impact of unanticipated discretionary military expenditures, they may not serve as an accurate measure of the overall potency of various government expenditure types.

In this section, we adopt the methodology introduced by Auerbach and Gorodnichenko (2012) and later extended by Berg (2015) to integrate professional forecasts into our TVP-VAR model. This approach offers a robust framework for addressing the challenge of shock anticipation in our analysis. However, unlike Auerbach and Gorodnichenko (2012) and Berg (2015), we do not examine the dynamic effects of expectations within our model. Instead of incorporating professional forecasts into the endogenous vector of variables, we focus solely on eliminating the anticipated component from our structural shocks. To achieve this, we control only for the current forecast made in the previous quarter.

This modeling decision represents a careful balancing act between upholding the model's integrity, as per the arguments presented by Mertens and Ravn (2010) and Ramey (2011a), and the practical computational constraints stemming from an expanded TVP-VAR model size. In alignment with the methodology outlined in Auerbach and Gorodnichenko (2012), we opt to employ forecasted growth rates instead of levels, given the historical volatility and revisions in the Survey of Professional Forecasters.

Given the relatively limited number of observations in our dataset and the significant parameterization inherent in our TVP-VAR model, the deliberate choice is made to exclude forecasts related to government revenues, as a comprehensive measure of such forecasts is unavailable for a significant portion of our sample. Additionally, the decision is taken to exclude forecasts of output. This choice aligns with the belief that output shocks are anticipated by market agents to a relatively limited extent compared to discretionary fiscal policy shocks. Consequently, the omission does not compromise the overall validity of our model.

At this point, it should be mentioned that the inclusion of forecasts in the TVP-VAR will lower the significance of our results. We expect most of the predicting power of our endogenous lag polynomial to be mirrored in the professional forecast. We estimate this extension on the same sample as in Auerbach and Gorodnichenko (2012): 1966Q4—2010Q3. Figure 6 presents the structural government expenditure shocks from our baseline along with shocks from the model extension controlling for professional forecasts.

Controlling for professional forecasts modifies our baseline results in three key ways. First, impact multipliers increase, with an unanticipated US$1 discretionary government expenditure shock raising output by an average of US$1.23. Second, all cumulative multipliers are lower than those in the baseline model. Third, as shown in Table 2, the maximum and minimum values of the estimated multipliers shift to different dates. Despite these changes, our main finding remains intact: the highest multiplier values occur in the earlier part of the sample (before the late 1980s), while the 2-year, 4-year, and 5-year cumulative multipliers reach their lowest levels around the global financial crisis.

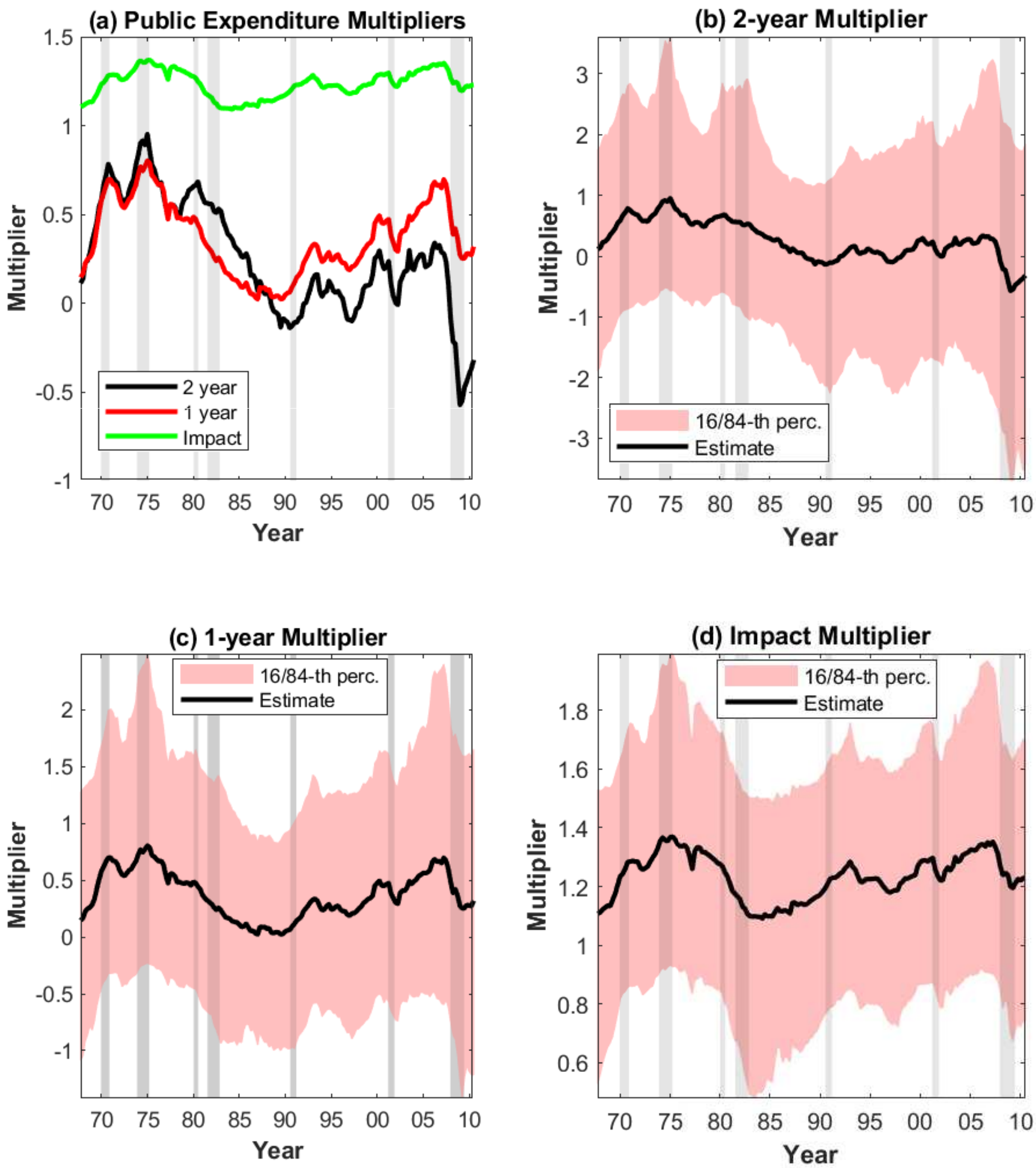

Figure 7: Unanticipated public expenditure multipliers over NBER recession dates. Median multiplier values are presented in the figure (a). The rest of the figures contains multipliers along with respective confidence bands, for 2-year (b) and 1-year (c) cumulative multipliers as well as the impact multiplier (d). Confidence bands are in red, calculated as 16th and 84th percentiles of the posterior multiplier distributions.

Figure 7 presents the dynamics of the estimated multipliers calculated using this extension. The shift in the state-dependent relationship is also evident in the case of the extended model. The 1-year and 2-year multiplier series exhibit local peaks around recessions before the late 1980s;

Table 2: Descriptive statistics for the estimated multiplier series: the case of unanticipated discretionary shocks.

Multiplier	Average	Min		Max	
		date	value	date	value
Impact	1.23***	1984Q2	1.09*	1975Q2	1.37***
Sum (1-year)	0.36	1987Q1	0.02	1975Q1	0.80
Sum (2-year)	0.29	2009Q1	-0.57	1975Q1	0.95
Sum (4-year)	0.36	2008Q4	-0.57	1974Q4	0.85
Sum (5-year)	0.32	2007Q2	-0.39	1975Q1	0.81

* $p < 0.32$, ** $p < 0.1$, *** $p < 0.05$

however, these peaks are substantially less pronounced than those in the baseline specification. After the 1980s, during the last two US recessions, a prominent decline can be observed. We can conclude that controlling for policy anticipation does not provide sufficient basis to challenge the result of the baseline setup—the relationship between the stage of the business cycle and government expenditure multiplier is not constant over time.

6 Multipliers' dynamics

6.1 Drivers

The estimated fiscal multiplier series during the post-WWII period exhibits significant variation, raising a critical question: what drives this fluctuation? Previous studies have noted a decrease in average multiplier values, often attributing this trend to broader economic phenomena such as the Great Moderation (Kirchner et al., 2010; Berg, 2015) and shifts in the U.S. policy mix (Leeper et al., 2017). However, the existing literature on the determinants of fiscal multipliers falls short in explaining the cyclical patterns observed in post-1980s estimates. This gap indicates that additional or evolving factors may be influencing these variations, necessitating further investigation into the underlying causes.

In this section, we advance two primary arguments. First, we establish a connection between the average multiplier size estimated by our model and existing research on fiscal-monetary interactions; for example, Davig and Leeper (2011); Leeper et al. (2017); Bianchi and Ilut (2017); Bianchi et al. (2023); Ascari et al. (2023). By aligning regime probabilities from Bianchi and Ilut (2017) with the multiplier trajectory generated by our TVP-VAR model, we find that multipliers tend to be larger during periods characterized by a fiscally-led policy mix compared to those dominated by monetary-led policies. This observation corroborates Leeper et al. (2017)'s conclusions, independent of the general equilibrium frameworks typically employed.

Furthermore, we argue that the peaks in fiscal multipliers observed during earlier U.S. recessions were likely due to shifts in the policy mix; from conflicting monetary and fiscal strategies to a more fiscally-led mix. This highlights the crucial role of policy changes in amplifying government expenditure multipliers during economic downturns.

Second, we present a new empirical challenge to the existing literature on fiscal multipliers. Our analysis of the post-1980s period reveals an unexpected pattern: in some instances, multipliers decline during recent recessions, contradicting the widely held assumption that they tend to rise in such conditions. We suggest that this anomaly may be related to the changing role of financial frictions in the recent U.S. economic history.

6.2 Fiscal multiplier and the fiscal-monetary policy mix

From a theoretical perspective, an extreme case of a fiscally-led policy mix is characterized by a central bank that remains largely unresponsive to inflation, coupled with a government that does not adhere to a fiscal rule aimed at stabilizing the debt-to-GDP ratio. In this scenario, fiscal policies are active while monetary policies are passive.[20] Under these conditions, the government is able to run substantial primary deficits without triggering an increase in the debt-to-GDP ratio over the long term. In such a framework, fiscal stimulus is perceived by market participants as unfunded, implying that larger fiscal deficits today are not expected to result in higher fiscal surpluses in the future; a characteristic of a non-Ricardian fiscal regime.

Without the classical Ricardian equivalence at work, households and businesses view additional fiscal stimulus as less costly. This prompts them to boost spending. As monetary policy does not tighten in response, inflation expectations rise, but nominal interest rates remain low. This combination stabilizes the debt-to-GDP ratio because inflation reduces the real value of debt while interest payments remain low due to the passive stance of monetary policy.

In a policy mix dominated by monetary policy, the central bank prioritizes stable inflation and at the same time, fiscal policy is required to generate surpluses to keep the debt-to-GDP ratio around a predetermined target. Under this arrangement, fiscal stimulus is perceived by market participants as funded, as they trust in the authorities' commitment to their respective goals. The central bank's focus on inflation stabilization anchors inflation expectations, while the private sector believes that the government will eventually need to run fiscal surpluses to maintain public debt stability relative to GDP. This belief, in line with Ricardian equivalence, influences consumption and savings behavior, leading to a muted response in private demand to fiscal stimulus. As a result, both the debt-to-GDP ratio and inflation expectations remain stable.[21]

Conversely, in a policy mix with conflicting objectives, the central bank's efforts to control inflation clash with a fiscal authority that does not prioritize debt stabilization. Here, the unsustainable trajectory of primary fiscal balances can fuel inflationary pressures, directly through increased demand and indirectly through concerns about potential monetary financing. This situation forces the central bank to raise real interest rates more aggressively to anchor inflation

[20] As defined by Leeper (1991).

[21] Importantly, the realized reaction of the central bank to any inflationary pressures from the fiscal shock also contributes to a lower multiplier value.

expectations. While this limits the overall impact on private demand, inflation expectations remain stable only if the central bank remains firmly committed to its inflation target. Meanwhile, the debt-to-GDP ratio rises, along with the interest payments on the accumulated debt.

Following this view, the government expenditure multipliers should be higher under the fiscally-led regime as the expansion in private demand is more pronounced. Therefore, we expect shifts to a fiscally led regime to amplify government expenditure multipliers.

We utilize regime probabilities from Bianchi and Ilut (2017) to evaluate how the average fiscal multiplier value evolved with shifts in the policy mix in the U.S. Bianchi and Ilut (2017) employ a Markov-Switching New Keynesian model to segment the period spanning from 1954Q4 to 2009Q3 into three regimes, aligning with the fiscally led, monetary led, and conflicting policy mixes described above. The estimated fiscal multipliers, derived from unanticipated fiscal shocks, are plotted against these regime probabilities in Figure 8. Three key observations are concluded from this figure as follows.

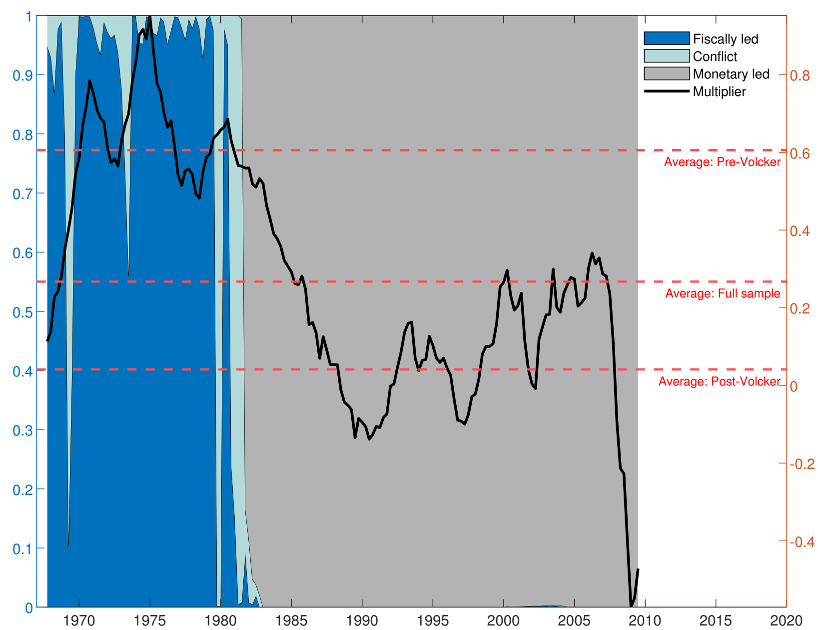

Figure 8: Average Fiscal Multiplier and Regime Probabilities from Bianchi and Ilut (2017)

First, multiplier values are notably higher during the pre-Volcker era (before 1979Q3), a period typically marked by high probabilities of a fiscally led policy mix. Second, a consistent decline in these values occurs during Paul Volcker's tenure as Fed Chairman (1979Q3-1987Q3), aligning with an initial phase of high conflict regime probabilities, followed by a shift to a monetary-led regime. Third, in the post-Volcker period (after 1987Q3), fiscal multipliers further diminish, corresponding with sustained high probabilities of a monetary-led policy mix.

This analysis corroborates the findings of Leeper et al. (2017). This shows that fiscal multipliers are indeed larger under a fiscally-led policy mix. This conclusion is drawn from an empirical model that imposes minimal structural constraints. However, our estimates deviate from those presented in Leeper et al. (2017)'s structural model across both policy regimes. Specifically, our model estimates a 10-quarter cumulative multiplier average of 0.6 for the pre-Volcker period, which is significantly lower than the 1.2-1.6 range suggested by Leeper et al. (2017). In the post-Volcker period, our average multiplier is 0.1, in contrast to the 0.5-0.7 range projected by their simulations. Our estimates also suggest that unanticipated shocks can be contractionary (i.e. produce negative multipliers) in the monetary-led regime, thus deviating from results in Ascari et al. (2023).

Shifts in the policy mix help explain why fiscal multipliers were higher during recessions in the earlier part of our sample period. Before Volcker's tenure, peaks in fiscal multipliers often followed periods of monetary policy loosening. The pronounced term spread inversions between the late 1960s and early 1980s, shown in Figure 10, coincide with the conflict regimes identified by Bianchi and Ilut (2017) in Figure 8, as well as periods of political pressure on the Federal Reserve documented by Drechsel (2024). This suggests that brief disinflationary efforts encountered political resistance, prompting a return to a more accommodative monetary stance and ultimately shifting the policy mix back to a fiscally led regime.

These policy shifts often coincided with U.S. recessions, which explain why some studies, such as Auerbach and Gorodnichenko (2012), attribute the rise in multiplier values to recessions rather than to underlying policy changes, as discussed by Leeper et al. (2017). This interpretation is supported by our multiplier estimates in Figure 9 which shows that before Volcker, negative term spreads, typically preceding policy shifts, were associated with increases in future multiplier values. However, this relationship vanishes in the post-Volcker era.

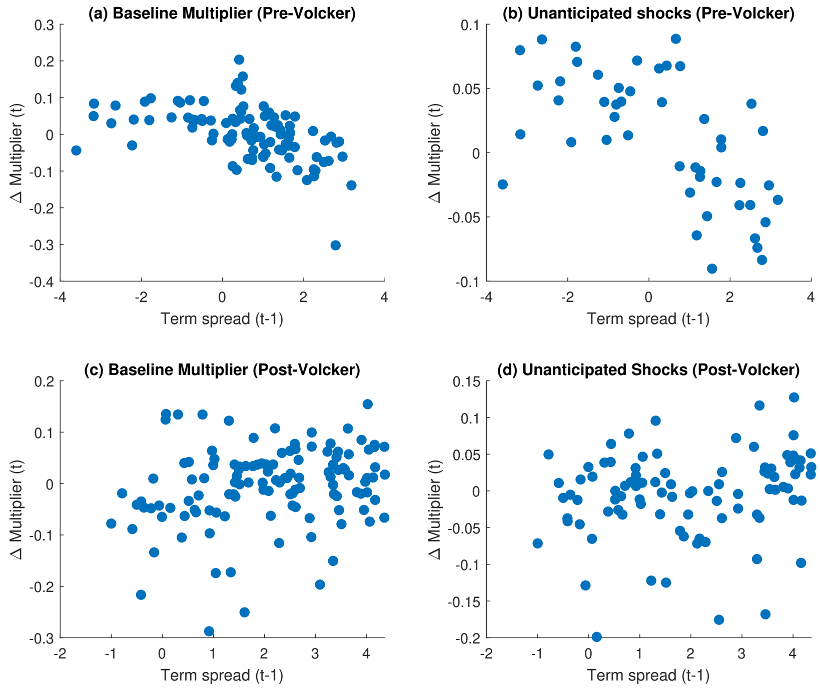

Figure 9: Change in the fiscal multiplier and lagged term spread

In the post-Volcker period, term spreads were no longer deeply negative before recessions, the likelihood of a fiscally led regime during U.S. recessions approached zero, and political pressure on the Federal Reserve diminished significantly (Drechsel, 2024). While this explains why multipliers did not increase during recessions in the latter half of our sample, it does not fully clarify why fiscal multipliers declined during some of the more recent recessions, particularly during the global financial crisis. This unresolved issue is explored further in the following section.

6.3 The financial friction puzzle

Our model reveals medium-term variations in fiscal multipliers that challenge current literature, especially studies on countercyclical fiscal multipliers. Contrary to the findings of Auerbach and Gorodnichenko (2012), our results do not show a consistent increase in fiscal multipliers during recessions. While this pattern holds for the pre-Volcker era, it diverges during the recessions of 1990-1991, 2001, and 2007-2009 across various model specifications. As discussed earlier, the higher fiscal multipliers observed in earlier recessions were linked to changes in policy mix, which were absent in the more recent recessions. However, the sharp decline in multiplier values during the global financial crisis remains an unresolved issue.

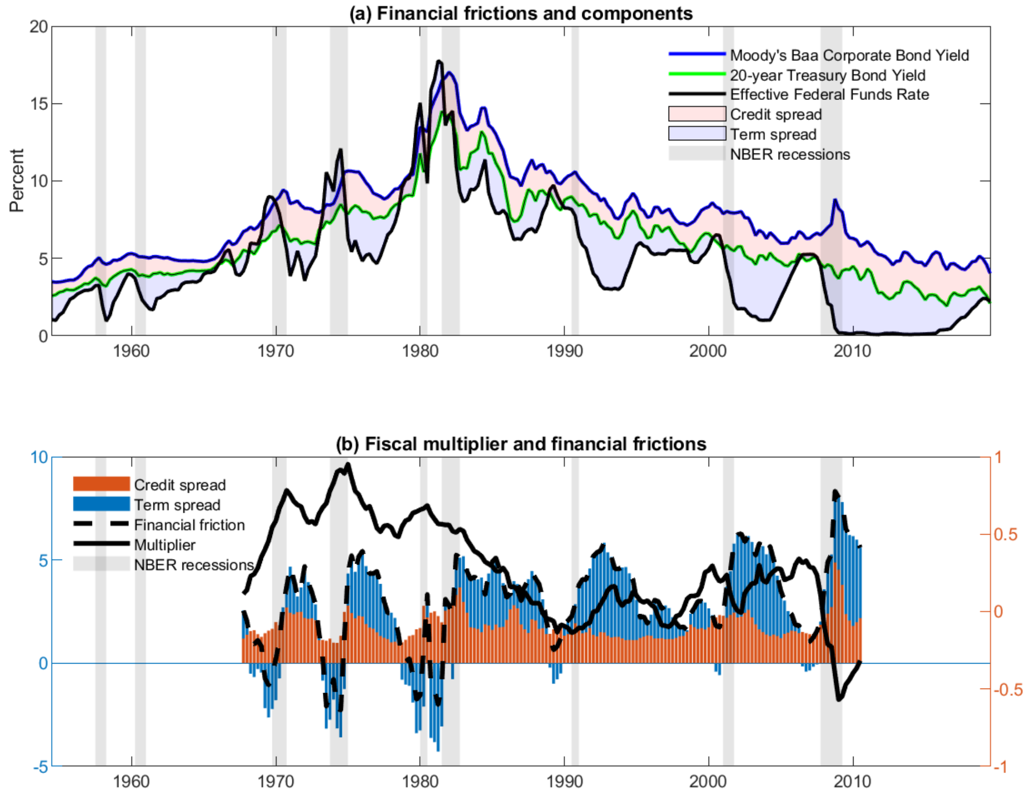

Figure 10: Fiscal multiplier, yield curve, and financial frictions in the post-WWII U.S.

The global financial crisis, however, provides a crucial context for understanding this decline. The 2008 crisis, representing the largest asset bubble burst in the U.S. since WWII, led to a significant increase in financial frictions, as evidenced by a dramatic widening of credit spreads (see Figure 10), and a substantial slowdown in credit growth. The ratio of private debt to income more than doubled compared to the earlier part of our sample, highlighting the severe constraints on borrowing. This scenario aligns with empirical findings, such as those by Gilchrist and Zakrajšek (2012), which show that increased financial frictions can substantially weaken the effectiveness of fiscal policy. Additionally, the high levels of household debt exacerbated the recession's impact, as constrained consumer spending further reduced the fiscal multiplier's effectiveness, echoing the arguments made by Mian and Sufi (2010), Andrés et al. (2015) and Justiniano et al. (2015).

To investigate whether financial frictions could have played a role in the significant declines in multiplier values during the global financial crisis and the aftermath of the dot-com bubble burst, we segment our sample into four distinct periods: (a) the post-WWII era of relative stability (1950s-60s); (b) the inflationary and economic turbulence of the 1970s; (c) the deregulation and globalization of the 1980s-90s; and (d) the disruptions in financial markets during the 2000s-10s.

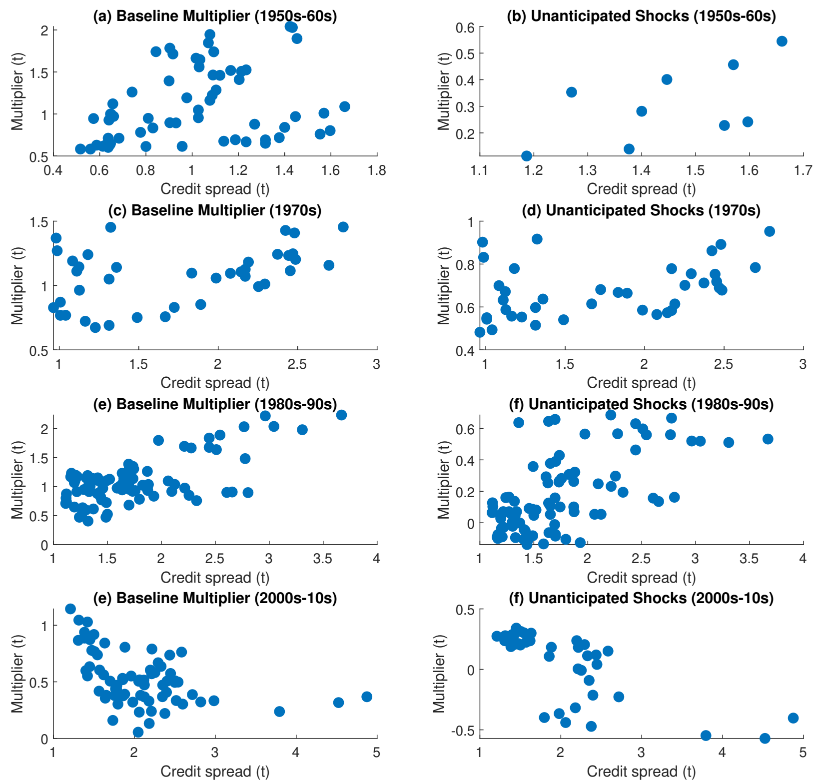

Figure 11: Fiscal multiplier and financial frictions

Figure 11 provides some evidence that the relationship between fiscal multipliers and financial frictions changed during the 2000s-10s. The historically positive relationship between fiscal multipliers and credit spreads appears to have reversed in recent years. This shift indicates that transformations in the financial environment, namely deeper global financial integration, monetary policy constraints at the zero lower bound, and rising household debt, may have fundamentally altered the transmission of fiscal policy.[22]

Our findings indicate a departure from the dominant academic view that financial frictions enhance fiscal multipliers by promoting economic growth and reducing financial intermediation costs.[23] While evidence of this relationship emerges in the earlier part of our sample, it appears to reverse in the 2000s, raising important questions about the evolving nature of fiscal transmission. Moreover, the forward-looking nature of credit and term spreads complicates the statistical

[22]The increased interconnectedness of global financial markets also has amplified the transmission of financial shocks, as noted by Reinhart and Rogoff (2008), causing that financial frictions in one part of the world can quickly affect fiscal multipliers elsewhere, particularly in highly integrated economies like the U.S.

[23]See, for example, Carrillo and Poilly (2013) and Canzoneri et al. (2016).

identification of their connection to fiscal multipliers, as endogeneity concerns pose significant challenges.

7 Conclusion

In conclusion, this paper introduces government expenditure multipliers that vary based on the phase of business cycles by employing a TVP-VAR framework. This challenges the prevailing literature that assumes a stable relationship between fiscal multipliers and business cycle phases over time. Our findings reveal that the relationship between multipliers and the stages of business cycles is not constant, with average multipliers becoming significantly smaller since the 1980s. We attribute these changes to policy shifts rather than the inherent characteristics of recessions, as suggested in previous literature. Furthermore, we uncover a new puzzle: while financial frictions historically amplified multipliers, this relationship has reversed in the 21st century. This is particularly evident during the global financial crisis when our estimated multiplier is much lower than the post-WWII average.

While our findings highlight distinct patterns in fiscal multipliers and introduce a novel transmission channel, further research is needed to expand the scope of this study. One potential direction is to employ less flexible but more efficient estimators, similar to those used by Ramey and Zubairy (2018) and Bernardini and Peersman (2018), to validate the relationships identified in this paper. Another avenue for future research is to test these findings using a larger fiscal VAR that incorporates additional variables, such as public debt and inflation, or to extend the analysis to a broader set of advanced economies.

Investigating the changing role of financial frictions within a theoretical framework, similar to Canzoneri et al. (2016), could help explain why multiplier values became negatively correlated with financial frictions during the Dot-com recession and the global financial crisis. Additionally, a more detailed examination of the evolving fiscal-monetary policy mix, particularly in the context of unconventional monetary policies, as explored by Bi and Traum (2023), could offer valuable insights into how these dynamics influence fiscal multipliers.

References

Abiad, A., Furceri, D., and Topalova, P. (2016). The macroeconomic effects of public investment: Evidence from advanced countries. *Journal of Macroeconomics*, 50:224–240.

Afonso, A., Baxa, J., and Slavík, M. (2018). Fiscal developments and financial stress: a threshold var analysis. *Empirical Economics*, 54:395–423.

Alloza, M. (2022). Is fiscal policy more effective during recessions? *International Economic Review*, 63(3):1271–1292.

Andrés, J., Boscá, J. E., and Ferri, J. (2015). Household debt and fiscal multipliers. *Economica*, 82:1048–1081.

Ascari, G., Beck-Friis, P., Florio, A., and Gobbi, A. (2023). Fiscal foresight and the effects of government spending: It's all in the monetary-fiscal mix. *Journal of Monetary Economics*, 134:1–15.

Auerbach, A. and Gorodnichenko, Y. (2012). Measuring the output responses to fiscal policy. *American Economic Journal: Economic Policy*, 4(2):1–27.

Auerbach, A. and Gorodnichenko, Y. (2013). Output spillovers from fiscal policy. *American Economic Review Papers and Proceedings*, 103(3):141–146.

Auerbach, A., Gorodnichenko, Y., and Murphy, D. (2020). Local fiscal multipliers and fiscal spillovers in the usa. *IMF Economic Review*, 68:195–229.

Bachman, R. and Sims, E. R. (2012). Confidence and the transmission of government spending shocks. *Journal of Monetary Economics*, 59(3):235–249.

Bachmann, R. and Sims, E. R. (2012). Confidence and the transmission of government spending shocks. *Journal of Monetary Economics*, 59(3):235–249.

Barro, R. and Redlick, C. J. (2011). Macroeconomic effects from government purchases and taxes. *Quarterly Journal of Economics*, 126(1):51–102.

Belmonte, M., Koop, G., and Korobilis, D. (2014). Hierarchical shrinkage in time-varying parameter models. *Journal of Forecasting*, 33:80–94.

Berg, T. (2015). Time varying fiscal multipliers in germany. *Review of Economics*, 66(1):13–46.

Bernanke, B. S. (2022). *21st Century Monetary Policy: The Federal Reserve from the Great Inflation to COVID-19*. WW Norton & Company.

Bernardini, M. and Peersman, G. (2018). Private debt overhang and the government spending multiplier: Evidence for the united states. *Journal of Applied Econometrics*, 33:485–508.

Bhar, R. and Malliaris, A. G. (2021). Modeling us monetary policy during the global financial crisis and lessons for covid-19. *Journal of Policy Modeling*, 43(1):15–33.

Bi, H. and Traum, N. (2023). Unconventional monetary policy and local fiscal policy. *European Economic Review*, 156:104476.

Bianchi, F., Faccini, R., and Melosi, L. (2023). A fiscal theory of persistent inflation. *The Quarterly Journal of Economics*, 138:2127–2179.

Bianchi, F. and Ilut, C. (2017). Monetary/fiscal policy mix and agents' beliefs. *Review of economic Dynamics*, 26:113–139.

Binning, A. (2013). Underidentified svar models: A framework for combining short and long-run restrictions with sign-restrictions. *Available at SSRN 2350094*.

Blanchard, O. and Perotti, R. (2002). An empirical characterization of the dynamic effects of changes in government spending and taxes on output. *Quarterly Journal of Economics*, 107(4):1329–1368.

Blinder, A. (1981). Thoughts on the laffer curve, the supply-side effects of economic policy. *Economic Policy Conference Series*, 1.

Callegari, M., Melina, G., and Batini, N. (2012). Successful austerity in the united states, europe and japan. *IMF Working Papers 2012/190*.

Canova, F. and Pappa, E. (2007). Price differentials in monetary unions: The role of fiscal shocks. *The Economic Journal*, 117(520):713–737.

Canzoneri, M., Collard, F., Dellas, H., and Diba, B. (2016). Fiscal multipliers in recessions. *The Economic Journal*, 126(590):75–108.

Carrillo, J. A. and Poilly, C. (2013). How do financial frictions affect the spending multiplier during a liquidity trap? *Review of Economic Dynamics*, 16(2):296–311.

Carter, C. K. and Kohn, R. (1994). On gibbs sampling for state space models. *Biometrika*, 81(3):541–553.

Chung, H. and Leeper, E. M. (2009). What has financed government debt? *NBER Working Paper*, WP 13425.

Cogley, T. and Sargent, T. J. (2005). Drifts and volatilities: monetary policies and outcomes in the post wwii us. *Review of Economic Dynamics*, 8(2):262–302.

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

Drechsel, T. (2024). Estimating the effects of political pressure on the fed: a narrative approach with new data. Technical report, National Bureau of Economic Research.

Eisenstat, E., Chan, J., and Strachan, R. W. (2016). Stochastic model specification search for time-varying parameter vars. *Econometric Reviews*, 35(8-10):1638–1665.

Frühwirth-Schnatter, S. and Wagner, H. (2010). Stochastic model specification search for gaussian and partial non-gaussian state space models. *Journal of Econometrics*, 154:85–100.

Furceri, D. and Li, B. G. (2017). The macroeconomic (and distributional) effects of public investment in developing economies. *International Monetary Fund Working Paper*, WP/17/217.

Furceri, J., Ge, J., Loungani, P., and Melina, G. (2018). The distributional effects of government spending shocks in developing economies. *International Monetary Fund Working Paper*, WP/18/57.

Ghiaie, H. and Rouillard, J.-F. (2022). Housing tax expenditures and financial intermediation. *Canadian Journal of Economics/Revue canadienne d'économique*, 55(2):937–970.

Gilchrist, S. and Zakrajšek, E. (2012). Credit spreads and business cycle fluctuations. *American economic review*, 102(4):1692–1720.

Hamilton, J. (2018). Why you should never use the hodrick-prescott filter. *Review of Economics and Statistics*, 100(5):831–843.

Hwang, I. and Kim, J. (2021). Oil price shocks and the us stock market: A nonlinear approach. *Journal of Empirical Finance*, 64:23–36.

Iwata, Y. and IIboshi, H. (2023). The nexus between public debt and the government spending multiplier: fiscal adjustments matter. *Oxford Bulletin of Economics and Statistics*.

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

Justiniano, A., Primiceri, G. E., and Tambalotti, A. (2015). Household leveraging and deleveraging. *Review of Economic Dynamics*, 18(1):3–20.

Kim, S., Burd, H., and Milligan, G. (1998). Model testing of closely spaced tunnels in clay. *Geotechnique*, 48(3):375–388.

Kirchner, M., Cimadomo, J., and Hauptmeier, S. (2010). Transmission of government spending shocks in the euro area: Time variation and driving forces.

Leeper, E. M. (1991). Equilibria under 'active'and 'passive'monetary and fiscal policies. *Journal of monetary Economics*, 27(1):129–147.

Leeper, E. M., Traum, N., and Walker, T. B. (2017). Clearing up the fiscal multiplier morass. *American Economic Review*, 107(8):2409–2454.

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

Levin, A. T. and Sinha, A. (2020). Limitations on the effectiveness of monetary policy forward guidance in the context of the covid-19 pandemic. Technical report, National Bureau of Economic Research.

Lucas Jr, R. E. (1976). Econometric policy evaluation: A critique. In *Carnegie-Rochester conference series on public policy*, volume 1, pages 19–46. North-Holland.

Mertens, K. and Ravn, M. O. (2010). Measuring the impact of fiscal policy in the face of anticipation: A structural var approach. *The Economic Journal*, 120(544):393–413.

Mian, A. R. and Sufi, A. (2010). Household leverage and the recession of 2007 to 2009. Technical report, National Bureau of Economic Research.

Mountford, A. and Uhlig, H. (2005). nwhat are the effects of fiscal policy shocks. Technical report, oHumboldt University, mimeo.

Mountford, A. and Uhlig, H. (2009). What are the effects of fiscal policy shocks? *Journal of Applied Econometrics*, 24(6):960–992.

Pereira, M. and Lopes, A. S. (2014). Time-varying fiscal policy in the u.s. *Studies in Nonlinear Dynamics & Econometrics*, 18(2):157–184.

Primiceri, G. (2005). Time varying structural vector autoregressions and monetary policy. *Review of Economic Studies*, 72(3):821–852.

Ramey, V. (2011a). Can government purchases stimulate the economy. *Journal of Economic Literature*, 49(3):673–685.

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

Ramey, V. and Shapiro, M. D. (1998). Costly capital reallocation and the effects of government spending. *Carnegie-Rochester Conference Series on Public Policy*, 48:145–194.

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

Reinhart, C. M. and Rogoff, K. S. (2008). Is the 2007 us sub-prime financial crisis so different? an international historical comparison. *American Economic Review*, 98(2):339–344.

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

Shaheen, R. and Turner, P. (2020). Fiscal multipliers and the level of economic activity: a structural threshold var model for the uk. *Applied Economics*, 52(17):1857–1865.

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

Stock, J. H. and Watson, M. W. (2012b). Generalized shrinkage methods for forecasting using many predictors. *Journal of Business & Economic Statistics*, 30(4):481–493.

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

Trabant, M. and Uhlig, H. (2011). The laffer curve revisited. *Journal of Monetary Economics*, 58(4):305–327.

Tsangarides, C. G., Arizala, F., Gonzalez-Garcia, J. R., and Yenice, M. (2017). The impact of fiscal consolidations on growth in sub-saharan africa. *International Monetary Fund Working Paper*, WP/17/281.

Appendices

A Estimation results

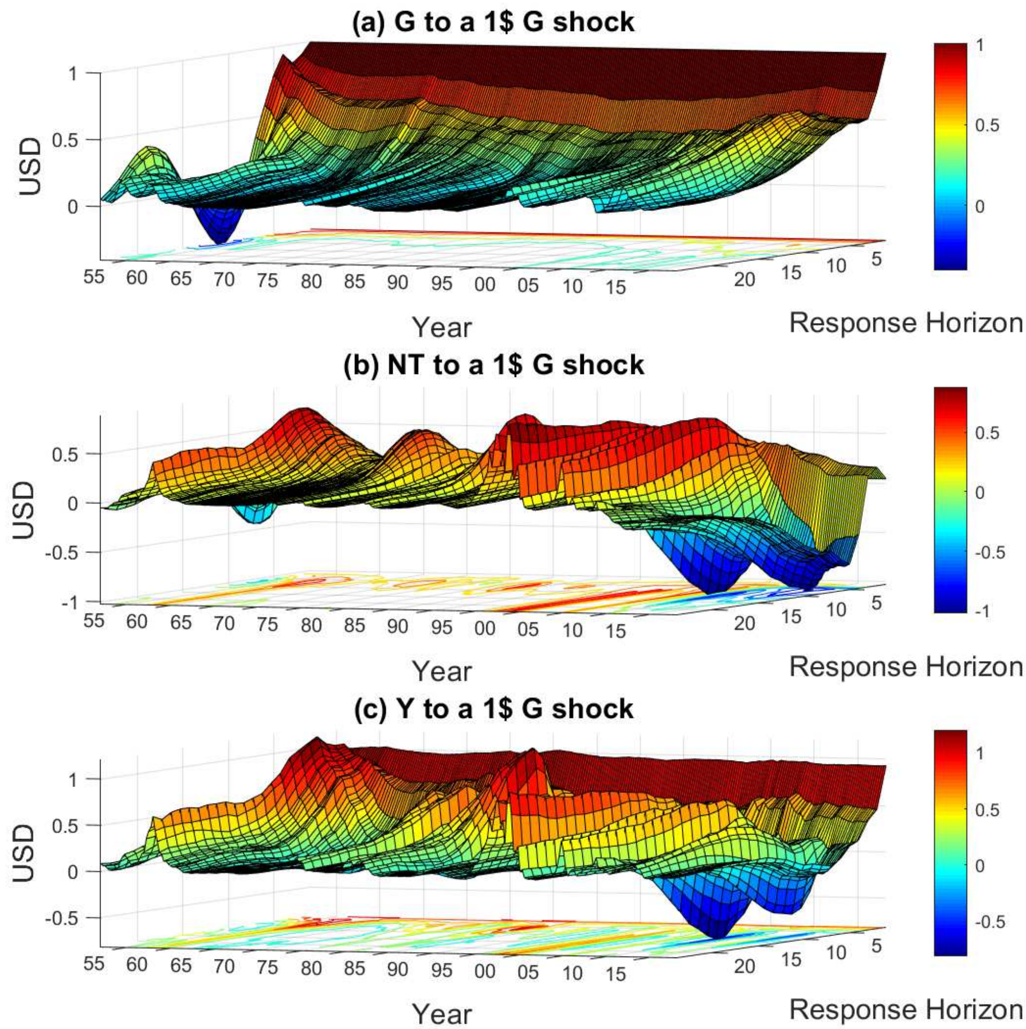

Figure 12: Baseline estimation. Median responses to discretionary government expenditure shocks as functions of time. Government expenditure response (a), net taxes' response (b), and output responses (c) to a US$1 government expenditure shock. Responses are measured in real U.S. dollars.

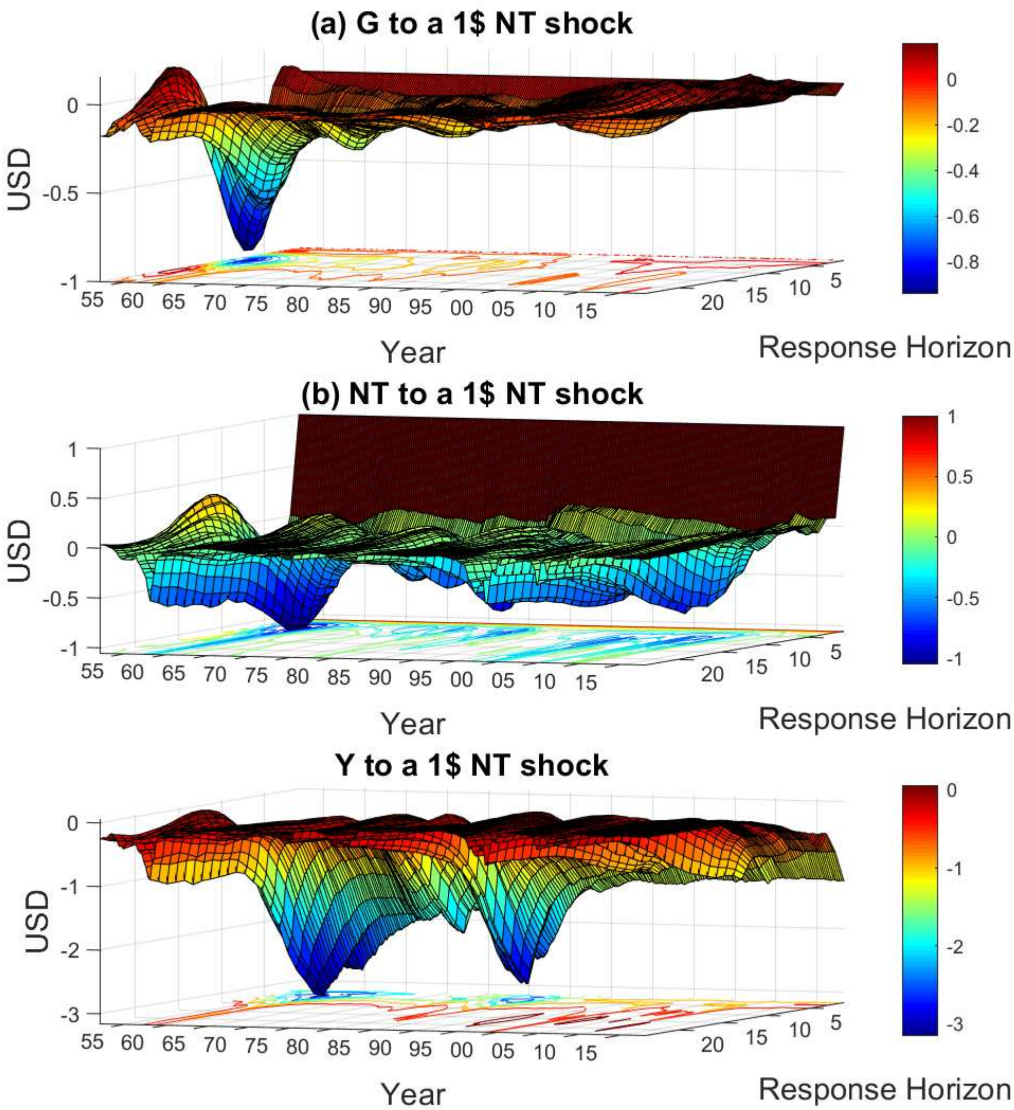

Figure 13: Baseline estimation. Median responses to tax shocks as functions of time. Government expenditure response (a), net taxes' response (b), and output responses (c) to a US$1 tax shock. Responses are measured in real U.S. dollars.

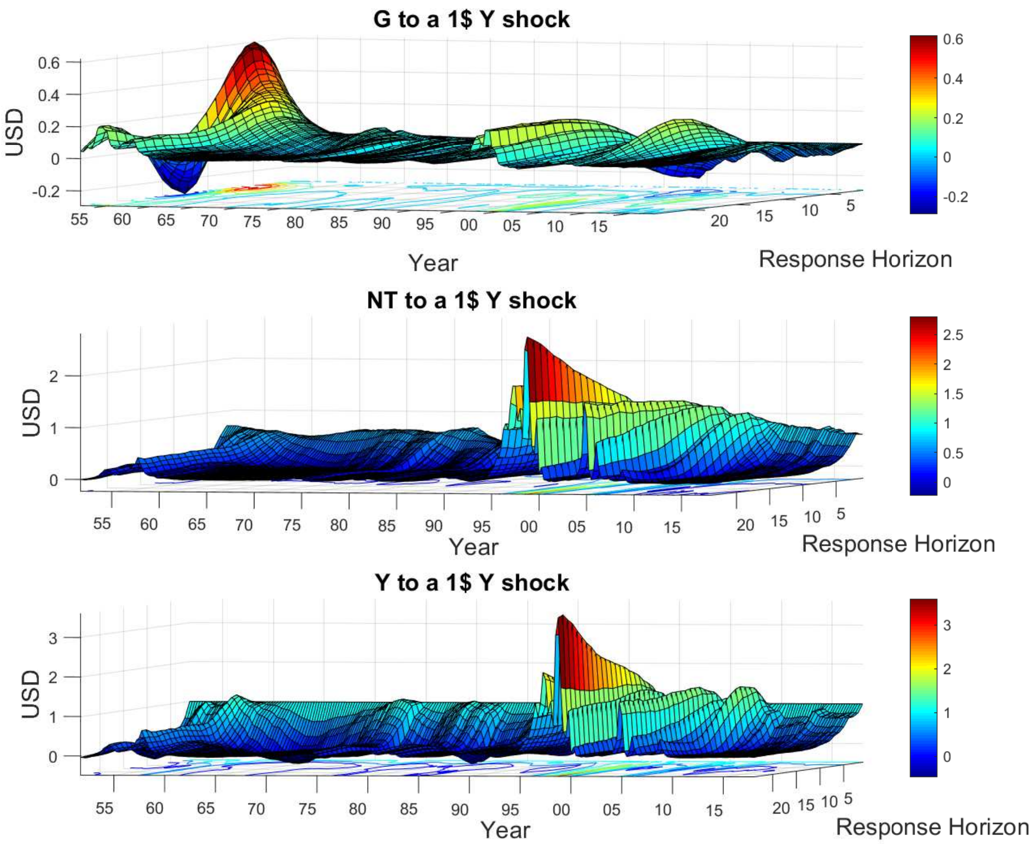

Figure 14: Baseline estimation. Median responses to output shocks as functions of time. Government expenditure response (a), net taxes' response (b), and output responses (c) to a US$1 output shock. Responses are measured in real U.S. dollars.

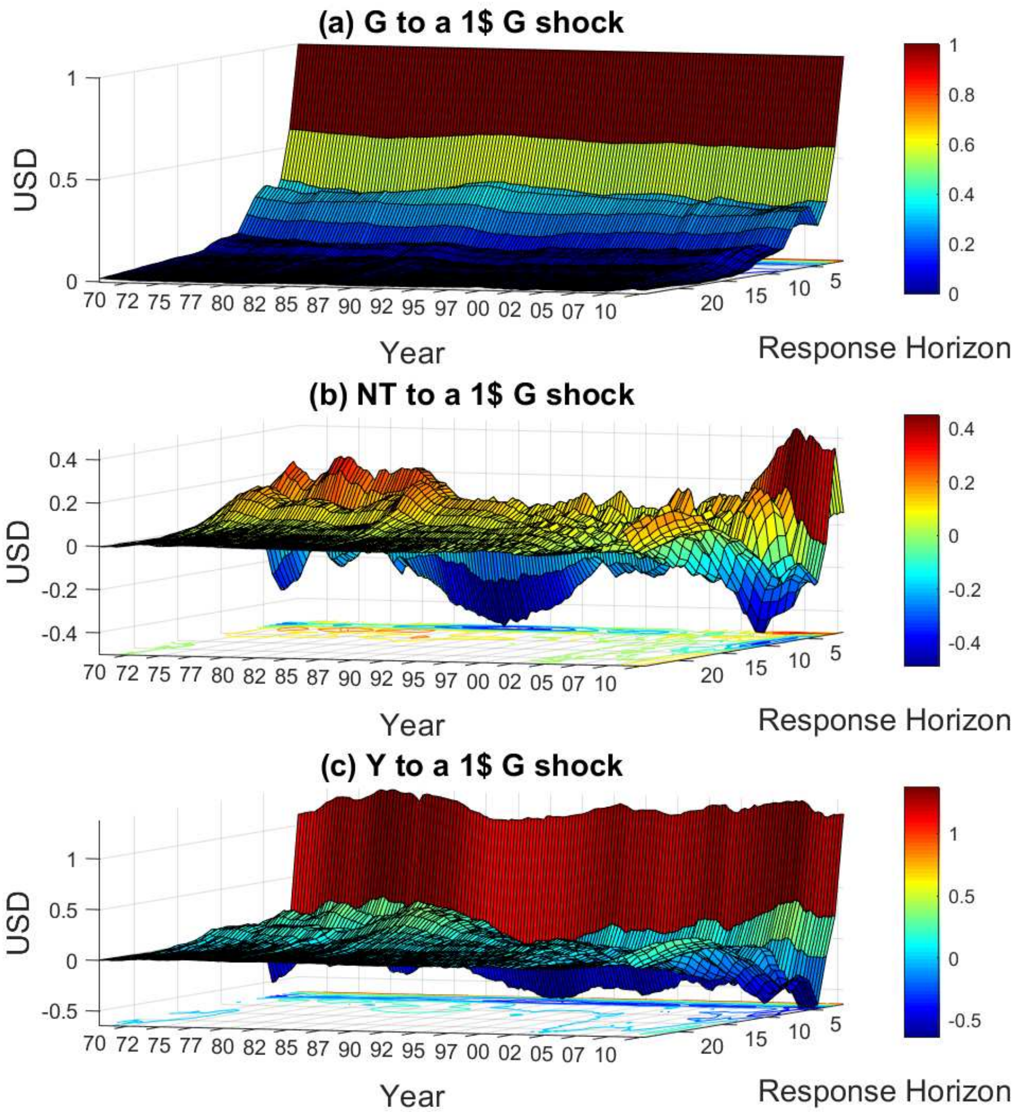

Figure 15: Unanticipated policy shocks. Median responses to government expenditure shocks as functions of time. Government expenditure response (a), net taxes' response (b), and output responses (c) to a US$1 unanticipated government expenditure shock. Responses are measured in real U.S. dollars.

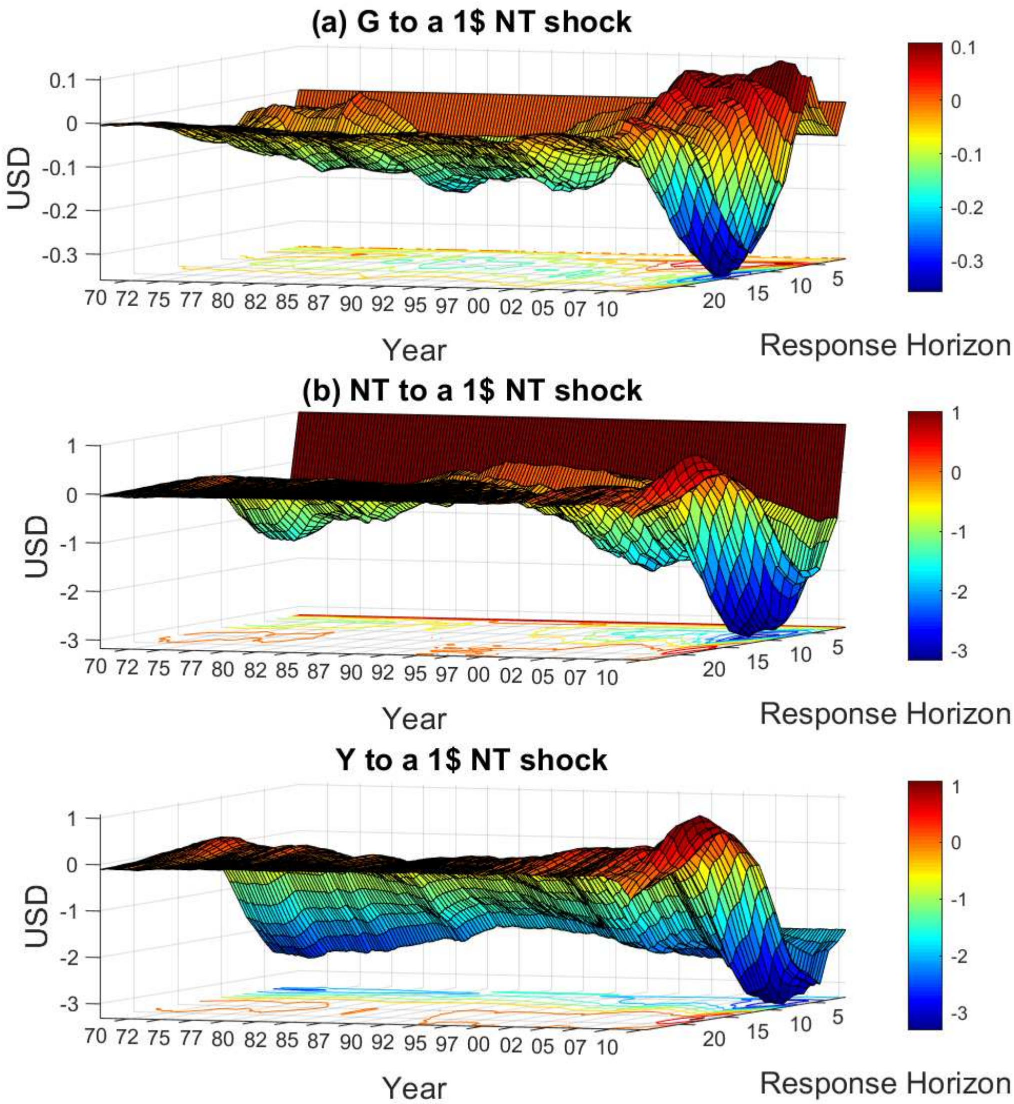

Figure 16: Unanticipated policy shocks. Median responses to tax shocks as functions of time. Government expenditure response (a), net taxes' response (b), and output responses (c) to a US$1 tax shock. Responses are measured in real U.S. dollars.

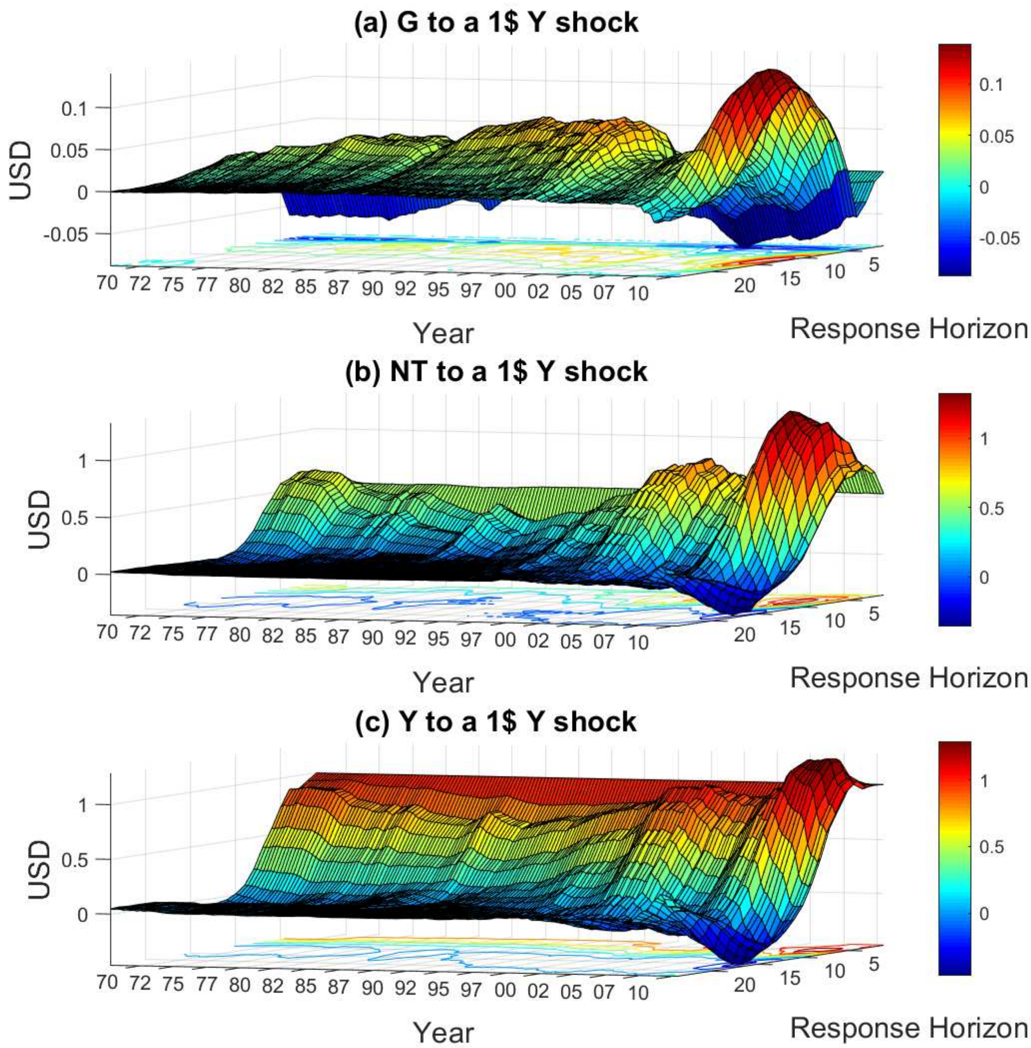

Figure 17: Unanticipated policy shocks. Median responses to output shocks as functions of time. Government expenditure response (a), net taxes' response (b), and output responses (c) to a US$1 output shock. Responses are measured in real U.S. dollars.

B Data preparation and sources

This study introduces Hamilton (2018)'s detrending procedure to the fiscal multiplier debate. The procedure makes use of a linear projection model, similar in spirit to direct forecasting or Jorda (2005) local projection method:

$$y_{t+h} = B(L)y_t + v_{t+h}, \qquad v_{t+h} \sim i.i.d.N(0, \sigma^2) \tag{B.1}$$

where $B(L)$ is the lag polynomial of the variable being detrended (y_t), h is the prediction horizon and v_{t+h} is an i.i.d. error term. In the case of the quarterly data, Hamilton (2018) recommends to set $h = 8$ and estimate the model with a lag polynomial of order 4. Resulting series of residuals ($\hat{v_{t+h}}$) would then represent a stationary zero-mean cyclical component. In essence, the method identifies the cyclical component as the forecast error that is due to macroeconomic developments taking place along those eight quarters.

Hamilton (2018)'s method has several advantages over alternative detrending procedures. First, it produces a non-linear trend estimate without the necessity to guess the functional form of such non-linearity. Second, due to the peculiarities of the method, it allows the trend to be influenced by macroeconomic events taking place in the past; as can be seen in Figure 18, the trend estimate experiences a pronounced dip in the aftermath of the global financial crisis.[24] Third, as argued by Hamilton (2018), the method does not produce spurious correlations between the resulting cyclical components and other macroeconomic series, as in the case of the Hodrick and Prescott (HP) filter. Fourth, it produces stationary series in levels. Estimating the TVP-VAR on such data allows us to interpret resulting IRFs as multipliers, not elasticities. Therefore, without the need to resort to the use of growth rates, this approach allows us to avoid the rescaling bias described in Ramey and Zubairy (2018). Finally, the method allows us to preserve a larger share of low-frequency variation in the target series.

Figure 19 presents the cyclical components of the real per capita GDP derived using Hamilton's method, HP filter and by removing a simple linear trend; to support the analysis we also plot first differences of the data and recession dates identified by NBER along with the cyclical components. First differences lack a significant share of the lower-frequency variation that contains crucial information necessary for correct fiscal policy evaluation.[25] Removal of the linear trend produces a cyclical components that does not revert to its mean for long periods of time; TVP-VAR models can be sensitive to the use of such time series. The cyclical component obtained using Hamilton's method delivers a compromise. The method preserves the mean-reverting nature of the series and allows for a sufficient share of low-frequency variation.

[24] One would be surprised if an event such a the global financial crisis had not affected the potential output or shifted the economy to a new growth path.

[25] This is precisely the reasoning used by Auerbach and Gorodnichenko (2012) to justify estimation of their ST-VAR in levels

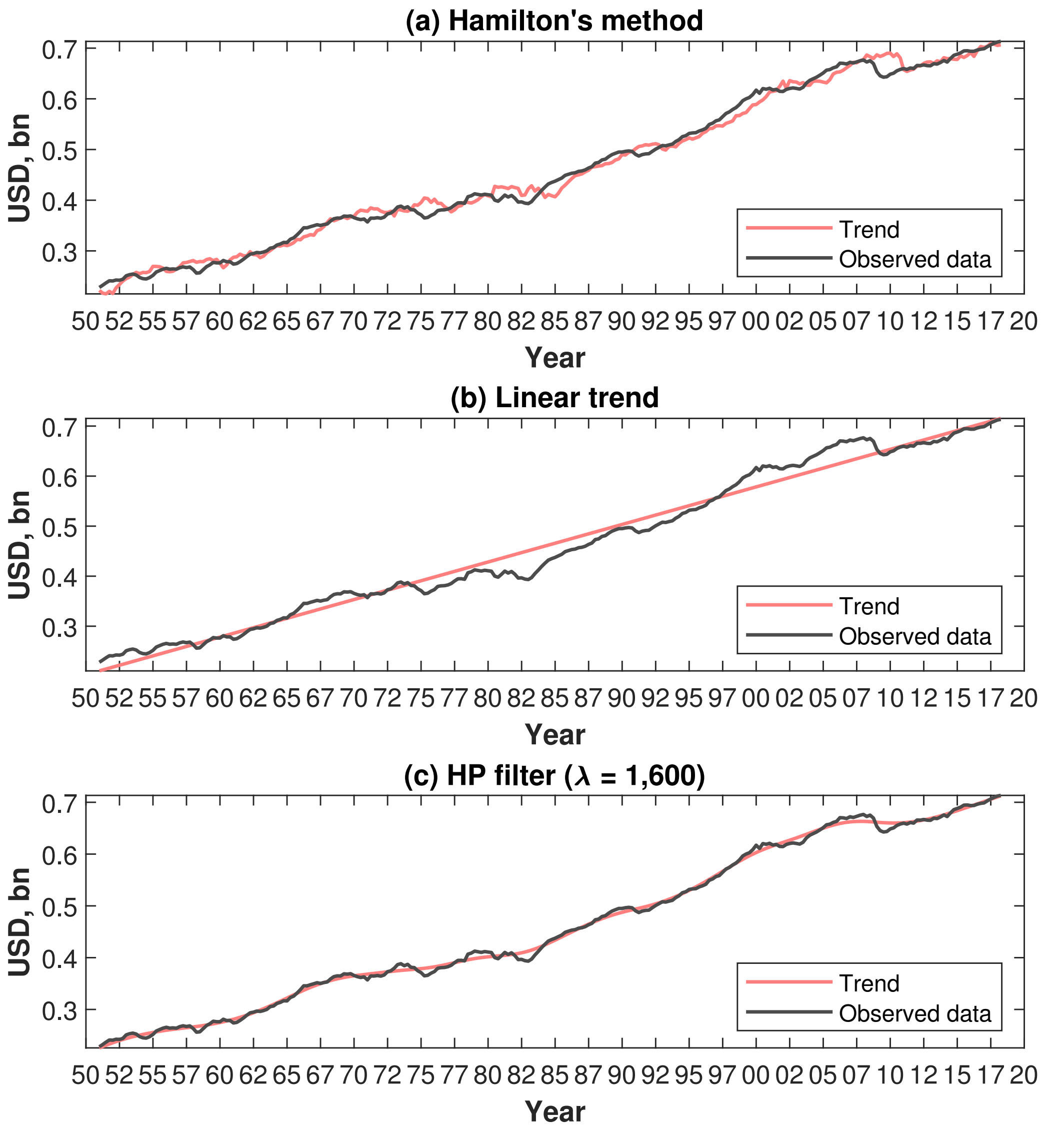

Figure 18: Detrending procedures and resulting trends. Observed US real per capita public expenditure (Investment+Consumption) along with trends produced by the Hamilton's linear projection technique (a), linear trend estimation (a) and the Hodrick-Prescott filter at $\lambda = 1,600$ (c).

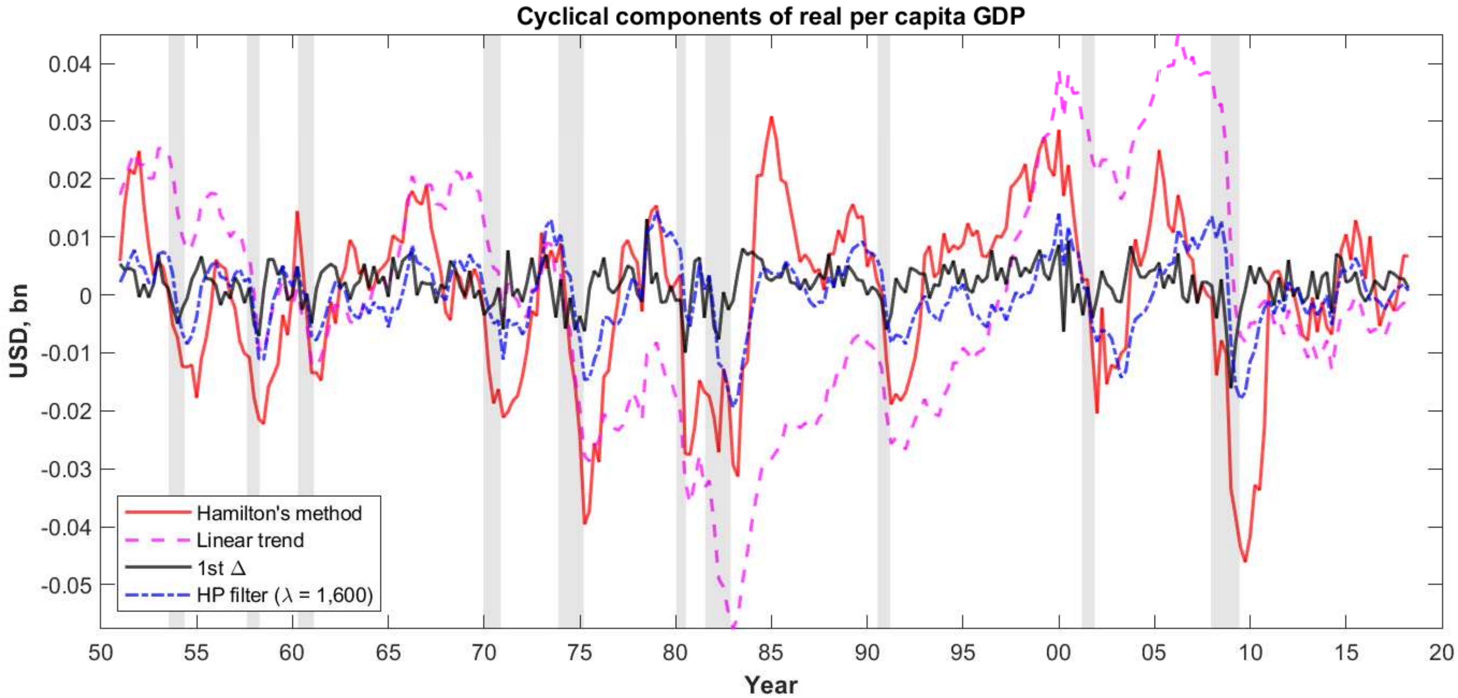

Figure 19: Transformations of GDP and low-frequency variation. Figure presents first differences of GDP, cyclical components obtained using the HP filter and Hamilton's method, and GDP without a linear time trend. Shaded areas are recessions defined by the NBER.

Table 3: Data transformations and sources.

Variable	Sample	Transformation and codes	Source
R.p.c* GDP	1948Q1 - 2018Q2	$\frac{T1.1.5L1}{T1.1.4L1 \times CNP16OV}$	BEA, FRED
R.p.c* public consumption	1948Q1 - 2018Q2	$\frac{T3.9.5L2}{T1.1.4L1 \times CNP16OV}$	BEA, FRED
R.p.c* public investment	1948Q1 - 2018Q2	$\frac{T3.9.5L3}{T1.1.4L1 \times CNP16OV}$	BEA, FRED
R.p.c* tax receipts	1948Q1 - 2018Q2	$\frac{T3.1L2+T3.1L7}{T1.1.4L1 \times CNP16OV}$	BEA, FRED
R.p.c* net taxes	1948Q1 - 2018Q2	$\frac{T3.1L2+T3.1L7+T3.1L10+T3.1L16-T3.1L30-T3.1L27-T3.1L22}{T1.1.4L1 \times CNP16OV}$	BEA, FRED
Forecast of public expenditure growth rate	1966Q4 - 2010Q3	$\Delta G_{t\|t-1}$	AG12
NBER recessions	1948Q1 - 2018Q2	$USRECQ$	FRED
10-Year to FFR spread	1962Q1 - 2018Q2	$GS10 - DFF$	FRED
5-Year to FFR spread	1962Q1 - 2018Q2	$DGS5 - DFF$	FRED
1-Year to FFR spread	1962Q1 - 2018Q2	$WGS1YR - DFF$	FRED
6-Month to FFR spread	1962Q1 - 2018Q2	$TB6MS - DFF$	FRED
3-Month to FFR spread	1962Q1 - 2018Q2	$TB3MS - DFF$	FRED

* Real per capita terms

$TxLy$ is the format used to denote a BEA NIPA Table x line y. BEA - Bureau of Economic Analysis, FRED - Federal Reserve Economic Database, AG12 - Auerbach and Gorodnichenko (2012).

C Fiscal VAR and Cholesky factorization.

Cholesky decomposition, used in Auerbach and Gorodnichenko (2012) and Ramey (2011b), although being the least cumbersome identification approach, is misleading in its application to a fiscal VAR where tax revenues are used instead of marginal tax rates. Although the resulting shock series are assumed to be independent, the immediate effect of the shock ordered first on the variable ordered last will contain immediate effects of shocks ordered in between. It is, therefore, crucial to ensure that shocks ordered in between are identified correctly.

Imposing a lower-unitriangular structure on the contemporaneous relations in $[G_t \quad T_t \quad Y_t]$ assumes that only innovations in T_t can contemporaneously affect Y_t, and not the other way around. Although such an approach seems to be justified in the case of discretionary government expenditure or marginal tax rates, it is not clear why output shocks cannot affect tax revenues in the same quarter. Let us consider a generic SVAR model with three endogenous variables:

$$Y_t = A_0 Y_t + B(L)Y_{t-1} + \varepsilon_t, \qquad \varepsilon_t \sim \mathcal{N}(0, \Theta) \tag{C.1}$$

where Θ is diagonal. The model in (8) has a reduced form representation:

$$Y_t = A(L)Y_{t-1} + u_t, \qquad u_t \sim \mathcal{N}(0, \Omega) \tag{C.2}$$

where Ω is a full symmetric matrix, that can be decomposed into a product of lower-unitriangular and diagonal matrices.

$$\Omega = P'P = C\Sigma\Sigma'C'$$

where the lower unitriangular matrix C contains the immediate responses of endogenous variables to the structural shocks. It is well known that the following relationship holds in this set-up:

$$C = (I_3 - A_0)^{-1}$$

Alternatively we can represent elements of C as functions of elements of A_0:

$$A_0 = \begin{bmatrix} 0 & 0 & 0 \\ \alpha_{21} & 0 & 0 \\ \alpha_{31} & \alpha_{32} & 0 \end{bmatrix} \qquad C = \begin{bmatrix} 1 & 0 & 0 \\ \alpha_{21} & 1 & 0 \\ \alpha_{31} + \alpha_{21}\alpha_{32} & \alpha_{32} & 1 \end{bmatrix}$$

Now let us return to the Auerbach and Gorodnichenko (2012) case; the immediate output response to a tax shock is given by α_{32}. It is equivalent to the coefficient of current period net taxes in the output equation in C.1. Therefore, it is impossible to constitute the direction of causality between output and taxes, as the coefficient can represent effects in both directions. The full set of restrictions imposed by Cholesky ordering is not sufficient to identify the tax shock. Restricting the reverse channel, discussed above, results in the α_{32} coefficient representing a mix between the two effects. That is precisely why Auerbach and Gorodnichenko (2012) had to rely on the elasticity measure defined by Blanchard and Perotti (2002) instead of depending solely on Cholesky in calculating IRFs to a tax shock, as they did with government expenditure shock. Figure 20 presents the time-varying response of output to a tax shock, the equivalent of α_{32} in our setup, identified via Cholesky decomposition; it is clear that the shock is not identified correctly. The immediate response captures the positive effect of output on the tax base, instead of the negative effect of the tax rate on output. Moreover, as it is evident from the matrix C above, α_{32} enters the calculation of the response of output to a government expenditure shock. Thus, one should not rely on the output responses identified using Cholesky in the Auerbach and Gorodnichenko (2012) setup, even if only output responses to a government expenditure shock are being investigated.

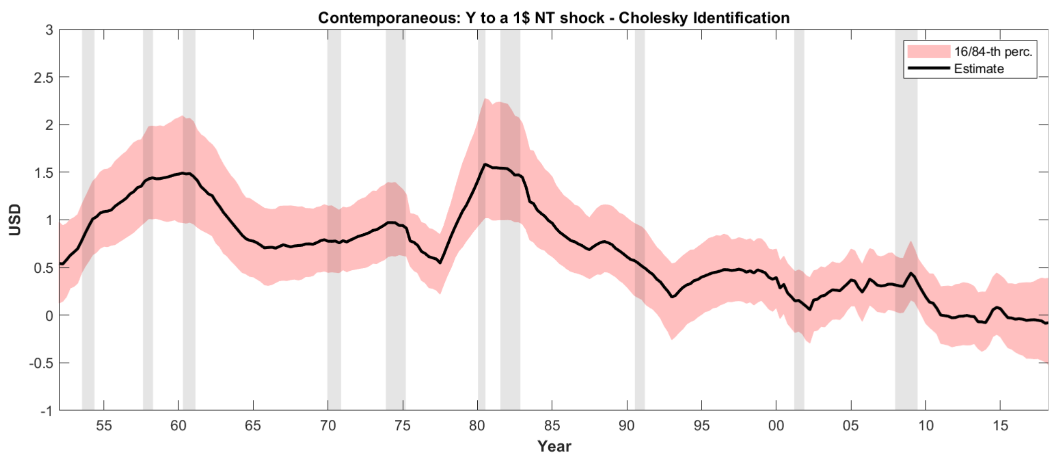

Figure 20: Immediate output response to a tax shock over time. Obtained using Cholesky decomposition.

Sign restrictions (e.g. Mountford and Uhlig (2009) and Canova and Pappa (2007)) or "narrative" (IV/Proxy) identification (e.g. Mertens and Ravn (2010), Stock and Watson (2012b), Stock and Watson (2012a) and Stock and Watson (2018)) should be preferred to a simple Cholesky decomposition in case of a fiscal VAR with tax revenues ordered after government expenditure.

D Identification via the mixture of short-term zero and sign restrictions

Our choice of the identification strategy, depicted in 3.1, constitutes an alternative way to solve the rebus discussed in Blanchard and Perotti (2002). Following the authors, let us assume that the reduced form shocks identified by our TVP-VAR are linear functions that can be expressed by the system of equations below:

$$\begin{aligned} u_t^G &= \alpha_t^{gy} u_t^Y + \beta_t^{gt} \varepsilon_t^T + \varepsilon_t^G \\ u_t^T &= \alpha_t^{ty} u_t^Y + \beta_t^{tg} \varepsilon_t^G + \varepsilon_t^T \\ u_t^Y &= \alpha_t^{yg} u_t^G + \alpha_t^{yt} u_t^T + \varepsilon_t^Y \end{aligned} \tag{D.1}$$

where u_t^G, u_t^T, and u_t^Y are reduced form shocks and ε_t^G, ε_t^T, and ε_t^Y are structural shocks. α_t^{gy} and α_t^{ty} capture the automatic response of the fiscal variables to changes in output (the automatic stabiliser effects) and the systematic discretionary response of fiscal variables to changes in output. We are interested in estimating the IRFs to the random discretionary shocks, in our case we focus on ε_t^G. In order to solve the above system of equations we need to impose a set of assumptions.

1. Following Blanchard and Perotti (2002) we assume $\alpha_t^{gy} = 0$. Such an assumption implies no automatic nor systematic discretionary responses of government expenditure to developments in output.[26] Since our set-up rests on Blanchard and Perotti we rely on their results on the role of automatic stabilisers; authors were not able to identify any automatic feedback from economic activity to government purchases.

2. Another restriction inspired by Blanchard and Perotti (2002) is that $\beta^{gt} = 0$. Authors argued that either β_t^{gt} or β_t^{tg} should be set to zero; since the correlation between government expenditure and net taxes is low, both restrictions produced similar results.

3. α_t^{ty} is positive. Allowing α_t^{ty} to be non-zero we imply that output shocks can affect net taxes through the tax base. Since we set it to be positive, we believe that a positive shock to output will expand the tax base and *vice verse*. Blanchard and Perotti (2002) directly estimate the coefficient as a function of two elasticities–elasticity of taxes to their respective tax bases and elasticity of the tax bases to GDP. Average value of α_t^{ty}, estimated on various sub-samples, remained positive.[27]

4. α_t^{yt} is negative. Blanchard and Perotti (2002) estimate a time-invariant coefficient directly for two cases–deterministic and stochastic trends. In both cases, the coefficient is negative and equals to -0.868 and -0.876, respectively.

Given the above-mentioned assumptions, we can show that the system in D.1 can be represented

[26] The absence of a systematic discretionary response is a consequence of the policy implementation lag; the policy-maker will need at least a quarter to come up and execute a discretionary government expenditure package in response to a surprise recession.

[27] Authors acknowledged that they focused on an average value of α_t^{ty}, while in reality, it should vary over time; our approach allows accounting for that.

as:

$$
\begin{aligned}
u_t^G &= \varepsilon_t^G \\
u_t^T &= \frac{\alpha_t^{ty}\alpha_t^{yg} + \beta_t^{tg}}{1-\alpha_t^{ty}\alpha_t^{yt}}\varepsilon_t^G + \frac{1}{1-\alpha_t^{ty}\alpha_t^{yt}}\varepsilon_t^T + \frac{\alpha_t^{ty}}{1-\alpha_t^{ty}\alpha_t^{yt}}\varepsilon_t^Y \\
u_t^Y &= \frac{\alpha_t^{yt}\beta_t^{tg} + \alpha_t^{yg}}{1-\alpha_t^{ty}\alpha_t^{yt}}\varepsilon_t^G + \frac{\alpha_t^{yt}}{1-\alpha_t^{ty}\alpha_t^{yt}}\varepsilon_t^T + \frac{1}{1-\alpha_t^{ty}\alpha_t^{yt}}\varepsilon_t^Y
\end{aligned}
$$

For simplicity, let us use the following matrix notation:

$$
\begin{bmatrix} u_t^G \\ u_t^T \\ u_t^Y \end{bmatrix} = \begin{bmatrix} c_t^{11} & 0 & 0 \\ c_t^{21} & c_t^{22} & c_t^{23} \\ c_t^{31} & c_t^{32} & c_t^{33} \end{bmatrix} \begin{bmatrix} \varepsilon_t^G \\ \varepsilon_t^T \\ \varepsilon_t^Y \end{bmatrix}
$$

If α_{ty} is positive and α_{yt} is negative, then c_{22} and c_{33} are both positive time-varying coefficients. These assumptions also imply that c_{32} is negative and c_{23} is positive. It is not necessary to impose the latter assumption, since imposing the former already results in c_{23} being positive; we show this in Figure 21. Finally, c_{11} is positive by definition. Our set of assumptions is not sufficient to identify signs of c_{21} and c_{31}.

This set of assumptions leaves us with an underidentified system and a partial understanding of coefficients' signs. We solve the system using a mixture of sign and short-term zero restrictions. We follow the technique described in Binning (2013) to identify structural shock with the following set of restrictions on the impact matrix:

$$
Z_t = \begin{array}{c} \\ G_0 \\ T_0 \\ Y_0 \end{array} \begin{array}{c} \begin{array}{ccc} \varepsilon_t^G & \varepsilon_t^T & \varepsilon_t^Y \end{array} \\ \begin{pmatrix} + & 0 & 0 \\ \times & + & \times \\ \times & - & + \end{pmatrix} \end{array} \tag{D.2}
$$

As depicted by our identification framework in 3.1, we allow for effects in both directions–output shocks affect tax revenues via the tax base, and tax shocks affect output by changing the tax rates. In doing so, we only restrict the contemporaneous response of output to a tax shock to be negative, without assuming any sign for the reverse channel. As can be seen in Figure 27, we do not need to impose the second restriction as our approach delivers contemporaneous effects of opposite signs.

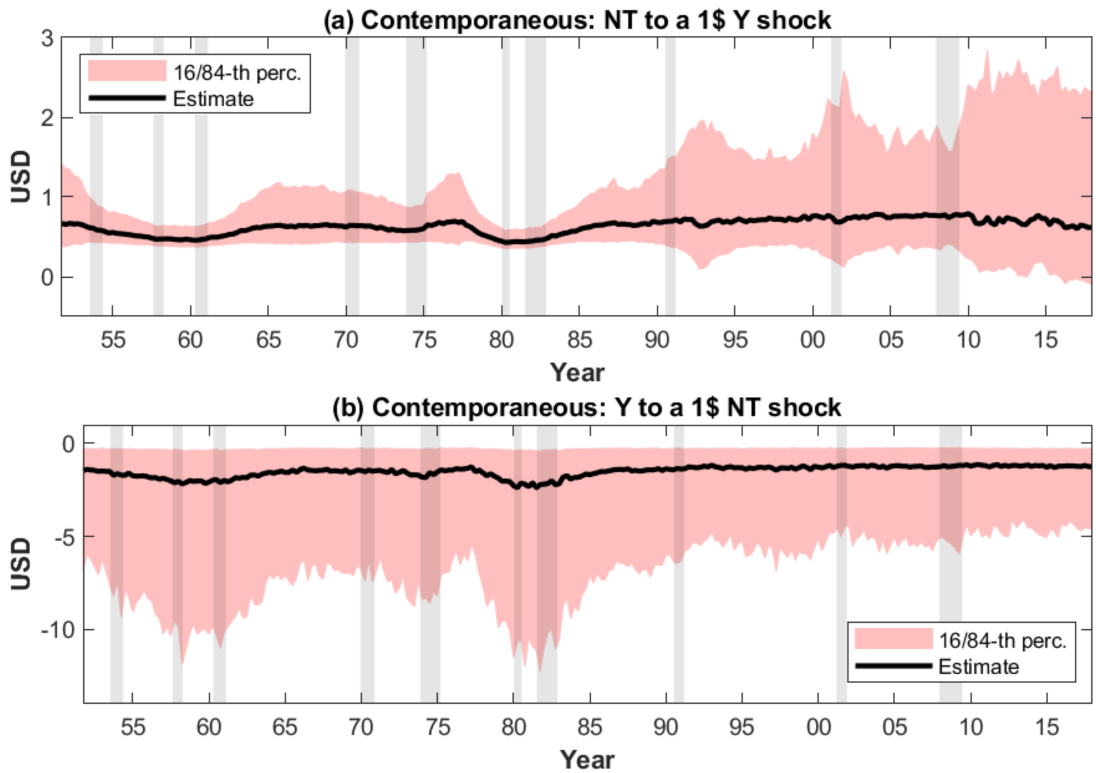

Figure 21: Identification of the two-way channel between taxes and output. Immediate responses of net taxes to an output shock (a) and output to a net tax shock (b) as functions of time. Obtained using the baseline identification strategy.

E Model setup

The overall methodology used in this paper follows the setup of Eisenstat et al. (2016). The model in 2.1-2.3 can be transformed into:

$$Y_t = X_t\alpha + X_t\Phi\tilde{\Omega}^{\frac{1}{2}}\gamma_t + \Sigma_t u_t, \qquad u_t \sim \mathcal{N}(0, I) \tag{E.1}$$

$$\gamma_t = \gamma_{t-1} + v_t^* \qquad v_t^* \sim \mathcal{N}(0, I) \tag{E.2}$$

$$log(\sigma_t) = log(\sigma_{t-1}) + \theta_t \qquad \theta_t \sim \mathcal{N}(0, W) \tag{E.3}$$

where α contains coefficients from a time-invariant version of the VAR, $\tilde{\Omega}^{\frac{1}{2}}$ and Φ are obtained from a factorization of the variance covariance matrix Ω and $\gamma_{j,t} = (\beta_{j,t} - \alpha_j)/\omega_j$ for $j = 1, \ldots, m$. The above model can be broken down into N separate state space representation models and estimated recursively using the Gibb's sampler. γ_t and σ_t are estimated via following state-space representation models:

Model 1:

$$\begin{aligned} \tilde{Y}_t &= W_t \gamma_t + \varepsilon_t \\ \gamma_t &= \gamma_{t-1} + v_t^* \end{aligned} \tag{E.4}$$

Model 2:

$$\begin{aligned} \varepsilon_t^{**} &= 2 \times log(\sigma_t) + log(u_t u_t') \\ log(\sigma_t) &= log(\sigma_{t-1}) + \theta_t \end{aligned} \tag{E.5}$$

where $\tilde{Y}_t = Y_t - X_t \alpha$ and $W_t = X_t \tilde{\Omega}^{\frac{1}{2}} \Phi$. Model 1 is a linear Gaussian state space representation model, thus, it can be solved using the Carter and Kohn (1994) approach. Model 3, on the other hand, is a linear but non-Gaussian state space representation model. It is solved via the Kim et al. (1998) approach, which uses the fact that $log(u_t u_t')$ has a χ_1^2 distribution, which can be approximated by the mixture of log-normals.
The variance covariance matrix W from the state equation is sampled from $\mathcal{IW}(\bar{W}^{-1}, \bar{T})$:

$$\begin{aligned} \bar{Q} &= \underline{\mathrm{W}} + \Sigma_{t=1}^{T} \theta_t \theta_t' \\ \bar{T} &= \underline{\mathrm{T}} + T \end{aligned}$$

α and Φ are estimated using simple linear regression techniques. In case of Φ, equation E.1 is further rearranged into:

$$Y_t^* = Z_t \phi + e_t \tag{E.6}$$

where ϕ contains all the non-zero off-diagonal elements of Φ, $Y_t^* = Y_t - X_t \alpha - X_t \tilde{\Omega}^{\frac{1}{2}} \gamma_t$, $Z_t = X_t \tilde{\Omega}^{\frac{1}{2}} F_t$ and F_t is defined as:

$$F_t = \begin{pmatrix} 0 & 0 & \dots & 0 \\ \gamma_{1,t}' & 0 & \dots & 0 \\ 0 & \gamma_{[1,\dots,2],t}' & & \vdots \\ \vdots & & \ddots & 0 \\ 0 & \dots & 0 & \gamma_{[1,\dots,m-1],t}' \end{pmatrix}$$

ω is obtained by defining:

$$Y = \begin{pmatrix} Y_1 \\ \vdots \\ Y_T \end{pmatrix}, \quad X = \begin{pmatrix} X_1 \\ \vdots \\ X_T \end{pmatrix}, \quad \gamma = \begin{pmatrix} \gamma_1 \\ \vdots \\ \gamma_T \end{pmatrix}, \quad \varepsilon = \begin{pmatrix} \varepsilon_1 \\ \vdots \\ \varepsilon_T \end{pmatrix}$$

and re-arranging (12) into:

$$v_j = g_j \omega_j + \varepsilon \tag{E.7}$$

where $v_j = Y - X\alpha - G_{\setminus j}\omega_{\setminus j}$, g_j denotes the jth column of G, $G_{\setminus j}$ denotes a $Tn \times (m-1)$ matrix obtained by deleting the jth column from G, $\omega_{\setminus j}$ is ω with the jth row removed and G is $G_t = X_t diag(\gamma_t)$ stacked in a similar way as above:

$$G = \begin{pmatrix} G_1 \\ \vdots \\ G_T \end{pmatrix}$$

ω is then sampled from a conditional posterior density that follows a 2-component mixture of truncated normals:

$$p(\omega_j | Y, \alpha, \gamma, \omega_{\setminus j}, \Sigma, \tau, \lambda) = \hat{\pi}_j \phi_{(-\infty, o)}(\omega_j | \mu_j, \tau_j^2) + (1 - \hat{\pi}_j)\phi_{(0,\infty)}(\omega_j | \hat{\mu}_j, \hat{\tau}_j^2)$$

τ_j^2 and λ are sampled in the same manner as in Belmonte, Koop and Korobolis (2014):

$$(\tau_j^{-2} | \lambda, \omega_j) \sim IG\left(\sqrt{\frac{\lambda^2}{(\omega_j - \mu_j)^2}}, \lambda^2\right)$$

$$(\lambda^2 | \tau) \sim G\left(\lambda_{01} + m, \lambda_{02} + \frac{1}{2}\Sigma_{j=1}^m \tau_j^2\right)$$

For further peculiarities, please see Eisenstat et al. (2016) and Frühwirth-Schnatter and Wagner (2010) and Belmonte et al. (2014).

The Gibb's sampler procedure takes the following from:

1. Draw α from $p(\alpha | Y^T, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;
2. Draw γ^T from $p(\gamma^T | Y^T, \alpha, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;
3. Draw Σ^T from $p(\Sigma^T | Y^T, \alpha, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;
4. Draw W from $p(W | Y^T, \alpha, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;
5. Draw ω from $p(\omega | Y^T, \alpha, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;

6. Draw τ from $p(\tau|Y^T, \alpha, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;
7. Draw λ from $p(\omega|Y^T, \alpha, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;
8. Draw Φ from $p(\Phi|Y^T, \alpha, \gamma^T, \Sigma^T, W, \omega, \tau, \lambda, \Phi)$;

The methodology in use requires specification of prior distributions. Standard independent priors are assumed for:

$$\alpha_0 \sim \mathcal{N}(0, I_m)$$
$$\beta_0 \sim \mathcal{N}(0, I_m)$$
$$\Sigma_0 \sim \mathcal{N}(0, I_n)$$

A Tobit prior for ω:

$$\omega_j^* \sim \mathcal{N}(0, \tau_j^2)$$
$$\omega_j = \begin{cases} 0 & \text{if } \omega_j^* \leq 0 \\ \omega_j^* & \text{if } \omega_j^* > 0 \end{cases}$$

A Lasso prior for tau_j^2:

$$\tau_j^2 \sim \mathcal{E}\left(\frac{\lambda^2}{2}\right)$$
$$\lambda^2 \sim \mathcal{G}(0.1, 0.1)$$

$$W \sim \mathcal{IW}(n + 11, 0.01^2(n + 11 - n - 1)I_n)$$

PUBLICATIONS

Cyclical Fiscal Multipliers: Policy Mix and Financial Friction Puzzle

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