Labor Market Institutions, Fiscal Multipliers, and Macroeconomic Volatility

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Abstract

How do labor market institutions shape the transmission of government spending shocks and macroeconomic volatility? We develop a theoretical model in which labor market institutions affect fiscal transmission through their effect on wage rigidity, job separation, and matching frictions. We estimate an interacted panel vector autoregressive model for 16 OECD economies and study how macroeconomic responses to government spending shocks vary with institutional labor market characteristics. In line with our theoretical predictions, we show that institutions that stabilize employment and wages tend to dampen output volatility and attenuate the response of output and employment to government spending shocks.

Keywords: Fiscal policy; labor market institutions; interacted panel VAR

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1. Introduction

Government spending is a central policy tool during economic downturns, yet its macroeconomic effects vary substantially across countries and over time. A growing literature documents large differences in the response of output, employment, and wages to fiscal shocks, even among advanced economies. Understanding the sources of this heterogeneity is crucial for assessing the effectiveness of fiscal policy and its role in macroeconomic stabilization.[1]

This paper studies how labor market institutions shape the transmission of government spending shocks and the volatility of macroeconomic outcomes. Structural characteristics of labor markets help explain differences in the absorption of external shocks across economies (Gnocchi, Lagerborg and Pappa, 2015; Abbritti and Weber, 2018). Labor market institutions influence wage setting, job separation, and matching efficiency, all of which affect how fiscal expansions propagate through the economy. While existing work has emphasized the role of individual institutions (most prominently, employment protection), much less is known about how multiple labor market institutions jointly shape fiscal transmission and macroeconomic volatility.

We address this gap by combining a structural model of the labor market with cross-country empirical evidence for OECD economies. On the theoretical side, we develop a model in which labor market institutions affect fiscal transmission through three main channels: wage rigidity, job separation dynamics, and matching frictions. These channels generate testable implications for the responses of output, employment, and real wages following government spending shocks. On the empirical side, we estimate impulse responses to fiscal shocks for a panel of 16 OECD economies and examine how their magnitude and persistence vary systematically with labor market characteristics.

Our theoretical framework combines a standard New Keynesian model with search and matching frictions in the labor market (Thomas and Zanetti, 2009; Christoffel, Kuester and Linzert, 2009). We focus on three measurable institutional dimensions. *Union density* determines workers' bargaining power in wage negotiations, the *unemployment benefit replacement rate* enters as a subsidy to unemployed workers, and *employment protection* is captured by firing costs per displaced worker. In the model, higher unemployment

[1] Labor market outcomes are often a primary target of macroeconomic stabilization policies. In the wake of the financial crisis of 2007–09, the American Recovery and Reinvestment Act (ARRA) explicitly aimed to raise employment (Romer, 2009). During the COVID-19 pandemic, the Coronavirus Aid, Relief, and Economic Security (CARES) Act expanded unemployment insurance to support displaced workers (Ganong, Noel and Vavra, 2020; Ganong et al., 2024). In Europe, the Next Generation EU program allowed member states greater fiscal flexibility and supported national relief measures.

benefits and greater union density increase real wages, which dampens the employment response to an expansionary fiscal shock. Higher employment protection reduces employment adjustment by lowering job separation rates.[2]

We validate the model's predictions using a semi-structural Bayesian interacted panel VAR estimated for 16 OECD countries (Towbin and Weber, 2013; Sá, Towbin and Wieladek, 2014). Identification relies on implementation lags of government spending (Blanchard and Perotti, 2002; Ilzetzki, Mendoza and Végh, 2013) and exogeneity restrictions on the variables measuring labor market institutions. The empirical results broadly corroborate the qualitative predictions of the theoretical model. We find that higher union density and stronger employment protection are associated with smaller output and employment responses to fiscal shocks. At the same time, these institutions reduce output volatility, whereas higher unemployment benefit replacement rates increase volatility in output and employment.

To shed light on the transmission mechanisms underlying these findings, we perform an impulse-response matching exercise linking empirical responses to the theoretical channels implied by different labor market institutions. This analysis shows that employment protection primarily operates through job separation risk, while union density additionally affects matching frictions in the labor market.

Our contribution advances the literature in several respects. First, we contribute to the extensive literature on the macroeconomic effects of fiscal policy shocks, and in particular to studies analyzing labor market responses to government spending (Monacelli, Perotti and Trigari, 2010; Brückner and Pappa, 2012). While a large body of work emphasizes state-dependent fiscal effects across business cycle regimes (Auerbach and Gorodnichenko, 2012), our analysis highlights how structural labor market characteristics shape the effectiveness of fiscal policy.

Second, we contribute to the literature on the macroeconomic consequences of labor market regulation. A growing theoretical literature studies the interaction between fiscal policy and labor market frictions in New Keynesian environments with search and matching (Christoffel, Kuester and Linzert, 2009; Thomas and Zanetti, 2009; Zanetti, 2009; Zanetti, 2011; Campolmi, Faia and Winkler, 2011). Complementing this work, empirical studies document that labor market institutions affect macroeconomic dynamics and shock absorption (Gnocchi, Lagerborg and Pappa, 2015; Abbritti and Weber, 2018; Hantzsche, Savsek and

[2] Ghassibe and Zanetti, 2022 study goods market frictions and show that tighter product markets are associated with smaller fiscal responses.

Weber, 2018). Our analysis extends this literature by examining how labor market institutions condition the transmission of exogenous fiscal policy shocks.

Our work is closely related to Cacciatore et al. (2021), who study the role of employment protection legislation in shaping fiscal spending responses. While their focus is on a single labor market institution, we consider a broader set of institutional features and assess their joint effects. This allows us to study potential complementarities and substitutabilities between labor market institutions in shaping fiscal transmission. Moreover, whereas existing work primarily focuses on fiscal responses, we explicitly analyze how labor market institutions affect macroeconomic volatility, highlighting trade-offs between stabilization and fiscal effectiveness that are not captured by multiplier estimates alone.

The remainder of the paper is structured as follows. Section 2 provides a descriptive overview of labor market institutions across selected OECD countries. Section 3 presents the theoretical model and its main predictions. Section 4 connects the theoretical framework to the empirical analysis and presents the econometric results. Section 5 concludes.

2. Structural Labor Market Indicators in OECD Economies

We begin by describing the evolution of labor market institutions (LMIs) in OECD countries using available data for the period 1960–2020.[3] We focus on three variables that correspond to structural dimensions of the labor market playing a central role in our theoretical model: (a) union density (UD, η), (b) the unemployment benefit replacement rate (BRR, φ), and (c) employment protection legislation (EPL, ς).[4] We provide a descriptive analysis of cross-country and within-country variation in these LMIs in Figure 1.[5]

In the theoretical model, the parameter η captures workers' bargaining power, which we proxy empirically using union density. This indicator is primarily derived from survey data or adjusted administrative records and represents the share of trade union members among wage and salary earners.[6] Higher values of union

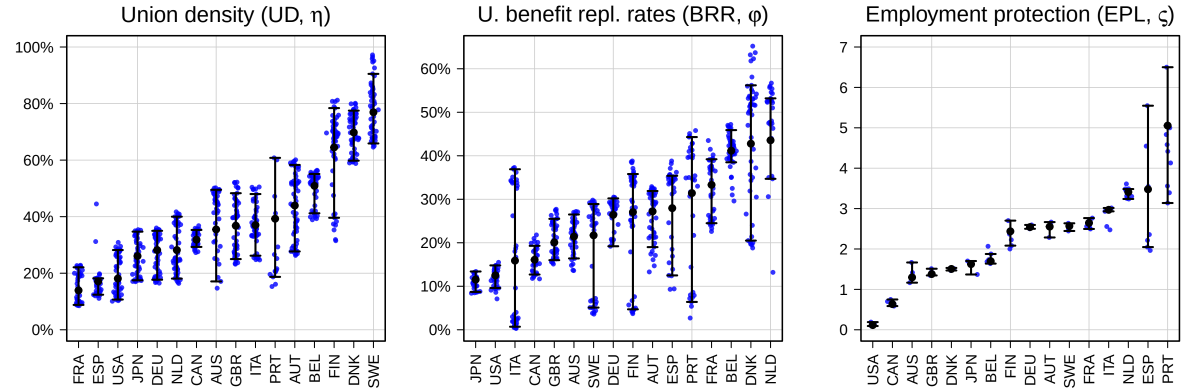
[3] Our sample consists of 16 countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Italy, Japan, the Netherlands, Portugal, Spain, Sweden, and the United States. Further details on the data sources are provided in Appendix D.

[4] Other structural labor market characteristics may also affect the transmission of fiscal shocks. Examples include openness to foreign labor (see, for instance Amuedo-Dorantes and Rica, 2013; Godøy, 2017; Schiman, 2021), long-run trends in labor productivity (see, for instance Policardo, Punzo and Sánchez Carrera, 2019; Li et al., 2021), and the demographic structure of the workforce (see, for instance Docquier et al., 2019). We limit our analysis to the dimensions listed above due to data availability and their suitability for the theoretical framework.

[5] Additional evidence on the time variation of these LMIs is presented in Figure D1 in the appendix.

[6] An alternative indicator of wage bargaining power is collective bargaining coverage. This measure is considerably more persistent than union density and does not necessarily reflect changes in effective bargaining power at the firm or worker level.

Figure 1: Labor Market Institutions in OECD Economies: Cross-Country Variation



Notes: The figure reports labor market institution variables (union density, unemployment benefit replacement rate, and employment protection legislation) by country. The black dot represents the mean, while the whiskers correspond to the 10th and 90th percentiles of the distribution. Blue points denote observed country-year data. The y-axis is measured in percent for union density and the benefit replacement rate; the EPL index is unit-free.

density indicate greater trade union influence in wage negotiations. Union density displays substantial cross-country variation, ranging from around 15% in France to approximately 80% in Sweden. It also exhibits considerable within-country variation, amounting to changes of 10–20 percentage points over the period considered.

The unemployment benefit replacement rate measures the proportion of income maintained during unemployment, defined as the ratio of net household income while unemployed to income prior to job loss. Higher values indicate more generous unemployment insurance systems. In the theoretical model, the parameter φ captures unemployment benefit payments relative to pre-dismissal wages and is measured empirically using the replacement rate. Similar to union density, the replacement rate shows substantial variation both across countries and over time.

We proxy the degree of employment protection (ς in the model) using the EPL index. This index measures the regulatory strictness of dismissals and temporary contracts based on regulations in force on January 1 of each year. In addition to severance payments, it incorporates one-time payments to the social security system related to the burden imposed on unemployment insurance by dismissals.[7] Higher index values indicate stricter employment protection. While the median level of EPL has remained broadly stable across countries, its dispersion has declined markedly over time, reflecting convergence, particularly among European Union member states (see Figure D1).

[7] The EPL index is qualitative in nature. Quantitative cross-country data on firing costs are not readily available, and monetary measures alone would omit non-pecuniary elements of employment protection, such as administrative and judicial procedures.

3. The Theoretical Model

For our theoretical model, we combine a Diamond–Mortensen–Pissarides search-and-matching labor market with an otherwise standard business-cycle framework, building on Merz (1995), Andolfatto (1996) and Krause and Lubik (2007) and Monacelli, Perotti and Trigari (2010). The model is deliberately parsimonious and is designed to isolate the role of labor market institutions as determinants of the fiscal multiplier. Appendix B discusses extensions that allow for richer interactions and that can be useful for other applications.

We assume representative firms and households. Each firm employs n_t workers and posts v_t vacancies. Posting a vacancy costs κ per vacancy, and endogenous job separations are associated with firing costs of b_t^s per separated worker. Unemployment is given by $u_t = 1 - n_t$.

New hires are determined by a standard matching function,

$$m_t = \bar{m} u_t^{\gamma} v_t^{1-\gamma}, \tag{3.1}$$

with $\bar{m} > 0$ and $\gamma \in (0,1)$. Labor market tightness is $\theta_t = v_t/u_t$. The probability that a vacancy is filled is $q_t = m_t/v_t = \bar{m}\theta_t^{-\gamma}$, and the job-finding probability is $p_t = m_t/u_t = \bar{m}\theta_t^{1-\gamma}$.

Each period, a fraction $\varrho(\tilde{a}_t)$ of existing matches is destroyed. Separations have an exogenous component $\bar{\varrho}$ and an endogenous component. Following Krause and Lubik (2007), idiosyncratic job productivities a_t are drawn each period from a distribution with c.d.f. $F(a_t)$ and density $f(a_t)$. A match is endogenously destroyed if $a_t < \tilde{a}_t$, where $\tilde{a}_t$ is a threshold chosen endogenously. The overall separation rate is

$$\varrho(\tilde{a}_t) = \bar{\varrho} + (1 - \bar{\varrho})F(\tilde{a}_t). \tag{3.2}$$

3.1 Intermediate Goods-Producing Firms

A representative intermediate-goods producer uses labor to produce output

$$y_t = \bar{A} n_t A(\tilde{a}_t), \tag{3.3}$$

where $\bar{A} > 0$ is a common aggregate productivity factor and

$$A(\tilde{a}_t) = E[a \mid a \geq \tilde{a}_t] = \frac{1}{1 - F(\tilde{a}_t)} \int_{\tilde{a}_t}^{\infty} a \, dF(a) \tag{3.4}$$

is the conditional expectation of productivity larger than the endogenously determined threshold.

Employment evolves according to

$$n_t = (1 - \varrho(\tilde{a}_t))(n_{t-1} + m_{t-1}). \tag{3.5}$$

Per period profits are given by

$$\pi_t^F = \frac{y_t}{\mu_t} - w_t n_t - \kappa v_t - F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b_t^s, \tag{3.6}$$

where the output price is normalized to one, μ_t is the price markup (so $mc_t = \mu_t^{-1}$), w_t is the average real wage across match productivities, and κ denotes the vacancy posting cost. The last term captures firing costs incurred for endogenously separated matches: cost per laid off worker, b_t^s, times the number of existing and new workers who were laid off due to endogenous job separation.[8] Firms maximize expected discounted profits, $\max_{n_t, v_t, \tilde{a}_t} \mathbb{E}_t \sum_{k \geq 0} \Lambda_{t,t+k} \pi_{t+k}^F$, where $\Lambda_{t,t+k}$ denotes the firm's stochastic discount factor. The optimization problem is subject to the production function and Eq. (3.5).

The first-order conditions imply

$$\begin{aligned}
F_t^n &= mc_t mpl_t - w_t + \mathbb{E}_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], && (3.7) \\
\frac{\kappa}{q_t} &= \mathbb{E}_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], && (3.8) \\
mc_t A(\tilde{a}_t) &= \frac{1}{\bar{A}} \left(w_t - b_t^s - \frac{\kappa}{q_t} \right), && (3.9)
\end{aligned}$$

where mpl_t is the marginal product of labor and F_t^n is the multiplier on Eq. (3.5).[9]

Eq. (3.7) defines the (shadow) value of an additional worker to the firm, which is given by the current marginal revenue product net of the wage, plus the discounted continuation value, net of expected firing costs. Eq. (3.8) is the free-entry condition, which equates the expected value of a vacancy to its posting cost. Eq. (3.9) pins down the endogenous separation threshold: higher wages increase the productivity cutoff required for continuing a match, while higher firing costs and higher expected hiring costs (lower q_t) make firms more willing to retain lower-productivity matches.

3.2 Final Goods-Producing Firms

Final-goods firms buy intermediate goods and sell differentiated varieties to households. There is a continuum of final-goods producers indexed by $i \in [0, 1]$. Firms are perfectly competitive in input markets and

[8] Appendix B considers the case where firing costs accrue to the government.

[9] Details for the condition determining $\tilde{a}_t$ are given in the appendix; the resulting expression implies Eq. (3.9).

monopolistically competitive in output markets. Prices are subject to Calvo stickiness à la Calvo (1983). Following Christiano, Eichenbaum and Evans (2005), firms that cannot re-optimize partially index prices to lagged CPI inflation, so that $P_t(j) = (1+\pi_{t-1})^{\gamma_p} P_{t-1}(j)$ with $\gamma_p \in [0,1]$. If $\gamma_p = 1$, non-re-optimizing firms fully index to lagged inflation; if $\gamma_p = 0$, they keep prices unchanged (see Uribe, 2020).

Let ξ denote the probability of not being able to re-optimize in a given period. Standard profit maximization yields a New Keynesian Phillips curve that is given by

$$(1+\gamma_p \xi \beta)\pi_t = \beta \mathbb{E}_t \pi_{t+1} + \gamma_p \pi_{t-1} + \frac{(1-\xi\beta)(1-\xi)}{\xi}\hat{mc}_t, \tag{3.10}$$

where $\hat{mc}_t$ denotes real marginal costs in log-deviations.

3.3 Households

Households are modeled as in Merz (1995). A representative household consists of a continuum of members of unit mass who pool income and consumption. Income accrues from labor earnings for employed members and unemployment benefits for unemployed members.

The household budget constraint is

$$c_t + B_t = R_{t-1}B_{t-1} + (1-\tau)w_t n_t + b_t^u(1-n_t) + T_t^S + \pi_t^F, \tag{3.11}$$

where B_t denotes holdings of government bonds paying a gross return of R_t, τ is the wage tax rate, b_t^u represents unemployment benefits per unemployed member, and T_t^S refers to lump-sum subsidies.

Employment evolves according to

$$n_t = (1-\varrho(\tilde{a}_t))n_{t-1} + p_t(1-n_{t-1}), \tag{3.12}$$

and the household maximizes expected lifetime utility $\mathbb{E}_t \sum_{k\geq 0} \beta^k u(c_{t+k}, n_{t+k})$ subject to both equations, where λ_t is the Lagrange multiplier attached to Eq. (3.11) and $\lambda_t H_t^n$ the one attached to Eq. (3.12). Optimality implies

$$1 = R_t \mathbb{E}_t \Lambda_{t,t+1}, \tag{3.13}$$

$$H_t^n = \tilde{w}_t^b - mrs_t + \mathbb{E}_t\left[1-\varrho(\tilde{a}_{t+1}) - p_{t+1}\right]\Lambda_{t,t+1}H_{t+1}^n, \tag{3.14}$$

where $\tilde{w}_t^b = (1-\tau)w_t - b_t^u$ and $mrs_t = -u_{n,t}/\lambda_t$. Hence, mrs_t captures both the marginal rate of substitution between consumption and work and the marginal value of non-work activities. The stochastic discount factor is given by $\Lambda_{t,t+1} = \beta\frac{\lambda_{t+1}}{\lambda_t}$.

Eq. (3.14) defines the (shadow) value of having one additional employed member, which is given by the increase in utility due to higher income, net of the disutility from work, plus the discounted continuation value of being employed next period.

3.4 Nash Wage Bargaining

Wages are set each period via Nash bargaining over the pre-tax average wage w_t between firms and workers, leading to

$$w_t = \arg\max_{w_t} (H_t^n)^\eta (F_t^n)^{1-\eta}, \tag{3.15}$$

where $0 < \eta \leq 1$ captures workers' bargaining power. The Nash solution implies

$$w_t = (1-\eta)\frac{mrs_t + b_t^u}{1-\tau} + \eta\left(mc_t mpl_t + \mathbb{E}_t\Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right]\right). \tag{3.16}$$

The wage is a weighted average of the household's outside option (captured by mrs_t and unemployment benefits) and the firm's surplus from the match (captured by marginal revenue product net of expected hiring and firing costs).

3.5 Policies, Aggregate Resource Constraint, and Government Budget Constraint

The government budget constraint is

$$\tau w_t n_t + B_t = R_{t-1}B_{t-1} + b_t^u u_t + T_t^S + g_t, \tag{3.17}$$

where g_t denotes government consumption.

Fiscal policy is specified as follows. Government consumption g_t follows an exogenous AR(1) process (in log-deviations). Unemployment benefits are given by

$$b_t^u = \bar{\varphi} + \varphi w_{t-1}, \tag{3.18}$$

where φ is the replacement rate relative to the last wage received. Firing costs are specified as

$$b_t^s = \bar{\varsigma} + \varsigma w_{t-1}. \tag{3.19}$$

Finally, lump-sum subsidies satisfy

$$T_t^S = \bar{T}^S + \varphi_{T^S} B_t. \tag{3.20}$$

The constants $\bar{T}^S$ and $\bar{\varsigma}$ facilitate steady-state computations, while the feedback term $\varphi_{T^S} B_t$ ensures debt stationarity.

Monetary policy follows a Taylor rule,

$$i_t = \rho_i i_{t-1} + (1 - \rho_i) \left(\phi_\pi \pi_t + \phi_y \hat{y}_t \right), \tag{3.21}$$

where $i_t = \ln(R_t)$ and a hat denotes the log-deviation from the steady state.

Finally, combining Eq. (3.17), the household budget constraint in Eq. (3.11), and firms' profits π_t^F yields the aggregate resource constraint,

$$y_t = c_t + g_t + \kappa v_t + F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b_t^s. \tag{3.22}$$

This equation closes the model.

3.6 Equilibrium, Model Solution, and Dynamic Simulations

We collect the LMI parameters of interest in the vector $\boldsymbol{\vartheta} = (\eta, \varphi, \varsigma)'$. To assess how labor market institutions shape fiscal transmission, we study how changes in $\boldsymbol{\vartheta}$ affect the impulse responses to a government spending shock.

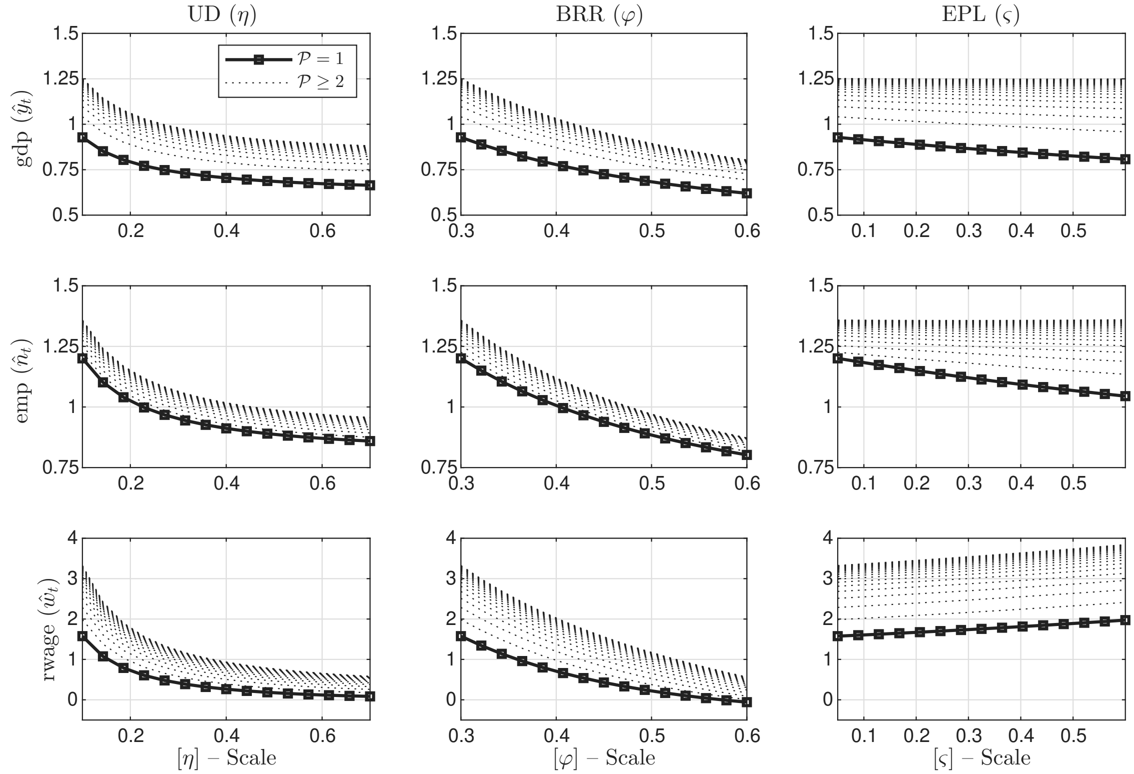
The model is solved by log-linearizing the system around its deterministic steady state. The equilibrium dynamics can be summarized by

$$\boldsymbol{\Psi}_0(\boldsymbol{\vartheta}) z_t = \boldsymbol{\Psi}_1(\boldsymbol{\vartheta}) z_{t-1} + \boldsymbol{\varepsilon}_t, \tag{3.23}$$

where z_t is the vector of endogenous variables and $\boldsymbol{\varepsilon}_t = \hat{g}_t$ denotes the vector of exogenous shocks, which in our case consists only of government spending shocks. The matrix $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ governs the endogenous dynamics, while $\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})^{-1}$ determines the contemporaneous impact of fiscal shocks. The dependence of these matrices on $\boldsymbol{\vartheta}$ allows us to trace how labor market institutions affect the propagation of fiscal policy.

We compute impulse response functions (IRFs) based on a calibrated version of the model, as described in Appendix A.2. Since the IRFs are continuous functions of $\boldsymbol{\vartheta}$, we evaluate them over a range of institutional configurations.

Figure 2: Fiscal Spending Multipliers and the LMIs ($\mathcal{M}(\boldsymbol{\vartheta})$).



Notes: The figure illustrates the sensitivity of the fiscal spending multipliers to variations in the three LMI parameters (UD–union density, BRR–(unemployment) benefit replacement rate and EPL–employment protection (legislation)). The multipliers are defined in Eq. (3.25) are presented for different horizons, with the contemporaneous multiplier ($\mathcal{P} = 1$) represented by a solid black square line, and the higher horizon multipliers ($\mathcal{P} \geq 2$) indicated by black dotted lines.

Following Crespo Cuaresma and Glocker (2023), we summarize the dynamic response of a variable x using the cumulative elasticity with respect to government spending over the horizon $\mathcal{P}$,

$$\mu_x(\vartheta_l) = \frac{\sum_{i=1}^{\mathcal{P}} \mathrm{IRF}_i^x(\vartheta_l)}{\sum_{j=1}^{\mathcal{P}} \mathrm{IRF}_j^{\hat{g}}(\vartheta_l)}, \qquad l \in \{1, 2, 3\}. \tag{3.24}$$

To facilitate comparison with the fiscal multiplier literature, we report the corresponding spending multiplier in levels,

$$\mathcal{M}_x(\vartheta_l) = \frac{\mu_x(\vartheta_l)}{\bar{g}/\bar{y}}, \qquad l \in \{1, 2, 3\}, \tag{3.25}$$

where $\bar{g}/\bar{y}$ denotes the steady-state government spending share in output (Ramey and Zubairy, 2018). In our sample, the average value of $\bar{g}/\bar{y}$ is approximately 0.20.

Figure 2 reports the multipliers for output ($\hat{y}_t$), employment ($\hat{n}_t$), and the real wage ($\hat{w}_t$) at different horizons. We distinguish between the contemporaneous response ($\mathcal{P} = 1$) and longer-horizon cumulative responses ($\mathcal{P} \geq 2$) to highlight dynamic adjustment. Each column shows how fiscal multipliers vary with the corresponding LMI parameter in $\boldsymbol{\vartheta}$.

3.7 *The Role of Differences in Labor Market Institutions*

We now examine how the three labor market institutions–the benefit replacement rate (φ), workers' bargaining power (η), and firing costs (ς)–shape fiscal spending multipliers.

Benefit replacement rate. We begin with the unemployment benefit replacement rate, φ. Lower replacement rates are associated with larger output and employment multipliers due to the mechanism that operates through wage bargaining. Since unemployment benefits raise workers' outside option, they increase the reservation wage of households and raise the equilibrium wage negotiated between firms and workers.

The reservation wages of firms and households are obtained by setting the marginal value of employment equal to zero, $F_t^n = 0$ and $H_t^n = 0$. Using Eq. (3.7) and (3.14), and setting $\tau = 0$ for simplicity, the firm's and household's reservation wages are given by

$$\overline{w}_t^F = mc_t mpl_t + \mathbb{E}_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], \tag{3.26}$$

$$\underline{w}_t^H = mrs_t + b_t^u - (1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}) \mathbb{E}_t \Lambda_{t,t+1} H_{t+1}^n. \tag{3.27}$$

The equilibrium wage is a weighted average of these reservation wages,

$$w_t = (1 - \eta) \overline{w}_t^F + \eta \underline{w}_t^H. \tag{3.28}$$

An increase in unemployment benefits raises the household reservation wage $\underline{w}_t^H$ by improving workers' outside option. This increases the negotiated wage and reduces the surplus from employment for firms. As a result, vacancy creation declines, and firms raise the productivity threshold $\tilde{a}_t$, thus increasing endogenous job separations. Both channels reduce the employment response to a fiscal expansion. Since output is proportional to employment, higher replacement rates also reduce the output multiplier. Therefore, in the model, economies with more generous unemployment benefits exhibit weaker employment and output responses to government spending shocks.[10]

[10] Albertini and Poirier (2015) emphasize that these effects may differ at the zero lower bound, where higher unemployment benefits can stimulate activity by lowering real interest rates through inflation expectations.

Workers' bargaining power. We next consider workers' bargaining power, η. Higher bargaining power increases the weight placed on the household reservation wage in Nash bargaining and therefore raises the equilibrium wage. This reduces the value of employment to firms and increases the endogenous job separation threshold $\tilde{a}_t$, leading to lower employment.

As a consequence, when government spending increases, employment and output rise by less in economies where workers have greater bargaining power. This mechanism is illustrated in the first column of Figure 2, which shows that fiscal multipliers decline as η increases.

Employment protection. Finally, we consider firing costs, ς, which capture employment protection. Unlike the previous institutions, firing costs do not directly affect wage bargaining. Instead, they influence firms' hiring and separation decisions.

Lower firing costs increase both job creation and job destruction by making employment adjustment less costly. This raises the responsiveness of employment to aggregate shocks. Consequently, employment reacts more strongly to fiscal spending shocks when firing costs are low, resulting in larger employment and output multipliers. Conversely, stricter employment protection dampens employment adjustment and reduces fiscal multipliers.[11]

4. The Empirical Evidence

This section empirically evaluates the theoretical prediction concerning how labor market institutions shape the transmission of fiscal policy shocks. In particular, we examine how the responses of output, employment, and wages to government spending shocks vary systematically with the benefit replacement rate, workers' bargaining power, and firing costs.

To do so, we estimate an interacted panel vector autoregressive (IPVAR) model, following Towbin and Weber (2013) and Sá, Towbin and Wieladek (2014), for a panel of advanced economies. The IPVAR framework allows the dynamic responses to fiscal shocks to depend explicitly on labor market institutions. This provides an empirical counterpart to the theoretical model, in which the propagation matrices $\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})$ and $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ depend on the institutional parameters $\boldsymbol{\vartheta}$.

[11] Appendix A.3 evaluates these mechanisms under alternative calibrations, and Appendix B considers several model extensions, including the absence of price indexation, real wage rigidities, limited asset market participation, firing costs accruing to the government, productivity-enhancing government spending, and complementarity between consumption and leisure. With the exception of the fiscal treatment of firing costs, these extensions do not alter the qualitative effects of labor market institutions on fiscal multipliers.

In the following subsections, we present the econometric specification, describe the data and identification strategy, and discuss the empirical results.

4.1 The Econometric Model

Let $\boldsymbol{y}_{it}$ denote the n-dimensional vector of endogenous variables for country $i = 1, \ldots, N$ at time t, and let $\boldsymbol{\vartheta}_{it}$ denote the d-dimensional vector of labor market institution indicators. The reduced-form IPVAR is given by

$$\boldsymbol{y}_{it} = \boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) + \sum_{j=1}^{p} \boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it})\boldsymbol{y}_{it-j} + \boldsymbol{u}_{it}, \quad \boldsymbol{u}_{it} \sim \mathcal{N}_n\big(\boldsymbol{0}, \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it})\big), \tag{4.1}$$

where $\boldsymbol{c}_i(\boldsymbol{\vartheta}_{it})$ is the intercept, $\boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it})$ are the autoregressive coefficient matrices, and $\boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it})$ is the covariance matrix. All coefficients are allowed to vary with the institutional variables $\boldsymbol{\vartheta}_{it}$, so that the propagation of shocks depends on the labor market environment. Appendix C provides full details of the parameterization.

For identification and interpretation, it is useful to consider the structural representation,

$$\tilde{\boldsymbol{\Psi}}_{i0}\boldsymbol{y}_{it} = \sum_{j=1}^{p} \tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it})\boldsymbol{y}_{it-j} + \boldsymbol{e}_{it}, \quad \boldsymbol{e}_{it} \sim \mathcal{N}_n(\boldsymbol{0}, \boldsymbol{I}), \tag{4.2}$$

where $\boldsymbol{e}_{it}$ denotes structural shocks. The structural coefficients $\tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it})$ capture how the transmission of shocks varies with labor market institutions.

The panel dimension allows us to estimate a common underlying economic structure while allowing for country-specific deviations. To this end, we employ a hierarchical prior which assumes that country-specific coefficients are drawn from a common Gaussian distribution, as in Jarociński (2010). Such a prior implies that

$$\tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it}) \sim \mathcal{N}\big(\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t), \boldsymbol{V}_j\big), \quad j = 1, \ldots, p, \tag{4.3}$$

where $\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t)$ denotes the cross-country mean and $\boldsymbol{V}_j$ governs cross-country dispersion. We employ global-local shrinkage priors to regularize estimation and avoid overfitting (Griffin and Brown, 2010; Huber and Feldkircher, 2019). Details of the prior specification are provided in Appendix C.

To connect the empirical specification to the theoretical model, we assume that the structural coefficients vary smoothly with labor market institutions. Specifically, we approximate the coefficient matrices using a first-order Taylor expansion,

$$\mathbf{\Psi}_j(\boldsymbol{\vartheta}_t) = \bar{\mathbf{\Psi}}_j + \sum_{l=1}^{d} \frac{\partial \mathbf{\Psi}_j}{\partial \vartheta_{lt}} (\vartheta_{lt} - \bar{\vartheta}_l), \quad j = 0, \dots, p. \tag{4.4}$$

In practice, we evaluate impulse responses by varying one institutional variable at a time while holding the others fixed at their median values. This approach isolates the contribution of each labor market institution to fiscal transmission.

Two potential limitations of this approach merit discussion. First, labor market institutions may be endogenous to macroeconomic conditions. However, their evolution appears to reflect long-run structural changes rather than cyclical dynamics (see Appendix D1). Consistent with this interpretation, impulse response analysis shows that labor market institutions do not respond significantly to cyclical shocks when estimating a standard panel VAR model that includes the institutional variables and examining their response to macroeconomic shocks.

Second, the specification assumes that coefficient variation with respect to institutions is linear. While more flexible specifications are possible, the available sample size limits the feasibility of estimating highly nonlinear models without overfitting.

4.2 Data and Specification

We use quarterly data for 16 OECD countries over the period 1960Q1–2020Q4 to estimate our IPVAR model.[12] The empirical specification follows the theoretical framework and focuses on government spending, output, employment, and wages.

The vector of endogenous variables is given by $\boldsymbol{y}_{it} = (g_{it}, x_{it}, er_{it}, \omega_{it})'$, where g_{it} denotes the growth rate of real government consumption per capita, x_{it} denotes the growth rate of real GDP per capita, er_{it} denotes the first difference of the employment rate, and ω_{it} denotes the growth rate of real wages.[13] Detailed definitions and data sources are provided in Appendix D.

Labor market institutions enter the model as interaction variables. Specifically, we define $\boldsymbol{\vartheta}_{it} = (\eta_{it}, \varphi_{it}, \varsigma_{it})'$, where η_{it} denotes union density, φ_{it} denotes the unemployment benefit replacement rate,

[12] The sample includes Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Italy, Japan, the Netherlands, Portugal, Spain, Sweden, and the United States.

[13] Following Brückner and Pappa (2012), all variables are expressed in per capita terms. As a robustness check, we also apply the transformation proposed by Gordon and Krenn (2010) and divide each variable by its trend estimated using the filter of Hamilton (2018). Consistent with Ramey (2016), this transformation produces impulse responses similar to those obtained using log-level specifications. Results are reported in Appendix E.2.

and ς_{it} denotes the employment protection index. These data are obtained from the CEP-OECD institutions database. Since institutional data are available annually, we construct quarterly series by assigning each annual observation to all quarters within the corresponding year, thus reflecting the slow evolution of labor market institutions over time.

Prior to estimation, all interaction variables are standardized to facilitate interpretation and comparability across countries. The IPVAR model includes all three institutional variables simultaneously. When evaluating impulse responses with respect to a particular institutional variable ϑ_l, we hold the remaining institutional variables fixed at their median values.

The lag length is set to $p = 1$ based on the Bayesian information criterion.[14] Estimation is based on 20,000 posterior draws, discarding the first 10,000 draws as burn-in.

4.3 Shock Identification

We identify fiscal spending shocks by imposing a recursive identification scheme based on the Cholesky decomposition of the reduced-form IPVAR shocks. We follow Blanchard and Perotti (2002) and assume that fiscal spending does not react contemporaneously to shocks arising from GDP or labor market variables in the system. These three variables are assumed to respond within the same quarter to the fiscal spending shock. Such a recursive structure is the most conventional strategy used to identify fiscal spending shocks in the established structural VAR literature (see, for instance, the discussion in Čapek and Crespo Cuaresma, 2020). We utilize this particular approach for two reasons. First, this approach is in line with recent studies that use panel VAR or country VAR methods to analyze the effects of fiscal policy (see, inter alia, Beetsma and Giuliodori, 2011; Bénétrix and Lane, 2013; Ilzetzki, Mendoza and Végh, 2013; Huidrom et al., 2020). Second, alternative identification approaches are infeasible in the context of a large cross-country analysis.[15]

One potential drawback in the context of a recursive approach to identification concerns the issue of fiscal foresight. Economic agents constantly receive information and update their expectations regarding news on fiscal policy issues. As econometricians, we may only observe a smaller information set. This misalignment between the information sets of the economic agents and the econometrician can generate equilibria with

[14] Appendix E.2 shows that the results are robust to alternative lag lengths, including two and four lags.

[15] Alternative approaches can be summarized in three distinct groups: (i) event-study approaches based on defense spending changes (Ramey and Shapiro, 1998; Ramey, 2011), (ii) sign-restrictions (Mountford and Uhlig, 2009), or (iii) narrative approaches (Romer and Romer, 2010; Guajardo, Leigh and Pescatori, 2014). These approaches are not feasible because they rely on additional data that are not available (e.g., detailed institutional information on fiscal spending plans) or are not practical for a large panel of countries. An interesting alternative is provided by the approach in Miyamoto, Nguyen and Sheremirov (2019), which uses military spending data in a large panel of countries but only on a yearly frequency.

a non-fundamental moving average representation (Ramey, 2011; Leeper, Walker and Yang, 2013; Ellahie and Ricco, 2017). To resolve this issue, the information set in the VAR is enlarged to contain a variable proxying for agents' expectations. These expectations are not available for the full sample of countries and time span considered here. Therefore, we show that our results are robust to this concern in a smaller setting in Appendix E.3.[16]

4.4 *The Effect on Fiscal Spending Effectiveness*

This subsection examines empirically how labor market institutions affect fiscal spending multipliers. Following the theoretical analysis, we report both the impact multiplier ($\mathcal{P} = 0$) and the one-year cumulative multiplier ($\mathcal{P} = 4$). Figure 3 shows how these multipliers vary with union density (η), the benefit replacement rate (φ), and employment protection (ς). The solid black and dashed red lines denote the median impact and one-year multipliers, respectively, and shaded areas indicate 80% credible sets. The horizontal axis spans the 10th to 90th percentiles of each institutional variable.

Fiscal multipliers for output decline markedly with union density. At low levels of union density, the impact elasticity is approximately 0.3, corresponding to a level multiplier of around 1.5 when scaled by the government spending share in output. As union density increases, the multiplier declines steadily and approaches zero at the upper end of the distribution. This decline is economically large and statistically significant.

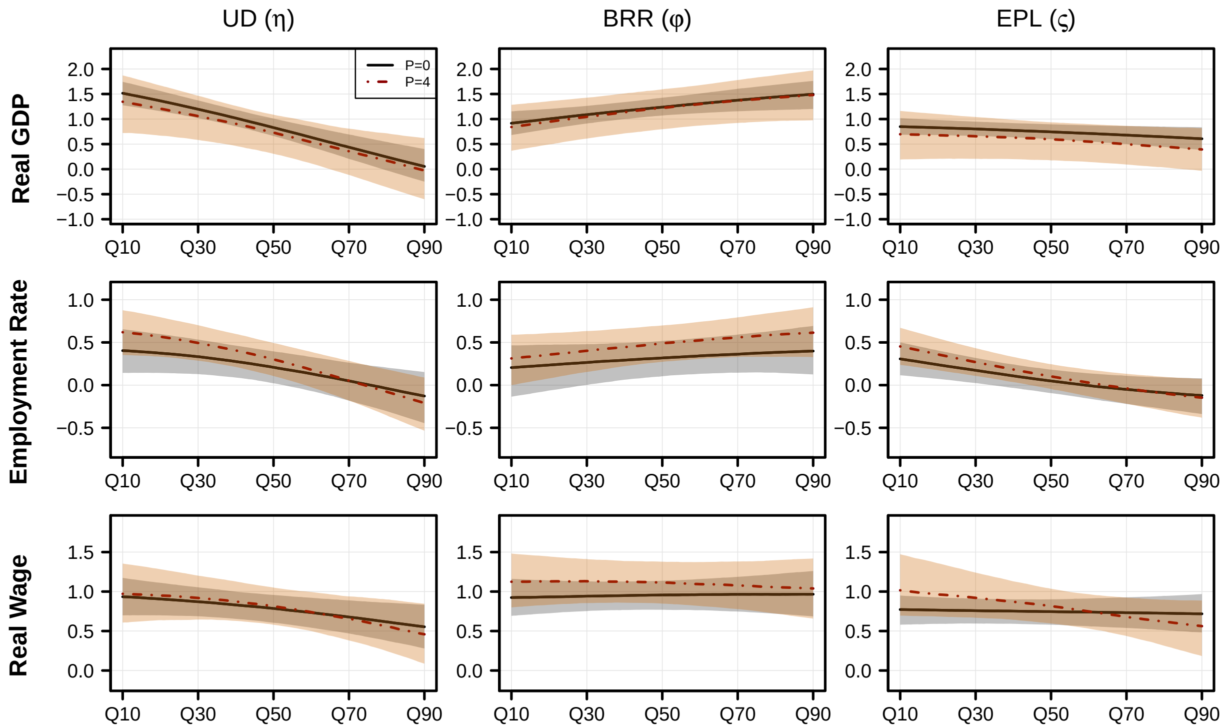
In contrast, employment protection has a weaker effect on output multipliers. While there is some decline as employment protection increases, the effect is modest and not precisely estimated. For the benefit replacement rate, we do not observe a decline in output multipliers. Instead, the estimates suggest a slight increase, although the differences are not statistically significant.[17]

The employment multiplier is consistently positive but declines with union density and employment protection. The impact employment multiplier decreases from approximately 0.60 at low levels of these institutional variables to values close to zero at high levels. These effects (which are statistically significant

[16] We follow the suggestions in Born, Juessen and Müller (2013), Born, Müller and Pfeifer (2020), or Ilori, Paez-Farrell and Thoenissen (2022) and use two different sets of government spending forecasts. To control for fiscal foresight, we rely on data from professional forecasters at Oxford Economics and the OECD. We restrict our sample to the G7 countries in the cross-section, starting only in the mid 1980s (OECD) or late 1990s (Oxford Economics). See Appendix E.3 for further details.

[17] Appendix E.1 reports impulse responses evaluated at different points in the institutional distribution and examines the posterior distribution of their differences.

Figure 3: Fiscal Multipliers.



Notes: The figure show the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is the unemployment benefit replacement rate, and ς is employment protection legislation). The y-axis gives the size of the multiplier while the x-axis runs from the 10th to the 90th quantile in terms of the respective LMI. The multipliers are defined in Eq. (3.25) and shown for different horizons: on impact ($\mathcal{P}$ = 0, solid black line) and one-year ($\mathcal{P}$ = 4, dashed-dotted red line) multiplier. The lines denote the median and the colored area refers to the 80% credible set.

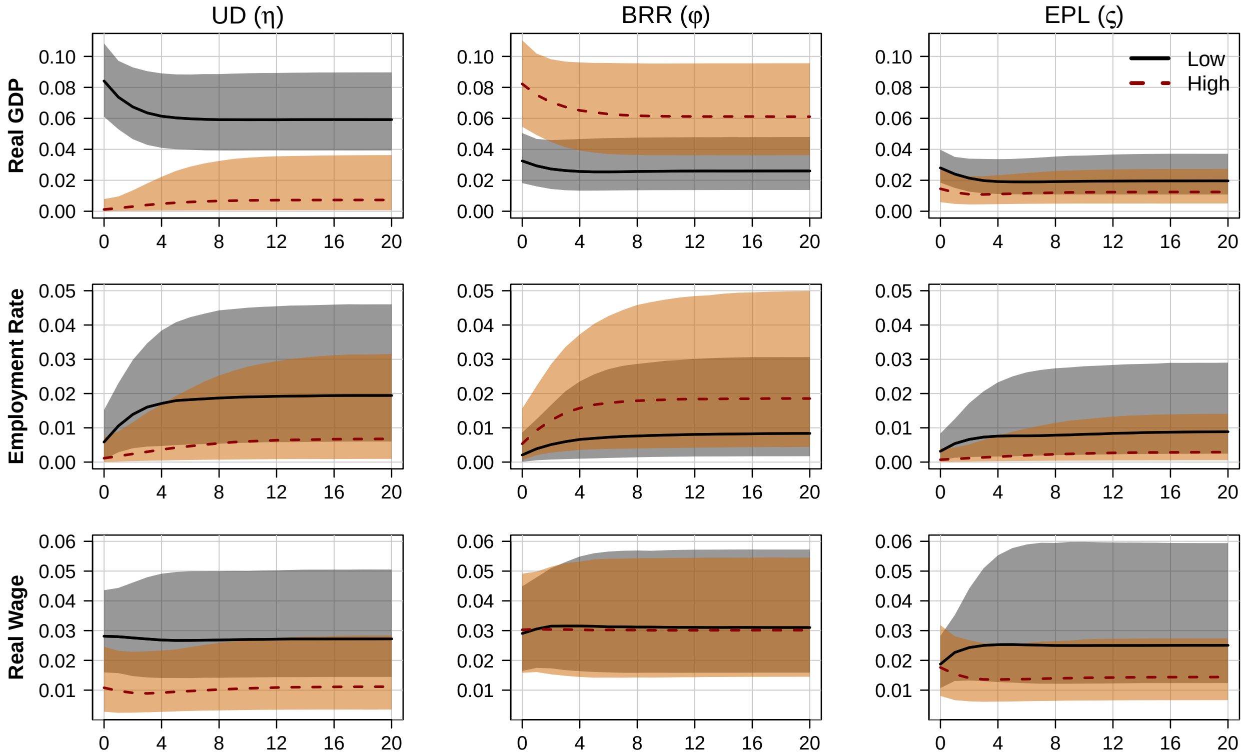
from a frequentist perspective) indicate that stricter labor market institutions dampen the employment response to fiscal spending shocks.

In contrast, the employment multiplier does not exhibit a systematic decline with the benefit replacement rate. Although point estimates suggest a modest increase, the associated uncertainty is substantial. One-year multipliers are generally larger than impact multipliers, reflecting gradual adjustment in labor market outcomes following fiscal shocks.

Fiscal multipliers for real wages are smaller and estimated less precisely. There is some evidence of declining multipliers with union density and employment protection at longer horizons, but credible intervals are wide. In contrast, real wage multipliers show little systematic variation with the benefit replacement rate.

Overall, the empirical results support the theoretical prediction that labor market institutions shape fiscal transmission. In particular, union density and employment protection significantly attenuate output

Figure 4: Forecast Error Variance Decomposition.



Notes: The sub-figures show the sensitivity of the explained forecast error variance to changes in the structural parameters (η is union density, φ is unemployment benefit replacement rate, and ς is employment protection). The y-axis gives the share of explained forecast error variance while the x-axis is the forecast horizon and runs up to 5 years (=20 quarters). The FEVD is shown for the respective LMI at the 10th quantile (low, black dashed line) and 90th quantile (high, dashed red line). Confidence bounds refer to the 10/90 quantile of the posterior distribution.

and employment multipliers. These findings closely mirror the predictions of the DSGE model shown in Figure 2.

One difference between the empirical and theoretical results concerns the benefit replacement rate. While the theoretical model predicts weaker multipliers with more generous unemployment benefits, the empirical estimates do not show such a decline. One possible explanation is that unemployment insurance reduces precautionary savings and strengthens aggregate demand (Challe, 2020; McKay and Reis, 2021; Kekre, 2023; Gorn and Trigari, 2024). This demand channel is not fully captured in the baseline theoretical framework. Another difference concerns the dynamic adjustment of employment, which appears less persistent in the empirical estimates than in the theoretical model.

Figure 4 reports the forecast error variance decomposition. Fiscal spending shocks explain a larger share of output fluctuations in economies with more flexible labor markets. For example, fiscal shocks account for up to 8% of output variance at longer horizons when union density is low, but substantially less when union density is high. Similar but weaker patterns are observed for employment protection. In contrast, the benefit replacement rate does not reduce the contribution of fiscal shocks to output variability. Consistent with the impulse response analysis, differences across institutional regimes are small for employment and real wages.

Our results are consistent with existing estimates of fiscal multipliers (Ramey, 2019; Ilzetzki, Mendoza and Végh, 2013) and confirm that employment responds less strongly than output to fiscal shocks (Monacelli, Perotti and Trigari, 2010). Our main contribution is to show that fiscal multipliers depend systematically on labor market institutions. Additional results and robustness checks are reported in Appendix E. These include results based on controlling for fiscal foresight, the use of alternative labor market indicators, and an assessment of cross-country heterogeneity.

4.5 The Effect on Macroeconomic Volatility

While stricter LMIs may reduce the effectiveness of discretionary fiscal policy, they may also dampen macroeconomic volatility. The LMIs we consider capture structural features of the labor market and can, to some extent, act as automatic stabilizers. This is particularly evident for the benefit replacement rate: more generous unemployment insurance smooths household income and provides insurance against unemployment risk. Employment protection and unionization can also affect cyclical adjustment by altering job separation, hiring, and wage dynamics. This raises the question of whether LMIs and discretionary fiscal policy act as substitutes in stabilizing macroeconomic outcomes.

To address this question, we use the IPVAR to quantify changes in unconditional macroeconomic volatility as a function of LMIs. Specifically, we compute the unconditional covariance matrix of the endogenous variables implied by the estimated IPVAR, conditional on different levels of our institutional variables.[18]

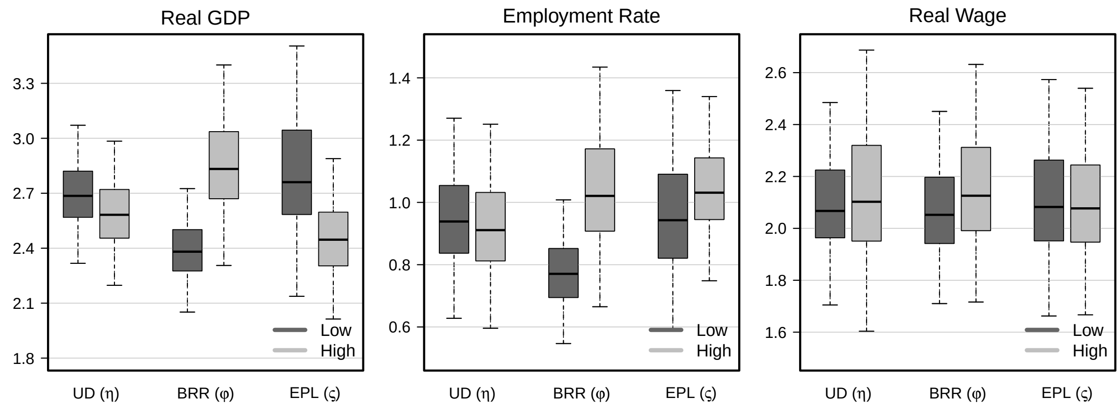
Let $\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)$ denote the unconditional covariance matrix of the IPVAR state vector in companion form. Using standard results for stable VARs (see, e.g., Hamilton (1994)), the unconditional covariance satisfies

$$\operatorname{vec}(\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)) = (\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}_t) \otimes \boldsymbol{F}(\boldsymbol{\vartheta}_t))^{-1} \operatorname{vec}(\boldsymbol{Q}(\boldsymbol{\vartheta}_t)) \,, \tag{4.5}$$

where $\boldsymbol{I}$ is the identity matrix of dimension $K^2 = (np)^2$, $\boldsymbol{F}(\boldsymbol{\vartheta}_t)$ is the $K \times K$ companion matrix implied by the VAR coefficient matrices, and $\boldsymbol{Q}(\boldsymbol{\vartheta}_t)$ is the $K \times K$ covariance matrix of the VAR innovations in companion form. We do not impose a pooling prior on the country-specific covariance matrices. For the variance calculations, we therefore work with a common-mean covariance matrix obtained by averaging the country-specific estimates, $\overline{\boldsymbol{\Sigma}}(\boldsymbol{\vartheta}) = N^{-1} \sum_{i=1}^{N} \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it})$. Eq. (4.5) makes explicit that unconditional volatilities depend

[18] For an analogous analysis in the theoretical model; see Appendix A.4. In the DSGE model, we consider volatilities conditional on specific shocks (government spending and technology shocks), whereas the IPVAR-based volatilities are unconditional.

Figure 5: Macroeconomic Volatilities Along LMIs.



Notes: Each sub-figure reports the implied unconditional standard deviation of the respective macroeconomic variable under a low (10th percentile, dark gray) and a high (90th percentile, light gray) value of the relevant labor market institution, holding the remaining LMIs fixed at their median values. The institutions are union density (UD, η), the unemployment benefit replacement rate (BRR, φ), and employment protection legislation (EPL, ς).

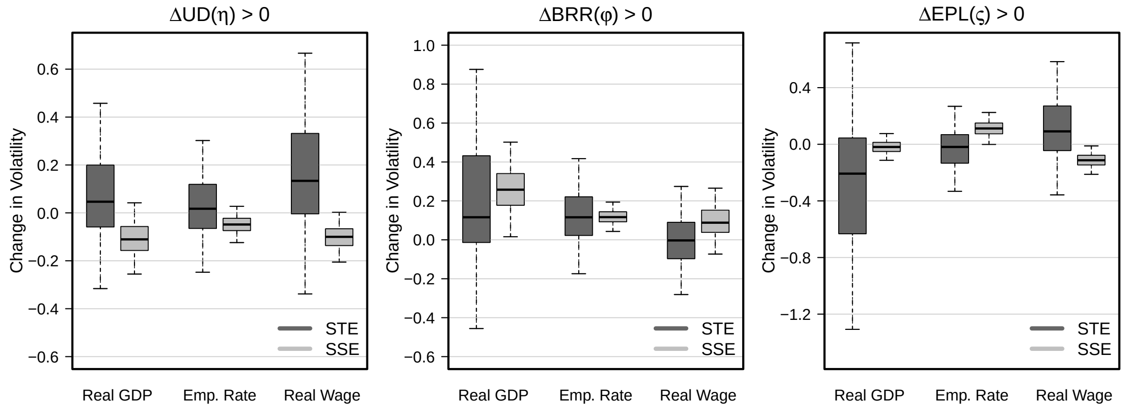
on institutional characteristics through both the propagation matrix $\boldsymbol{F}(\boldsymbol{\vartheta}_t)$ and the innovation covariance $\boldsymbol{Q}(\boldsymbol{\vartheta}_t)$.

Figure 5 summarizes the implied volatilities by comparing low (10th percentile) and high (90th percentile) values of each LMI. Two patterns stand out. First, higher union density and stronger employment protection tend to reduce output volatility, although the effect is more pronounced for employment protection. Second, a higher benefit replacement rate is associated with higher output volatility and also increases the volatility of the employment rate. Real wage volatility exhibits comparatively small differences across institutional regimes. These results refer to unconditional volatilities and are consistent with the general message in Zanetti (2011). At the same time, the effect of LMIs on volatility depends on the source of shocks, as discussed in Appendix A.4.

4.6 How Do Labor Market Institutions Affect Macroeconomic Volatility?

We next investigate through which channels LMIs affect (unconditional) macroeconomic volatility. Eq. (4.5) shows that institutional variables can influence volatility (i) by changing the propagation of shocks through the VAR dynamics, or (ii) by changing the size and covariance structure of reduced-form innovations. We refer to these components as the *shock transmission effect* (STE) and the *shock size effect* (SSE), respectively.

Figure 6: Change in Macroeconomic Volatilities Along LMIs.



Notes: Each sub-figure shows the change in the unconditional standard deviation of the respective macroeconomic variable when moving from the low (10th percentile) to the high (90th percentile) regime of the respective labor market institution, holding the remaining LMIs fixed at their median values. Dark gray bars represent the shock transmission effect (STE) and light gray bars represent the shock size effect (SSE), as defined in Eq. (4.7). The institutions are union density (UD, η), the unemployment benefit replacement rate (BRR, φ), and employment protection legislation (EPL, ς).

Let $\omega_{kk}(\boldsymbol{\vartheta}_t)$ denote the kth diagonal element of $\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)$, i.e., the unconditional variance of the kth variable. Define

$$\tilde{\boldsymbol{F}}(\boldsymbol{\vartheta}_t) \equiv (\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}_t) \otimes \boldsymbol{F}(\boldsymbol{\vartheta}_t))^{-1}, \qquad \tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}_t) \equiv \operatorname{vec}(\boldsymbol{Q}(\boldsymbol{\vartheta}_t)). \tag{4.6}$$

Using $\operatorname{vec}(\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)) = \tilde{\boldsymbol{F}}(\boldsymbol{\vartheta}_t)\tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}_t)$, the partial effect of an institutional variable ϑ_{lt} on ω_{kk} can be written as

$$\frac{\partial \omega_{kk}(\boldsymbol{\vartheta}_t)}{\partial \vartheta_{lt}} = \sum_j \left(\underbrace{\tilde{q}_j(\boldsymbol{\vartheta}_t) \frac{\partial \tilde{f}_{kj}(\boldsymbol{\vartheta}_t)}{\partial \vartheta_{lt}}}_{\text{STE}} + \underbrace{\tilde{f}_{kj}(\boldsymbol{\vartheta}_t) \frac{\partial \tilde{q}_j(\boldsymbol{\vartheta}_t)}{\partial \vartheta_{lt}}}_{\text{SSE}} \right), \tag{4.7}$$

where $\tilde{f}_{kj}(\boldsymbol{\vartheta}_t)$ denotes element (k, j) of $\tilde{\boldsymbol{F}}(\boldsymbol{\vartheta}_t)$ and $\tilde{q}_j(\boldsymbol{\vartheta}_t)$ denotes element j of $\tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}_t)$. This decomposition separates changes in volatility due to propagation (STE) from changes due to shock variances and covariances (SSE).

Figure 6 reports the change in volatility when moving from low to high values of each institutional variable and decomposes this change into STE and SSE. For union density, the STE tends to increase volatility, whereas the SSE tends to reduce it. These effects partially offset each other, implying a small net effect on unconditional volatility, particularly for real wages. For the benefit replacement rate, both STE and SSE increase volatility, most clearly for output and, to a lesser extent, employment. For employment protection, the reduction in output volatility is primarily driven by the STE, while the SSE contribution is close to zero; effects on employment and wages are comparatively small.

Table 1: Estimated model parameters.

	UD (η)		BRR (φ)		EPL (ς)	
	Low ($\eta = 0.1$)	High ($\eta = 0.7$)	Low ($\varphi = 0.3$)	High ($\varphi = 0.6$)	Low ($\varsigma = 0.1$)	High ($\varsigma = 0.6$)
ϱ_g	0.42	0.45	0.41	0.44	0.41	0.43
ξ	0.56	0.58	**0.56**	**0.91**	0.56	0.55
γ_p	0.02	0.03	**0.00**	**0.52**	0.04	0.03
γ	**0.56**	**0.88**	0.58	0.69	0.71	0.76
μ_a	**-0.97**	**0.66**	1.78	1.83	**1.30**	**-1.48**
σ_a	2.88	3.03	2.32	3.08	2.39	2.99

Notes: Structural parameters: ϱ_g is the autoregressive coefficient of the AR(1)-government consumption spending shock, ξ captures the degree of price stickiness, γ_p denotes the share of inflation indexed prices, μ_a and σ_a are the first and second moments of the distribution of the idiosyncratic job productivity ($\tilde{a}_t$). The values in parentheses in the rows for the Low and High LMIs represent the specific LMI values used in the matching process. All other parameters are fixed at their baseline calibration.

Overall, both channels matter for how institutions shape volatility. The decomposition also shows that STE and SSE can operate in opposite directions, which helps explain why net effects are sometimes small despite sizable underlying components.[19]

4.7 Accounting for the Evidence

We next explore which structural features of the theoretical model account for the empirical differences across institutional regimes. To this end, we estimate a subset of structural parameters by matching IRFs from the DSGE model to those estimated from the IPVAR.

Let $\widehat{\text{IRF}}$ denote the empirical IRFs obtained from the IPVAR (for the common-mean dynamics).[20] Let $\text{IRF}(\boldsymbol{\theta})$ denote the model-implied IRFs from the DSGE model. We fix all structural parameters at their baseline calibration (Table A1), except for the persistence of the government spending shock (ϱ_g), price stickiness (ξ), the degree of inflation indexation (γ_p), the matching elasticity with respect to unemployment (γ), and the first two moments of the idiosyncratic productivity distribution (μ_a, σ_a). Idiosyncratic productivity is i.i.d. log-normally distributed with c.d.f. $F(\cdot)$. We estimate the moments of $\ln(\tilde{a})$, i.e., $\mu_a = E[\ln(\tilde{a})]$ and $\sigma_a = \sqrt{\text{Var}[\ln(\tilde{a})]}$. These parameters are collected in $\boldsymbol{\theta}$ and are selected because they play a central role in shaping fiscal multipliers and labor market adjustment (Galí, López-Salido and Valles, 2007; Dupaigne and Fève, 2016).

[19] These results are robust to replacing the employment rate with the unemployment rate; see Appendix E.4.

[20] Throughout, we focus on impulse responses of the *common mean* in the hierarchical IPVAR, which pools information across countries while allowing for idiosyncratic deviations. The corresponding IRFs are reported in Appendix E.1.

We match the first 25 periods of the responses of government spending ($\hat{g}_t$), output ($\hat{y}_t$), employment ($\hat{n}_t$), and real wages ($\hat{w}_t$). The parameter vector is obtained by minimizing the distance between empirical and theoretical IRFs

$$\hat{\boldsymbol{\theta}} = \arg\min_{\boldsymbol{\theta}} \left\| \widehat{\mathrm{IRF}} - \mathrm{IRF}(\boldsymbol{\theta}) \right\|. \tag{4.8}$$

We restrict attention to parameter values that imply equilibrium determinacy. We perform this exercise separately for each institutional dimension and for low and high values of the corresponding LMIs (10th and 90th percentiles), holding the remaining LMIs fixed at their median values.

The resulting estimates are reported in Table 1. The persistence of the government spending shock (ϱ_g) is moderate and varies little across institutional regimes, and we find similarly limited variation in the dispersion of idiosyncratic productivity (σ_a). In contrast, several parameters show economically meaningful changes across regimes. For union density, the estimates point to substantial changes in the matching elasticity γ and the mean of the idiosyncratic productivity distribution μ_a, both of which shape job creation and separation dynamics. For the benefit replacement rate, the matching results primarily indicate stronger nominal rigidities, reflected in changes in price stickiness and inflation indexation. Finally, employment protection mainly affects the mean of idiosyncratic job destruction, consistent with the interpretation that stricter EPL lowers the propensity for separations. These structural differences help rationalize the cross-regime variation in empirical fiscal responses.

5. Concluding Remarks

This paper studies how labor market institutions shape the transmission of fiscal spending shocks and macroeconomic volatility. We combine a structural New Keynesian model with search and matching frictions and an interacted panel vector autoregressive (IPVAR) model estimated for 16 OECD economies. This integrated theoretical and empirical framework allows us to quantify how institutional differences affect fiscal multipliers and the volatility of macroeconomic outcomes.

Our first main finding is that labor market institutions significantly affect the transmission of fiscal policy shocks. Higher union density and stronger employment protection reduce the output and employment responses to government spending shocks, thereby attenuating fiscal multipliers. The benefit replacement rate exhibits weaker and less precisely estimated effects on multipliers. These empirical patterns closely align with the predictions of the theoretical model and can be traced to institutional effects on wage bargaining, job separation, and matching dynamics.

Our second main finding is that labor market institutions also shape macroeconomic volatility. Higher union density and stronger employment protection are associated with lower output volatility, while the benefit replacement rate is associated with higher volatility in output and employment. Using a decomposition of unconditional variances, we show that these effects arise through both changes in shock transmission and changes in the size and covariance structure of reduced-form innovations.

Taken together, these results highlight a systematic relationship between labor market institutions, fiscal transmission, and macroeconomic stability. Labor market rigidities weaken the short-run effectiveness of discretionary fiscal policy but can also reduce output volatility. This implies that structural labor market characteristics influence both the need for and the effectiveness of countercyclical fiscal policy.

More broadly, our findings underscore the importance of accounting for institutional heterogeneity when evaluating fiscal policy effectiveness across countries. Differences in labor market institutions help explain cross-country variation in fiscal multipliers and macroeconomic volatility, and therefore constitute a key structural determinant of fiscal transmission.

Declaration of Interest

The authors declare that they have no conflict of interest.

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Online Appendix: Labor Market Institutions, Fiscal Multipliers, and Macroeconomic Volatility*

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Appendix Material

This appendix provides additional material that complements the analysis in the main text. Appendix A provides further details on the theoretical model, while we present extensions to the theoretical model in Appendix B. Appendix C provides further details on the econometrical model. Data sources, availability, and transformations are listed in Appendix D. We present a number of further empirical results in Appendix E.

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A. Further Details on the Theoretical Model

This section provides further details on the solution of the baseline model. The model extensions considered also rest upon the solution procedure outlined here.

A.1 Equilibrium Equations

The following provides an overview as regards the equations that characterize the equilibrium. The particular functional form of the instantaneous utility function is given by: $u(c,n) = \frac{c^{1-\sigma}(1+(\sigma-1)\phi n)^{\sigma}-1}{1-\sigma}$.

Production

- $y_t = \bar{A} n_t A(\tilde{a}_t)$, with $A(\tilde{a}_t) = \int_{\tilde{a}_t}^{\infty} \frac{a}{1-F(\tilde{a}_t)} dF(a)$
- $\frac{\kappa}{q_t} = \mathbb{E}_t \Lambda_{t,t+1} \left[(1-\varrho(\tilde{a}_{t+1}))F^n_{t+1} - b^s_{t+1}(1-\bar{\varrho})F(\tilde{a}_{t+1}) \right]$
- $F^n_t = mc_t mpl_t - w_t + \mathbb{E}_t \Lambda_{t,t+1} \left[(1-\varrho(\tilde{a}_{t+1}))F^n_{t+1} - b^s_{t+1}(1-\bar{\varrho})F(\tilde{a}_{t+1}) \right]$
- $mc_t mpl_t = w_t - b^s_t - \frac{\kappa}{q_t}$
- $(1+\gamma_p \xi \beta)\pi_t = \beta \mathbb{E}_t \pi_{t+1} + \gamma_p \pi_{t-1} + \frac{(1-\xi\beta)(1-\xi)}{\xi} \hat{mc}_t$

Households

- $1 = \mathbb{E}_t \left[\Lambda_{t,t+1} \right] R_t$ with $\Lambda_{t,t+1} = \beta \lambda_{t+1}/\lambda_t$ and $\lambda = u_{c,t}$
- $mrs_t = -u_{n,t}/\lambda_t$

Labor market and Nash wage

- $n_t = (1-\varrho(\tilde{a}_t))(n_{t-1} + q_{t-1}v_{t-1})$ with $\varrho(\tilde{a}_t) = \bar{\varrho} + (1-\bar{\varrho})F(\tilde{a}_t)$
- $q_t = m_t/v_t$, $p_t = m_t/u_t$ with $u_t = 1 - n_t$ and $\theta_t = v_t/u_t$
- $m_t = \bar{m} u_t^{\gamma} v_t^{1-\gamma}$
- $w_t = (1-\eta)\frac{mrs_t + b^u_t}{1-\tau} + \eta \left(mc_t mpl_t + \mathbb{E}_t \Lambda_{t,t+1} \left[\kappa \theta_{t+1} - b^s_{t+1}(1-\bar{\varrho})F(\tilde{a}_{t+1}) \right] \right)$

Constraints and Policies

- $\tau w_t n_t + B_t = R_{t-1}B_{t-1} + b^u_t u_t + \bar{T}^s_t + g_t$

- $y_t = c_t + g_t + \kappa v_t + F(\tilde{a}_t)(1-\bar{\varrho})(n_{t-1} + q_{t-1}v_{t-1})b_t^s$

- Monetary policy: $i_t = \rho_i i_{t-1} + (1-\rho_i)(\phi_\pi \pi_t + \phi_y \hat{y}_t)$

- Fiscal policy: $\hat{g}_t \sim \text{AR}(1)$, $b_t^s = \bar{\varsigma} + \varsigma w_{t-1}$, $b_t^u = \bar{\varphi} + \varphi w_{t-1}$ and $T_t^s = \bar{T}^s + \varphi_{T^s} B_t$

where $mpl_t = y_t/n_t$, $\hat{mc}_t$ indicates the relative deviation of (real) marginal costs from the steady state level $\bar{mc} = \frac{\varepsilon}{\varepsilon-1}$, and a is log-normally distributed of which F is the c.d.f. The expression for the total surplus is finally given by: $S_t = F_t^n + H_t^n$ where $H_t^n = (1-\tau)w_t - b_t^u - mrs_t + (1-\varrho(\tilde{a}_{t+1}) - p_{t+1})\mathbb{E}_t\left[\Lambda_{t,t+1}H_{t+1}^n\right]$.

A.2 Calibration and the Steady State

We compute the steady state for the purpose of simulating the model. Variables without a time subscript denote steady state values. We start by considering an ex-ante calibration of the probability of an unemployed person finding a job (p_t), the labor market tightness (θ_t), and the ratio between the marginal rate of substitution between consumption and labor on the side of the households and the marginal product of labor on the side of the firms ($\zeta_t = mrs_t/mpl_t$). Additionally, we calibrate the steady-state separation rate $\varrho(\tilde{a})$ and, following the argument in den Haan, Ramey and Watson (2000), we also calibrate the exogenous job destruction rate $\bar{\varrho}$. The idiosyncratic productivity $\tilde{a}$ is assumed to be i.i.d. log-normally distributed with c.d.f. F of which we calibrate the first and second moments ($\mu_a = E[\ln(\tilde{a})]$ and $\sigma_a = \sqrt{Var[\ln(\tilde{a})]}$, where $\mu_a \in \mathbb{R}$ and $\sigma_a \in \mathbb{R}^+$). Given steady state values for p_t, θ_t, ζ_t and values for the structural parameters outlined in Table A1 in Section A.2, we then compute values for κ and $\bar{m}$ and the remaining variables of the model.

In particular, from $\bar{m} = p/\theta^{1-\gamma}$ we get the probability of a vacancy being filled $q = \bar{m}\theta^{-\gamma}$, the number of employed and unemployed persons $n = (1-\varrho(\tilde{a}))p/((1-\varrho(\tilde{a}))p + \varrho(\tilde{a}))$ and $u = 1-n$, the number of vacancies posted $v = \theta \cdot u$, and the number of matches $m = \bar{m}u^\gamma v^{1-\gamma}$ in the steady state. Given the assumptions on the steady-state separation rate $\varrho(\tilde{a})$ and the exogenous job destruction rate $\bar{\varrho}$, the endogenous separation rate is then given by $F(\tilde{a}) = \varrho^n = (\varrho(\tilde{a}) - \bar{\varrho})/(1-\bar{\varrho})$. From this we can obtain the steady-state threshold for the idiosyncratic productivity: $\tilde{a} = F^{-1}(\varrho^n)$, which allows us to compute the conditional expectation $A(\tilde{a}) = \int_{\tilde{a}}^{\infty} \frac{a}{1-F(\tilde{a})} dF(a)$. Given employment n, we can then compute the level of output in the steady state $y = \bar{A} \cdot n \cdot A(\tilde{a})$, the marginal product of labor $mpl = y/n$ and the level of government spending $g = g_y y$.

Using Eq. (3.7), Eq. (3.8) and Eq. (3.16) and the marginal product of labor, the vacancy posting cost parameter κ can be computed by $\kappa = b_1 \cdot mpl$ where b_1 is a parameter composed of the various structural model parameters (φ, η, β, τ, $\bar{\varrho}$, ζ, ...). Given κ and the marginal rate of substitution ($mrs = \zeta \cdot mpl$), the steady state real wage rate is then given by $w = b_1 \cdot mpl + b_2\kappa$. Finally, using Eq. (3.9), we calibrate $\bar{\varsigma}$ such that $A(\tilde{a}) = (w - b^s - \kappa/q)/\bar{A}$.

Household consumption is given by $c = y - g - \kappa v$. Using the steady state values for consumption and labor, the marginal utilities of consumption and labor and the parameter $\phi = mrs/(\sigma c - mrs \cdot (\sigma - 1)n)$ can then be computed. Finally, assuming net-government debt to be zero in the steady state ($B = 0$), the amount of lump-sum transfers $\bar{T}^s$ is then given by $\bar{T}^s = \tau wn - \varphi w(1 - n) - g$. If $\bar{T}^s < 0$, it can be interpreted as lump-sum taxes and as lump-sum subsidies if $\bar{T}^s > 0$.

Our benchmark calibration is summarized in Table A1. Given that our focus is on the role of the LMIs in the transmission of fiscal spending shocks, we do not calibrate our model to a particular economy. Regarding the parameters specific to the labor market institutions (LMI) in the steady state, we set $\eta = 0.5$ (UD), $\varphi = 0.5$ (BRR), and $\varsigma = 0.3$ (EPL), where we used empirical evidence provided in OECD (2007) and Wesselbaum (2020). The latter two values imply that the variable for unemployment benefits (b_t^u) and the one capturing firing costs (b_t^s) are connected to the wage in the baseline setup. The chosen value of the union density parameter (UD), η, ensures an equal bargaining weight for workers and firms in the wage negotiation process. We follow Krause and Lubik (2007), Christoffel, Kuester and Linzert (2009) and Zanetti (2011) among others for the choice of various values of the structural parameters.[1] The complementarity coefficient σ in the households' instantaneous utility function $u(c, n)$ is set to 1, which corresponds to the separable utility case. We also need to calibrate the shock process, for which we assume that the logarithm of fiscal spending $\hat{g}_t$ follows an AR(1) process with autocorrelation equal to 0.85. We calibrate the standard deviations of the two shocks (std(ϵ_t^G) = 0.48 and std(ϵ_t^A) = 0.39) in line with Christoffel, Kuester and Linzert (2009).

A.3 A Quantitative Evaluation Based on a More General Calibration

While the purpose of this exercise is to highlight the general effects of the LMIs on fiscal spending multipliers, the results presented in Section 3 might, however, be due to the specific calibration chosen. In order to assess the validity of the model's implications in a more general setting, we now extend the analysis.

[1] We opted for a calibration of the model that is not anchored to any particular country, but instead adopts a general specification with parameter values that fall within the ranges commonly reported in the literature. In this context, we refer, for example, to Christoffel, Kuester and Linzert (2009), whose model includes many of the parameters relevant to our basic setup and incorporates similar structural frictions; moreover, their estimates are based on euro area data, which closely aligns with the composition of our country sample. For the interest rate smoothing parameter, we also consider the values reported by Lansing (2002), Zanetti (2011) and Kim and Park (2019), among others. With respect to price rigidity (the Calvo parameter), we are guided by the empirical findings of Cagliarini, Robinson and Tran (2011) and Lhuissier and Zabelina (2015). For the elasticity of matching, γ, we rely on the evidence presented by Burda (2014) and Pissarides and Petrongolo (2001).

Table A1: Calibration of the Model.

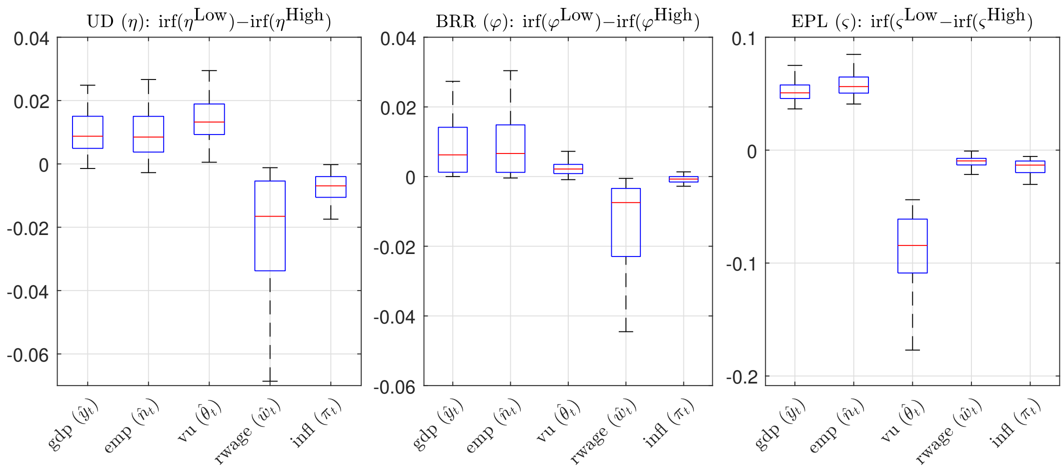
Parameter	Description	Value	Range
β	Discount factor	0.992	[0.95 – 0.999]
γ	Elasticity of matching of unemployed persons	0.45	[0.05 – 0.95]
g_y	Government consumption share in total output	0.2	[0.1 – 0.5]
ζ	Ratio of mrs to mpl	0.7	[0.55 – 0.95]
θ	Labor market tightness	0.43	[0.05 – 0.95]
p	Probability of an unemployed person finding a job	0.30	[0.05 – 0.95]
τ	Labor tax rate	0	[0 – 0.5]
μ_a	Steady state mean of idiosyncratic productivity	0.0	[-2 – 2]
σ_a	Steady state standard-deviation of idiosyncratic productivity	0.15	[0.01 – 5]
$\bar{\varrho}$	Exogenous job separation rate	0.03	[0.01 – 0.15]
$\varrho(\tilde{a})$	(Overall) Job separation rate	0.05	[0.03 – 0.2]
σ	Complementarity coefficient	0.2	[0.05 – 1]
ϕ_π	Inflation sensitivity in the Taylor rule	1.5	[1.0 – 2.5]
ϕ_y	Output sensitivity in the Taylor rule	0.5	[0.0 – 0.9]
ρ_i	Nominal interest rate smoothing	0.3	[0.0 – 0.7]
ρ_G	Government consumption smoothing	0.95	[0.85 – 0.99]
ξ	Calvo price stickiness	0.7	[0.45 – 0.95]
γ_p	Share of inflation indexed prices	0.3	[0.0 – 0.5]
η	Bargaining power of workers (UD)	0.4	—
φ	Unemployment benefit replacement rate (BRR)	0.0	—
ς	Firing costs in relation to last wage (EPL)	0.0	—

We consider a continuum of values for all parameters other than η, φ and ς for which Table A1 provides the details. We simulate the model over a wide range of different values for the parameters. To this purpose, we attach a uniform distribution to each parameter and define upper and lower bounds as indicated in the fourth column (*Range*) in Table A1. We simulate the model 2000 times and compute the difference of the impulse response functions for the following two scenarios: low value of ϑ_i versus high value of ϑ_i where ϑ_i refers to one of the three parameters of interest (η, φ and ς). We focus on the impact responses. The three scenarios (UD, η; BRR, φ; and EPL, ς) are depicted in the sub-panels in Figure A1.[2] The box-plots show the difference in the impact response for each of the three cases for the following variables: output, employment, unemployment, the labor market tightness (v_t/u_t) and real wage. The difference is computed by considering the impulse response functions with a low value of the parameter of interest relative to a high value.

Considering the output response ($\hat{y}_t$) in the left sub-panel as an example, we notice that it is positive throughout due to the fact that the impact response of output when workers have a low power within the wage

[2] We draw values for the structural parameters shown in Table A1. For instance, in the case for η: for a particular draw, we solve the model for $\eta = 0.5$ and compute impulse response functions. For the same draw we also solve the model using $\eta = 0.6$ – in both cases holding the remaining parameters fixed. We use the average value over the first four periods of the impulse response functions and consider the difference thereof which is depicted in Figure A1. By this procedure we can uniquely attach the difference in the impulse response functions to changes in η, while at the same time allowing for flexibility in the model calibration. We carry out the same exercise for ς and φ.

Figure A1: Fiscal Spending Shocks and the LMIs.



negotiations ($\eta = 0.3$), is systematically higher than that when they have a high power ($\eta = 0.7$). The positive range of values in this particular plot replicates the path of the fiscal spending multipliers shown in Figure 2. The employment response replicates that of output, unemployment shows instead a negative reaction, that is, in response to an expansionary fiscal spending shock, unemployment declines by more if workers' power within the wage negotiations is low. The figure highlights also that the impact response of the real wage is hardly affected by the η, and the reaction in the labor market tightness (θ_t) is ambiguous due to the different effects of η on the job creation and job destruction activities by firms on the one hand, and labor supply decisions of households on the other hand.

The remaining two sub-panels show the results for the unemployment benefits replacement rate (φ) and the extent of employment protection (ς). In both cases, the box-plots for output and employment are positive throughout, highlighting the extent to which values of φ and/or ς attenuate fiscal spending multipliers.

We conclude that the general results provided here confirm those put forward in Section 3. The assessment carried out in this section only concerns the calibration of the model's parameters; however, it ignored the extent to which the structure of the model might shape the overall results. Against this background, the following Sections will address specific extensions of the model.

A.4 The LMIs and Macroeconomic Volatility

The current section serves to assess the consequences of the LMIs on the overall macroeconomic volatility. To this purpose, we consider a government spending shock next to a technology shock so that the model

Table A2: Volatility of output, employment and the real wage.

LMI:	UD (η)	BRR (φ)	EPL (ς)
	Government spending shock (ϵ_t^G)		
Output ($\hat{y}_t$)	2.00	2.40	1.00
Employment ($\hat{n}_t$)	2.00	2.40	0.99
Real wage ($\hat{w}_t$)	3.27	2.93	0.75
Consumption ($\hat{c}_t$)	0.68	0.63	1.34
	Technology shock (ϵ_t^A)		
Output ($\hat{y}_t$)	1.02	1.52	0.90
Employment ($\hat{n}_t$)	1.15	3.08	0.46
Real wage ($\hat{w}_t$)	2.25	2.83	0.73
Consumption ($\hat{c}_t$)	1.01	1.54	0.79

Notes: The table shows the sensitivity of the output, employment and the real wage volatilities to changes in the LMIs. The shocks considered are a government spending and a technology shock. The values indicate the standard deviation of output, employment and the real wage (x_t's) when the respective LMIs take on a low value relative to the standard deviations when the LMIs are set at a high value ($\mathrm{Var}(x_t(\mathrm{LMI}_{\mathrm{low}}))/\mathrm{Var}(x_t(\mathrm{LMI}_{\mathrm{high}}))$). The shocks considered are a technology shock (ϵ_t^A) and the government spending shock (ϵ_t^G).

comprises a demand and supply shock. We decompose the effect of LMIs on the volatility of the endogenous variables to the two shocks. That is, we examine the extent to which LMIs affect macroeconomic volatility in the wake of demand and supply shocks and analyze each shock separately.

This is motivated by the fact that distinct shocks give rise to distinct cross-correlations (sign and values). While such a setting is admittedly unrealistic, it allows us to assess whether our results are driven by a specific shock and if the heterogeneity in the volatility of the endogenous variables conditional on the LMIs is important. Considering Eq. (3.23) and following Hamilton (1994, Chapter 10.2), the variance-covariance matrix $\mathbf{\Sigma}_z(\boldsymbol{\vartheta})$ of the vector of endogenous variables z_t of the solved DSGE model is given by

$$\mathrm{vec}(\mathbf{\Sigma}_z(\boldsymbol{\vartheta})) = \left(I - \mathbf{\Psi}_0^{-1}\mathbf{\Psi}_1 \otimes \mathbf{\Psi}_0^{-1}\mathbf{\Psi}_1\right)^{-1} \mathrm{vec}\left(\mathbf{\Psi}_0^{-1}\Sigma_\epsilon \left(\mathbf{\Psi}_0'\right)^{-1}\right) \tag{A.1}$$

where the dependency of $\mathbf{\Psi}_0$ and $\mathbf{\Psi}_1$ on $\boldsymbol{\vartheta}$ has been omitted to preserve notational simplicity. This expression explicitly accounts for the fact that the volatility depends on the structural parameters of interest, UD (η), BRR (φ) and EPL (ς), in $\boldsymbol{\vartheta} = (\eta, \varphi, \varsigma)'$. In what follows we again confine the analysis to the volatility of output ($\hat{y}_t$), employment ($\hat{n}_t$) and the real wage ($\hat{w}_t$). We use the estimated values put forth in Christoffel, Kuester and Linzert (2009) to calibrate the parameters for the autocorrelation coefficients (ρ_i) of the AR(1) shocks. The calibration of the idiosyncratic variances (σ_i^2) is described in Section A.2 and we consider two values (high and low) for each LMI in this respect.

The results are depicted in Table A2. The table only shows the values of the variance of the endogenous variables (output, employment, and the real wage) for a low value of the LMIs relative to their variances when a high value is used. A few results emerge from the analysis. First, in the case of a technology shock, the ability of UD to influence fluctuations in output and consumption is limited, whereas its impact is sizable in the case of a government spending shock. In contrast, employment protection legislation (EPL) exhibits the opposite pattern, having a greater effect on output and consumption fluctuations in the case of a technology shock. Second, in the case of government spending shocks, high values of UD and BRR tend to decrease output volatility while exacerbating consumption volatility. In the case of a technology shock, both output and consumption volatility drop with higher values of UD and BRR. Again, the opposite pattern is observed with EPL, where the volatility dynamics are reversed. Third, the quantitative impact of LMIs on macroeconomic volatility in response to technology shocks exceeds the corresponding impact of a government spending shock only when EPL is considered, whereas this is not the case for UD and BRR. In summary, these findings suggest that LMIs can potentially mitigate macroeconomic volatility. However, this effect is highly dependent on the type of shock that prevails and the specific LMI in question.

B. Extensions to the Theoretical Model

This section considers various extensions to the baseline model outlined in Section 3. These include shutting down the inflation indexation of prices, real wage rigidity, limited asset market participation, the case when firing costs accrue to the government as revenues, productivity enhancing government spending, and a deeper look into the consumption and leisure complementarity. We consider one extension at a time, as otherwise the precise role of the additional frictions considered becomes difficult to assess. Note that we report the outcomes in this section in terms of elasticities according to Eq. (3.24) and not in terms of multipliers.

B.1 Inflation indexation of prices

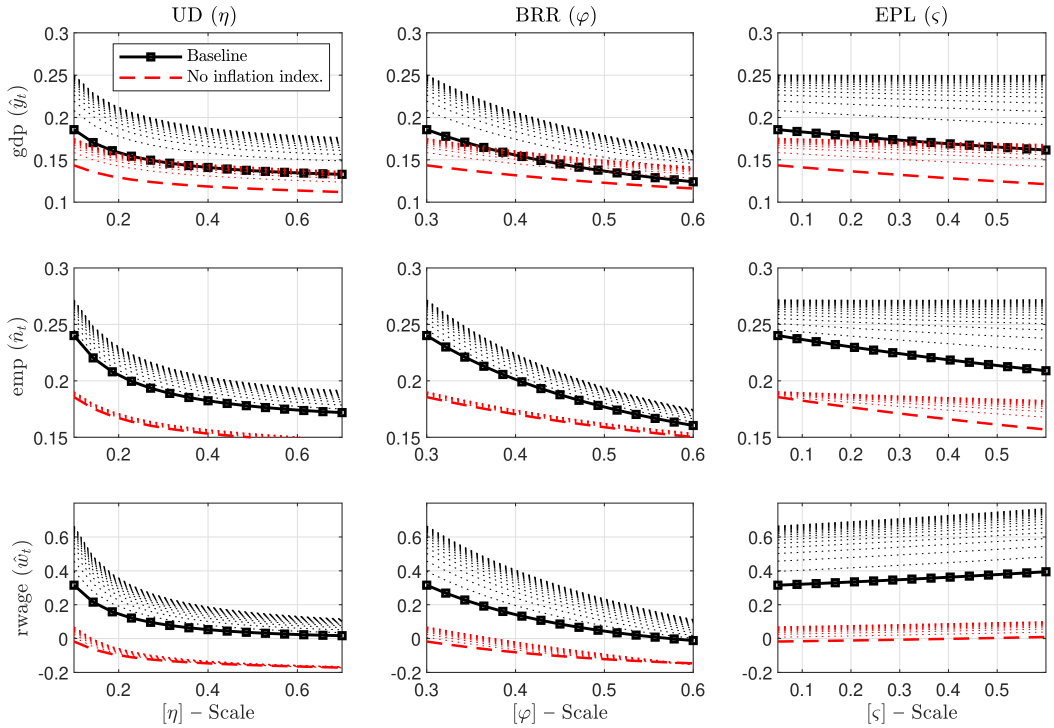
Our baseline model incorporates nominal frictions along two dimensions. The first dimension pertains to the infrequent adjustment of prices towards their optimal levels. The second involves a rule-of-thumb price adjustment mechanism for prices that cannot be optimally adjusted within a given period. The latter represents a commonly observed price-setting behavior. This extension increases the degree of inflation inertia and amplifies the real costs associated with prices being set sub-optimally.

This section examines the immediate consequences of this rule-of-thumb price adjustment mechanism on fiscal spending multipliers. The results are presented in Figure B1. The figure contrasts the baseline results (depicted in black), which assume ($\gamma_p = 0.3$) (indicating that 30% of prices are adjusted based on the previous period's inflation rate), with those in which inflation indexation is absent ($\gamma_p = 0.0$), shown by the red lines. As observed, the fiscal spending elasticity for output is smaller when inflation indexation is absent, underscoring the significance of delayed price adjustments in determining the overall impact of an expansionary demand shock on output.

B.2 Real Wage Rigidity

The existence of real wage rigidities has been pointed out by many authors as a feature needed to account for a number of labor market facts (see, among others, Hall, 2005). Krause and Lubik (2007) stress the role of real wage rigidity to improve the predictions of the labor market. Real wage rigidity comprises a particularly important aspect for our case: A rigid real wage strongly increases the incentive to create jobs in the wake of an expansionary fiscal spending shock (or expansionary demand shock in more general terms), since firms share less of the benefit with their workers. However, at the same time, as vacancies rise and unemployment

Figure B1: Fiscal spending elasticities and the LMIs ($\mu(\boldsymbol{\vartheta})$) – no inflation indexation of prices



Notes: The figure shows the sensitivity of the fiscal spending elasticity to variations in the three LMI parameters (UD–union density, BRR–unemployment benefit replacement rate and EPL–employment protection legislation). The elasticities are defined in Eq. (3.24) and shown for different horizons, with the contemporaneous multiplier ($\mathcal{P} = 1$) represented by a solid black square line for the baseline model ($\gamma_p = 0.3$) and by a red dashed line for the model without inflation indexation of prices ($\gamma_p = 0.0$). The higher-horizon elasticity ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model without inflation indexation).

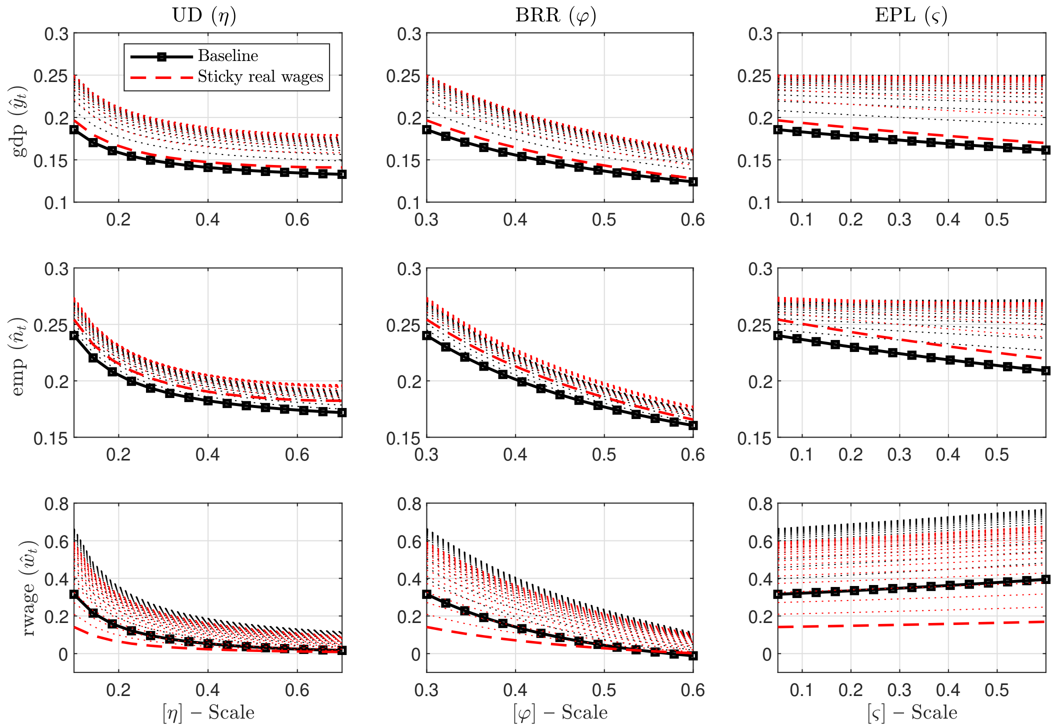
falls, there is a substantial increase in the cost of hiring workers (κ/q_t rises since q_t falls on the back of an increase in vacancies v_t), which is a component of firms' real marginal costs. Hence, the role of rigid real wages can be confined to two elements, one of which becomes more rigid while the other becomes more volatile.

We assume that real wages (w_t) respond sluggishly to changes in labor market conditions. To simplify the exposition, we proceed by considering real wage inertia as a result of some imperfections or frictions in labor markets, which are modeled in a reduced form. Specifically, we assume the partial adjustment model, which extends Eq. (3.16) to the following

$$w_t = \varrho_w w_{t-1} + (1 - \varrho_w)\check{w}_t \tag{B.1}$$

where $\check{w}_t = (1-\eta)\frac{mrs_t + b_t^u}{1-\tau} + \eta\left(mpl_t + \mathbb{E}_t\Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right]\right)$. The parameter ϱ_w captures the extent of real wage rigidity, and we choose a value equal to 0.4. Eq. (B.1) can be considered a parsimonious

Figure B2: Fiscal spending elasticities and the LMIs ($\mu(\boldsymbol{\vartheta})$) – real wage rigidity



Notes: The figures shows the sensitivity of the fiscal spending elasticity to variations in the three LMI parameters (UD–union density, BRR–unemployment benefit replacement rate and EPL–employment protection legislation). The elasticities are defined in Eq. (3.24) and shown for different horizons, with the contemporaneous elasticity ($\mathcal{P}$ = 1) represented by a solid black square line for the baseline model (ϱ_w = 0.0) and by a red dashed line for the model with real wage rigidity (ϱ_w = 0.4). The higher-horizon elasticity ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model with real wage rigidity).

and ad hoc way of modeling the sluggish adjustment of real wages to changes in labor market conditions. Alternative formalizations, explicitly derived from the staggering of real wage decisions and the like, are presented in Blanchard and Galí (2007), Zanetti (2007), and Gertler, Huckfeldt and Trigari (2020).

The results of the model extended for real wage rigidities are shown and compared to the baseline model in Figure B2. Let us first consider the dependency of the output elasticity on BRR (φ) and EPL (ς), shown in the sub-panels in the second and third columns. It can be seen that the shape of the output elasticity with respect to the two LMIs does not change; instead, the extent of real wage rigidity causes a more or less proportional drop in the size of the elasticity. This highlights that the rise in hiring costs in the wake of the expansionary demand shock dominates the drop in the benefit firms have to share with workers. This attenuates firms incentives to create jobs. The output and employment elasticity are thus smaller when real wage rigidities are present.

In the case of UD (η), the elasticities for output and employment are affected more profoundly when real wage rigidities are present. Both elasticities now show a concave pattern with respect to η: when η is low, increases therein raise fiscal spending elasticities, while the opposite occurs when η is already high. The intuition is that when η is low, the drop in the benefits firms have to share with workers now dominates the increase in hiring costs, giving rise to a positive dependency between η and the output and employment elasticities. For higher values of η, the dominance structure changes, and the baseline results (higher η causes a smaller output multiplier) apply again. Nevertheless, a concave pattern emerges only modestly and is confined to small values of η.

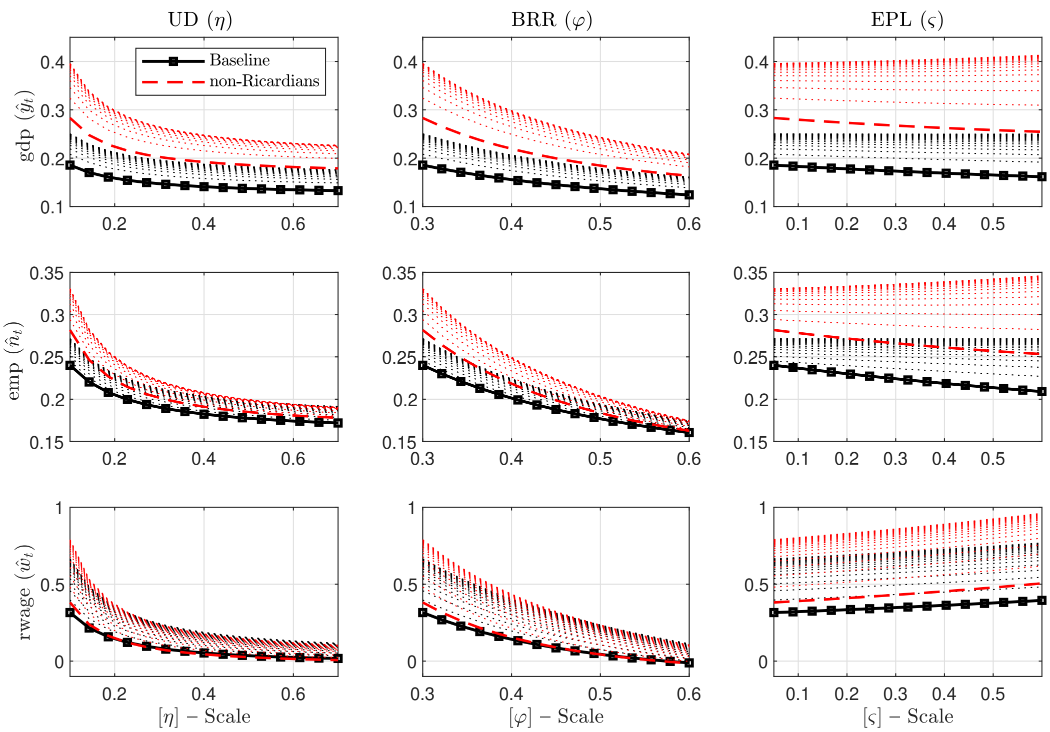
B.3 Limited Asset Market Participation

Galí, López-Salido and Valles (2007) show how the interaction of rule-of-thumb consumers with sticky prices and deficit financing can account for the existing evidence on the effects of government spending. In this context, rule-of-thumb consumers are characterized by limited asset market participation, which implies that they lack the ability to smooth their consumption profile; as a consequence, they spend (consume) all their income each period. This rule-of-thumb gives rise to a consumption pattern that strongly aligns with wage income. This results in a positive consumption response in the wake of an expansionary fiscal spending shock.

We follow Galí, López-Salido and Valles (2007) and add the second consumer type into the baseline model. The consumers outlined in the baseline model are now referred to as *Ricardian* consumers, and their consumption is henceforth referred to as c_t^r (the same for their labor supply n_t^r). On the other hand, *rule-of-thumb* households are assumed to behave in a "hand-to-mouth" fashion, fully consuming their current labor income. Their period utility is given by $u(c_t^{nr}, n_t^{nr})$, and they are subject to the budget constraint $c_t^{nr} = (1-\tau)w_t n_t^{nr} + b_t^u(1-n_t^{nr}) + T_t^{s,nr}$. Aggregate consumption and employment are given by a weighted average of the corresponding variables for each consumer type. Formally, $c_t = \lambda c_t^{nr} + (1-\lambda)c_t^r$, $n_t = \lambda n_t^{nr} + (1-\lambda)n_t^r$. It is further assumed that the labor market is characterized by a structure that gives rise to wages negotiated in a centralized manner by an economy-wide union with firms.

Figure B3 shows the results of the LMIs on the elasticities for output, employment, and real wages in the extended model (labeled "non-Ricardian") and compares them to the baseline model. The simulations are based on a share of one-quarter of non-Ricardian households ($\lambda = 0.25$). The elasticities are generally higher; this applies to both the output and employment elasticities, as well as the real wage elasticity. The reason

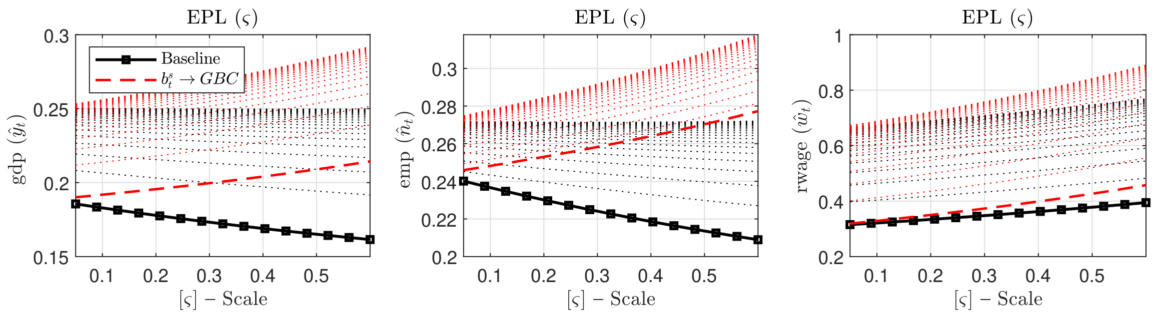
Figure B3: Fiscal spending elasticities and the LMIs ($\mu(\boldsymbol{\vartheta})$) – limited asset market participation



Notes: The figure shows the sensitivity of the fiscal spending elasticity to variations in the three LMI parameters (UD–union density, BRR–unemployment benefit replacement rate and EPL–employment protection legislation). The elasticities are defined in Eq. (3.24) and shown for different horizons, with the contemporaneous elasticity ($\mathcal{P} = 1$) represented by a solid black square line for the baseline model ($\lambda_p = 0.0$) and by a red dashed line for the model with limited asset market participation ($\lambda_p = 0.25$). The higher-horizon elasticities ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model without inflation indexation).

for the higher elasticity throughout is due to the different reaction of consumption. In the baseline model, consumption declines due to the negative wealth effect that accompanies the (deficit financed) increase in fiscal spending. The (absolute) size of the decline is, however, decreasing in λ, reflecting the offsetting role of rule-of-thumb behavior on the conventional negative wealth and intertemporal substitution effects triggered by the fiscal expansion. The figure thus illustrates the amplifying effects of the introduction of rule-of-thumb consumers. Most importantly, though, is the fact that the introduction of limited asset market participation does not change the dependency of the elasticities on the LMIs. With respect to the output multiplier, the negative relation with the LMIs still applies. Even more, the negative relation now turns out stronger than in the baseline model.

Figure B4: Fiscal spending elasticities and the LMIs ($\mu(\boldsymbol{\vartheta})$) – firing costs accrue to the government



Notes: The figures shows the sensitivity of the fiscal spending elasticity to variations in the three LMI parameters (UD–union density, BRR–unemployment benefit replacement rate and EPL–employment protection legislation). The elasticities are defined in Eq. (3.24) and shown for different horizons, with the contemporaneous elasticity ($\mathcal{P} = 1$) represented by a solid black square line for the baseline model and by a red dashed line for the model with in which firms' firing costs accrue to the government as revenue. The higher-horizon elasticities ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model in which firms' firing costs accrue to the government as revenue).

B.4 Firing Costs as Government Revenues

The baseline model specifies firing costs as real resource costs. This is quite a strong assumption, as in many countries, firing costs arise in the context of severance payments that are eventually re-distributed back to households. Against this background, we now assess the implications of ς, once firms' expenses on firing accrue to the government as revenues. These additional revenues will eventually be re-distributed back to households in the form of lump-sum subsidies or the like. Hence, in this case, the government budget constraint Eq. (3.17) and the real resource constraint Eq. (3.22) are given by:

$$F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1}v_{t-1})b_t^s + \tau w_t n_t + B_t = R_{t-1}B_{t-1} + b_t^u u_t + T_t^s + g_t \tag{B.2}$$

$$y_t = c_t + g_t + \kappa v_t \tag{B.3}$$

We extend the baseline model in this respect. Since the simulations for η and φ are based on zero firing costs ($\varsigma = 0$), this extension has no effect on the shape of the elasticities with respect to η and φ.

The results are shown in Figure B4. As can be seen, when firing costs accrue to the government, fiscal spending elasticities are notably higher. In particular, the reaction in employment and output is more positive (for values of $\varsigma > 0$), while at the same time, the contraction in the real wage is augmented too. The key element behind this pertains to the redistributive element that operates in the background. When firing costs accrue to the government, they are re-distributed back to households, giving rise to a smaller drop in consumption in response to the fiscal spending shock, which in turn raises the output elasticity. In contrast, when firing costs enter the aggregate resource constraint, this implies that they are real resource costs that cannot be uncovered. This loss attenuates the output elasticity; the attenuation effect increases with ς, which captures the firing costs per laid off worker. While this attenuation effect is also present when firing costs are re-distributed back to households via the government, the re-distribution channel raises the output multiplier. This effect is absent in the other case.

B.5 Productivity Enhancing Government Spending

The standard assumption in macroeconomics is that government spending is unproductive. An even more extreme but common assumption is that government spending is entirely purposeless, with purchases comprising real resource costs. These assumptions contrast with the observation that various public goods indeed enhance the productivity of the economy. Examples include the extensive rail system in Europe, public education, government-funded research, among other projects (Daniel and Gao, 2015). Against this background, we extend the baseline model to allow for productivity enhancing public spending. The literature considers distinct approaches in this respect. Daniel and Gao (2015), for instance, model productive government spending as subsidies to education, which build up the human capital stock. Kumhof et al. (2010) consider a set-up in which government spending accumulates a productive capital stock that enters the production function. We proceed by assuming that government spending g_t builds up the public capital stock, which then enters the production function. The public capital stock evolves according to

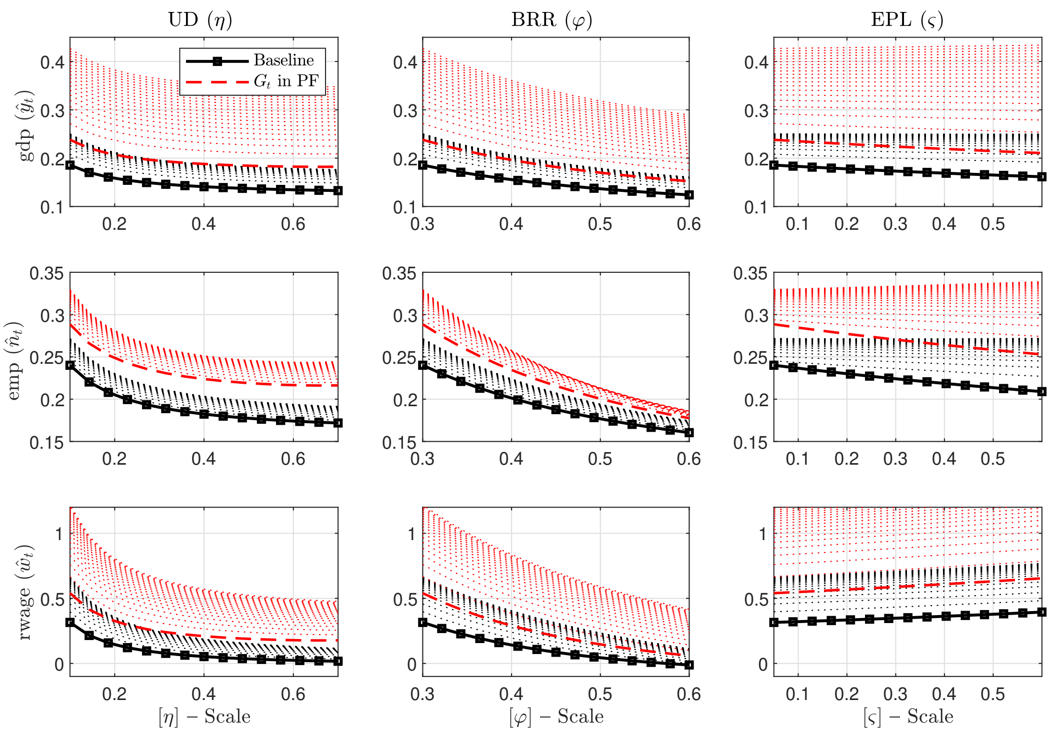
$$k_t^G = (1-\delta)k_{t-1}^G + g_t \tag{B.4}$$

where δ is the depreciation rate of the public capital stock, which we set equal to 0.01 (quarterly data frequency). Importantly, the public capital stock is identical for all firms and provided free of charge to the end user (but not to the taxpayer). This approach conforms with the set-up in Kumhof et al. (2010). We modify the production function as follows

$$y_t(g_t) = \bar{A}n_t A(\tilde{a}_t) \cdot \left(\frac{k_t^G}{\bar{k}^G}\right)^{\alpha_g} \tag{B.5}$$

The parameter $\alpha_g \in [0,1]$ captures the sensitivity (elasticity) of aggregate production with respect to changes in the public capital stock, and $\bar{k}^G$ is the steady state value for k_t^g. Note that this production function exhibits constant returns to scale in private inputs (n_t), while public spending enters externally in an analogous manner to exogenous technology. Hence, government spending augments labor productivity

Figure B5: Fiscal spending elasticities and the LMIs ($\mu(\boldsymbol{\vartheta})$) – productivity enhancing government spending.

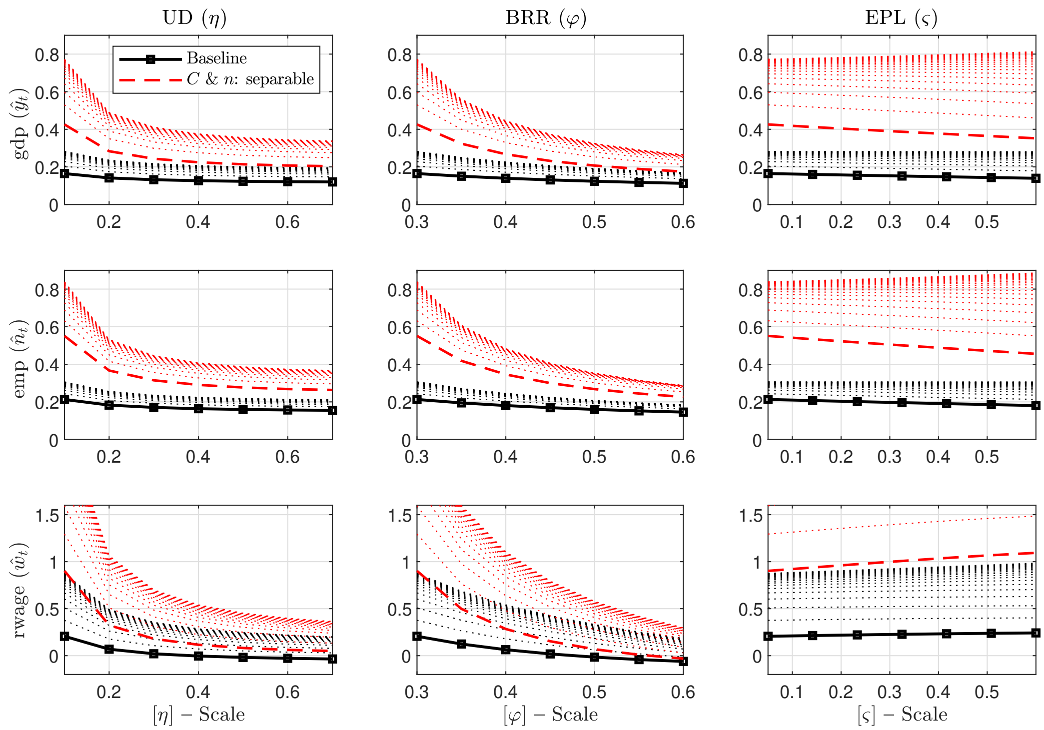


Notes: The figure shows the sensitivity of the fiscal spending elasticities to variations in the three LMI parameters (UD–union density, BRR–unemployment benefit replacement rate and EPL–employment protection legislation). The elasticities are defined in Eq. (3.24) and shown for different horizons, with the contemporaneous elasticity ($\mathcal{P} = 1$) represented by a solid black square line for the baseline model ($\alpha_g = 0.0$) and by a red dashed line for the model with a productive public capital stock ($\alpha_g = 0.1$). The higher-horizon elasticities ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model with a productive public capital stock).

directly: $mpl_t(k_t^g) = y_t(k_t^g)/n_t$. We chose a conservative value for the elasticity $\alpha_g = 0.1$, which implies that a one percent increase in the public capital stock (relative to the steady state) raises labor productivity by 0.1 percent.

We carry out the same simulations as in Section 3. The results are shown in Figure B5. Productive government spending leads to a significantly higher output elasticity. At the same time, the employment elasticity is attenuated owing to the rise in labor productivity and the higher real wage. The latter comprises the most noteworthy change compared to the baseline results. The higher labor productivity causes a rise in the real wage already at impact. For us, the most important factor, however, is the impact of the LMIs on the output elasticity. With a view to Figure B5, while the output response increases with the extent of productive government spending, the dependency of the output elasticity with respect to the three LMIs

Figure B6: Fiscal spending elasticities and the LMIs ($\mu(\boldsymbol{\vartheta})$) – consumption/leisure complementarity.



Notes: The figure shows the sensitivity of the fiscal spending elasticities to variations in the three LMI parameters (UD–union density, BRR–unemployment benefit replacement rate and EPL–employment protection legislation). The elasticities are defined in Eq. (3.24) and shown for different horizons, with the contemporaneous elasticity ($\mathcal{P}$ = 1) represented by a solid black square line for the baseline model ($sigma$ = 0.2) and by a red dashed line for the model with additively separable consumption and leisure (σ = 1). The higher-horizon elasticities ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model with a productive public capital stock).

remains, however, unchanged compared to the baseline results. In each of the three cases (UD, BRR, and EPL), a higher value attenuates the output reaction in response to a government spending increase.

B.6 Consumption and leisure complementarity

The utility function specification allows for both complementarity and separability between consumption and leisure (or, equivalently, the negative of labor supply), with the relationship determined by the parameter σ. When $\sigma = 1$, consumption and leisure are additively separable, while for $0 < \sigma < 1$, they exhibit varying degrees of complementarity. The degree of complementarity is particularly relevant in the context of fiscal spending shocks. In addition to price rigidities, a deficit-financed increase in government spending generates a negative wealth effect. The magnitude of this wealth effect depends on whether consumption and leisure are separable.

The negative wealth effect arises because, in response to a future anticipated tax burden, forward-looking households adjust their behavior by reducing current consumption and increasing labor supply (i.e., reducing leisure). The relative strength of the response in consumption and leisure is determined by the degree of complementarity between the two, as captured by the parameter σ. In what follows, we consider the case of $\sigma = 1$ as an alternative scenario to the baseline result, which relies on $\sigma = 0.2$.

Figure B6 provides the results. The decline in consumption and leisure is attenuated when complementarity between the two prevails, thereby moderating the overall negative wealth effect. Specifically, when $\sigma < 1$, the negative wealth effect is relatively weak. As a result, the reduction in leisure and the corresponding increase in labor supply are modest. This leads to a relatively small increase in output, implying a smaller fiscal spending elasticity compared to the case where $\sigma = 1$, which corresponds to the separability between consumption and leisure and generates a more pronounced negative wealth effect.

Intuitively, when $\sigma < 1$, agents substitute leisure for consumption, which mitigates the negative wealth effect on consumption. The extent of this mitigation depends on the degree of complementarity between labor and consumption, with stronger complementarity leading to a more pronounced response in consumption.

C. Bayesian Interacted Panel Vector Autoregressions

In this section, we provide estimation details on the Bayesian Interacted Panel Vector Autoregression (IPVAR). The model is similar to the model proposed by Towbin and Weber (2013) and Sá, Towbin and Wieladek (2014). The model is estimated in its recursive form to allow for contemporaneous interactions. Structural analysis (e.g., IRFs or FEVDs) is then carried out given a particular value of the interaction term.

Let $\{\boldsymbol{y}_{it}\}_{t=1}^{T_i}$ and $\{\boldsymbol{\vartheta}_{it}\}_{t=1}^{T_i}$ denote an n- and d-dimensional time series process for country $i = 1, \ldots, N$, respectively. Note that we allow for differing sample lengths for country i, specified with sample length T_i. We write the IPVAR as follows

$$\boldsymbol{J}_{it}\boldsymbol{y}_{it} = \boldsymbol{a}_i + \sum_{j=1}^{p}\left(\boldsymbol{A}_{ij}\boldsymbol{y}_{it-j} + \sum_{l=1}^{d}\boldsymbol{B}_{ijl}\boldsymbol{y}_{it-j} \times \vartheta_{ilt}\right) + \tilde{\boldsymbol{u}}_{it}, \quad \tilde{\boldsymbol{u}}_{it} \sim \mathcal{N}_M(\boldsymbol{0}, \boldsymbol{\Omega}_i). \tag{C.1}$$

We denote with $\boldsymbol{a}_i$ the $n \times 1$ country-specific intercept vector, while $\boldsymbol{A}_{ij}$ denotes the $n \times n$ country-specific autoregressive coefficient matrix for lag $j = 1, \ldots, p$. The $n \times 1$ vector of residuals $\boldsymbol{u}_{it}$ is assumed to be uncorrelated across countries and normally distributed with mean zero and a $n \times n$ covariance matrix $\boldsymbol{\Omega}_i$. Due to the recursive structure of the VAR, the covariance matrix is diagonal, i.e., $\boldsymbol{\Omega}_i = \text{diag}(\omega_{i1}, \ldots, \omega_{in})$. The interaction term $\boldsymbol{\vartheta}_{it}$ is allowed to influence the dynamic relationship between the endogenous variables of the system via the $n \times n$ coefficient matrices $\boldsymbol{B}_{ijl}$ for lag $j = 1, \ldots, p$ and interaction variable $l = 1, \ldots, d$. Lastly, we have to discuss the nature of the $n \times n$ matrix $\boldsymbol{J}_{it}$, which is a lower unitriangular matrix. This matrix exhibits a time index t because we allow the interaction term to affect the contemporaneous relationships between equations. The contemporaneous effect of the q-th ordered variable on the w-th ordered variable is given by $-[\boldsymbol{J}_{it}]_{wq}$, where we denote the scalar element in the w-th row and q-th column of the matrix $\boldsymbol{J}_{it}$ as $[\boldsymbol{J}_{it}]_{wq}$. The elements are modeled as follows

$$[\boldsymbol{J}_{it}]_{wq} = \begin{cases} [\tilde{\boldsymbol{J}}_{i0}]_{wq} + \sum_{l=1}^{d}[\tilde{\boldsymbol{J}}_{il}]_{wq}\vartheta_{ilt}, & \text{if } q < w, \\ 1, & \text{if } q = w, \\ 0, & \text{if } q > w. \end{cases} \tag{C.2}$$

The model parameters can be re-written as a function of $\boldsymbol{\vartheta}_{it}$. Hence, this results in

$$\boldsymbol{y}_{it} = \boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) + \sum_{j=1}^{p}\boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it})\boldsymbol{y}_{it-j} + \boldsymbol{u}_{it}, \quad \boldsymbol{u}_{it} \sim \mathcal{N}_n\left(\boldsymbol{0}, \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it})\right), \tag{C.3}$$

where $\boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}\boldsymbol{a}_i$, $\boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}\left(\boldsymbol{A}_{ij} + \sum_{l=1}^{d}\boldsymbol{B}_{ijl}\vartheta_{ilt}\right)$, and $\boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it}) = (\boldsymbol{J}_{it}^{-1})\boldsymbol{\Omega}_i(\boldsymbol{J}_{it}^{-1})'$. From this representation, it is straightforward to derive impulse response functions (IRFs) or compute the forecast error variance decomposition (FEVD) *given* a particular value of the interaction term $\boldsymbol{\vartheta}_{it}$.

We pursue a Bayesian approach to the estimation of the model. Therefore, we discuss our prior setup next. The prior setup is similar in spirit to the one presented in Jarociński (2010) but we additionally impose regularization with global-local shrinkage priors (Griffin and Brown, 2010). This has been shown to be beneficial when applied to VARs (Huber and Feldkircher, 2019). We use a variant of the Normal-Gamma (NG) shrinkage prior for each level of the model. In particular, we use the lagwise version of the Normal-Gamma prior, inducing more shrinkage for higher-order lags. Furthermore, we shrink coefficients in the estimation equation to their common mean and the common mean towards zero. For the specification of the prior distribution, we start by stacking to a $k = (1+d)n^2$-dimensional vector $\boldsymbol{\beta}_{ij} = \text{vec}(\boldsymbol{A}_{ij}, \boldsymbol{B}_{ij1}, \ldots, \boldsymbol{B}_{ijd})$ for lag j and country i and specify the prior distribution as follows

$$[\boldsymbol{\beta}_{ij}]_s \mid \lambda_{ij}^2, [\boldsymbol{\theta}_{ij}]_s \sim \mathcal{N}\left([\boldsymbol{b}_j]_s, 2/\lambda_{ij}^2[\boldsymbol{\theta}_{ij}]_s\right), \quad [\boldsymbol{\theta}_{ij}]_s \sim \mathcal{G}\left(\vartheta_\theta, \vartheta_\theta\right), \quad s = 1, \ldots, k. \tag{C.4}$$

Here $[\boldsymbol{\beta}_{ij}]_s$, $[\boldsymbol{b}_j]_s$, and $[\boldsymbol{\theta}_{ij}]_s$ denote the s-th element of the respective vector. The latter is the local-shrinkage component on which we specify a Gamma distribution with hyperparameter ϑ_θ. This hyperparameter governs the strength of the regularization towards the specified mean. For instance, centering the hyperparameter on unity translates into the Bayesian LASSO (Park and Casella, 2008). Instead, we allow for additional flexibility and put a hyperprior on $\vartheta_\theta \sim Exp(1)$, centered a priori on unity. λ_{ij}^2 denotes the global-shrinkage component. The lagwise NG prior setup features one global-shrinkage component per lag to impose more shrinkage for higher-order lags (similar in spirit to the Minnesota prior setup of Doan, Litterman and Sims, 1984). Hence, the prior distribution on λ_{ij}^2 is a multiplicative Gamma prior

$$\lambda_{ij}^2 = \prod_{g=1}^{j}\zeta_{ig}^{\lambda}, \quad \zeta_{ig}^{\lambda} \sim \mathcal{G}\left(c_0, d_0\right), \tag{C.5}$$

with $c_0 = d_0 = 0.01$. As long as the global-shrinkage parameter λ_{ij}^2 exceeds unity, this prior shrinks coefficients associated with higher lags more towards zero. This implies that the coefficient vector $\boldsymbol{\beta}_{ij}$ becomes increasingly sparse for higher lags. Next, we impose an NG prior on the free off-diagonal elements of $\tilde{\boldsymbol{J}}_{it}$

$$[\tilde{\boldsymbol{J}}_{il}]_{st} \mid \delta_{il}^2, [\boldsymbol{\theta}_{il}^{\tilde{J}}]_{st} \sim \mathcal{N}\left([\boldsymbol{g}_l]_{st}, 2/\delta_{il}^2[\boldsymbol{\theta}_{il}^{\tilde{\boldsymbol{J}}}]_{st}\right), \quad [\boldsymbol{\theta}_{il}^{\tilde{\boldsymbol{J}}}]_{st} \sim \mathcal{G}\left(\vartheta_\theta^{\tilde{\boldsymbol{J}}}, \vartheta_\theta^{\tilde{\boldsymbol{J}}}\right), \tag{C.6}$$

with $s = 2, \ldots, n$ and $t = 1, \ldots, s-1$, denoting the respective row or column index. Again, we specify a hyperprior on $\vartheta_\theta^{\tilde{\boldsymbol{J}}} \sim Exp(1)$, allowing for additional flexibility. Similar to before, we assume a Gamma prior on $\delta_{il}^2 \sim \mathcal{G}(c_0, d_0)$. For the intercept vector, $\boldsymbol{a}_i$, we specify a simple Gaussian $\mathcal{N}(0, 100)$ for each element to be uninformative. We have not yet discussed the prior distributions of the common means, $\boldsymbol{b}_j$ and $\boldsymbol{g}_l$. They no longer feature a country-indicator i, establishing linkages between the country models. This constitutes the second layer of the prior setup, in which we shrink coefficients towards zero. The prior setup looks as follows

$$[\boldsymbol{b}_j]_s \mid \kappa_j^2, [\boldsymbol{\phi}_j]_s \sim \mathcal{N}\left(0, 2/\kappa_j^2[\boldsymbol{\phi}_j]_s\right), \quad [\boldsymbol{\phi}_j]_s \sim \mathcal{G}\left(\vartheta_\phi, \vartheta_\phi\right), \quad s = 1, \ldots, k. \tag{C.7}$$

As before, $[\boldsymbol{b}_j]_s$, and $[\boldsymbol{\phi}_j]_s$ denote the s-th element of the respective vector. We put a hyperprior on $\vartheta_\phi \sim Exp(1)$. Also, similar to before, we use the lagwise NG prior setup for the global component. Therefore, the prior distribution on κ_j^2 looks as follows

$$\kappa_j^2 = \prod_{g=1}^{j} \zeta_g^\kappa, \quad \zeta_g^\kappa \sim \mathcal{G}\left(c_0, d_0\right). \tag{C.8}$$

We conclude the second layer by specifying the NG prior for the off-diagonal elements of $\boldsymbol{g}_l$, which is given by

$$[\boldsymbol{g}_l]_{st} \mid \tau_l^2, [\boldsymbol{\phi}_l^{\boldsymbol{g}}]_{st} \sim \mathcal{N}\left(0, 2/\tau_l^2[\boldsymbol{\phi}_l^{\boldsymbol{g}}]_{st}\right), \quad [\boldsymbol{\phi}_l^{\boldsymbol{g}}]_{st} \sim \mathcal{G}\left(\vartheta_\phi^{\boldsymbol{g}}, \vartheta_\phi^{\boldsymbol{g}}\right), \tag{C.9}$$

with $s = 2, \ldots, n$ and $t = 1, \ldots, s-1$. Again, $\boldsymbol{\vartheta}_\phi^{\boldsymbol{g}} \sim Exp(1)$ and $\tau_l^2 \sim \mathcal{G}(c_0, d_0)$. We conclude the prior setup by specifying a prior on the diagonal elements of $\boldsymbol{\Omega}_i$,

$$\omega_{is} \sim \mathcal{IG}(c_0, d_0), \quad s = 1, \ldots, n, \quad i = 1, \ldots, N. \tag{C.10}$$

D. Data Sources

All series were gathered from the sources listed below, including OECD Main Economic Indicators, OECD National Accounts Quarterly, Eurostat, Annual Macroeconomic (AMECO) database, FRED database, or a national source. All time series cover the period from 1960Q1 to 2020Q4. All series are seasonally adjusted. The gathered data consists of $N = 16$ countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Italy, Japan, the Netherlands, Portugal, Spain, Sweden, and the United States.

In Table D1, we define the exact transformations of the variables used in the estimation. Note that we use year-on-year growth rates. In Table D2, we show the exact sample coverage for each of the estimated models. In particular, we use $N = 16$ countries for the model featuring employment/unemployment, while we only use $N = 13$ countries for the model with the labor market tightness indicator. Sample sizes also reduce for this indicator compared to the other two labor market variables.

Table D1: Variable Definitions.

Variable	Transformation	Details
g_{it}	$100 \times \left[\ln\left(\frac{\mathtt{RGOVC}_{it}}{\mathtt{POP}_{it}}\right) - \ln\left(\frac{\mathtt{RGOVC}_{it-4}}{\mathtt{POP}_{it-4}}\right)\right]$	$\mathtt{RGOVC}_{it}$ is *General Government Final Consumption Expenditure, Constant Prices, seasonally adjusted*
x_{it}	$100 \times \left[\ln\left(\frac{\mathtt{RGDP}_{it}}{\mathtt{POP}_{it}}\right) - \ln\left(\frac{\mathtt{RGDP}_{it-4}}{\mathtt{POP}_{it-4}}\right)\right]$	$\mathtt{RGDP}_{it}$ is *Gross Domestic Product, Constant Prices, seasonally adjusted*
er_{it}	$100 \times \left[\ln\left(\frac{\mathtt{EMP}_{it}}{\mathtt{EMP}_{it}+\mathtt{UE}_{it}}\right) - \ln\left(\frac{\mathtt{EMP}_{it-4}}{\mathtt{EMP}_{it-4}+\mathtt{UE}_{it-4}}\right)\right]$	$\mathtt{EMP}_{it}$ is *Total Employment, Persons, seasonally adjusted*
ur_{it}	$100 \times \left[\ln\left(\frac{\mathtt{UE}_{it}}{\mathtt{EMP}_{it}+\mathtt{UE}_{it}}\right) - \ln\left(\frac{\mathtt{UE}_{it-4}}{\mathtt{EMP}_{it-4}+\mathtt{UE}_{it-4}}\right)\right]$	$\mathtt{UE}_{it}$ is *Harmonised Unemployment, Persons, seasonally adjusted*
v_{it}/u_{it}	$\ln\left(\frac{\mathtt{VAC}_{it}}{\mathtt{UE}_{it}}\right)$	$\mathtt{VAC}_{it}$ is *Vacancies, Persons, seasonally adjusted*
ω_{it}	$100 \times [\ln(\mathtt{RWAGE}_{it}) - \ln(\mathtt{RWAGE}_{it-4})]$	$\mathtt{RWAGE}_{it}$ is *Wages & Salaries, Constant Prices, seasonally adjusted*
η_{it}	$\frac{\mathtt{UD}_{it}-\overline{\mathtt{UD}_i}}{\sigma^2_{\mathtt{UD},i}}$	$\mathtt{UD}_{it}$ is *Trade Union Density*
φ_{it}	$\frac{\mathtt{BRR}_{it}-\overline{\mathtt{BRR}_i}}{\sigma^2_{\mathtt{BRR},i}}$	$\mathtt{BRR}_{it}$ is *Average Gross Unemployment Benefit Replacement Rates*
ς_{it}	$\frac{\mathtt{EPL}_{it}-\overline{\mathtt{EPL}_i}}{\sigma^2_{\mathtt{EPL},i}}$	$\mathtt{EPL}_{it}$ is *Employment Protection*

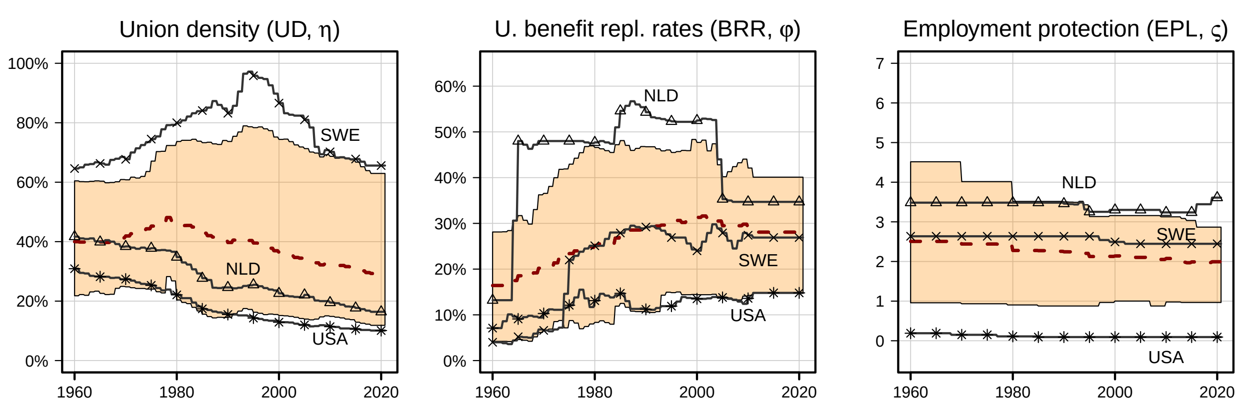
Notes: $\mathtt{POP}_{it}$ refers to *Total Population (Persons)*, $\mathtt{PRICE}_{it}$ refers to *Gross Domestic Product Deflator*.

Table D2: Sample Coverage in Different Models.

Countries / Model with...	Employment Rate	Unemployment Rate	Tightness
Australia	1966Q3-2020Q2	1964Q1-2020Q4	1978Q2-2020Q4
Austria	1970Q1-2020Q4	1970Q1-2020Q4	1970Q1-2020Q4
Belgium	1980Q4-2020Q4	1980Q4-2020Q4	no data
Canada	1961Q1-2020Q4	1961Q1-2020Q4	no data
Denmark	1980Q1-2020Q4	1980Q1-2020Q4	2009Q1-2020Q4
Finland	1965Q1-2020Q4	1960Q1-2020Q4	1960Q1-2020Q4
France	1960Q1-2020Q4	1960Q1-2020Q4	1995Q1-2020Q4
Germany	1991Q1-2020Q4	1991Q1-2020Q4	1991Q1-2020Q4
Great Britain	1971Q1-2020Q4	1971Q1-2020Q4	1970Q1-2020Q4
Italy	1960Q1-2020Q4	1960Q1-2020Q4	no data
Japan	1960Q1-2020Q4	1960Q1-2020Q4	1960Q1-2020Q4
Netherlands	1975Q1-2020Q4	1975Q1-2020Q4	1996Q1-2020Q4
Portugal	1995Q1-2020Q4	1995Q1-2020Q4	1995Q1-2020Q4
Spain	1961Q1-2020Q4	1976Q3-2020Q4	196Q1-2020Q4
Sweden	1960Q1-2020Q4	1960Q1-2020Q4	1960Q3-2020Q4
United States	1960Q1-2020Q4	1960Q1-2020Q4	2000Q1-2020Q4

Notes: The sample refer to data availability. In the estimation we loose four observations due to the applied transformation.

Figure D1: Labor Market Institutions in OECD Economies: Time Variation.



Notes: The figure shows the time variation in labor market institutions (union density, unemployment benefit replacement rate, and employment protection legislation) across all countries. The dark-red dashed line reports the mean and the orange colored region one standard deviation of the respective labor market institution. The y-axis refers to percent (union density, benefit replacement rate) or has no unit of scaling attached (employment protection legislation). The x-axis refers to time.

E. Further Empirical Results

This subsection presents further empirical results. First, we report the dynamic effects of government spending shocks conditional on the LMIs. Second, we provide robustness to various choices for the baseline model specification. Third, we investigate the issue of controlling for fiscal foresight. Fourth, we inspect other labor market indicators such as the unemployment rate and labor market tightness, given by the vacancy-unemployment ratio. Fifth, we examine cross-country heterogeneity and cluster countries into two groups according to their overall level of the LMIs or their consumption share in total GDP.

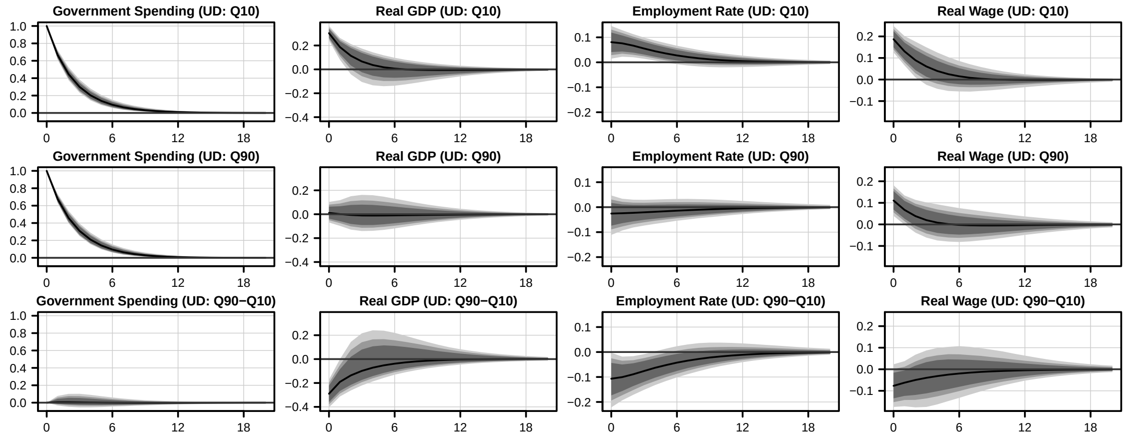
E.1 Dynamic Effects

In this subsection, we report the dynamic effects of government spending shocks conditional on the labor market institutions. In Figure E1 we show the impulse response functions of all the variables in the system by varying one LMI and keeping the others constant. They correspond to the first and last point in Figure 3 where we focus on the evolution of the marginal effects. Here, we focus more on the dynamic evolution of the IRFs themselves.

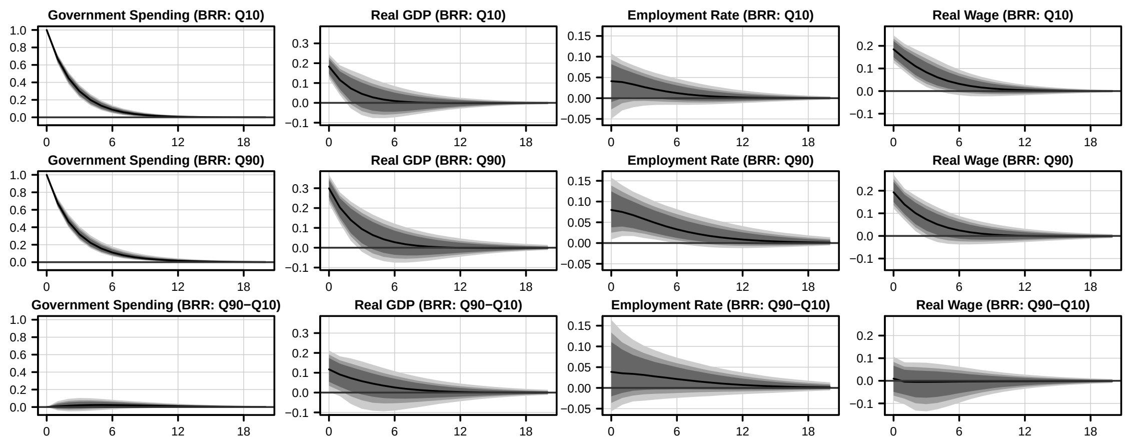
We report three panels, one for each LMI: union density (UD), unemployment benefit replacement rate (BRR), and employment protection legislation (EPL). We vary the value of the LMI as follows: Either we set it to the 10th (upper row) or the 90th (middle row) of its respective empirical distribution. The remaining LMIs are set to their median value. In the lower row, we also report the full posterior distribution of the differences to investigate statistical significance.

While confirming the overall picture and findings presented in the main text, a few remarks are in order. First, note that the difference in the impulse response of government spending is not affected at all by varying the intensity of the LMIs. This is reassuring in that the government spending shock and its transmission is not driving the results. Second, the dynamic propagation of the impulse responses confirms conventional wisdom. Output, employment, and real wages increase in response to a government spending shock. In some instances, the effect on the employment rate is not statistically significantly different from zero. This holds particularly true for the models using high levels of the LMIs. Third, we can examine whether the differences are statistically significant. From Figure 1a and Figure 1c we conclude that the differences are strongly statistically significant. For BRR, we only find that the difference for output is statistically significant.

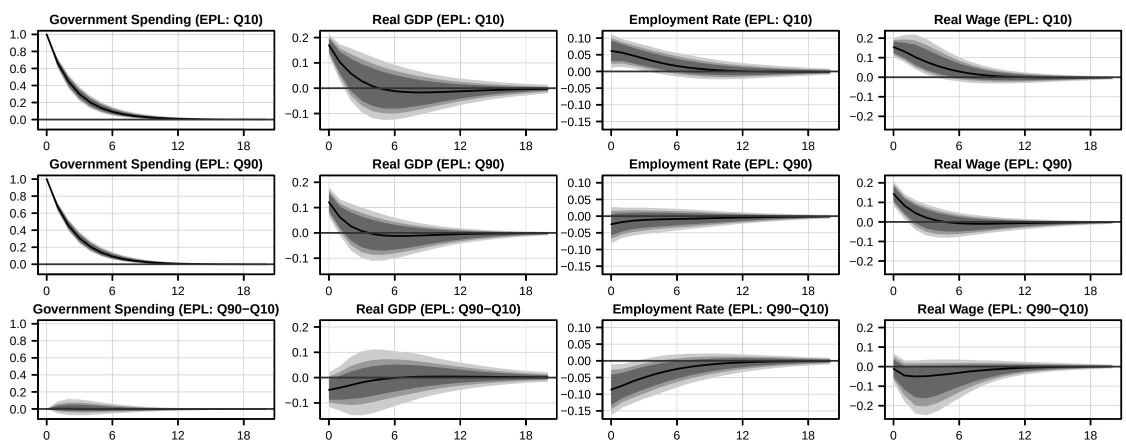
Figure E1: Dynamic Effects of Government Spending Shocks.



(a) Impulse Responses: Union Density.



(b) Impulse Responses: Unemployment Benefit Replacement Rate.



(c) Impulse Responses: Employment Protection Legislation.

Notes: This figure reports impulse response functions to a government spending shock. The black solid line refers to the median estimate while the gray areas refer to 68/80/90% credible sets. In each panel, we report the value of the LMI between its 10th (upper row) and 90th quantile (middle row) while setting the other LMI values to its median. We also report the difference (lower row). Panel (a) reports the conditional effects of union density, panel (b) the conditional effects of the unemployment benefit replacement rate, and panel (c) the conditional effects of employment protection legislation.

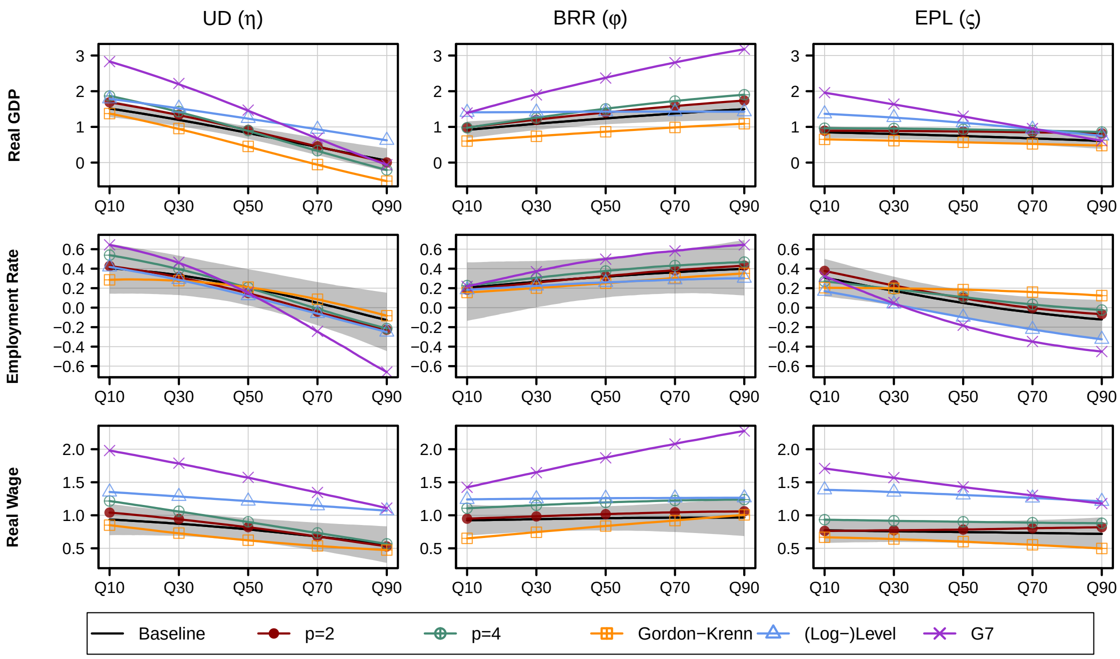
E.2 Robustness of Model Specification

In this subsection, we explore the robustness of the baseline specification in more detail. We conduct a number of robustness checks to the baseline model specification to assess the stability of the results. The results are presented in Figure E2, where we report the on impact (panel (a)) and one-year (panel (b)) fiscal multipliers. The baseline results (black solid line) with 80% credible sets (gray area) are those from the baseline model reported in Figure 3. Additionally, we add colored lines with distinct points to the figure for the alternative models.

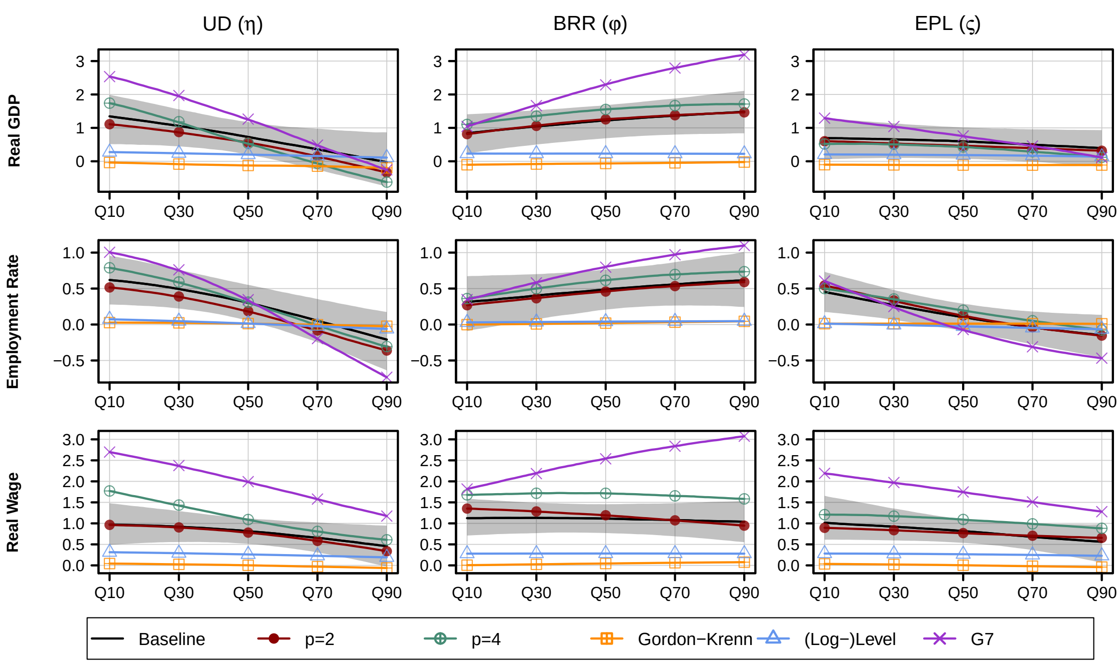
We have conducted the following robustness checks. First, we have varied the number of lags, using $p = 2$ and $p = 4$ lags. Results are robust to this choice, with the median estimates falling within the 80% posterior credible sets. In almost all cases, both median estimates are within the credible sets of the baseline model. As a next check, we follow Ramey (2016) and transform variables with a trend using the procedure proposed by Gordon and Krenn (2010). The original Gordon and Krenn (2010) procedure involves estimating a polynomial trend and dividing the respective series by this trend. However, due to the end-point problem in both the polynomial trend estimation and in using the Hodrick-Prescott filter, we use the Hamilton (2018) filter. By using this transformation instead of differentiation to remove the unit root, we retrieve similar fiscal multipliers. Some smaller differences arise, such as a stronger reduction in the fiscal multiplier for UD or a less pronounced multiplier for BRR in the case of real GDP. Furthermore, one-year multiplier for real wages is, in all cases, relatively constant and zero. Overall, however, the dynamics are similar to those of the baseline model, particularly for the contemporaneous multipliers. We also estimate the model in (log-)levels instead of growth rates. Results are robust to this choice as well, apart from minor differences. The strongest differences arise again with respect to the one-year fiscal multiplier of real wages, which is quite subdued.

As a last check, we restrict our sample to a smaller set of countries. We only estimate the model for the G7 countries: Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States. This comparison is interesting for two reasons. The first reason is to investigate the strong cross-country heterogeneity, while the second reason is that this comparison will be important when we control for fiscal foresight in the next subsection. Generally, the overall qualitative pattern is similar for the G7 countries. Particularly, it is not only similar, but fiscal multipliers are even more pronounced. We not only find higher fiscal multipliers (e.g., up to 0.6 for real GDP) but also more negatively sloped marginal effects.

Figure E2: Robustness: Model Specification and Sample.



(a) Fiscal Multipliers: On Impact.



(b) Fiscal Multipliers: One Year.

Notes: The figures show the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is the unemployment benefit replacement rate, and ς is employment protection legislation). The y-axis gives the size of the multiplier while the x-axis runs from the 10th to the 90th quantile in terms of the respective LMI. The multipliers are defined in Eq. (3.25) and shown for different horizons: on impact (upper panel) and one-year (lower panel) multiplier. The black solid line denotes the median and the gray area refers to the 80% credible set of the baseline model. The colored lines with different points refer to alternative models.

To summarize, the baseline model is robust to a number of choices in a qualitative sense. Quantitatively, results are, in some cases, different. When using different transformations, effects are somewhat subdued, but when looking at a smaller country set, the effects are even magnified. The overall pattern, however, is remarkably robust.

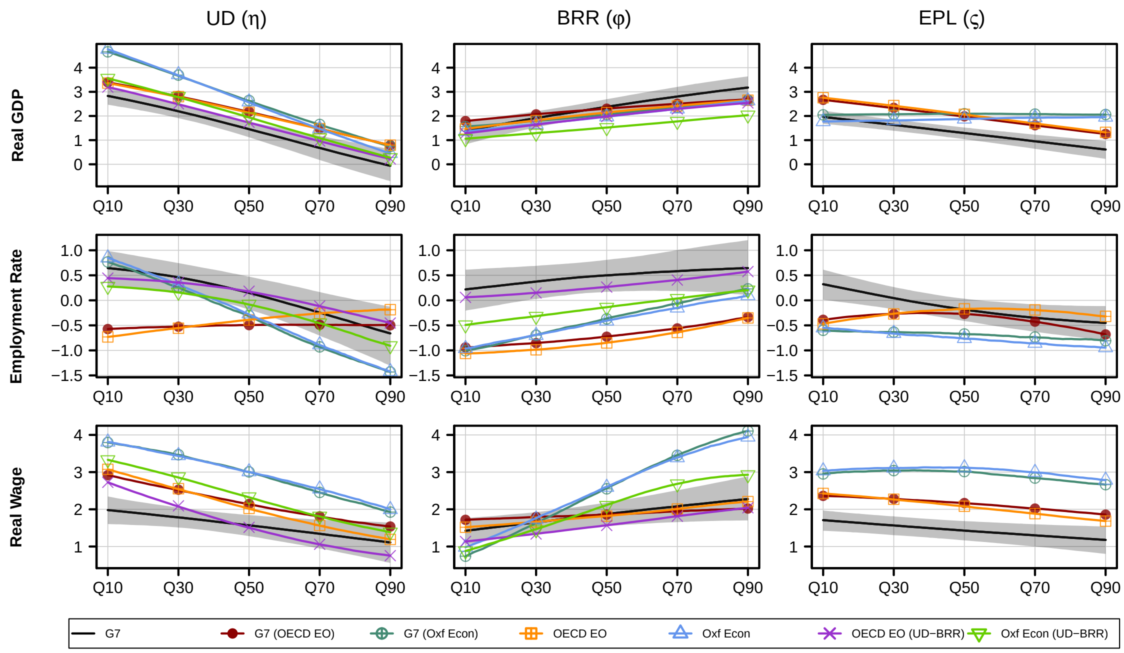
E.3 Controlling for Fiscal Foresight

In this subsection, we explore the issue of fiscal foresight in more detail. As pointed out by Ramey (2011) or Leeper, Walker and Yang (2013), econometric models can suffer from informational insufficiency due to a misalignment between the information sets of the economic agents and the econometrician. The identification of government spending shocks can be clouded by potential anticipation effects of fiscal policy changes due to their lagged implementation. This is particularly the case for structural vector autoregressions (SVARs) identifying fiscal shocks because they rely mostly on a small number of endogenous variables, as in our case of four variables. In order to add this missing information to the model, the literature has suggested two strategies to circumvent this issue. Either one controls directly for anticipation effects by including expectations data to the model or by enlarging the information set to mirror that of the economic agents, which they use to predict fiscal policy changes.

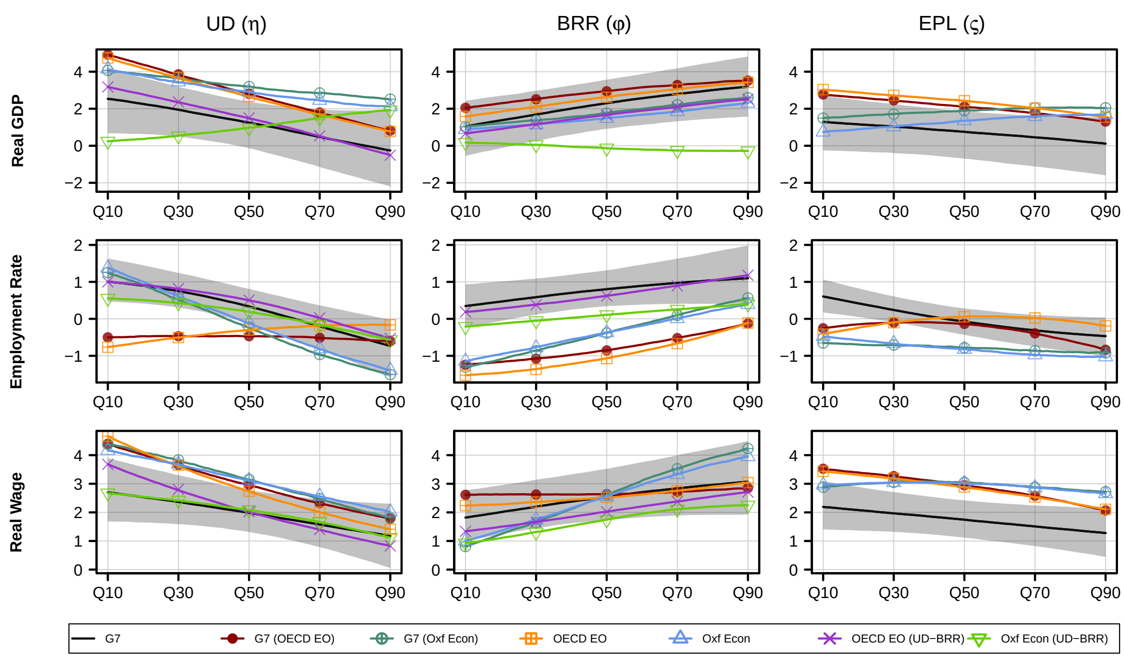
Both approaches are problematic in our setup. A large information set and data on expectations for a broad panel of countries over a long time period (1960Q1 to 2020Q4 in this case) are hardly possible to gather. However, by reducing the sample of countries and the time frame, it allows us to control for fiscal foresight. To fix ideas, we include government spending forecasts, Δg_{it}^{e}, to the vector of endogenous variables, which yields $\boldsymbol{y}_{it} = (g_{it}, \Delta g_{it}^{e}, x_{it}, er_{it}, \omega_{it})'$.[3] We rely on two different data sources for the measurement of government spending forecasts, following suggestions in the literature on cross-country analyzes (Auerbach and Gorodnichenko, 2013; Born, Juessen and Müller, 2013; Born, Müller and Pfeifer, 2020; Ilori, Paez-Farrell and Thoenissen, 2022). First, we rely on government spending forecasts provided by Oxford Economics, a large forecasting firm that serves 1,500 clients, among them international corporations, financial institutions, government organizations, and universities. These forecasts are available starting in the late 1990s. Second, we also use semi-annual professional forecasts by the OECD disseminated in its Economic Outlook. These forecasts are consistently available for all G7 countries since the mid 1980s[4] and tend to perform comparably

[3] Results are robust when exchanging the order of the variables, ordering the government spending forecast first.
[4] For the remaining countries in our sample, we have forecasts available starting in the mid 1990s.

Figure E3: Controlling for Fiscal Foresight.



(a) Fiscal Multipliers: On Impact.



(b) Fiscal Multipliers: One Year.

Notes: The figure show the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is the unemployment benefit replacement rate, and ς is employment protection legislation). The y-axis gives the size of the multiplier while the x-axis runs from the 10th to the 90th quantile in terms of the respective LMI. The multipliers are defined in Eq. (3.25) and shown for different horizons: on impact (upper panel) and one-year (lower panel) multiplier. The black solid line denotes the median and the gray area refers to the 80% credible set of the baseline model. The colored lines with different points refer to alternative models.

well compared to other professional forecasts (Auerbach and Gorodnichenko, 2012). These forecasts are prepared twice a year, in June and December. We follow the suggestion in Ilori, Paez-Farrell and Thoenissen (2022) and interpolate the semi-annual series to a quarterly series using linear interpolation. Both series are used in growth rates rather than levels due to irregular base-year changes for the countries in our sample.

We have to restrict our sample in terms of country coverage and time frame. We reduce the set of countries to the G7 countries: Canada, France, Germany, Italy, Japan, United Kingdom, and United States. For a comparison, we re-estimate the model over the full sample using the G7 countries. The results are presented in Figure E3, where we report the on impact (panel (a)) and one-year (panel (b)) fiscal multipliers. The model outcome of the G7 countries is reported by the median (black solid lines) together with 80% credible set (gray area). We have discussed the comparison to the baseline model in the previous subsection. As a first test, we cut the sample to the availability of government spending forecasts from the OECD Economic Outlook (labeled *G7 (OECD EO)*) and Oxford Economics (labeled *G7 (Oxf Econ)*) without including the forecasts themselves in the model. Shortening the sample to start in the mid 1980s (OECD) or late 1990s (Oxf Econmics) also affects the cross-sectional dimension. In the OECD model, we have five countries left (Germany, France, Italy, Japan, and the United Kingdom)[5], while in the Oxford Economics model we only have four countries left (France, Italy, Japan, and the United Kingdom).[6] The reason is that for the other countries no time variation is left in one of the LMIs, and thus we have excluded these countries from the model. This is driven by the relatively weak time variation in the EPL, a point to which we return later. On the same country coverage and sample size, we have then added the government spending forecasts to the model (labeled as *OECD EO* and *Oxf Econ* in the figure). A few interesting observations arise. First, overall the findings are quite robust to the comparison model. For output and real wages, the marginal effects are even more pronounced, while for employment more subdued. Qualitatively, the shape of the marginal effects is robust with some exceptions (mostly for the employment rate). Second, the marginal effects for the models including and excluding the government spending forecasts (e.g., *G7 (OECD EO)* and *OECD EO*) are extremely similar. Hence, differences across models seem not to be driven by anticipation effects but rather due to subsample stability and heterogeneity issues. This point is also highlighted by Ellahie and Ricco (2017).

[5] In more detail, the estimation sample spans 1986Q3 to 2020Q4 except for Germany which only starts in 1992Q1. Due to the unavailability of data on wages, Germany also starts in the baseline model only at this point, as detailed in Table D2.

[6] In more detail, the estimation sample spans 2000Q1 to 2020Q4.

As a last check, we exclude the EPL from the set of LMIs and re-estimate the model. This leaves us with all seven G7 countries in the sample (labeled *OECD EO (UD-BRR)* and *Oxf Econ (UD-BRR)*). The results are qualitatively stable. Interestingly, we even find that the subdued response of the employment rate for the *G7 (OECD EO)* and *OECD EO* model vanishes when we extend the sample. This is a clear indication that not anticipation effects are an issue, but rather sub-sample (in-)stability. In some instances, the marginal effects of fiscal multipliers are more subdued when using the Oxford Economics government spending forecasts (e.g., output for the one-year multiplier). However, we do note that the sample is relatively short, and we lose a lot of information in the LMIs, which could drive the results as well.

E.4 Different Labor Market Indicators

In this subsection, we explore other labor market indicators. We exchange the employment rate in the baseline model for either the unemployment rate or the labor market tightness (v_{it}/u_{it} locus). Similar to Figure 3, we report the contemporaneous and one-year fiscal multipliers.

Results are provided in Figure E4. We report the results for the unemployment rate in panel (a) and for the labor market tightness in panel (b). The results for the unemployment rate confirm our main findings. We find a strongly downward-shaped curve for the marginal effects for UD and now an upward-shaped curve for the unemployment rate. The effects on the EPL are dampened, while the output multiplier for BRR is upward sloping. Similar to the baseline model, there are no statistically significant differences for the unemployment rate for BRR. The real wage does not strongly react to the stringency of the LMIs.

The model including labor market tightness is characterized by relatively large volatility around the estimates, pointing to instability. We have also added the employment rate to the model, as in the baseline. While the multipliers on impact are strongly centered around zero, we see an attenuation of marginal effects for the one-year multiplier. However, none of these outcomes is statistically significant, and we are thus not too confident about the outcomes of the model.

In Figure E5 we report the implied volatilities and the change in the volatilities for the model featuring unemployment. In comparison to the baseline model, we do not find stark differences. The only difference is that the volatility for the unemployment rate increases when moving from the low to the high regime when looking at UD. Similarly, there is a sign change for the change in the volatilities when looking at the unemployment rate.

E.5 Cross-Country Heterogeneity

In this subsection, we explore heterogeneous effects utilizing between-country variation since the analysis in the main text is based on within-country variation of the LMIs. When discussing Figure 1 we have already noticed that there is considerable cross-country heterogeneity in the LMIs. We conduct two robustness exercises: First, we investigate whether effects differ in countries with more rigid LMIs deployed (e.g., in Scandinavia) to countries with more flexible labor markets (e.g., Anglo-Saxon countries). Second, we examine whether there are cross-country differences with respect to the consumption share in GDP. For both approaches we cluster the countries in two groups. Then, we re-estimate the model in Eq. (4.1) for both groups. We only allow for two groups to have enough variation in both country groups to estimate the IPVAR.

The clustering for the first approach is done via k-means clustering of the LMIs.[7] We standardize the data (over all countries) before using the algorithm so that no variable has a stronger influence due to its scaling. In case a country is not classified entirely into one group, we apply a 50% rule: If more than 50% of the observations of one country are classified into one group, the country is classified into the same group. From the clustering algorithm, we get two groups which we label as follows. *Upper group*: Austria, Belgium, Denmark, Finland, the Netherlands, Portugal, and Sweden. *Lower group*: Australia, Canada, Germany, Great Britain, Italy, Japan, Spain, and the United States. The groups align well with various definitions of welfare regimes and are depicted in Figure E6. The second classification is done via average household consumption shares (as percent of GDP). Those range from 43.55% to 68.06% and yield the following groups: *Upper group*: Canada, France, Great Britain, Italy, Japan, Portugal, Spain, and the United States. *Lower group*: Australia, Austria, Belgium, Denmark, Finland, Germany, the Netherlands, and Sweden.

In Figure E7, we examine the fiscal multipliers based on both within- and between-country variation, where we report the on impact (panel (a)) and one-year (panel (b)) fiscal multipliers. The baseline results (black solid line) with 80% credible sets (gray area) are those from the baseline model reported in Figure 3. There is considerable cross-country heterogeneity, although the overall picture is qualitatively similar. This exercise reveals some interesting insights. First, the upper consumption share group and the lower LMI group yield generally larger fiscal multipliers than the baseline model. The marginal effects along the stringency of the LMIs also show more time variation, i.e., a stronger attenuation effect of fiscal multipliers for UD or EPL. The effect on the shape of the fiscal multipliers is particularly pronounced for the upper consumption

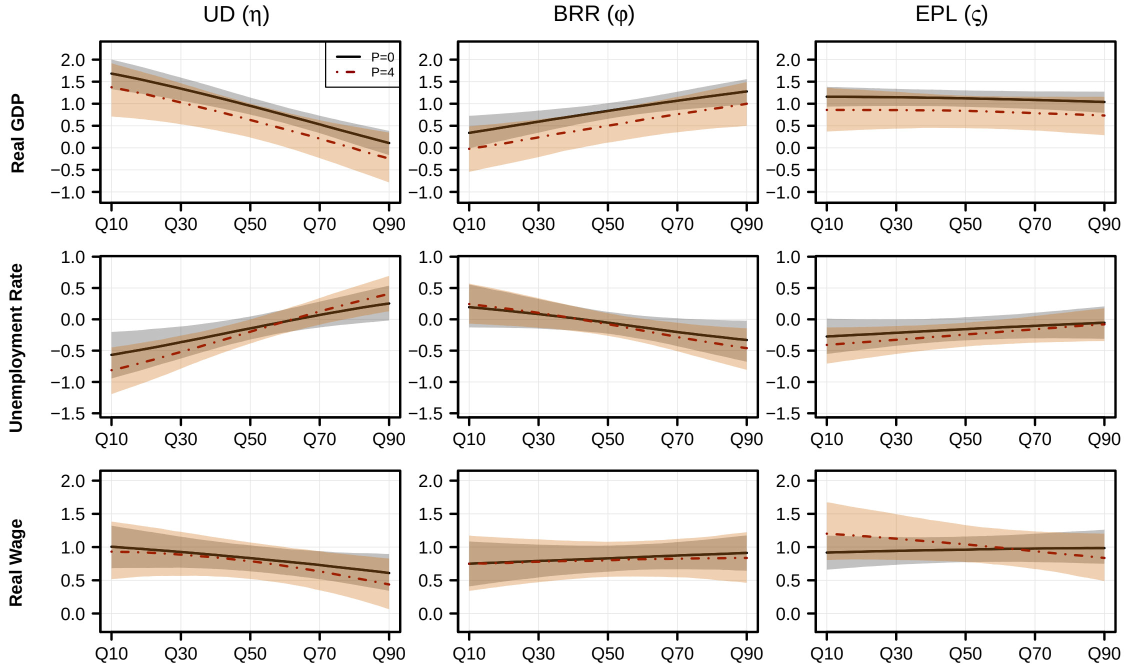
[7] This is a frequently employed clustering algorithm based on the idea that each observation belongs to the cluster with the nearest mean (or cluster centroid).

group, whereas we find less strong evidence on the shape of the marginal effects in the lower LMI group. On the contrary, the lower consumption share group and the upper LMI group yield less pronounced fiscal multipliers. Similarly, the marginal effects are dampened for this group. This holds not only for the impact multipliers, but also for the one-year multipliers.

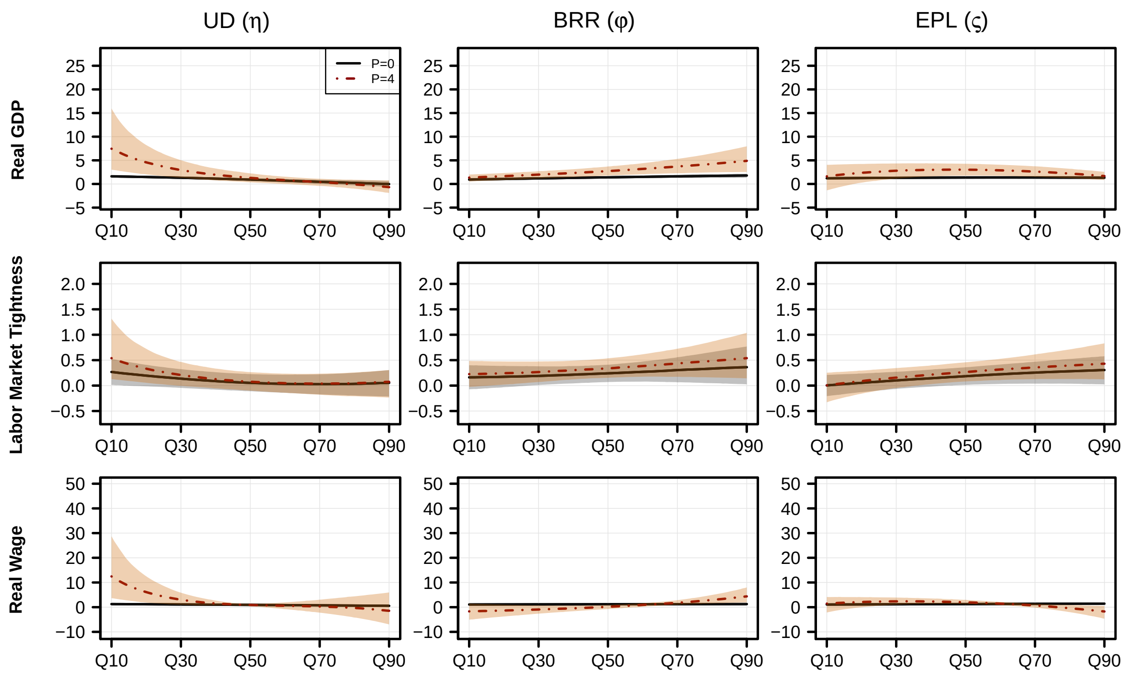
These results are in line with conventional predictions. Countries with a higher consumption share are more strongly dependent upon output fluctuations due to demand-side shocks, such as a government spending shock. Similarly, in countries with a generally lower level of LMIs, the findings are more pronounced. This provides an indication that the *level* of the LMIs matters as well, and not only their *change* over time. These results must be interpreted with a grain of salt since we use the same identifying restrictions across all models.

Overall, the results outlined here strengthen the implications of our baseline results as of Section 4. In the "upper" group of countries, cyclical policies do not have a strong effect on labor market variables. Cyclical policies still affect the fiscal multipliers in the "lower" group of countries, with a clear downward-sloping effect along the within-country variation.

Figure E4: Fiscal Multipliers Using Other Labor Market Indicators.



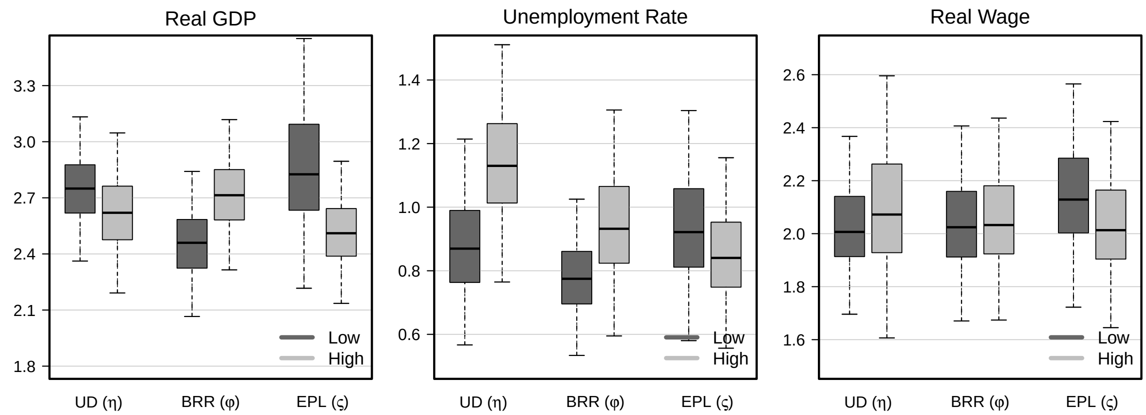
(a) Unemployment Rate.



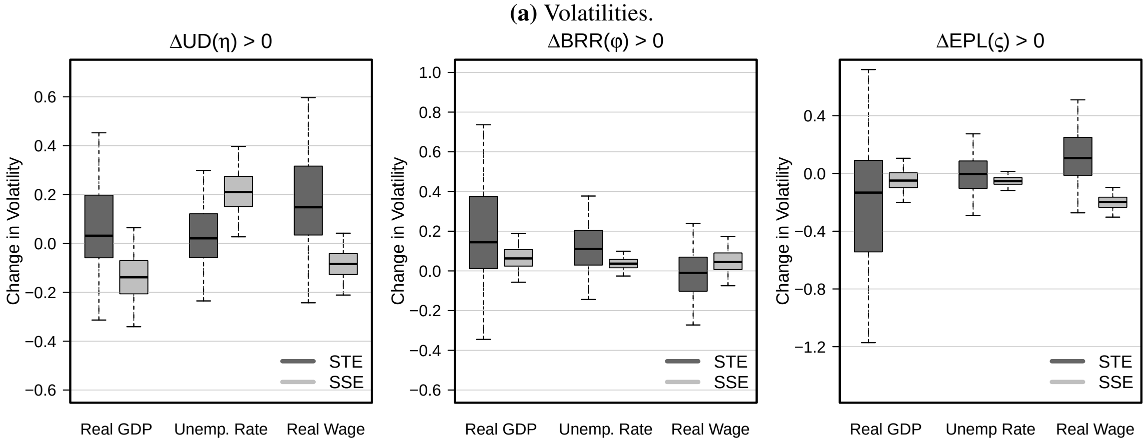
(b) Labor Market Tightness (v_{it}/u_{it}).

Notes: The figure show the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is the unemployment benefit replacement rate, and ς is employment protection legislation). The y-axis gives the size of the multiplier while the x-axis runs from the 10th to the 90th quantile in terms of the respective LMI. The multipliers are defined in Eq. (3.25) and shown for different horizons: on impact ($\mathcal{P}$ = 0, solid black line) and one-year ($\mathcal{P}$ = 4, dashed-dotted red line) multiplier. The lines denotes the median and the colored area refers to the 80% credible set.

Figure E5: Volatilities.



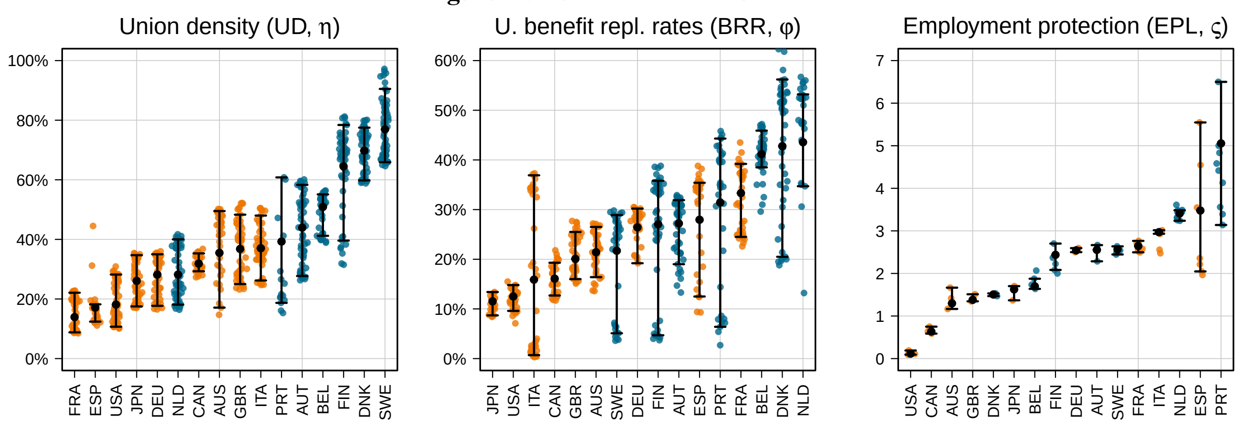
(a) Volatilities.



(b) Change in Volatilities.

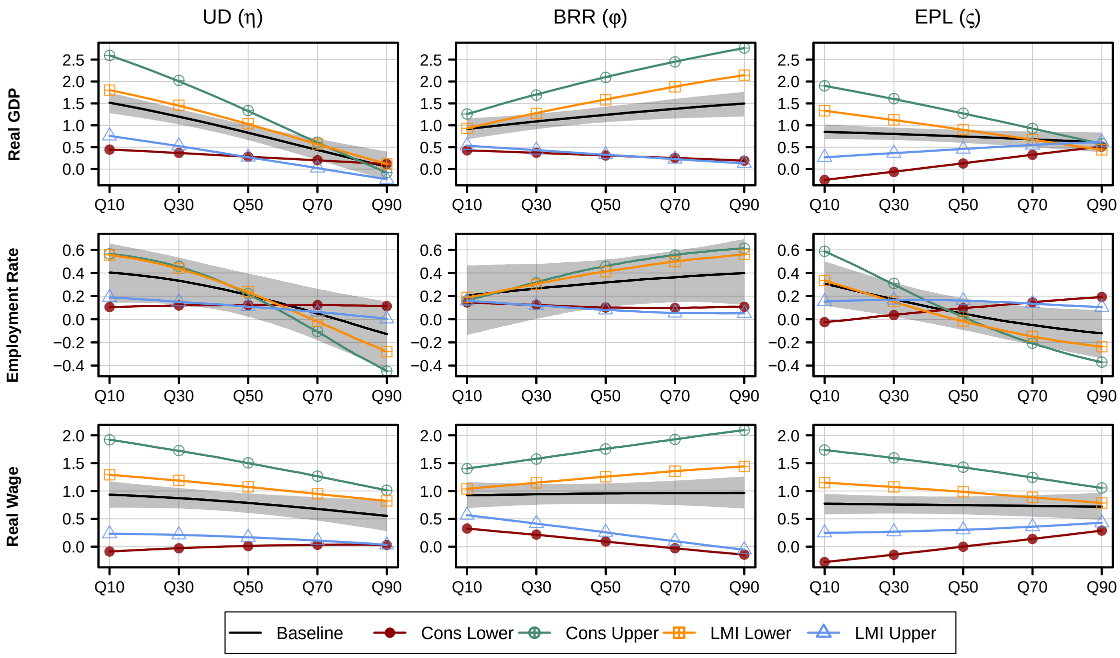
Notes: The upper panel shows the standard deviation of the respective macroeconomic variable in a regime with low (10th quantile) and high (90th quantile) labor market institution (LMIs) while the remaining LMIs are at their median. The lower panel shows the change in the standard deviation of the respective macroeconomic variable when going from the high (90th quantile) to the low (10th quantile) regime. STE refers to the *shock transmission effect*, while SSE refers to the *shock size effect* as depicted in Eq. (4.7). The LMIs under consideration are union density (UD, η), unemployment benefit replacement rate (BRR, φ), and employment protection legislation (EPL, ς).

Figure E6: Classification of Countries.

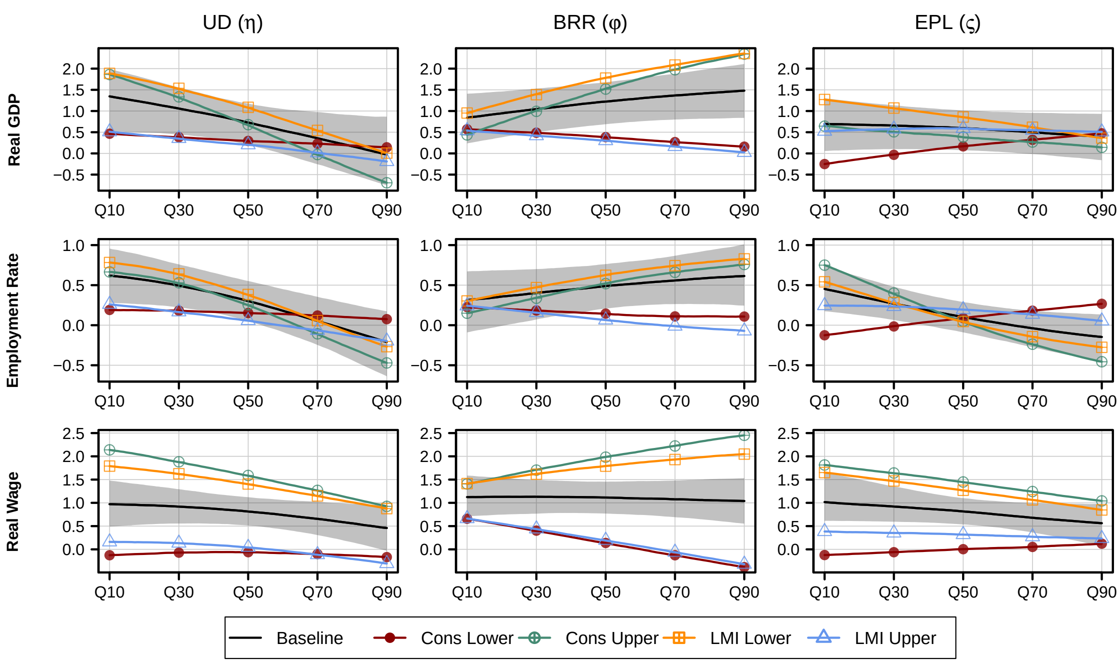


Notes: Each figure shows the mean of each labor market institution (union density, unemployment benefit replacement rate, and employment protection legislation) for each country, together with the 10th and 90th quantile of its distribution. The points are observed data for the respective country. Color shadings differentiate countries belonging to the *Upper Group* (blue) and *Lower Group* (orange).

Figure E7: Cross-Country Heterogeneity.



(a) Fiscal Multipliers: On Impact.



(b) Fiscal Multipliers: One Year.

Notes: The figure show the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is the unemployment benefit replacement rate, and ς is employment protection legislation). The y-axis gives the size of the multiplier while the x-axis runs from the 10th to the 90th quantile in terms of the respective LMI. The multipliers are defined in Eq. (3.25) and shown for different horizons: on impact (upper panel) and one-year (lower panel) multiplier. The black solid line denotes the median and the gray area refers to the 80% credible set of the baseline model. The colored lines with different points refer to alternative models.

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