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NEW EVIDENCE ON STATE-DEPENDENT FISCAL MULTIPLIERS

by Fabrizio Renzi*

Abstract

Building on the dataset and identification strategy of Ramey and Zubairy (2018), this paper investigates the dynamic effects of fiscal shocks using a state-dependent local projection approach. By interacting fiscal shocks with a continuous state variable, the analysis captures how fiscal multipliers evolve over the business cycle. This approach is novel within this literature, as most previous studies rely on two-regime estimates to account for non-linearity. The results indicate that the government spending multiplier varies significantly, ranging from near zero during expansions to slightly below one in downturns. While this evidence highlights meaningful state dependence in the transmission of fiscal policy, the relatively modest size of the estimated multipliers calls for caution in drawing strong conclusions about the effectiveness of countercyclical policies.

JEL Classification: E62, E32.

Keywords: fiscal multipliers, government spending, state-dependent local projections.

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1 Introduction[1]

The size of fiscal multipliers is a central issue in empirical macroeconomics, with notable growth in interest from policymakers and academics over the past two decades. Questions about the size of fiscal multipliers were brought into sharp focus by the two major economic crises of the 2000s, the Global Financial Crisis of 2008 and the COVID-19 pandemic, both of which required large-scale government intervention through fiscal stimulus to stabilize output, support employment, and prevent deeper recessions. Furthermore, as interest rates reached the zero lower bound, monetary policy became less effective in stimulating demand, prompting policymakers to rely more heavily on fiscal measures to support recovery.

A key question in the literature is whether fiscal multipliers vary depending on the state of the economy, in line with the traditional Keynesian argument that government spending is more effective during economic downturns. This idea has been extensively explored in recent theoretical work, which highlights several mechanisms capable of generating state-dependent multipliers. Notably, a number of contributions emphasize that labor market frictions can make fiscal policy particularly effective in periods of slack[2]. Another mechanism operates through financial frictions, as tighter financial conditions in recessions can amplify the effects of government spending, leading to larger multipliers during downturns[3].

Despite these theoretical insights, the empirical literature − largely based on comparing multipliers prevailing in good and bad times − has produced mixed results. In particular, two really influential papers report contrasting findings: while Auerbach and Gorodnichenko (2012) find fiscal multipliers remarkably higher in recessions, Ramey and Zubairy (2018) conclude that there is virtually no difference between good and bad times. In addition, a recent methodological work by Gonçalves et al. (2024b) has pointed to potential limitations in recent empirical estimates, showing that standard recession–expansion LP frameworks can produce inconsistent results when the state variable is endogenous. Using a non-parametric alternative, they obtain substantially larger multipliers in downturns when replicating the Ramey and Zubairy (2018) estimates.

Against this background, this paper contributes to the debate by reassessing the results in Ramey and Zubairy (2018) using the more recent approach proposed by Cloyne et al.

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[2] See, among others, Michaillat (2014), Shen and Yang (2018), Albertini et al. (2021) for models in which idle labor, downward nominal wage rigidity, or unemployment-risk channels amplify the effects of government spending during downturns.

[3] See, for example, Canzoneri et al. (2016).

(2023), which models state dependence through a continuous state variable rather than imposing discrete regime splits between "good" and "bad" times. By avoiding arbitrary sample partitions, the continuous-state approach is, by construction, more robust and mitigates concerns related to threshold selection, loss of information, and potential misclassifications of economic states. Moreover, this framework allows the impact of fiscal shocks to be traced smoothly across the entire spectrum of economic conditions — from deep recessions to strong expansions — thereby delivering a more nuanced and data-driven characterization of how fiscal multipliers vary. In doing so, it addresses the econometric issues that may have affected earlier estimates and provides a richer understanding of state-dependent fiscal policy effects. While this methodology is novel within this strand of literature, recent applications have appeared in studies examining non-linearities in the transmission of monetary policy shocks (Alessandri et al. (2025); Caramp and Feilich (2024)).

The results support the view that the government spending multiplier is higher during downturns, ranging from 0 to values slightly below 1 depending on the state of the economy. Notably, the variation becomes more pronounced as unemployment declines below its median level, indicating that fiscal policy is markedly less effective when the economy operates closer to full capacity.

The remainder of the paper is organized as follows. Section 2 provides a brief review of the related literature, with a particular focus on key contributions and methodological papers that are especially relevant to the empirical strategy adopted in this study. Section 3 outlines the methodology, while Section 4 presents the main results. Section 5 reports a series of robustness checks. Finally, Section 6 concludes.

2 Literature review

The literature on state-dependent fiscal multipliers is extensive. This section focuses on a few key contributions relevant for the subsequent analysis, beginning with studies on the main strategies for identifying fiscal policy shocks, followed by selected works that examine how the effects of government spending vary over the business cycle and concluding with recent methodological papers on state-dependent local projection methods.

The identification of fiscal shocks.

The identification of fiscal shocks is one of the main challenges in estimating the size of fiscal multipliers. This is primarily due to endogeneity concerns, as fiscal policy aggregates are closely tied to the business cycle developments. To overcome this issue, several strategies have been proposed in the literature.

A well-known identification approach was introduced in the seminal paper by Blanchard and Perotti (2002). This strategy relies on the assumption that government spending responds with a lag to contemporaneous macroeconomic variables. Under this assumption, fiscal shocks can be recovered through a Cholesky decomposition, ordering government spending first. However, a common critique of this method is that the shocks may be anticipated by economic agents, undermining the assumption of exogeneity (Leeper et al. (2013); Ramey (2016)). Another SVAR-based identification strategy is proposed by Mountford and Uhlig (2009), who employ sign restrictions to identify fiscal shocks while controlling for generic business cycle and monetary policy shocks.

An alternative approach relies on fiscal forecast errors, e.g. using data from the Survey of Professional Forecasters (Ramey (2011)); under this strategy, forecast errors are interpreted as unanticipated changes in fiscal variables and treated as exogenous shocks.

Finally, a now widely used strategy is the narrative or external instrument approach, which uses external information as a proxy for the exogenous variation in government spending or taxes. A prominent example is the series of US defense news shocks introduced by Ramey and Shapiro (1998) and extended in Ramey (2011) and in Ramey and Zubairy (2018). The authors produced an estimate of changes in the expected present value of government defense spending, based on news sources. The underlying assumption is that decisions about future increases in spending are driven by wars or geopolitical developments, and therefore represent shocks that are exogenous to the current state of the economy. These news shocks have been shown to be relevant external instruments for identifying exogenous movements in government spending.

Fiscal multipliers in good and bad times.

As emphasized by Ramey (2019), linear estimates of the spending multiplier - which do not account for any sort of state dependence - are broadly consistent across various methodologies and samples. In particular, once a common definition is adopted − specifically, the cumulative change in output divided by the cumulative change in government spending − estimates for multipliers on general government purchases tend to cluster within a narrow range between 0.6 and 1[4]. In contrast, no such consensus exists when it comes to nonlinear estimates, that assess whether fiscal multipliers are different depending on the cyclical position of the economy.

A very influential paper by Auerbach and Gorodnichenko (2012) find higher multipliers

[4]Some results in the literature are based on the definition of the fiscal multiplier adopted by Blanchard and Perotti (2002), in which the multiplier is calculated by comparing the peak output response to the initial increase in government spending. Another potential source of disagreement among estimates stems from the procedure used to convert estimated elasticities into fiscal multipliers (see footnote 11).

during recessions. Their estimates are based on a smooth-transition VAR model, in which each observation is modeled as a weighted average of two regimes: one prevailing during deep recessions and the other during strong expansions; the transition between regimes is driven by a logistic function, depending on GDP growth rate. The estimates are based on U.S. quarterly data spanning from 1947 to 2008; fiscal shocks are identified using the Blanchard and Perotti approach, while mitigating anticipation effects by incorporating professional forecast data into the SVAR. However, later contributions has raised concerns about the robustness of these findings. As pointed out by Ramey and Zubairy (2018), this result is mainly driven by medium term response of GDP, which keeps rising indefinitely after a spending shock during a recession, while government spending does not keep rising. This pattern arises from the assumption embedded in their model, that the economy stays in its current state after the shock. Consequently, GDP response in bad states is overestimated, since the counterfactual implies that the economy stays in recession indefinitely, whereas in reality recessions typically last only 3.3 quarters[5]. Furthermore, Alloza (2022) highlighted that the results are sensitive to the somewhat arbitrary definition of the variable that determines the transition between good and bad states[6].

Weaker evidence for nonlinearity is found in Caggiano et al. (2015). The authors employ a STVAR approach close to that of Auerbach and Gorodnichenko (2012), but use a Generalized Impulse Response Function, which allow to endogenize the transition between regimes following the shock, and explicitly consider the role of fiscal foresight by including a measure of fiscal news[7]. Their results show that differences in spending multipliers arise only when focusing on "extreme" events, i.e. deep recessions vs. strong expansionary periods.

A key contribution to this debate is Ramey and Zubairy (2018). In this paper, the authors construct a comprehensive dataset covering the period from 1889 to 2015, and identify fiscal shocks using an external instrument − namely, the news about future military spending, as previously discussed. The estimates are obtained using a non-linear version of Jordà's Local

[5]This issue is naturally addressed in LPs (Jordà (2005)). Since they are essentially direct forecasting methods, the estimated impulse responses inherently reflect how the average shock is likely to change the state, while natural transitions between states that are independent of the shock should be captured by the state-dependent control variables. Indeed, Ramey and Zubairy (2018) replicate the analysis by applying the Jordà method to the same dataset, sample period, and identification approach, finding no evidence of higher multipliers during recessions.

[6]In Auerbach and Gorodnichenko (2012), the transition between expansion and recession regimes is governed by a logistic function that depends on z_t, defined as a centered moving average of order 7 of the GDP growth rate. Alloza (2022) shows that the results of the estimation are highly sensitive to the choice of centering and to the ordering of the moving average.

[7]This variable is defined as the sum of revisions of expectations about future government spending collected by the Survey of Professional (which collects forecasts of variables up to time t+3). The estimates are based on U.S. data spanning the period 1981:Q3-2013:Q1, with 1981:Q3 being the first available quarter to construct the news variable.

Projections method, in which two distinct regimes are defined according to a dummy variable that takes the value of one when a selected state variable exceeds a specified threshold. In the benchmark specification, which distinguishes between periods of high and low unemployment, the authors find that the government spending multiplier is broadly similar across regimes — around 0.7 at a four-year horizon. Overall, the results provide no evidence of significantly higher multipliers during periods of economic slack.

This strand of literature has faced a recurring econometric challenge: time series data typically include only a limited number of recession episodes. For this reason, the contribution by Ramey and Zubairy (2018) is particularly valuable. By employing a long historical dataset, the study incorporates a greater number of slack periods and captures much larger variations in government spending, whereas earlier contributions typically rely on samples starting in the mid-1940s[8]. A different strategy is adopted by Mumtaz and Sunder-Plassmann (2021), who use U.S. state-level data and employs a mixed-frequency panel VAR model which allows them to incorporate spending data from 1950:Q1. The model features two regimes for each U.S. state, with regime switches determined by a state-specific threshold process. Fiscal shocks are identified using the Blanchard-Perotti approach. This long time span, combined with the cross-sectional dimension of the panel, enables estimates in recessions and expansions to be based on substantially more observations than those in a time series model. The authors find that fiscal multipliers are larger during recessions. Moreover, this panel-based approach reveals substantial heterogeneity across states, with the degree of nonlinearity in the effects of spending shocks being more pronounced in states experiencing higher levels of financial frictions.

Non-linear Local projections.

Regarding the methodology, most recent applied macroeconomic studies, including those focused on non-linear estimates of fiscal multipliers, commonly employ the Local Projections (LP) approach introduced by Jordà (2005). The widespread adoption of this approach likely stems from its straightforward implementation, especially in non-linear settings, as the estimates are simply obtained via a set of OLS/IV regressions. However, beyond its practicality, LPs also offer notable econometric properties that can make them preferable to traditional Vector Autoregressions (VARs) in certain applications.

As shown by Plagborg-Møller and Wolf (2021) and Montiel Olea and Plagborg-Møller (2021), LPs and VARs estimate the same impulse responses in the population under certain conditions, and structural identification schemes commonly used in VARs can be equivalently

[8] Furthermore, the authors show that the news on defense spendins serves as a relevant instrument only when considering the whole dataset.

implemented within the LP framework. However, in finite samples, LPs and VARs fall on opposite ends of bias-variance spectrum. VARs impose stronger parametric restrictions, making them more efficient but also more prone to misspecification. In contrast, LPs offer greater flexibility but can suffer from higher sampling variability. In practice, let p denote the maximum lag length and h the horizon at which the impulse response is evaluated. Local Projections (LP) and Vector Autoregressions (VAR) produce approximately equivalent estimates up to $h = p$, the main difference between the two methods emerges at longer horizons. LPs are generally preferable for estimating impulse responses at extended horizons because they employ a direct forecasting approach: each horizon is estimated independently using the relevant data, thereby avoiding reliance on model-based iteration. In contrast, a VAR(p) uses the first p autocovariances to estimate the entire system and generates responses for $h > p$ by iterating on the estimated model dynamics. This iterative procedure can amplify model misspecification over time, leading to biased estimates at longer horizons (Jordà et al., 2024).[9].

When it comes to non-linear estimates, a number of recent contributions have raised concerns about the use of Local Projections (LPs). Gonçalves et al. (2021) examine two prominent examples of economically meaningful nonlinear transformations $f(\cdot)$ of the shock variable ϵ_t. The first is a censored variable, defined as $\epsilon_t^+ \equiv \max(0, \epsilon_t)$, which captures possible asymmetries in the response to positive versus negative shocks. The second involves polynomial transformations, such as including both ϵ_t and ϵ_t^3 in the regression. This allows the size of the shock—regardless of sign—to have a nonlinear impact on the outcome variable, with larger shocks producing stronger effects. The authors show that the conventional LP approach fails to recover the population IRFs in the presence of such nonlinearities and propose a new plug-in estimator that delivers consistent estimates.

Gonçalves et al. (2024b) critically examines commonly used state-dependent LPs that rely on a dichotomous state variable to distinguish between different regimes, such as recessions and expansions. The authors show that when the state variable is endogenous, state-dependent LPs can yield inconsistent estimates of the IRFs when the size of the shock ϵ_t is large. To address this issue, they propose a non-parametric estimator that yields consistent estimates regardless of the endogeneity of the state variable. Using this method to replicate the analysis in Ramey and Zubairy (2018), they find substantially different results: fiscal multipliers are significantly larger during economic downturns. These findings further highlight the uncertainty surrounding state-dependent fiscal multiplier estimates.

[9]The authors show that LPs are robust to lag truncation - even when $h > p$ - as this has asymptotically negligible effects on estimator consistency. In contrast, impulse responses derived from truncated VARs depend on the full dynamic structure of the model. As a result, small-sample inconsistencies can accumulate over time, leading to potentially significant biases, particularly at longer horizons.

A new methodological framework to account for state-dependence is proposed by Cloyne et al. (2023). Building on the Kitagawa–Oaxaca–Blinder decomposition, this approach allows the overall impact of an exogenous shock to be separated into direct and indirect effects. The direct effects capture the average response to the shock, comparable to estimates from a standard linear model, while the indirect effects measure how the response is influenced by a set of covariates. Empirically, this framework translates into a straightforward extension of linear LPs, in which interaction terms between the shock and selected covariates (that potentially affect the shock's transmission) are included to estimate the indirect effects. In their application, the authors primarily use this framework to examine the interaction between fiscal and monetary policy, finding that the fiscal multiplier is larger when monetary policy is more accommodative, a result consistent with standard New Keynesian models. This methodology provides a more nuanced understanding of how the effect of an exogenous shock vary across different economic conditions. A key strength of this approach is that the state variable is a continuous variable, thereby avoiding the need to divide the sample into discrete regimes. This results in a more flexible and realistic alternative to traditional non-linear models. Given its flexibility and ability to assess non-linearities, this framework provides an ideal tool for revisiting and reassessing earlier findings in the literature, particularly in light of the significant differences observed in recent studies and of the concerns about the reliability of two-regime LPs.

3 Data and methodology

To quantify the dynamic effects of fiscal shocks across different states of the economy, I employ a state-dependent local projection approach based on the framework introduced by Cloyne et al. (2023). In particular, the impulse response functions (IRFs) are derived through the following sequence of non-linear Local Projections (LPs):

$$y_{t+h} = \alpha_h + \beta_h \epsilon_t + (\lambda_h + \gamma_h \epsilon_t) S_{t-1} + \sum_{j=1}^{p} \phi_j x_{t-j} + u_{t+h}, h = 0, 1, ..., H \tag{1}$$

where y_{t+h} denotes either GDP or Government spending h period ahead, following an exogenous fiscal shock ϵ_t; S_{t-1} is a state variable that may influence the transmission of the fiscal shock; x_{t-j} are a set of lagged controls and u_{t+h} is a residual term.

The estimates are based on the fiscal shocks and the dataset constructed by Ramey and Zubairy (2018), which spans quarterly U.S. data from 1889 to 2015. I also kept the same control variables and applied the same transformation in order to highlight the role played by the new methodological framework hereby implemented. More in detail: the fiscal shocks are obtained from news on defense spending[10]; the control variables in the model include four lags of real GDP, government spending and the shock. All NIPA variables are divided by an estimate of potential GDP, according to the Gordon-Krenn transformation [11]. In the benchmark specification, the state variable S_{t-1} is defined as the unemployment gap, computed as the deviation of the observed unemployment rate from its trend component estimated using a Hodrick-Prescott (HP) filter[12]. The state variable is therefore continuous, and it is standardized to have a mean of zero and a standard deviation of one in order to facilitate the interpretation of the coefficients (see section 3.1).

The interaction term between the fiscal shock and the state variable is constructed by lagging the latter, such that the state variable measures the level of economic activity prevailing one quarter prior to the shock. From a practical standpoint, this timing ensures that the state variable reflects the initial macroeconomic conditions before the shock hits and aligns with the government's budgeting and implementation process, which typically relies on data from the previous quarter. From an econometric perspective, using a lagged variable helps mitigate concerns about endogeneity: as long as it is reasonable to assume that the shock ϵ_t is fully exogenous and not anticipated by the economic agents, S_{t-1} can be considered orthogonal to ϵ_t. This approach has been recently employed in Caramp and Feilich (2024), who used a specification close to (1) to assess how predetermined values of U.S. public debt affect the transmission of monetary policy shocks. The econometric strategy also appears robust to the concerns raised by Gonçalves et al. (2024a), which highlights that in a model with a nonlinear interaction term the IRFs are potentially biased when S_t is a nonlinear function of the shock ϵ_t. However, in the present model there are no strong reasons to expect that the contemporaneous response of unemployment to a fiscal shock might be

[10] The fiscal shocks are identified using a narrative approach based on news about anticipated future changes in military spending. Since such news typically stems from geopolitical events or emerging conflicts, it is orthogonal to the current state of the business cycle and thus provides an exogenous source of changes in government spending dynamics. Additionally, through a statistical analysis, the authors also showed that these shocks serve as a relevant instrument for government spending, strengthening their validity in empirical analysis.

[11] As discussed in Ramey (2019), dividing both GDP and government spending by potential GDP places the variables in the same units, allowing the fiscal multiplier to be estimated directly. In contrast, when the variables are expressed in logarithmic form, the estimated coefficients correspond to elasticities, which must be converted into level effects by multiplying them by the average ratio of GDP to government spending over the sample period. However, this practice can introduce substantial bias, as the ratio Y/G may vary considerably over time.

[12] In the benchmark specification of Ramey and Zubairy (2018), the state variable S_t is based on unemployment, which is used to distinguish between two regimes: one where unemployment is above a fixed threshold (6.5), one where it is below. This approach allows for the estimation of fiscal multipliers during both good and bad times using standard non-linear LPs. In a robustness check, the threshold is defined as an HP filter, obtained by fitting the HP filter over a split sample, 1889–1929 and 1947–2015, and then linearly interpolating the gap. This estimate of the HP filter is considered in this paper to compute the unemployment gap.

nonlinear, and a brief data analysis supports this assumption[13].

3.1 State-dependent IRFs

In a standard two-regime approach, as the one implemented in Ramey and Zubairy (2018), the parameters are allowed to vary between regimes, directly providing a different response in each state of the economy. Conversely, this model differs in that the coefficients do not depend on the state, but the impulse response does.

More in detail, in the model described in (1) the state-dependence is captured by the interaction term $\gamma_h \epsilon_t S_{t-1}$, which assesses how the effects of the fiscal intervention varies according to the state of the economy. In this setting, the overall effect of a fiscal shock can be decomposed into two components: β_h measures the direct effects, which are comparable to the average response obtained in a linear estimate; the coefficient γ_h captures the indirect effect arising from the interaction between the fiscal shock and the state variable S_{t-1}. Indeed, the state-dependent IRFs can be expressed as follows:

$$IRF_h(S_{t-1}) = \frac{\partial y_{t+h}}{\partial \epsilon_t} = \beta_h + \gamma_h S_{t-1} \tag{2}$$

In other terms, the response to a fiscal shock is allowed to be heterogeneous, with the heterogeneity measured through an observable state variable S_{t-1}. As the latter is a continuous variable, the IRF_h can theoretically be computed for any value of S_{t-1}, allowing to assess the responses to fiscal shocks across different values of the state variable, which correspond to various phases of the business cycle. Furthermore, as S_{t-1} is standardized: when the state variable stays at its average level ($S_{t-1} = 0$), the response to a shock is simply given by β_h, which therefore measures the average response as in a linear model. The second term captures the heterogeneity around the average, specifically the additional (indirect) response that operates through S_{t-1}, becoming larger the more the state variable deviates from its mean; particularly, γ_h measures the additional effect when the state variable is one standard deviation greater than the mean (namely, $S_{t-1} = 1$). In a linear model the coefficient γ_h is implicitly assumed to be zero; hence, not including the interaction term when a

[13] More in details, Gonçalves et al. (2024a) consider a data generating process where $y_t = \beta_{21}x_t + \beta_{23}r_{t-1} + \alpha_{21}x_t \times r_t + \gamma_{21}y_{t-1} + \epsilon_{2t}$, with $r_t = f(x_{t-1}) + \epsilon_{3t}$ and x_t being an exogenous shock. Intuitively, if $f(x_t)$ is non-linear, a shock in t affects the state variable in the next period, inducing a nonlinear relationship between y_{t+1} and x_t. Therefore, the LP estimator will fail to recover the true impulse response for $h > 0$, with the bias depending on the size of the coefficient β_{23} and the function $f()$. Conversely, when $f()$ is linear, the true IRF can be recovered from the usual LP. It can be shown that the same holds if the state variable is lagged, but react contemporaneously to a shock, namely if $y_t = \beta_{21}x_t + \beta_{23}r_{t-1} + \alpha_{21}x_t \times r_{t-1} + \gamma_{21}y_{t-1} + \epsilon_{2t}$, with $r_t = f(x_t) + \epsilon_{3t}$, which is closer to the setup considered here. A more detailed discussion of these considerations is provided in the Appendix A.3.

state-dependence is statistically significant might result in an omitted variable bias.

3.2 State-dependent fiscal multipliers

In line with previous work, I consider the cumulative spending multiplier, defined as the ratio between the cumulative impulse response of real GDP and that of government spending, computed in response to the same shock and over the same horizon. In this setting, estimates of the fiscal multiplier at horizon k can be obtained as the ratio of the cumulative response obtained through equation (2), for $h = 0, 1, ..., k$. This approach represents a state-dependent version of the "3-step" approach proposed by Ramey and Zubairy (2018)[14]:

$$m_k^{3s}(S_{t-1}) = \frac{\sum_{h=0}^{k} \beta_h^y + \sum_{h=0}^{k} \gamma_h^y S_{t-1}}{\sum_{h=0}^{k} \beta_h^g + \sum_{h=0}^{k} \gamma_h^g S_{t-1}} \tag{3}$$

where the multiplier is a non-linear combination of the coefficients β_h and γ_h, and of the state variable S_{t-1} which affects both the numerator and the denominator. Using different input values for S_{t-1} (e.g., using different percentiles), it is possible to assess how the fiscal multiplier varies according to the state of the economy.

The "3-step" approach offers a straightforward and flexible way to obtain point estimates of the fiscal multiplier. However, it is not ideal for drawing inference conclusions. At medium horizon, equation (3) depends on a large number of parameters, each contributing its own uncertainty, which might therefore results in wide confidence intervals.

A more efficient alternative is provided by the 1-step approach, extended to account for state dependence:

$$Y_{t+h} = \alpha_h + m_h G_{t+h} + \Gamma_h G_{t+h} S_{t-1} + \sum_{j=1}^{p} \phi_j x_{t-j} + u_{t+h}, h = 0, 1, ..., H \tag{4}$$

where $Y_{t+h} = \sum_{j=0}^{h} y_{t+j}$ and $G_{t+h} = \sum_{j=0}^{h} g_{t+j}$, namely the cumulative value of output and government spending; $G_{t+h}S_{t-1}$ is a cumulative interaction term that is meant to capture how the state of the economy affects the size of the fiscal multiplier.

To address the endogeneity of G_{t+h} and retrieve its exogenous variation, the variable is

[14] The term "3-step" refers to the procedure used to compute the multiplier, which involves three stages: first, estimating the cumulative response of GDP; second, estimating the cumulative response of government spending; third, taking the ratio of the two.

replaced with the fitted-values $\hat{G}_{t+h}$ obtained from the first stage regression[15]:

$$G_{t+h} = \alpha_h + \phi_h \epsilon_t + \sum_{j=1}^{p} \phi_j x_{t-j} + u_{t+h}, h = 0, 1, ..., H \tag{5}$$

The state-dependent 1-step fiscal multiplier is then given by the following:

$$m_h^{1s}(S_{t-1}) = \frac{\partial Y_{t+h}}{\partial \hat{G}_{t+h}} = m_h + \Gamma_h S_{t-1} \tag{6}$$

where m_h is a direct estimate of the average multiplier and Γ_h measures how the multiplier varies according to the state of economy[16]. More precisely, as already pointed out for equation (2), as S_{t-1} is standardized, Γ_h measures the additional effect when the state variable is one standard deviation greater than the mean (namely, $S_{t-1} = 1$).

In the context of the two-regime LPs employed by Ramey and Zubairy (2018), the 1-step approach is generally preferred, as it yields the same point estimates with more reliable inference. However, in the current setting the two approaches do not coincide. Unlike the "3-step" approach described in equation (3), the "1-step" approach imposes a linear relationship between the state variable and the fiscal multiplier. While this assumption simplifies estimation and facilitates inference, it may overlook more complex nonlinear interactions between the state of the economy and the effects of fiscal policy. By contrast, the 3-step approach offers greater flexibility, allowing for a richer and potentially more realistic depiction of state dependence. Taken together, these two approaches complement each other and can be used to check the internal consistency of the estimates.

4 Results

I begin the analysis by examining the dynamic responses of GDP and government spending to a fiscal shock, as modelled by equation 2. The two sets of coefficients of primary interest are β_h, which measure the direct effects, and γ_h, which capture the state-dependence of the response.

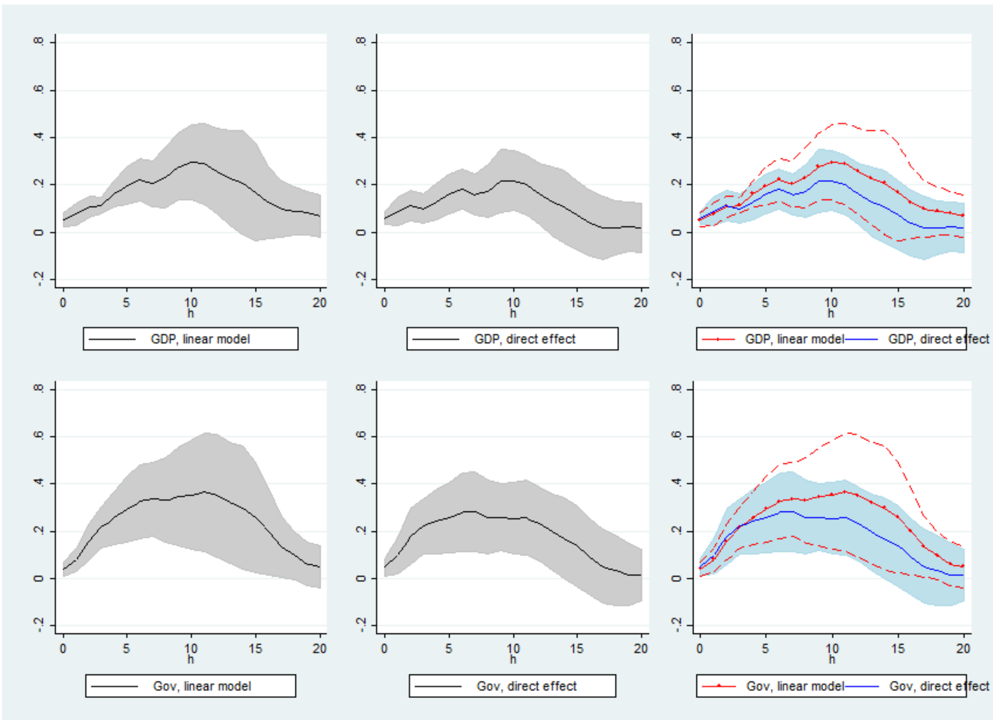
As previously discussed, the direct effects should, in principle, be consistent with those estimated in a linear model. As shown in figure 1, the estimates exhibit broadly similar patterns to those from the linear specification, though notable differences emerge. First, the average estimates differ, likely due to omitted variable bias in the linear model stemming from

[15] A similar strategy is employed in Alessandri et al. (2025)

[16] To take into account the first stage variability, confidence bands for m_h, Γ_h, m_h^{1s} are estimated via bootstrap.

the exclusion of the interaction term that accounts for state-dependent dynamics[17]. Second, the confidence bands in the state-dependent model are markedly narrower, suggesting that a significant share of the variability in the estimates is now explained by the interaction term introduced in this specification.

Figure 1: Direct effects (β_h) on GDP and Government spending



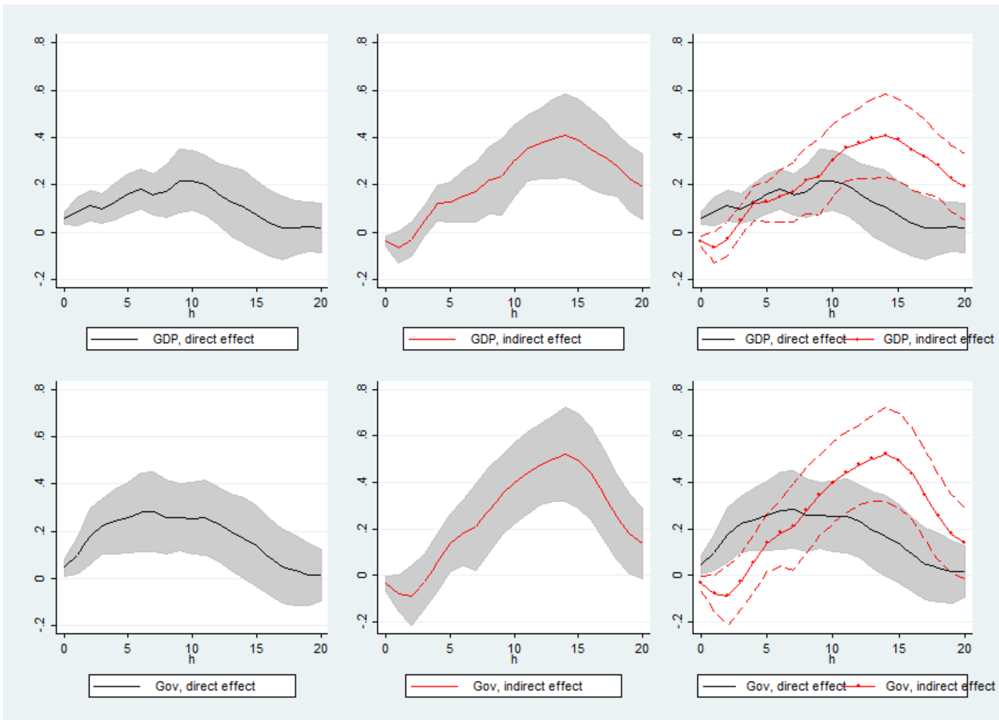
Note: The figure displays the estimates of β_h as defined in equation 1. The first row illustrates the impact on GDP, while the second row refers to government spending. The left column presents estimates from a linear model; the middle column shows the estimates obtained from equation 1; the right column plots both sets of estimates together for comparison.

The coefficients γ_h, which capture the indirect effects, are central to this analysis, as they reflect the state-dependent nature of the response. Three key observations emerge from figure 2. First, indirect effects are statistically significant (for $h > 4$), suggesting that the interaction term plays a significant role in explaining the dynamic of the response as modeled in Equation 2. Second, the sign of the coefficients is positive: consistent with prior findings, both GDP and government spending exhibit stronger responses when unemployment is high. Lastly, the magnitude of γ_h is remarkable, increasing with the horizon h and exceeding in the medium term that of β_h. In contrast, the β_h coefficients are significant on impact but decay more rapidly towards zero. This implies that when unemployment is one standard deviation

[17]Minor differences also stem from the inclusion of an additional control, namely the lagged value of the state variable S_{t-1}.

above its trend value, the overall response is considerably higher in the medium run; in contrast, it is close to zero when unemployment is one standard deviation below trend. The increasing response over the horizon h likely reflects the nature of the instrument, which captures news about future government spending.

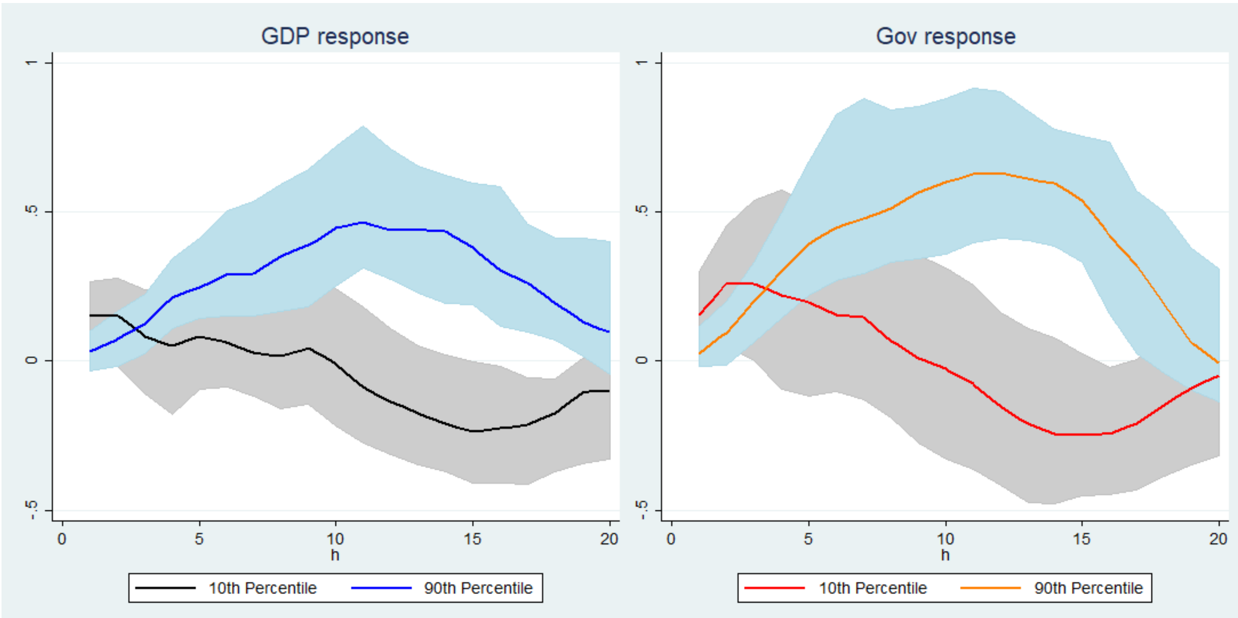
Figure 2: Direct (β_h) and Indirect effects (γ_h) on GDP and Government spending



Note: The figure presents the estimates of γ_h, as defined in equation 1. The first row illustrates the impact on GDP, while the second row refers to government spending. The first column reproduces the direct effects previously discussed, the second column displays the indirect effects and the third column overlays the two sets of estimates to facilitate direct comparison.

Using the estimates of β_h and γ_h it is possible to compute the state-dependent impulse response functions as defined in equation 2. The results are displayed in figure 3, which shows the IRFs for GDP and government spending, computed at the 10^{th} and 90^{th} percentiles of S_{t-1}, which can be considered a proxy for the state of the economy in good and bad times. The graphs clearly illustrate the presence of strong state dependence: when unemployment is low (i.e., S_{t-1} is at the 10^{th} percentile), the response of GDP to a government spending shock is close to zero, with its highest value on impact. In contrast, when unemployment is high (90^{th} percentile), the response becomes substantially more pronounced, especially in the medium term, when the indirect effects are more pronounced (see figure 2). A similar pattern is observed in the response of government spending, which also varies with the level of unemployment[18]. Nevertheless, what ultimately matters for the evaluation of fiscal policy is their relative response, captured by the fiscal multiplier.

Figure 3: State-dependent IRF - estimates at 10th and 90th percentile of S_{t-1}



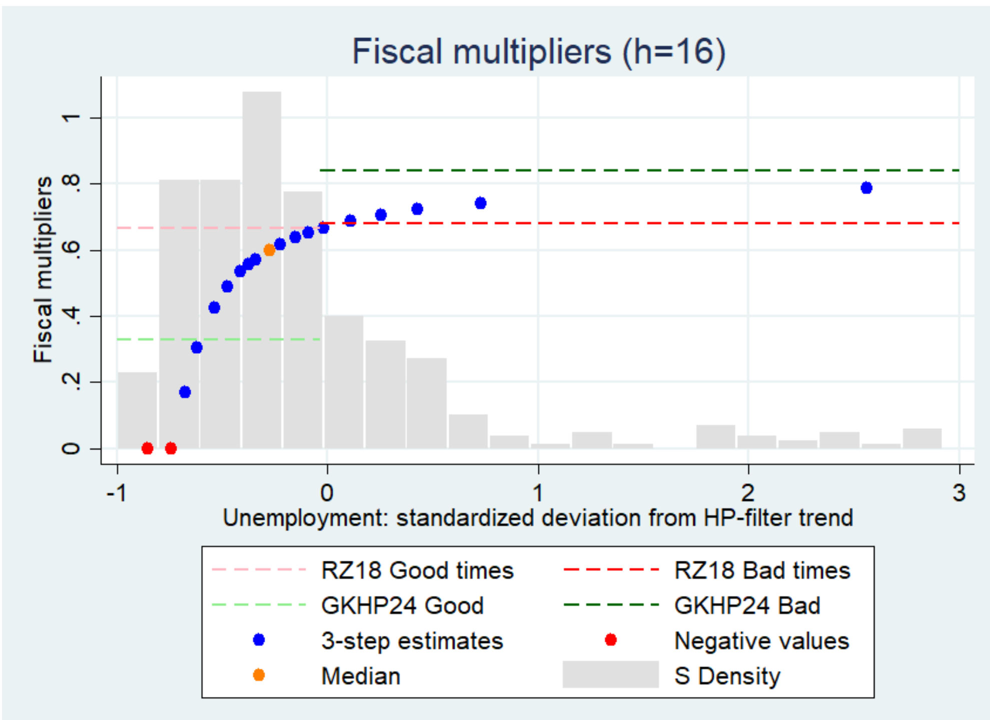
Note: Note: The figure presents estimates of the state-dependent IRFs based on equation 2. Confidence bands are defined by the 5$^{\text{th}}$ and 95$^{\text{th}}$ percentiles of the bootstrapped distribution of the IRFs.

The estimates of the fiscal multiplier obtained through the "3-step approach" (equation 3) are presented in figure 4. The figure plots the 4-year cumulative fiscal multiplier ($h = 16$) across values of unemployment ranging from the 5$^{\text{th}}$ to the 95$^{\text{th}}$ percentile and includes, for reference, previous estimates based on the same dataset and fiscal shocks, but obtained using two-regime LPs. In addition, the graph displays the empirical distribution of the state variable S_{t-1}, which is notably positively skewed: while unemployment can rise sharply during downturns, it cannot fall far below its median during expansions. The central result is straightforward: the fiscal multiplier varies markedly across the distribution of the state variable, ranging from values lower than zero to a maximum of approximately 0.8. This heterogeneity is most pronounced on the left-hand side of the distribution. Specifically, the multiplier declines steeply from its median value (approximately 0.6) to negative values when

[18] The stronger reaction of government spending in bad times —as in Ramey and Zubairy (2018)— can be understood from two complementary standpoints. From an econometric perspective, this may reflect differences in the correlation between the instrument and government spending across states, as the identified shocks can capture policy actions of slightly different intensity in recessions. From an economic perspective, the observed response may reflect spending components of a different nature that the instrument does not fully isolate—such as automatic stabilizers and additional discretionary measures that systematically expand during downturns.

unemployment deviations fall below the 10$^{\text{th}}$ percentile[19]. Conversely, as unemployment rises, the multiplier increases only moderately — even at very high levels of slack. Compared to previous estimates, this approach reveals greater heterogeneity than reported by Ramey and Zubairy (2018), who find no significant difference between multipliers across regimes of high and low unemployment. In contrast, the results are broadly consistent with the recent findings of Gonçalves et al. (2024b), who employ a non-parametric estimator to address endogeneity concerns and find multipliers significantly higher in bad times (0.8) compared to good times (0.3).

Figure 4: 3-step estimates of the fiscal multiplier



Note: The figure presents the estimates of the state-dependent fiscal multiplier, obtained using the "3-step approach" described in equation 3. Each dot represents the estimated multiplier at different levels of unemployment, ranging from the 5$^{\text{th}}$ to the 95$^{\text{th}}$ percentile. The median estimate is highlighted in orange, while negative values are shown in red. Dashed lines indicate the benchmark estimates reported in Ramey and Zubairy (2018) —labelled "RZ18"— and the more recent estimates by Gonçalves et al. (2024b) —labelled "GKHP24".

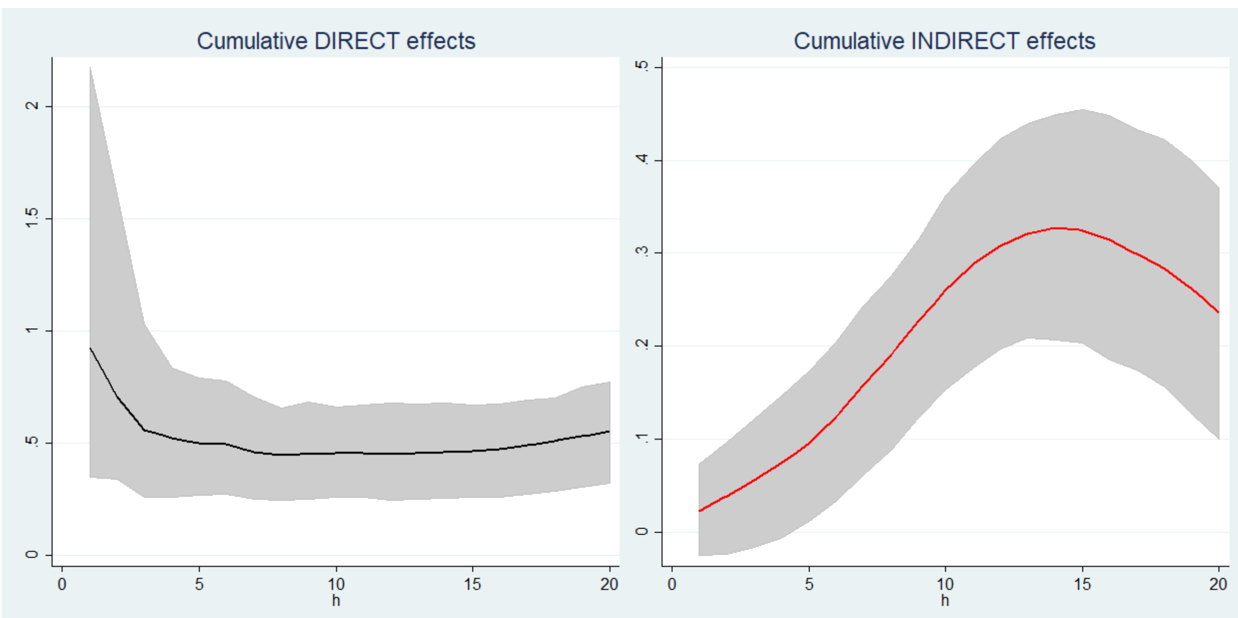
As previously noted, the "3-step" approach provides a highly flexible framework to assess state-dependence in fiscal multipliers. However, it is not ideally suited for drawing inferential conclusions, as the multiplier at $t+h$ is a nonlinear combination of $4h$ parameters obtained

[19] For these low values of unemployment, the numerator of the multiplier turns slightly negative, while the denominator remains positive but very close to zero. As a result, the ratio becomes mechanically strongly negative.

from $2h$ separate regressions. Furthermore, for low levels of unemployment, the fiscal multiplier is computed as the ratio of two values that are both close to zero; therefore, even minor estimation errors can be greatly amplified, leading to wide confidence bands. To evaluate — also in statistical terms — the differences in fiscal multipliers across economic conditions, alternative estimates can be derived using the "1-step" approach outlined in equations 4 and 6. This method allows for direct inference, as it estimates the fiscal multiplier in a single regression framework that explicitly accounts for the interaction between the cumulative variation of government spending and the state of the economy.

Estimates of cumulative direct and indirect effects are shown in figure 5. The cumulative direct effects are broadly in line with the linear estimates reported in Ramey and Zubairy (2018), with an impact value close to one and a medium-term value slightly above 0.5; the cumulative interaction term, which is meant to measure the state-dependence, remains statistically significant at all horizons. Notably, its magnitude is larger in the medium run, the typical horizon at which the fiscal multiplier is assessed.

Figure 5: Cumulative direct (m_h) and indirect (Γ_h) effects

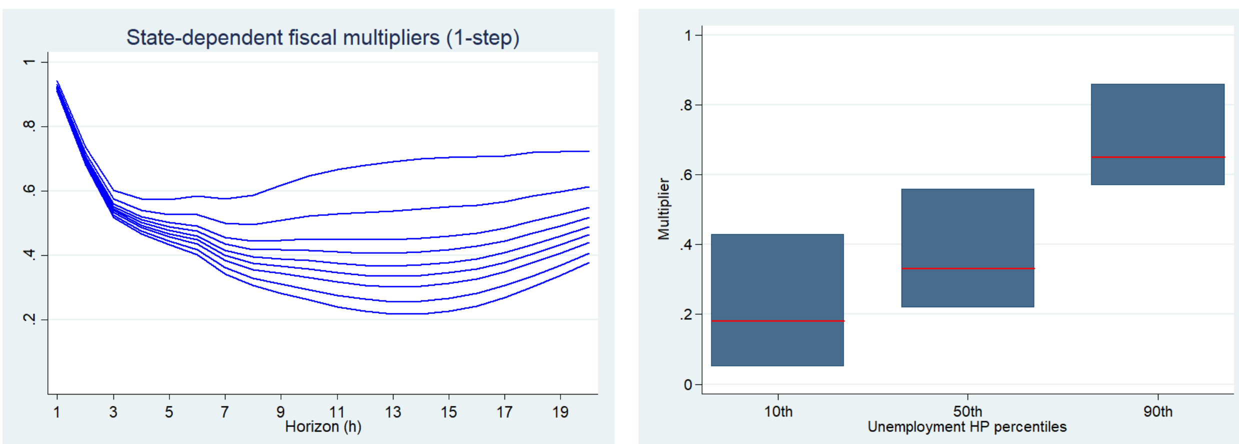


Note: The graph shows estimates of cumulative direct (left panel) and indirect effects (right panel), estimated through 4.

The fiscal multiplier – constructed as a linear combination of direct and indirect effects – varies significantly across different states of the economy, thereby reinforcing the evidence obtained through the "3-step" approach. As shown in the left panel of figure 6, the variability of the fiscal multiplier is greater in the medium term, when the indirect effects play a more prominent role. Focusing on $t + 16$, the right panel of figure 6 displays three different estimates of the fiscal multiplier at three distinct percentiles of unemployment[20]. The figure suggests that, even when accounting for the confidence bands, the fiscal multiplier is higher when unemployment is high compared to when it is low.

To formally assess the difference between multipliers across different values of the state variable, it is appropriate to test the hypothesis through a statistical procedure. Specifically, under this setting, a test on Γ_h would in principle be sufficient, as the difference between multipliers is proportional to this parameter and the difference between the values at which the multiplier is computed [21]. However, this method would not account for the uncertainty associated with the coefficient m_h, which itself contributes to the confidence bands shown in figure 6. To address this, I also report a more conservative inference procedure that accounts for the full uncertainty in the multiplier estimates by simulating the difference between states using independent bootstrap draws (see Appendix 6). This alternative and more cautious approach further confirms the results.

Figure 6: Fiscal multiplier under 1-step approach



Note: In the left panel, each line represents a different value of the fiscal multiplier obtained through equation 6, corresponding to values of S_t ranging from the 10^{th} to the 90^{th} percentile. Right panel shows estimate of the fiscal multiplier at t+16, including 90% bootstrapped confidence bands.

Overall, the findings in this section are broadly consistent using two different estimation approaches, showing a similar range of variability in the fiscal multiplier across different values of unemployment. It is worth emphasizing that the baseline specifications presented here are directly derived from Ramey and Zubairy (2018), who find no evidence of higher

[20] I consider the 10^{th}, 50^{th} and 90^{th} percentile. Since this approach imposes a linear relationship between the multiplier and S_{t-1}, it might be sentitive to extreme values of the state variable (below the 10^{th} or above the 90^{th}).

[21] Under the 1-step approach the fiscal multiplier is obtained as $\text{multi}_{p,h} = m_h + \Gamma_h \cdot S_p$ and the difference in multipliers at horizon h between two percentiles p_2 and p_1 of the state variable S is given by $\text{Diff}_{h,p_2,p_1} = (S_{p_2} - S_{p_1}) \cdot \Gamma_h$.

multipliers during periods of slack. Therefore, the contrasting results stem entirely from the adoption of a more recent methodological framework, echoing the concern raised by Gonçalves et al. (2024b) on the importance of using more robust methodological tools to handle state dependence in local projections.

5 Robustness

In the baseline, the state variable S_{t-1} is defined as the standardized unemployment gap, computed as the deviation of the observed unemployment rate from its trend component estimated using a Hodrick-Prescott (HP) filter (S_t^B). Here I perform a robustness exercise using different alternatives for the state variable [22]. Firstly, I consider a very simple measure of slack, namely, standardized unemployment (S_t^I). Secondly, I consider a standardized deviation of unemployment from a backward moving average or order 4 (S_t^{II}). Finally, an alternative and natural measure of the business cycle conditions is given by the output gap, obtained as the difference between real GDP and potential output (S_t^{III})[23]. To recap:

$$S_t^B = \frac{U^* - \overline{U^*}}{\sigma_{U^*}}, \quad U^* = U - U^{HP}; \qquad S_t^I = \frac{U - \overline{U}}{\sigma_U};$$
$$S_t^{II} = \frac{U^s - \overline{U^s}}{\sigma_{U^s}}, \quad U^s = U - U^{MA(4)}; \qquad S_t^{III} = \frac{\tilde{y} - \overline{\tilde{y}}}{\sigma_{\tilde{y}}}, \quad \tilde{y} = y - y^*$$

While the first three measures based on unemployment exhibit similar patterns, the fourth diverges markedly over the time series considered (see figure A.3).

The "1-step" estimates (obtained from equation 6) remain stable across different specifications of the state variable S_t. The cumulative indirect effects, Γ_h, are statistically significant, and their estimated values fall within a narrow range. In each specification, the fiscal multiplier computed at the 10th percentile of the state variable differs significantly from the value at the 90th percentile. Turning to the "3-step" estimates, results remain broadly unchanged when using alternative measures of unemployment. The evidence of state dependence is weaker when employing S_t^{II}, as the coefficients γ_h are not statistically significant at the 5% level for most horizons, although they become so at approximately one standard deviation. Nevertheless, the range of estimated fiscal multipliers remains broadly in line with the baseline (see Figure A.5). Furthermore, asymmetry persists, with the fiscal multiplier

[22] I haven't considered measures of output growth, as they are conceptually different from states of slack. Following a deep recession, an economy can experience positive growth for several quarters while remaining below its potential.

[23] I consider the value obtained in Ramey and Zubairy (2018), which estimates real potential GDP based on a 6th-degree polynomial fit from 1889:1 to 2015:4, omitting the Great Depression and WWII.

varying more in bad times while remaining broadly unchanged when real GDP exceeds the potential output. Across all specifications considered, using the news on defense spending as a fiscal shock (or instrument), the estimated multiplier ranges from values close to zero to values somewhat below one (see Table A.2).

Using BP shocks, the evidence remains – albeit weaker – only under the 1-step approach. Given the econometric framework implemented, which includes an interaction term between a lagged state variable and the fiscal shock, this attenuation may reflect the well-known anticipation effects. Indeed, lagging the state variable might not ensure orthogonality between the shock and the state variable if the latter is influenced by future values of the shock. This underscores the importance of relying on truly unanticipated fiscal shocks and, therefore, orthogonal to lagged macroeconomic variables. Without such exogeneity, the estimates might be biased, as the interaction term could reflect endogenous variation rather than genuine state dependence.

6 Conclusions

This paper re-evaluates the influential findings of Ramey and Zubairy (2018), who provided estimates of the government spending multiplier using two-regime local projections. Their original results suggested no evidence of state dependence in fiscal multipliers across economic conditions.

By applying the more recent methodological framework proposed by Cloyne et al. (2023), which introduces a novel approach to assessing state dependence, I have found that the fiscal multiplier is significantly larger during periods of economic slack – that is, in "bad times" characterized by high unemployment or below-trend output. Overall the paper shows that the fiscal multiplier ranges between 0 and 1 across different specifications, with the multiplier tending to decline more quickly as the economy moves toward full capacity, while it increases only marginally during economic downturns. This asymmetry is consistent with the argument that, when the economy is operating above potential, fiscal interventions are less effective at stimulating demand due to limited slack in the labor market and other productive resources. In such conditions, additional government spending may simply result in crowding-out effects rather than fostering real economic growth.

This central result — a larger fiscal multiplier in bad times — is robust to the choice of estimation approach. Specifically, it holds both when using the more flexible "3-step" estimation procedure, which derives the multiplier as the ratio between the cumulative IRFs for GDP and government spending, and the more efficient "1-step" approach, which directly estimates the fiscal multiplier. The evidence of state dependence is not sensitive to the specific definition of the economic state. Whether the state variable is based on the unemployment gap (measured in different ways) or the output gap, the finding of stronger fiscal effects in bad times remains. The evidence is weaker when using Blanchard-Perotti (BP) shocks instead of narrative shocks, which is consistent with Ramey and Zubairy (2018), who show that BP shocks may not be a relevant instrument for identifying government spending surprises.

Fiscal multipliers below one indicate that fiscal policy is only relatively effective at stimulating the economy. This finding is consistent with trends observed following the last two major crises− the global financial crisis and the COVID-19 pandemic − after which governments across all major economies responded with massive fiscal stimuli, leading to a dramatic rise in government debt.

However, there are a few important considerations to keep in mind. First, the rationale for fiscal intervention during periods of economic distress lies not only in stimulating aggregate demand but also in protecting the most vulnerable and preventing a further rise in inequality. This objective becomes even more important when monetary policy is constrained—such as at the zero lower bound—or faces other limitations. Therefore, a relatively low multiplier does not mean that the government should remain passive during a major crisis. Moreover, different fiscal instruments — other than government spending — can have very different effects on output. In this regard, it is important to consider well-targeted policies that are more likely to generate larger multipliers, as suggested by the existing literature (e.g., Oh and Reis (2012)).

Regarding long-term and sustainability implications, the combination of large stimuli during crises and moderate growth has inevitably led to rising government debt. Fiscal consolidation measures may therefore be required to ensure that debt remains on a sustainable trajectory, particularly in economies with limited growth potential. If we assume that fiscal shocks exhibit no asymmetry with respect to their sign[24], the results of this paper suggest that such consolidation efforts - made necessary by the high levels of debt reached across most advanced economies — can be implemented with relatively low economic cost when the economy is operating near full capacity, as the fiscal multiplier tends to decline rapidly toward zero under those conditions.

Lastly, while this paper provides evidence of state dependence in the magnitude of fiscal multiplier, this research would benefit from a further exploration of the underlying transmission mechanisms. Investigating whether the asymmetry in the multiplier is driven by

[24]As shown in Ben Zeev et al. (2023), who found differences in short-run impulse responses, whereas the overall government spending multiplier is symmetric, with increases and decreases in spending producing comparable effects on economic output.

differences in household behavior, firm investment decisions, or other channels is essential for a deeper understanding of fiscal policy effectiveness. A detailed examination of this important aspect is left for future research.

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Appendix

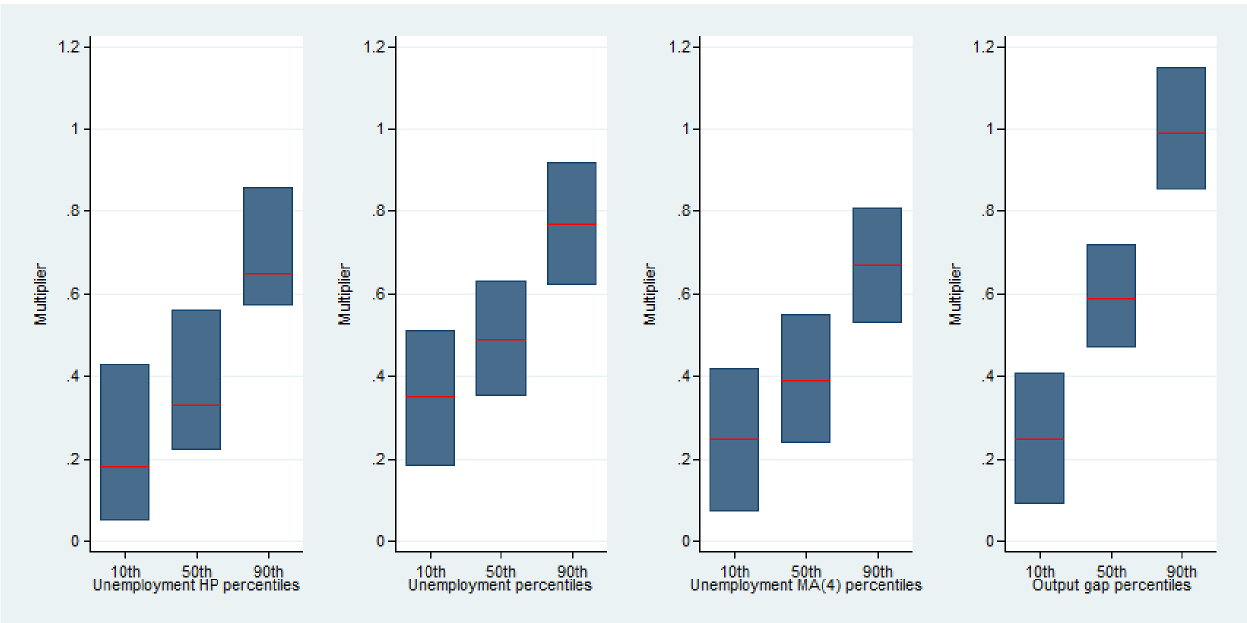
A.1 Estimate results using different definitions of S_{t-1}

Table A.1: Bootstrap estimates of 1-step equation using different S_{t-1}

	Baseline			S^{I} (std unemp.)			S^{II} (dev from MA(4))			S^{III} (output gap)		
	Avg.	LB	UB	Avg.	LB	UB	Avg.	LB	UB	Avg.	LB	UB
m_{16}	0.56	0.42	0.70	0.42	0.32	0.62	0.60	0.48	0.73	0.46	0.32	0.61
Γ_{16}	0.26	0.17	0.35	0.32	0.22	0.42	-0.28	-0.34	-0.22	0.26	0.18	0.34

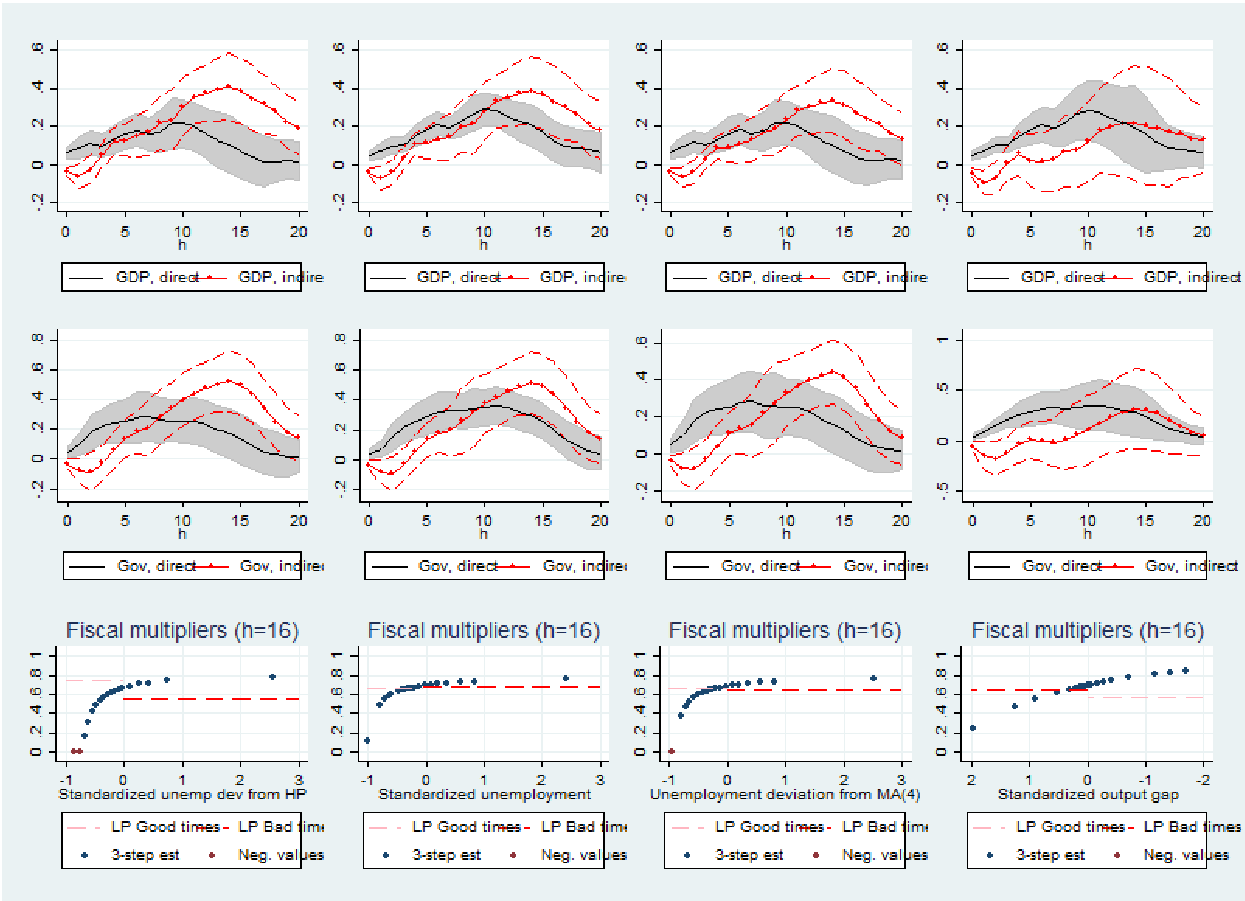
Note: The table reports the average, 10th percentile (lower bound), and 90th percentile (upper bound) of bootstrapped estimates for the direct (m_h) and indirect (Γ_h) cumulative effect.

Figure A.1: Bootstrap estimates of 1-step fiscal multipliers using different S_{t-1}



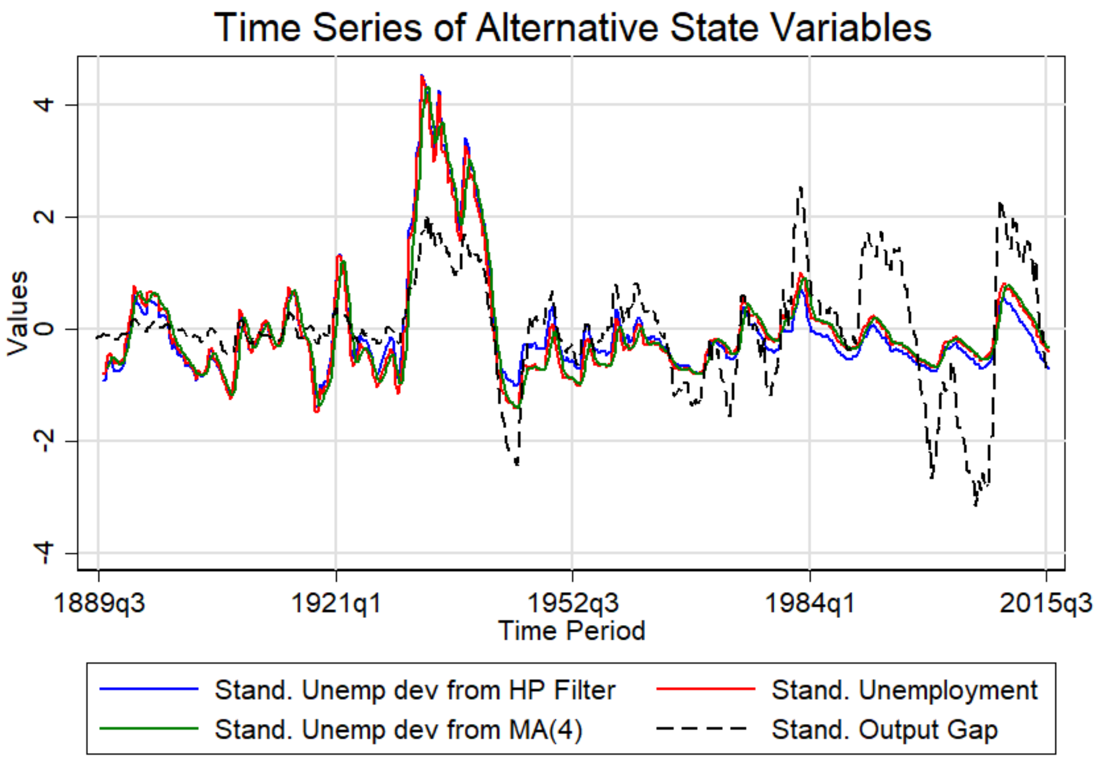
Note: The figure reports the state-dependent fiscal multipliers obtained with bootstrapped 1-step estimates, for the 10th, 50th and 90th percentile of the state variable. Estimates are obtained using news on military spending as an instrument for Government spending. For S^{III}_{t-1}, the fiscal multiplier at the 10th percentile and 90th percentile are reversed.

Figure A.2: β_h, γ_h and 3-step fiscal multipliers using different S_{t-1}



Note: This figure shows the results obtained using the 3-step approach with different values of S_t and news on defense spending as a fiscal shock. In the fourth specification, lagged unemployment is kept as a control (in addition to lagged output gap); furthermore, the x-axis is reversed to facilitate the comparison.

Figure A.3: State variables S_{t-1}



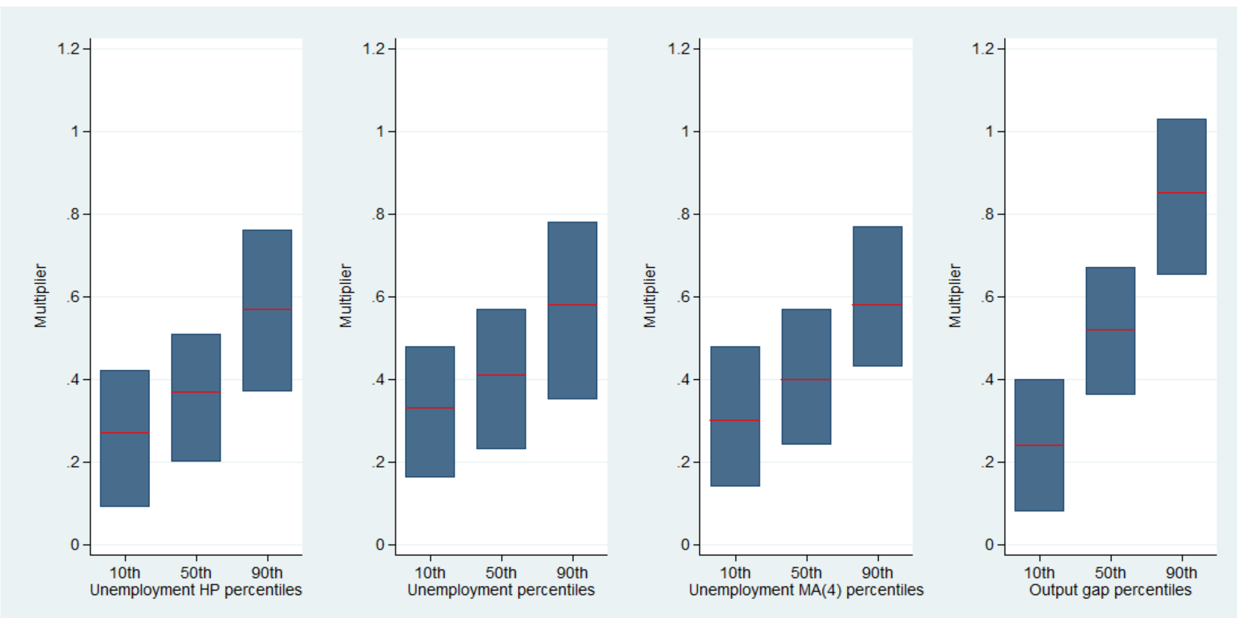
A.2 Estimates results using Blanchard-Perotti shocks

Table A.2: Bootstrap estimates of 1-step equation using different S_{t-1}

	Baseline			S^{I} (std unemp.)			S^{II} (dev from MA(4))			S^{III} (output gap)		
	Avg.	LB	UB	Avg.	LB	UB	Avg.	LB	UB	Avg.	LB	UB
Direct	0.45	0.27	0.61	0.42	0.25	0.58	0.52	0.37	0.68	0.44	0.30	0.62
Indirect	0.16	0.07	0.24	0.20	0.11	0.30	-0.23	-0.28	-0.17	0.17	0.08	0.27

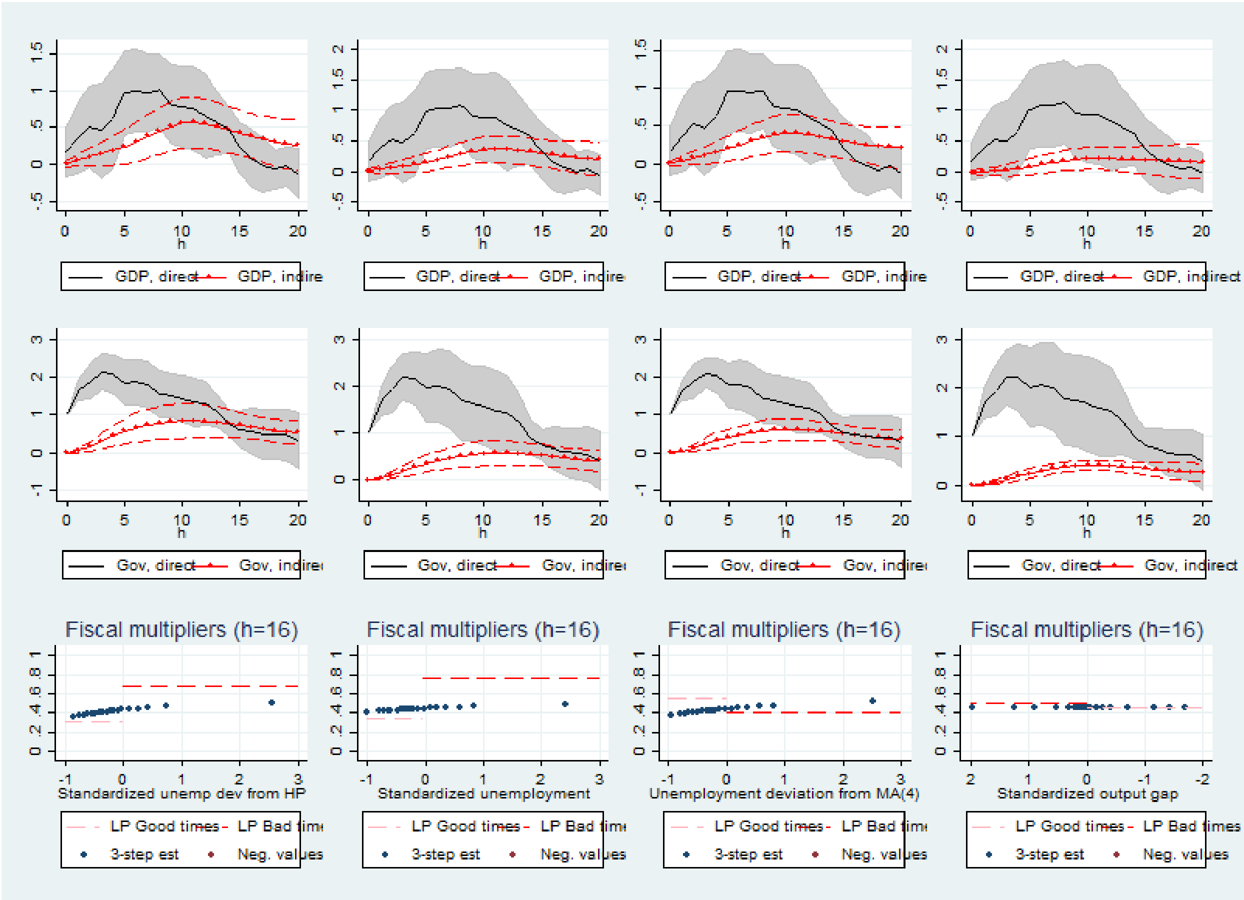
Note: The table reports the average, 10th percentile (lower bound), and 90th percentile (upper bound) of bootstrapped estimates for the direct (m_h) and indirect (Γ_h) cumulative effect.

Figure A.4: Bootstrap estimates of 1-step fiscal multipliers using different S_{t-1} - BP shocks



Note: The figure reports the state-dependent fiscal multipliers obtained with bootstrapped 1-step estimates, for the 10th, 50th and 90th percentile of the state variable. Estimates are obtained using BP shocks as an instrument for Government spending. For S^{III}_{t-1} the fiscal multiplier at the 10th percentile and 90th percentile are reversed.

Figure A.5: β_h, γ_h and 3-step fiscal multipliers using different S_{t-1}



Note: This figure shows the results obtained using the 3-step approach with different values of S_t and BP shocks. In the fourth specification, lagged unemployment is kept as a control (in addition to lagged output gap); furthermore, the x-axis is reversed to facilitate the comparison.

A.3 A note on State-dependent Local projections

The methodology employed in this paper is based on Cloyne et al. (2023), who use the Kitagawa-Blinder-Oaxaca decomposition to explore the response's heterogeneity over states of the economy. Empirically, this approach entails a straightforward extension of the standard local projection, by introducing interaction terms obtained as a product of the shock and a set of contemporaneous (demeaned) covariates x_t:

$$y_{t+h} - y_{t-1} = \alpha_h + \Gamma(L)ctrl_{t-1} + \beta_h \epsilon_t + \gamma_h (x_t - \overline{x})\epsilon_t + u_{t+h}, h = 0, 1, ..., H \tag{7}$$

The authors emphasize that, when assessing state dependence, both the shock and the state variable (i.e., one of the covariates) must be exogenous to ensure unbiased estimates and enable a causal interpretation. This condition can be satisfied, for example, by relying on an exogenous shock and instrumenting the state variable using a valid instrument. Using a contemporaneous state variable is crucial for specific exercises, such as the one proposed by the authors, which assesses how monetary-policy interaction affects the size of the fiscal multiplier. In this paper, since the research question focuses on how the effectiveness of fiscal policy varies with the business cycle, it is more natural to define the interaction term using a lagged state variable capturing the initial level of economic activity before the shock occurs; furthermore, this choice aligns with the government's budgeting and decision-making process, which relies on the most recent data available—typically from the previous quarter. Thus, lagging the state variable offers a more realistic representation of how policymakers should account for business cycle conditions when designing and implementing fiscal policy. Furthermore, from an econometric point of view, as long as the state variable can be considered independent of future values of the shocks, using its lagged value should be sufficient to mitigate potential endogeneity issues. This approach has been recently employed in Caramp and Feilich (2024), who used a specification close to equation (1) to assess how predetermined values of U.S. public debt affect the transmission of monetary policy shocks.

A recent paper by Gonçalves et al. (2024a) provides an in-depth analysis of state-dependent local projections, explicitly addressing the endogeneity issue in different settings including the one hereby implemented.

In particular, the authors consider an example inspired by Cloyne et al. (2023) and Caramp and Feilich (2024) and highlight that in this setting the LPs might fail to recover the true IRF. In particular, the DGP considered is the following:

$$\begin{cases} x_t = \epsilon_{1t} \\ r_t = f(x_{t-1}) + \epsilon_{3t} \\ y_t = \beta_{21}x_t + \beta_{23}r_{t-1} + \alpha_{21}x_t \times r_t + \gamma_{21}y_{t-1} + \epsilon_{2t} \end{cases}$$

where ϵ_{it} are $N(0,1), iid$. x_t can be interpreted as an exogenous shock, r_t as a state variable[25].

The definition of the impulse response function is based on the Conditional Average Response (CAR), which compares the baseline value $Y_{t+h}(\varepsilon_{1t})$ with the counterfactual value of Y at $t+h$ that would have been observed if ε_{1t} had been subject to a shock of size δ, denoted $Y_{t+h}(\varepsilon_{1t}+\delta)$, conditioning on the information set Ω_t:

$$\mathrm{CAR}_h(\delta,\omega) \equiv E\left[Y_{t+h}(\varepsilon_{1t}+\delta) - Y_{t+h}(\varepsilon_{1t}) \mid \Omega_t = \omega\right],$$

Given this definition, it is possible to compute the true IRF for any horizon and highlight potential estimation issues. At t, the effect of the shock is a linear function of the true shock, and can therefore be recovered consistently via LPs:

$$\begin{aligned} IRF_t = CAR_0(\delta,\omega) =& E[y_t(e+\delta) - y_t(e)|r_t] \\ =& E[(\beta_{21}(e+\delta) + \beta_{23}r_{t-1} + \alpha_{21}(e+\delta) \times r_t + \gamma_{21}y_{t-1} + \epsilon_{2t}) - \\ & (\beta_{21}(e) + \beta_{23}r_{t-1} + \alpha_{21}(e) \times r_t + \gamma_{21}y_{t-1} + \epsilon_{2t})|r_t] \\ =& \delta(\beta_{21} + \alpha_{21}r_t) \end{aligned}$$

Starting from $t+1$, a shock on ϵ_{1t} has a primary effect on y_t through β_{21} and α_{21}; a secondary effect through β_{23} due to the relationship between the state variable r_t and the shock:

$$\begin{aligned} IRF_{t+1} = CAR_1(\delta,\omega) &= E[y_{t+1}(e+\delta) - y_{t+1}(e)|r_t] = \\ &= \beta_{23}E(f(e+\delta) - f(e)] + \gamma_{21} \times CAR_{0,\delta} \end{aligned}$$

as $E[\beta_{21}x_{t+1}(e+\delta)] = E[\beta_{21}x_{t+1}(e)|r_t] = 0$ and $E[\alpha_{21}x_{t+1}f(e+\delta) - \alpha_{21}x_{t+1}f(e)|r_t] = 0$.

For $t+h$, it is possible write:

$$IRF_{t+h} = CAR_h(\delta,\omega) = \gamma_{21}CAR_{h-1}(\delta,\omega)$$

[25] The same results hold if the state variable is lagged but responds contemporaneously to the shock – specifically, if the data-generating process is given by $y_t = \beta_{21}x_t + \beta_{23}r_{t-1} + \alpha_{21}x_t \times r_{t-1} + \gamma_{21}y_{t-1} + \epsilon_{2t}$, with $r_t = f(x_t) + \epsilon_{3t}$. This structure is closer to the specification considered in equation (2).

In cases where f is linear, the difference $f(\epsilon_{1t} + \delta) - f(\epsilon_{1t})$ is also a linear function of δ, and the IRF can still be recovered using standard local projections. However, when $f(\epsilon_t)$ is nonlinear, local projections yield biased estimates.

Having said that, it is worthwhile to assess the relationship between the fiscal shock and the state variables considered in this paper to check for potential nonlinearities that could undermine the validity of our estimates. First, one could argue that there are no strong priors to expect the contemporaneous response of unemployment to a fiscal shock to be nonlinear, and to the best of my knowledge, there is also no empirical evidence in the literature supporting such a nonlinearity. Second, we conducted a brief data analysis to test this hypothesis by estimating a set of regressions of the form $S_t = a\epsilon_t + bf(\epsilon_t) + \Phi ctrl_{t-1}$, where $f(\cdot)$ denotes a nonlinear function. As shown in Table A.3, the coefficient b associated with the nonlinear term is not statistically significant in any of the specifications considered. While these tests may not be exhaustive in principle, they can be considered sufficient for the purposes of this exercise.

Table A.3: Testing non-linear relationship between S_t and ϵ_t

	S^b				S^I				S^{II}				S^{III}			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
ϵ	-0.428	-1.397*	-1.287	-0.640	-0.730*	-1.130	-1.260	-0.855	-0.457	-1.490*	-1.372	-0.682	0.885***	0.757	0.940	1.014*
	(0.345)	(0.652)	(0.703)	(0.624)	(0.340)	(0.644)	(0.694)	(0.615)	(0.368)	(0.695)	(0.750)	(0.665)	(0.237)	(0.449)	(0.484)	(0.428)
$\|\epsilon\|$		1.170				0.482				1.248				0.155		
		(0.669)				(0.660)				(0.713)				(0.460)		
ϵ^2			1.905				1.175				2.032			-0.120		
			(1.360)				(1.342)				(1.451)			(0.936)		
ϵ^3				0.797				0.471				0.850				-0.484
				(1.961)				(1.932)				(2.091)				(1.346)

Notes: Standard errors in parentheses. * p<0.05, ** p<0.01, *** p<0.001.
All regressions include four lags of the shock variable, output (y), and government spending (g), as in the benchmark specification.

A.4 Testing the difference between fiscal multipliers

To assess whether fiscal multipliers differ significantly across economic states, I consider the difference in multipliers at horizon h between two percentiles p_2 and p_1 of the state variable S:

$$\text{Diff}_{h,p_2,p_1} = \text{multi}_{p_2,h} - \text{multi}_{p_1,h} \tag{8}$$

Recalling the linear form of the multiplier, under the 1-step approach:

$$\text{multi}_{p,h} = m_h + \Gamma_h \cdot S_p \tag{9}$$

where m_h denotes the intercept term representing the average multiplier at horizon h when the state variable is zero, and Γ_h captures the indirect effect of the state variable on the multiplier at that horizon. This difference can be rewritten as:

$$\text{Diff}_{h,p_2,p_1} = (S_{p_2} - S_{p_1}) \cdot \Gamma_h \tag{10}$$

with associated variance:

$$\text{Var}(\text{Diff}_{h,p_2,p_1}) = (S_{p_2} - S_{p_1})^2 \cdot \text{Var}(\Gamma_h) \tag{11}$$

Hence, inference on Diff_{h,p_2,p_1} reduces to inference on Γ_h. If the latter is significantly different from zero, the difference in multipliers will also be statistically significant, also when $S_{p_2} - S_{p_1}$ is small. Consequently, testing the difference between multipliers at different values of the state variables provides no additional insight beyond the significance of Γ_h. [26]

However, one might want to take into account also the uncertainty surrounding the estimates of the average effect. To this end, it is possible to consider two independent estimates $\text{multi}_{p_2,h}$ and $\text{multi}_{p_1,h}$ from two independent bootstrap simulations. The difference is then obtained between two multipliers computed using independently estimated parameters:

$$\text{Diff}^*_{h,p_2,p_1} = \left(m_h^{(1)} + \Gamma_h^{(1)} \cdot S_{p_2}\right) - \left(m_h^{(2)} + \Gamma_h^{(2)} \cdot S_{p_1}\right) \tag{12}$$

This approach allows for variation in both m_h and Γ_h, yielding a more realistic distribution of Diff_h. The variance becomes:

$$\text{Var}(\text{Diff}^*_{h,p_2,p_1}) = 2 \cdot \text{Var}(m_h) + (S_{p_2} - S_{p_1})^2 \cdot 2 \cdot \text{Var}(\Gamma_h) \tag{13}$$

[26] This can be done in a standard bootstrap setting where the difference is computed within the same bootstrap draw. As the parameters m_h and Γ_h are fixed in both percentiles, the terms m_h cancel out.

assuming m_h and Γ_h are uncorrelated and identically distributed across the two simulations.

This more conservative approach offers a cautious framework for inference, as it fully incorporates the uncertainty associated with the direct effect. This strategy prevents automatic rejection of the null based solely on the significance of Γ_h and offers a prudent assessment of the statistical difference between fiscal multipliers for different percentiles.

Figure A.6: Testing the difference between $\text{multi}_{90,16}$ and $\text{multi}_{10,16}$ using different S_t