CESifo Working Papers

9749
2022

May 2022

Labor Market Institutions, Fiscal Multipliers, and Macro-economic Volatility

Maximilian Boeck, Jesús Crespo Cuaresma, Christian Glocker

CESifo

Impressum:

CESifo Working Papers
ISSN 2364-1428 (electronic version)
Publisher and distributor: Munich Society for the Promotion of Economic Research - CESifo GmbH
The international platform of Ludwigs-Maximilians University's Center for Economic Studies and the ifo Institute
Poschingerstr. 5, 81679 Munich, Germany
Telephone +49 (0)89 2180-2740, Telefax +49 (0)89 2180-17845, email office@cesifo.de
Editor: Clemens Fuest
https://www.cesifo.org/en/wp
An electronic version of the paper may be downloaded
· from the SSRN website: www.SSRN.com
· from the RePEc website: www.RePEc.org
· from the CESifo website: https://www.cesifo.org/en/wp

CESifo *Working Paper No. 9749*

Labor Market Institutions, Fiscal Multipliers, and Macroeconomic Volatility

Abstract

We study empirically how various labor market institutions – (i) union density, (ii) unemployment benefit remuneration, and (iii) employment protection – shape fiscal multipliers and output volatility. Our theoretical model highlights that more stringent labor market institutions attenuate both fiscal spending multipliers and macroeconomic volatility. This is validated empirically by an interacted panel vector autoregressive model estimated for 16 OECD countries. The strongest effects emanate from employment protection, followed by union density. While some labor market institutions mitigate the contemporaneous impact of shocks, they, however, reinforce their propagation mechanism. The main policy implication is that stringent labor market institutions render cyclical fiscal policies less relevant for macroeconomic stabilization.

JEL-Codes: E620, C330, J210, J380.

Keywords: fiscal policy, fiscal multipliers, labor market institutions, interacted panel VAR.

*Maximilian Boeck**
Department of International Economics
Vienna School of International Studies
Favoritenstrasse 15a
Austria - 1040 Vienna
maximilian.boeck@da-vienna.ac.at

Jesús Crespo Cuaresma
Department of Economics
Vienna University of Economics and Business
Welthandelsplatz 1
Austria - 1020 Vienna
jesus.crespo.cuaresma@wu.ac.at

Christian Glocker
Austrian Institute of Economic Research
Arsenal Objekt 20
Austria - 1030 Vienna
christian.glocker@wifo.ac.at

*corresponding author

May 2022

1 Introduction

In this paper, we contribute to a recent literature that examines the role of labor market institutions (henceforth *LMIs*) as a determinant of business cycle fluctuations. The motivation arises from the discussion about the respective roles of structural reforms and cyclical policies for macroeconomic stabilization. The debate is centered around the question whether an enhanced fiscal architecture that fosters the conduct and effectiveness of cyclical policies is to be preferred over labor market reforms (see Banerji *et al.*, 2017; Arnold *et al.*, 2018; Sondermann, 2018; Masuch *et al.*, 2018; Duval and Furceri, 2018; Aiyar *et al.*, 2019, for instance). Several recent studies (see Zanetti, 2009; 2011; Abbritti and Weber, 2018; Cacciatore *et al.*, 2016, among others) highlight the ability of LMIs to mitigate macroeconomic volatility. Such results challenge the use of traditional cyclical fiscal policy, whose main objective is to smooth fluctuations in economic activity. The success of fiscal policy depends crucially on the size of fiscal multipliers, which, however, change considerably over time (see Auerbach and Gorodnichenko, 2012; Ilzetzki *et al.*, 2013; Leeper *et al.*, 2017, among others). In turn, the ability of LMIs to mitigate macroeconomic fluctuations underscores their important role in shaping the transmission channel of exogenous shocks. This, however, also raises the question of the extent to which LMIs affect cyclical fiscal policy and in particular whether they reinforce or attenuate the effects of discretionary fiscal spending policy on the economy.

We provide an analysis of the role of LMIs as determinants of cyclical macroeconomic fluctuations that centers around the following questions: How do LMIs shape fiscal spending multipliers? How do LMIs affect macroeconomic volatility and along which channel – by shaping the transmission of exogenous shocks, or by affecting the contemporaneous impact of shocks? What is the role of cyclical fiscal spending policy when stringent LMIs are in place? The answer to the first question will allow us to assess the role of LMIs in shaping the effectiveness of fiscal policy, while the answer to the second one helps us to judge their ability to mute cyclical fluctuations. This eventually allows an assessment of the degree of complementarity (or substitutability) among different LMIs and cyclical spending policies and hence provides an answer to the third question. Likewise, this will enable a direct comparison of structural versus cyclical policies in smoothing fluctuations in economic activity.

We start by developing a theoretical model to assess the role of LMIs in shaping (i) fiscal spending multipliers and (ii) macroeconomic volatility. We consider a set-up that combines the characteristics of a Diamond-Mortensen-Pissarides model with those of a standard real business cycle model structure. We capture union density in the theoretical model by means of workers' bargaining power within the wage negotiations, the unemployment benefit replacement rate by means of subsidies to the unemployed which are proportional to their previous wage, and employment protection by the level of firing costs per displaced worker. We confront the predictions of the model with the data by estimating a Bayesian panel vector-autoregressive (PVAR) specification for 16 OECD economies, where we identify the effects of LMIs on fiscal spending multipliers and macroeconomic volatility by means of interaction terms. The structural interpretation of the econometric model relies on two main building blocks. First, 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. Additionally, the structure of the empirical model allow us to estimate and analyze the variation of LMIs on a lower frequency together with time-series data on a higher frequency. Importantly to note, we only utilize within-country variation to estimate our baseline model. In addition, we also inspect cross-country heterogeneity. We abstain from estimating the theoretical model directly because due to the availability of LMIs only on a lower frequency within a limited time sample renders the panel approach particularly useful. Nevertheless, we show the strong similarity between the outcomes of the DSGE model and the interacted PVAR model. Both allow for an evaluation of the sensitivity of the endogenous shock transmissions with respect to changing structural properties. Second, 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 *et al.* (2013).

Our theoretical model predicts that more stringent LMIs attenuate both fiscal spending multipliers and output volatility. However, the latter depends crucially on the source of the shocks under consideration, and there is a trade-off with the volatility of the real wage and other variables. The strongest mitigation emanates from employment protection, followed by union density and the unemployment benefit replacement rate. Our empirical model provides evidence for this. The reduction of the size of fiscal multipliers is not limited to output, but carries over to employment and unemployment multipliers. The drop in the size of the output multiplier is up to 40 percent and depends on the particular LMI indicator considered. These differences highlight the loss in effectiveness of fiscal spending policy once stringent LMIs are in place.

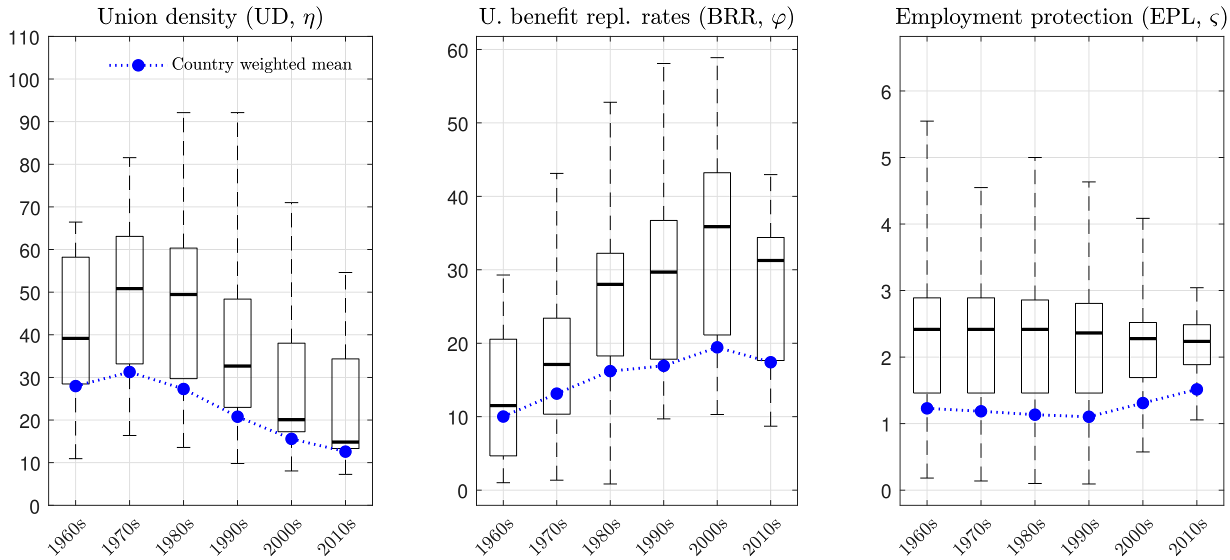
While our results highlight the limited ability of fiscal policy in attenuating cyclical fluctuations once stringent LMIs are deployed, we find that LMIs by themselves mute cyclical fluctuations. The mitigation of macroeconomic volatility (measured by the standard deviation of output) amounts to up to 25 percent regarding employment protection. The other two LMI indicators have the same qualitative effect, but to a lesser extent quantitatively. The distinct quantitative effects on cyclical volatility are due to the fact that the extent of employment protection 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.

Our contribution relates to various strands of the literature. First, we add to the literature on the effects of fiscal spending policy (e.g., Blanchard and Perotti, 2002; Čapek and Crespo Cuaresma, 2020; Corsetti *et al.*, 2012; Ilzetzki *et al.*, 2013) and specifically to the strand that evaluates their variation over time (Auerbach and Gorodnichenko, 2012; Hernández de Cos and Moral-Benito, 2016). Existing results indicate that the fiscal multiplier is higher during periods of financial turmoil (Bernardini *et al.*, 2020), when household leverage is higher (Demyanyk *et al.*, 2019), when the interest rates are at the zero lower bound (Ramey and Zubairy, 2018) or, in general, when the fiscal expansion is accommodated by monetary policy easing. Giambattista and Pennings (2017) argue that the multiplier is larger for direct transfers to financially constrained households than for government purchases and multipliers also depend on the way spending is financed (Hagedorn *et al.*, 2019). Hence, there are both cyclical and structural factors that shape the size of spending multipliers and hence the effectiveness of fiscal policy. The few contributions that assess the effects of fiscal policy on the labor market tend to ignore country-specific labor market characteristics (see for instance Monacelli *et al.*, 2010; Brückner and Pappa, 2012; Turrini, 2013). In this context, Ball *et al.* (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.

Second, our contribution is also related to the literature on the macroeconomic effects of labor market regulation. While this literature has traditionally centered on the long-run implications of market deregulation, our study is related to a more recent research program that examines their short-run effects. For instance, Zanetti (2009, 2011) or Cacciatore *et al.* (2016) assess theoretically the macroeconomic effects of market reforms featuring the removal of labor and product market frictions. An increasing number of empirical studies estimate the aggregate impact of changes in regulation (Pérez and Yao, 2015; Ordine and Rose, 2016; Duval *et al.*, 2019). The main difference between these studies and our contribution is that the former address the short-run macroeconomic effects of market reforms, while our focus rests upon the impact of labor market regulation on the fiscal multiplier. Cacciatore *et al.* (2021), Abbritti and Weber (2018) and Hantzsche *et al.* (2018) count among the few contributions that explicitly acknowledge the role played by labor market institutions in the macroeconomic responses to exogenous shocks. Cacciatore *et al.* (2021) analyze the role of employment protection for fiscal spending shocks, while Abbritti and Weber (2018) assess how the resilience of an economy with respect to external shocks is affected by its structural labor market characteristics. Hantzsche *et al.* (2018), on the other hand, analyze the interaction between labor market institutions and contractionary financing shocks.

Third, our contribution is also related to the vast empirical literature that investigates the determinants of the Great Moderation in an attempt to disentangle the relative contribution of good policies and good luck. The good luck hypothesis has been advocated by a number of authors, including Stock and Watson (2002) and Sims and Zha (2006), which conclude that the bulk of the volatility reduction in macroeconomic variables associated with the Great Moderation is due to a decrease in the variances of structural shocks. Other contributions (see, for instance, Lubik and Schorfheide, 2004; Boivin and Giannoni, 2006) find evidence for the good policy hypothesis through a shift in the systematic component of monetary policy, materialized in a change in the parameters that govern the dynamic interaction of the macroeconomic variables in a structural model rather than changes in the variances of the structural shocks. Our empirical set-up allows disentangling the relative contributions of LMIs to the parameters that govern the dynamic interaction and the contemporaneous impact of shocks. This allows for an assessment of the role of LMIs to attenuate exogenous shocks by reducing their contemporaneous impact (*good luck*) or by mitigating

Figure 1: Labor Market Institutions in OECD Economies: Time Variation.



Note: The figure shows the distribution of three labor market indicators across the sample of 16 OECD countries and their variation across countries and over time. The Greek letters attached to each LMI indicator refer to their parametric counterpart in the theoretical model. The blue line shows the weighted mean across countries for each decade.

their propagation mechanism (*good policy*). In contrast to the approaches considered in the literature, our empirical set-up not only allows uniquely to identify the source, but also to assess the sign and quantitative importance of LMIs in this respect.

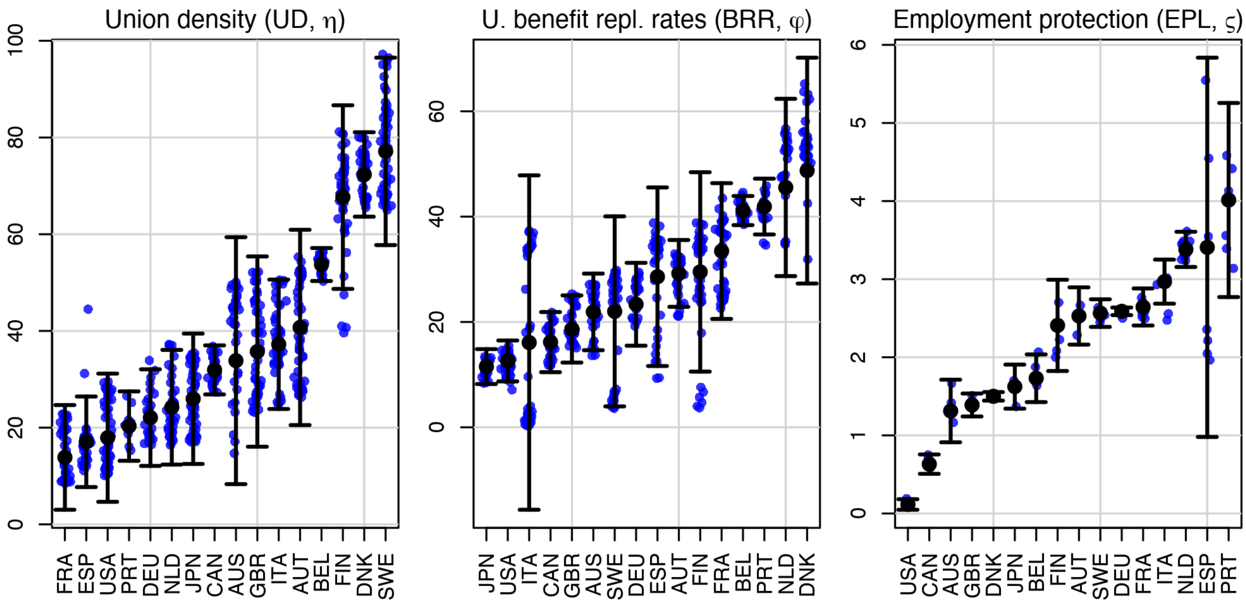
The remainder of the paper is structured as follows. In Sec. 2 we provide a descriptive overview of LMIs across selected OECD countries. Sec. 3 introduces the theoretical model and presents its main predictions. Sec. 4 discusses the connection between the theoretical and the empirical model, and presents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

2 Structural Labor Market Indicators in OECD Economies

The measurement of labor market institutions is limited by data availability, especially in the cross-country dimension. In order to balance cross-country coverage and the availability of relatively long time series information, we use OECD labor market statistics for the following three categories: (i) union density (UD), (ii) unemployment benefit replacement rates (BRR), and (iii) employment protection legislation (EPL). These three categories capture structural characteristics of the labor market across distinct dimensions. We collect these data for a total of 16 OECD countries (see Appendix C).[1]

[1] 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 *et al.*, 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 2: Labor Market Institutions in OECD Economies: Cross-Country Variation.



Notes: Each sub-plot shows the mean of each LMI variable for each country, together with two standard deviations in each direction (black). The blue points are observed data points for the respective country.

The measure of union density (UD) is based on survey data wherever possible, and administrative data adjusted for non-active and self-employed members otherwise. It is computed as the ratio of wage and salary earners that are trade union members to the total number of wage and salary earners. Higher values imply a higher extent of trade union membership and consequently a higher weight of trade unions in the context of wage negotiations. The unemployment benefit replacement rates (BRR) measure the proportion of income that is maintained after a given number of months of unemployment. The indicator is the ratio of net household income during a selected month of the unemployment spell to the net household income before the job loss. Higher values imply more generous unemployment benefit systems. These two measures, UD and BRR, affect employment, vacancies and unemployment through their effect on the price of labor. In contrast to this, employment protection legislation (EPL) is likely to affect the quantity of labor directly, changing wages only in the wake of second-round effects. The EPL index is a synthetic indicator and captures the extent of regulatory strictness on dismissals and on the use of temporary contracts. For each year, the employment protection indicator refers to the regulation in force on the 1st of January, and higher values of the indicator imply a higher extent of employment protection.

Fig. 1 shows the distributions of the three LMI indicators across countries and their variation over time. We compute 10-year averages of the three indicators for each country. The extent of time-variation in the distribution is displayed by means of boxplots for each decade starting in the 1960s. The boxplots are extended by a weighted mean (blue dotted line) where the weights correspond to the GDP share of each country (in PPP terms) relative to the aggregate of the country group as a whole.

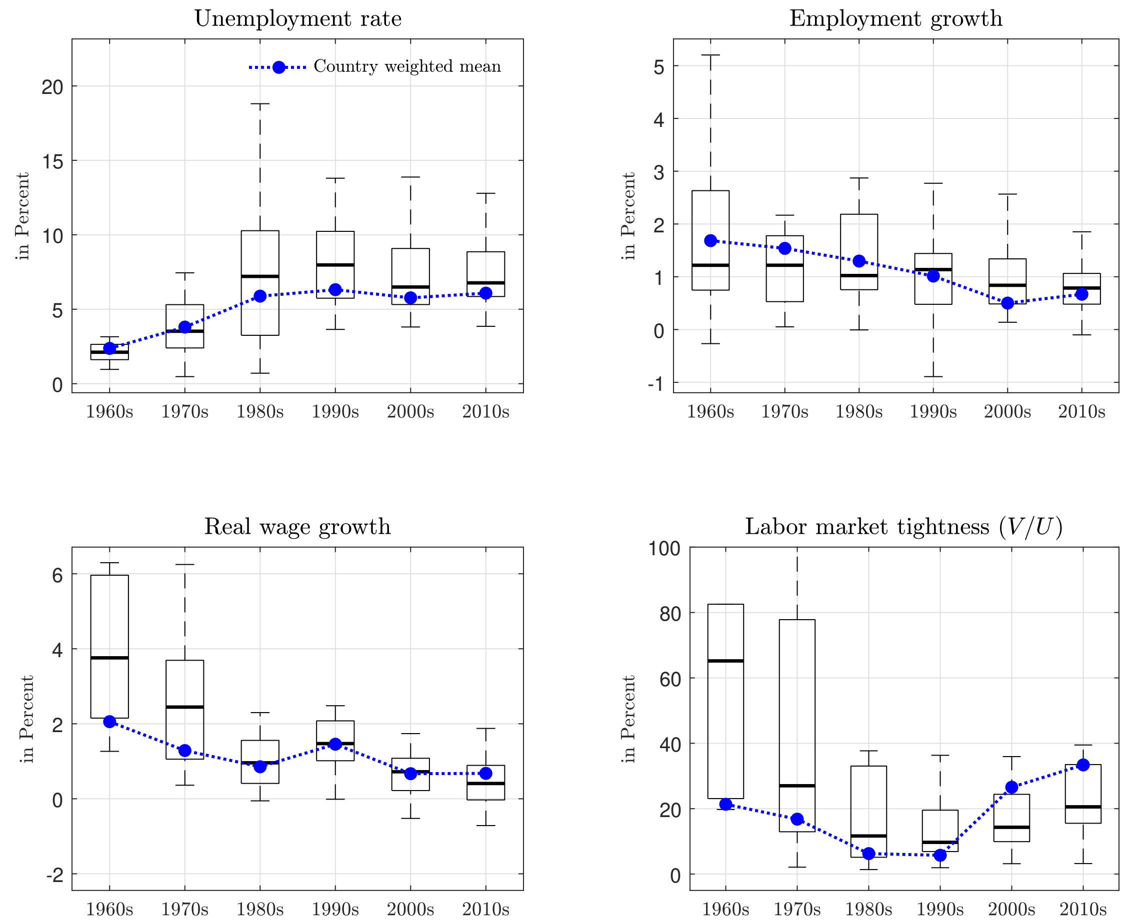
For UD, an upward shift of the distribution within the 1970s relative to the 1960s was followed by subsequent downward movements driven by a steady drop in union membership rates. This development is visible in both the median (thick black line) and the country-weighted mean (blue dotted line). The dispersion of UD also exhibits a pronounced variation over time, It was the largest in the 1980s and 1990s, and noticeably smaller before and after. The BRR experienced a steady increase up to the 2000s. Since then, the median across countries has dropped. The significant time variation in the median is accompanied by a change in the shape of the distribution as a whole. A fairly narrow distribution in the 1960s contrasts with a wide distribution in the 1990s and 2000s. The last decade was characterized by a tendency towards convergence of the unemployment benefit replacement rates across countries. The distribution of the EPL index displays an interesting pattern. While its median has remained rather constant over time, the dispersion of the distribution has narrowed significantly over time. As in the case of the BRR, the EPL index has also experienced a strong convergence across countries. This is primarily attributable to the convergent dynamics of those countries in our sample that are part of the European Union. In many other countries, the extent of employment protection has changed little over the period under scrutiny.

In Fig. 2, we examine the cross-country variation of the three indicators of LMIs. The LMI indicators exhibit strong country heterogeneity, most notably for UD and BRR and less so for EPL. In particular, Scandinavian and continental European countries present the highest levels of UD and BRR. In countries with a liberal tradition of social systems, such as the United States or Great Britain, all indicators for LMIs tend to show relatively low values.

Finally, we inspect the distribution and variation of key labor market outcome variables over time in Fig. 3. The growth rates of employment and real wages[2] were buoyant in the 1960s, followed by a subsequent moderation. The moderation in employment growth aligns with a steady rise in the unemployment rate. While the median increases continuously, the distribution of unemployment rates across countries experiences both periods of widening (1970s and 1980s) and periods of narrowing (from the 1990s onwards), highlighting significant variation over time and across countries. The fourth sub-panel shows the ratio of vacancies relative to the number of unemployed persons, that is, the slope of the Beveridge curve, which is a standard measure of tightness in the labor market. The variable presents a close co-movement with the growth rate of real wages. The number of vacancies was large in relation to the number of unemployed persons in the 1960s and partly in the 1970s, when in some countries (for instance in Germany), the number of vacancies even exceeded the number of unemployed persons. The scarcity of labor triggered a strong upward pressure on real wages. The subsequent drop in the labor market tightness (caused by both a decline in the number of vacancies and an increase in the number of unemployed persons) aligns with a significant moderation in the growth rate of real wages starting in the 1980s.

[2] The real wage rate is measured in terms of the number of employed persons, rather than in terms of hours worked.

Figure 3: Labor Market Outcomes in OECD Economies



Note: The figure shows the distribution of four key labor market variables based on 16 OECD countries and the variation over time.

The descriptive analysis highlights that increases in the BRR (which increased across most countries) and the EPL align with lower employment growth and less variation across decades. The reductions in UD, in turn, align with moderation phases in real wage growth. In what follows, we study in detail the implications of differences in these LMIs as determinants of differences in fiscal spending multipliers and macroeconomic volatility. The theoretical model built in the next section centers on the interaction between labor market institutions and outcomes when assessing the size of 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), Krause and Lubik (2007), and Monacelli *et al.* (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 Sec. A.4 of the Appendix.

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 Firms

The representative firm produces output y_t, for which it uses labor as the only input factor of production according to $y_t = \bar{A} n_t A(\tilde{a}_t)$, where $\bar{A} > 0$ is a common productivity factor and $A(\tilde{a}_t) = \int_{\tilde{a}_t}^{\infty} \frac{a}{1-F(\tilde{a}_t)} dF(a)$, where the conditional expectation is given by $E[a | a \geq \tilde{a}_t] = \int_{\tilde{a}_t}^{\infty} a f(a) da$ and $1/(1 - F(\tilde{a}_t))$ is a constant term shaping the level of $A(\tilde{a}_t)$. 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 $\pi_t^F = y_t - w_t n_t - \kappa v_t - F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b_t^s$ where the output price is normalized to unity, $w_t = \int_{\tilde{a}_t}^{\infty} \frac{\tilde{w}_t(a)}{1-F(\tilde{a}_t)} dF(a)$ is the (average) real wage weighted according to the idiosyncratic job productivity, 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[3]. The firm maximizes the present discounted value of expected profits: $\max_{n_t, v_t, \tilde{a}_t} E_t \sum_{k \geq 0} \Lambda_{t,t+k} \pi_{t+k}^F$, subject to the production function and Eq. (3.1). E_t is the expectation conditional on the information up to and including time t; $\Lambda_{t,t+k}$ denotes the firm's stochastic discount factor, defined below. The first order

[3] We consider the case where firing costs accrue to the government in Appendix A.4.4.

conditions give rise to[4]

$$F_t^n = mpl_t - w_t + E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right] \tag{3.2}$$

$$\frac{\kappa}{q_t} = E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right] \tag{3.3}$$

$$A(\tilde{a}_t) = \frac{1}{\bar{A}}\left(w_t - b_t^s - \frac{\kappa}{q_t}\right) \tag{3.4}$$

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 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, i.e., the overall household utility.

The budget constraint is given by

$$c_t + B_t = R_{t-1}B_{t-1} + (1-\tau)w_t n_t + b_t^u(1-n_t) - T_t^S + \pi_t^F, \tag{3.5}$$

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 the budget constraint, the household takes into account the flow of employment by its members according

[4] 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), Eq. (3.2), and Eq. (3.1), this equation can be further simplified to: $(1-\bar{\varrho})(1-F(\tilde{a}_t)\left(\frac{\partial A(\tilde{a}_t)}{\partial \tilde{a}_t} - \frac{\partial w_t}{\partial \tilde{a}_t}\right) = b_t^s(1-\bar{\varrho})f(\tilde{a}_t) + \left(mpl_t - w_t + \frac{\kappa}{q_t}\right)\frac{\partial \varrho(\tilde{a}_t)}{\partial \tilde{a}_t}$. 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).

to

$$n_t = (1 - \varrho(\tilde{a}_t))n_{t-1} + p_t(1 - n_{t-1}). \tag{3.6}$$

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: $\max_{c_t,n_t} E_t \sum_{k\geq 0} \beta^k u(c_{t+k}, n_{t+k})$, subject to the budget constraint, Eq. (3.5) and the employment flow Eq. (3.6). Optimization leads to the following conditions

$$\begin{aligned} 1 &= R_t E_t \Lambda_{t,t+1}, && (3.7)\\ H_t^n &= \tilde{w}_t^b - mrs_t + E_t \left[1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}\right] \Lambda_{t,t+1} H_{t+1}^n, && (3.8) \end{aligned}$$

where λ_t is the Lagrange multiplier attached to Eq. (3.5) and $\lambda_t H_t^n$ the one attached to equation Eq. (3.6). 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 equation Eq. (3.8), 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.

3.3 Nash Wage Bargaining

Wages are set each period based by 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

$$w_t = (1-\eta)\frac{mrs_t + b_t^u}{1-\tau} + \eta \left(mpl_t + E_t \Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b_{t+1}^s (1-\bar{\varrho}) F(\tilde{a}_{t+1})\right]\right). \tag{3.9}$$

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.

3.4 Fiscal Policy, Aggregate Resource Constraint, and Government Budget Constraint

The government budget constraint satisfies

$$\tau w_t n_t + B_t = R_{t-1} B_{t-1} + b_t^u u_t + T_t^s + g_t, \tag{3.10}$$

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 $b_t^u = \varphi w_{t-1}$ 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: $T_t^S = \bar{T}^S + \varphi_{T^s} B_t$; $\bar{T}^S$ and $\bar{\varsigma}$ serve the purpose to simplify the steady state computations and $\varphi_{T^s} B_t$ ensures that the necessary stability conditions are satisfied.

Finally, using Eq. (3.10), Eq. (3.5), and the expression for firms' profits (π_t^F), we obtain the aggregate resource constraint

$$y_t = c_t + g_t + \kappa v_t + F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b_t^s \tag{3.11}$$

This equation closes the model.

3.5 Embedding LMIs in the Model

The indicators for the LMIs as discussed in Section 2 are embedded in the model by the three structural parameters ς, η and φ. Considering ς first, the government's ability to shape the extent of 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.

The parameter η captures the bargaining power of workers. It is thus a measure of the implicit advantage that employees benefit from within the wage-setting process. 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). The extent of variation of this quantity across countries and across time is remarkable. Moreover, some countries also adjust the extent of unemployment remuneration in relation to the severity of crises (Ganong *et al.*, 2020).

The mapping between the empirical LMIs and their theoretical counterparts in the DSGE model can only be qualitative. In particular, when considering the case of employment protection (EPL) for instance, 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. Similar limitations arise in case of the measure for union density (UD) and the unemployment benefit replacement rates (BRR).

3.6 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-linearised solution of the rational expectations model around its steady state,

$$z_t = \boldsymbol{\Psi}_1(\boldsymbol{\vartheta}) z_{t-1} + \boldsymbol{\Psi}_0(\boldsymbol{\vartheta}) \boldsymbol{\varepsilon}_t \tag{3.12}$$

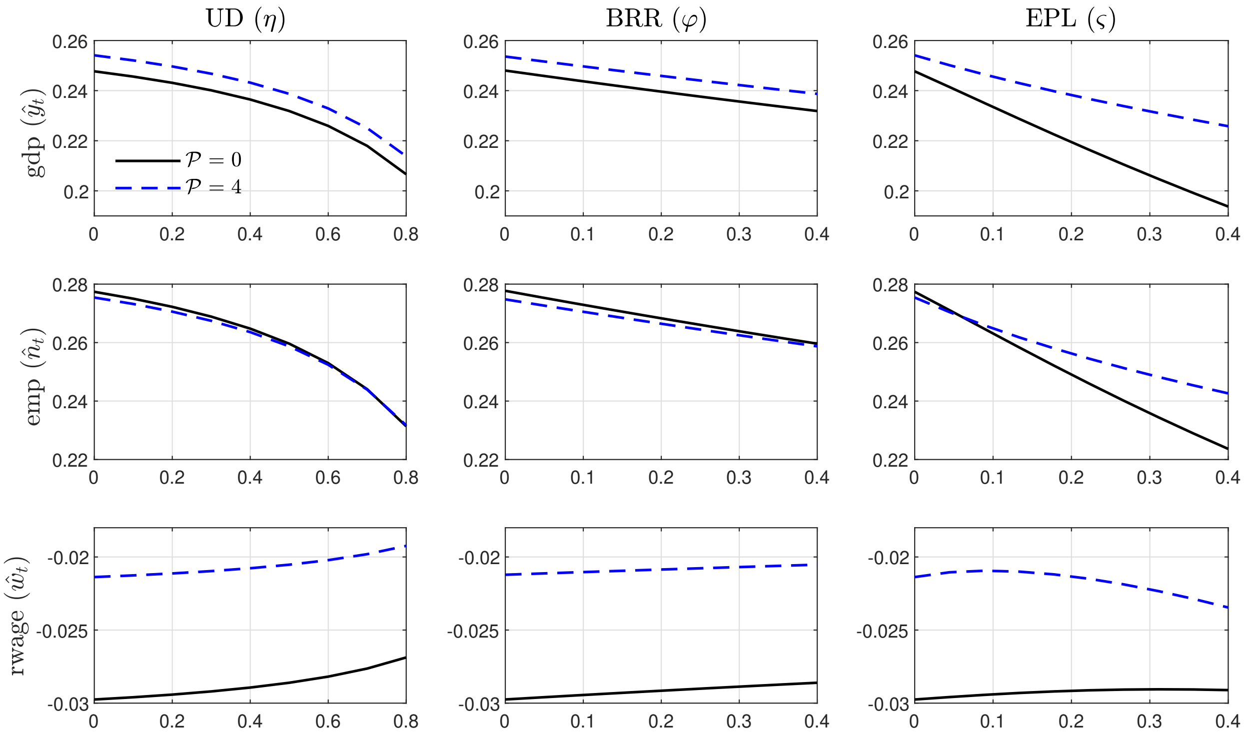
where the vector z_t contains the endogenous variables and the vector of exogenous shocks simplifies to $\boldsymbol{\varepsilon}_t = \hat{g}_t$ in our case, denoting log-deviation of our variables from the steady state with a hat. The matrix $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ governs the dynamics among the dependent variables and the vector/matrix $\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})$ determines the contemporaneous impact of the fiscal spending shock on the endogenous variables. Eq. (3.12) 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 functions of $\boldsymbol{\vartheta}$, we can display them over a whole range of values of $\boldsymbol{\vartheta}$. We do so considering the following definition of the fiscal spending multiplier for some variable x

$$\mu_x(\vartheta_l) = \frac{\sum_{i=1}^{\mathcal{P}} \mathrm{IRF}_i^x(\vartheta_l)}{\sum_{j=1}^{\mathcal{P}} \mathrm{IRF}_j^{\hat{g}}(\vartheta_l)} \quad \forall\, l = \{1, 2, 3\} \tag{3.13}$$

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 Fig. 4 for output ($\hat{y}_t$), employment ($\hat{n}_t$), and the real wage ($\hat{w}_t$). The multipliers for each variable are displayed for two distinct horizons ($\mathcal{P} = \{0, 4\}$); 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,

Figure 4: Fiscal spending multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$).



Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for different horizons: contemporaneous multiplier ($\mathcal{P} = 0$) and four quarters ($\mathcal{P} = 4$). The acronyms (UD, BRR, and EPL) refer to union density, (unemployment) benefit replacement rates and employment protection (legislation).

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. Unemployment declines in response to the rise in employment. 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.

3.7 Implications of 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 Fig. 4. 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 to decrease labor supply and demand. Using Eq. (3.2) and Eq. (3.3), and setting τ equal to zero for simplicity, the reservation wages are given by

$$\overline{w}_t^F = mpl_t + E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right], \tag{3.14}$$

$$\underline{w}_t^H = mrs_t + b_t^u - (1-\varrho(\tilde{a}_{t+1}) - p_{t+1})E_t\Lambda_{t,t+1}H_{t+1}^n, \tag{3.15}$$

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 ($\partial H_t^n/\partial\varphi < 0$). Intuitively, an increase in unemployment benefits raises workers' outside option (i.e., the present value of being unemployed) and improves their wage bargaining position. In addition, the decrease in the search intensity of workers reduces the job finding rate and the bargaining position of firms. The rise in the reservation wage of households causes the equilibrium wage (w_t) to increase, which in turn decreases the value to the firm of having an additional worker employed ($\partial F_t^n/\partial w_t < 0$). Hence higher unemployment benefit remuneration attenuates the expansion in employment[5] in response to the expansionary fiscal spending shock.

We move now to the effects of the weight of workers in the wage bargaining process (η). As highlighted above, the equilibrium wage w_t is a weighted average of the two reservations wages with the weights being determined by η. From equation (3.9), $\partial w_t/\partial\eta = \overline{w}_t^F - \underline{w}_t^H$ (for $\tau = 0$). Since $\overline{w}_t^F > \underline{w}_t^H$, as otherwise no worker-firm employment match would be created, then $\partial w_t/\partial\eta > 0$. As long as $\underline{w}_t^H < \overline{w}_t^F$, increases in the bargaining power of workers hence bring their reservation wages closer to those of firms: $\underline{w}_t^H \rightarrow \overline{w}_t^F$. This exerts upward pressure on equilibrium wages, which in turn (i) reduces the

[5] 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.

value to the firm of having an additional worker employed, and (ii) induces firms to require a higher idiosyncratic productivity of workers. The latter hence raises the endogenous job separation ($\tilde{a}_t$ rises). Consequently, both effects contract employment. Thus, in response to an expansionary fiscal spending shock, the increase in employment and output is smaller as the bargaining power of workers within wage negotiations increases. This is highlighted in the sub-panels in the first column in Fig. 4.

Finally, we turn to the effect of the extent of employment protection (ς). The key mechanism through which employment protection affects fiscal spending multipliers is related to the fact that the extent of firing costs affects the sensitivity of job destruction and job creation, as implied by equations (3.4) and (3.2). Considering the latter first, low firing costs raise the value of an additional worker to the firm ($\partial F_t^n / \partial \varsigma < 0$). In other words, low firing costs promote job creation. Hence, in response to an expansionary fiscal spending shock, low firing costs give rise to relatively higher job creation. As regards job destruction, equation (3.4) implies that low firing costs raise the job destruction rate ($\partial \tilde{a}_t / \partial \varsigma > 0$). While this effect works opposite to the job creation effect, both render employment more sensitive to aggregate shocks. Hence, in response to an expansionary fiscal spending shock, employment shows a larger reaction when firing costs are low, as can be seen in the sub-panels in the third column in Fig. 4.

The discussion so far centered on the role of the LMIs for fiscal multipliers. However, they are likely to also shape macroeconomic volatility. To this purpose, we consider Eq. (3.12) and compute the variance of the endogenous variables in z_t and extend the vector of structural shocks (ϵ_t) by additional ones to get a more reliable picture. For the sake of brevity, we describe the analysis in detail in Section A.4.6 of the Appendix and limit the discussion here to most important implications. First, as regards output, a higher value of any LMI tends to lead to lower output volatility. This applies, however, to only four out of the seven shocks considered in case of UD and EPL, while to six in case of the BRR. Second, the impact of the LMIs on output and employment volatility tends to be qualitatively the same for most shocks. Finally, there tends to be a trade-off as regards the impact of the LMIs on output and employment volatility on the one hand, and real wage volatility on the other. All these results, however, crucially depend on the source of the shock. This highlights that the LMIs can potentially mitigate output volatility, however, the nature and dominance of specific shocks is crucial.

3.8 Key Messages and Extensions of the Theoretical Model

The results of the previous exercises illustrate how LMIs affect the functioning of fiscal spending policy. On the one hand, a lack of labor market flexibility attenuates the ability of the government to provide an economic stimulus via expansionary spending policies. The reduced effectiveness of fiscal spending policies is due to a weakening of the fiscal spending multipliers by the LMIs. Stringent LMIs themselves can dampen output volatility, which, however, crucially depends on the source of the shocks.

While the theoretical results emanate from a specific model based on a particular calibration, we provide various extensions of the theoretical setting in the Appendix. Appendix A.3, for instance, reassesses the results provided in this section by considering a more general calibration of the model parameters. Sections A.4.1 – A.4.5 consider various model extensions in the form of (i) monopolistic competition and markup pricing, (ii) real wage rigidities, (iii) limited asset market participation, (iv) the case when firing costs accrue to the government as revenues, and (v) productivity-enhancing government spending. Across all extensions, the qualitative impact of the LMI parameters on the multipliers of output and employment remain identical and only the size of the multipliers are changed.

4 The Econometric Model

We empirically validate the results of our theoretical model by examining the conditional response to fiscal spending shocks for different levels of LMI indicators and their effect on macroeconomic volatility in a panel of developed countries using an interacted panel vector-autoregressive (IP-VAR) specification as popularized by Towbin and Weber (2013) and Sá *et al.* (2014). The IP-VAR 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. (3.12) 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. (3.13) 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 estimation of the econometric model.

4.1 Econometric Model

We estimate the following reduced-form IP-VAR (see Eq. (B.1) in App. B) model

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

where $\boldsymbol{y}_{it}$ denotes the M-dimensional vector of macroeconomic time series for country i and $\boldsymbol{\vartheta}_{it}$ denotes the d-dimensional interaction term, with $i = 1, \ldots, N$ denoting the country and $t = 1, \ldots, T$ the time period. Coefficients of the model are a country-specific intercept vector $\boldsymbol{c}_i$, a coefficient matrix $\boldsymbol{\Phi}_{ij}$ for lag j, and a variance-covariance matrix of the vector error term, given by $\boldsymbol{\Sigma}_i$. Note that all these reduced-form coefficients are a linear function of the interaction term. Hence, the reduced-form model is a panel VAR specification whose parameters change depending on the exact value taken by the interaction variable. The details of the model framework are presented in App. B. The structural identification of fiscal spending shocks is performed by imposing a recursive identification scheme based on the Cholesky decomposition of the variance-covariance matrix $\boldsymbol{\Sigma}_i$ of the reduced-form IP-VAR shocks. We discuss shock identification in more detail in the next subsection.

The structural IP-VAR representation of the DSGE model comprised by Eq. (3.12) is given by

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

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) and explained in detail in App. B. The prior assumes that the structural individual-level coefficients have a common underlying Gaussian distribution,

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

with a variance-covariance matrix $\boldsymbol{V}_j$. We exert regularization via this variance-covariance matrix towards the common mean model with the help of Bayesian global-local shrinkage priors (Griffin and Brown, 2010; Huber and Feldkircher, 2019). The exact specification can be found in App. B. The correspondence between the observable LMIs $\boldsymbol{\vartheta}_{it}$ (depicted in Fig. 1) and the structural LMI parameters $\boldsymbol{\vartheta}$ of the DSGE model can be made explicit by defining $\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}) = \boldsymbol{\Psi}_{jt}$, $\boldsymbol{\Psi}_j(\bar{\boldsymbol{\vartheta}}) = \bar{\boldsymbol{\Psi}}_j - \partial\boldsymbol{\Psi}_j(\boldsymbol{\vartheta})/\partial\boldsymbol{\vartheta} \cdot \bar{\boldsymbol{\vartheta}}$, $\boldsymbol{\Gamma}_j^{\boldsymbol{\Psi}} = \partial\boldsymbol{\Psi}_j(\boldsymbol{\vartheta})/\partial\boldsymbol{\vartheta}$ for $j = 0, 1, \ldots, p$. This implies that the coefficients of the $\boldsymbol{\Psi}_{jt}$ matrix vary as follows

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

which relates the empirical set-up directly to Eq. (4.1) of the theoretical model. The full IP-VAR model is given by Eq. (4.3), and its equivalence with the solution of the DSGE model depicted in Eq. (3.12) and Eq. (4.1) is evident when considering a lag length of one.

In the IP-VAR 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 for varying levels of one interactive variable, while keeping the remaining ones at a given level.

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

Fig. 1, structural rather than cyclical factors appear to determine their dynamics.[6] A second potential limitation is the linearity assumption (in the parameters) embedded in the IP-VAR 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 $\boldsymbol{\vartheta}$. However, depending on the number of observations and parameters of interest in 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 Shock Identification

We identify fiscal spending shocks by imposing a recursive identification based on the Cholesky decomposition of the reduced-form IP-VAR shocks. We follow Blanchard and Perotti (2002) and assume that fiscal spending does not react contemporaneously to shocks arising from GDP or labor market variables in the system. These three variables are 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 analyse the effects of fiscal policy (Beetsma and Giuliodori, 2011; Bénétrix and Lane, 2013; Ilzetzki *et al.*, 2013; Huidrom *et al.*, 2020, to mention a few). Second, alternative identification approaches are infeasible in the context of our research question. In particular, event-study approaches based on defence spending changes (Ramey and Shapiro, 1998; Ramey, 2011) are not really suitable in our context, as defense spending is negligibly small in most of the countries in our data set. The approach by Blanchard and Perotti (2002) would additionally require institutional information on the elasticity of government spending and revenues to output and inflation, which is not practical for a large panel of countries. Mountford and Uhlig (2009) use sign-restrictions to identify fiscal policy shocks which is not practical in a dataset with a large panel of countries either. Finally, the narrative approach (Romer and Romer, 2010; Guajardo *et al.*, 2014) requires the availability of detailed legislative records in order to extract policy shocks.[7] Such approaches would require collecting detailed institutional information and data on fiscal spending plans for 16 countries for a sufficient long time horizon, thus rendering these approaches infeasible for our purposes.

One potential drawback in the context of a recursive identification scheme concerns the extent of unpredictability of changes in government spending from the point of view of a statistician. This might stand in contrast to economic agents, who might well have anticipated at least parts of the fiscal shock.

[6] 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 signifcantly react to cyclical shocks.

[7] Kraay (2012) describes yet another approach whose applications are essentially limited to developing countries as it relies on two features which are unique to low-income countries: (1) borrowing from the World Bank and (2) spending on World Bank–financed projects.

Legal processes usually create a time gap between the announcement and the implementation of a given fiscal policy measure. A statistician, relying on the data a posteriori, would attach the implementation as starting point of the policy, while economic agents might have already reacted to the mere announcement of the policy change. Ignoring this aspect results in a potential underestimation of the effects of fiscal spending shocks. The role played by such anticipation effects is ultimately an empirical question. It relates to the extent of liquidity constrained households and the share of consumption in GDP. Mertens and Ravn (2010) highlight that a simple Cholesky decomposition delivers practically correct impulse responses for a large class of theoretical models even if shocks are anticipated by the private sector.

4.3 Data and Specification

We use quarterly data ranging from 1960:Q1 to 2020:Q4 for 16 OECD countries to estimate the IP-VAR model. The sample includes information for Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Italy, Japan, the Netherlands, Portugal, Spain, Sweden, and the United States. We specify $\boldsymbol{y}_{it} = \{\text{govc}_{it}, \text{gdp}_{it}, \text{emp}_{it}, \text{rwage}_{it}\}$ for our baseline specification, where we follow Brückner and Pappa (2012) and express all variables in per-capita terms. The variable govc_{it} denotes the growth rate of real government consumption per capita, gdp_{it} is the growth rate of real GDP per capita, emp_{it} is the growth rate of employment per capita, and rwage_{it} is the growth rate of the real wage (see Tab. C1 and Tab. C2 in the Appendix for further details). We consider various extensions to the baseline specification in which we substitute employment (emp_{it}) by (i) the growth rate of unemployed per capita (unemp_{it}) and (ii) the labor market tightness indicator given by the ratio of vacancies to unemployment (vu_{it})[8]; see App. D.

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 (see Tab. C1 in the Appendix for further details) for union density (η_{it}), unemployment benefit replacement rates (φ_{it}) and employment protection legislation (ς_{it}). 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 (see also Abbritti and Weber, 2018). We estimate the IP-VAR model using all three interaction variables at once. Additionally, we standardize each interaction variable prior to estimation. This serves the purpose of comparability across countries, otherwise the proposed common-mean prior specification runs into (numerical) troubles. We abstain from the alternative – taking differences of the interaction variables – to have a more direct interpretation with respect to the respective country levels. Either way, this estimation strategy only utilizes the within-country variation of the LMIs. Hence, our estimates are of a more conservative nature due to the strong cross-country heterogeneity in the LMIs (see Fig. 2). We investigate cross-country heterogeneity by splitting the sample of countries in two distinct groups in Sec. 4.6. Given the standardization of the interaction variables, the interpretation is as follows: A unit rise in ϑ_l corresponds to a one standard

[8] In the IP-VAR model in which the labor market tightness indicator is used instead of employment, we have to reduce the country coverage of our sample to $N = 13$, as data for vacancies are unavailable for Belgium, Canada, and Italy.

deviation increase of the respective LMI within countries. When simulating the IP-VAR model along a particular interaction variable ϑ_l, we set the remaining ones ϑ_ℓ ($\ell \neq l$) equal to zero (the mean).

The baseline model (and so too the remaining two) is estimated with one lag ($p = 1$) as proposed by the Bayesian information criterion. The estimation is based on 20,000 posterior draws, where we discard the first 10,000 as burn-ins.

4.4 The Effect on Fiscal Spending Effectiveness

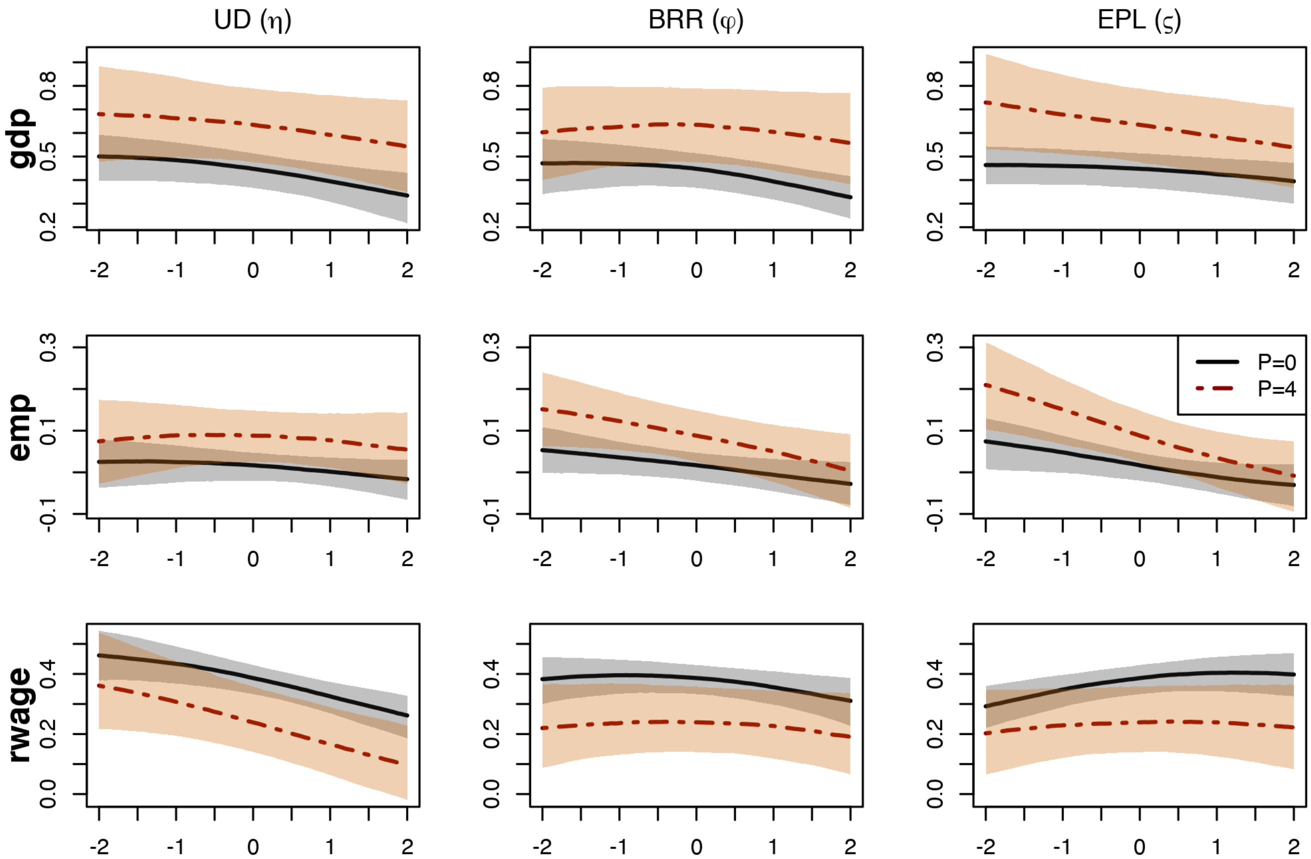
In this section, we present the effects of the LMIs on fiscal spending effectiveness. In line with the theoretical results, we use Eq. (3.13) to compute multipliers for the initial impact ("impact multiplier", horizon $P = 0$) and for the effect after four quarters ("one-year multiplier", horizon $P = 4$). Fig. 5 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 multipliers and are complemented with the 68% confidence bound in each case. The horizontal axis ranges from -2 to +2 standard deviations of the respective LMIs while the vertical axis depicts the value of the fiscal multiplier for the respective variable.

Considering for instance the impact multiplier for output and its dependency on union density (first panel), we find that at a low value of UD (η) a one percent increase in fiscal spending raises output by 0.5 percent; the value of the output multiplier, however, drops to around 0.3 when the UD is at a high value. This gives rise to a decline in the output multiplier of up to 40% which is substantial and statistically significantly different from zero. A drop of a similar size also applies in case of the BRR (φ), while in case of the EPL (ς) the decline of the output multiplier is weaker (around 20%; from 0.5 to 0.4).

The impact multipliers for employment, while consistently positive, are also negatively affected by the LMIs. The drop is sizeable in case of the EPL and amounts to around 55% (from 0.9 down to 0.4), while moderate for the remaining two LMIs. For both output and employment, the one-year multipliers consistently exceed the impact multipliers which highlights the inertia of the impact of government spending shocks on economic activity. Moreover, the one-year multiplier for output displays a lower sensitivity to the BRR than the impact multiplier. The opposite applies to the EPL – the drop is now close to 30%. In case of the UD the sensitivity does not change across impact and one-year multiplier.

The noticeably higher value of the output multiplier of the IP-VAR model compared to the DSGE model stresses the role of nominal rigidities (see Section A.4.1 in the Appendix); with a view to the employment multiplier, the fact that the one-year multiplier consistently exceeds the impact multiplier highlights the role of limited asset market participation in this context (see Section A.4.3 in the Appendix). 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 A.4.4 in the 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.

Figure 5: Fiscal Multipliers.

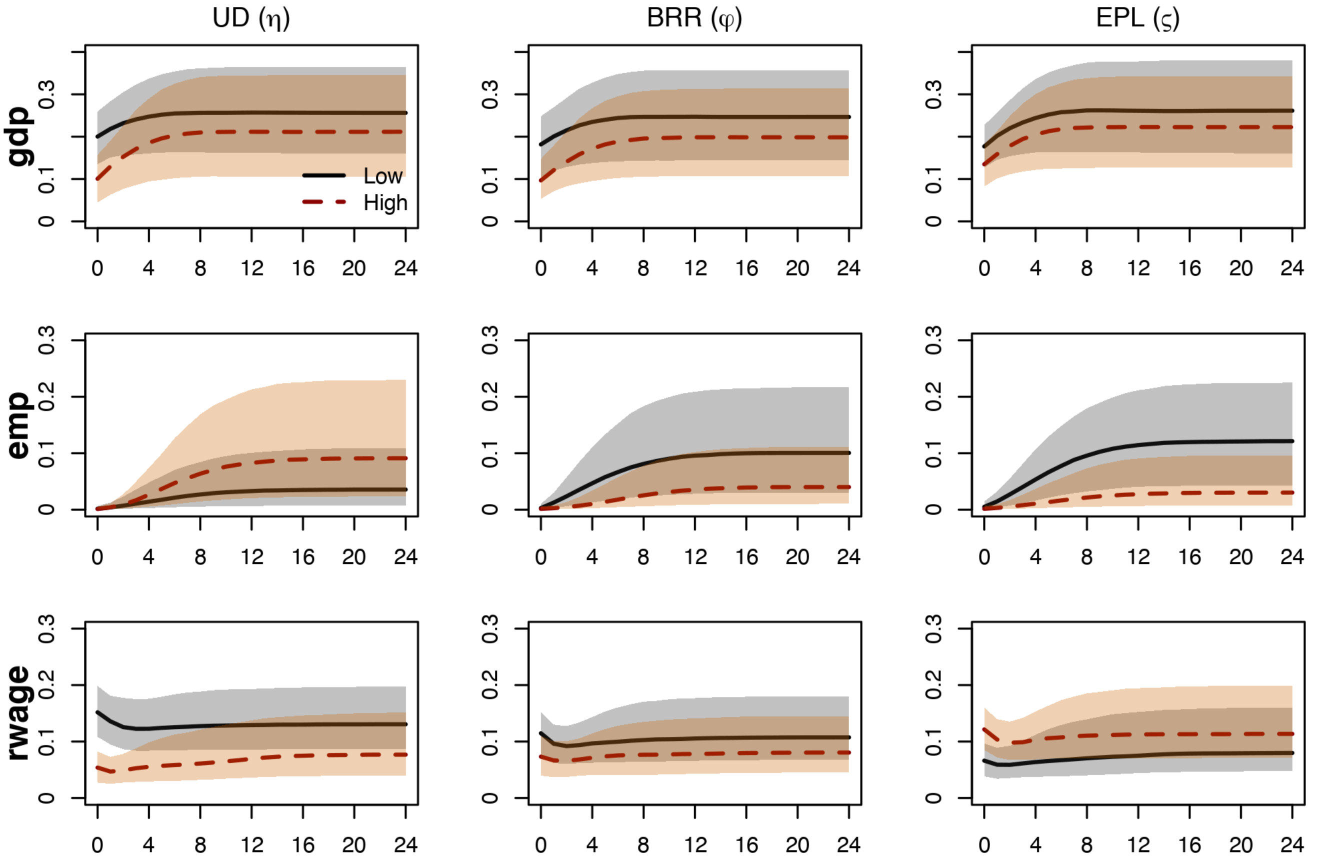


Notes: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is unemployment benefit replacement rate, and ς is employment protection). The y-axis gives the size of the multiplier while the x-axis runs from -/+ 2 standard deviations in terms of the respective LMI. The multipliers are shown for different horizons: contemporaneous multiplier ($\mathcal{P}$ = 0, solid line) and four quarters ($\mathcal{P}$ = 4, dash-dotted line). Confidence bounds refer to the 16/84 quantile of the posterior distribution.

The most noteworthy deviations from the (baseline) theoretical predictions apply to the real wage. The IP-VAR model gives rise to a positive real wage multiplier; moreover, the multiplier abates with the horizon quickly. At the same time the real wage multiplier increases with the EPL, while the opposite applies for the UD and to a lesser extent for the BRR. This highlights the presence of nominal ridigities and productivity enhancing government spending (see Section A.4.5 in the Appendix) for explaining the positive value of the real wage multiplier – both frictions give rise to a theoretic real wage multiplier replicating the course of its empirical counterpart.

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 *et al.*, 2013), as well as the lower value of the employment relative to the output multiplier (Monacelli *et al.*, 2010). Not least the positive real wage multiplier aligns with the findings in Brückner and Pappa (2012) who identify both negative and positive multipliers across distinct countries. Our key contribution in this context concerns the assessment of the size and shape of

Figure 6: Forecast Error Variance Decomposition.



Notes: The sub-plots 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 6 years (=24 months). The FEVD is shown for a regime with low (−2sd) and high (+2sd) LMIs.

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 App. D. There, we also show the impulse response functions and additionally provide results for the alternative two models featuring unemployment and the vacancy-to-unemployment ratio (labor market tightness, v_t/u_t) instead of employment. These alternative models confirm the size of the output multipliers and their dependency on the LMIs of the baseline model. The same applies to the real wage multipliers. Moreover, the additional models highlight the negative (positive), though, sluggish effects on unemployment (labor market tightness), both of which are in line with the theoretical model.

So far our results highlight the role of the LMIs for the transmission of discretionary fiscal policy shocks. While this finding is interesting for itself, it though does not tell anything about the extent of the variation in output that can be affected by fiscal policy, and hence on the fiscal authority's quantitative ability in attenuating output fluctuations. To this end, we extend the previous analysis to include the forecast error variance decomposition (FEVD). The results are provided in Fig. 6.

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 stringent LMIs are deployed ("high"). In contrast, they explain up to 28% at horizons of eight quarters and beyond when the LIMs are, however, less stringent ("low"). This share is substantially lower for the variables characterizing the labor market, employment and real wages. In particular, while a higher level of the LMIs reduce the explained forecast error variance of output only slightly, the attenuation is sizable for employment in case of the BRR (φ) and the EPL (ς), and for the real wage in case of the UD (η). From an economic point of view, more stringent LMIs abate the amount of variation explained in labor market variables by fiscal spending shocks. Put differently, when stringent LMIs are deployed, discretionary fiscal policy only has a limited potential in affecting labor market outcomes.

As a robustness check, we re-do the analysis with other labor market variables: unemployment and our measure for the labor market tightness (v_t/u_t). The results are provided in Fig. D5 in the Appendix. In both models, we observe no stark differences to the baseline results. Furthermore, the reduction in the explained forecast error variance is even stronger to some extent.

Overall, the LMIs are found to play a role for the effectiveness of discretionary fiscal policy, though most of the results are borderline significant only. This applies to both the goods and the labor market. Among the three LMIs considered, the UD is found to have the strongest effect on real wages, while the EPL on employment. As the BRR is targeting both, quantity and prices, it hence shapes both employment and real wages as our results highlight. Most importantly, more stringent LMIs limit the fiscal authority's ability in affecting cyclical swings in economic activity. This, however, does not indicate anything as to whether discretionary policy measures attenuate or reinforce cyclical fluctuations. In the end, this crucially depends on the timing of fiscal interventions (if timed adequately, a counter-cyclical policy stance emerges which smooths cyclical fluctuations) and the size of the interventions (if sized too big, an overshooting might occur which by itself exacerbates cyclical fluctuations). While these factors shape the success of fiscal policy, the LMIs counteract its effectiveness. This raises the question of whether the LMIs themselves contribute to mitigating cyclical fluctuations which is what we focus on in the following section.

4.5 The Effect on Macroeconomic Volatility

The analysis so far has shown that the effects of discretionary fiscal policy potentially decrease the more stringent the LMIs are. However, this raises a fundamental question: Is there a need for discretionary fiscal policy in an environment of stringent LMIs? After all, the main objective of discretionary (spending) policies is to stimulate aggregate demand in the event of a negative demand shock or to tighten in the opposite case. In other words, aggregate demand is smoothed over the business cycle. However, it may well be that in an environment with already stringent LMIs, these very elements already contribute significantly to cyclical smoothing by which they render any discretionary spending policy obsolete.

Hence, we want to assess whether there is a degree of substitutability between the LMIs and cyclical spending policies.

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 hence over households' employment/unemployment states which in turn stabilizes consumption at the individual and aggregate level. The EPL and the UD work through similar channels. Hence, the LMIs possibly have the potential to prevent an economy from recessionary tendencies being self-reinforcing to some extent. We expect output volatility to be lower in an economy with a more rigid labor market. Our econometric setting allows for an assessment of macroeconomic volatilities with respect to the LMIs. With this in mind, we analyze the impact of the LMIs on macroeconomic volatility. To this end, we determine the variance of the endogenous variables of the IP-VAR model[9] and examine the impact of the LMIs. The variance covariance matrix of the endogenous variables $\mathbf{y}_{it}$ in the IP-VAR system is given by

$$\operatorname{vec}(\boldsymbol{\Omega}(\boldsymbol{\vartheta})) = (\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}) \otimes \boldsymbol{F}(\boldsymbol{\vartheta}))^{-1} \operatorname{vec}(\boldsymbol{Q}(\boldsymbol{\vartheta})) \tag{4.6}$$

where $\boldsymbol{I}$ is an identity matrix of dimension $K^2 = (Mp)^2$, $\boldsymbol{F}(\boldsymbol{\vartheta})$ denotes the $K \times K$ companion matrix form of $\boldsymbol{\Psi}_j(\boldsymbol{\vartheta})$ with $j = 1, \ldots, p$, and $\boldsymbol{Q}(\boldsymbol{\vartheta})$ denotes the $K \times K$ companion matrix form of the common-mean variance-covariance matrix $\overline{\boldsymbol{\Sigma}}(\boldsymbol{\vartheta}) = N^{-1} \sum_{i=1}^{N} \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta})$.[10] As can be seen from Eq. (4.6), the variance covariance matrix of the endogenous variables thus depends on the interaction term $\boldsymbol{\vartheta}$, which are the LMIs in our setting. It follows that $\boldsymbol{\Omega}(\boldsymbol{\vartheta})$ is the $K \times K$ variance covariance matrix of the VAR system in stacked form. The results are depicted in Fig. 7. We focus on output and employment only and provide further details in the Appendix. In Fig. 7, we measure volatility with the model-implied standard deviations of output and employment from the baseline model and compare the volatilities for high (+2sd) and low (-2sd) values of the respective LMIs.[11] We observe that the LMIs have a potentially volatility-reducing effect. For instance, a high UD attenuates output volatility by around 10% with a probability of 67%, while the effect on employment is smaller in size (reduction of around 3% with a probability of 56%). In case of the BRR, the effects are rather muted for both variables. While the output volatility tends to be negatively affected by a higher BRR (the probability that the volatility declines is 56%), the opposite applies to the employment volatility – the probability that the employment volatility declines with a higher BRR is only 35%. The largest effects emanate from the EPL. A more stringent EPL gives rise to a drop

[9] The results of the DSGE model of this exercise are presented in Section A.4.6 in the Appendix.

[10] 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 IP-VAR.

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

Figure 7: Macroeconomic Volatilities Along LMIs.

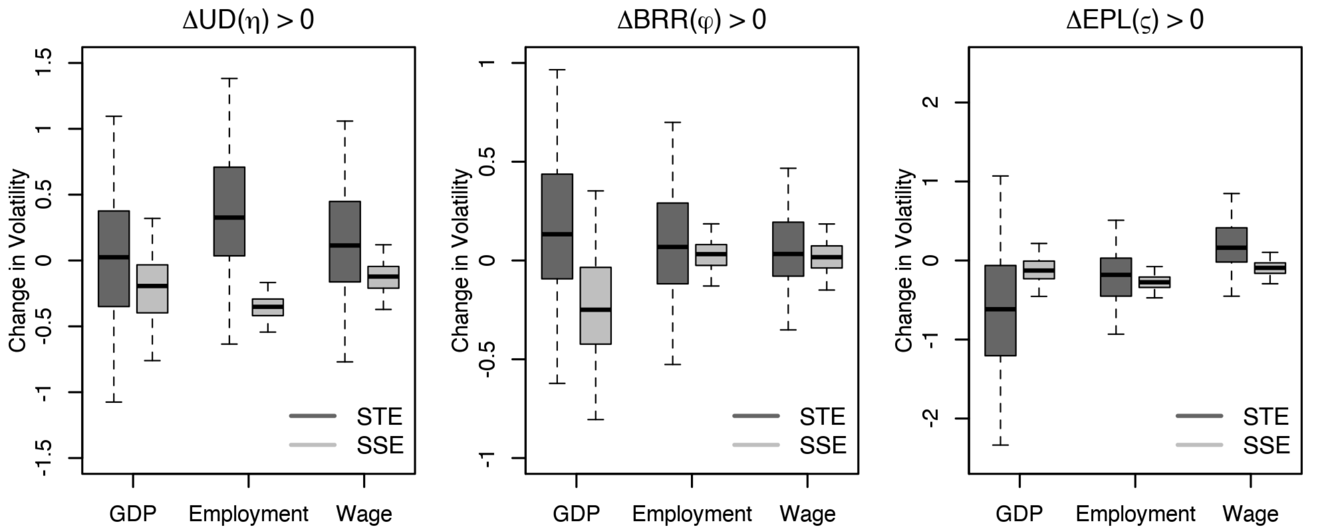
Notes: Each sub-plot shows the standard deviations of the respective macroeconomic variable in a regime with low (−2sd) and high (+2sd) LMIs. The LMIs under consideration are union density (UD(η)), unemployment benefit replacement rate (BRR(φ)), and employment protection (EPL(ς)).

in the volatilities of output and employment of around 25% (with a probability of 85%) and 30% (with a probability of 93%), respectively. These results align with the theoretical predictions put forth in Section A.4.6 in the Appendix. Most importantly, the theoretical results in this context highlight that while the ability of the LMIs in abating macroeconomic volatility crucially depends on the shocks' sources, the EPL still tends to mitigate output and employment volatility to the largest extent, compared to the UD and BRR.

How do the LMIs shape the effects of shocks? The ability of the LMIs to dampen macroeconomic volatility thus classifies them as important elements for the purpose of smoothing cyclical fluctuations. At the same time, however, this raises a crucial question: How do 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 $\vartheta_l \in \boldsymbol{\vartheta}$ on ω_{kk}, which is the k,k-th element in the matrix $\boldsymbol{\Omega}(\boldsymbol{\vartheta})$ and denotes the volatility of the kth variable in the vector of endogenous variables $\boldsymbol{y}_{it}$. Furthermore, we denote with $\tilde{\boldsymbol{F}}(\boldsymbol{\vartheta}) \equiv (\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}) \otimes \boldsymbol{F}(\boldsymbol{\vartheta}))^{-1}$ and $\tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}) \equiv \mathrm{vec}\,(\boldsymbol{Q}(\boldsymbol{\vartheta}))$, where small letters denote scalars and refer to an element of the corresponding vector $\tilde{q}_j(\boldsymbol{\vartheta}) \in \tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta})$ and

Figure 8: Change in Macroeconomic Volatilities Along LMIs.



Notes: Each sub-plot shows the change in the standard deviations of the respective macroeconomic variable when going from a regime with high (+2sd) to low (−2sd) LMIs. 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 (EPL(ς)).

matrix $\tilde{f}_{kj}(\boldsymbol{\vartheta}) \in \tilde{\boldsymbol{F}}(\boldsymbol{\vartheta})$. Then $\omega_{kk}(\boldsymbol{\vartheta}) = \sum_j \tilde{f}_{kj}(\boldsymbol{\vartheta}) \cdot \tilde{q}_j(\boldsymbol{\vartheta})$ holds, and the partial effect is given by

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

which allows us to decompose the overall change in the volatility with respect to the LMIs along the proposed dimensions: STE and SSE. In other words, we are now able to give an answer whether the LMIs affect the size of the contemporaneous impact or the transmission channel of exogenous shocks. 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 Fig. 8, where we show the change in the volatility of output and employment ($\Delta\omega_{kk}(\boldsymbol{\vartheta})$) when moving from loose to stringent LMIs ($\Delta\vartheta_l > 0$). The overall effect ($\Delta\omega_{kk}(\boldsymbol{\vartheta})/\Delta\vartheta_l$) is decomposed into the STE (dark box-plots in each panel) and SSE (bright box-plots in each panel). For a better 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 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. The effects are sizeable, however, since they drift in opposite directions, the overall effect as depicted in Fig. 7 is hence negligibly small.

In case of the BRR, the results for the output volatility are similar as for the UD – the BRR exacerbates the propagation mechanism of shocks while at the same time it attenuates the output volatility due to a smaller contemporaneous impact of shocks. With a view to employment, both effects, the STE and the SSE, promote volatility which explains the increase in employment volatility with a higher BRR as shown in Fig. 7. The most uniform effects emanate from the EPL. Both the STE and the SSE induce a mitigation of the volatility. For output the shock transmission effect dominates while for employment the shock size effect (SSE).

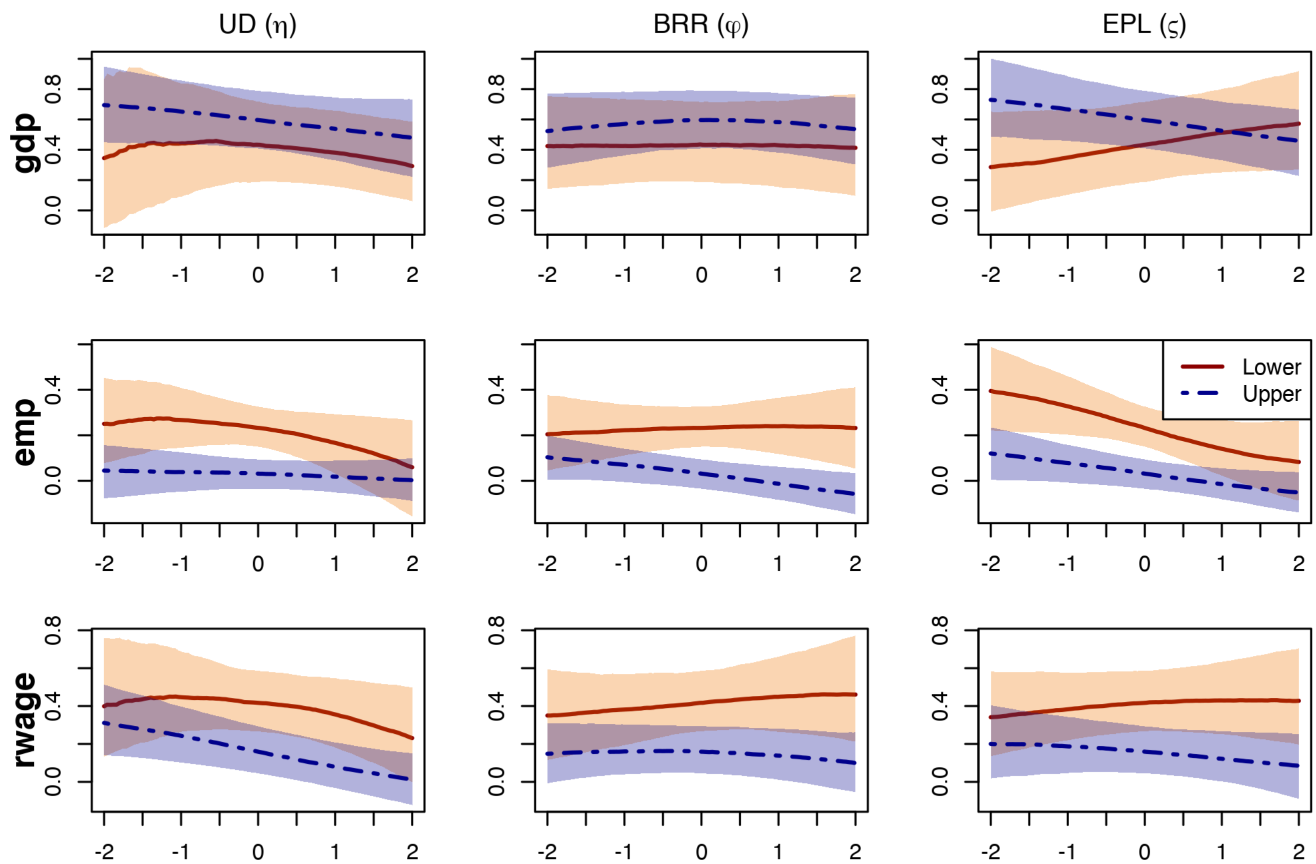
As a robustness check, we again examine the effects on the change in the volatility when variables other than employment are used in the IP-VAR model, for instance, unemployment and the measure for the labor market tightness (v_t/u_t). The results are presented in Fig. D6 and Fig. D7 in the Appendix. We observe that the previous findings also apply in the two alternative specifications. The reduction of the volatility is again strongest with the EPL for output, and, across all specifications, a higher BRR tends to raise volatility in the labor market. Regarding the dimensions that matter in this respect, we find that the BRR consistently reinforces the propagation mechanism of exogenous shocks while at the same time it attenuates output volatility due to a smaller contemporaneous impact of shocks. In case of the EPL, both dimensions (STE and SSE) apply and work in the same direction.

Overall, 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. In all cases, both STE and SSE play a significant role. However, a crucial detail in this respect is the fact that the direction of the effect is in some cases opposite (for instance the output volatility in case of the BRR) which gives rise to an overall small effect.

4.6 Inspecting the Between-Country Variation

So far, the analysis is based on within-country variation of the LMIs. In this section, we want to explore possible effects of between-country variation. As we have noticed in Fig. 2, there is considerable cross-country heterogeneity in the respective indicators. It is thus an interesting exercise to check whether effects differ in countries with more stringent LMIs deployed (e.g., in Scandinavia) to countries with more flexible labor markets (e.g., Anglo-Saxon countries). We proceed as follows. All countries of the baseline model are clustered into two groups based on their LMIs. 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 IP-VAR. The clustering is done via k-means clustering. 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). 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

Figure 9: Fiscal Multipliers Across Groups.



Notes: The sub-plots show the sensitivity of the fiscal spending multipliers 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 size of the multiplier while the x-axis runs from -/+ 2 standard deviations in terms of the respective LMI (within-country variation). The multipliers are shown for a horizon of four quarters and for the *Lower* and *Upper* group (between-country variation) Confidence bounds refer to the 16/84 quantile of the posterior distribution.

label as follows. "Upper" group: Austria, Belgium, Germany*, Denmark, Spain, Finland*, France, the Netherlands, Portugal, and Sweden*. "Lower" group: Australia, Canada, Great Britain, Italy*, Japan, and the United States.[12] The groups align well with various definitions of welfare regimes and are depicted in Fig. D8 in the Appendix.

In Fig. 9, we examine the fiscal multipliers both on a within- and between-country variation basis. In line with the theoretical results and the analysis conducted in Sec. 4.4, we use Eq. (3.13) to compute multipliers for the effects after four quarters ("one-year multiplier", horizon $P = 4$) and compare both groups. The results already discussed are robust to the sample split. But if both groups are homogeneous, then fiscal multipliers would overlap. While this roughly holds for the fiscal multiplier for output, where no significant differences accrue, we observe strong differences for the labor market variables. In particular, the employment fiscal multiplier is almost zero in any case and only slightly downward shaped for the "upper" group. In the "lower" group, however, we observe a positive employment multiplier throughout;

[12] A star indicates that the 50% rule applies.

most importantly, it depends negatively on the LMIs, that is, more stringent LMIs attenuate employment multipliers. A similar pattern arises for the real wage rate. In Fig. D9 we observe qualitatively similar results from the two alternative models which feature unemployment and the labor market tightness instead of employment.

Overall, the results outlined here strengthen the implications of our baseline results as of Section 4.4. 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.

4.7 Discussion

So far our results have highlighted the important role LMIs have in (i) determining the effectiveness of discretionary fiscal spending and (ii) for influencing macroeconomic volatility directly through distinct channels. In what follows we discuss each aspect from a more general point of view.

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. The finding that more stringent LMIs render discretionary fiscal policy less effective cannot be classified as an unpleasant consequence of a less flexible labor market per se. While this is clearly bad news for a government which – when confronted with a negative private demand shock – intents to stabilize total demand by raising public demand, the same, however, is good news for a government ahead of reducing the public budget deficit and debt when trying to bring public finances in order. Hence, the subjective perception of this result – the negative effects of the LMIs on the effectiveness of discretionary fiscal policy – depends on the specific purpose of fiscal spending as a macroeconomic policy instrument.

While more stringent LMIs limit a fiscal authority's ability in affecting output fluctuations, our second key result highlights the extent to which the LMIs shape output volatility directly. There is a vast literature that tries to identify the factors that cause volatility changes in macroeconomic time series and to assess the consequences of that for policymakers. For instance, in case of a volatility attenuation, one of the key questions concerns whether the increased stability can be associated with an alteration of the transmission mechanism of exogenous shocks. The answer depends on the origin of the volatility changes: whether the reduced volatility is due to smaller or less frequent macroeconomic shocks, or whether the propagation of these shocks has changed. In this context, the literature on the Great Moderation has focused on the role of good policy and good luck. The inconclusive results of this literature have raised several points of critique. First, the good luck interpretation of existing VAR evidence suffers from the logical drawback that it implies a simultaneous reduction in the volatilities of conceptually *orthogonal* shocks. Second, the good luck and good policy interpretations are close to observationally equivalent (Benati and Surico, 2009).

Our approach enables a more accurate assessment of distinct explanations in this respect. First of all, our approach permits to consider distinct observable structural elements as explanations for volatility changes in macroeconomic time series. Secondly, our approach permits 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).

From the perspective of a policy maker, the second point (STE versus SSE) is particularly relevant. For instance, if the main cause of the decreased economic volatility is a reduction in the size of the contemporaneous impact of shocks (i.e., "good luck", SSE), then a re-emergence of large successive shocks would eventually lead the economy to becoming more volatile again. Alternatively, if the reduced volatility is due to a change in the propagation of shocks (i.e., "good policy", STE), then it is reasonable to expect that the low-volatility regime will continue.

In this respect, our results highlight the role of the EPL in attenuating the propagation mechanism of exogenous shocks and the opposite applies for the UD and the BRR. While the latter mitigate the contemporaneous impact of shocks they, however, reinforce the propagation mechanism of shocks. This in turn casts doubts on the ability of the UD and the BRR for attenuating macroeconomic volatility in general. Moreover, by being able to moderate the contemporaneous impact of shocks, they create the conditions for a seemingly tranquil macroeconomic surface to emerge, beneath which, however, the buildup of imbalances and macroeconomic risks can quickly be overlooked. The pitfall in this respect is to assign a seemingly high shock absorption capacity, to an economy that is merely in a temporary phase of moderate cycles. This applies to all structural elements that shape macroeconomic volatility along the SSE rather than the STE.

5 Concluding Remarks

We have shown, both theoretically and empirically, the eminent role of labor market institutions for fiscal policy and macroeconomic volatility alike.

A detailed descriptive overview of key indicators of labor market institutions for which we focus on (i) union density, (ii) unemployment benefit replacement rates, and (iii) employment protection legislation, highlights their enormous variation across time and countries. This fact raises the question of the extent to which the LMIs trigger effects beyond the labor market. In this context, our analysis identifies two key findings.

First, we find that the labor market institutions affect the propagation mechanism of discretionary fiscal spending. To the extent that more stringent labor market institutions decrease fiscal spending multipliers, they hence mitigate a government's ability to use fiscal policy as a macroeconomic stabilization tool. These effects turn out strongest in case of the employment protection legislation while weaker with respect to the union density and the unemployment benefit replacement rates.

Second, we find that the labor market institutions by themselves mute output volatility. The mitigation of cyclical fluctuations – measured by the standard deviation of output – amounts to up to 25 percent in the case employment protection. The other two labor market institutions have the same qualitative effect, but to a significantly lesser extent quantitatively. The reason for the distinct quantitative effects on the volatility is because the employment protection legislation attenuates output volatility by mitigating both the propagation mechanism and the size of the contemporaneous impact of shocks. The union density and unemployment benefit replacement rates, however, exacerbate the propagation mechanism of exogenous shocks while moderating their contemporaneous impact.

These results emerge, on the one hand, from a theoretical model which combines the characteristics of a Diamond-Mortensen-Pissarides model with a standard real business cycle set up; and, on the other hand, from an interacted panel vector auto-regressive (IP-VAR) model estimated for a panel data of 16 OECD economies. The empirical findings confirm the theoretical results and are robust to various extensions.

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.

Our results have important implications for the implementation of stabilization policies and our analysis contributes to the macroeconomic literature dealing with the interaction between labor market characteristics and economic policy reactions. Unveiling the particular channels at work in shaping the relationship between labor market regulation and fiscal impulses in individual economies requires the use of firm-level data and may prove a fruitful avenue of future research that builds upon the results presented here.

Acknowledgments

We thank Astrid Czaloun for excellent research assistance. Furthermore, Maximilian Böck and Christian Glocker gratefully acknowledge financial support by funds of the Österreichische Nationalbank (Austrian National Bank, Anniversary Fund, project number: 18466). We would like to thank Thomas Url for valuable comments and helpful discussions.

References

ABBRITTI M, AND WEBER S (2018), "Reassessing the Role of Labor Market Institutions for the Business Cycle," *International Journal of Central Banking* **14**(1), 1–34. [2, 4, 12, 20]

AIYAR S, BLUEDORN J, DUVAL R, FURCERI D, GARCIA-MACIA D, JI Y, MALACRINO D, QU H, SIMINITZ J, AND ZDZIENICKA A (2019), "Strengthening the Euro Area: The Role of National Structural Reforms in Building Resilience," IMF Staff Discussion Note 19/05, International Monetary Fund. [2]

ALBERTINI J, AND POIRIER A (2015), "Unemployment Benefit Extension at the Zero Lower Bound," *Review of Economic Dynamics* **18**(4), 733–751. [15]

AMUEDO-DORANTES C, AND RICA S (2013), "The immigration surplus and the substitutability of immigrant and native labor: evidence from Spain," *Empirical Economics* **44**(2), 945–958. [5]

ANDOLFATTO D (1996), "Business cycles and labor-market search," *American Economic Review* **86**(1), 112–132. [8, 38]

ARNOLD N, BARKBU B, TURE E, WANG H, AND YAO J (2018), "A Central Fiscal Stabilization Capacity for the Euro Area," IMF Staff Discussion Note 18/03, International Monetary Fund. [2]

AUERBACH AJ, AND GORODNICHENKO Y (2012), "Measuring the Output Responses to Fiscal Policy," *American Economic Journal: Economic Policy* **4**(2), 1–27. [2, 3]

BALL L, JALLES JT, AND LOUNGANI P (2015), "Do Forecasters Believe in Okun's Law? An Assessment of Unemployment and Output Forecasts," *International Journal of Forecasting* **31**(1), 176–184. [4]

BANERJI A, CRISPOLTI V, DABLA-NORRIS E, DUVAL R, EBEKE C, FURCERI D, KOMATSUZAKI T, AND POGHOSYAN T (2017), "Labor and Product Market Reforms in Advanced Economies: Fiscal Costs, Gains, and Support," IMF Staff Discussion Note 17/03, International Monetary Fund. [2]

BEETSMA R, AND GIULIODORI M (2011), "The effects of government purchases shocks: Review and estimates for the EU," *Economic Journal* **121**(550), 4–32. [19]

BENATI L, AND SURICO P (2009), "VAR Analysis and the Great Moderation," *The American Economic Review* **99**(4), 1636–1652. [30]

BÉNÉTRIX A, AND LANE P (2013), "Fiscal Shocks and the Real Exchange Rate," *International Journal of Central Banking* **9**(3), 6–37. [19]

BERNARDINI M, DE SCHRYDER S, AND PEERSMAN G (2020), "Heterogeneous Government Spending Multipliers in the Era Surrounding the Great Recession," *The Review of Economics and Statistics* **102**(2), 304–322. [3]

BLANCHARD O, AND GALÍ J (2007), "Real wage rigidities and the New Keynesian model," *Journal of Money, Credit and Banking* **39**(s1), 35–65. [43]

BLANCHARD O, AND PEROTTI R (2002), "An Empirical Characterization of the Dynamic Effects of Changes in Government Spending and Taxes on Output," *the Quarterly Journal of economics* **117**(4), 1329–1368. [3, 19]

BOIVIN J, AND GIANNONI M (2006), "Has Monetary Policy Become More Effective?" *The Review of Economics and Statistics* **88**(3), 445–462. [4]

BRÜCKNER M, AND PAPPA E (2012), "Fiscal Expansions, Unemployment, and Labor Force Participation: Theory and Evidence," *International Economic Review* **53**(4), 1205–1228. [4, 20, 22]

CACCIATORE M, DUVAL R, FIORI G, AND GHIRONI F (2016), "Short-term pain for long-term gain: Market deregulation and monetary policy in small open economies," *Journal of International Money and Finance* **68**, 358–385. [2, 4]

CACCIATORE M, DUVAL R, FURCERI D, AND ZDZIENICKA A (2021), "Fiscal multipliers and job-protection regulation," *European Economic Review* **132**, 103616. [4, 9]

ČAPEK J, AND CRESPO CUARESMA J (2020), "We Just Estimated Twenty Million Fiscal Multipliers," *Oxford*

Bulletin of Economics and Statistics **82**(3), 483–502. [3, 19]

Christoffel K, Kuester K, and Linzert T (2009), "The role of labor markets for euro area monetary policy," *European Economic Review* **53**(8), 908–936. [48]

Corsetti G, Meier A, and Müller GJ (2012), "What Determines Government Spending Multipliers?" *Economic Policy* **27**(72), 521–565. [3]

Daniel BC, and Gao S (2015), "Implications of productive government spending for fiscal policy," *Journal of Economic Dynamics and Control* **55**, 148–175. [46]

Demyanyk Y, Loutskina E, and Murphy D (2019), "Fiscal Stimulus and Consumer Debt," *The Review of Economics and Statistics* **101**(4), 728–741. [4]

Doan T, Litterman R, and Sims C (1984), "Forecasting and Conditional Projection Using Realistic Prior Distributions," *Econometric reviews* **3**(1), 1–100. [51]

Docquier F, Kone ZL, Mattoo A, and Ozden C (2019), "Labor market effects of demographic shifts and migration in OECD countries," *European Economic Review* **113**, 297–324. [5]

Duval R, and Furceri D (2018), "The Effects of Labor and Product Market Reforms: The Role of Macroeconomic Conditions and Policies," *IMF Economic Review* **66**(1), 31–69. [2]

Duval R, Furceri D, and Jalles J (2019), "Job protection deregulation in good and bad times," *Oxford Economic Papers* **72**(2), 370–390. [4]

Galí J, López-Salido D, and Valles J (2007), "Understanding the Effects of Government Spending on Consumption," *Journal of the European Economic Association* **5**(1), 227–270. [41, 44]

Ganong P, Noel P, and Vavra J (2020), "US unemployment insurance replacement rates during the pandemic," *Journal of Public Economics* **191**(C). [12]

Gertler M, Huckfeldt C, and Trigari A (2020), "Unemployment Fluctuations, Match Quality, and the Wage Cyclicality of New Hires," *Review of Economic Studies* **87**(4), 1876–1914. [43]

Giambattista E, and Pennings S (2017), "When is the government transfer multiplier large?" *European Economic Review* **100**, 525–543. [4]

Godøy A (2017), "Local labor markets and earnings of refugee immigrants," *Empirical Economics* **52**(1), 31–58. [5]

Griffin JE, and Brown PJ (2010), "Inference with normal-gamma prior distributions in regression problems," *Bayesian Analysis* **5**(1), 171–188. [18, 50]

Guajardo J, Leigh D, and Pescatori A (2014), "Expansionary Austerity? International Evidence," *Journal of the European Economic Association* **12**(4), 949–968. [19]

Hagedorn M, Manovskii I, and Mitman K (2019), "The Fiscal Multiplier," Working Paper 25571, National Bureau of Economic Research. [4]

Hall RE (2005), "Employment Fluctuations with Equilibrium Wage Stickiness," *American Economic Review* **95**(1), 50–65. [42]

Hamilton JD (1994), *Time Series Analysis*, Princeton University Press. [25, 48]

Hantzsche A, Savsek S, and Weber S (2018), "Labour market adjustments to financing conditions under sectoral rigidities in the Euro Area," *Open Economies Review* **29**(4), 769–794. [4]

Hernández de Cos P, and Moral-Benito E (2016), "Fiscal multipliers in turbulent times: The case of Spain," *Empirical Economics* **50**(4), 1589–1625. [3]

Huber F, and Feldkircher M (2019), "Adaptive shrinkage in Bayesian vector autoregressive models," *Journal of Business & Economic Statistics* **37**(1), 27–39. [18, 50]

Huidrom R, Kose MA, Lim JJ, and Ohnsorge FL (2020), "Why do fiscal multipliers depend on fiscal

Positions?" *Journal of Monetary Economics* **114**(C), 109–125. [19]

ILZETZKI E, MENDOZA EG, AND VÉGH CA (2013), "How big (small?) are fiscal multipliers?" *Journal of Monetary Economics* **60**(2), 239–254. [2, 3, 19, 22]

JAROCIŃSKI M (2010), "Responses to Monetary Policy Shocks in the East and the West of Europe: A Comparison," *Journal of Applied Econometrics* **25**(5), 833–868. [18, 50]

KILIAN L, AND LÜTKEPOHL H (2017), *Structural vector autoregressive analysis*, Cambridge University Press. [25]

KRAAY A (2012), "How large is the Government Spending Multiplier? Evidence from World Bank Lending," *The Quarterly Journal of Economics* **127**(2), 829–887. [19]

KRAUSE MU, AND LUBIK TA (2007), "The (ir)relevance of real wage rigidity in the New Keynesian model with search frictions," *Journal of Monetary Economics* **54**(3), 706–727. [9, 42]

KUMHOF M, LAXTON D, MUIR D, AND MURSULA S (2010), "The Global Integrated Monetary and Fiscal Model (GIMF) – Theoretical Structure," IMF Working Papers 10/34, International Monetary Fund. [46]

LEEPER EM, TRAUM N, AND WALKER TB (2017), "Clearing up the fiscal multiplier morass," *American Economic Review* **107**(8), 2409–54. [2]

LI LB, ZHANG CC, HU JL, AND CHIU CR (2021), "Disaggregate productivity growth sources of regional industries in China," *Empirical Economics* **60**(3), 1531–1557. [5]

LUBIK T, AND SCHORFHEIDE F (2004), "Testing for Indeterminacy: An Application to U.S. Monetary Policy," *American Economic Review* **94**(1), 190–217. [4]

MASUCH K, ANDERTON R, SETZER R, AND BENALAL N (2018), "Structural policies in the euro area," Occasional Paper Series 210, European Central Bank. [2]

MERTENS K, AND RAVN MO (2010), "Measuring the impact of fiscal policy in the face of anticipation: A structural VAR approach," *Economic Journal* **120**(544), 393–413. [20]

MERZ M (1995), "Search in the labor market and the real business cycle," *Journal of Monetary Economics* **36**(2), 269–300. [8, 10, 38]

MONACELLI T, PEROTTI R, AND TRIGARI A (2010), "Unemployment fiscal multipliers," *Journal of Monetary Economics* **57**(5), 531–553. [4, 9, 22, 38]

MOUNTFORD A, AND UHLIG H (2009), "What are the effects of fiscal policy shocks?" *Journal of applied econometrics* **24**(6), 960–992. [19]

ORDINE P, AND ROSE G (2016), "Two-tier labor market reform and entry wage of protected workers: Evidence from Italy," *Empirical Economics* **51**(1), 339–362. [4]

PARK T, AND CASELLA G (2008), "The Bayesian LASSO," *Journal of the American Statistical Association* **103**(482), 681–686. [51]

PÉREZ E, AND YAO Y (2015), "Can institutional reform reduce job destruction and unemployment duration? Yes it can," *Empirical Economics* **49**(3), 961–983. [4]

POLICARDO L, PUNZO LF, AND SÁNCHEZ CARRERA E (2019), "On the wage–productivity causal relationship," *Empirical Economics* **57**(1), 329–343. [5]

RAMEY VA (2011), "Identifying government spending shocks: It's all in the timing," *The Quarterly Journal of Economics* **126**(1), 1–50. [19]

——— (2019), "Ten Years After the Financial Crisis: What Have We Learned from the Renaissance in Fiscal Research?" *Journal of Economic Perspectives* **33**(2), 89–114. [22]

RAMEY VA, AND SHAPIRO MD (1998), "Costly capital reallocation and the effects of government spending," *Carnegie-Rochester Conference Series on Public Policy* **48**(1), 145–194. [19]

RAMEY VA, AND ZUBAIRY S (2018), "Government spending multipliers in good times and in bad: Evidence from US historical data," *Journal of Political Economy* **126**(2), 850–901. [4]

ROMER CD, AND ROMER DH (2010), "The Macroeconomic Effects of Tax Changes: Estimates Based on a New Measure of Fiscal Shocks," *American Economic Review* **100**(3), 763–801. [19]

SÁ F, TOWBIN P, AND WIELADEK T (2014), "Capital Inflows, Financial Structure And Housing Booms," *Journal of the European Economic Association* **12**(2), 522–546. [17, 50]

SCHIMAN S (2021), "Labor supply shocks and the Beveridge Curve: Empirical evidence from EU enlargement," *Review of Economic Dynamics* **40**, 108–127. [5]

SIMS C, AND ZHA T (2006), "Were There Regime Switches in U.S. Monetary Policy?" *American Economic Review* **96**(1), 54–81. [4]

SONDERMANN D (2018), "Towards more resilient economies: The role of well-functioning economic structures," *Journal of Policy Modeling* **40**(1), 97–117. [2]

STOCK J, AND WATSON M (2002), "Has the Business Cycle Changed and Why?" NBER Working Papers 9127, National Bureau of Economic Research, Inc. [4]

TOWBIN P, AND WEBER S (2013), "Limits of floating exchange rates: The role of foreign currency debt and import structure," *Journal of Development Economics* **101**(C), 179–194. [17, 50]

TRIGARI A (2009), "Equilibrium Unemployment, Job Flows, and Inflation Dynamics," *Journal of Money, Credit and Banking* **41**(1), 1–33. [41]

TURRINI A (2013), "Fiscal Consolidation and Unemployment: Does EPL Matter?: A Look at EU Countries," *IZA Journal of Labor Policy* **2**(1), 1–18. [4]

ZANETTI F (2007), "A non-Walrasian labor market in a monetary model of the business cycle," *Journal of Economic Dynamics and Control* **31**(7), 2413–2437. [43]

——— (2009), "Effects of product and labor market regulation on macroeconomic outcomes," *Journal of Macroeconomics* **31**(2), 320–332. [2, 4]

——— (2011), "Labor market institutions and aggregate fluctuations in a search and matching model," *European Economic Review* **55**(5), 644–658. [2, 4]

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 $A(\tilde{a}_t) = \int_{\tilde{a}_t}^{\infty} \frac{a}{1-F(\tilde{a}_t)} dF(a)$
- $\frac{\kappa}{q_t} = E_t \Lambda_{t,t+1} \left[(1-\varrho(\tilde{a}_{t+1})) F^n_{t+1} - b^s_{t+1}(1-\bar{\varrho}) F(\tilde{a}_{t+1}) \right]$
- $F^n_t = mpl_t - w_t + E_t \Lambda_{t,t+1} \left[(1-\varrho(\tilde{a}_{t+1})) F^n_{t+1} - b^s_{t+1}(1-\bar{\varrho}) F(\tilde{a}_{t+1}) \right]$
- $A(\tilde{a}_t) = \frac{1}{\bar{A}} \left(w_t - b^s_t - \frac{\kappa}{q_t} \right)$

Households

- $1 = E_t \left[\Lambda_{t,t+1} \right] R_t$ 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 $\theta_t = v_t / u_t$
- $m_t = \bar{m} u_t^{\gamma} v_t^{1-\gamma}$
- $w_t = (1-\eta) \frac{mrs_t + b^u_t}{1-\tau} + \eta \left(mpl_t + E_t \Lambda_{t,t+1} \left[\kappa \theta_{t+1} - b^s_{t+1}(1-\bar{\varrho}) F(\tilde{a}_{t+1}) \right] \right)$

Constraints and Policy

- $\tau w_t n_t + B_t = R_{t-1} B_{t-1} + b^u_t u_t + \bar{T}^s_t + g_t$
- $y_t = c_t + g_t + \kappa v_t + F(\tilde{a}_t)(1-\bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b^s_t$
- $\hat{g}_t \sim$ AR(1), $b^s_t = \bar{\varsigma} + \varsigma w_{t-1}$, $b^u_t = \varphi w_{t-1}$ and $T^s_t = \bar{T}^s + \varphi_{T^s} B_t$

where $mpl_t = y_t / n_t$ and a is log-normally distributed of which F is the c.d.f. The expression for the total surplus is finally given by: $S_t = F^n_t + H^n_t$ where $H^n_t = (1-\tau) w_t - b^u_t - mrs_t + (1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}) E_t \left[\Lambda_{t,t+1} H^n_{t+1} \right]$.

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 et al. (2000), we also calibrate the exogenous job destruction rate $\bar{\varrho}$. The idiosyncratic productivity is assumed to be i.i.d. log-normally distributed with c.d.f. F of which we calibrate the first and second moments ($\mu_a = E[\ln(a)]$ and $\sigma_a = \sqrt{Var[\ln(a)]}$). 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 $n = p/(1 - \varsigma + p)$ 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 (3.2), (3.3) and (3.9) 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 (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 $\bar{T}^s = \tau w n - \varphi w(1 - n) - g$. 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 Tab. 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. The specific values chosen are standard in the literature (Merz, 1995; Andolfatto, 1996; Monacelli *et al.*, 2010). 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 0.6.

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

Our baseline results rest upon a general calibration which does not address a specific country. While the purpose of this exercise is to highlight the general effects of the LMIs on fiscal spending multipliers, the results presented in

Table A1: Calibration of the Model.

Parameter	Description	Value	Range
α	Elasticity of labor in the production function	0.7	[0.5 – 1.0]
β	Discount factor	0.997	[0.95 – 0.999]
γ	Elasticity of matching of unemployed persons	0.5	[0.05 – 0.95]
g_y	Government consumption share in total output	0.3	[0.1 – 0.5]
ζ	Ratio of mrs to mpl	0.8	[0.65 – 0.95]
θ	Labor market tightness	0.5	[0.05 – 0.95]
p	Probability of an unemployed person finding a job	0.45	[0.05 – 0.95]
τ	Labor tax rate	0	[0.0 – 0.5]
μ_a	Steady state mean of idiosyncratic productivity	0.0	[0.0 – 2]
σ_a	Steady state standard-deviation of idiosyncratic productivity	0.15	[0.05 – 3.0]
$\bar{\varrho}$	Exogenous job separation rate	0.05	[0.01 – 0.15]
$\varrho(\tilde{a})$	(Overall) Job separation rate	0.10	[0.05 – 0.30]
η	Bargaining power of workers (UD)	0.5	—
φ	Unemployment benefit replacement rate (BRR)	0.0	—
ς	Firing costs in relation to last wage (EPL)	0.0	—

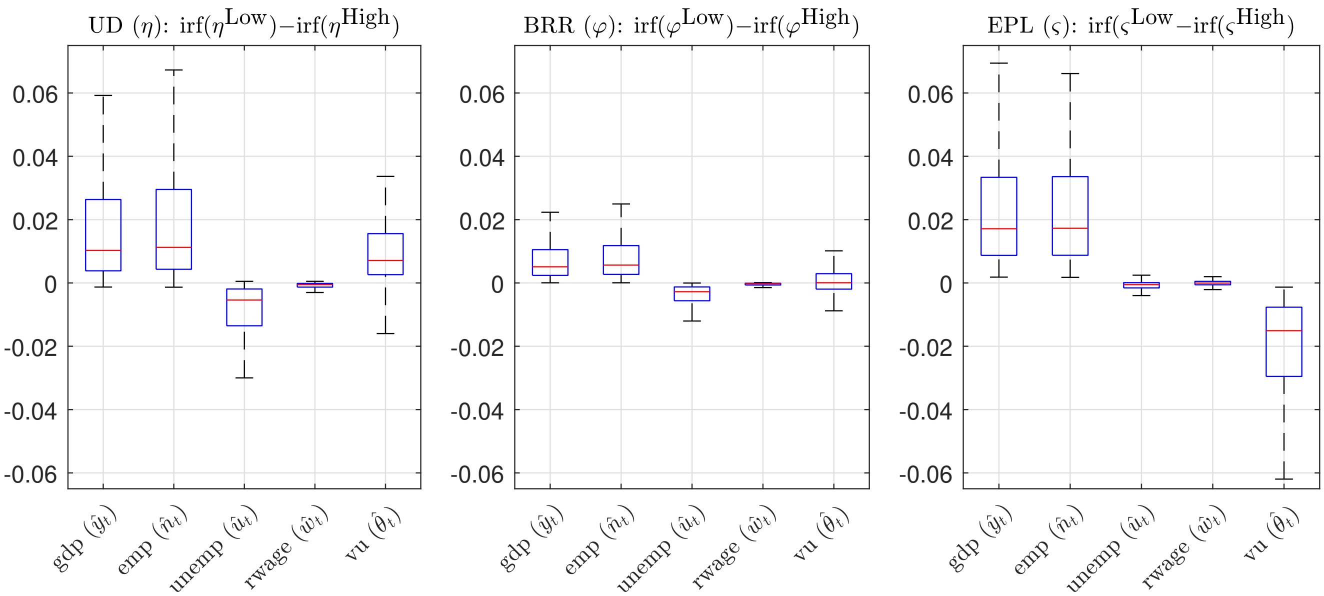
Section 3.7 might, however, be due to the specific calibration chosen. In order to assess the validity of the model's implications in a more general setting, we now extend the analysis.

We consider a continuum of values for all parameters other than η, φ and ς for which Tab. 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 Tab. 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 Fig. A1.[13] The box-plots show the difference in the impact response for each of the three cases for the following variables: output, employment, unemployment, the labor market tightness (v_t/u_t) and real wage. The difference is computed by considering the impulse response functions with a low value of the parameter of interest relative to a high value.

Considering the output response ($\hat{y}_t$) in the left sub-panel as an example, we notice that it is positive throughout due to the fact that the impact response of output when workers have a low power within the wage 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 Fig. 4 and Fig. D1. The employment response replicates the one 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

[13] We draw values for the structural parameters shown in Tab. A1. For instance, in the case for η: for a particular draw, we solve the model for $\eta = 0.5$ and compute impulse response functions. For the same draw we also solve the model using $\eta = 0.6$ – in both cases holding the remaining parameters fixed. The difference in the impact values of the impulse response functions is depicted in Fig. 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.



creation and job destruction activities by firms on the one hand, and labor supply decisions of households on the other hand.

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

We conclude that the general results provided here confirm those put forward in Section 3.7. 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 Extensions to the Theoretical Model

This section considers various extensions to the baseline model outlined in Section 3. These include price stickiness, real wage rigidity, limited asset market participation of one group of households and the case when firing costs accrue to the government as revenues. We always consider one extension at a time, as otherwise the precise role of the additional frictions considered becomes difficult to assess.

A.4.1 Markup Pricing and Monopolistic Competition

While a neo-classical set-up, as considered in Section 3, is commonly considered when analyzing the effects of fiscal policy, it misses one important element. In particular, in the neo-classical set-up, the expansionary effect of higher government spending on output results from the wealth effect on leisure. This channel has been viewed critically especially also because it is unrelated to Keynesian arguments based on increases in private demand that are aggravated by the slow (rather than fast) increase in prices in response to an expansionary fiscal spending shock. Against this background, we now consider an extension of the model of Section 3 that features sticky rather than flexible prices.

The key element of price stickiness in the context of fiscal policy shocks pertains to the behavior of markups. Rigid prices render markups counter-cyclical in light of any shock that boosts output and there fore nominal marginal costs. Markups shift the standard marginal product of labor schedule, which reinforces the effect on employment stemming from the wealth effect on labor supply. As highlighted in Galí *et al.* (2007), the fiscal spending multiplier increases with the extent of price stickiness.

We use the approach put forth in Trigari (2009) to introduce monopolistic competition and nominal price rigidity. This extends the model of Section 3 by a standard monopolistically competitive retail sector in which we locate inertia in price setting. The firm sector where search frictions are located is kept unchanged and is re-labeled as intermediate goods sector for convenience. Retailers acquire goods from intermediate goods firms in competitive markets, differentiate them with a technology that transforms one unit of intermediate goods into one unit of retail goods, and re-sell them to the households. Prices are adjusted according to a conventional Calvo approach by retailers where $1-\kappa$ denotes each period fixed probability of re-setting the price.

Within this extension, the price of intermediate goods in terms of final goods corresponds to the real marginal cost of production faced by the retailers, that is, to the inverse price markup. This implies that the marginal product of labor in the intermediate goods sector expressed in terms of final goods is $\widetilde{mpl}_t = mpl_t/\mu_t$ where μ_t is the price marktup. Hence movements in the markup affect the (shadow) value accruing to the firm when employing one additional worker via the marginal product of labor and hence the equilibrium wage equation (3.9)

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

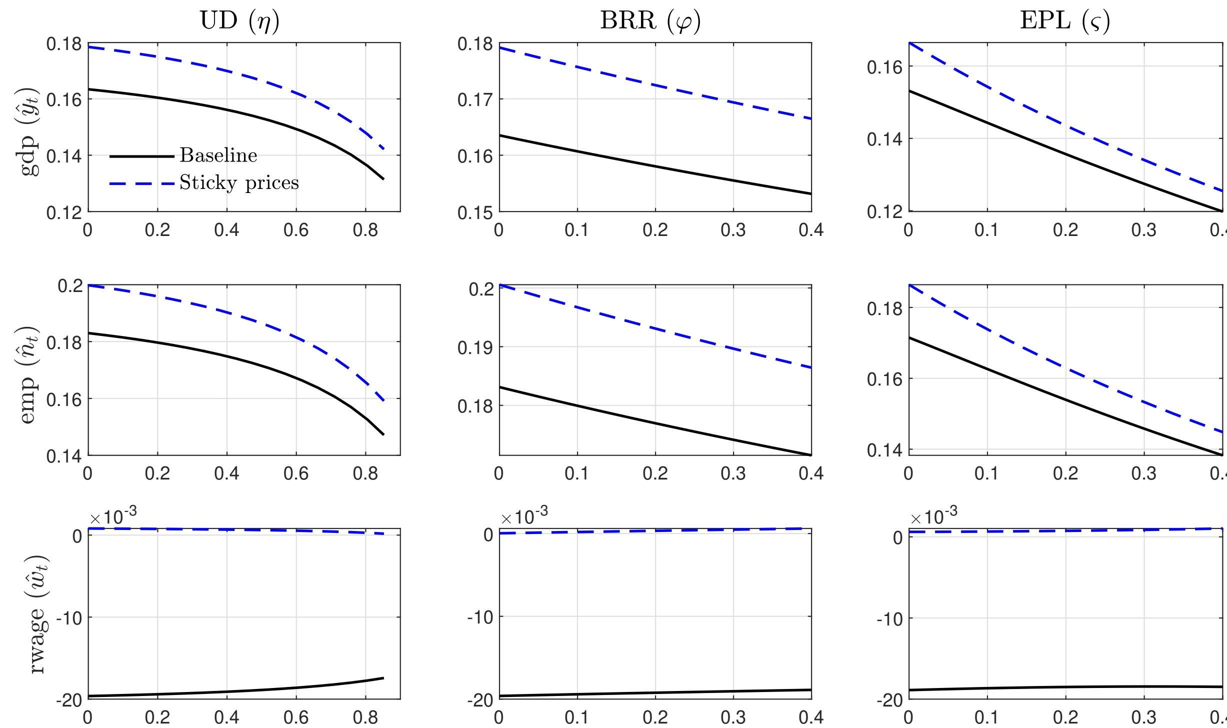
Due to price rigidity, counter-cyclical movements in the markup (μ_t) raise the effective marginal product of labor. In equilibrium, since hiring depends on the current and the expected future values of the marginal product of labor, this boosts the real wage rate, hiring and employment.

Finally, the model is closed by adding an interest rate rule according to which the monetary authority adjusts the short-term nominal interest rate i_t to the inflation rate π_t as follows: $1+i_t = (1+\pi_t)^{\varphi_\pi}$ with $\varphi_\pi > 1$.

We continue to assume $\kappa = 0.7$ and $\varphi_\pi = 1.5$. Fig. A2 displays the multipliers of selected variables from the model for the one-quarter horizon. As can be seen the output multiplier is higher when price stickiness is present. Most importantly, though, is the fact that the reaction in the real wage rate is now even positive which conforms with the empirical results put forth in Section 4.

The reaction of the monetary authority comprises an important additional element in this context: a high degree of nominal rigidities leads to a smaller increase in the real interest rate in response to the higher inflation rate induced by the fiscal expansion (this depends crucially on the reaction of monetary policy); as a result consumption declines by less (or even reacts positively), which in turn aggravates the positive effect on output. Hence, when prices are fully flexible, consumption is always crowded out in response to a rise in government spending, independently of the degree of persistence of the latter. The size of the response of output is increasing in the degree of price rigidities,

Figure A2: Fiscal spending multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – The role of nominal rigidities



Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for a horizons of $\mathcal{P} = 0$, i.e. contemporaneous multiplier.

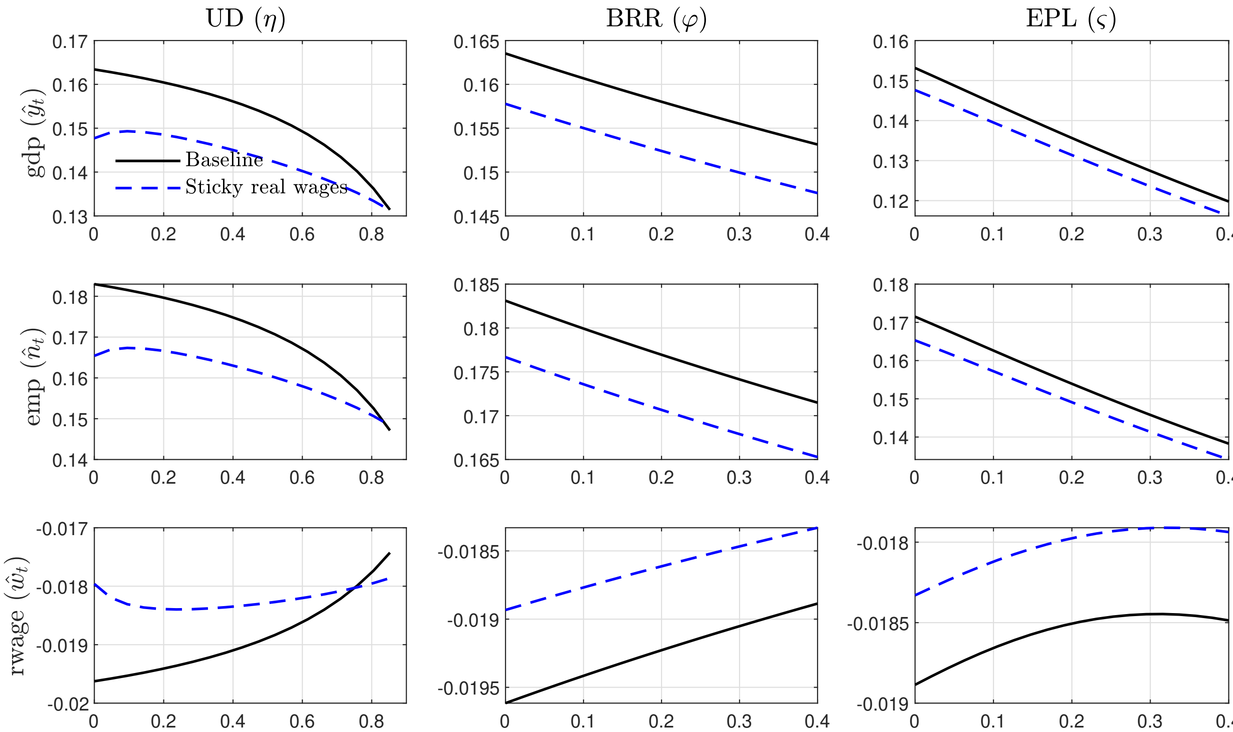
which emerges as a result of a stronger multiplier effect on consumption. The less-negative multiplier effect on consumption transfers naturally to employment and to output.

Despite a noteworthy effect of price stickiness on the size of the multiplier for output and other variables, the implications of the three LMIs for the multiplier as outlined in Section 3.7 still apply. This also holds for the multiplier of the real wage rate even though the sign has changed. It continues to be the case that higher values of the LMIs attenuate the multipliers.

A.4.2 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 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.

Figure A3: Fiscal spending multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – The role of real wage rigidity



Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for a horizons of $\mathcal{P} = 0$, i.e. contemporaneous multiplier.

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 (3.9) to the following

$$w_t = \varrho_w w_{t-1} + (1 - \varrho_w)\check{w}_t \tag{A.2}$$

where $\check{w}_t = (1-\eta)\frac{mrs_t + b_t^u}{1-\tau} + \eta\left(mpl_t + E_t\Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right]\right)$. The parameter ϱ_w captures the extent of real wage rigidity. Equation (A.2) can be considered as 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), Gertler *et al.* (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 Fig. A3. 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.

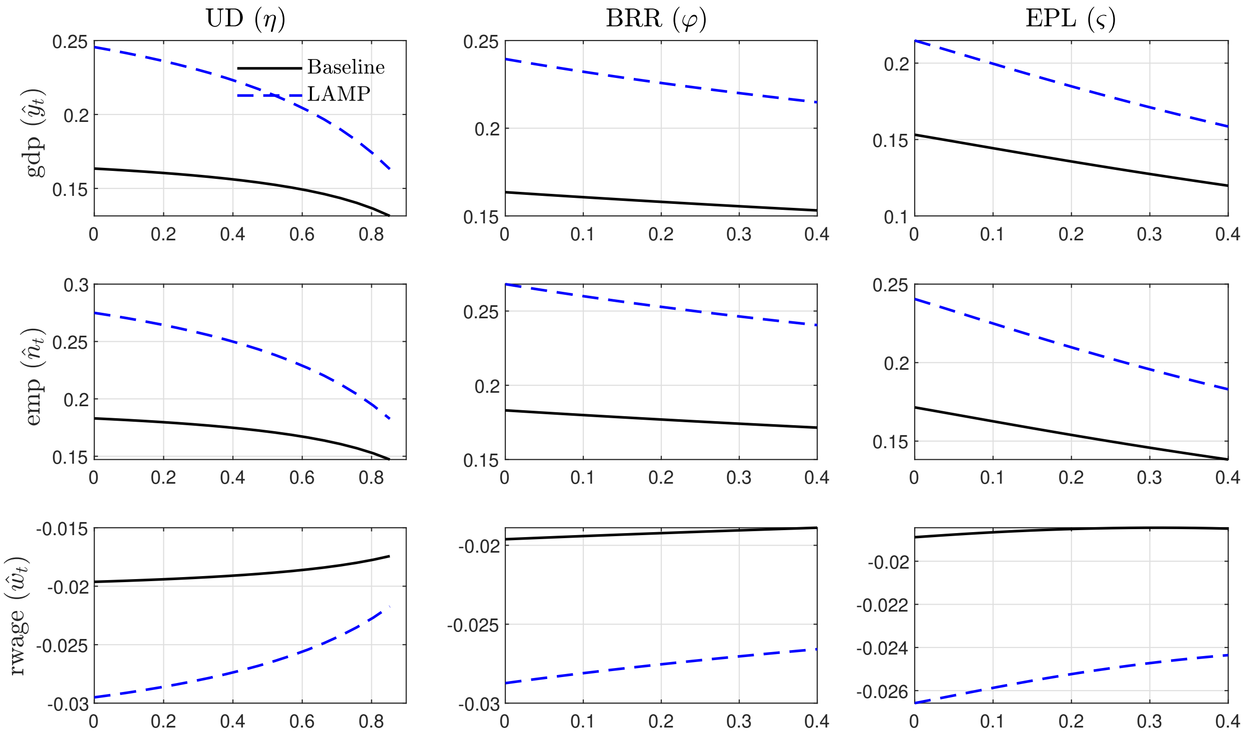
A.4.3 Limited Asset Market Participation

Galí *et al.* (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 *et al.* (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.

Fig. A4 shows the results of the LMIs on the multipliers for output, employment, etc. in the extended model (labeled "LAMP") 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 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 A4: Fiscal spending multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – The role of limited asset market participation



Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for a horizons of $\mathcal{P} = 0$, i.e. contemporaneous multiplier.

A.4.4 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 (equation (3.10)) and the real resource constraint (equation (3.11)) are then given by:

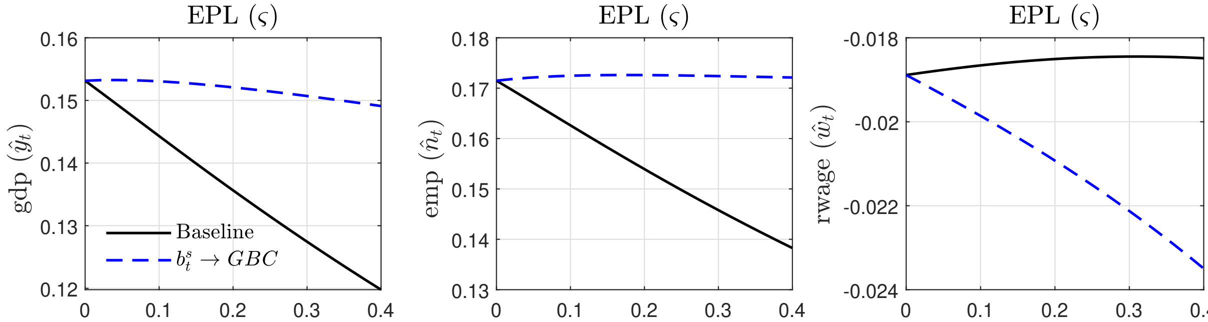
$$F(\tilde{a}_t)(1-\bar{\varrho})(n_{t-1}+q_{t-1}v_{t-1})b_t^s + \tau w_t n_t + B_t = R_{t-1}B_{t-1} + b_t^u u_t + T_t^s + g_t \tag{A.3}$$

$$y_t = c_t + g_t + \kappa v_t \tag{A.4}$$

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 Fig. A5 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

Figure A5: Fiscal spending multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – When firing costs accrue to the government



Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for a horizons of $\mathcal{P} = 0$, i.e. contemporaneous multiplier.

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.

A.4.5 Productivity Enhancing Government Spending

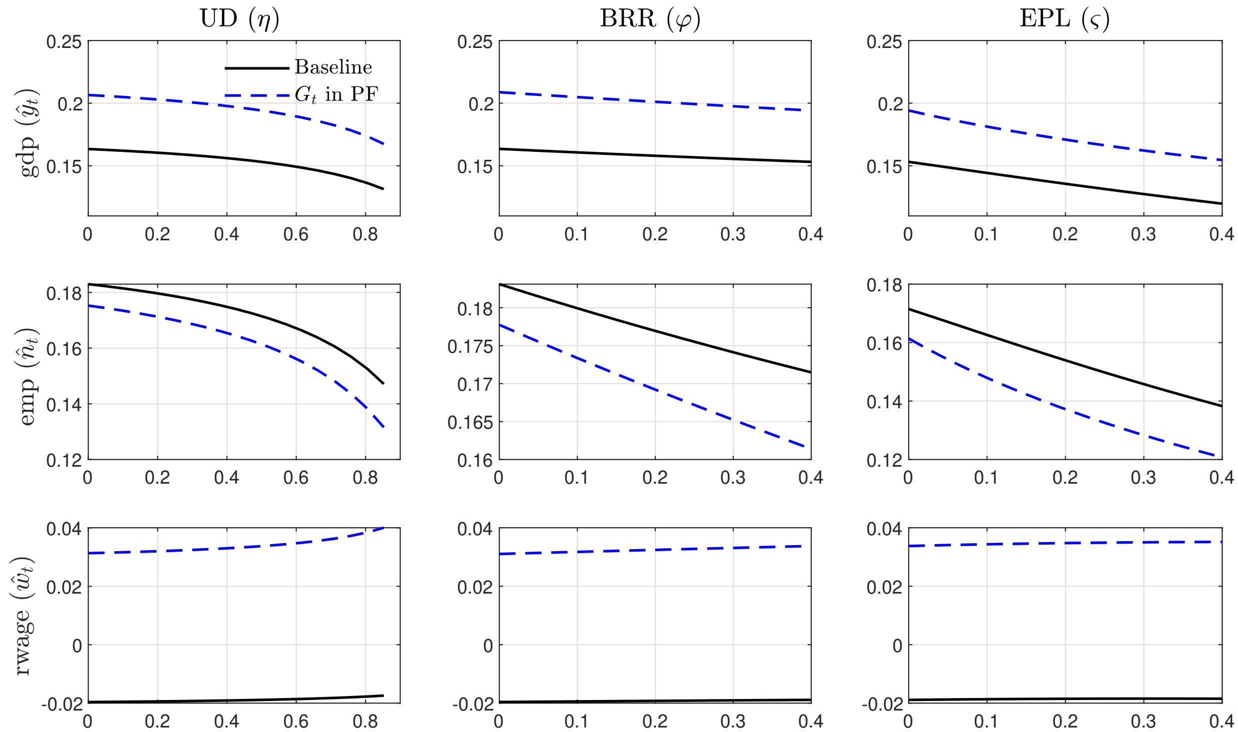
The standard assumption in macroeconomics is that government spending is unproductive. An even more extreme but common assumption is that government spending is entirely purposeless with purchases comprising real resource costs. These assumptions contrast with the observation that various public goods 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 enters the production function directly. Importantly, the public spending 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), though with full depreciation of the public capital stock in each period. We modify the production function as follows

$$y_t(g_t) = \bar{A} n_t A(\tilde{a}_t) \cdot \left(\frac{g_t}{\bar{g}}\right)^{\xi} \tag{A.5}$$

The parameter $\xi \in [0, 1]$ captures the sensitivity (elasticity) of the aggregate production with respect to changes in government spending and $\bar{g}$ is the steady state value for g_t. 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

Figure A6: Fiscal spending multipliers and the LMIs ($\mu(\boldsymbol{\vartheta})$) – Productivity enhancing government spending



Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for a horizons of $\mathcal{P} = 0$, i.e. contemporaneous multiplier.

exogenous technology. Hence government spending augments labor productivity directly: $mpl_t(g_t) = y_t(g_t)/n_t$. We chose a conservative value for the elasticity $\xi = 0.05$ which implies that a one percent increase in government spending (relative to the steady state) raises labor productivity by 0.05 percent.

We carry out the same simulations as in Section 3.7. The results thereof are shown in Fig. A6. 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. The positive reaction in the real wage even surpasses the rise therein when nominal price stickiness is present (compare Fig. A2). Most important aspect for us concerns, though, the impact of the LMIs on the output multiplier. With a view to Fig. A6, while the output response increases with the extent of productive government spending, the dependency of the output multiplier with respect to the three LMIs 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.

A.4.6 The LMIs and macroeconomic volatility – Implications of the theoretical model

The previous Sections highlighted the consequences of various model extensions on the relationship between fiscal multipliers and the LMIs. The current section serves to assess the consequences of the LMIs on the overall macroeconomic volatility. To this purpose, we extend the baseline model to incorporate further shocks. In particular, we add a technology shock ($\bar{A} \rightarrow \bar{A}_t$ with $E(\bar{A}_t) = \bar{A}$), a time-preference shock ($\beta \rightarrow \beta_t$ with $E(\beta_t) = \beta$), a labor-supply preference shock ($\phi \rightarrow \phi_t$ with $E(\phi_t) = \phi$), a shock to the matching technology ($\bar{m} \rightarrow \bar{m}_t$ with $E(\bar{m}_t) = \bar{m}$), a shock to the vacancy posting costs ($\kappa \rightarrow \kappa_t$ with $E(\kappa_t) = \kappa$), and a shock to the exogenous job separation rate ($\bar{\varrho} \rightarrow \bar{\varrho}_t$ with $E(\bar{\varrho}_t) = \bar{\varrho}$). Each shock is specified as an AR(1) with auto-correlation ρ_i and idiosyncratic variance $\sigma_i^2 \; \forall \; i \in \{G, A, \beta, \phi, m, \kappa, \varrho\}$. This set of structural disturbances captures a wide range of distinct shocks involving the goods market (demand and supply), household preferences (time and labor) and the labor market (demand, supply, vacancy posting costs and matching efficiency).

We intend to analyze to what extent macroeconomic volatility is affected by the LMIs. To this purpose, we 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 (3.12) and following Hamilton (1994), the variance covariance matrix $\mathbf{\Sigma}_z(\boldsymbol{\vartheta})$ of the vector of endogenous variables z_t is given by

$$\text{vec}(\mathbf{\Sigma}_z(\boldsymbol{\vartheta})) = (I - \mathbf{\Psi}_1(\boldsymbol{\vartheta}) \otimes \mathbf{\Psi}_1(\boldsymbol{\vartheta}))^{-1} \,\text{vec}\left(\mathbf{\Psi}_0(\boldsymbol{\vartheta})\Sigma_{\epsilon}\mathbf{\Psi}_0'(\boldsymbol{\vartheta})\right) \tag{A.6}$$

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 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 *et al.* (2009) to calibrate the parameters for the auto-correlation coefficients (ρ_i) of the AR(1) shocks. The idiosyncratic variances (σ_i^2) are calibrated such that each shock delivers the same value of the output volatility ($Var(\hat{y}_t)$) at a low value of the LMIs. The benefit of this normalization is that we can then compare the effects of changes in the LMIs on the output and employment volatilities across distinct shocks *directly*. This allows to answer the following questions: Which shocks are most sensitive to the LMIs? We consider two values (high and low) for each LMI in this respect.

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

	Exogenous shocks						
x_t	ϵ_t^G	ϵ_t^A	ϵ_t^β	ϵ_t^m	ϵ_t^ϕ	ϵ_t^κ	ϵ_t^ρ
UD (η): $\text{Var}(x_t(\eta_{\text{low}})) - \text{Var}(x_t(\eta_{\text{high}}))$							
Output ($\hat{y}_t$)	-	+	-	+	-	+	+
Employment ($\hat{n}_t$)	-	-	-	+	+	+	-
Real wage ($\hat{w}_t$)	+	-	-	+	+	+	-
BRR: (φ): $\text{Var}(x_t(\varphi_{\text{low}})) - \text{Var}(x_t(\varphi_{\text{high}}))$							
Output ($\hat{y}_t$)	+	+	-	+	+	+	+
Employment ($\hat{n}_t$)	+	-	-	+	+	+	-
Real wage ($\hat{w}_t$)	+	-	-	+	-	-	-
EPL (ς): $\text{Var}(x_t(\varsigma_{\text{low}})) - \text{Var}(x_t(\varsigma_{\text{high}}))$							
Output ($\hat{y}_t$)	+	+	-	+	-	+	-
Employment ($\hat{n}_t$)	+	-	-	+	-	+	+
Real wage ($\hat{w}_t$)	+	-	+	+	+	-	+

Note: The table shows the sensitivity of the output, employment and the real wage volatilities to changes in the LMIs. The shocks considered are the following ones: technology shock (ϵ_t^A), a time-preference shock (ϵ_t^β), a labor-supply preference shock (ϵ_t^ϕ), a shock to the matching technology (ϵ_t^m), a shock to the vacancy posting costs (ϵ_t^κ), a shock to the exogenous job separation rate (ϵ_t^ϱ), and the government spending shock (ϵ_t^G).

The results are depicted in Tab. A2. The table only shows the sign of the difference in the variance. For instance, in case of output ($\hat{y}_t$) a negative entry implies that the variance of output with respect to a specific shock is high if the respective LMI takes on a high value ($\text{Var}(\hat{y}_t(\vartheta^i_{\text{Low}}) < \text{Var}(\hat{y}_t(\vartheta^i_{\text{High}})))$). A few results emerge from the analysis. First, as regards output, a higher value of any LMI gives rise to lower output volatility. This applies to four out of the seven shocks in case of UD (η) and EPL (ς), while even to six in case of the BRR (φ). Second, the impact of the LMIs on output and employment volatility is qualitatively the same for most shocks. In particular, this applies to each LMI for four out of seven shocks. Finally, there is a trade-off as regards the impact of the LMIs on output and employment volatility on the one hand, and real wage volatility on the other. This applies to four out of seven shocks. To sum up, these results highlight that the LMIs can potentially mitigate macroeconomic volatility. However, this crucially depends on which shocks dominate.

B Bayesian IP-VAR

In this section, we provide estimation details on the Bayesian IP-VAR. The model is similar to the model proposed by Towbin and Weber (2013) and Sá *et al.* (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 M- 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 can write the Interacted Panel Vector Autoregression (IP-VAR) as follows

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

We denote with $\boldsymbol{a}_i$ the $M \times 1$ country-specific intercept vector, while $\boldsymbol{A}_{ij}$ denotes the $M \times M$ country-specific autoregressive coefficient matrix for lag $j = 1, \ldots, p$. The $M \times 1$ vector of residuals $\boldsymbol{u}_{it}$ is assumed to be uncorrelated across countries and normally distributed with mean zero and a $M \times M$ covariance matrix $\boldsymbol{\Omega}_i$. Due to the recursive structure of the VAR, the covariance matrix is diagonal, i.e., $\boldsymbol{\Omega}_i = \text{diag}(\omega_{i1}, \ldots, \omega_{iM})$. The interaction term $\boldsymbol{\vartheta}_{it}$ is allowed to influence the level of the endogenous variables via $\boldsymbol{b}$ ($M \times q$) and the dynamic relationship between the endogenous variables of the system via the $M \times M$ 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 $M \times M$ matrix $\boldsymbol{J}_{it}$, which is a lower unitriangular matrix. This matrix exhibits a time index t because we also allow the interaction term to affect the contemporaneous relationships between equations. In particular, 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

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

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

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

where $\boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}(\boldsymbol{a}_i + \boldsymbol{b}_i\boldsymbol{\vartheta}_{it})$, $\boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}\left(\boldsymbol{A}_{ij} + \sum_{l=1}^{d}\boldsymbol{B}_{ijl}\vartheta_{ilt}\right)$, and $\boldsymbol{\Sigma}_i(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}\boldsymbol{\Omega}_i\boldsymbol{J}_{it}^{-1\prime}$. 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 $\boldsymbol{\vartheta}_{it}$.

The model is estimated in a Bayesian fashion, and thus 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 $k = (1+d)M^2$-dimensional vector $\boldsymbol{\beta}_{ij} = \text{vec}(\boldsymbol{A}_{ij}, \boldsymbol{B}_{ij1}, \ldots, \boldsymbol{B}_{ijd})$ for lag j and country i and specify the prior distribution as follows

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

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 *et al.*, 1984). Hence, the prior distribution on λ_{ij}^2 is a multiplicative Gamma prior

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

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 $\tilde{\boldsymbol{J}}_{it}$

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

with $s = 2, \ldots, M$ 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 vectors / matrices, $\boldsymbol{a}_i$ and $\boldsymbol{b}_i$, we specify for each element a simple Gaussian $\mathcal{N}(0, 10)$ to be uninformative. We have not yet talked about 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

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

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

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

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

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

with $s = 2, \ldots, M$ 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 $\boldsymbol{\Omega}_i$,

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

C Data

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 $N = 19$ countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Italy, Japan, Netherlands, Norway, Portugal, South Korea, Spain, Sweden, Switzerland, United States.

In Tab. C1, we define the exact transformations of the variables used in the estimation. Note that we use year-on-year growth rates. In Tab. C2, 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 C1: Variable Definitions.

Variable	Transformation	Details
$\mathbf{govc}_{it}$	$100 \times \left[\ln\left(\frac{\mathtt{GOVC}_{it}}{\mathtt{POP}_{it}\times\mathtt{PRICE}_{it}}\right) - \ln\left(\frac{\mathtt{GOVC}_{it-4}}{\mathtt{POP}_{it-4}\times\mathtt{PRICE}_{it-4}}\right)\right]$	$\mathtt{GOVC}_{it}$ is *General Government Final Consumption Expenditure*
$\mathbf{gdp}_{it}$	$100 \times \left[\ln\left(\frac{\mathtt{GDP}_{it}}{\mathtt{POP}_{it}\times\mathtt{PRICE}_{it}}\right) - \ln\left(\frac{\mathtt{GDP}_{it-4}}{\mathtt{POP}_{it-4}\times\mathtt{PRICE}_{it-4}}\right)\right]$	$\mathtt{GDP}_{it}$ is *Gross Domestic Product (Current Prices)*
$\mathbf{emp}_{it}$	$100 \times \left[\ln\left(\frac{\mathtt{EMP}_{it}}{\mathtt{POP}_{it}}\right) - \ln\left(\frac{\mathtt{EMP}_{it-4}}{\mathtt{POP}_{it-4}}\right)\right]$	$\mathtt{EMP}_{it}$ is *Total Employment (Persons)*
$\mathbf{unemp}_{it}$	$100 \times \left[\ln\left(\frac{\mathtt{UNEMP}_{it}}{\mathtt{POP}_{it}}\right) - \ln\left(\frac{\mathtt{UNEMP}_{it-4}}{\mathtt{POP}_{it-4}}\right)\right]$	$\mathtt{UNEMP}_{it}$ is *Harmonised Unemployment (Persons)*
$\mathbf{vu}_{it}$	$\ln\left(\frac{\mathtt{VAC}_{it}}{\mathtt{UNEMP}_{it}}\right)$	$\mathtt{VAC}_{it}$ is *Vacancies*
$\mathbf{rwage}_{it}$	$100 \times \left[\ln\left(\frac{\mathtt{WAGE}_{it}}{\mathtt{PRICE}_{it}\times\mathtt{EMP}_{it}\times\mathtt{POP}_{it}}\right) - \ln\left(\frac{\mathtt{WAGE}_{it-4}}{(\mathtt{PRICE}_{it-4}\times\mathtt{EMP}_{it-4}\times\mathtt{POP}_{it-4}}\right)\right]$	$\mathtt{WAGE}_{it}$ is *Wages & Salaries (Current Prices)*
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 C2: Sample Coverage in Different Models.

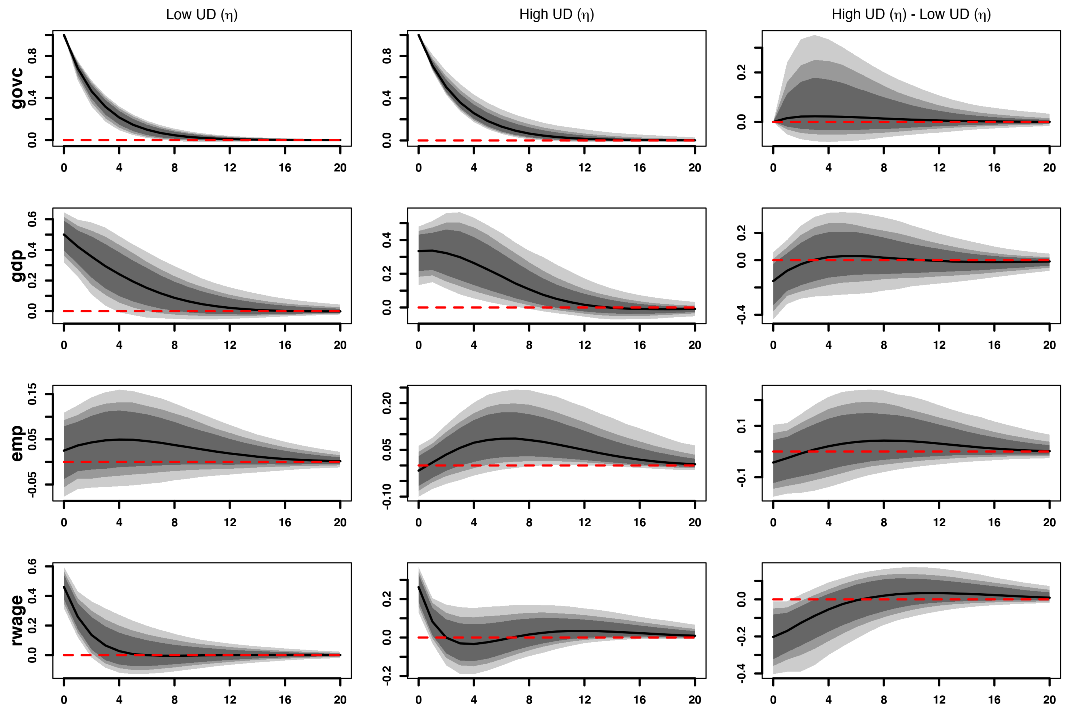
Countries / Model with...	Employment	Unemployment	Tightness
Australia	1966Q3-2020Q2	1964Q1-2020Q4	1978Q2-2020Q4
Austria	1970Q1-2020Q4	1970Q1-2020Q4	1970Q1-2020Q4
Belgium	1980Q4-2020Q4	1980Q4-2020Q4	no data
Canada	1961Q1-2020Q4	1961Q1-2020Q4	no data
Denmark	1980Q1-2020Q4	1980Q1-2020Q4	2009Q1-2020Q4
Finland	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.

D Additional Results

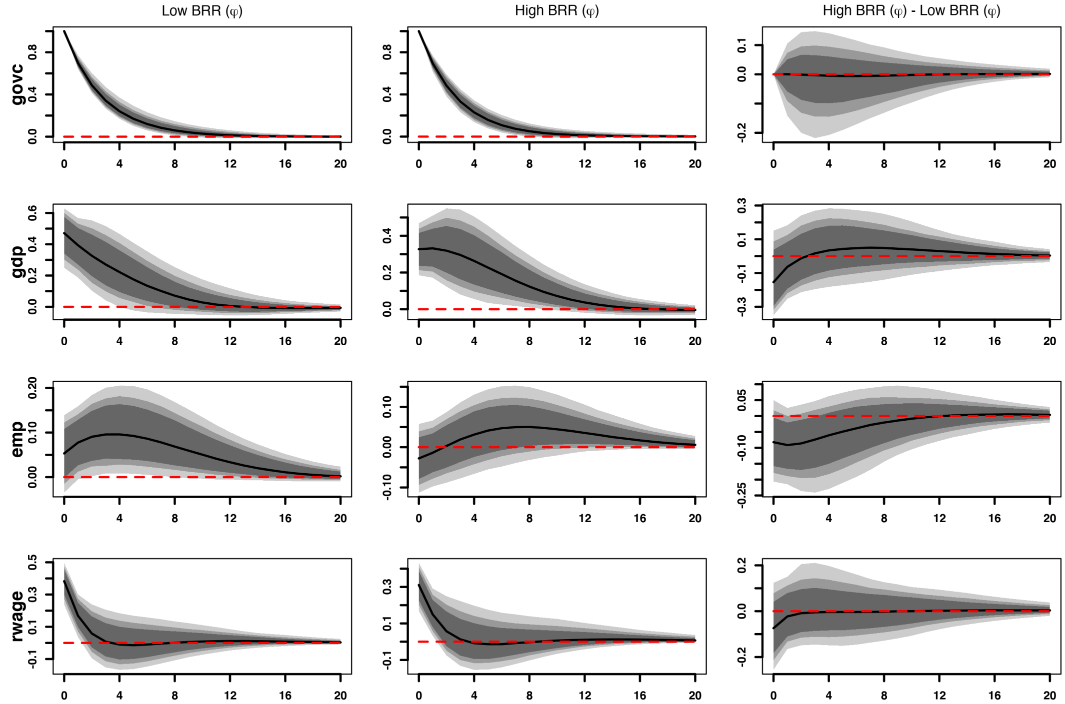
D.1 Additional Results: Effects on Fiscal Spending

Figure D1: Impulse Responses of the Model with Employment - UD (η).



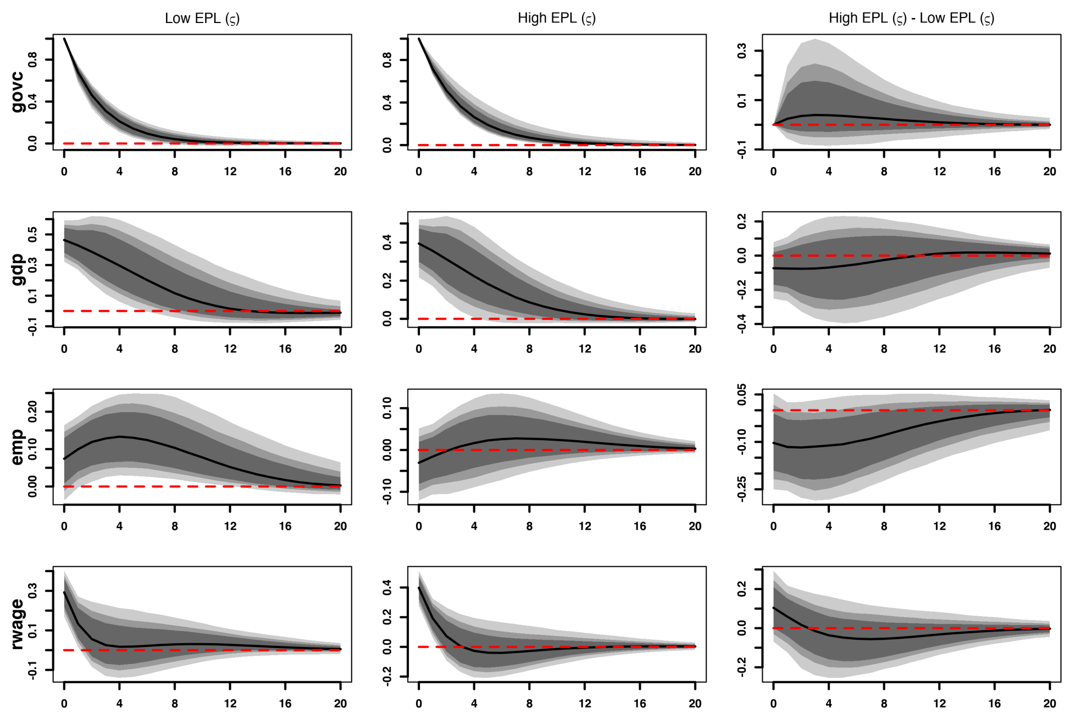
Notes: Impulse response function are shown with 68/80/90 % confidence bounds. Responses are scaled in growth rates for government consumption (govc), gross domestic product (gdp), employment (emp), and real wage (rwage). "Low" indicates −2 standard deviation and "High" indicates +2 standard deviations.

Figure D2: Impulse Responses of the Model with Employment - BRR (φ).



Notes: Impulse response function are shown with 68/80/90 % confidence bounds. Responses are scaled in growth rates for government consumption (govc), gross domestic product (gdp), employment (emp), and real wage (rwage). "Low" indicates −2 standard deviation and "High" indicates +2 standard deviations.

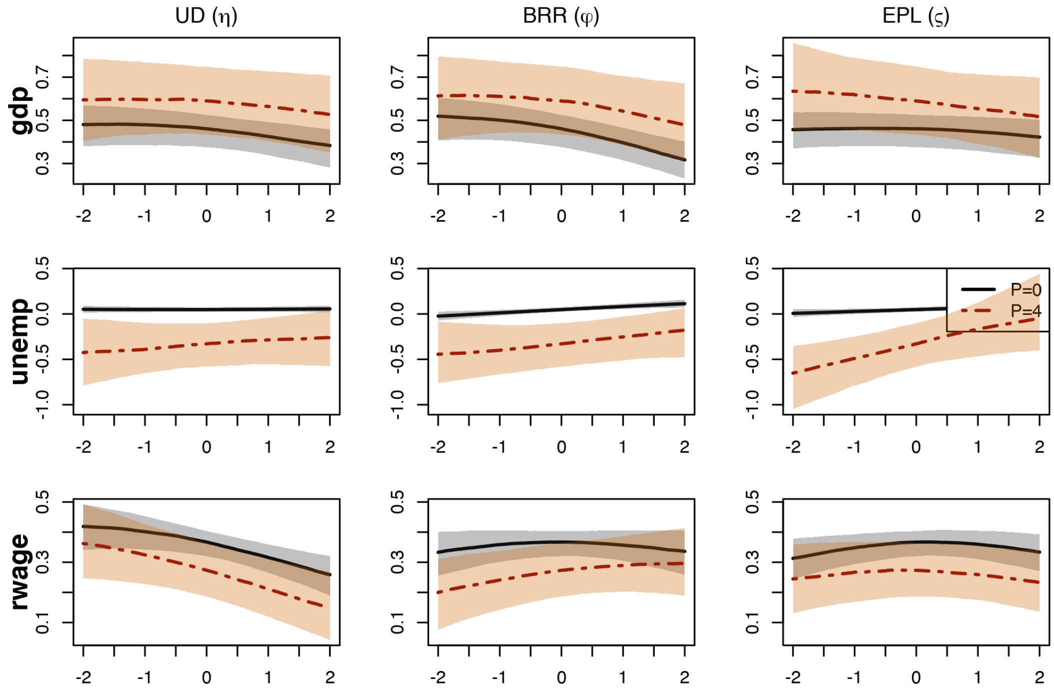
Figure D3: Impulse Responses of the Model with Employment - EPL (ς).



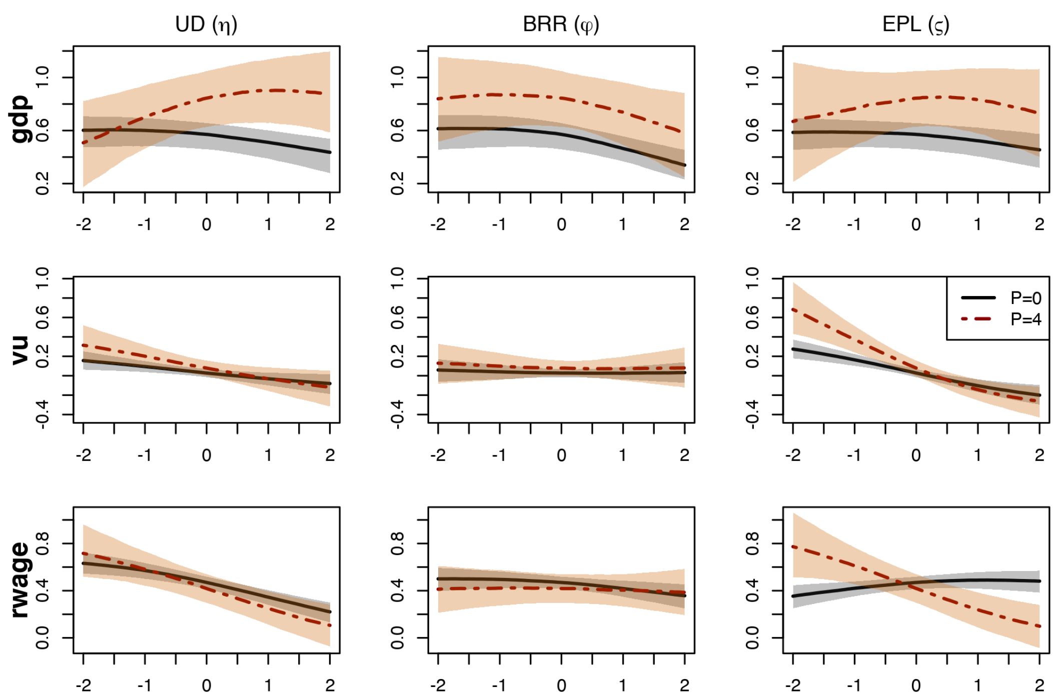
Notes: Impulse response function are shown with 68/80/90 % confidence bounds. Responses are scaled in growth rates for government consumption (govc), gross domestic product (gdp), employment (emp), and real wage (rwage). "Low" indicates −2 standard deviation and "High" indicates +2 standard deviations.

Figure D4: Fiscal Multipliers in the Models with Other Variables.

(a) Unemployment.



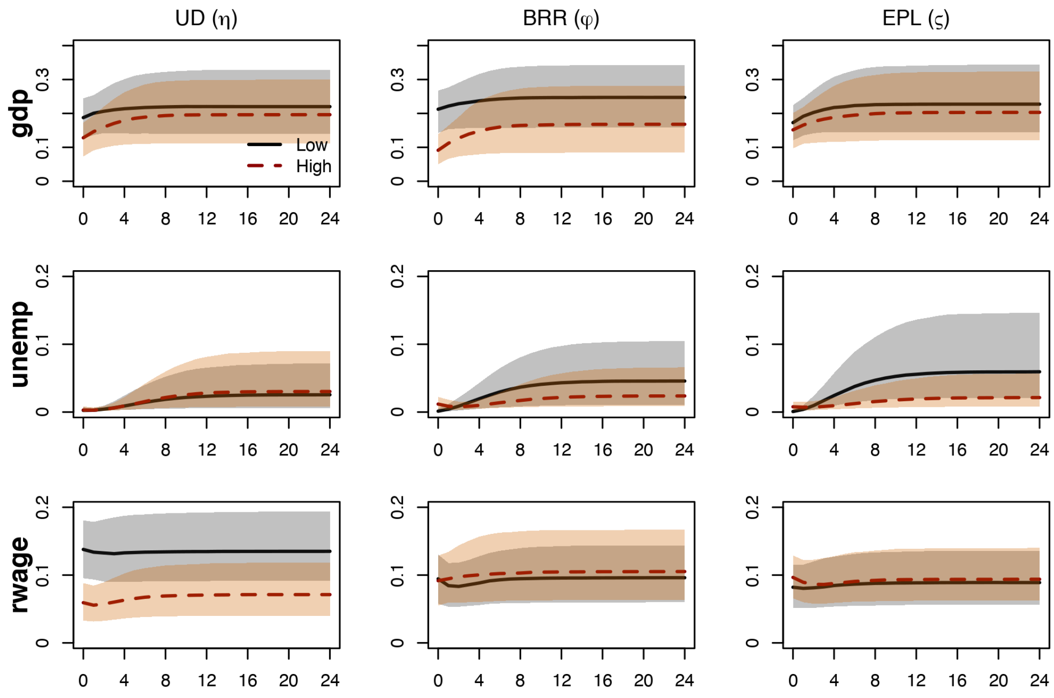
(b) Labor Market Tightness.



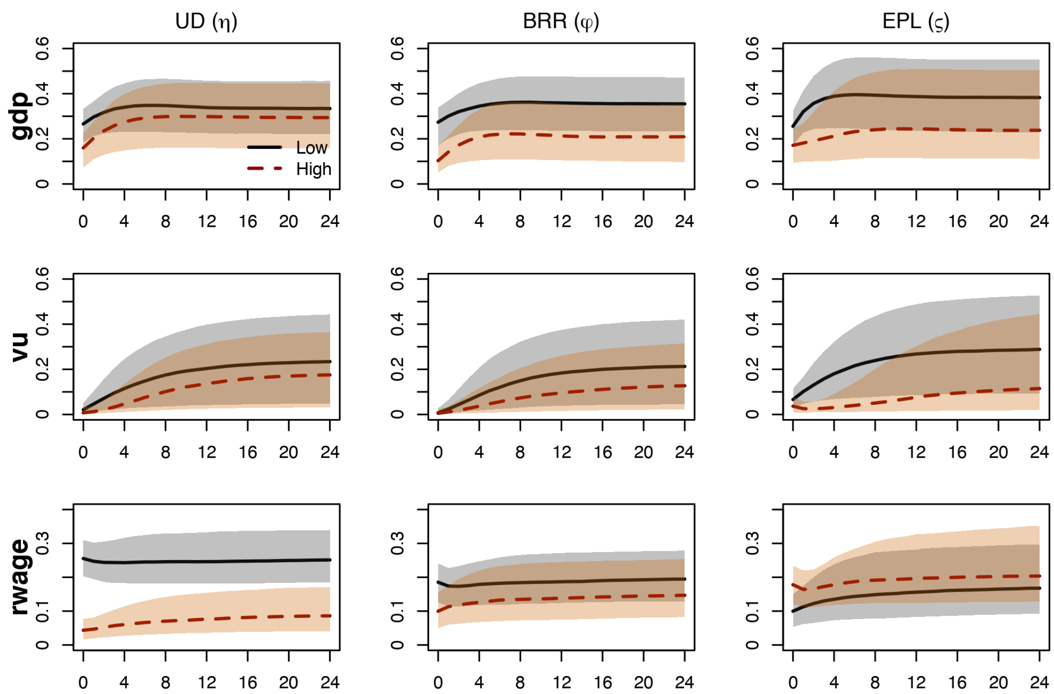
Notes: The sub-plot shows the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is unemployment benefit replacement rate, and ς is employment protection). The y-axis gives the size of the multiplier while the x-axis runs from -/+ 2 standard deviations in terms of the respective LMI. The multipliers are shown for different horizons: contemporaneous multiplier ($\mathcal{P}$ = 0) and four quarters ($\mathcal{P}$ = 4). Confidence bounds refer to the 16/84 quantile of the posterior distribution.

Figure D5: Forecast Error Variance Decomposition in the Models with Other Variables.

(a) Unemployment.



(b) Labor Market Tightness.

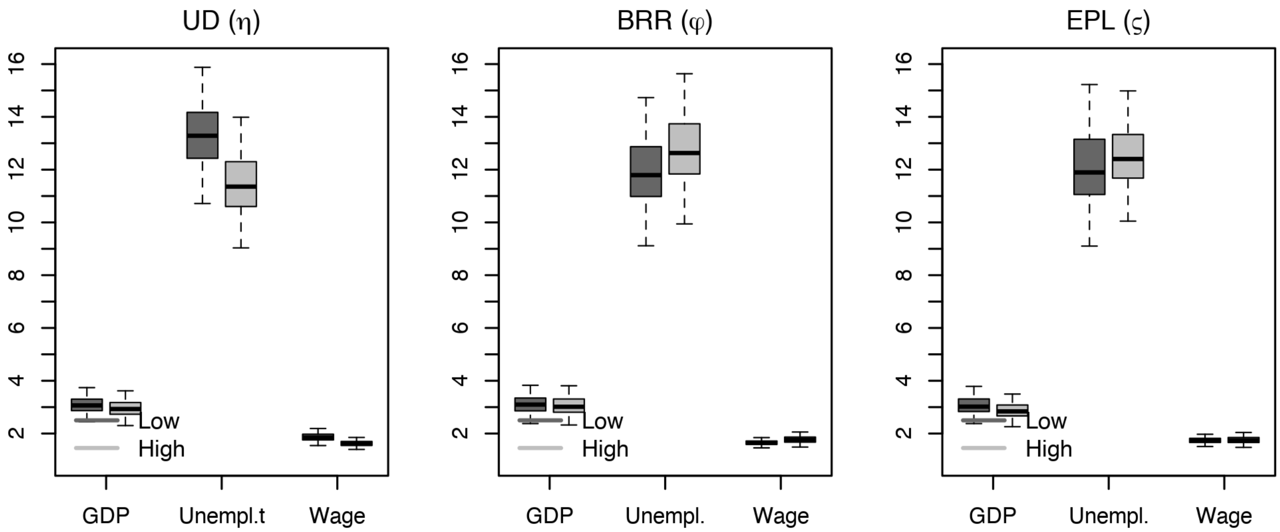


Notes: The sub-plots show the sensitivity of the explained forecast error variance to changes in the structural parameters (η is union density, φ is unemployment benefit replacement rate, and ς is employment protection). The y-axis gives the share of explained forecast error variance while the x-axis is the forecast horizon and runs up to 6 years (=24 months). The FEVD is shown for a regime with low (−2sd) and high (+2sd) LMIs.

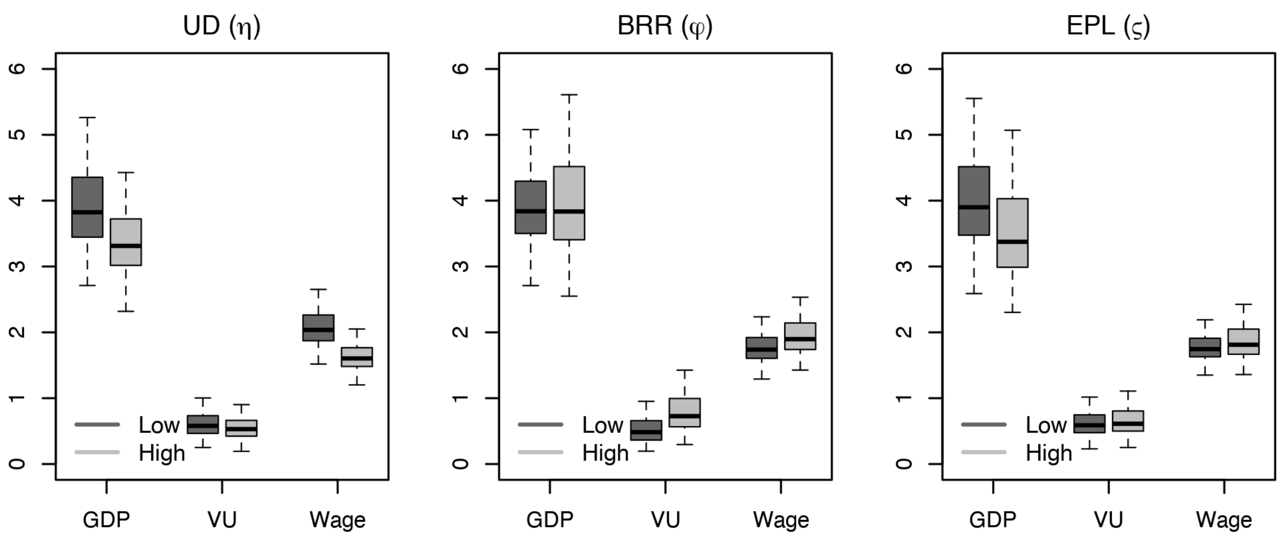
D.2 Additional Results: Effects on Macroeconomic Volatility

Figure D6: Macroeconomic Volatilities Along LMIs with Other Variables.

(a) Unemployment.



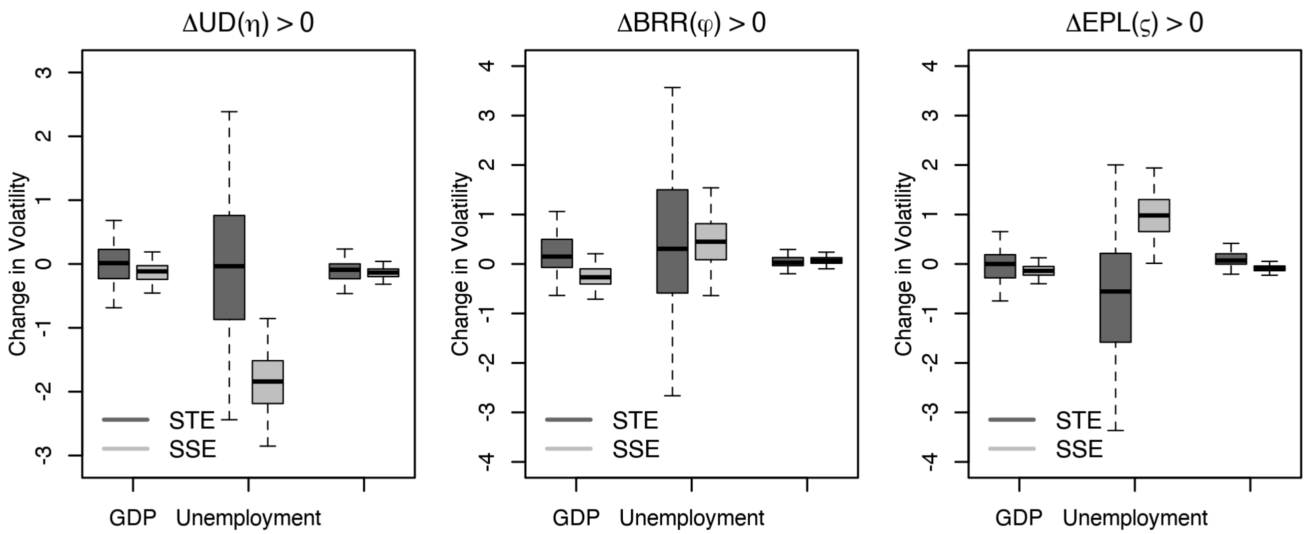
(b) Labor Market Tightness.



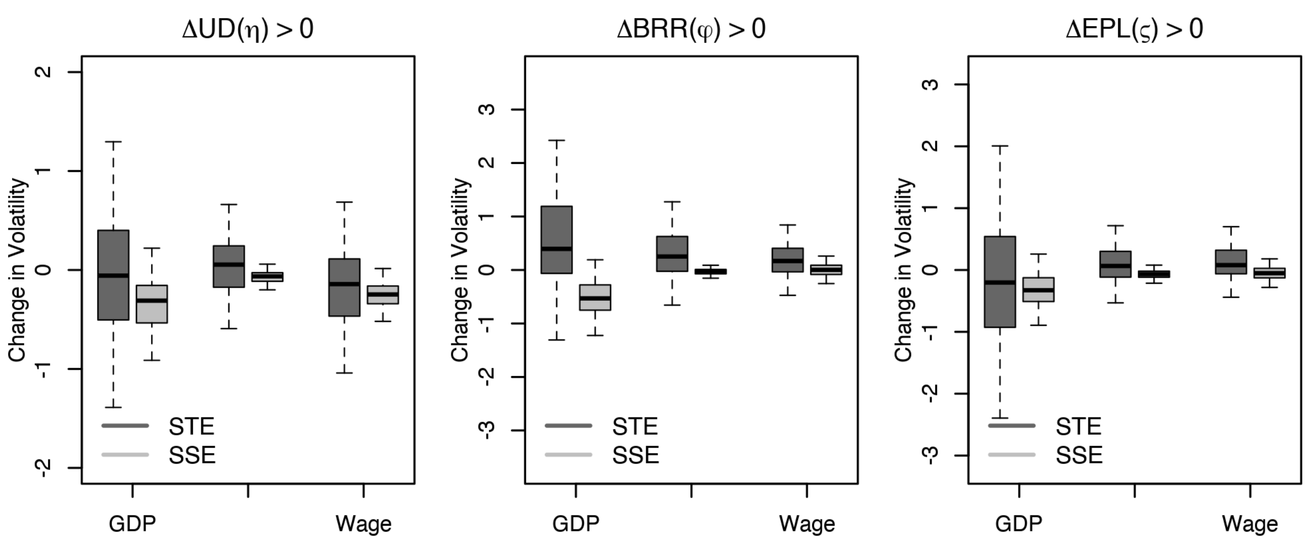
Notes: Each sub-plot shows the standard deviations of the respective macroeconomic variable in a regime with low (−2sd) and high (+2sd) LMIs. The LMIs under consideration are union density (UD(η)), unemployment benefit replacement rate (BRR(φ)), and employment protection (EPL(ς)).

Figure D7: Change in Macroeconomic Volatilities Along LMIs with Other Variables.

(a) Unemployment.



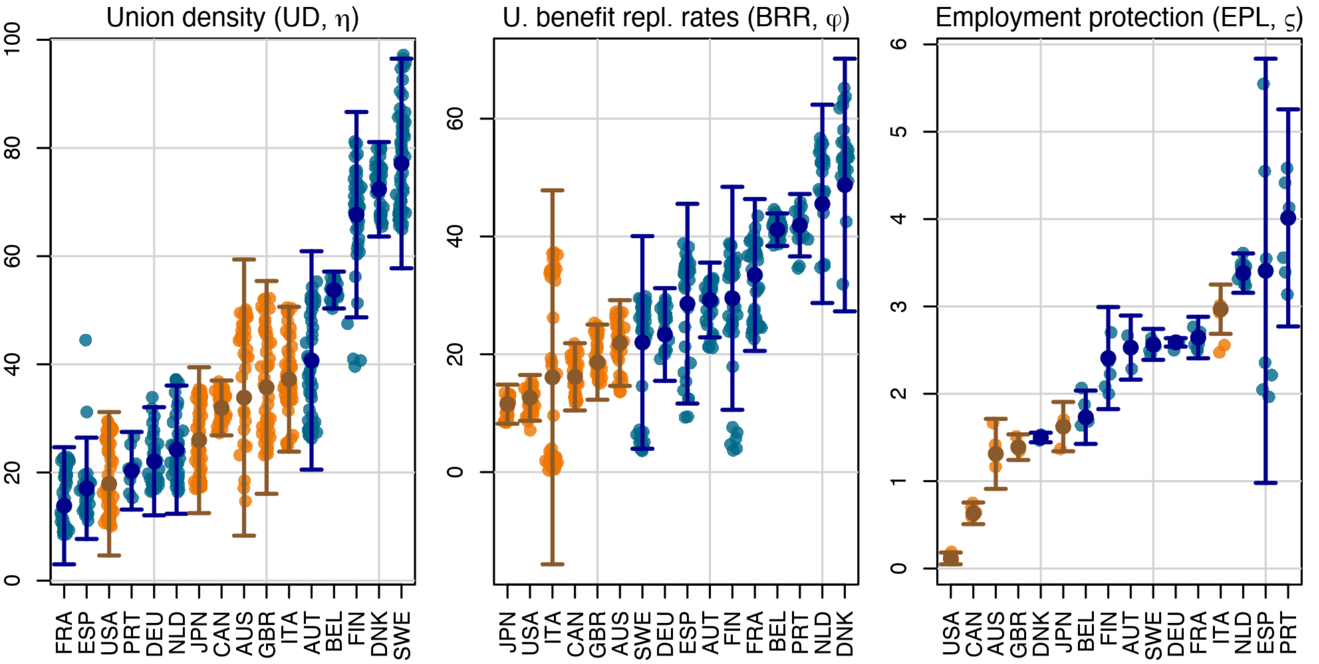
(b) Labor Market Tightness.



Notes: Each sub-plot shows the change in the standard deviations of the respective macroeconomic variable when going from a regime with high (+2sd) to low (−2sd) LMIs. 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 (EPL(ς)).

D.3 Additional Results: Between Country Variation

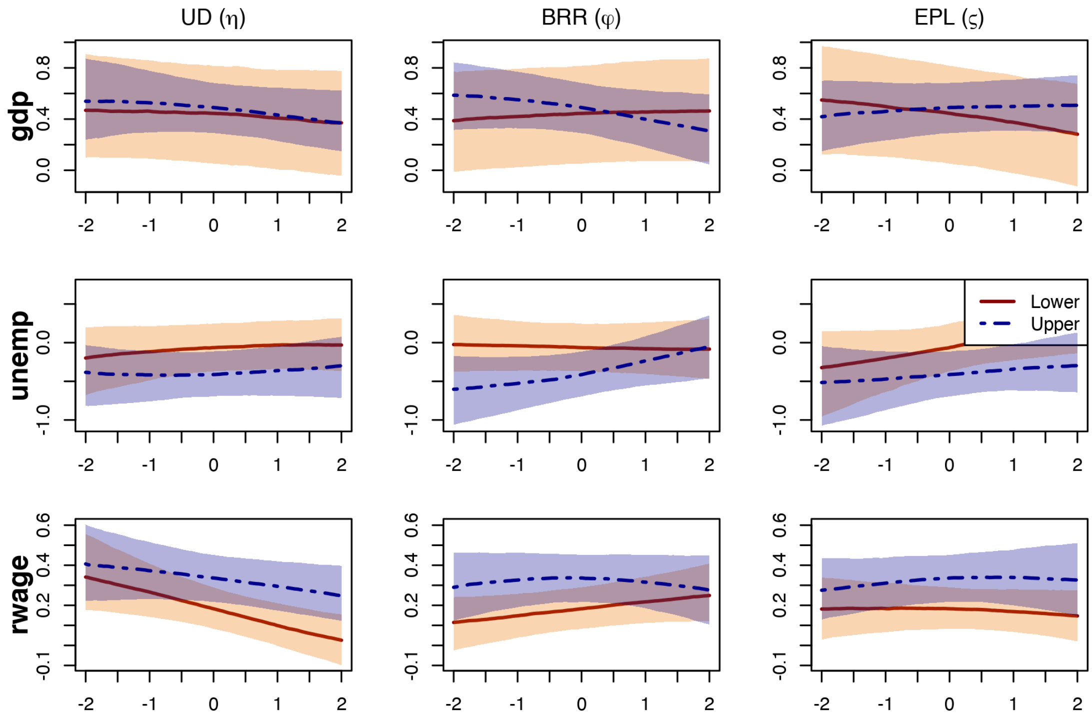
Figure D8: Classification of Countries



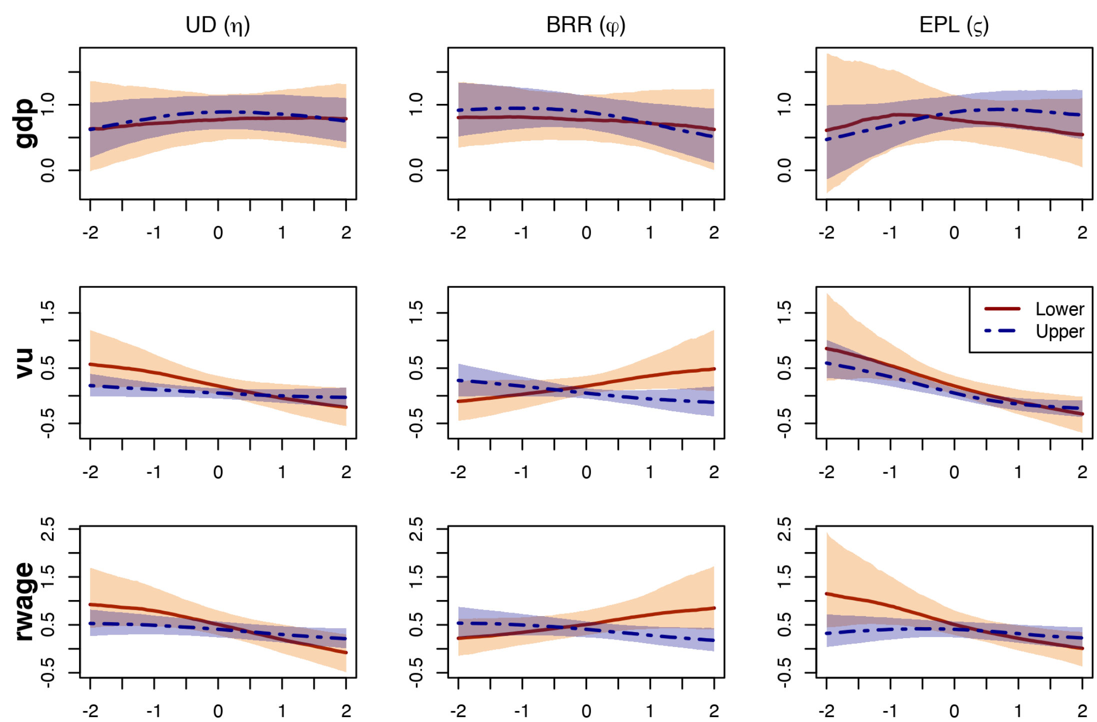
Notes: Each sub-plot shows the mean of each LMI for each country, together with two standard deviation in each direction. The points are observed data for the respective country. Color shadings differentiate countries belonging to the *Upper Group* (blue) and *Lower Group* (orange).

Figure D9: Fiscal Multipliers Across Groups in the Models with Other Variables.

(a) Unemployment.



(b) Labor Market Tightness.



Notes: The sub-plots show the sensitivity of the fiscal spending multipliers 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 size of the multiplier while the x-axis runs from -/+ 2 standard deviations in terms of the respective LMI (within-country variation). The multipliers are shown for a horizon of four quarters and for the *Lower* and *Upper* group (between-country variation) Confidence bounds refer to the 16/84 quantile of the posterior distribution.