Labor Market Institutions, Fiscal Multipliers, and Macroeconomic Volatility

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Abstract

We empirically examine how labor market institutions (LMI) – union density (UD), unemployment benefit replacement rates (BRR), and employment protection legislation (EPL) – shape fiscal multipliers and macroeconomic volatility. Our theoretical model predicts that stringent labor market institutions attenuate fiscal spending multipliers and reduce macroeconomic volatility, with UD showing the strongest effects, followed by EPL and BRR. These predictions are validated through an interacted panel vector autoregressive (IPVAR) model estimated for 16 OECD countries. Results underscore that stringent LMIs render cyclical fiscal policies less effective for macroeconomic stabilization. The findings highlight heterogeneities across LMIs, and while some LMIs mitigate the contemporaneous impact of fiscal shocks, others amplify the propagation mechanism. Our findings offer important implications for policymakers navigating labor market and fiscal policy trade-offs.

Keywords: Fiscal policy; Labor market institutions; Interacted panel VAR

JEL Codes: E62, C33, J21, J38

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1. Introduction

Large economic downturns, such as the global financial crisis 2008 or the Covid-19 pandemic, trigger strong fiscal responses to stabilize the economy. Policymakers are interested in the efficacy of fiscal expansions, particularly whether they support the labor market.[1] However, the academic discussion on fiscal multipliers mostly centers on output and less on (un)employment multipliers (Monacelli, Perotti and Trigari, 2010). The debate strongly focuses on the state-dependency of fiscal multipliers (e.g., in *recessions* vs. *expansions*).[2] The potential state-dependency cannot only arise due to cyclical fluctuations but also due to structural features of the economy. As such, labor market frictions critically affect labor market outcomes beyond business cycle fluctuations. While it is a well-established fact that labor market institutions (LMI) matter for business cycles (see, for instance, Gnocchi, Lagerborg and Pappa, 2015 or Abbritti and Weber, 2018), there is less evidence on the ability of different LMIs to affect fiscal (labor market) multipliers.

This paper contributes to this literature by examining how different labor market institutions affect the size of fiscal multipliers on the labor market and their role as determinants of business cycle fluctuations. We consider a range of labor market rigidities such as union density (UD), unemployment benefit replacement rates (BRR), and employment protection legislation (EPL). This allows us to assess the role of different LMIs in shaping the effectiveness of fiscal policy and in judging their ability to dampen cyclical fluctuations. Eventually, we also evaluate the degree of complementarity (or substitutability) among these LMIs and cyclical spending policies. This enables a direct comparison of structural and cyclical policies of fiscal multipliers and the ability to smooth macroeconomic fluctuations, which has not been addressed in a joint framework featuring multiple LMIs.

We start by developing a theoretical model to assess qualitatively the role of LMIs in shaping fiscal spending multipliers and macroeconomic volatility. We combine a standard New Keynesian model with search and matching frictions in the labor market (Thomas and Zanetti, 2009; Christoffel, Kuester and Linzert, 2009). We measure union density by means of workers' bargaining power within the wage negotiations, the unemployment benefit replacement rate by means of subsidies to the unemployed, and employment

[1] In the wake of the financial crisis 2007-09, the American Recovery and Reinvestment Act (ARRA) (Romer, 2009) explicitly promoted to raise employment. The Coronavirus Aid, Relief, and Economic Security (CARES) Act, on the other hand, expanded unemployment insurance in order to help unemployed workers (Ganong, Noel and Vavra, 2020; Ganong et al., 2024). The European Union Recovery Instrument (or Next Generation EU) allowed member states flexibility in terms of the fiscal rules and supported national relief programs.

[2] See, inter alia, the contributions by Blanchard and Perotti (2002), Ramey (2011), Corsetti, Meier and Müller (2012), Auerbach and Gorodnichenko (2012), Ilzetzki, Mendoza and Végh (2013), Leeper, Walker and Yang (2013), Ramey and Zubairy (2018), Born, Müller and Pfeifer (2020), Born et al. (2024) and Jo and Zubairy (forthcoming).

protection by the level of firing costs per displaced worker. In a search and matching framework on the labor market, a positive change in the unemployment benefit replacement rates and union density crucially affects real wages positively, which attenuates the expansion of employment following an expansionary fiscal shock.[3] Employment protection, however, affects directly job creation and destruction in opposing directions. Higher employment protection leads then to an attenuated response of the fiscal multiplier on employment. Taken together, the model predicts that more stringent LMIs attenuate both fiscal spending multipliers and macroeconomic volatility. The strongest mitigation emanates from employment protection, followed by union density and the unemployment benefit replacement rate.[4] To evaluate these channels, we conduct an impulse response matching exercise after the empirical outcomes are presented.

In a next step, we confront the predictions of the theoretical model by estimating a semi-structural Bayesian interacted panel vector autoregressive (IPVAR) model for 16 OECD countries (Towbin and Weber, 2013; Sá, Towbin and Wieladek, 2014). Conditional on the interaction terms (the three LMIs in our case), we identify and trace out the dynamic effects of a government spending shock and their effect on macroeconomic volatility. The structural interpretation of the econometric model relies on two main building blocks. First, our approach to identification relies on an implementation lag of government spending as outlined in Blanchard and Perotti (2002) and applied in a panel setting by Ilzetzki, Mendoza and Végh (2013).[5] Second, we assume the exogeneity of the LMIs with respect to the interacted current and lagged values of the endogenous variables in the system. LMIs change slowly over time and correlations to cyclical variables are rather low, which renders the choice of the interacted panel VAR model particularly convenient.[6] Due to the panel dimension and the constrained availability of the LMIs, we do not estimate the theoretical model. However, we do not only show the qualitative similarity between the outcomes of the DSGE model and the IPVAR but also perform impulse response matching to look into the exact channels driving the differences in the shock transmission.

[3] This is contrast to Ghassibe and Zanetti (2022), which investigate goods market tightness as a potential source of the size of the fiscal multiplier. However, similar to their study, the fiscal multiplier decreases with higher goods market tightness.

[4] Recent contributions by Challe (2020), McKay and Reis (2021), or Kekre (2023) highlight the precautionary savings motive in models with heterogeneous agents who are exposed to unemployment risk. In these models, higher unemployment benefits cause an additional aggregate demand channel which boosts output and employment. We abstract from these developments in the model presented here.

[5] In a robustness exercise, we control for fiscal foresight, which can pose problems when identifying government spending shocks as pointed out by Ramey (2011), Auerbach and Gorodnichenko (2012), or Leeper, Walker and Yang (2013). Although this issues has been resolved in panel settings (see, for instance, Born, Müller and Pfeifer, 2020) the necessary data is not available for our full sample. Hence, we only provide a robustness check to a sample with less country and time coverage.

[6] Note that we estimate the model on a quarterly frequency while the LMI interaction terms are on a yearly frequency. The model exploits only within-country variation by standardizing the interaction terms appropriately.

The results can be summarized as follows. The empirical model grossly corroborates the qualitative theoretical predictions. In terms of fiscal multipliers, we find the strongest attenuation of output fiscal multipliers for UD, followed by EPL. Similarly, both UD and EPL show a sizable reduction in the employment fiscal multiplier along their stringency. The attenuation for the fiscal multiplier of the real wage is less pronounced and without statistical significance. On the contrary, we find a slight increase in the output and employment multipliers for the BRR with varying statistical significance. Although outside of our model, we explain these findings through a reduction of the precautionary savings motive induced by unemployment risk, which leads to an additional demand channel. These differences highlight the heterogeneous outcomes of fiscal spending effectiveness once stringent LMIs are in place. For macroeconomic volatility, we find that UD and EPL lead to a decrease in volatility for output, while BRR increases the volatility for output and employment. The distinct quantitative effects on cyclical volatility are due to the fact that the extent of EPL attenuates macroeconomic volatility by mitigating both the propagation mechanism and the contemporaneous impact of shocks. The extent of union density and the size of the unemployment benefit replacement rates, in turn, exacerbate the propagation mechanism of shocks while moderating their contemporaneous impact.

This paper contributes to two strands of the literature. First, we add to the vast literature which investigates the effects of fiscal policy shocks.[7] An important and lively area of research in this field is the investigation of the potential *state-dependent* effects of fiscal spending policy. This was initiated by the seminal paper by Auerbach and Gorodnichenko (2012) which examines the state-dependency of fiscal multipliers in recessions and expansions.[8] Most of these papers focus exclusively on cyclical factors affecting the state. To this end, we extend this analysis in the direction of structural policies and investigate how labor market rigidities shape the effectiveness of fiscal multipliers. Additionally, we relate to a number of contributions that focus more strongly on labor market outcomes/multipliers (Monacelli, Perotti and Trigari, 2010; Brückner and Pappa, 2012).[9]

[7] Important contributions in the realm of identifying fiscal spending shocks in a US and panel-country context are, inter alia, Blanchard and Perotti (2002), Mountford and Uhlig (2009), Ramey (2011), Corsetti, Meier and Müller (2012), Born, Juessen and Müller (2013), Ilzetzki, Mendoza and Végh (2013), Miyamoto, Nguyen and Sheremirov (2019), Born, Müller and Pfeifer (2020), or Ilori, Paez-Farrell and Thoenissen (2022).

[8] In subsequent contributions, others show state-dependency due to the accommodation of monetary policy (Christiano, Eichenbaum and Rebelo, 2011; Coenen et al., 2012; Fernández-Villaverde et al., 2015; Rendahl, 2016), the exchange rate regime (Born, Juessen and Müller, 2013; Born et al., 2024), the zero lower bound (Ramey and Zubairy, 2018; Di Serio, Fragetta and Gasteiger, 2020), fiscal financing (Hagedorn, Manovskii and Mitman, 2019), household leverage (Demyanyk, Loutskina and Murphy, 2019), financial turmoil (Bernardini, De Schryder and Peersman, 2020), and asymmetric effects due to downward nominal wage rigidity (Shen and Yang, 2018; Jo and Zubairy, forthcoming) or due to market incompleteness (Barnichon, Debortoli and Matthes, 2022).

[9] In this context, Ball, Jalles and Loungani (2015) highlight that the link between GDP and the labor market strongly depends on the idiosyncratic labor market institutions in place in a given economy.

Our paper also complements the literature on the macroeconomic effects of labor market regulation. An increasing theoretical (Christoffel, Kuester and Linzert, 2009; Thomas and Zanetti, 2009; Zanetti, 2009; Zanetti, 2011; Campolmi, Faia and Winkler, 2011) and empirical literature (Gnocchi, Lagerborg and Pappa, 2015; Abbritti and Weber, 2018; Hantzsche, Savsek and Weber, 2018; Cacciatore et al., 2021) assesses the macroeconomic implications of market reforms, such as the removal of labor and product market frictions. The empirical evidence points broadly to an attenuation of the *unconditional* business cycle volatility[10], while there is less work on the *conditional* effects to exogenous shocks. Notable exceptions are Abbritti and Weber (2018), Hantzsche, Savsek and Weber (2018), and Cacciatore et al. (2021). While the first two papers investigate the conditional effects of oil prices, world demand, and financing shocks, Cacciatore et al. (2021) is closer to our work. They analyze the role of employment protection legislation for fiscal spending shocks. However, in contrast to their work, we highlight significant heterogeneities across different LMIs. While we corroborate their findings in terms of the EPL, we find even stronger effects for UD, which works through differences in matching elasticities and not only job loss probability in the case of EPL. Furthermore, the investigation of the BRR even suggests an additional aggregate demand effect (Challe, 2020; McKay and Reis, 2021; Kekre, 2023).

The remainder of the paper is structured as follows. In Section 2 we provide a descriptive overview of LMIs across selected OECD countries. Section 3 discusses the theoretical model and presents its main predictions. Section 4 shows the connection between the theoretical and the empirical model, and presents the results of the econometric analysis. Finally, Section 5 concludes.

2. Structural Labor Market Indicators in OECD Economies

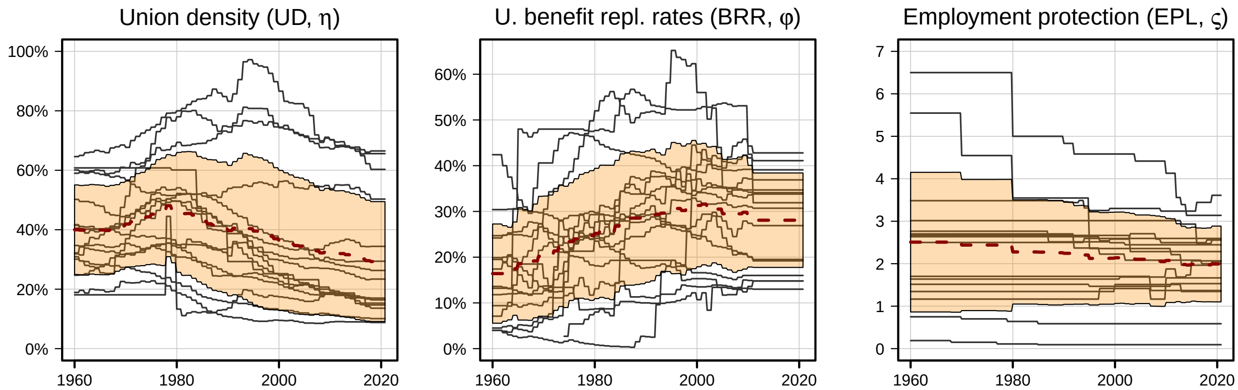
We focus on the following labor market institutions in OECD countries[11]: union density (UD, η), the unemployment benefit replacement rate (BRR, φ), and employment protection legislation (EPL, ς).[12] For an overview, we provide in Figure 1 and Figure 2 descriptive evidence on the time variation and cross-country heterogeneity.

[10] See, inter alia, also the contributions by Merkl and Schmitz (2011), Rumler and Scharler (2011), or Ferraro and Fiori (2023).

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

[12] There are, of course, many other important structural labor market characteristics which affect the transmission channel of fiscal shocks. Prominent ones concern the degree of labor market openness to foreign workers (see, for instance Amuedo-Dorantes and Rica, 2013; Godøy, 2017; Schiman, 2021), the declining trend in labor productivity (see, for instance Policardo, Punzo and Sánchez Carrera, 2019; Li et al., 2021), or demographic changes (see, for instance Docquier et al., 2019). We limit our analysis to the categories mentioned above due to data availability.

Figure 1: 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.

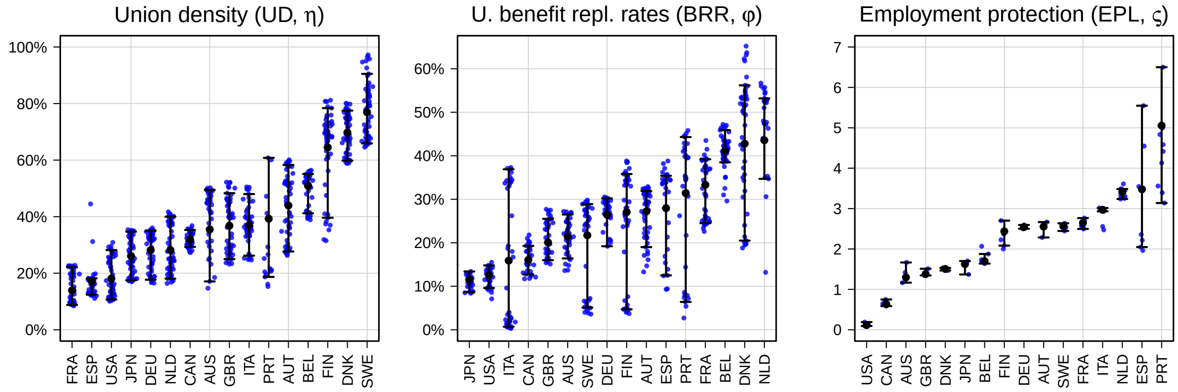
UD is primarily derived from survey data or adjusted administrative data for non-active and self-employed members. It represents the ratio of trade union members among wage and salary earners. Higher values indicate greater trade union influence in wage negotiations. In the 1970s, UD distribution shifted upward compared to the 1960s, followed by a steady decline in union membership rates. While this trend is reflected in the mean (dashed dark red line), it conceals rising cross-country dispersion from the 1980s onward (15% in France and 80% in Sweden).

The BRR measures the proportion of income maintained during unemployment as a ratio of net household income during unemployment to income before job loss. Higher values indicate more generous systems. BRR steadily increased until the 2000s but has since seen a slight median decline across countries. A narrow distribution in the 1960s widened in the 1990s and 2000s, with recent trends showing some convergence across countries.

The EPL index measures regulatory strictness on dismissals and temporary contracts, based on regulations in force on January 1 each year. Higher values indicate stricter employment protection. While the median EPL index has stayed constant, its dispersion has significantly narrowed, reflecting convergence, especially among EU countries. In contrast, non-EU countries showed little change over time.

The LMIs affect different macroeconomic variables. While UD and BRR affect employment, vacancies, and unemployment indirectly through their effect on the price of labor, EPL is likely to affect those variables

Figure 2: Labor Market Institutions in OECD Economies: Cross-Country Variation.



Notes: The figure shows the cross-country variation in labor market institutions (union density, unemployment benefit replacement rate, and employment protection legislation) for each country. The black dot reports the mean and the wishkers correspond to the 10th and 90th quantile of the distribution. The blue points are observed data. 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 each individual country.

directly. In what follows, we study in detail the implications of different LMIs as determinants of differences in fiscal spending multipliers and macroeconomic volatility.

3. The Theoretical Model

In the theoretical model, we merge the structure of a Diamond-Mortensen-Pissarides model with a standard real business cycle framework and rely on the setting put forward by Merz (1995), Andolfatto (1996) and Krause and Lubik (2007), and Monacelli, Perotti and Trigari (2010). The model is intended to be parsimonious and focus on the role played by labor market institutions. We consider various extensions of the set-up that can accommodate more complex interactions in Appendix B.

We assume representative firms and households. Each firm employs n_t workers and posts v_t vacancies to attract new workers. Firms incur a cost κ per vacancy posted and firing costs b_t^s per laid off worker from endogenous job separations. The total number of unemployed workers searching for a job is $u_t = 1 - n_t$. The number of new hires m_t is determined according to the matching function $m_t = \bar{m} u_t^{\gamma} v_t^{1-\gamma}$, with $\bar{m} > 0$ and $\gamma \in (0, 1)$. The probability that a firm fills a vacancy is given by $q_t = m_t/v_t = \bar{m}\theta_t^{-\gamma}$, where $\theta_t = v_t/u_t$ is the extent of labor market tightness. The probability that an unemployed worker finds a job is given by $p_t = m_t/u_t = \bar{m}\theta_t^{1-\gamma}$. Firms and workers take q_t and p_t as given. Finally, each firm separates from a fraction $\varrho(\tilde{a}_t)$ of existing workers each period. This quantity involves an exogenous component, $\bar{\varrho}$, and an endogenous one. Following Krause and Lubik (2007), job destruction probabilities a_t are drawn every period from a distribution with c.d.f $F(a_t)$ with positive support and density $f(a_t)$. $\tilde{a}_t$ is an endogenously determined threshold value and a job is destroyed if $a_t < \tilde{a}_t$. This gives rise to an endogenous job separation rate $F(\tilde{a}_t)$. The total separation rate is given by: $\varrho(\tilde{a}_t) = \bar{\varrho} + (1 - \bar{\varrho})F(\tilde{a}_t)$.

3.1 Intermediate Goods-Producing Firms

The (representative, intermediate goods producing) firm uses labor to produce output y_t according to $y_t = \bar{A}n_t A(\tilde{a}_t)$, where $\bar{A} > 0$ is a common productivity factor and $A(\tilde{a}_t) = E[a|a \geq \tilde{a}_t] = \frac{1}{1-F(\tilde{a}_t)} \int_{\tilde{a}_t}^{\infty} a dF(a)$ is the conditional expectation of productivity being larger than the endogenously determined critical threshold. To raise the workforce in turn, firms need to post vacancies. Hence, the firm can influence employment along two dimensions: the number of vacancies posted and the number of endogenously destroyed jobs. This gives rise to the following employment dynamics

$$n_t = (1 - \varrho(\tilde{a}_t))(n_{t-1} + m_{t-1}). \tag{3.1}$$

Current period profits are given by $\pi_t^F = y_t/\mu_t - w_t n_t - \kappa v_t - F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1}v_{t-1})b_t^s$ where the output price is normalized to unity, $w_t = \int_{\tilde{a}_t}^{\infty} \frac{w_t(a)}{1-F(\tilde{a}_t)} dF(a)$ is the (average) real wage weighted according to the idiosyncratic job productivity, μ_t is the price markup (its inverse equals real marginal costs mc_t), and the last term captures firing costs (Cacciatore et al., 2021). In detail, $(n_{t-1} + m_{t-1})(1 - \bar{\varrho})F(\tilde{a}_t)$ represents the number of existing (n_{t-1}) and new (m_{t-1}) workers who survived the exogenous job separation $(1 - \bar{\varrho})$, but got laid off due to the endogenous job separation $(F(\tilde{a}_t))$. b_t^s captures the cost per laid off worker. Firm expenses from firing are modeled as real resource costs[13]. The firm maximizes the present discounted value of expected profits: $\max_{n_t, v_t, \tilde{a}_t} \mathbb{E}_t \sum_{k \geq 0} \Lambda_{t,t+k} \pi_{t+k}^F$, subject to the production function and Eq. (3.1). $\mathbb{E}_t$ is the expectation conditional on the information up to and including time t; $\Lambda_{t,t+k}$ denotes the firm's stochastic

[13] We consider the case where firing costs accrue to the government in Appendix B.

discount factor, defined below. The first order conditions give rise to[14]

$$
\begin{aligned}
F_t^n &= mc_t mpl_t - w_t + \mathbb{E}_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], && (3.2)\\
\frac{\kappa}{q_t} &= \mathbb{E}_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], && (3.3)\\
mc_t A(\tilde{a}_t) &= \frac{1}{\bar{A}} \left(w_t - b_t^s - \frac{\kappa}{q_t} \right), && (3.4)
\end{aligned}
$$

where mpl_t is the marginal product of labor and F_t^n is the Lagrange multiplier associated with Eq. (3.1). In Eq. (3.2), F_t^n captures the (shadow) value accruing to the firm when employing one additional worker at time t and consists of four components: (i) the marginal product of a worker, (ii) the (marginal) cost of employing one additional worker, (iii) the continuation value of keeping the worker employed and (iv) the cost per laid off worker of the endogenous job separation. Eq. (3.3) is the free entry condition. It relates the value of employing an additional worker ($(1 - \varrho(\tilde{a}_{t+1}))F_{t+1}^n$) to the cost per vacancy (κ/q_t) and the cost per laid off worker ($b_{t+1}^s(1 - \bar{\varrho})F(\tilde{a}_{t+1})$). Finally, Eq. (3.4) sets the conditions for the idiosyncratic job productiveness ($\tilde{a}_t$) and hence for endogenous job destruction. Firms accept a lower idiosyncratic job productivity from workers when (i) firing costs (b_t^s) and/or (ii) search costs (κ/q_t) increase; however, (iii) higher wages induce firms to require higher productivity from workers.

3.2 Final Goods-Producing Firms

Final goods producers buy intermediate goods and sell them to the households. We assume that there is a continuum of final goods producers indexed by $i \in [0, 1]$. They are perfectly competitive in their input markets and monopolistically competitive in their output market. Their price setting is subject to nominal rigidities à la Calvo (1983). In line with Christiano, Eichenbaum and Evans (2005), firms that cannot re-optimize their price in a given period partially index their price to the previous period's CPI inflation rate π_t,

[14] The first order condition with respect to $\tilde{a}_t$ is given by: $n_t \left(\bar{A} \frac{\partial A(\tilde{a}_t)}{\partial \tilde{a}_t} - \frac{\partial w_t}{\partial \tilde{a}_t} \right) = (n_{t-1} + q_{t-1} v_{t-1}) \left(b_t^s (1 - \bar{\varrho}) f(\tilde{a}_t) + F_t^n \frac{\partial \varrho(\tilde{a}_t)}{\partial \tilde{a}_t} \right)$. Using Eq. (3.3), (3.2), and (3.1), this equation can be further simplified to: $(1 - \bar{\varrho})(1 - F(\tilde{a}_t) \left(\frac{\partial A(\tilde{a}_t)}{\partial \tilde{a}_t} - \frac{\partial w_t}{\partial \tilde{a}_t} \right) = b_t^s (1 - \bar{\varrho}) f(\tilde{a}_t) + \left(\mu_t mpl_t - w_t + \frac{\kappa}{q_t} \right) \frac{\partial \varrho(\tilde{a}_t)}{\partial \tilde{a}_t}$. Using the derivatives of $\frac{\partial A(\tilde{a}_t)}{\partial \tilde{a}_t}$, $\frac{\partial w_t}{\partial \tilde{a}_t}$ and $\frac{\partial \varrho(\tilde{a}_t)}{\partial \tilde{a}_t}$ yields the following expression: $\tilde{w}_t(a) = b_t^s + \frac{\kappa}{q_t} + \bar{A}a$. Finally, operating on both sides with $\int_{\tilde{a}_t}^{\infty} \frac{dF(a)}{1-F(\tilde{a}_t)}$ and using the definition of the production function gives Eq. (3.4).

implying $P_t(j) = (1+\pi_{t-1})^{\gamma_p} P_{t-1}(j)$, where $\gamma_p \in [0,1]$ captures the extent to which prices are indexed to the past inflation rate.[15]

The probability that a firm cannot re-optimize its price for k periods is given by ξ^k. Profit maximization by the intermediate goods producer j, which is allowed to re-optimize its price at time t, implies that it chooses a target price P_t^* that maximizes the following stream of future profits $\mathbb{E}_t \sum_{k\geq 0} \xi^k \Lambda_{t,t+k} \int_0^{y_{t+k}(j)} (P_t^* \Psi_{t+k} - mc_{t+k} P_{t+k}) dq$, subject to the demand constraint $y_{t+k}(j) = (P_t^*/P_{t+k})^{-\varepsilon} y_{t+k}$ and $\Psi_{t+k} = \left(\frac{P_{t-1+k}/P_{t-1}}{P_{t+k}/P_t}\right)^{\gamma_p}$. The first order condition with respect to the price P_t^* implies that the following condition has to hold $\mathbb{E}_t \sum_{k\geq 0} \xi^k \Lambda_{t,t+k} y_{t+k}(j) \left(P_t^* \Psi_{t+k} - \frac{\varepsilon}{\varepsilon-1} mc_{t+k} P_{t+k}\right) = 0$. This expression highlights that the price P_t^* set by firm j at time t, is a markup over expected future marginal costs. If prices can be adjusted at any point in time ($\xi = 0$), the markup is equal to $\frac{\varepsilon}{\varepsilon-1}$. With sticky prices, the markup varies over time. Finally, the definition of the price index P_t for domestic goods implies that its law of motion is given by

$$P_t = \left(\xi((1+\pi_{t-1})^{\gamma_p} P_{t-1})^{1-\varepsilon} + (1-\xi)(P_t^*)^{1-\varepsilon}\right)^{\frac{1}{1-\varepsilon}}, \tag{3.5}$$

which then results in the following expression for the New Keynesian Phillips curve

$$(1+\gamma_p \xi \beta)\pi_t = \beta \mathbb{E}_t \pi_{t+1} + \gamma_p \pi_{t-1} + \frac{(1-\xi\beta)(1-\xi)}{\xi} \hat{mc}_t. \tag{3.6}$$

3.3 Households

We model households following the approach proposed by Merz (1995). We consider an infinitely lived representative household consisting of a continuum of individuals of mass one. Household members pool income which accrues from labor income and unemployment benefit remuneration from employed and unemployed household members, respectively. Household members pool consumption to maximize the sum of utilities, that is, the overall household utility.

The budget constraint is given by

$$c_t + B_t = R_{t-1} B_{t-1} + (1-\tau) w_t n_t + b_t^u (1-n_t) + T_t^S + \pi_t^F, \tag{3.7}$$

where c_t is household consumption and B_t are period t holdings of government bonds, for which a rate of return R_t accrues. b_t^u and T_t^S denote unemployment benefits per unemployed household member and lump-sum subsidies. Finally, $(1-\tau)w_t$ is the after-tax wage, corresponding to the tax rate τ. In addition to

[15] If $\gamma_p = 1$, then the firms that cannot re-optimize their price in period t instead adjust their price to the lagged inflation rate; if $\gamma_p = 0$, the firms that cannot re-optimize their price in period t leave their price unchanged. Uribe (2020) describes indexation as rule-of-thumb for adjusting prices.

the budget constraint, the household takes into account the flow of employment by its members in line with

$$n_t = (1 - \varrho(\tilde{a}_t))n_{t-1} + p_t(1 - n_{t-1}). \tag{3.8}$$

In a given period, the household derives utility from consumption c_t and dis-utility from working n_t. The instant utility function is $u(c_t, n_t)$. The household discounts instant utility with a discount factor β and maximizes the expected lifetime utility function: $\max_{c_t,n_t} \mathbb{E}_t \sum_{k\geq 0} \beta^k u(c_{t+k}, n_{t+k})$, subject to the budget constraint, Eq. (3.7) and the employment flow Eq. (3.8). Optimization leads to the following conditions

$$\begin{aligned} 1 &= R_t \mathbb{E}_t \Lambda_{t,t+1}, & (3.9)\\ H_t^n &= \tilde{w}_t^b - mrs_t + \mathbb{E}_t \left[1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}\right] \Lambda_{t,t+1} H_{t+1}^n, & (3.10) \end{aligned}$$

where λ_t is the Lagrange multiplier attached to Eq. (3.7) and $\lambda_t H_t^n$ the one attached to equation Eq. (3.8). Furthermore, $\tilde{w}_t^b = (1 - \tau)w_t - b_t^u$, $mrs_t = -u_{n,t}/\lambda_t$ and $u_{n,t} < 0$ is the marginal dis-utility of working. Note that λ_t is equal to the marginal utility of consumption in this case but also the marginal utility of wealth because it is the (Lagrange) multiplier on the household's budget constraint. Hence, mrs_t captures both the marginal rate of substitution between consumption and work and the marginal value of non-work activities. Assuming efficient financial markets implies that the stochastic discount factor, given by $\Lambda_{t,t+k} = \beta^k \frac{\lambda_{t+k}}{\lambda_t}$, applies to both households and firms.

Considering Eq. (3.10), H_t^n captures the household's (shadow) value of having one additional employed member. It consists of three components: (i) the increase in utility owing to the higher income when having an additional member employed, (ii) the decrease in utility from lower leisure captured by the marginal dis-utility of work, and (iii) the continuation utility value, given by the contribution of a current match a household's employment in the next period.

3.4 Nash Wage Bargaining

Wages are set each period based on Nash-bargaining of the pre-tax (average) wage w_t between firms and workers. The Nash wage satisfies: $w_t = \arg\max_{w_t} (H_t^n)^\eta (F_t^n)^{1-\eta}$ where $0 < \eta \leq 1$ captures workers' bargaining power. Optimization yields: $\eta F_t^n = (1 - \eta)H_t^n/(1 - \tau)$, which can be rearranged to

$$w_t = (1 - \eta)\frac{mrs_t + b_t^u}{1 - \tau} + \eta \left(mc_t mpl_t + \mathbb{E}_t \Lambda_{t,t+1} \left[\kappa\theta_{t+1} - b_{t+1}^s (1 - \bar{\varrho})F(\tilde{a}_{t+1})\right]\right). \tag{3.11}$$

The wage per worker is a weighted average of the unemployment benefit and the marginal rate of substitution on the one hand; and the marginal product of labor, the expected search cost and the firing costs (per worker) on the other. Higher unemployment benefits (b_t^u) and labor tax rates (τ) render non-work activities more attractive, inducing a rise in the equilibrium wage rate from the side of households. Conversely, a higher current marginal product of labor, higher expected search costs, and lower expected firing costs cause upward pressure on the equilibrium wage from the side of firms.

3.5 Policies, Aggregate Resource Constraint, and Government Budget Constraint

The government budget constraint satisfies

$$\tau w_t n_t + B_t = R_{t-1}B_{t-1} + b_t^u u_t + T_t^s + g_t, \tag{3.12}$$

where g_t is government consumption. Fiscal policy is governed by (i) an exogenous AR(1) process g_t (in log-deviations), (ii) a specification for unemployment benefits according to $b_t^u = \bar{\varphi} + \varphi w_{t-1}$ where φ is the replacement rate of a worker with respect to his last wage received, (iii) a specification for firing costs according to $b_t^s = \bar{\varsigma} + \varsigma w_{t-1}$, and (iv) government subsidies: $T_t^S = \bar{T}^S + \varphi_{T^s} B_t$. The parameters $\bar{T}^S$ and $\bar{\varsigma}$ serve the purpose to simplify the steady state computations and $\varphi_{T^s} B_t$ ensure that the necessary stability conditions are satisfied.

Monetary policy is governed by a Taylor rule according to $i_t = \rho_i i_{t-1} + (1-\rho_i)(\phi_\pi \pi_t + \phi_y \hat{y}_t)$ where $i_t = \ln(R_t)$ and the hat-notation refers to the log-deviation from the steady state.

Finally, using Eq. (3.12), Eq. (3.7), and the expression for firms' profits (π_t^F), we obtain the aggregate resource constraint

$$y_t = c_t + g_t + \kappa v_t + F(\tilde{a}_t)(1-\bar{\varrho})(n_{t-1} + q_{t-1}v_{t-1})b_t^s. \tag{3.13}$$

This equation closes the model.

3.6 Embedding LMIs in the Model

The indicators for the LMIs are embedded in the model by the three structural parameters ς, η, and φ. The mapping between the empirical LMIs and their theoretical counterparts in the DSGE model can only be qualitative.

We proxy the government's ability to shape the extent of employment protection, ς, with the EPL index. While employment protection can take a variety of forms, such as strict layoff rules for individual occupational groups, short-time work models that allow companies to forego layoffs due to (temporary) subsidies, and also the existence of payments that may arise with a dismissal. In addition to severance payments, the latter also includes, as is customary in many countries, one-time payments to the social security system due to the burden on the unemployment insurance caused by the dismissal.[16] The parameter η captures the bargaining power of workers and is proxied by UD. It is thus a measure of the implicit advantage that employees benefit from within the wage-setting process.[17] In a more general interpretation, this can also be viewed as a measure of union strength or as a measure of the degree of centralization of wage bargaining since a higher degree of centralization of wage bargaining is typically considered beneficial for workers in the wage bargaining process (Abbritti and Weber, 2018). Finally, the parameter φ captures the amount of unemployment benefit payments in relation to the wage received before dismissal. This value is usually set directly by governments and is comparatively less ambiguous than the other two LMI parameters (η and ς).

3.7 Equilibrium, Model Solution, and Dynamic Simulations

We collect the LMI parameters of interest in the vector $\boldsymbol{\vartheta} = (\eta, \varphi, \varsigma)'$ and assess the implications of changes in these for fiscal policy by assessing their effects on the impulse response functions to a shock in government spending (g_t). To this purpose, we consider a log-linearized solution of the rational expectations model around its steady state,

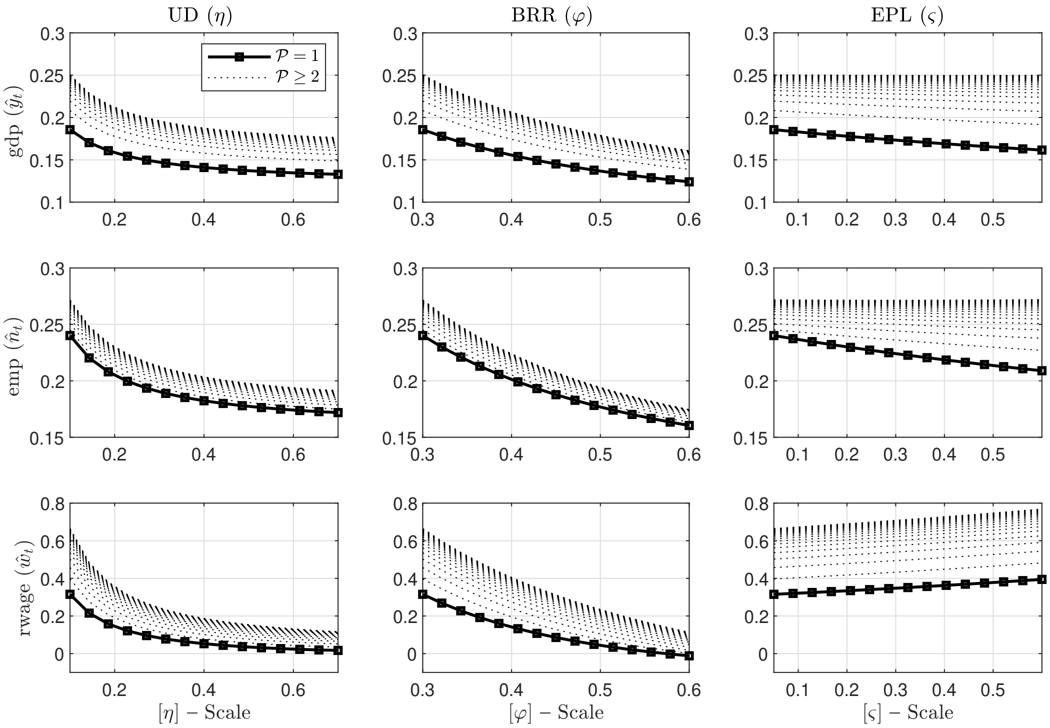
$$\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})z_t = \boldsymbol{\Psi}_1(\boldsymbol{\vartheta})z_{t-1} + \boldsymbol{\varepsilon}_t, \tag{3.14}$$

where the vector z_t contains the endogenous variables and the vector of exogenous shocks simplifies to $\boldsymbol{\varepsilon}_t = \hat{g}_t$ in our case. The matrix $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ governs the dynamics among the dependent variables and the matrix $\boldsymbol{\Psi}_0^{-1}(\boldsymbol{\vartheta})$ determines the contemporaneous impact of the fiscal spending shock on the endogenous variables. Eq. (3.14) explicitly depicts the dependency of the coefficient matrices on the three LMI parameters. We assess the consequences of each of the three parameters individually by computing impulse response functions (IRFs) based on a calibration of the model's parameters as outlined in Appendix A.2. As the IRFs are continuous

[16] However, the index remains qualitative. Quantitative data on firing costs are not readily available at the country level. Moreover, as they cover only severance payments and the length of the notice period, they omit non-monetizable elements of employment protection, as for instance administrative and judicial procedures.

[17] An alternative indicator for wage bargaining power is coverage of collective bargaining. This measure is far more persistent than union density and thus does not adequately reflect the underlying bargaining power of wage earners.

Figure 3: Fiscal Spending Multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$).



Notes: The sub-figures illustrate 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 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.

functions of $\boldsymbol{\vartheta}$, we can display them over a whole range of values of $\boldsymbol{\vartheta}$. We do so by considering the following definition of the fiscal spending multiplier for some variable x

$$\mu_x(\vartheta_l) = \frac{\sum_{i=1}^{\mathcal{P}} \mathrm{IRF}_i^x(\vartheta_l)}{\sum_{j=1}^{\mathcal{P}} \mathrm{IRF}_j^{\hat{g}}(\vartheta_l)}, \qquad \forall\, l = \{1, 2, 3\}, \tag{3.15}$$

where $\mathrm{IRF}_t^x(\vartheta_l)$ and $\mathrm{IRF}_t^{\hat{g}}(\vartheta_l)$ denote the impulse response functions of some variable x and government spending $\hat{g}$ to the fiscal spending shock over the horizon $\mathcal{P}$. The definition of $\mu_x(\vartheta_l)$ considers the response of a variable relative to the size and persistence of the shock. In what follows, we will refer to $\mu_x(\vartheta)$ as the multiplier for a specific variable x and focus on distinct horizons $\mathcal{P}$. The results are shown in Figure 3 for output ($\hat{y}_t$), employment ($\hat{n}_t$), and the real wage ($\hat{w}_t$). The multipliers for each variable are displayed distinctively for the first horizon ($\mathcal{P} = 1$) and higher horizons ($\mathcal{P} \geq 2$); this has the advantage to depict also the dynamics. The columns consider the dependency of the multipliers on the respective LMI parameters ($\boldsymbol{\vartheta}$).

An intuitive understanding of the working of the model can be gained by considering the negative wealth effect caused by higher government spending. Consumption and leisure are both normal goods, hence they both fall as a result of the negative wealth effect from higher expected taxation. The drop in consumption raises the marginal utility of consumption, which gives rise to a drop in the marginal rate of substitution between consumption and labor ($mrs_t = -u_{n,t}/u_{c,t}$) or, in other words, a decrease in the current value of non-work activities. As a consequence of the drop in leisure, the associated increase in employment raises output and leads to a positive fiscal spending multiplier. The effect on the equilibrium wage is in principle ambiguous: the drop in the marginal product of labor and the marginal rate of substitution (or equivalently, the value of non-work activities) contrasts with a rise in the expected search cost. The comparably larger reaction of the former two triggers a drop in the equilibrium wage rate. In a similar vein, the response of the labor market tightness variable (θ_t) is ambiguous despite the decrease in unemployment. The drop in the equilibrium wage raises the value to the firm of an additional worker ($\partial F_t^n/\partial w_t < 0$) which creates incentives for firms to increase vacancy postings and hiring activities. This contrasts with the rise in expected search costs. The overall effect on vacancies v_t and labor market tightness θ_t is thus ambiguous.

In what follows, we focus on the role of the relative bargaining power of workers (UD, η), the extent of employment protection (EPL, ς) and the unemployment benefit replacement rate (BRR, φ) in shaping the responses of interest.

3.8 Implications of the LMIs

We start by considering the BRR (φ) and its role as a determinant of the shape of the employment and output response to fiscal shocks, as depicted in Figure 3. While employment increases in response to the fiscal spending rise, the positive response is larger when the unemployment remuneration is low. This can be explained by considering the reservation wages for households and firms ($\underline{w}_t^H$ and $\overline{w}_t^F$).

The reservation wage of a household (firm) is given by the minimum (maximum) wage acceptable. Since H_t^n (F_t^n) describes the marginal value to the household (firm) of having one further worker employed, the reservation wages of a household and a firm are hence determined by $H_t^n = 0$ and $F_t^n = 0$. In this situation, the household and the firm are not willing to increase or decrease labor supply and demand. Using Eq. (3.2) and (3.10), and setting τ equal to zero for simplicity, the reservation wages are given by

$$\overline{w}_t^F = mc_t mpl_t + \mathbb{E}_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], \tag{3.16}$$

$$\underline{w}_t^H = mrs_t + b_t^u - (1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}) \mathbb{E}_t \Lambda_{t,t+1} H_{t+1}^n, \tag{3.17}$$

from which $w_t = (1-\eta)\overline{w}_t^F + \eta \underline{w}_t^H$ follows. Hence, higher unemployment benefits ($\partial b_t^u / \partial \varphi > 0$) raise the reservation wage for households ($\partial \underline{w}_t^H / \partial \varphi > 0$), which contracts their value of employment ($\partial H_t^n / \partial \varphi < 0$) giving rise to a negative impulse on household's labor supply. Intuitively, an increase in unemployment benefits raises workers' outside option (i.e., the present value of being unemployed rises) which improves their bargaining position for the wage negotiations. This in turn leads to an increase in the reservation wage of households and, given a non-zero weight of households/workers in the wage bargaining process ($\eta > 0$), to an increase in the equilibrium wage (w_t). The higher wage, however, causes the value to the firm of having an additional worker employed ($\partial F_t^n / \partial w_t < 0$) to decrease. Hence, higher unemployment benefit remuneration attenuates the expansion in employment in response to the expansionary fiscal spending shock.[18] This result conforms with Atkinson and Micklewright (1991), Gruber (1997), Fuller (2014) and Faig, Zhang and Zhang (2016) who argue that aggregate output might be affected negatively by unemployment benefit remuneration payments since resources are shifted away from their efficient use. The negative impact on employment and output is countered by a stabilizing effect on consumer spending, which is particularly pronounced when non-Ricardian consumer behavior is taken into account (see Appendix B for more details).

We now turn to the effects of workers' bargaining power in the wage negotiation process, denoted by η. As discussed earlier, the equilibrium wage w_t is a weighted average of the two reservation wages, with the weights determined by η. From Eq. (3.16) and (3.17), we have

$$\frac{\partial w_t}{\partial \eta} = \overline{w}_t^F - \underline{w}_t^H, \quad (\text{for } \tau = 0)$$

where $\overline{w}_t^F > \underline{w}_t^H$, as otherwise no worker-firm match would be feasible. Consequently, $\frac{\partial w_t}{\partial \eta} > 0$, the implication of which is that as long as $\underline{w}_t^H < \overline{w}_t^F$, increases in workers' bargaining power push their reservation wages closer to those of firms, i.e., $\underline{w}_t^H \to \overline{w}_t^F$. This generates upward pressure on equilibrium wages, which, in turn, has two effects. Firstly, it reduces the value to the firm of employing an additional worker, and secondly, it leads firms to require a higher idiosyncratic productivity from workers. The latter

[18] Albertini and Poirier (2015) stress the role of the zero lower bound in this context. While increases in unemployment benefits always raise unemployment in normal times, the opposite may occur at the zero lower bound as the inflationary pressure triggered by higher unemployment benefits reduces the real interest rate, which in turn promotes consumption, output and employment.

effect raises the endogenous job separation rate ($\tilde{a}_t$ increases), resulting in a contraction of employment. Therefore, in response to an expansionary demand shock, such as a fiscal spending shock, the increase in both employment and output is smaller when workers' bargaining power is high. This is illustrated in the sub-panels in the first column of Figure 3, where the positive output effects diminish as workers' bargaining power increases. These findings are consistent with the results in Zanetti (2009), who shows that an increase in workers' bargaining power leads to a contraction in macroeconomic activity.

Finally, we turn to the effect of the extent of employment protection, denoted by ς. Unlike the previous variables—workers' bargaining power in wage negotiations (η) and unemployment benefits (φ)—employment protection does not influence the equilibrium by affecting reservation wages or the wage bargaining process. Instead, it directly impacts equilibrium outcomes by imposing constraints on firms' ability to adjust employment levels optimally. This distinction is important because employment protection affects firms' decision-making processes directly, rather than indirectly through changes in reservation wages, thus playing a unique role in shaping labor market dynamics. The primary mechanism through which employment protection influences fiscal spending multipliers is related to its effect on the sensitivity of job destruction and job creation, as described by Eq. (3.2) and (3.4). Starting with the latter, low firing costs increase the value of an additional worker to the firm ($\partial F_t^n / \partial \varsigma < 0$). In other words, lower firing costs promote job creation. Therefore, in response to an expansionary fiscal spending shock, low firing costs lead to relatively higher job creation, resulting in a stronger increase in employment relative to output. Regarding job destruction, Eq. (3.4) suggests that low firing costs increase the job destruction rate ($\partial \tilde{a}_t / \partial \varsigma > 0$). While this effect works in the opposite direction to job creation, both mechanisms render employment more sensitive to aggregate shocks. Consequently, in response to an expansionary fiscal spending shock, employment exhibits a larger response when firing costs are low. This discussion also highlights the possible heterogeneities arising from the LMIs in altering the transmission of shocks.

We reassess these results by considering a more general calibration in Appendix A.3 and examine the robustness to several extensions – (i) no inflation indexed prices, (ii) real wage rigidities, (iii) limited asset market participation, (iv) firing costs as government revenues, (v) productivity-enhancing government spending, and (vi) complementarity between consumption and leisure – in Appendix B. In short, we find that inflation indexation of prices crucially shapes the qualitative response of the real wage to fiscal spending shocks. Moreover, with the exception of extension (iv), the qualitative impact of the LMIs on the output and employment multipliers remains unchanged.

4. The Empirical Evidence

We empirically validate the results of our theoretical model using an interacted panel vector-autoregressive (IPVAR) specification as popularized by Towbin and Weber (2013) and Sá, Towbin and Wieladek (2014). We examine the conditional response to fiscal spending shocks and the effect on macroeconomic volatility for different levels of LMI indicators in a panel of developed countries. The IPVAR model is employed to assess how the characteristics of the matrices $\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})$ and $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ of the system given by Eq. (3.14) depend on the LMIs in place. We consider a first-order Taylor expansion of these matrix functions around the sample average of $\boldsymbol{\vartheta}$, given by $\bar{\boldsymbol{\vartheta}}$

$$\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}) \approx \boldsymbol{\Psi}_j(\bar{\boldsymbol{\vartheta}}) + \sum_{l=1}^{3} \left[\frac{\partial \boldsymbol{\Psi}_j}{\partial \vartheta_l}(\bar{\boldsymbol{\vartheta}})(\vartheta_l - \bar{\vartheta}_l) \right], \quad j \in \{0, 1\}. \tag{4.1}$$

Substituting the matrices $\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})$ and $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ in Eq. (3.15) by the Taylor approximation given by Eq. (4.1) gives rise to an additive separable expression for the parameters ϑ_l, $l = \{1, 2, 3\}$, multiplied in each case by the endogenous variables. From an econometric point of view, this implies that interaction terms appear in the specification after this substitution is carried out. In the following, we describe the econometric model used to estimate $\boldsymbol{\Psi}_j(\boldsymbol{\vartheta})$, before presenting the results and providing a discussion of the insights gained from the empirical evidence.

4.1 The Econometric Model

We estimate the following reduced-form IPVAR for the n-dimensional vector of endogenous time series, $\boldsymbol{y}_{it}$, conditional on the d-dimensional vector of interaction variables, $\boldsymbol{\vartheta}_{it}$, for country $i = 1, \dots, N$

$$\mathbf{y}_{it} = \boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) + \sum_{j=1}^{p} \boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it})\mathbf{y}_{it-j} + \mathbf{u}_{it}, \quad \mathbf{u}_{it} \sim \mathcal{N}_n\big(\mathbf{0}, \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it})\big), \tag{4.2}$$

where all coefficients are country-specific and dependent on the interaction term. $\boldsymbol{c}_i$ denotes the intercept, $\boldsymbol{\Phi}_{ij}$ the coefficient matrix for lag $j = 1, \dots, p$, and $\boldsymbol{\Sigma}_i$ the covariance matrix. All these reduced-form coefficients are a linear function of the interaction term and the parameters change depending on the exact value taken by the interaction variable. The details of the model framework are presented in Appendix C.

The structural IPVAR representation is then given by

$$\tilde{\boldsymbol{\Psi}}_{i0}\boldsymbol{y}_{it} = \sum_{j=1}^{p} \tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it})\boldsymbol{y}_{it-j} + \boldsymbol{e}_{it}, \quad \boldsymbol{e}_{it} \sim \mathcal{N}_n(\mathbf{0}, \boldsymbol{I}), \tag{4.3}$$

where we have excluded the deterministic term for the sake of simplicity. The underlying idea of the panel setup is to estimate a common economic model for all countries in our sample. This is done via a pooling prior in the spirit of Jarociński (2010). The prior assumes that the structural country-specific coefficients have a common underlying Gaussian distribution,

$$\tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it}) \sim \mathcal{N}(\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t), \boldsymbol{V}_j), \quad j = 1, \ldots, p, \tag{4.4}$$

where $\boldsymbol{V}_j$ denotes the covariance matrix. We exert regularization via this covariance matrix towards the common-mean model through Bayesian global-local shrinkage priors (Griffin and Brown, 2010; Huber and Feldkircher, 2019). The pooling prior and the exact specification of the shrinkage prior are described in Appendix C.

The correspondence between the observable LMIs $\boldsymbol{\vartheta}_{it}$ and the structural LMI parameters $\boldsymbol{\vartheta}$ of the DSGE model can be made explicit by defining $\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t) = \boldsymbol{\Psi}_{jt}$ and $\boldsymbol{\Psi}_j(\bar{\boldsymbol{\vartheta}}_t) = \bar{\boldsymbol{\Psi}}_j + \partial \boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t)/\partial \boldsymbol{\vartheta}_t \cdot (\boldsymbol{\vartheta} - \bar{\boldsymbol{\vartheta}})$ for $j = 0, 1, \ldots, p$. This implies that the coefficients of the $\boldsymbol{\Psi}_{jt}$ matrix vary as follows

$$\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t) = \boldsymbol{\Psi}_{jt} = \bar{\boldsymbol{\Psi}}_j + \sum_{l=1}^{d} \frac{\partial \boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t)}{\partial \vartheta_{lt}} (\vartheta_{lt} - \bar{\vartheta}_l), \quad j = 0, ..., p. \tag{4.5}$$

This relates the empirical set-up directly to Eq. (4.1) of the theoretical model. The full IPVAR model is given by Eq. (4.3), and its equivalence with the solution of the DSGE model depicted in Eq. (3.14) and Eq. (4.1) is evident when considering a lag length of one.

In the IPVAR specification, interactions between the endogenous variables and labor market indicators are thus included in the specification and thus LMIs act as mediators of the effect of fiscal policy (and other) shocks. As a result, impulse response functions can be evaluated for varying values of ϑ_l. For the ease of interpretation, we examine changes in the structural coefficients only by varying one interaction variable, while keeping the remaining ones at a given level (i.e., at the median).

There are two potential limitations to the empirical approach adopted here. First of all, LMIs may be endogenous to shocks hitting the economy. Given the path of the LMI variables depicted in Figure 1, structural rather than cyclical factors appear to determine their dynamics.[19] A second potential limitation is the linearity assumption (in the parameters) embedded in the IPVAR model, which mimics the approximation considered in Eq. (4.1). In principle, the assumption of linearity could be relaxed by considering various non-linear extensions of $\boldsymbol{\vartheta}_t$. However, depending on the number of observations and parameters of interest in

[19] This is confirmed within a robustness check where we include each one of the LMI indicators in a standard panel VAR and calculate the impulse response functions of the LMI variables. They do not significantly react to cyclical shocks.

the estimation, overfitting of the model becomes a problem in our setting, so we stick to linear specifications with interactions in this piece instead of assessing more complex nonlinear parametrizations of the model.

4.2 Data and Specification

We use quarterly data ranging from 1960Q1 to 2020Q4 for 16 OECD countries to estimate the IPVAR model.[20] The vector of endogenous variables includes $\boldsymbol{y}_{it} = (g_{it}, x_{it}, er_{it}, \omega_{it})'$ which denotes the growth rate of real government consumption per capita, g_{it}, the growth rate of real GDP per capita, x_{it}, first difference of the employment rate, er_{it}, and the growth rate of real wages, ω_{it}.[21] We use year-on-year growth rates and first differences and refer to Table D1 and Table D2 in the Appendix for further details.

As regards the interaction variables, we specify $\boldsymbol{\vartheta}_{it} = (\eta_{it}, \varphi_{it}, \varsigma_{it})'$ and use data from the CEP-OECD institutions database for union density (η_{it}), unemployment benefit replacement rates (φ_{it}) and employment protection legislation (ς_{it}).[22] The original data set contains annual observations, which we interpolate to a quarterly frequency by assigning the annual value of a particular year to each quarter of the same year. We estimate the IPVAR model using all three interaction variables at once. Prior to estimation, we standardize each interaction variable. This serves the purpose of comparability across countries as we exploit *within*-country variation in the LMIs. We interpret our estimates thus as conservative due to the strong cross-country heterogeneity in the LMIs. When simulating the IPVAR model along a particular interaction variable ϑ_l, we set the remaining ones ϑ_ℓ ($\ell \neq l$) equal to its median.

All models are estimated with one lag ($p = 1$) as proposed by the Bayesian information criterion.[23] The estimation is based on 20,000 posterior draws, where we discard the first 10,000 as burn-ins.

4.3 Shock Identification

We identify fiscal spending shocks by imposing a recursive identification 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

[20] The sample includes information on the following OECD countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Italy, Japan, the Netherlands, Portugal, Spain, Sweden, and the United States.

[21] Following Brückner and Pappa (2012), we express all variables in per-capita terms. Additionally, we provide robustness by applying the procedure suggested by Gordon and Krenn (2010). We divide each variable with its Hamilton (2018)-filtered trend. As noted by Ramey (2016), the impulse responses from using this transformation are quite similar to those using log-levels and it produces relatively narrow confidence bands. We report results in Appendix E.2.

[22] For a descriptive description and further details of the LMIs, we refer the reader to Section 2 and Table D1 in the Appendix.

[23] We provide robustness to this choice in Appendix E.2. Our results are robust to using two or four lags.

system. These three variables are hence assumed to respond within the same quarter to the fiscal spending shock. This 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 recursive identification approach for fiscal spending shocks 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 (Beetsma and Giuliodori, 2011; Bénétrix and Lane, 2013; Ilzetzki, Mendoza and Végh, 2013; Huidrom et al., 2020, to mention a few). Second, alternative identification approaches are infeasible in the context of our research question.[24]

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 of the econometrician can generate equilibria with 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.[25]

4.4 The Effect on Fiscal Spending Effectiveness

In this section, we present the effects of the LMIs on fiscal spending effectiveness. In line with the theoretical results, we compute multipliers for the initial impact ("impact multiplier", horizon $P = 0$) and for the effect after four quarters ("one-year multiplier", horizon $P = 4$). Figure 4 depicts the multipliers for output, employment, and the real wage. In each panel, we display the sensitivity of the multipliers with respect to the LMIs. The black solid and the red dash-dotted lines refer to the median value of the impact and one-year

[24] 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). All these approaches are not feasible because they rely on additional data (e.g., detailed institutional information or data on fiscal spending plans) or are not practical for a large panel of countries. An interesting alternative is Miyamoto, Nguyen and Sheremirov (2019) who use military spending data in a large panel of countries but only on a yearly frequency.

[25] 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 by professional forecasters of Oxford Economics and the OECD. We restrict our sample to the G7 countries in the cross-section and starting only in the mid 1980s (OECD) or late 1990s (Oxford Economics). We refer to the Appendix E.3 for further details.

multipliers and are complemented with the 80% confidence bounds in each case. The horizontal axis ranges from the 10th to 90th quantile of the distribution of the respective LMIs, while the vertical axis depicts the value of the fiscal multiplier for the respective variable.

We start by discussing the impact multiplier for output and its dependency on LMIs. Across all LMIs, we find impact multipliers between 0.2-0.3 for low levels of the LMIs. At a low value of UD (η), we find that a one percent increase in fiscal spending raises real GDP by 0.3 percent; the value of the output multiplier drops to zero when UD is at the upper end of the distribution. This gives rise to a strong decline in the output multiplier, which is substantial and statistically significantly different from zero. In the case of the EPL (ς) we do not find a strong decline in the output multiplier, while we even find a statistically significantly larger output multiplier for higher levels of the BRR (φ).[26] This finding is interesting and deserves a more thorough discussion which we provide further below.

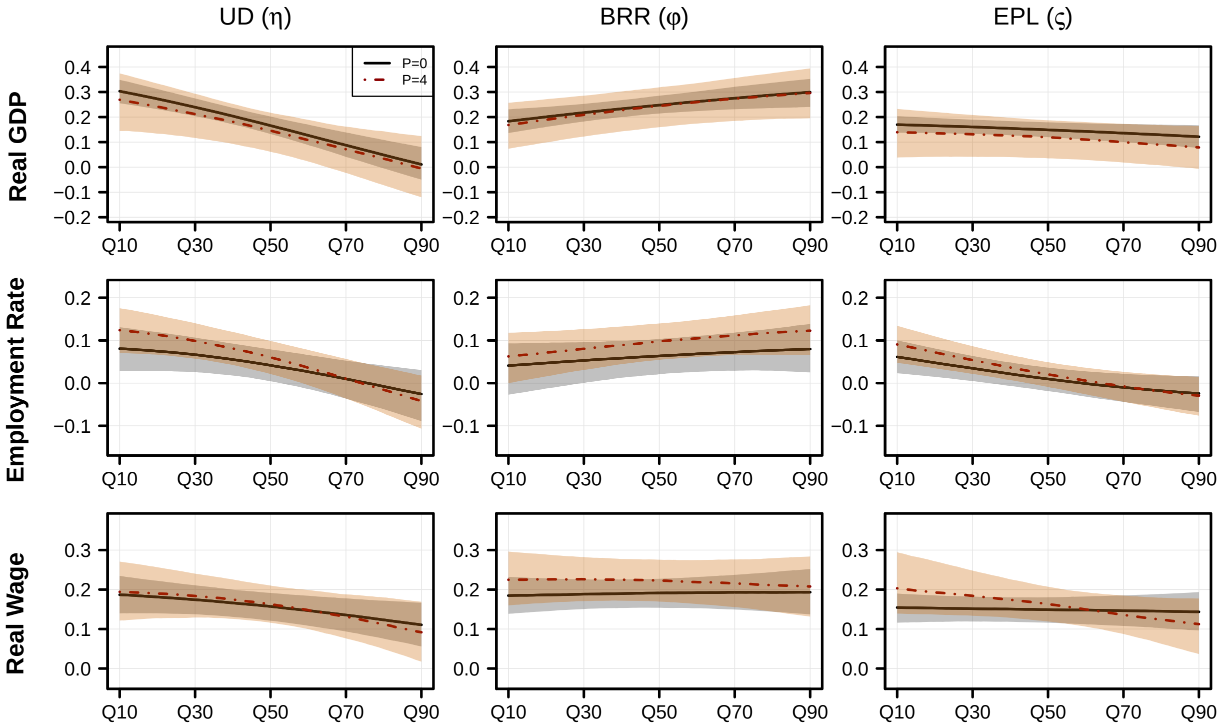
The impact multipliers for employment, while consistently positive, are negatively affected by two out of three LMIs. The drop is sizeable in the case of UD and the EPL and amounts to around 10 log points (from 0.10 down to 0.00), while being again moderately increasing for the BRR. The one-year multipliers consistently exceed the impact multipliers, which highlights the inertia of the impact of government spending shocks on economic activity. For UD and EPL, the decline of the employment multiplier is statistically significant, while the increase for the BRR is estimated relatively imprecisely.[27] For the real wage, we find the least pronounced effects. The multiplier declines for UD both for the on impact and one-year multiplier, while it only declines for the EPL for the one-year multiplier. However, these estimates face high uncertainty. In the case of the BRR, we find constant multipliers.

Overall, the LMIs are found to play a role for the effectiveness of discretionary fiscal policy. This applies to both the goods and the labor market. Among the three LMIs considered, the UD and EPL are found to have the strongest effects. The output and employment multipliers of the IPVAR model are comparable to those of the calibrated DSGE model. Similar to Figure 3, a strong decrease is visible for UD and EPL. The comparably mild decline of the output multiplier with respect to the EPL highlights the limited role severance payments and alike, which characterize the extent of employment protection (see Section B in the

[26] These differences are statistically significant, as we show in Appendix E.1. We examine the dynamic responses and show IRFs evaluated at the 10th and 90th quantile of the respective LMI distribution. Furthermore, we investigate the full posterior distribution of the differences in these IRFs. The evidence suggests that differences are strongly significant at the on impact and one-year horizon for UD (output and employment) and EPL (employment). For the BRR, the differences are not statistically significant. Furthermore, we also find that most of the dynamic response happens within the first six months.

[27] In fact, in subsection E.1 we show that differences are not statistically significant at the 68% level.

Figure 4: Fiscal Multipliers.



Notes: The sub-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 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.

Appendix), while at the same time give rise to a re-distribution to households which in turn attenuates the negative impact of a more stringent EPL on the output multiplier. Nevertheless, some differences arise. First and most noteworthy, we find an increase in the multiplier when imposing higher values of the BRR. This is clearly visible for the output multiplier. This could reflect that our model does not include unemployment risk, giving rise to a precautionary savings channel. Unemployment insurance can mitigate this channel and cause an additional demand channel (Challe, 2020; McKay and Reis, 2021; Kekre, 2023).[28] Furthermore, the null effect on the employment rate for BRR aligns well with micro-evidence from Boone et al. (2021) examining unemployment insurance extensions. Second, our baseline model shows no strong change in the multipliers with respect to real wages. Third, in the theoretical model, the one-year employment multiplier

[28] In fact, this literature differentiate between changing the level of unemployment benefits and extending unemployment benefits, where our results only speak to the former. While Challe (2020) investigates the interaction with monetary policy, McKay and Reis (2021) examine the optimal unemployment insurance level as automatic stabilizer, and Kekre (2023) looks into duration extensions. All, however, point to an additional demand channel arising through less precautionary savings when unemployment risk is *somehow* insured.

consistently exceeds the impact multiplier, which highlights the role of limited asset market participation in this context.[29] The empirical findings point, however, to rather little inertia in the dynamic responses.

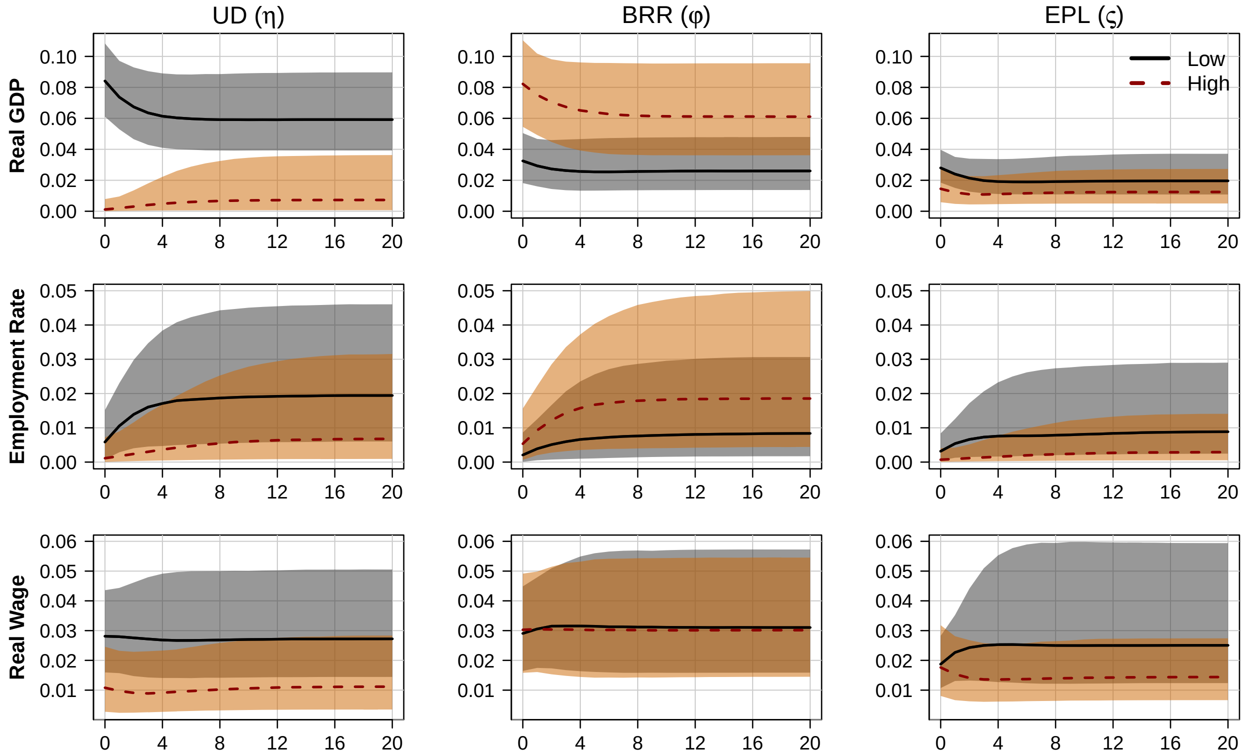
We provide the forecast error variance decomposition (FEVD) in Figure 5. The share of the variation in output explained by fiscal spending shocks depends both on the horizon and the LMIs. As can be seen, fiscal spending shocks explain a low fraction of the variance of output when the horizon considered is short and a stringent value of UD is deployed ("high"). In contrast, it explains up to 8% at horizons of eight quarters and beyond when UD are, however, less stringent ("low"). This share is substantially lower for the variables characterizing the labor market, employment, and real wages. The forecast error variance of output is reduced the most clearly, while less clear reductions are visible for the employment rate and the real wage. We find similar effects, albeit to a less pronounced degree, for the EPL. Lastly, mirroring the findings of the IRF analysis, for the BRR (φ) the picture changes. We find that a more stringent value of the BRR ("high") leads to a larger explained share of the forecast error variance for output. We do not find significant differences for the other two variables. Overall, we find that when stringent LMIs are deployed, discretionary fiscal policy only has a limited potential in affecting labor market outcomes.

These findings are in line with the literature as regards the size of fiscal spending multipliers for output (Ramey, 2019), the extent of inertia (Ilzetzki, Mendoza and Végh, 2013), as well as the lower value of the employment relative to the output multiplier (Monacelli, Perotti and Trigari, 2010). The positive real wage multiplier is consistent with the findings in Brückner and Pappa (2012) who identify both negative and positive multipliers across distinct countries. Note that our results refer to average effects across positive and negative fiscal spending shocks, although asymmetries can arise (Barnichon, Debortoli and Matthes, 2022). Our key contribution in this context concerns the assessment of the size and shape of fiscal multipliers with respect to the LMIs. In this regard, we find strong evidence in favor of a dependency of fiscal multipliers and hence of the effectiveness of discretionary fiscal policy on the LMIs.

Further results on the extent to which the LMIs shape the transmission channel of fiscal spending shocks are provided in Appendix E. We report the dynamic responses as well as robustness checks to the baseline model. In Figure E3 we control for fiscal foresight on a smaller sample which is not found to exert a strong influence on the baseline findings. We also re-do the analysis with other labor market variables, the

[29] We provide and discuss a number of extensions to the theoretical model in Appendix B. Limited asset market participation generally yield a higher level of fiscal multipliers due to the presence of non-Ricardians.

Figure 5: 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 months). 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.

unemployment rate and labor market tightness. Finally, we also investigate the issue of strong cross-country heterogeneity in the LMIs as depicted in Figure 2.

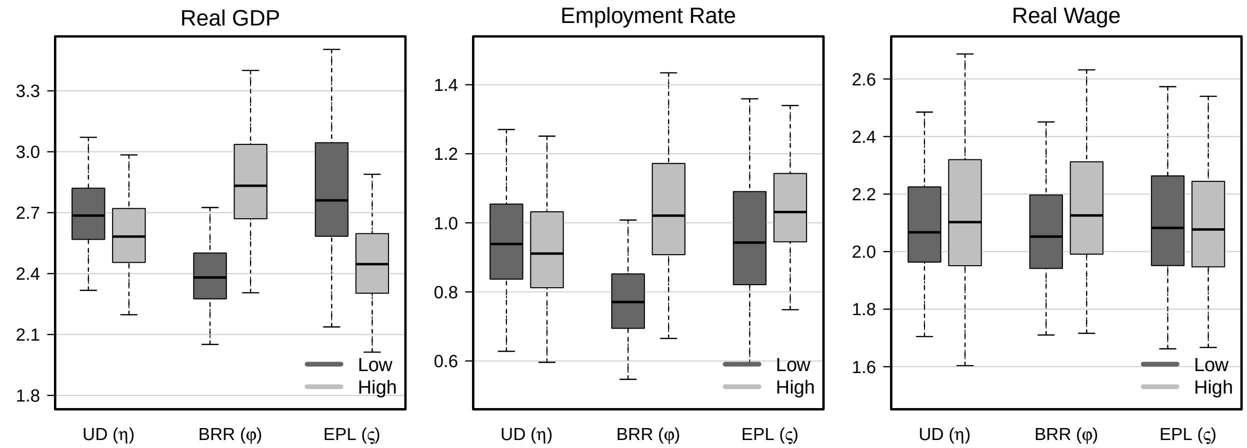
4.5 The Effect on Macroeconomic Volatility

While the effects of discretionary fiscal policy potentially are decreased due to more stringent LMIs, the question arises whether there is a strong need for these policies in an environment with stronger labor market frictions.[30] We are thus interested in the degree of substitutability between the LMIs and fiscal spending policies.

Therefore, we use our econometric setting to assess macroeconomic volatility with respect to the LMIs. We determine the variance of the endogenous variables of the IPVAR model conditional on the stringency of

[30] The LMIs that we consider capture structural labor market characteristics across distinct dimensions, however, they can, at least partly, be viewed as automatic stabilizers. Looking more closely at the individual LMIs, this is most evident with the BRR as an automatic stabilizer. In case of an adverse shock, a higher level of the BRR smooths household income over the business cycle and provides insurance for unemployment risk. The EPL and the UD work through similar channels. If these very elements already contribute significantly to cyclical smoothing by which they render any discretionary spending policy obsolete.

Figure 6: Macroeconomic Volatilities Along LMIs.



Notes: Each sub-figure shows the standard deviation of the respective macroeconomic variable in a regime with low (10th quantile, dark gray) and high (90th quantile, light gray) labor market institutions while the remaining LMIs are at their median. The labor market institutions under consideration are union density (UD, η), unemployment benefit replacement rate (BRR, φ), and employment protection legislation (EPL, ς).

the LMIs.[31] The covariance matrix of the endogenous variables $\mathbf{y}_{it}$ in the IPVAR system is given by[32]

$$\text{vec}\left(\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)\right) = \left(\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}_t) \otimes \boldsymbol{F}(\boldsymbol{\vartheta}_t)\right)^{-1} \text{vec}\left(\boldsymbol{Q}(\boldsymbol{\vartheta}_t)\right), \tag{4.6}$$

where $\boldsymbol{I}$ is an identity matrix of dimension $K^2 = (Mp)^2$, $\boldsymbol{F}(\boldsymbol{\vartheta}_t)$ denotes the $K \times K$ companion matrix form of $\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}_t)$ with $j = 1, \ldots, p$, and $\boldsymbol{Q}(\boldsymbol{\vartheta}_t)$ denotes the $K \times K$ companion matrix form of the common-mean covariance matrix.[33] It can directly be seen from the equation that the $K \times K$ covariance matrix of the endogenous variables, $\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)$, depends on the interaction terms, $\boldsymbol{\vartheta}_t$.

The results are depicted in Figure 6, where we report the model-implied volatility for a low (10th quantile, dark gray) and high (90th quantile, light gray) value of the respective LMI.[34] We observe that the LMIs have a potentially volatility-reducing effect. For output, UD and EPL reduce the volatility but significantly only for the latter. In contrast, higher levels of BRR clearly lead to higher volatility in output. While we do not see strong differences in volatilities regarding the employment rate or the real wage for UD and EPL, higher levels of BRR also lead to a higher volatility of the employment rate. We do not see strong differences for

[31] We redo this analysis in the theoretical model as well, see Appendix A.4. There, however, we condition on either a government spending or a technology shock.

[32] The definition of the companion form can be found in standard time series text books, e.g., Hamilton (1994) or Kilian and Lütkepohl (2017). The exact formula for the variance of the endogenous variables in the VAR system is 10.2.18 in Hamilton (1994), which we have adapted for the case of the IPVAR.

[33] We do not impose a pooling prior on the country-specific covariance matrices. Therefore, we use ex-post a common-mean approach by averaging across the country-specific covariance matrices, i.e., $\overline{\boldsymbol{\Sigma}}(\boldsymbol{\vartheta}) = N^{-1} \sum_{i=1}^{N} \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it})$.

[34] We abstain from reporting the implied volatilities when setting the respective LMI to its median value. Due to the standardization of the data, the median is zero and thus we only need half of the parameters for inspecting the median. The decreased number of involved parameters is another source of variance minimization which we do not want to exploit.

the real wage.[35] The theoretical model aligns with these results. It suggests sizable decreases for UD, and no strong effects for EPL. For BRR, the theoretical model points to higher volatility for less stringent LMI in contrast to the empirical findings. Most importantly, the theoretical results in this context highlight that the ability of the LMIs in attenuating macroeconomic volatility crucially depends on the shocks' sources, as we discuss in more detail in Appendix A.4. To conclude, UD and EPL reduce volatility, while BRR leads to higher levels of volatility.

4.6 How Do Labor Market Institutions Change the Effect on Macroeconomic Volatility?

In a next step, we are interested in how the LMIs dampen macroeconomic volatility. On the one hand, it is possible that they affect the propagation mechanism (transmission channel) of exogenous shocks but leave the size of the contemporaneous impact of exogenous shocks unaffected. On the other hand, the opposite could apply equally well. In the following, we discuss this issue in more detail.

As is evident from Eq. (4.6), the effect of the LMIs on the volatility of variable k occurs along two distinct dimensions. These concern the *transmission channel* of exogenous shocks (shock transmission effect, henceforth STE) or the *size* of the contemporaneous impact of exogenous shocks (shock size effect, henceforth SSE). More formally, we are interested in the partial effect of $\vartheta_{lt} \in \boldsymbol{\vartheta}_t$ on ω_{kk}, which is the k-th element on the main diagonal of the matrix $\boldsymbol{\Omega}(\boldsymbol{\vartheta}_t)$ denoting the volatility of the kth variable in the vector of endogenous variables $\boldsymbol{y}_{it}$. Furthermore, we denote with $\tilde{\boldsymbol{F}}(\boldsymbol{\vartheta}_t) \equiv (\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}_t) \otimes \boldsymbol{F}(\boldsymbol{\vartheta}_t))^{-1}$ and $\tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}_t) \equiv \text{vec}\,(\boldsymbol{Q}(\boldsymbol{\vartheta}_t))$. In the following, small letters denote scalars and refer to an element of the corresponding vector $\tilde{q}_j(\boldsymbol{\vartheta}_t) \in \tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}_t)$ or matrix $\tilde{f}_{kj}(\boldsymbol{\vartheta}_t) \in \tilde{\boldsymbol{F}}(\boldsymbol{\vartheta}_t)$. Then $\omega_{kk}(\boldsymbol{\vartheta}_t) = \sum_j \tilde{f}_{kj}(\boldsymbol{\vartheta}_t) \cdot \tilde{q}_j(\boldsymbol{\vartheta}_t)$ holds, and the partial effect is given by

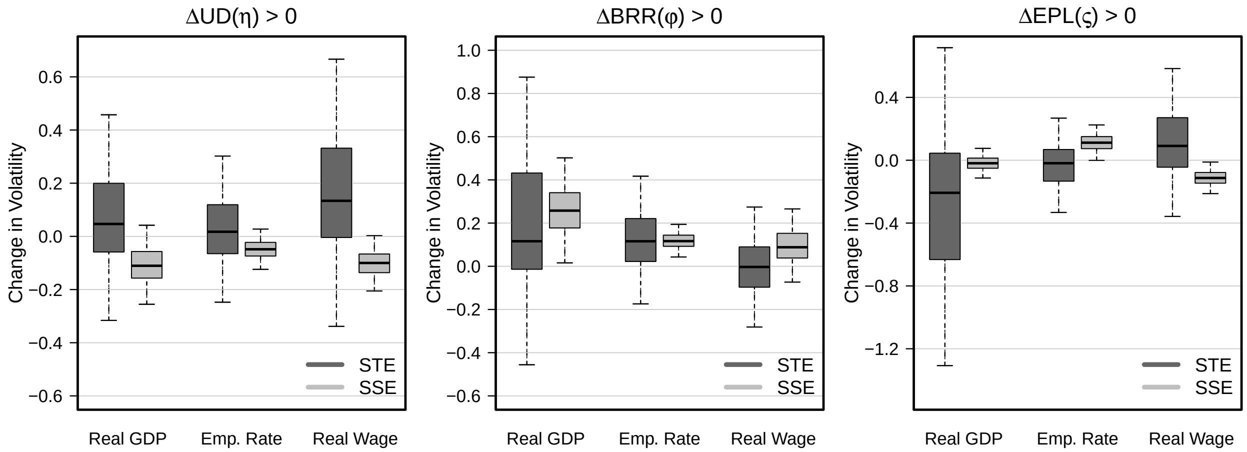
$$\frac{\partial \omega_{kk}(\boldsymbol{\vartheta}_t)}{\partial \vartheta_{lt}} = \sum_j \Bigg(\underbrace{\tilde{q}_j(\boldsymbol{\vartheta}_t) \frac{\partial \tilde{f}_{kj}(\boldsymbol{\vartheta}_t)}{\partial \vartheta_{lt}}}_{\text{STE}} + \underbrace{\tilde{f}_{kj}(\boldsymbol{\vartheta}_t) \frac{\partial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

which allows us to decompose the overall change in the volatility with respect to the LMIs along the proposed dimensions: STE and SSE. For each of the two cases, the signs of the partial derivatives allow for an exact identification of the partial effect.

The results are provided in Figure 7, where we show the change in the volatilities ($\Delta\omega_{kk}(\boldsymbol{\vartheta}_t)$) when moving from loose to stringent LMIs ($\Delta\vartheta_l > 0$). The overall effect ($\Delta\omega_{kk}(\boldsymbol{\vartheta}_t)/\Delta\vartheta_{lt}$) is decomposed into the STE (dark gray box-plots in each panel) and SSE (light gray box-plots in each panel). For a better

[35] This results are robust to exchanging the employment rate with the unemployment rate, as discussed in Appendix E.4.

Figure 7: Change in Macroeconomic Volatilities Along LMIs.



Notes: Each sub-figure 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 of the respective labor market institution. STE (dark gray) refers to the *shock transmission effect*, while SSE (light gray) refers to the *shock size effect* as depicted in Eq. (4.7). The labor market institutions under consideration are union density (UD, η), unemployment benefit replacement rate (BRR, φ), and employment protection legislation (EPL, ς).

understanding, consider the change in the output volatility that arises from an increase in the UD, which is displayed in the left panel. The higher UD affects output volatility both along the STE and SSE. The median change in the output volatility is slightly above zero according to the STE. This implies that a higher UD causes higher output volatility by reinforcing the propagation mechanism of shocks. While the STE gives rise to an endogenous reinforcement of shocks, the opposite applies to the SSE as according to which a higher UD attenuates the output volatility due to a smaller contemporaneous impact of shocks. The results are similar as regards the employment rate volatility: on the one hand, a higher UD causes higher volatility by reinforcing the propagation mechanism of shocks; on the other hand, a higher UD attenuates the employment volatility due to a smaller contemporaneous impact of shocks. Most clearly, we see these effects again for the real wage. The effects are sizeable, however, since they drift in opposite directions, the overall effect as depicted in Figure 6 is hence negligibly small.

In the case of the BRR, we find that volatilities are higher due to both STE and SSE. We observe this most clearly for output but to a lesser degree also for the employment rate volatility. Again, these results align well with the findings depicted in Figure 6. Lastly, we investigate the change in the volatilities for the EPL. Results suggest that the large decline in volatility for output visible in Figure 6 is mostly driven by the STE. The SSE for output volatility is basically null. The other two variables do not show a strong mitigation of the involved volatilities. The SSEs seem to matter with opposing signs but with small absolute effects.

To sum up, our results show that both the STE and the SSE matter with respect to the impact of the LMIs for the output and employment volatility.[36] The decline in the output volatility for UD and EPL is coming from different sources. UD seems to affect less the transmission of the effects while EPL does so sizably. On the contrary, the BRR paints a different picture. Both transmission and size play a role for the increase in output and employment rate volatility. This exercise reveals that the direction of the effect is in some cases opposite, which gives rise to an overall small effect.

4.7 Accounting for the Evidence

We move on to the exploration of structural differences that arise from different labor market institutions. Possibly, these differences may not only give rise to but also reinforce structural disparities in other dimensions. For example, a higher benefit replacement rate could strengthen the bargaining position of workers in wage negotiations, as the outside option (i.e., not working) becomes more attractive due to higher non-work income. This, in turn, could influence wage and price dynamics, thereby shaping the degree of nominal rigidities within the economy. Nominal rigidities, in turn, play a crucial role in determining the fiscal multiplier in the context of fiscal spending shocks.

In this regard, we assess the extent to which the theoretical model can explain the empirical evidence. We estimate key structural parameters of the theoretical model by matching the impulse response functions of the theoretical and empirical models. The purpose of this exercise is to contrast the theoretical model predictions for the effects of government spending shocks under different values (low vs. high) of the LMIs.

We fix all structural parameters to their values as depicted in Table A1, except the parameters capturing the persistence of the government spending shock (ϱ_g), the extent of price stickiness (ξ) and the share of prices that is adjusted in line with the previous period's inflation rate (γ_p), the elasticity of matching with respect to unemployed workers (γ), and the first and second moment of the distribution $F(\cdot)$ of idiosyncratic job destruction (μ_a, σ_a), which are all collected in the vector $\boldsymbol{\theta}$. The parameter selection is motivated by their significant role in shaping fiscal spending multipliers (Galí, López-Salido and Valles, 2007; Dupaigne and Fève, 2016).

In order to obtain estimates for these parameters, we match empirical (IPVAR) and theoretical impulse responses (see, for instance, Rotemberg and Woodford, 1997). Let $\widehat{\text{IRF}}$ be the empirical impulse response

[36] These results are robust to exchanging the employment rate with the unemployment rate, as discussed in Appendix E.4.

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 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.

functions obtained from estimating the IPVAR[37], and let IRF($\boldsymbol{\theta}$) be its counterpart from the theoretical model. We focus on the first 25 periods of the responses of government spending ($\hat{g}_t$), output ($\hat{y}_t$), employment ($\hat{n}_t$), and the real wage ($\hat{w}_t$). We estimate the parameter vector $\boldsymbol{\theta}$ by minimizing the distance[38] between empirical and theoretical impulse response functions under low and high values of the LMIs.

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

where $\widehat{\mathrm{IRF}}$ and IRF($\boldsymbol{\theta}$) are column vectors[39] of impulse responses. We establish estimates for $\hat{\boldsymbol{\theta}}$ for (i) each LMI individually, and (ii) separately for low and high values.

The parameter estimates are provided in Table 1. The estimated autocorrelation coefficient (ϱ_g) for the government spending shock reflects a medium-high persistence and a value at the lower bound compared to those commonly found in the literature (see, for instance, Born, Juessen and Müller, 2013). Most interestingly, the estimates are the same across labor market regimes (that is, whether a high or a low value of any LMI is considered). Moreover, we also observe that the second moment of the distribution function of idiosyncratic job destruction (σ_a) is rather stable across labor market regimes.

We find a noteworthy variation of the remaining parameter estimates across distinct labor market regimes. With respect to union density (UD), we observe a sizeable change in the estimates of γ and μ_a, both of which crucially shape employment dynamics. The estimates suggest that the transition to a labor market

[37] Specifically, as always done throughout the paper, we focus on the impulse response functions of the *common mean* in the model. The *common mean* is an estimate, which encompasses all the information from the respective single-country models but still allows for idiosyncrasies. We report the respective impulse response functions in Figure E1 in the Appendix.
[38] We stick to a procedure that produces parameter estimates which give rise to equilibrium determinacy (i.e., saddle-path stability).
[39] Their dimension is $(25 \cdot 4) \times 1$ in each case: 25 periods of the impulse response functions for four variables.

characterized by a higher union density renders employment more volatile owing to (i) a higher elasticity of matching unemployed workers (γ), and (ii) a higher idiosyncratic job loss probability (μ_a).[40] As a result, employment adjustments become more pronounced, which indirectly reflects the enhanced wage-setting power of workers and the corresponding response of firms to this shift. This contrasts with the effect of the benefit replacement rate (BRR), where a higher replacement rate increases nominal rigidities, while leaving most of the other parameter estimates largely unchanged. The higher extent of nominal rigidities is reflected on the one hand by a lower frequency of price adjustments (ξ) and a higher share of inflation indexed prices, both of which raise the extent of nominal rigidities and hence exert upward pressure on the fiscal spending multiplier in the short run. Finally, the degree of employment protection (EPL) primarily influences the mean of idiosyncratic job destruction (μ_a). Higher EPL is generally associated with a lower average probability of job destruction, which reflects the intended purpose of EPL—to provide greater job security. This reduction in job destruction, in turn, dampens the employment response to expansionary fiscal spending shocks, and consequently mitigates their effect on output. As a result, higher EPL moderates the fiscal spending multiplier.

5. Concluding Remarks

We have shown the eminent role of LMIs for fiscal policy and macroeconomic volatility alike. These results emerge from a theoretical model which combines the characteristics of a Diamond-Mortensen-Pissarides model with a standard New Keynesian setup; and from an interacted panel vector auto-regressive (IPVAR) model estimated for a panel data of 16 OECD economies. The empirical findings confirm the theoretical results and are robust to various extensions.

Our first key result highlights that the LMIs affect both the transmission channel of fiscal spending shocks, as well as the governments' quantitative ability in shaping output fluctuations. While this finding applies to all three LMIs under inspection, quantitative differences, though, emerge. These effects turn out strongest in the case of union density (UD) while being weaker with respect to employment protection legislation (EPL); unemployment benefit replacement rate (BRR) shows more mixed evidence. Using partial information matching, we find that the mechanism runs through variations in price rigidity, matching elasticity, and idiosyncratic job loss probability depending on the LMI.

[40] As outlined in the Appendix, the idiosyncratic productivity $\tilde{a}$ is assumed to be i.i.d. log-normally distributed with c.d.f. F of which we estimate 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 $\sigma_a \in \mathbb{R}^+$).

The second key result is that LMIs by themselves mute output volatility for UD and EPL. Our approach enables an assessment of distinct explanations in this respect. It permits to consider distinct observable structural elements for explaining volatility changes in macroeconomic time series and to assess whether distinct structural elements reduce macroeconomic volatility either by mitigating the propagation mechanism of shocks (STE) or by changing their contemporaneous impact characteristics (SSE). Results suggest that in several instances the STE is responsible for the reduction in volatility.

A key policy implication of our findings is that stringent labor market institutions render expansionary spending policies less effective while at the same time reduce the pain of fiscal consolidations. Moreover, while more stringent labor market institutions attenuate macroeconomic volatility, the fact that in some cases this occurs by attenuating the contemporaneous impact of shocks while concurrently exacerbating their propagation mechanism allows for the build up of risks and imbalances underneath a seemingly tranquil macroeconomic surface. This suggests a cautionary tale of stringent labor market institutions.

Declaration of Interest

The authors declare to 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 contains additional material not reported 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: $\hat{g}_t \sim$ 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 equations Eq. (3.2), Eq. (3.3) and Eq. (3.11) 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 equation Eq. (3.4), 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. We closely follow Christoffel, Kuester and Linzert (2009) for the choice of the 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 auto-correlation 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.

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

[1] Christoffel, Kuester and Linzert (2009) estimate a DSGE model with an extended labor market structure in their model based on the data for the euro zone. Since most of the countries in our sample are part of the euro zone, we hence rely on the estimates in Christoffel, Kuester and Linzert (2009).

[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

Table A1: Calibration of the Model.

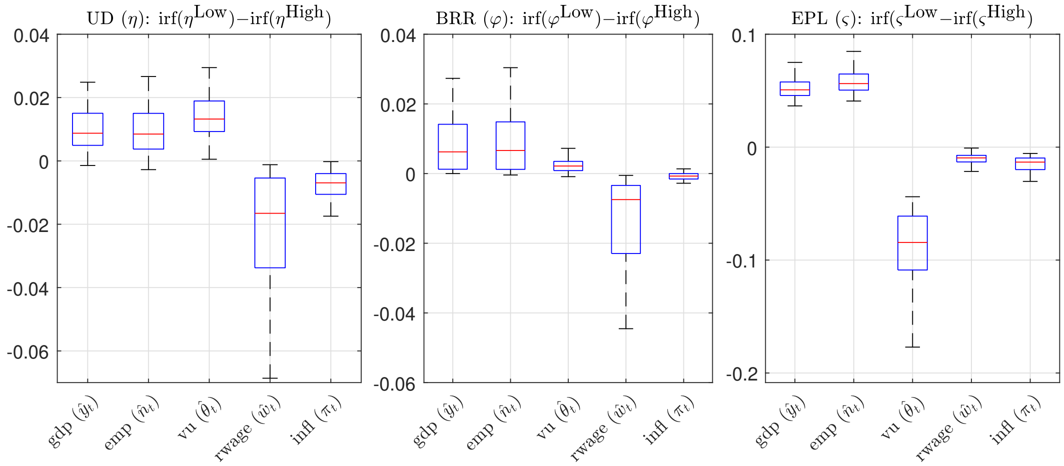
Parameter	Description	Value	Range
β	Discount factor	0.992	[0.95 – 0.999]
γ	Elasticity of matching of unemployed persons	0.45	[0.05 – 0.95]
g_y	Government consumption share in total output	0.3	[0.1 – 0.5]
ζ	Ratio of mrs to mpl	0.7	[0.55 – 0.95]
θ	Labor market tightness	0.43	[0.05 – 0.95]
p	Probability 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	—

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 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 3. 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.

cases holding the remaining parameters fixed. The difference in the impact values of the impulse response functions 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.



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 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 equation Eq. (3.14) and following Hamilton (1994, Chapter 10.2), the variance covariance matrix $\mathbf{\Sigma}_{\mathbf{z}}(\boldsymbol{\vartheta})$ of the vector of

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 ($\text{Var}(x_t(\text{LMI}_{\text{low}}))/\text{Var}(x_t(\text{LMI}_{\text{high}}))$). The shocks considered are a technology shock (ϵ_t^A) and the government spending shock (ϵ_t^G).

endogenous variables z_t of the solved DSGE model, is given by

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

where the dependency of $\boldsymbol{\Psi}_0$ and $\boldsymbol{\Psi}_1$ 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 auto-correlation 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 of one group of households, 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 always consider one extension at a time, as otherwise the precise role of the additional frictions considered becomes difficult to assess.

B.1 Inflation indexation of prices

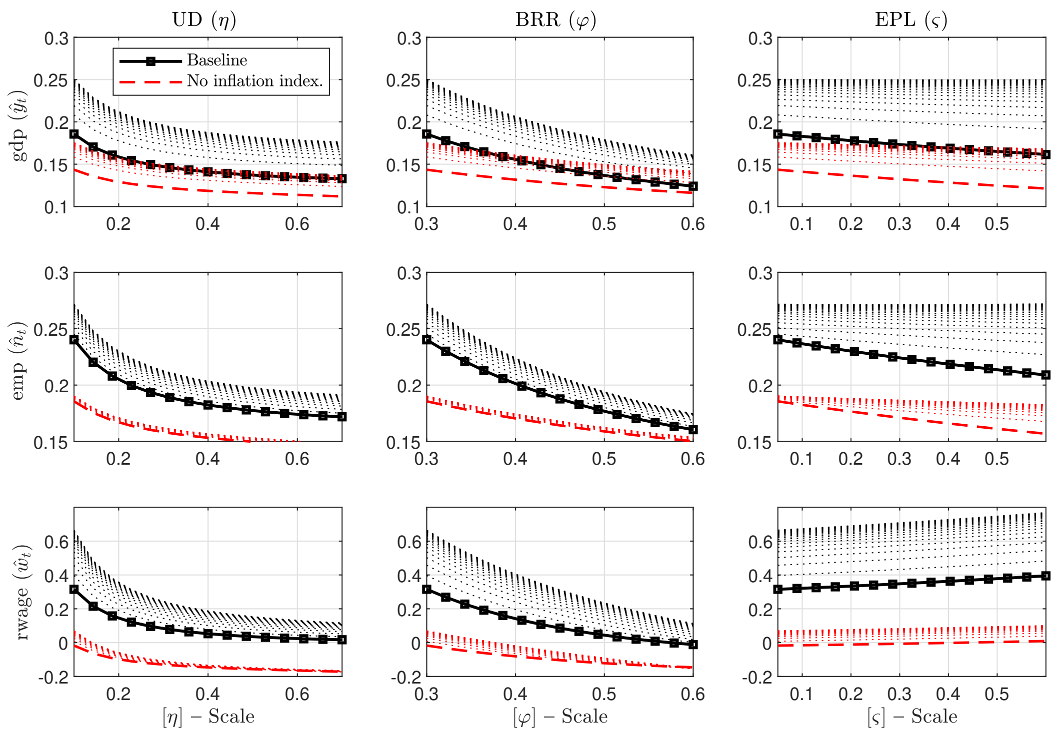
Our baseline model incorporates nominal frictions along two dimensions. The first dimension pertains to the infrequent adjustment of prices towards their optimal levels. The second involves a rule-of-thumb price adjustment mechanism for prices that cannot be optimally adjusted within a given period. 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 multiplier 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 to by many authors as a feature needed to account for a number of labor market facts (see Hall, 2005, among others). Krause and Lubik (2007) stress the role of real wage rigidity in the sort of models considered in Section 3 to improve the predictions of the labor market. Real wage rigidity might comprise 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

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



Notes: The sub-figures illustrate 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 presented 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 multipliers ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model without inflation indexation).

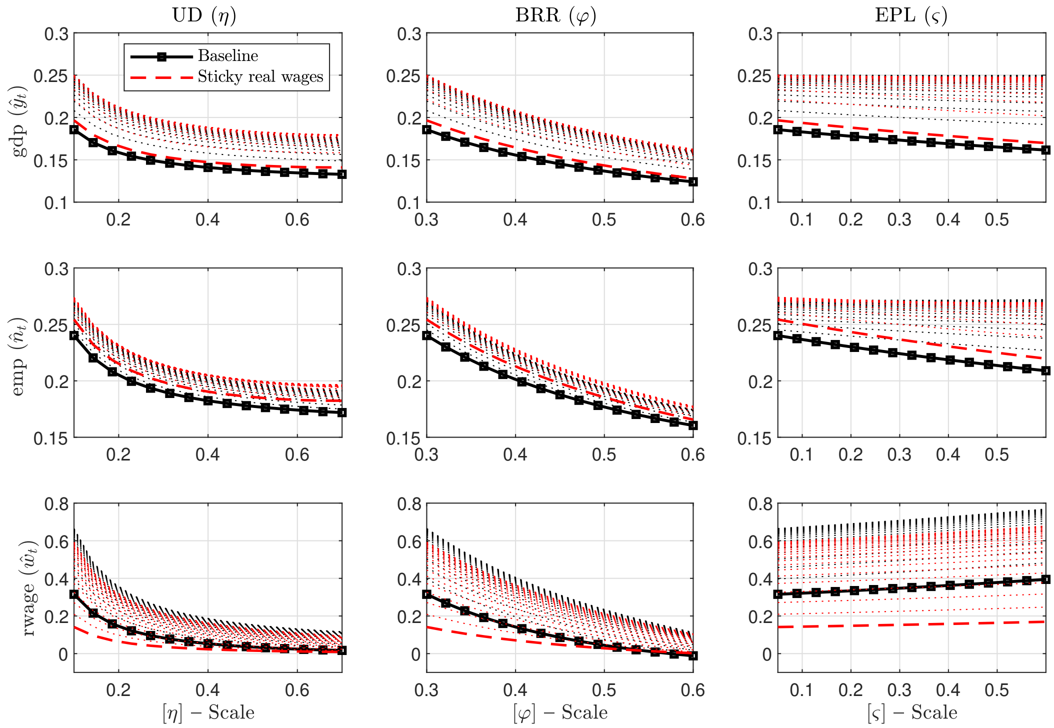
the same time, as vacancies rise and unemployment falls, there is a substantial increase in the cost of hiring workers (κ/q_t rises since q_t falls on the back of an increase in vacancies v_t) which are a component of firms' real marginal costs. Hence the role of rigid real wages can be confined to two elements, of which one becomes more rigid while the other 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 imperfection or friction in labor markets which are modeled in a reduced form. Specifically, we assume the partial adjustment model which extends equation Eq. (3.11) 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. Equation Eq. (B.1) can be considered as

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



Notes: The sub-figures illustrate 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 presented for different horizons, with the contemporaneous multiplier ($\mathcal{P}$ = 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 multipliers ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model with real wage rigidity).

a parsimonious but ad hoc way of modeling the sluggish adjustment of real wages to changes in labor market conditions, as found in a variety of models of real wage rigidities, without taking a stand on what the right model is. Alternative formalizations, explicitly derived from staggering of real wage decisions and alike, are presented in Blanchard and Galí (2007), Zanetti (2007), and Gertler, Huckfeldt and Trigari (2020) and the papers cited therein. The results of the model extended for real wage rigidities are shown and compared to the baseline model in Figure B2. Considering first the dependency of the output multiplier on φ and ς shown in in the sub-panels in the second and third columns, it can be seen that the shape of the output multiplier 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 multiplier. 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 multipliers are hence smaller when real wage rigidities are present.

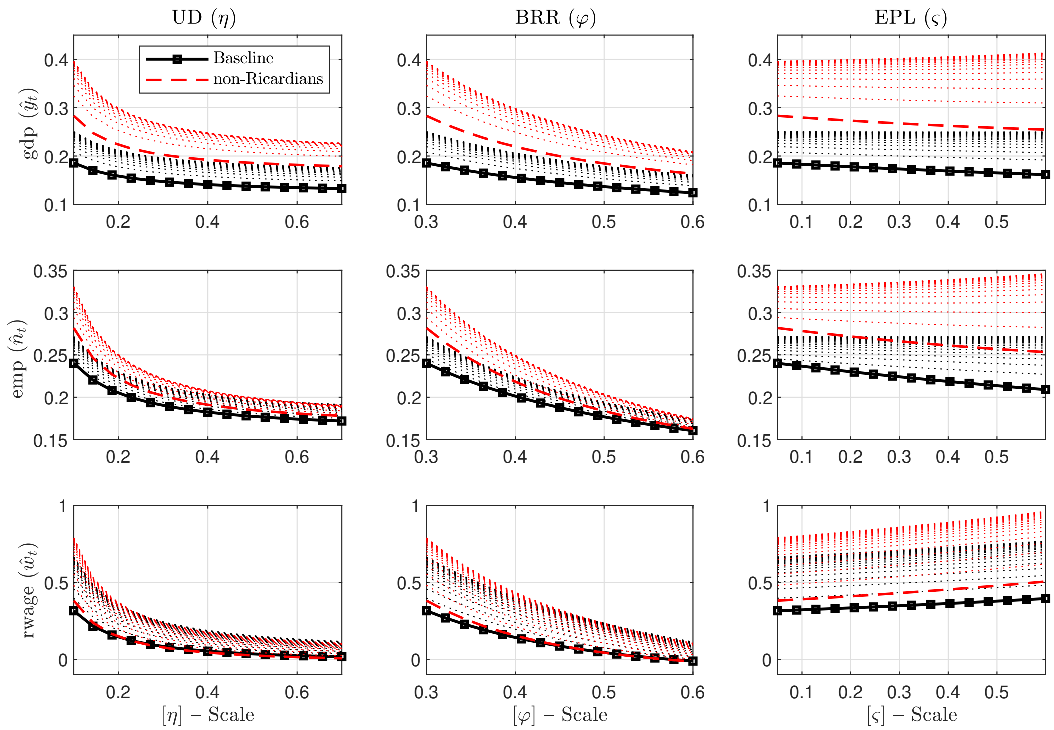
In case of η, the multipliers for output and employment are affected more profoundly when real wage rigidities are present. Both multipliers now show a concave pattern with respect to η: when η is low, increases therein raise fiscal spending multipliers, while the opposite occurs when η is already high. The intuition is that when η is low the drop in the benefits firms have to share with workers now dominate to increase in hiring costs giving rise to a positive dependency between η and the output and employment multipliers. For higher values of η, the dominance structure changes and the baseline results (higher η causes a smaller output multiplier) applies again. Nevertheless, a concave pattern shows up 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 any ability of smoothing their consumption profile; as a consequence, they spend (consume) each period all of their income. This rule-of-thumb gives rise to a consumption pattern that strongly aligns with wage income. This gives rise to 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 (same for their labor supply n_t). 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 which gives rise to wages being negotiated in a centralized manner by an economy-wide union with firms.

Figure B3 shows the results of the LMIs on the multipliers for output, employment, etc. 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 ($\lambda = 0.25$). As can be seen, the multipliers are throughout

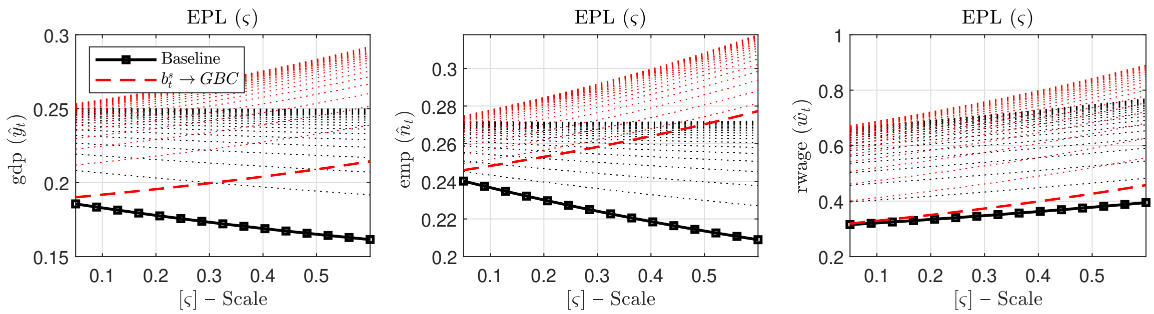
Figure B3: Fiscal Spending Multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – The Role of Limited Asset Market Participation



Notes: The sub-figures illustrate 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 presented for different horizons, with the contemporaneous multiplier ($\mathcal{P}$ = 1) represented by a solid black square line for the baseline model (λ_p = 0.0) and by a red dashed line for the model with limited asset market participation (λ_p = 0.25). The higher-horizon multipliers ($\mathcal{P} \geq 2$) are indicated by black dotted lines (baseline model) and red dotted lines (model without inflation indexation).

higher; this applies to both the output and employment multipliers, but also for the real wage. The reason for the higher multiplier throughout is due to the different reaction of consumption. In the baseline model, consumption declines owing to the negative wealth effect that comes along with 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 hence illustrates the amplifying effects of the introduction of rule-of-thumb consumers. Most important, though is the fact that the introduction of limited asset market participation does not change the dependency of the multipliers on the LMIs. With a view on 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 Multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – When Firing Costs Accrue to the Government



Notes: The sub-figures illustrate 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 presented for different horizons, with the contemporaneous multiplier ($\mathcal{P}$ = 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 multipliers ($\mathcal{P} \geq 2$) 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 a quite strong assumption, as in many countries firing costs arise in the context of severance payments, etc. which will eventually be 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 alike. Hence in this case, the government budget constraint (Eq. (3.12)) and the real resource constraint (Eq. (3.13)) are then 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 hence has no effect on the shape of the multipliers with respect to η and φ.

The results are shown in Figure B4 for output, employment and the real wage. As can be seen, when firing costs accrue to the government, fiscal spending multipliers 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 re-distributional element which 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 multiplier. In contrast to this, when firing costs enter the aggregate resource constraint, then this implies that they are real resource costs which cannot be uncovered. This loss attenuates the output multiplier; 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 get 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 which 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

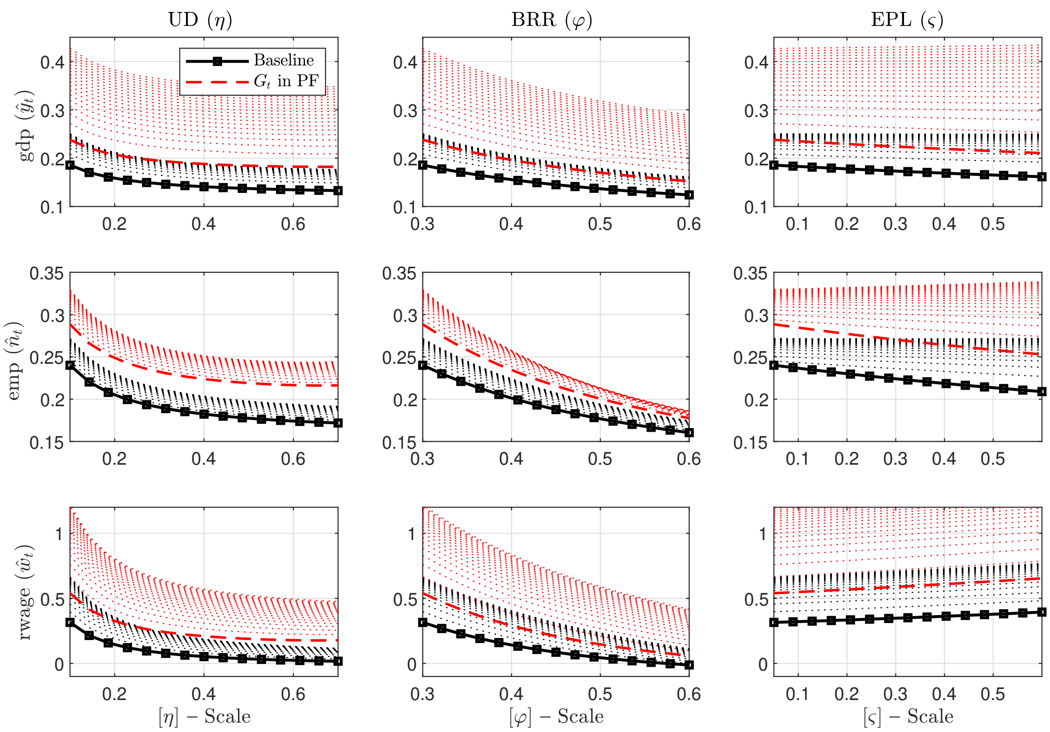
$$k_t^G = (1-\delta)k_{t-1}^G + g_t \tag{B.4}$$

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 of course 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 the 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 the public spending enters externally, in an analogous manner to exogenous technology. Hence government spending augments labor productivity

Figure B5: Fiscal Spending Multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – Productivity Enhancing Government Spending.

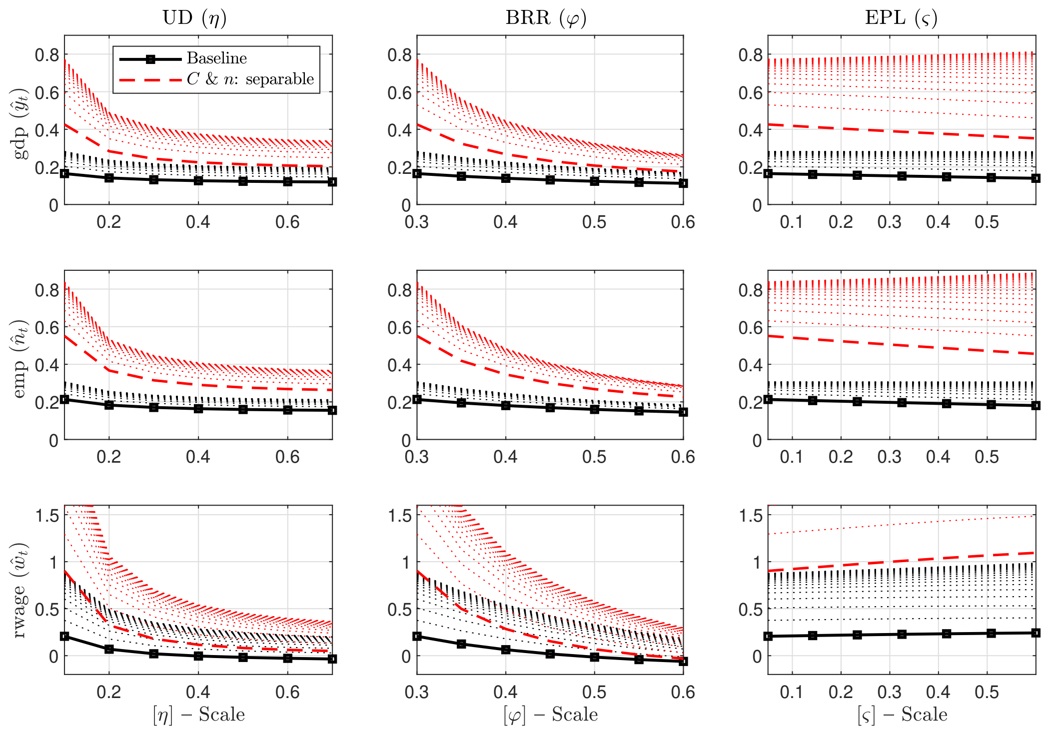


Notes: The sub-figures illustrate the sensitivity of the fiscal spending multipliers to variations in the three LMI parameters (UD–union density, BRR–(unemployment) benefit replacement rate and EPL–employment protection (legislation)). The multipliers are presented for different horizons, with the contemporaneous multiplier ($\mathcal{P} = 1$) 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 multipliers ($\mathcal{P} \geq 2$) 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 thereof are shown in Figure B5. As can be seen, productive government spending leads to a significantly higher output multiplier. At the same time, the employment multiplier 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, though, is the impact of the LMIs on the output multiplier. With a view to Figure B5, while the output response increases with the extent of productive government spending, the dependency of the output multiplier with respect to the three LMIs

Figure B6: Fiscal Spending Multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – Consumption/leisure complementarity.



Notes: The sub-figures illustrate 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 presented for different horizons, with the contemporaneous multiplier ($\mathcal{P} = 1$) 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 ($\sigma = 1$). The higher-horizon multipliers ($\mathcal{P} \geq 2$) 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 multiplier 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

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

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 $l = 1, \ldots, d$. Last, 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 into

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

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 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 such that we are inducing more shrinkage to higher-order lags. Furthermore, we shrink coefficients in the estimation equation to its common mean and the common mean towards zero. For the specification of the prior distribution, we start with stacking to a $k = (1 + d)n^2$-dimensional vector $\boldsymbol{\beta}_{ij} = \text{vec}(\boldsymbol{A}_{ij}, \boldsymbol{B}_{ij1}, \dots, \boldsymbol{B}_{ijd})$ for lag j and country i and specify the prior distribution as follows

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

Here $[\boldsymbol{\beta}_{ij}]_s$, $[\boldsymbol{b}_j]_s$, and $[\boldsymbol{\theta}_{ij}]_s$ denotes the s-th element of the respective vector. The latter one is the local-shrinkage component on which we specify a Gamma-distribution with hyperparameter ϑ_θ. This hyperparameter is governing 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 additionally 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 $\tilde{\boldsymbol{J}}_{it}$

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

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 for each element a simple Gaussian $\mathcal{N}(0, 100)$ to be uninformative. We have not yet discussed the prior distributions of the common means, $\boldsymbol{b}_j$ and $\boldsymbol{g}_l$. They do not feature a country-indicator i anymore, 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 thus 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 $[\boldsymbol{\phi}_j]_s$ denotes 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 as well 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 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, Netherlands, Portugal, Spain, Sweden, 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 for the model featuring employment / unemployment $N = 16$ countries while we only use $N = 13$ countries for the model with the labor market tightness indicator. Sample sizes also reduces 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.

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 in two groups according to their overall level of the LMIs or their consumption share in total GDP.

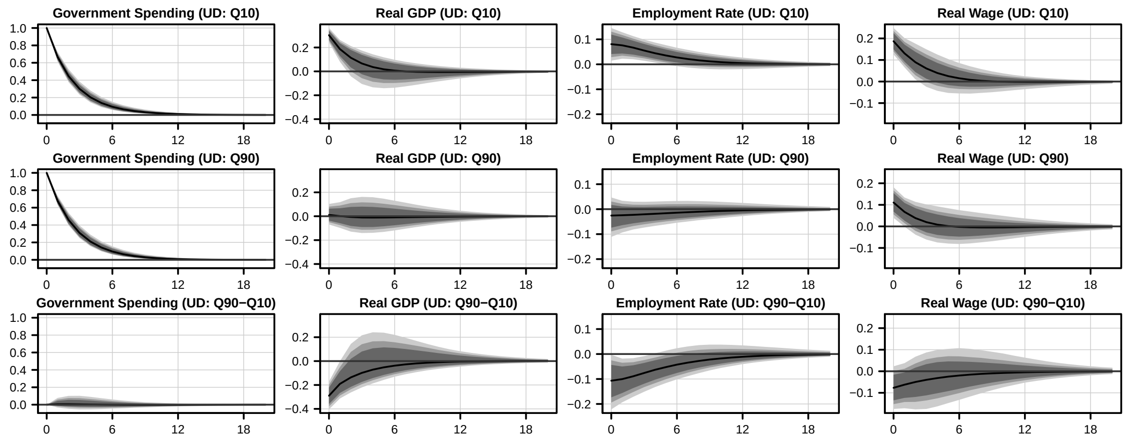
E.1 Dynamic Effects

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 4 where we focus on the evolution of the marginal effects. Here, we focus more on the dynamic evolution of the IRFs themselves.

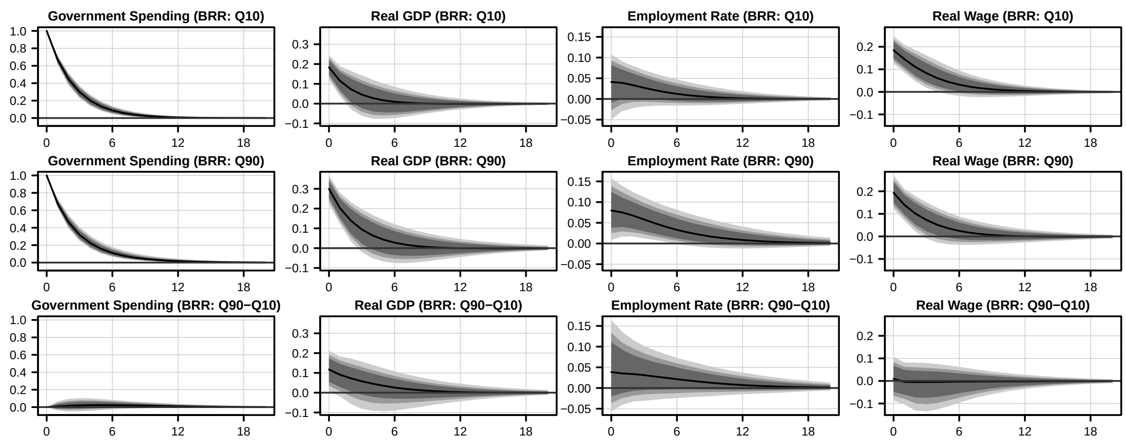
We report three panels, one for each LMI: union density (UD), unemployment benefit replacement rate (BRR), and employment protection legislation (EPL). We vary the value of the LMI as follows: Either we set it to the 10th (upper row) or the 90th (middle row) of its respective empirical distribution.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

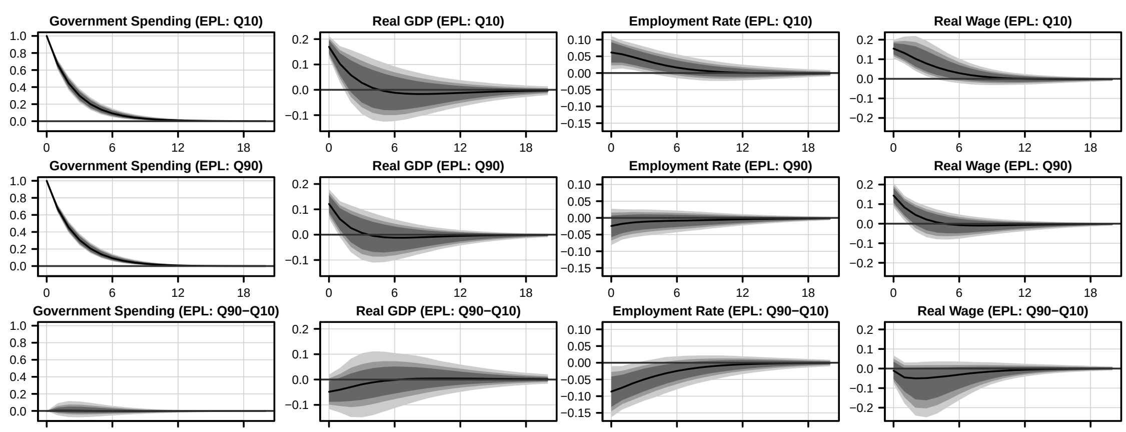
Figure E1: Dynamic Effects of Government Spending Shocks.



(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 replacment rate, and panel (c) the conditional effects of employment protection legislation.

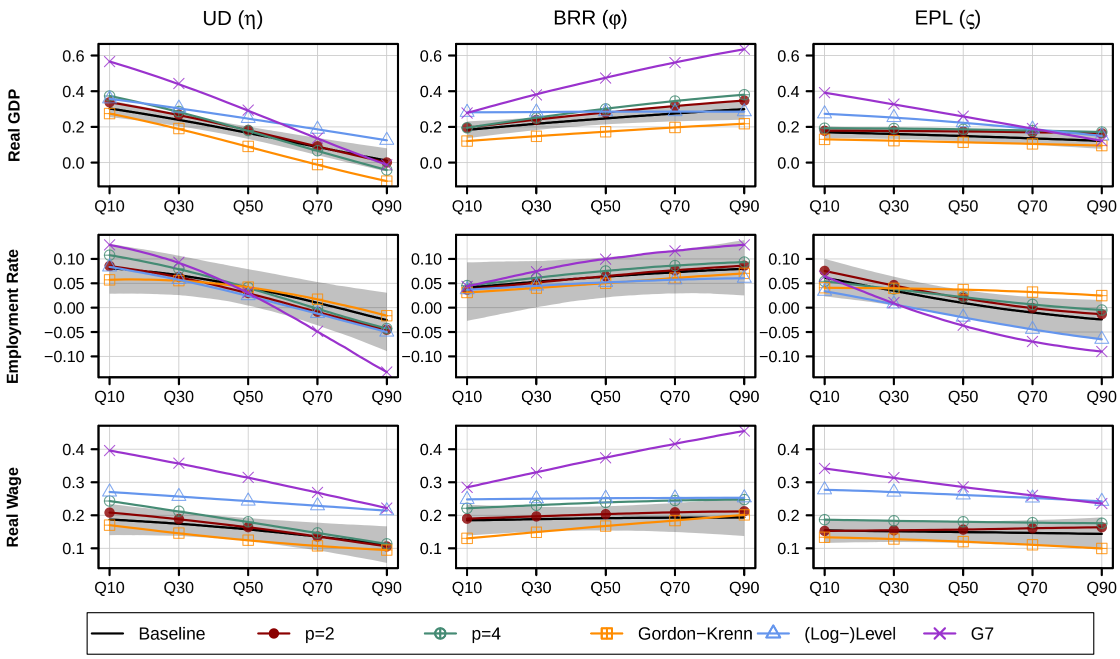
E.2 Robustness of Model Specification

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 check 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 the ones from the baseline model reported in Figure 4. Additionally, we add colored lines with distinct points to the figure for the alternative models.

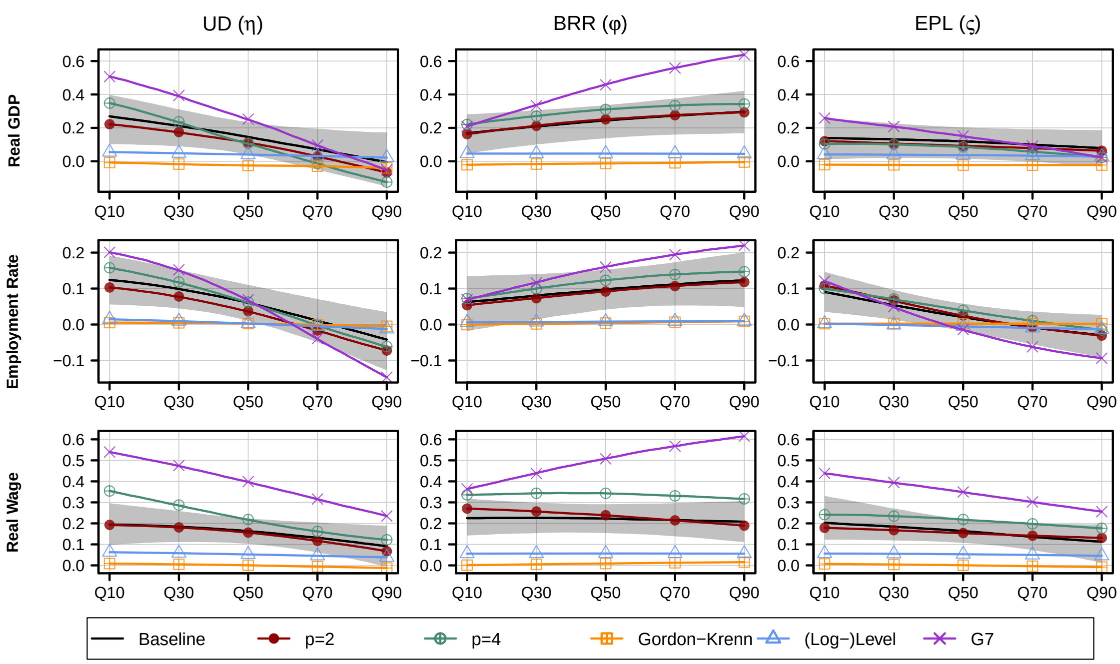
We have done the following robustness checks. First, we have varied the number of lags, using $p = 2$ and $p = 4$ lags. Results are robust to this choice, where the median estimates are within the 80% posterior credible sets. In almost all cases, both median estimates are within the credible sets of the baseline model. As a next check, we follow Ramey (2016) and transform variables with 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 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, besides 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, 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 do 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 sub-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 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, then effects are even magnified. The overall pattern, however, is remarkably robust.

E.3 Controlling for Fiscal Foresight

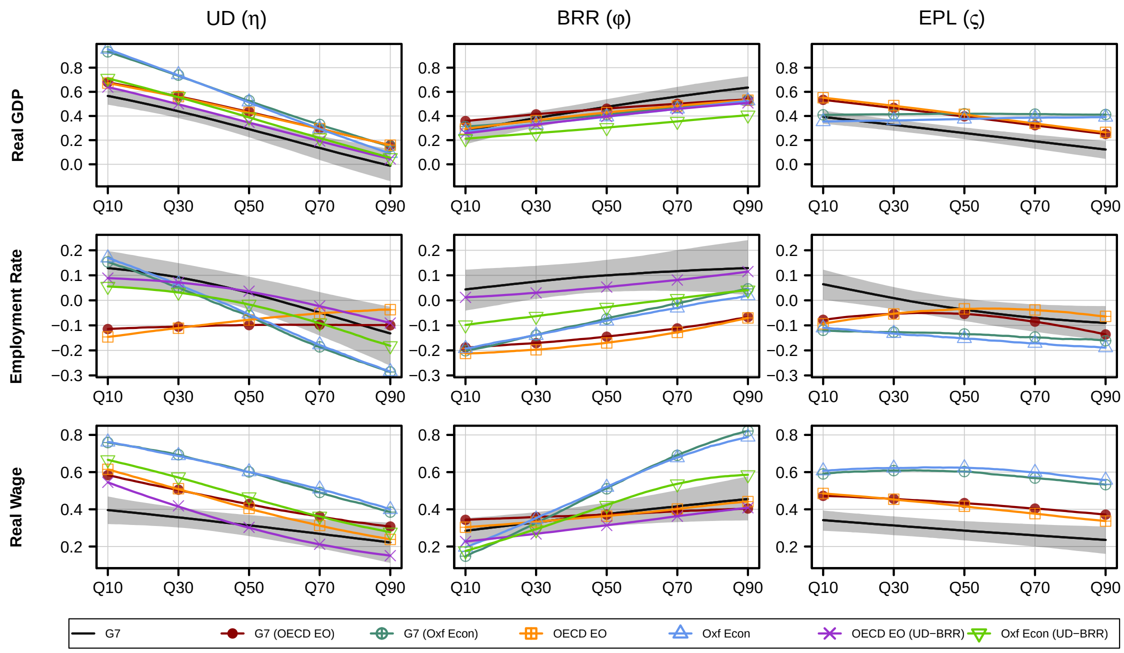
In this subsection, we explore the issue of fiscal foresight in more detail. As pointed out by Ramey (2011) or Leeper, Walker and Yang (2013), econometric models can suffer from informational insufficiency due to a misalignment between the information sets of the economic agents and the econometrician. 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 the one 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^e_{it}, to the vector of endogenous variables which yields $\boldsymbol{y}_{it} = (g_{it}, \Delta g^e_{it}, x_{it}, er_{it}, \omega_{it})'$.[3] We rely on two different data sources for the measurement of government spending forecasts, following suggestions in the literature on cross-country analyses (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

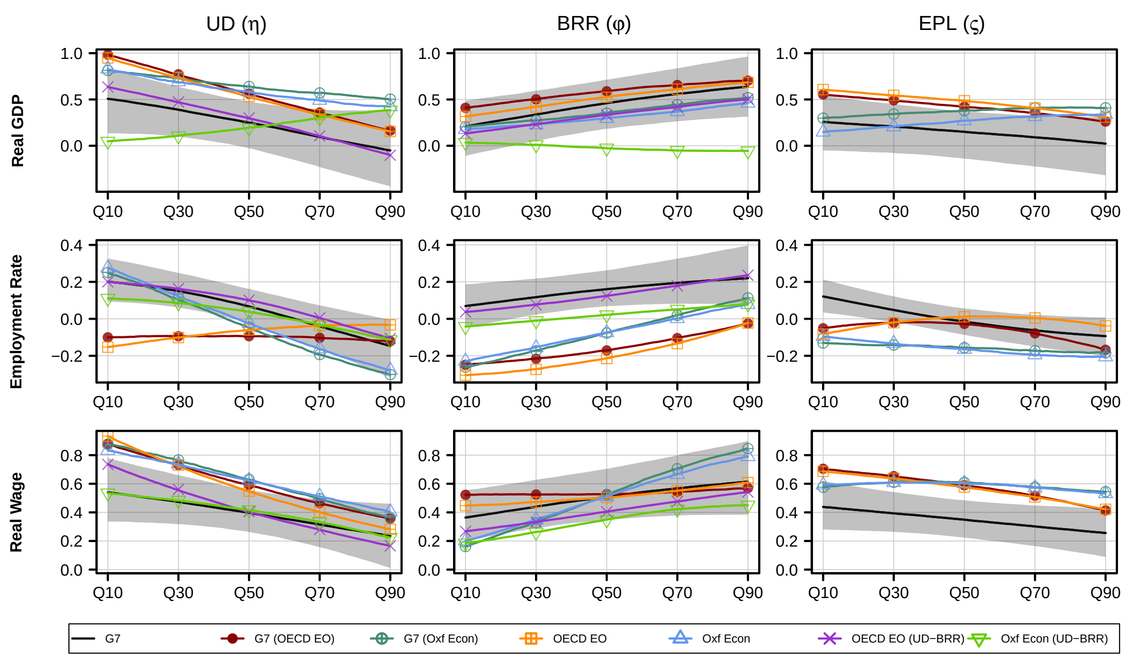
[3] Results are robust when exchanging the order of the variables, ordering the government spending forecast first.

[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 sub-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 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

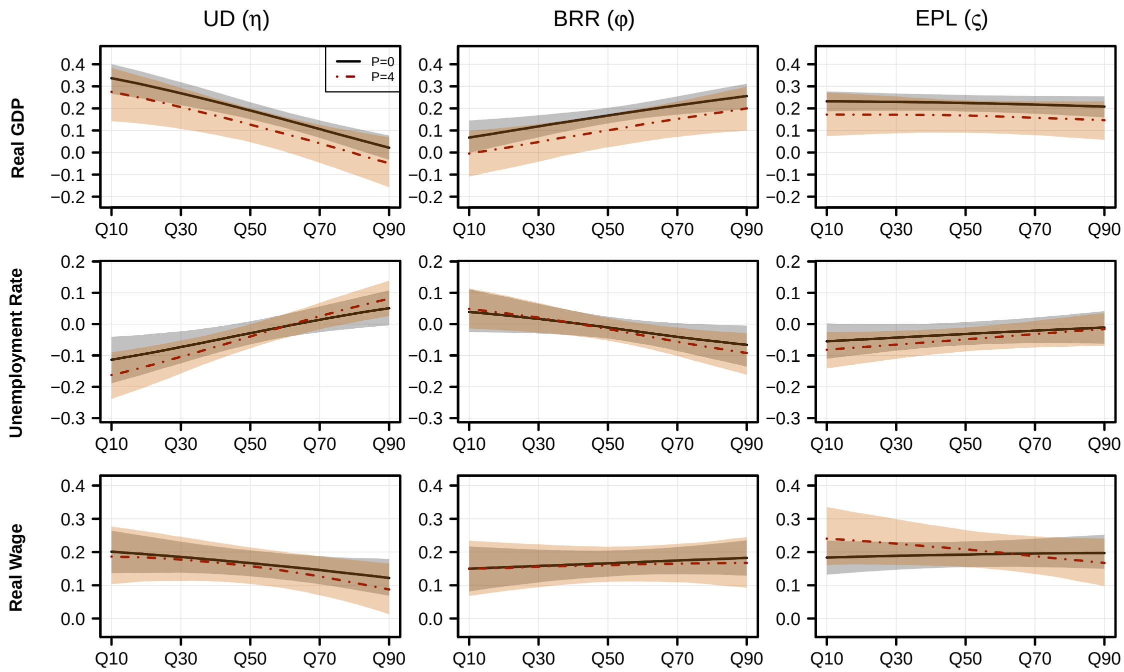
In this subsection, we explore other labor market indicators. We exchange the employment rate in the baseline model with either the unemployment rate or the labor market tightness (v_{it}/u_{it} locus). Similar to Figure 4, we report the contemporaneous and one-year fiscal multipliers.

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 strong 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 is not strongly reacting along the stringency of the LMIs.

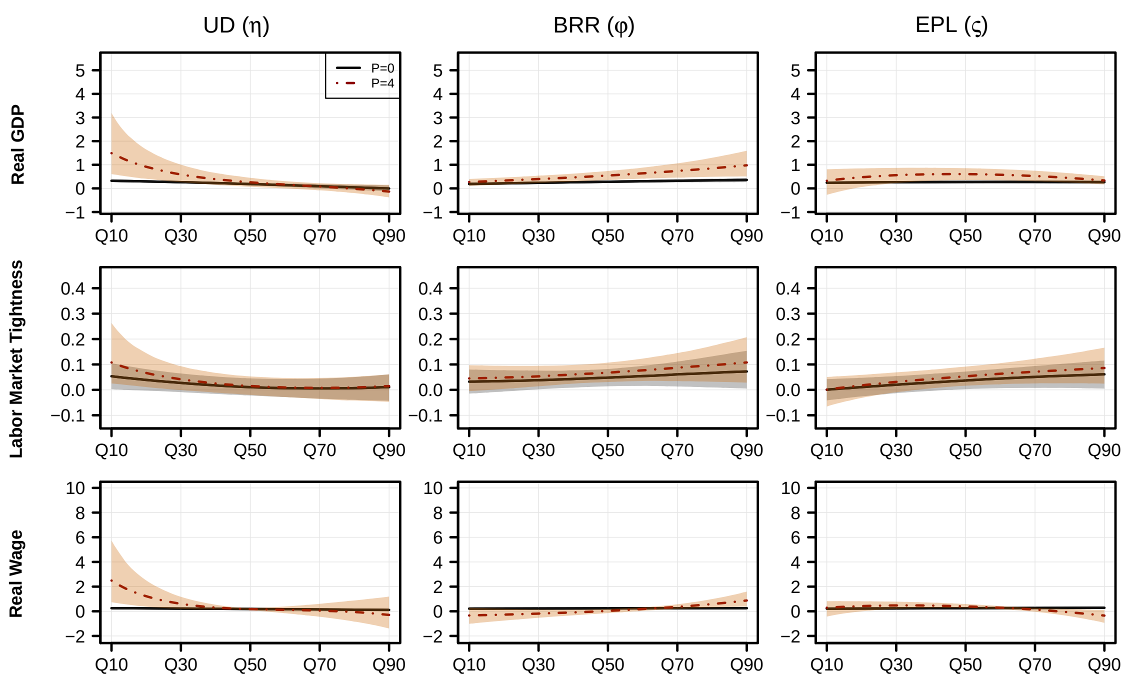
The model including labor market tightness is characterized by relatively large volatility around the estimates, pointing to instability. We have also added the employment rate to the model, as in the baseline. While the multipliers on impact are strongly centered around zero, we see an attenuation of 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.

Figure E4: Fiscal Multipliers Using Other Labor Market Indicators.



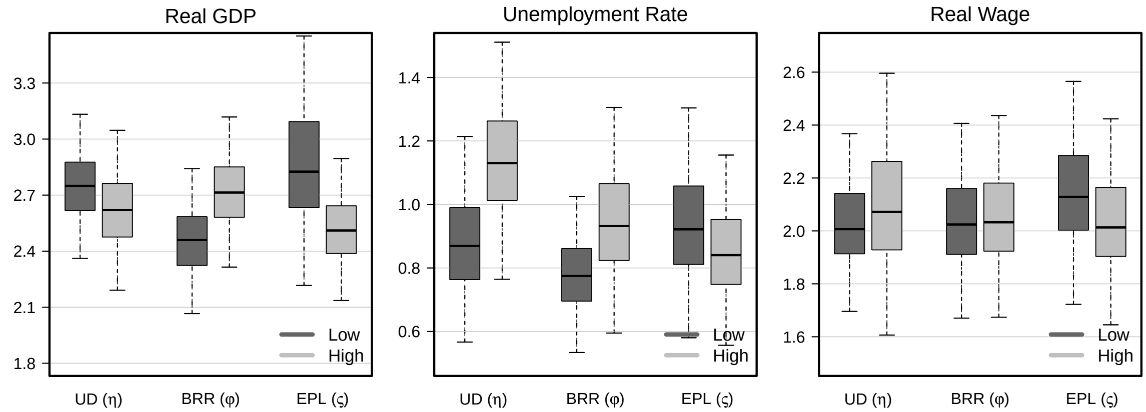
(a) Unemployment Rate.



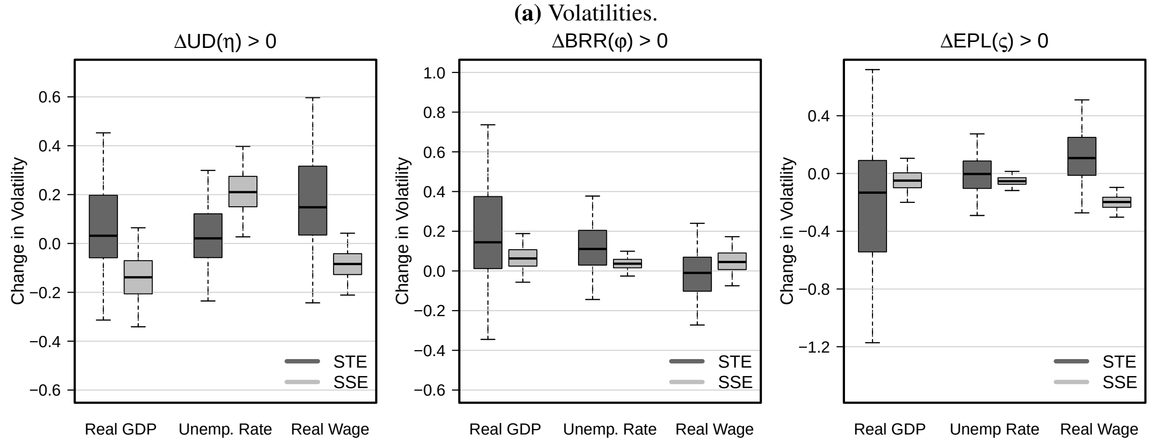
(b) Labor Market Tightness (v_{it}/u_{it}).

Notes: The sub-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 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.



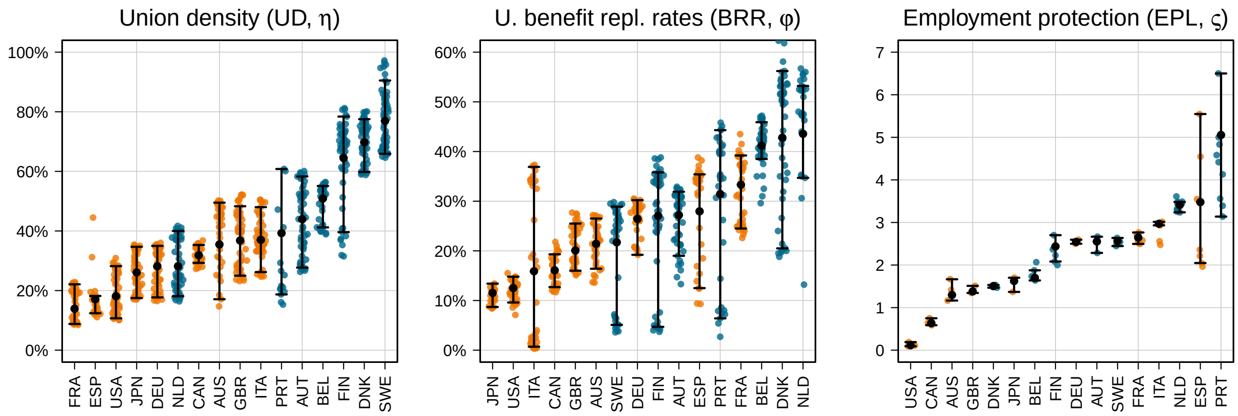
(b) Change in Volatilities.

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).

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 2 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 stringent 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

Figure E6: Classification of Countries.



Notes: Each sub-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).

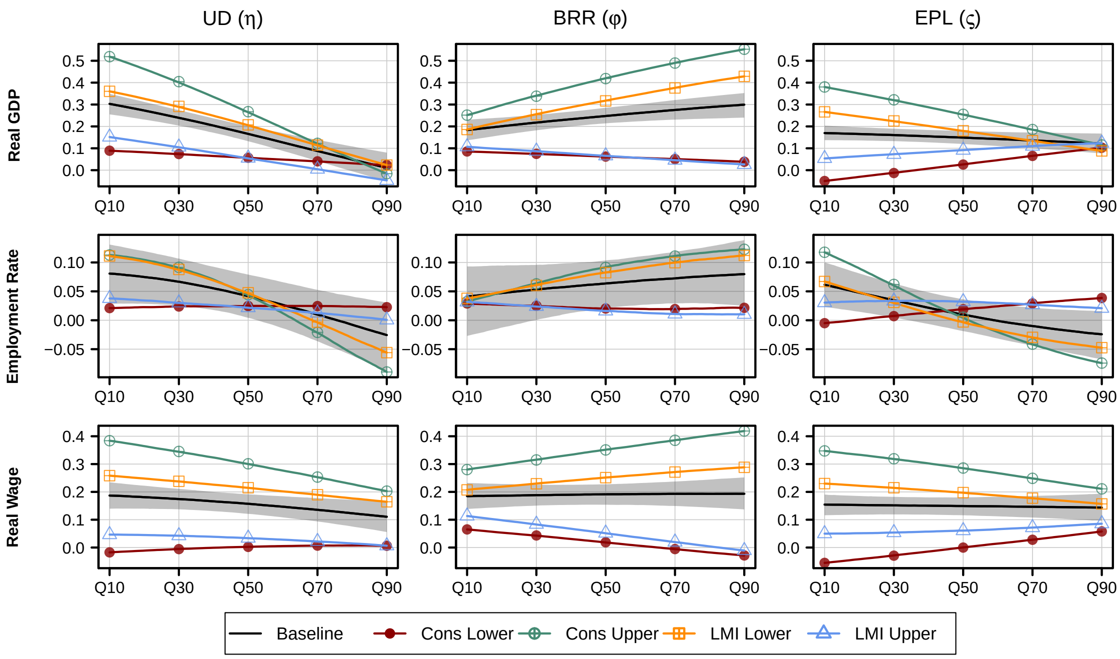
approaches we cluster the countries in two groups. Then, we re-estimate the model in Eq. (4.2) for both groups. We only allow for two groups to have enough variation in both country groups to estimate the IPVAR.

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 such that no variable has a stronger influence due to its scaling. In case a country is not classified entirely to one group, we apply a 50% rule: If more than 50% of the observations of one country are classified to one group, the country is classified to 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, Netherlands, and Sweden.

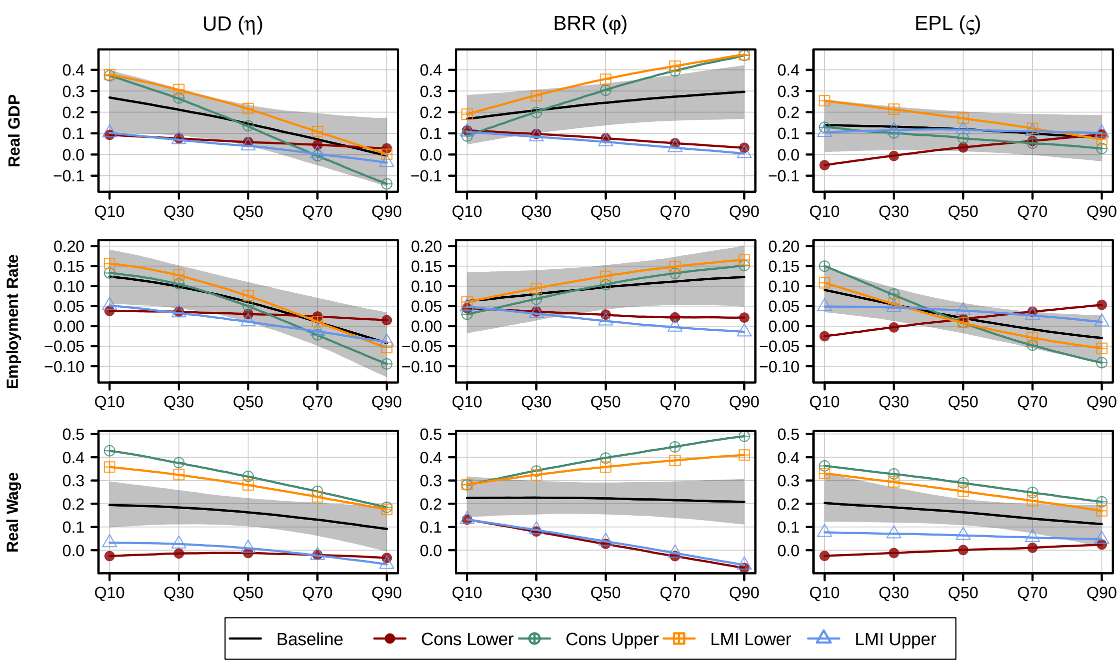
In Figure E7, we examine the fiscal multipliers both on a within- and between-country variation basis, where we report the on impact (panel (a)) and one-year (panel (b)) fiscal multipliers. The baseline results (black solid line) with 80% credible sets (gray area) are the ones from the baseline model reported in Figure 4. There is quite some 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

[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).

Figure E7: Cross-Country Heterogeneity.



(a) Fiscal Multipliers: On Impact.



(b) Fiscal Multipliers: One Year.

Notes: The sub-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 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.

yield generally larger fiscal multipliers than the baseline model. The marginal effects along the stringency of the LMIs shows also 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 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 on impact multipliers but also for the one-year multipliers.

These results are in line with conventional predictions. Countries with a higher consumption share are stronger dependent upon output fluctuations by 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 also the *level* of the LMIs matter and not only their *change* over time. These results have to 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.

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