Labor Market Institutions, Fiscal Multipliers, and Macroeconomic Volatility

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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.

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

JEL Codes: E62, C33, J21, J38

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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 Jung and Kuester, 2015; Banerji et al., 2017; Sondermann, 2018; Masuch et al., 2018; Duval and Furceri, 2018; Aiyar et al., 2019, for instance). Several recent studies (see Zanetti, 2009; Zanetti, 2011a; 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, Mendoza and Végh, 2013; Leeper, Traum and Walker, 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 qualitatively 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. In a search-and-matching framework with labor market tightness[1], a positive change in the unemployment benefit replacement rates and union density crucially affect the wages positively, attenuating the expansion of employment following an expansionary fiscal shock. Employment protection, however, affects directly job creation and destruction in opposing directions. Given our calibration, higher employment protection leads to an attenuated response of the fiscal multiplier on employment.

In a next step, we confront the predictions of the theoretical model by estimating an empirical model without imposing much structure on the data. In particular, we estimate a Bayesian interacted 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 allows 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 an additional exercise, we inspect the cross-country heterogeneity. We abstain from estimating the theoretical model directly due to the availability of LMIs only on a lower frequency within a limited time sample. This 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, Mendoza and Végh (2013).

The results can be summarized as follows. We use our theoretical model to inspect qualitatively the outcomes of the empirical model. The theoretical model predicts that more stringent LMIs attenuate both fiscal spending multipliers and output volatility. However, the latter depends on the source of the shocks under consideration. 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

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

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.

In a last step, we check whether the theoretical model can account for the evidence presented. Specifically, we employ impulse response matching to investigate the underlying mechanism driving the results. The autocorrelation coefficient of the government spending shock and the share of non-Ricardians is similar across regimes. We find, however, noteworthy variation in the parameters governing real wage and price rigidity.

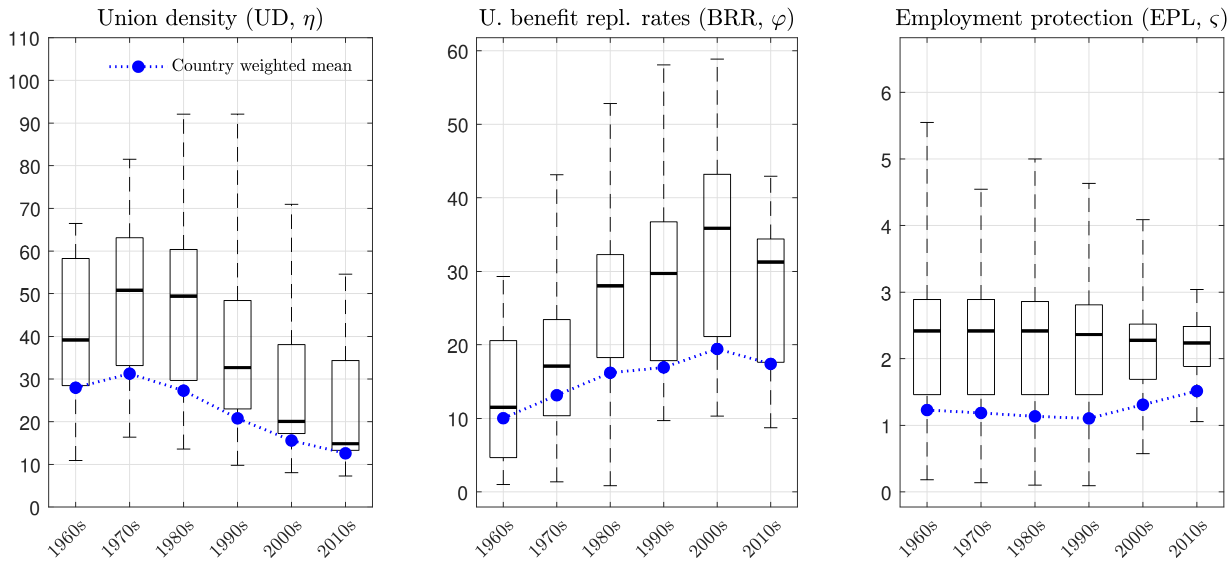
The remainder of the paper is structured as follows. The next session discusses the related literature. In Section 3 we provide a descriptive overview of LMIs across selected OECD countries. Section 4 introduces the theoretical model and presents its main predictions. Section 5 discusses the connection between the theoretical and the empirical model, and presents the results of the econometric analysis. Finally, Section 6 concludes.

2. Related Literature

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, Meier and Müller, 2012; Ilzetzki, Mendoza and Végh, 2013) and specifically to the strand that evaluates their variation over time (Auerbach and Gorodnichenko, 2012; Cos and Moral-Benito, 2016). Existing results indicate that the fiscal multiplier is higher during periods of financial turmoil (Bernardini, De Schryder and Peersman, 2020), when household leverage is higher (Demyanyk, Loutskina and Murphy, 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 (Christiano, Eichenbaum and Rebelo, 2011; Coenen et al., 2012; Fernández-Villaverde et al., 2015; Rendahl, 2016). Furthermore, Shen and Yang (2018) show that spending multipliers become countercyclical under downward nominal wage rigidity. In a next step, Jo and Zubairy (2022) show that in a New Keynesian model with nominal downward wage rigidity that spending multiplier differ markedly in demand-driven recessions than in supply-driven recessions. 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, Manovskii and Mitman, 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, Perotti and Trigari, 2010; Brückner and Pappa, 2012; Turrini, 2013). In this context, Ball, Jalles and Loungani (2015) highlight that the link between GDP and the labor market strongly depends on the idiosyncratic labor market institutions in place in a given economy.

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) and Zanetti (2011a) or Cacciatore et al. (2016) assess theoretically the macroeconomic effects of market reforms featuring the removal of labor and product market frictions. A rather small number of studies focuses also on the short-run effects of labor market institutions, such as price stability (Thomas and Zanetti, 2009), the cyclical behavior of vacancies and unemployment (Zanetti, 2011b) asymmetries in the business cycle (Ferraro and Fiori, 2023). An increasing number of empirical studies estimate the aggregate impact of changes in regulation (Pérez and Yao, 2015; Ordine and Rose, 2016; Duval, Furceri and Jalles, 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. Gnocchi, Lagerborg and Pappa (2015) show that labor market institutions matter for business cycle fluctuations in a large panel of countries by exploiting data on specific labor market reform episodes. Cacciatore et al. (2021), Abbritti and Weber (2018) and Hantzsche, Savsek and Weber (2018) count among the few contributions that explicitly acknowledge the role played by labor market institutions in the macroeconomic responses to exogenous shocks. Our contribution is most closely related to the studies of Cacciatore et al. (2016) and Abbritti and Weber (2018). Cacciatore et al. (2021) analyze the role of employment protection for fiscal spending shocks. To this end, our contribution corroborates the findings of Cacciatore et al. (2016) by considering a larger set of LMIs. Furthermore, our approach allows to examine the effects of LMIs in a joint framework. Abbritti and Weber (2018), however, study the LMIs' ability to mitigate business cycle shocks (oil price shock and global demand shock) without paying attention to fiscal multipliers and the extent to which they are shaped by the LMIs. Moreover, our analysis of the LMIs role for shaping macroeconomic volatility comprises an extension to both studies. Lastly, the present study highlights the link between fiscal multipliers and macroeconomic volatility and

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.

the associated trade-off between structural and cyclical oriented policies. Hantzsche, Savsek and Weber (2018), on the other hand, analyze the interaction between labor market institutions and contractionary financing shocks.

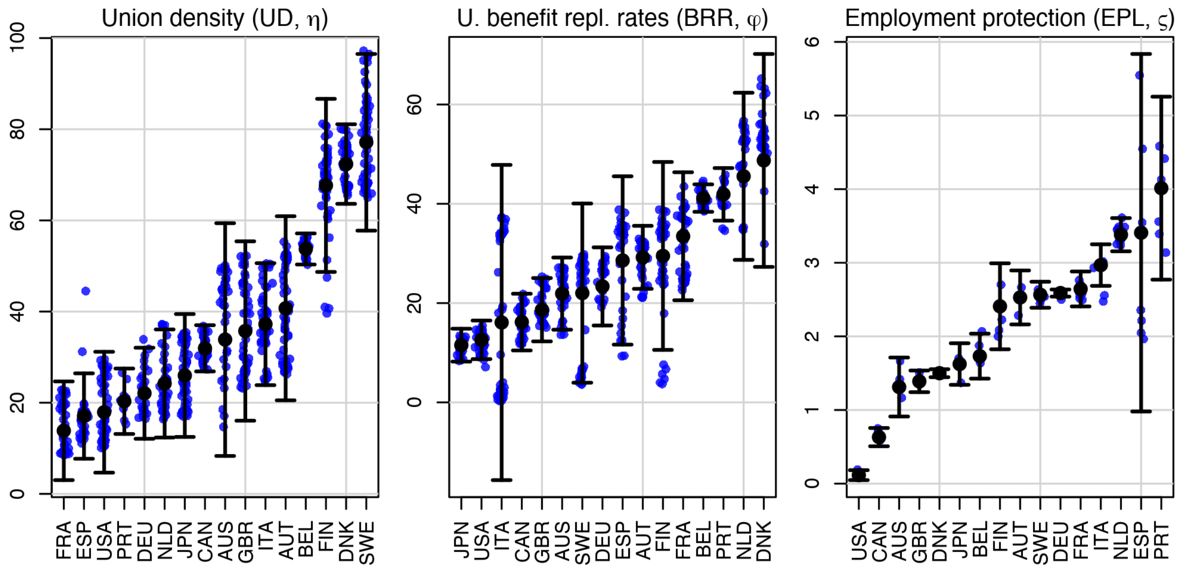
3. Structural Labor Market Indicators in OECD Economies

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).[2]

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

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

Figure 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.

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.

Figure 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.

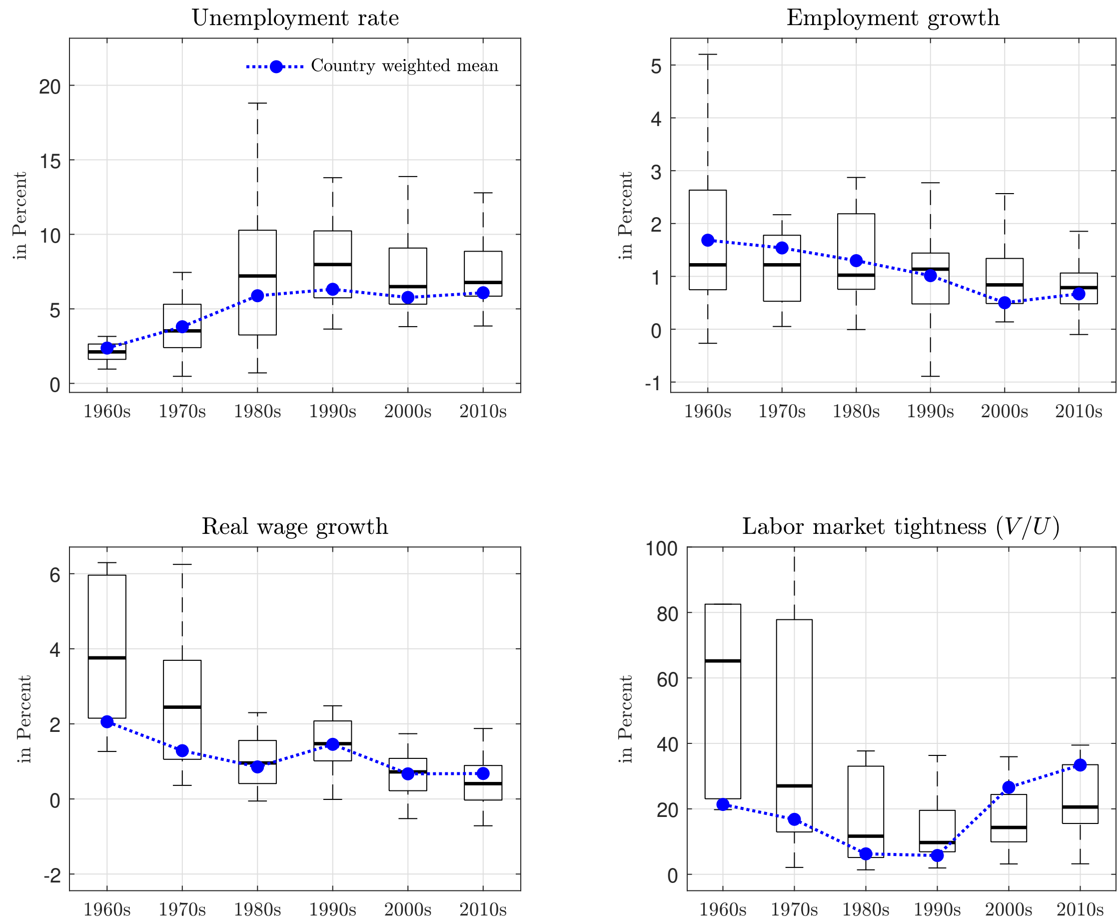
In Figure 2, we examine the cross-country variation of the three indicators of LMIs. The LMI indicators exhibit strong country heterogeneity, most notably for UD and BRR and less so for EPL. In particular, Scandinavian and continental European countries present the highest levels of UD and BRR. In countries 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 Figure 3. The growth rates of employment and real wages[3] 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.

The descriptive analysis highlights that increases in the BRR and the EPL align with lower employment growth and less variation across decades. The reductions in UD, in turn, align with

[3] 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.

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.

4. The Theoretical Model

In the theoretical model, we merge the structure of a Diamond-Mortensen-Pissarides model with a standard real business cycle framework and rely on the setting put forward by Merz (1995), Andolfatto (1996) and Krause and Lubik (2007), and Monacelli, Perotti and Trigari (2010). The model is intended to be parsimonious and focus on the role played by labor market institutions. We consider various extensions of the set-up that can accommodate more complex interactions in subsection 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: $\varrho(\tilde{a}_t) = \bar{\varrho} + (1 - \bar{\varrho})F(\tilde{a}_t)$.

4.1 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

$$n_t = (1 - \varrho(\tilde{a}_t))(n_{t-1} + m_{t-1}). \tag{4.1}$$

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[4]. 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 Equation 4.1.

[4] We consider the case where firing costs accrue to the government in Appendix A.4.

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 conditions give rise to[5]

$$F_t^n = mpl_t - w_t + E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right] \tag{4.2}$$

$$\frac{\kappa}{q_t} = E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right] \tag{4.3}$$

$$A(\tilde{a}_t) = \frac{1}{\bar{A}}\left(w_t - b_t^s - \frac{\kappa}{q_t}\right) \tag{4.4}$$

where mpl_t is the marginal product of labor and F_t^n is the Lagrange multiplier associated with Equation 4.1. In Equation 4.2, 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. Equation 4.3 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, Equation 4.4 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.

4.2 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

$$c_t + B_t = R_{t-1}B_{t-1} + (1-\tau)w_t n_t + b_t^u(1-n_t) + T_t^S + \pi_t^F, \tag{4.5}$$

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

[5] 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 Equation 4.3, Equation 4.2, and Equation 4.1, 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 Equation 4.4.

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 to

$$n_t = (1-\varrho(\tilde{a}_t))n_{t-1} + p_t(1-n_{t-1}). \tag{4.6}$$

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, Equation 4.5 and the employment flow Equation 4.6. Optimization leads to the following conditions

$$1 = R_t E_t \Lambda_{t,t+1}, \tag{4.7}$$

$$H_t^n = \tilde{w}_t^b - mrs_t + E_t\left[1-\varrho(\tilde{a}_{t+1}) - p_{t+1}\right]\Lambda_{t,t+1}H_{t+1}^n, \tag{4.8}$$

where λ_t is the Lagrange multiplier attached to Equation 4.5 and $\lambda_t H_t^n$ the one attached to equation Equation 4.6. 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 Equation 4.8, 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.

4.3 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

$$w_t = (1-\eta)\frac{mrs_t + b_t^u}{1-\tau} + \eta\left(mpl_t + E_t\Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right]\right). \tag{4.9}$$

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.

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

The government budget constraint satisfies

$$\tau w_t n_t + B_t = R_{t-1}B_{t-1} + b_t^u u_t + T_t^s + g_t, \tag{4.10}$$

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 Equation 4.10, Equation 4.5, and the expression for firms' profits (π_t^F), we obtain the aggregate resource constraint

$$y_t = c_t + g_t + \kappa v_t + F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1}v_{t-1})b_t^s \tag{4.11}$$

This equation closes the model.

4.5 Embedding LMIs in the Model

The indicators for the LMIs as discussed in Section 3 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, Noel and Vavra, 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).

4.6 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-linearized solution of the rational expectations model around its steady state,

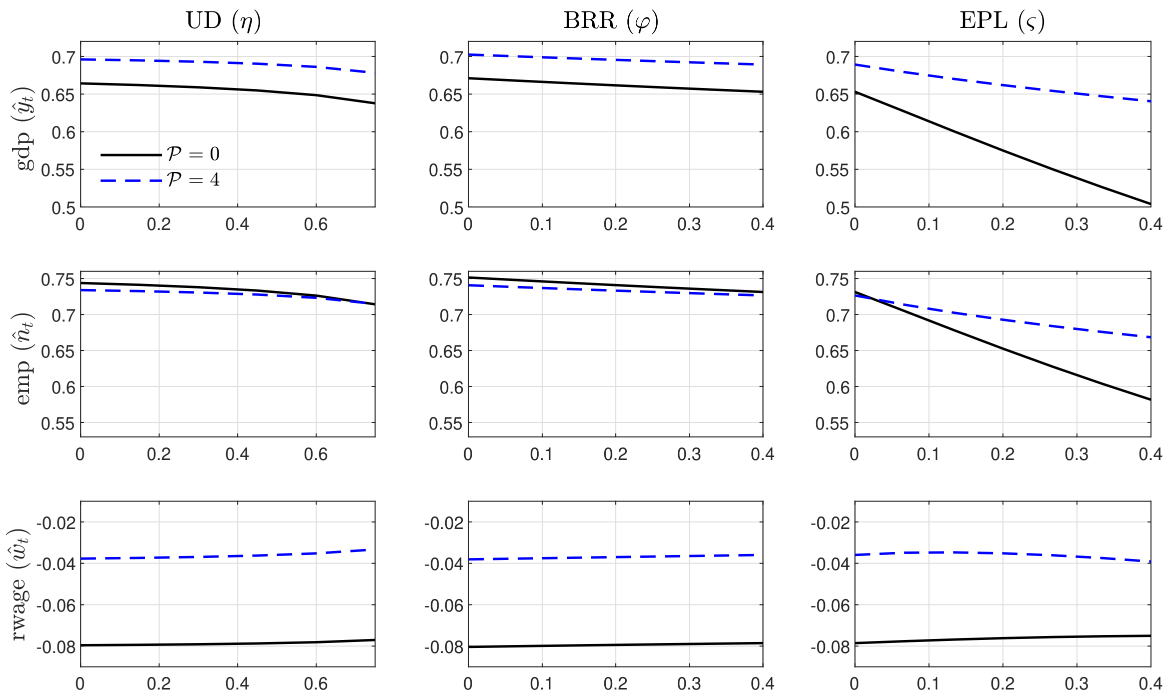
$$z_t = \boldsymbol{\Psi}_1(\boldsymbol{\vartheta})z_{t-1} + \boldsymbol{\Psi}_0(\boldsymbol{\vartheta})\boldsymbol{\varepsilon}_t \tag{4.12}$$

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. Equation 4.12 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 by considering the following definition of the fiscal spending multiplier for some variable x

$$\mu_x(\vartheta_l) = \frac{\sum_{i=1}^{\mathcal{P}} \text{IRF}_i^x(\vartheta_l)}{\sum_{j=1}^{\mathcal{P}} \text{IRF}_j^{\hat{g}}(\vartheta_l)} \quad \forall\, l = \{1, 2, 3\} \tag{4.13}$$

where $\text{IRF}_t^x(\vartheta_l)$ and $\text{IRF}_t^{\hat{g}}(\vartheta_l)$ 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 $\mathcal{P}$.

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).

The results are shown in Figure 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, 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.

4.7 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 Figure 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 Equation 4.2 and Equation 4.3, and setting τ equal to zero for simplicity, the reservation wages are given by

$$\overline{w}_t^F = mpl_t + E_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], \tag{4.14}$$

$$\underline{w}_t^H = mrs_t + b_t^u - (1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}) E_t \Lambda_{t,t+1} H_{t+1}^n, \tag{4.15}$$

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[6] in response to the expansionary fiscal spending shock.

[6] 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.

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 (4.9), $\partial w_t / \partial \eta = \overline{w}_t^F - \underline{w}_t^H$ (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 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 Figure 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 (4.4) and (4.2). 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 (4.4) 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 Figure 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 Equation 4.12 and compute the variance of the endogenous variables in z_t and extend the vector of structural shocks (ϵ_t) by a technology shock (ϵ_t^A) so that the set of exogenous shocks involves both a demand and a supply shock. We carry out this simulation to provide an answer to the following: How do the LMIs shape macroeconomic volatility? For the sake of brevity, we describe the analysis in detail in Section A.5 of the Appendix and limit the discussion here to the most important implications. We compute the variance of the vector of endogenous variables once when a low value of a LMI is considered and compare it to the variance once a high value is used. The results are provided in Table 1. For instance, in case of employment, a high value of the EPL gives rise to a lower employment volatility; this is indicated by the value of 1.16 in the last column of the second row. More generally, we find that a higher BRR and EPL attenuate output volatility, while the opposite emerges from a higher UD. Employment volatility is mitigated by a higher UD and EPL while it gets

Table 1: Volatility of output, employment and the real wage.

LMI:	UD (η)	BRR (φ)	EPL (ς)
Output ($\hat{y}_t$)	0.97	1.07	1.12
Employment ($\hat{n}_t$)	1.04	0.95	1.16
Real wage ($\hat{w}_t$)	0.99	0.96	0.99

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

exacerbated by a higher BRR. The volatility of the real wage in turn is hardly affected by the UD and EPL while gets attenuated by a higher BRR. Across all LMIs, the EPL tends to exert the strongest effects on the volatilities of output and employment while the real wage is primarily affected by the BRR. All these results, however, crucially depend on the source of the shock. We provide a deeper analysis in the Appendix where we examine the sensitivity of the effects of changes in the LMIs with respect to distinct exogenous shocks. The key finding of this is that the LMIs can potentially mitigate output volatility, however, the nature and dominance of specific shocks is crucial.

4.8 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. Section A.4 considers 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.

5. 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, Towbin and Wieladek (2014). The IP-VAR model is employed to assess how the characteristics of the matrices $\mathbf{\Psi}_0(\boldsymbol{\vartheta})$ and $\mathbf{\Psi}_1(\boldsymbol{\vartheta})$ of the system given by Equation 4.12 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 $\bar{\boldsymbol{\vartheta}}$

$$\mathbf{\Psi}_j(\boldsymbol{\vartheta}) \approx \mathbf{\Psi}_j(\bar{\boldsymbol{\vartheta}}) + \sum_{l=1}^{3} \left[\frac{\partial \mathbf{\Psi}_j}{\partial \vartheta_l}(\bar{\boldsymbol{\vartheta}})(\vartheta_l - \bar{\vartheta}_l) \right], \quad j \in \{0, 1\}. \tag{5.1}$$

Substituting the matrices $\mathbf{\Psi}_0(\boldsymbol{\vartheta})$ and $\mathbf{\Psi}_1(\boldsymbol{\vartheta})$ in Equation 4.13 by the Taylor approximation given by Equation 5.1 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 $\mathbf{\Psi}_j(\boldsymbol{\vartheta})$, before presenting the results and providing a discussion of the insights gained from the estimation of the econometric model.

5.1 Econometric Model

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

$$\mathbf{y}_{it} = \boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) + \sum_{j=1}^{p} \mathbf{\Phi}_{ij}(\boldsymbol{\vartheta}_{it})\mathbf{y}_{it-j} + \mathbf{u}_{it}, \quad \mathbf{u}_{it} \sim \mathcal{N}_M(\mathbf{0}, \mathbf{\Sigma}_i(\boldsymbol{\vartheta}_{it})), \tag{5.2}$$

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 $\mathbf{\Phi}_{ij}$ for lag j, and a variance-covariance matrix of the vector error term, given by $\mathbf{\Sigma}_i$. 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 Appendix 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 $\mathbf{\Sigma}_i$ 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 Equation 4.12 is given by

$$\boldsymbol{y}_{it} = \sum_{j=1}^{p} \tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it})\boldsymbol{y}_{it-j} + \boldsymbol{e}_{it}, \quad \boldsymbol{e}_{it} \sim \mathcal{N}_M(\boldsymbol{0}, \boldsymbol{I}), \tag{5.3}$$

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 Appendix B. The prior assumes that the structural individual-level coefficients have a common underlying Gaussian distribution,

$$\tilde{\boldsymbol{\Psi}}_{ij}(\boldsymbol{\vartheta}_{it}) \sim \mathcal{N}(\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}), \boldsymbol{V}_j), \quad j = 1, \ldots, p, \tag{5.4}$$

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 Appendix B. The correspondence between the observable LMIs $\boldsymbol{\vartheta}_{it}$ (depicted in Figure 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

$$\boldsymbol{\Psi}_j(\boldsymbol{\vartheta}) = \boldsymbol{\Psi}_{jt} = \bar{\boldsymbol{\Psi}}_j + \sum_{l=1}^{d} \boldsymbol{\Gamma}_{jl}^{\boldsymbol{\Psi}} \vartheta_{lt}, \quad j = 0, ..., p. \tag{5.5}$$

which relates the empirical set-up directly to Equation 5.1 of the theoretical model. The full IP-VAR model is given by Equation 5.3, and its equivalence with the solution of the DSGE model depicted in Equation 4.12 and Equation 5.1 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 Figure 1, structural rather than cyclical factors appear to determine their dynamics.[7] A second potential limitation is the linearity assumption (in the parameters) embedded in the IP-VAR model, which mimics the approximation considered in Equation 5.1. In principle, the assumption of

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

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.

5.2 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 analyze the effects of fiscal policy (Beetsma and Giuliodori, 2011; Bénétrix and Lane, 2013; Ilzetzki, Mendoza and Végh, 2013; Huidrom et al., 2020, to mention a few). Second, alternative identification approaches are infeasible in the context of our research question. In particular, event-study approaches based on defense 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, Leigh and Pescatori, 2014) requires the availability of detailed legislative records in order to extract policy shocks.[8] 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. Legal processes usually create a time gap between the announcement and the

[8] 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.

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.

5.3 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 Table C1 and Table 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})[9]; see Appendix 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 Table 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. 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 Figure 2). We investigate cross-country heterogeneity by splitting the sample of countries in two distinct groups in subsection D.3.

[9] 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.

Given the standardization of the interaction variables, the interpretation is as follows: A unit rise in ϑ_l corresponds to a one standard 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.

5.4 The Effect on Fiscal Spending Effectiveness

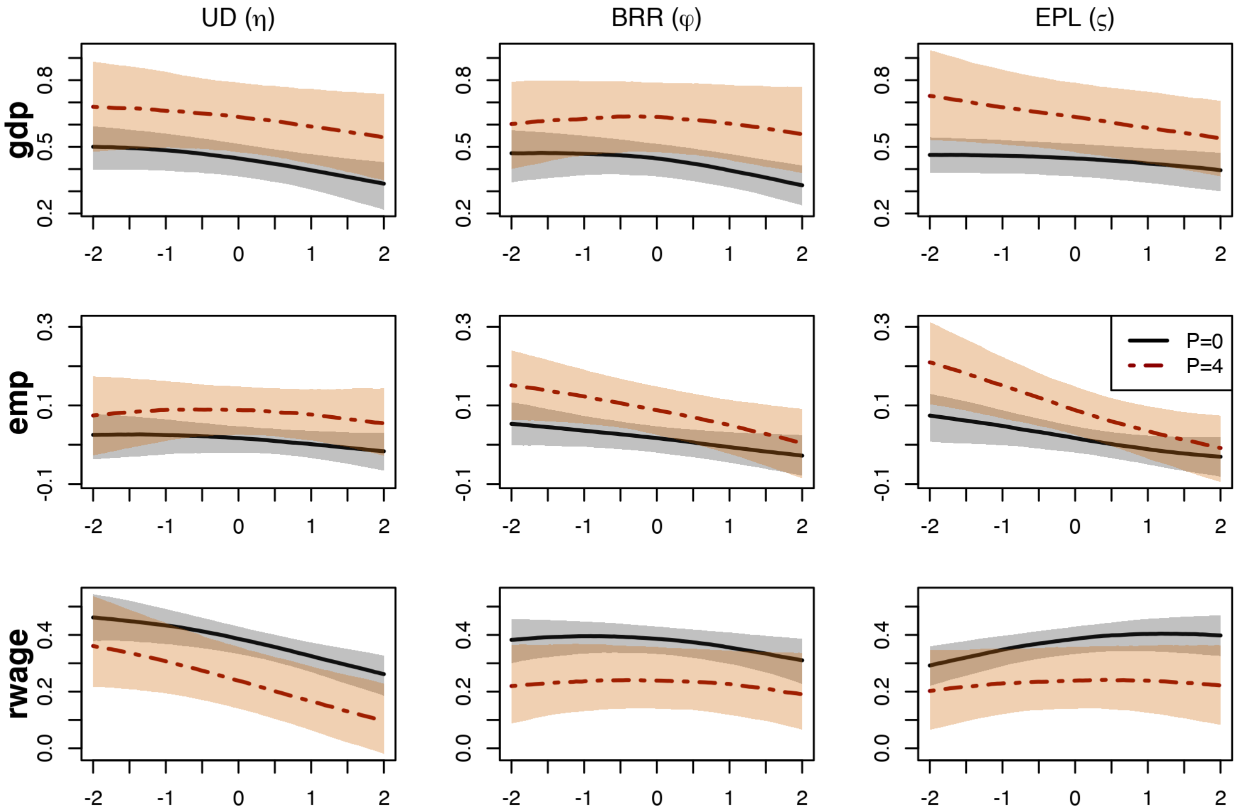
In this section, we present the effects of the LMIs on fiscal spending effectiveness. In line with the theoretical results, we use Equation 4.13 to compute multipliers for the initial impact ("impact multiplier", horizon $P = 0$) and for the effect after four quarters ("one-year multiplier", horizon $P = 4$). Figure 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 20 log points (from 0.25 down to 0.05), while being a 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 output multiplier of the IP-VAR model is comparable to those of the calibrated DSGE model. Similar to Figure 4, the strongest decrease is visible for EPL. Fiscal multipliers with respect to UD and BRR decrease to a lesser degree. Nevertheless, also some differences arise.

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.

Specifically, our baseline model shows no strong change in the multipliers with respect to real wages. Here, we stress the role of nominal rigidities (see Section A.4 in the Appendix). The one-year employment multiplier consistently exceeds the impact multiplier, which highlights the role of limited asset market participation in this context (see Section A.4 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 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.

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

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.

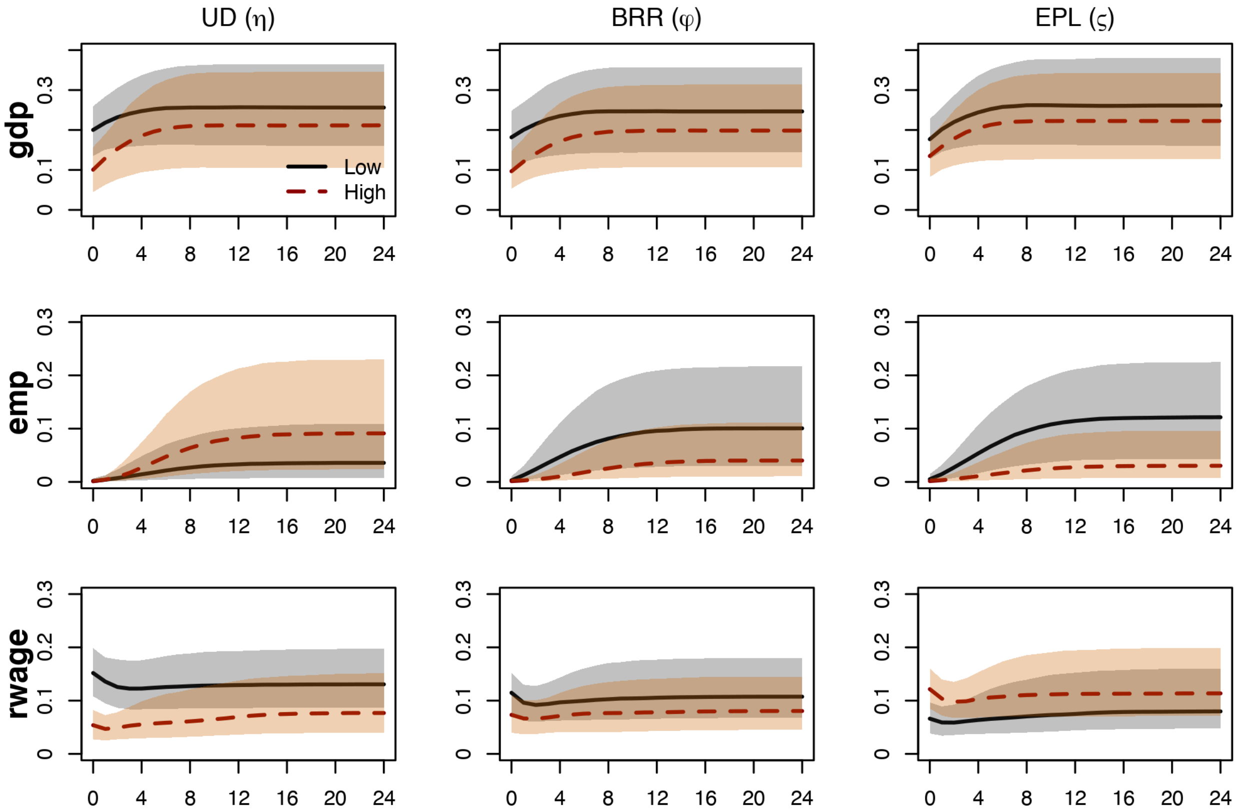
These findings are in line with the literature as regards the size of fiscal spending multipliers for output (Ramey, 2019), the extent of inertia (Ilzetzki, Mendoza and Végh, 2013), as well as the lower value of the employment relative to the output multiplier (Monacelli, Perotti and Trigari, 2010). 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 fiscal multipliers with respect to the LMIs. In this regard we find strong evidence in favor of a dependency of fiscal multipliers and hence of the effectiveness of discretionary fiscal policy on the LMIs.

Further results on the extent to which the LMIs shape the transmission channel of fiscal spending shocks are provided in Appendix 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.

We extend the previous analysis for the forecast error variance decomposition (FEVD). The results are provided in Figure 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 Figure 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.

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.

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.

5.5 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, 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[10] 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

$$\text{vec}\left(\boldsymbol{\Omega}(\boldsymbol{\vartheta})\right) = \left(\boldsymbol{I} - \boldsymbol{F}(\boldsymbol{\vartheta}) \otimes \boldsymbol{F}(\boldsymbol{\vartheta})\right)^{-1} \text{vec}\left(\boldsymbol{Q}(\boldsymbol{\vartheta})\right) \tag{5.6}$$

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 $\overline{\boldsymbol{\Sigma}}(\boldsymbol{\vartheta}) = N^{-1} \sum_{i=1}^{N} \boldsymbol{\Sigma}_i(\boldsymbol{\vartheta})$.[11] As can be seen from Equation 5.6, 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 Figure 7. We focus on output and employment only and provide further details in the Appendix. In Figure 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

[10] The results of the DSGE model of this exercise are presented in Section A.5 in the Appendix.

[11] 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.

Figure 7: Macroeconomic Volatilities Along LMIs.

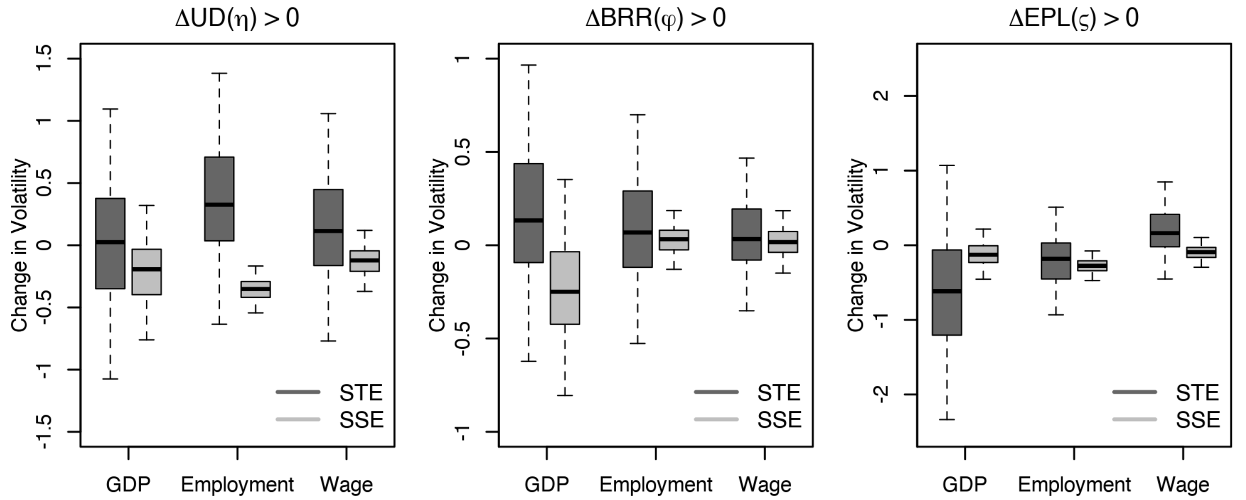
Notes: Each sub-plot shows the standard deviations of the respective macroeconomic variable in a regime with low (−2sd) and high (+2sd) LMIs. The LMIs under consideration are union density (UD(η)), unemployment benefit replacement rate (BRR(φ)), and employment protection (EPL(ς)).

LMIs.[12] 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 in the volatilities of output and employment of around 25% (with a probability of 85%) and 30% (with a probability of 93%), respectively. Again, these results align well with the theoretical predictions as shown in Table 1. Our theoretical model also suggests that EPL can mitigate macroeconomic volatility to the largest extent in terms of output and employment. 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, as we discuss in more detail in A.5. To conclude, EPL reduces volatility more pronounced 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 mechan-

[12] 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 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 Equation 5.7. The LMIs under consideration are union density (UD(η)), unemployment benefit replacement rate (BRR(φ)), and employment protection (EPL(ς)).

ism (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 Equation 5.6, 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 $\tilde{\boldsymbol{Q}}(\boldsymbol{\vartheta}) \equiv \text{vec}\,(\boldsymbol{Q}(\boldsymbol{\vartheta}))$, 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 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

$$\frac{\partial \omega_{kk}(\boldsymbol{\vartheta})}{\partial \vartheta_l} = \sum_j \Bigg(\underbrace{\tilde{q}_j(\boldsymbol{\vartheta}) \frac{\partial \tilde{f}_{kj}(\boldsymbol{\vartheta})}{\partial \vartheta_l}}_{\text{STE}} + \underbrace{\tilde{f}_{kj}(\boldsymbol{\vartheta}) \frac{\partial \tilde{q}_j(\boldsymbol{\vartheta})}{\partial \vartheta_l}}_{\text{SSE}} \Bigg), \tag{5.7}$$

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 Figure 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 Figure 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 Figure 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 Figure D6 and Figure 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.

5.6 *Accounting for the Evidence*

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

The model version used for this exercise considers the various extensions put forth in the Appendix (see Section A.4), involving nominal price and real wage rigidities and limited asset market participation. We fix all structural parameters to their values as depicted in Table A1, except the parameters capturing the persistence of the government spending shock (ϱ_g) the real wage inertia (ϱ_w), the extent of price stickiness (κ) and the share of non-Ricardian consumers λ which are collected in the vector $\boldsymbol{\theta}$. The parameter selection is motivated by their significant role in shaping fiscal spending multipliers. In particular, Galí, López-Salido and Valles (2007) emphasize the role of non-Ricardian agents next to nominal rigidities, Blanchard and Galí (2007) stress the role of real wage rigidities and finally, Dupaigne and Fève (2016) highlight the role of the persistence of government spending in shaping government spending multipliers.

In order to obtain estimates for these parameters, we match empirical (IP-VAR) and theoretical impulse responses (see, for instance, Rotemberg and Woodford, 1997). Let $\widehat{\text{IRF}}$ be the empirical impulse response functions obtained from estimating the IP-VAR[13] , and let $\text{IRF}(\boldsymbol{\theta})$ be its counterpart from the theoretical model. We focus on the first 25 periods of the responses of government spending ($\hat{g}_t$), output ($\hat{y}_t$), employment ($\hat{n}_t$), and the real wage ($\hat{w}_t$). We estimate the parameter vector $\boldsymbol{\theta}$ by minimizing the weighted distance[14] between empirical and theoretical impulse response functions under low and high values of the LMIs

$$\hat{\boldsymbol{\theta}} = \arg\min \left\| \widehat{\text{IRF}} - \text{IRF}(\boldsymbol{\theta}) \right\|, \tag{5.8}$$

where $\widehat{\text{IRF}}$ and $\text{IRF}(\boldsymbol{\theta})$ are column vectors[15] of impulse responses. We establish estimates for $\hat{\boldsymbol{\theta}}$ for (i) each LMI individually, and (ii) separately for low and high values. The parameter estimates are provided in Table 2.

The estimated autocorrelation coefficient for the government spending shock reflects a high persistence and a value commonly found in the literature (see, for instance, Born, Juessen and Müller, 2013). Most interestingly, the estimates are the same across labor market regimes (that is,

[13] Specifically, as always done throughout the paper, we focus on the impulse response functions of the *common mean* in the model. The *common mean* is an estimate, which encompasses all the information from the respective single-country models but still allows for idiosyncrasies. We report the respective impulse response functions in Figures D1, D2, and D3 in the Appendix.

[14] We stick to a procedure that produces parameter estimates which give rise to equilibrium determinacy (i.e. saddle-path stability).

[15] Their dimension is $(25 \cdot 4) \times 1$ in each case: 25 periods of the impulse response functions for four variables.

Table 2: Estimated model parameters

	UD (η)		BRR (φ)		EPL (ς)	
	Low	High	Low	High	Low	High
ϱ_g	0.65	0.67	0.66	0.65	0.65	0.67
ϱ_w	**0.74**	**0.64**	**0.65**	**0.78**	0.70	0.73
κ	**0.37**	**0.45**	0.40	0.36	**0.51**	**0.26**
λ	0.03	0.05	0.04	0.06	0.04	0.06

Notes: As regards the structural parameters, ϱ_g is the autoregressive of the AR(1)-government consumption spending shock, ϱ_w captures the degree of real wage inertia (see Section A.4), κ captures the degree of price stickiness (see Section A.4, and λ measures the share of non-Ricardian consumers A.4.

whether a high or a low value of any LMI is considered). Moreover, we also observe that the share of non-Ricardian households is unaffected by distinct labor market regimes. Our estimated values are at the lower edge of the estimates commonly found in the literature (Bilbiie, Meier and Müller, 2008), and hence imply modest financial frictions only.

We find a noteworthy variation of the remaining two parameter estimates across distinct labor market regimes. With respect to union density (UD), we observe a significant change in the estimates of ϱ_w and κ, with divergent effects on real wage and price rigidity. The estimates suggest that the transition to a labor market characterized by a higher union density renders real wages less sticky and causes less frequent price adjustments. As a result, nominal wage adjustments are likely to be more frequent, partly reflecting the increase in workers' wage-setting power. This contrasts with the effect of the benefit replacement rate (BRR), where a higher replacement rate increases real wage rigidity while leaving all the other parameter estimates largely unchanged. Finally, the degree of employment protection (EPL) primarily affects nominal frictions, with a minimal impact on real wage rigidity.

5.7 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 – intends 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. Our approach enables an assessment of distinct explanations in this respect. First of all, it permits to consider distinct observable structural elements for explaining 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 (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 (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.

6. 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. Using partial information matching, we find that the mechanism runs through variations in real wage and price rigidity depending on the LMI.

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.

Declaration of Interest

The authors declare to have no conflict of interest.

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A. Further Details on the Theoretical Model

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

A.1 Equilibrium Equations

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

Production

- $y_t = \bar{A} n_t A(\tilde{a}_t)$, with $A(\tilde{a}_t) = \int_{\tilde{a}_t}^{\infty} \frac{a}{1-F(\tilde{a}_t)} dF(a)$
- $\frac{\kappa}{q_t} = E_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right]$
- $F_t^n = mpl_t - w_t + E_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right]$
- $A(\tilde{a}_t) = \frac{1}{\bar{A}} \left(w_t - b_t^s - \frac{\kappa}{q_t} \right)$

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

Constraints and Policy

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

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: $S_t = F_t^n + H_t^n$ where $H_t^n = (1 - \tau) w_t - b_t^u - mrs_t + (1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}) E_t \left[\Lambda_{t,t+1} H_{t+1}^n \right]$.

A.2 Calibration and the Steady State

We compute the steady state for the purpose of simulating the model. Variables without a time subscript denote steady state values. We start by considering an ex-ante calibration of the probability of an unemployed person finding a job (p_t), the labor market tightness (θ_t), and the ratio between the marginal rate of substitution between consumption and labor on the side of the households and the marginal product of labor on the side of the firms ($\zeta_t = mrs_t/mpl_t$). Additionally, we calibrate the steady-state separation rate $\varrho(\tilde{a})$ and following the argument in den Haan, Ramey and Watson (2000), we also calibrate the exogenous job destruction rate $\bar{\varrho}$. The idiosyncratic productivity 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 (4.2), (4.3) and (4.9) 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 (4.4), we calibrate $\bar{\varsigma}$ such that $A(\tilde{a}) = (w - b^s - \kappa/q)/\bar{A}$.

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

Our benchmark calibration is summarized in Table A1. Given that our focus is on the role of the LMIs in the transmission of fiscal spending shocks, we do not calibrate our model to a particular economy. We closely follow Christoffel, Kuester and Linzert (2009) for the choice of the values

Table A1: Calibration of the Model.

Parameter	Description	Value	Range
α	Elasticity of labor in the production function	0.66	[0.5 – 1]
β	Discount factor	0.992	[0.95 – 0.999]
γ	Elasticity of matching of unemployed persons	0.68	[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	0.43	[0.05 – 0.95]
p	Probability of an unemployed person finding a job	0.30	[0.05 – 0.95]
τ	Labor tax rate	0	[0 – 0.5]
μ_a	Steady state mean of idiosyncratic productivity	0.0	[0 – 2]
σ_a	Steady state standard-deviation of idiosyncratic productivity	0.15	[0.05 – 3]
$\bar{\varrho}$	Exogenous job separation rate	0.03	[0.01 – 0.15]
$\varrho(\tilde{a})$	(Overall) Job separation rate	0.07	[0.03 – 0.3]
σ	Complementarity coefficient	1	[0.5 – 3]
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	—

of the structural parameters.[16] The complementarity coefficient σ in the households' instantaneous utility function $u(c, n)$ is set to 1, which corresponds to the separable utility case. We also need to calibrate the shock process, for which we assume that the logarithm of fiscal spending $\hat{g}_t$ follows an AR(1) process with auto-correlation equal to 0.85. We calibrate the standard deviations of the two shocks ($\text{std}(\epsilon_t^G) = 0.48$ and $\text{std}(\epsilon_t^A) = 0.39$) in line with Christoffel, Kuester and Linzert (2009).

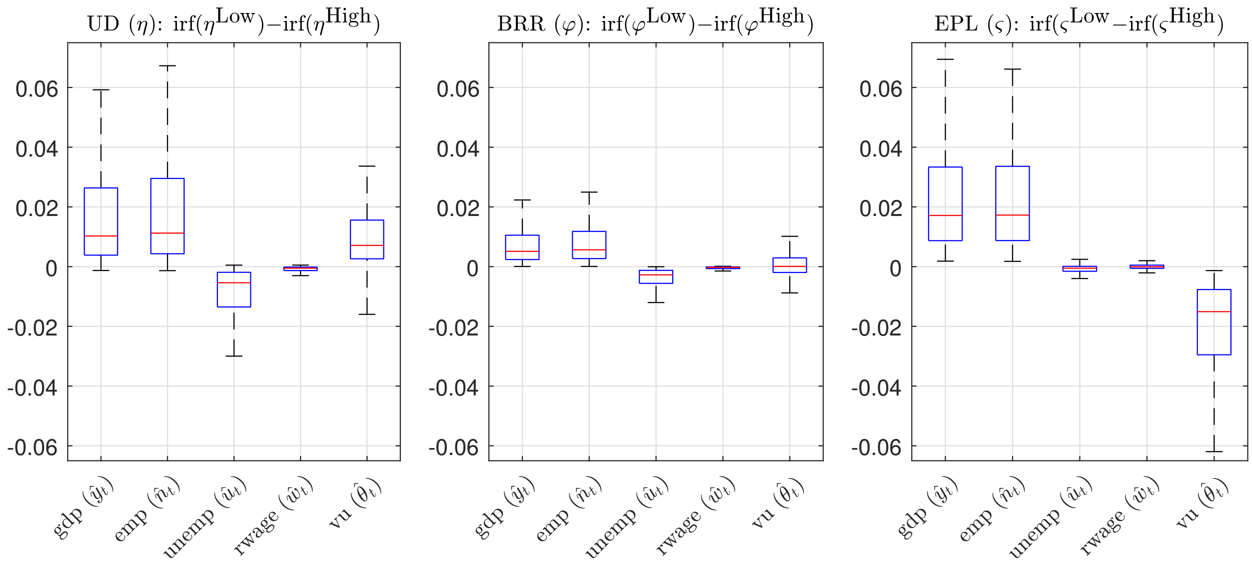
A.3 A Quantitative Evaluation Based on a more General Calibration

While the purpose of this exercise is to highlight the general effects of the LMIs on fiscal spending multipliers, the results presented in Section 4.7 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 Table A1 provides the details. We simulate the model over a wide range of different values for the parameters. To this purpose we attach a uniform distribution to each parameter and define upper and lower bounds as indicated in the fourth column (*Range*) in Table A1. We simulate the model 2000 times and compute the difference of the impulse response functions for the following two scenarios: low value of ϑ_i versus high value of ϑ_i where ϑ_i refers to one of the three parameters of interest (η, φ and ς). We focus on the impact responses. The three scenarios (UD, η; BRR, φ; and EPL, ς) are

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

Figure A1: Fiscal Spending Shocks and the LMIs.



depicted in the sub-panels in Figure A1.[17] 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 Figure 4 and Figure 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 creation and job destruction activities by firms on the one hand, and labor supply decisions of households on the other hand.

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

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 4.7. 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

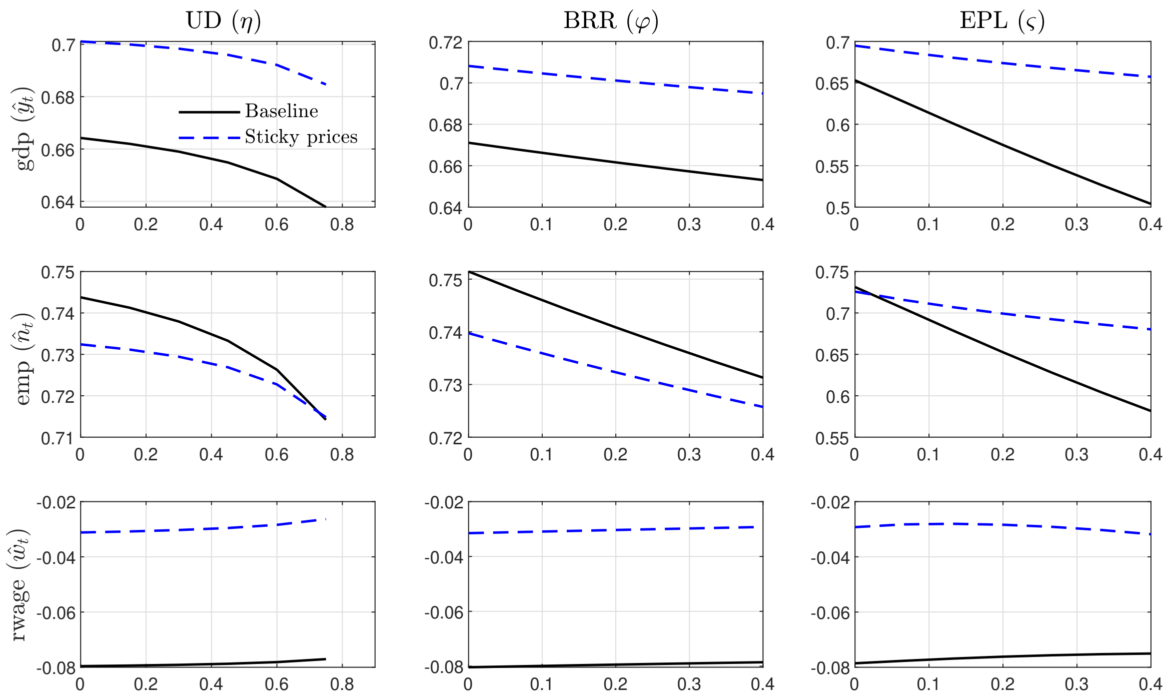
This section considers various extensions to the baseline model outlined in Section 4. These include price stickiness, real wage rigidity, limited asset market participation of one group of households and the case when firing costs accrue to the government as revenues. We always consider one extension at a time, as otherwise the precise role of the additional frictions considered becomes difficult to assess.

Markup Pricing and Monopolistic Competition. While a neo-classical set-up, as considered in Section 4, 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 4 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, López-Salido and Valles (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 4 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

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.

them to the households. Prices are adjusted according to a conventional Calvo approach by retailers where $1 - \kappa$ denotes the 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 (4.9)

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

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 $\varphi_\pi = 1.5$. Figure 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 5.

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, 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.

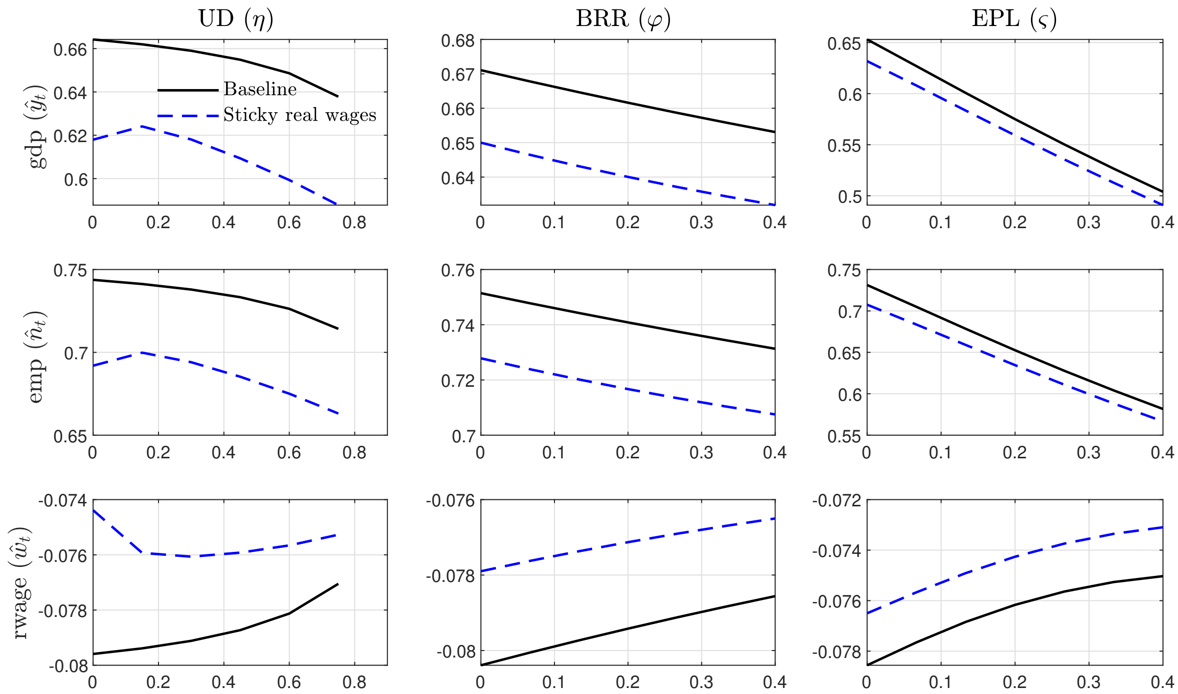
Despite a noteworthy effect of price stickiness on the size of the multiplier for output and other variables, the implications of the three LMIs for the multiplier as outlined in Section 4.7 still apply. This also holds for the multiplier of the real wage rate even though the sign has changed. It continues to be the case that higher values of the LMIs attenuate the multipliers.

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 4 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.

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 (4.9) to the following

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

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.

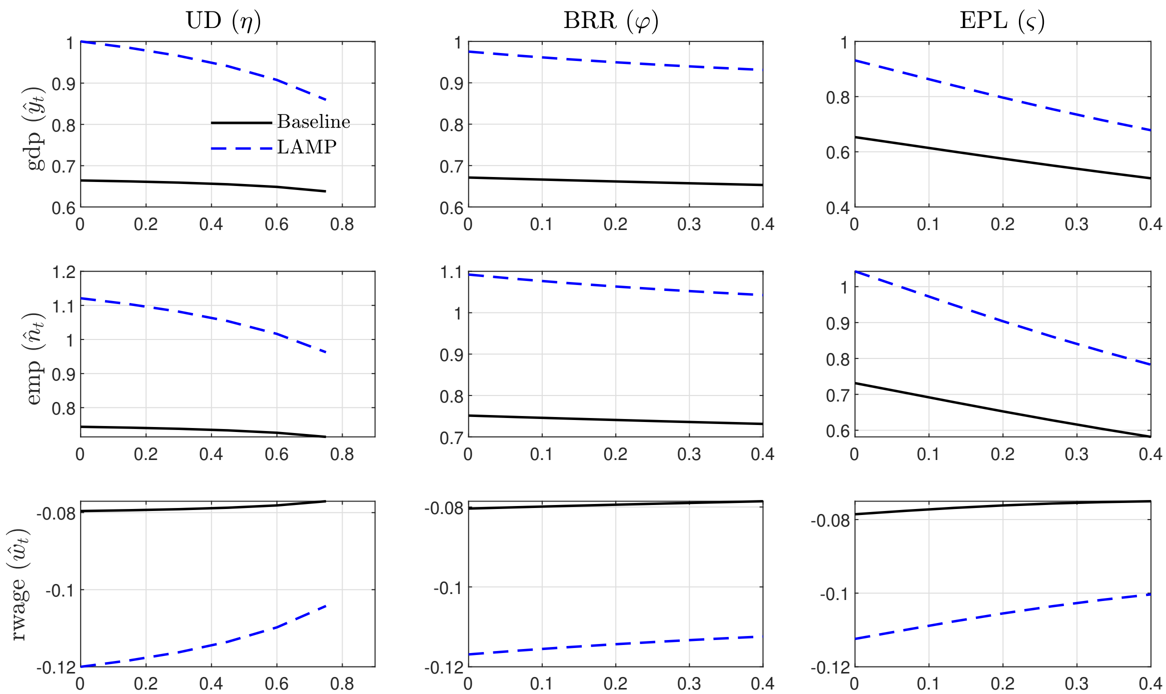
where $\check{w}_t = (1-\eta)\frac{mrs_t+b_t^u}{1-\tau} + \eta\left(mpl_t + E_t\Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b^s_{t+1}(1-\bar{\varrho})F(\tilde{a}_{t+1})\right]\right)$. The parameter ϱ_w captures the extent of real wage rigidity. Equation (A.2) can be considered as a parsimonious but ad 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) and Gertler, Huckfeldt and Trigari (2020) and the papers cited therein. The results of the model extended for real wage rigidities are shown and compared to the baseline model in Figure 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.

Limited Asset Market Participation. Galí, López-Salido and Valles (2007) show how the interaction of rule-of-thumb consumers with sticky prices and deficit financing can account for the existing evidence on the effects of government spending. In this context, rule-of-thumb consumers are characterized by limited asset market participation which implies that they lack any ability of smoothing their consumption profile; as a consequence, they spend (consume) each period all of their income. This rule-of-thumb gives rise to a consumption pattern that strongly aligns with wage income. This gives rise to a positive consumption response in the wake of an expansionary fiscal spending shock. We follow Galí, López-Salido and Valles (2007) and add the second consumer type into the baseline model. The consumers outlined in the baseline model are now referred to as *Ricardian* consumers and their consumption is henceforth referred to as c_t^r (same for their labor supply n_t). Rule-of-thumb households are assumed to behave in a "hand-to-mouth" fashion, fully consuming their current labor income. Their period utility is given by $u(c_t^{nr}, n_t^{nr})$ and they are subject to the budget constraint $c_t^{nr} = (1 - \tau)w_t n_t^{nr} + b_t^u(1 - n_t^{nr}) + T_t^{s,nr}$. Aggregate consumption and employment are given by a weighted average of the corresponding variables for each consumer type. Formally, $c_t = \lambda c_t^{nr} + (1 - \lambda)c_t^r$, $n_t = \lambda n_t^{nr} + (1 - \lambda)n_t^r$. It is further assumed that the labor market is characterized by a structure which gives rise to wages being negotiated in a centralized manner by an economy-wide union with firms.

Figure 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.

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.

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.

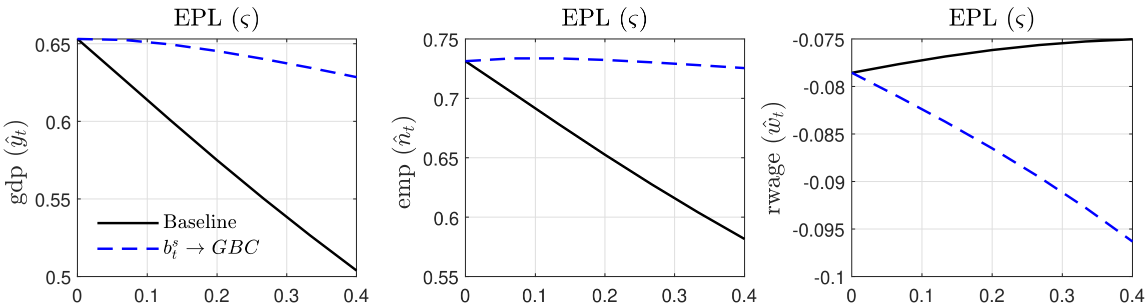
Firing Costs as Government Revenues. The baseline model specifies firing costs as real resource costs. This is a quite strong assumption, as in many countries firing costs arise in the context of severance payments, etc. which will eventually be re-distributed back to households. Against this background, 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 (4.10)) and the real resource constraint (equation (4.11)) 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

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



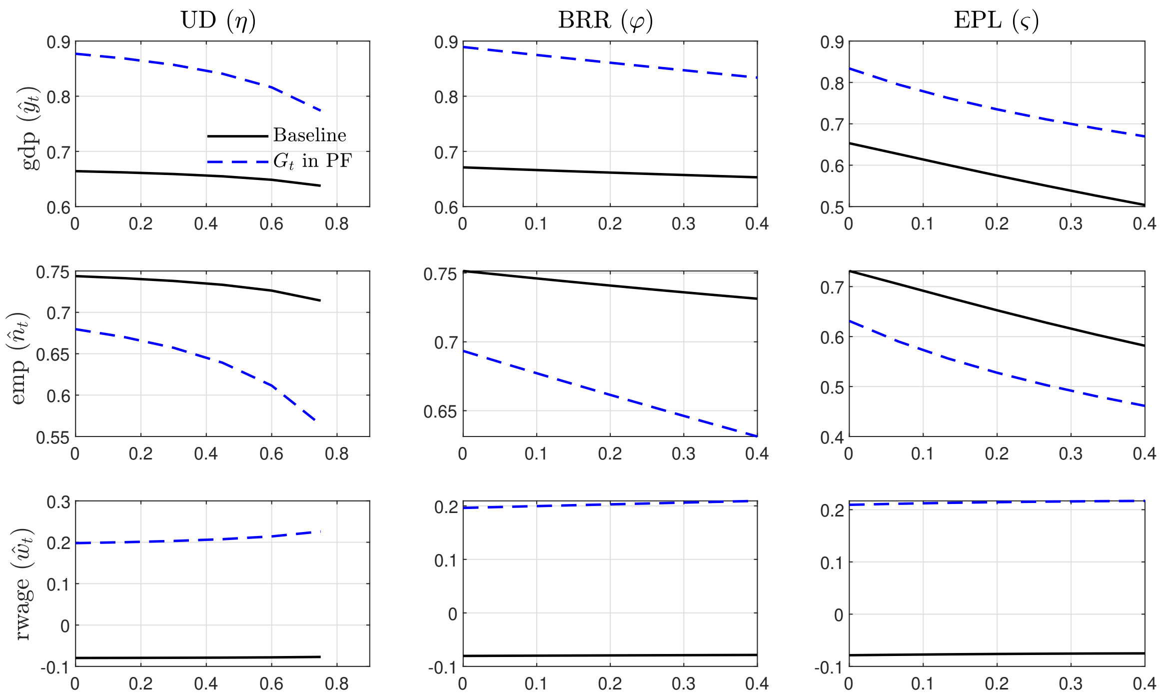
Note: The sub-plots show the sensitivity of the fiscal spending multipliers to changes in the structural parameters. The multipliers are shown for a horizons of $\mathcal{P} = 0$, i.e. contemporaneous multiplier.

firing costs ($\varsigma = 0$), this extension hence has no effect on the shape of the multipliers with respect to η and φ.

The results are shown in Figure 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 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.

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

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.

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 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 4.7. The results thereof are shown in Figure 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 Figure A2). Most important aspect for us concerns, though, the impact of the LMIs on the output multiplier. With a view to Figure 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.5 The LMIs and macroeconomic volatility

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 consider a the government spending shock next to a technology shock so that the model comprises a demand and supply shock. The volatility analysis carried out in Section 4.7 examined the change in the volatility of the vector of endogenous variables of the model with respect to changes in the LMIs. Most importantly, the two shocks were used jointly in the simulations. While this aligns with the empirical analysis done in Section 5.5, it could still be the case, however, that the impact of the LMIs on the volatility of the ednogensous varibales strongly depends on the particular shock considered. Against this background, we now perform the same analysis as in section 4.7, but decompose the effect of LMIs on the volatility of endogenous variables to the corresponding shocks. That is, we examine the extent to which LMIs affect macroeconomic volatility in the wake of demand and supply shocks. 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 (4.12) and following Hamilton (1994), the variance covariance matrix $\boldsymbol{\Sigma}_z(\boldsymbol{\vartheta})$ of the vector of endogenous variables $\boldsymbol{z}_t$ is given by

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

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 again confine the analysis to the volatility of output ($\hat{y}_t$), employment ($\hat{n}_t$) and the real wage ($\hat{w}_t$). We use the estimated values put forth in Christoffel, Kuester and Linzert (2009) to calibrate the parameters

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

LMI:	UD (η)	BRR (φ)	EPL (ς)
	Government spending shock (ϵ_t^G)		
Output ($\hat{y}_t$)	0.99	1.11	1.08
Employment ($\hat{n}_t$)	1.00	1.10	1.10
Real wage ($\hat{w}_t$)	1.02	1.05	1.05
	Technology shock (ϵ_t^A)		
Output ($\hat{y}_t$)	0.97	1.08	1.12
Employment ($\hat{n}_t$)	1.12	0.87	1.06
Real wage ($\hat{w}_t$)	0.99	0.96	0.99

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

for the auto-correlation coefficients (ρ_i) of the AR(1) shocks. The calibration of the idiosyncratic variances (σ_i^2) is described in Section A.2 and we consider two values (high and low) for each LMI in this respect.

The results are depicted in Table A2. The table only shows the values of the variance of the endogenous variables (output, employment and the real wage) for a low value of the LMIs relative to their variances when a high value is used. A few results emerge from the analysis. First, the LMIs' effect on macroeconomic volatility tends to strongly depend on the particular shock considered. In case of the government spending shock, higher values of the LMIs give rise to lower volatility in most cases, the opposite in turn applies for the technology shock. Second, the quantitative impact of the LMIs on macroeconomic volatility in the wake of a technology exceeds the corresponding impact from the government spending shock. 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, Towbin and Wieladek (2014). The model is estimated in its recursive form to allow for contemporaneous interactions. Structural analysis (e.g., IRFs or FEVDs) is then carried out given a particular value of the interaction term.

Let $\{\boldsymbol{y}_{it}\}_{t=1}^{T_i}$ and $\{\boldsymbol{\vartheta}_{it}\}_{t=1}^{T_i}$ denote an 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 $\boldsymbol{c}_i(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}\left(\boldsymbol{a}_i + \boldsymbol{b}_i\boldsymbol{\vartheta}_{it}\right)$, $\boldsymbol{\Phi}_{ij}(\boldsymbol{\vartheta}_{it}) = \boldsymbol{J}_{it}^{-1}\left(\boldsymbol{A}_{ij} + \sum_{l=1}^{d}\boldsymbol{B}_{ijl}\vartheta_{ilt}\right)$, 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, Litterman and Sims, 1984). Hence, the prior distribution on λ_{ij}^2 is a multiplicative Gamma prior

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

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

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

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

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

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 $s = 2, \dots, M$ and $t = 1, \dots, s-1$. 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 $\boldsymbol{\Omega}_i$,

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

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 Table C1, we define the exact transformations of the variables used in the estimation. Note that we use year-on-year growth rates. In Table 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.

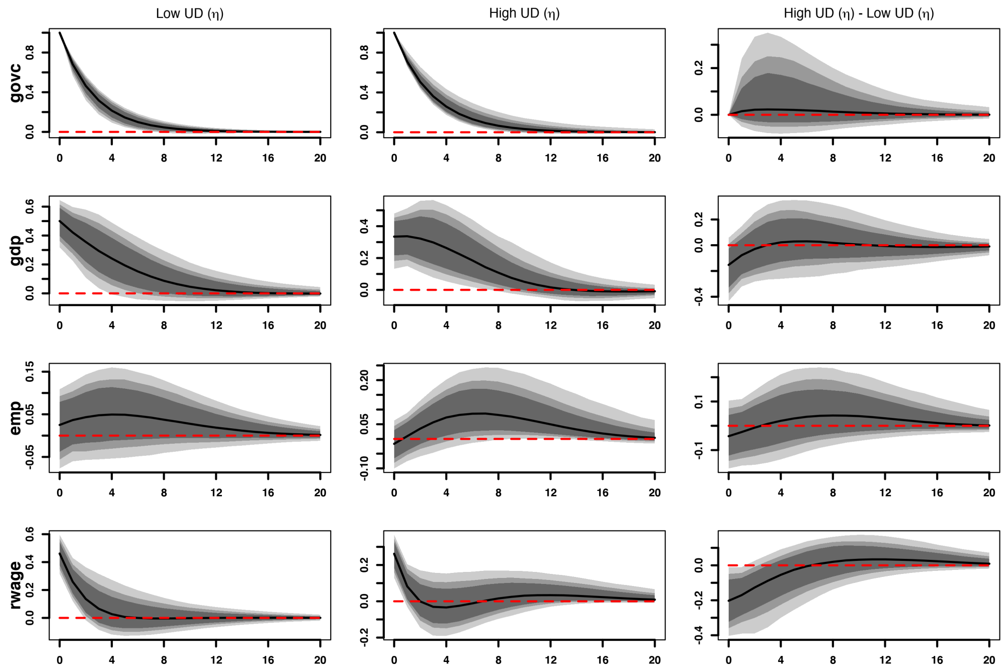
Countries / Model with...	Employment	Unemployment	Tightness
Australia	1966Q3-2020Q2	1964Q1-2020Q4	1978Q2-2020Q4
Austria	1970Q1-2020Q4	1970Q1-2020Q4	1970Q1-2020Q4
Belgium	1980Q4-2020Q4	1980Q4-2020Q4	no data
Canada	1961Q1-2020Q4	1961Q1-2020Q4	no data
Denmark	1980Q1-2020Q4	1980Q1-2020Q4	2009Q1-2020Q4
Finland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

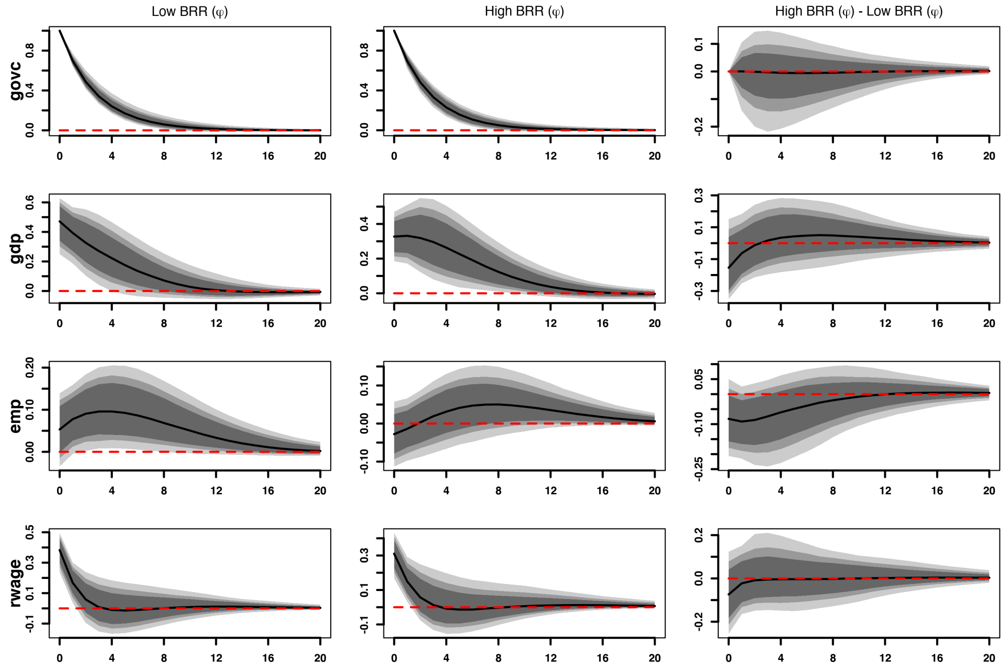
D.1 Additional Results: Effects on Fiscal Spending

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



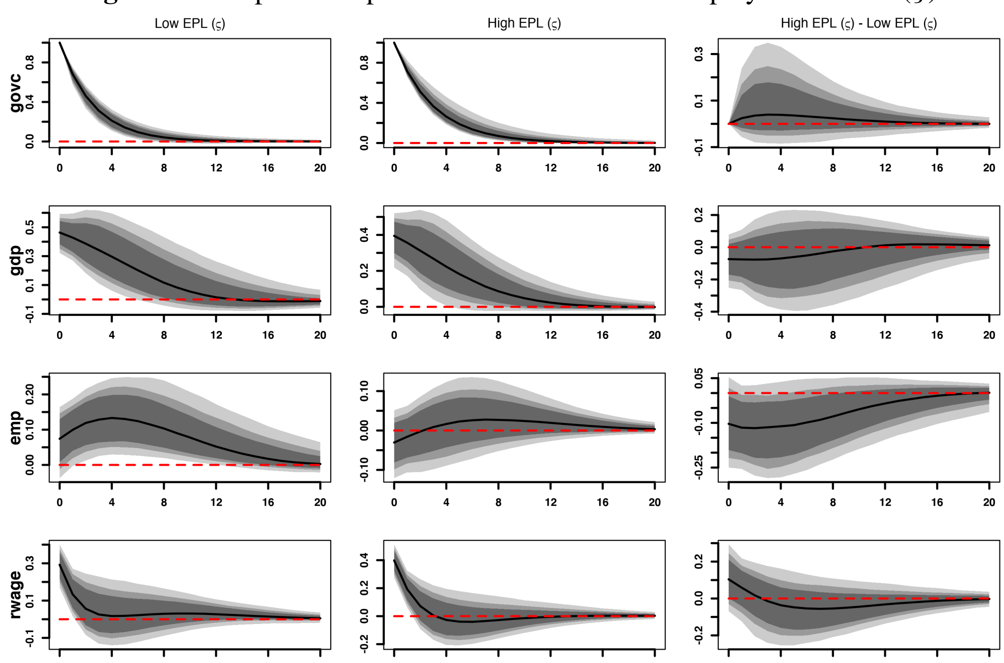
Notes: Impulse response function are shown with 68/80/90 % confidence bounds. Responses are scaled in growth rates for government consumption (govc), gross domestic product (gdp), employment (emp), and real wage (rwage). "Low" indicates −2 standard deviation and "High" indicates +2 standard deviations.

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.

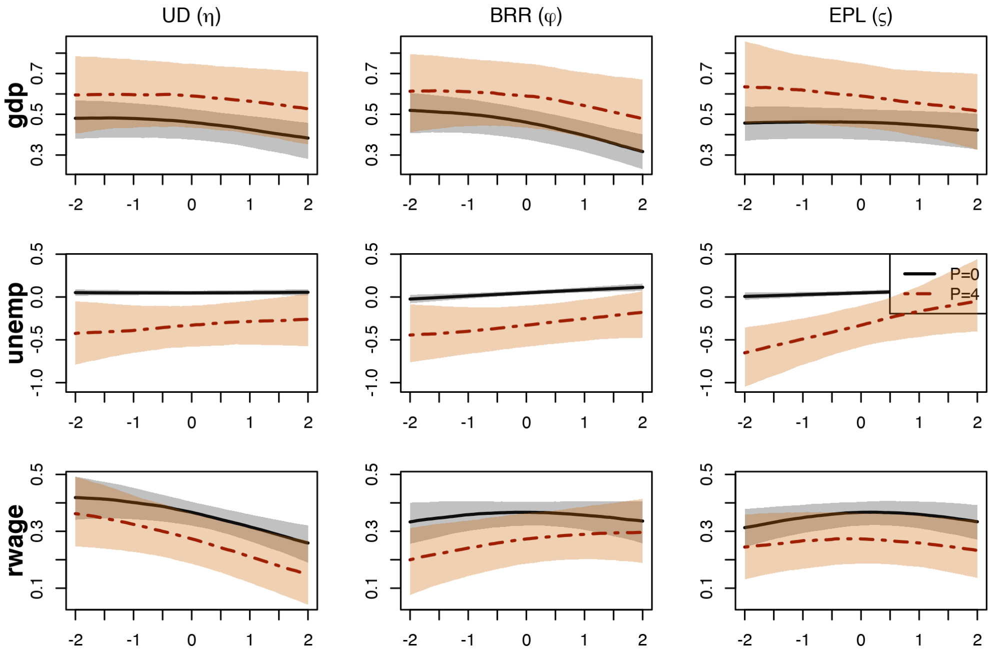
Figure D3: Impulse Responses of the Model with Employment - EPL (ς).



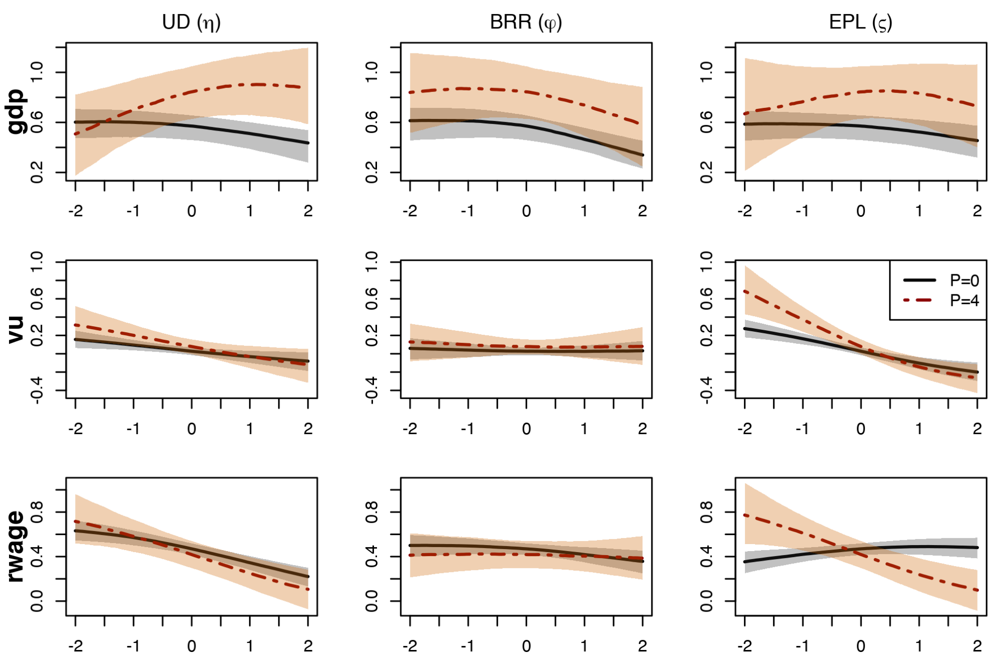
Notes: Impulse response function are shown with 68/80/90 % confidence bounds. Responses are scaled in growth rates for government consumption (govc), gross domestic product (gdp), employment (emp), and real wage (rwage). “Low” indicates −2 standard deviation and “High” indicates +2 standard deviations.

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

(a) Unemployment.



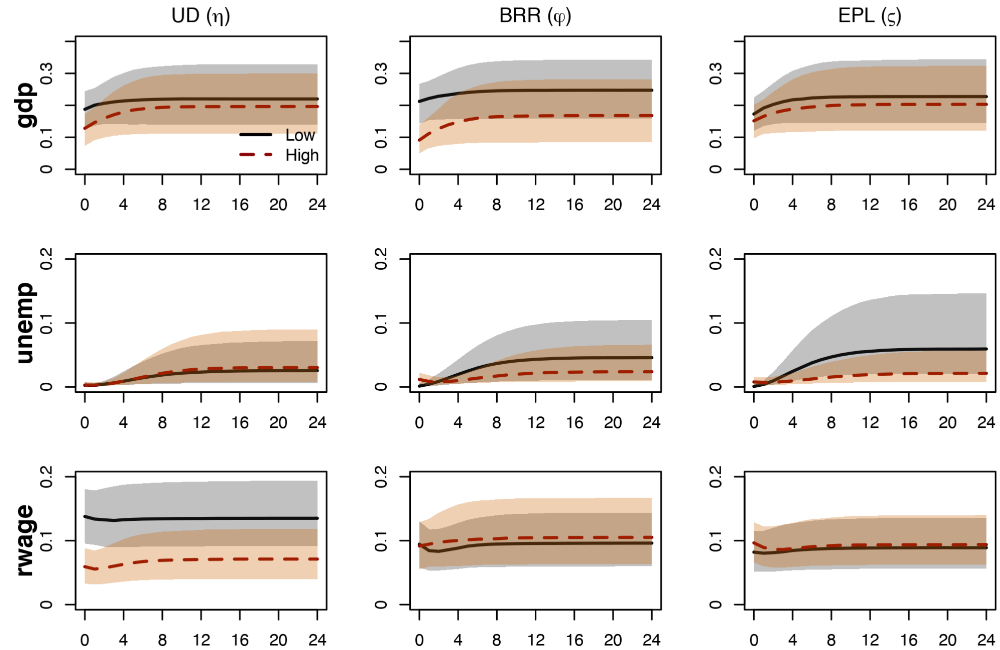
(b) Labor Market Tightness.



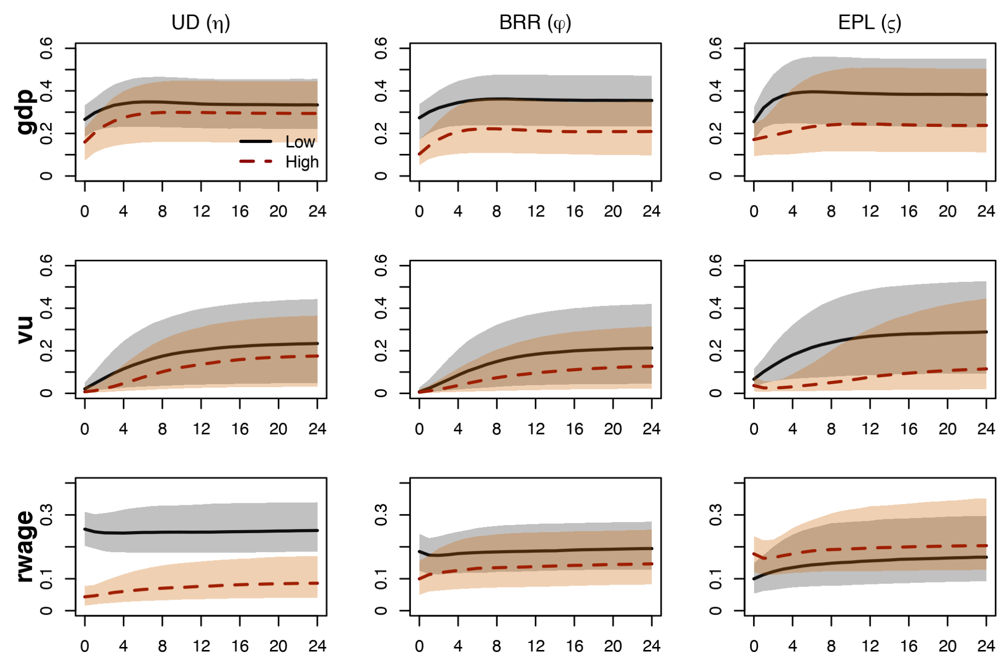
Notes: The sub-plot shows the sensitivity of the fiscal spending multipliers to changes in the structural parameters (η is union density, φ is unemployment benefit replacement rate, and ς is employment protection). The y-axis gives the size of the multiplier while the x-axis runs from -/+ 2 standard 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.

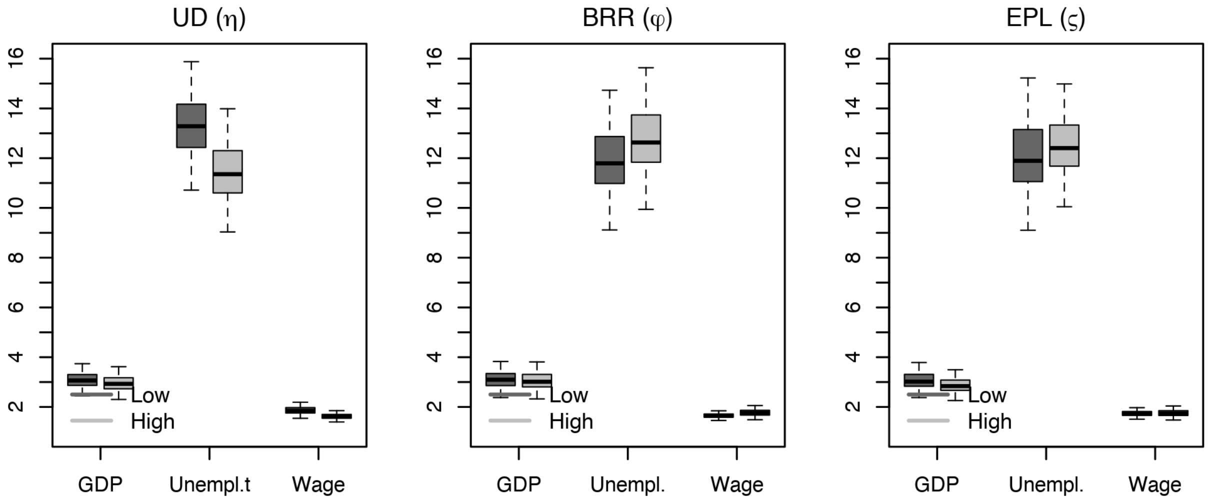


Notes: The sub-plots show the sensitivity of the explained forecast error variance to changes in the structural parameters (η is union density, φ is unemployment benefit replacement rate, and ς is employment protection). The y-axis gives the share of explained forecast error variance while the x-axis 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.

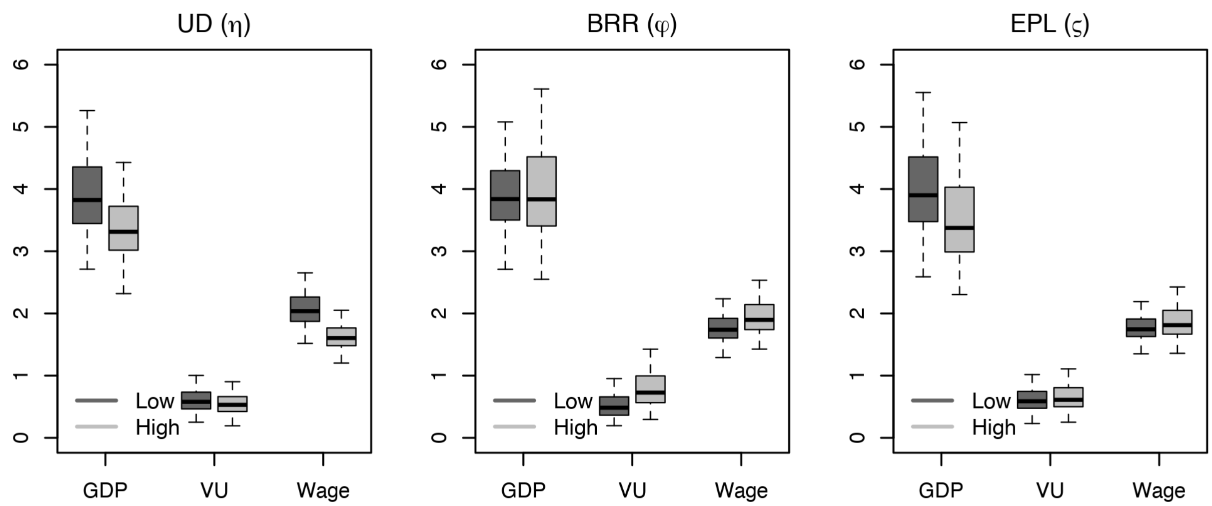
D.2 Additional Results: Effects on Macroeconomic Volatility

Figure D6: Macroeconomic Volatilities Along LMIs with Other Variables.

(a) Unemployment.



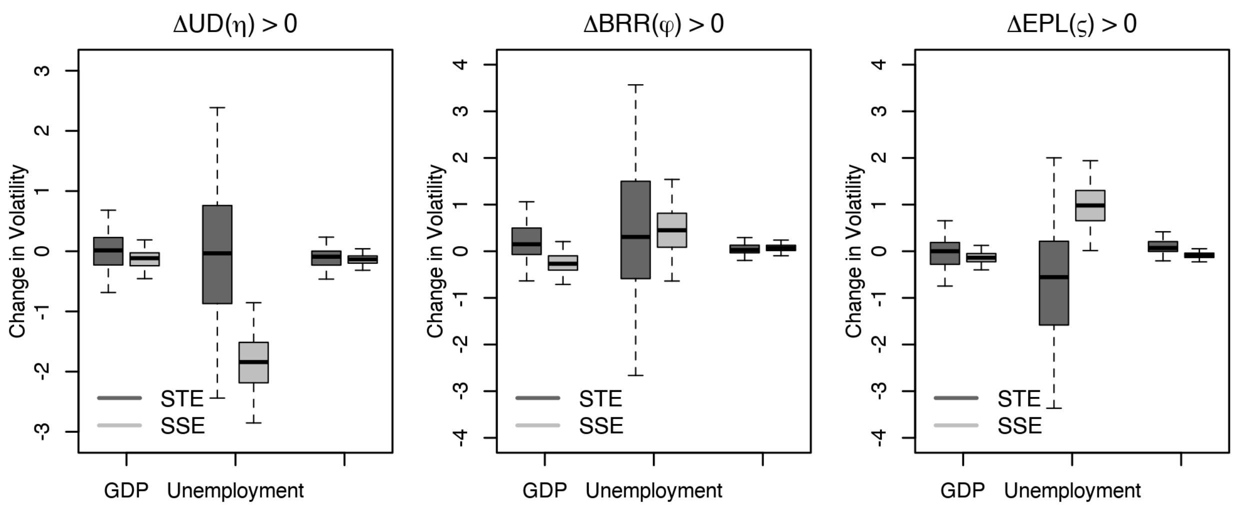
(b) Labor Market Tightness.



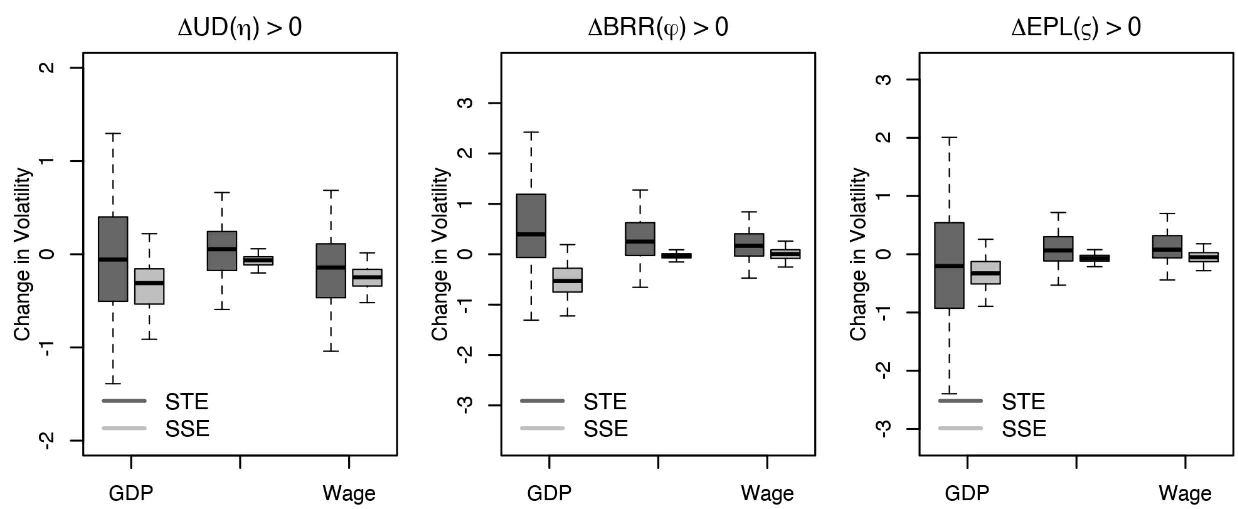
Notes: Each sub-plot shows the standard deviations of the respective macroeconomic variable in a regime with low (−2sd) and high (+2sd) LMIs. The LMIs under consideration are union density (UD(η)), unemployment benefit replacement rate (BRR(φ)), and employment protection (EPL(ς)).

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 Equation 5.7. The LMIs under consideration are union density (UD(η)), unemployment benefit replacement rate (BRR(φ)), and employment protection (EPL(ς)).

D.3 Additional Results: Inspecting the Between-Country Variation

The analysis in section 5.4 is based on within-country variation of the LMIs. In this additional section, we want to explore possible effects of between-country variation. As we have noticed in Figure 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 Equation 5.2 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 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.[18] The groups align well with various definitions of welfare regimes and are depicted in Figure D9 in the Appendix.

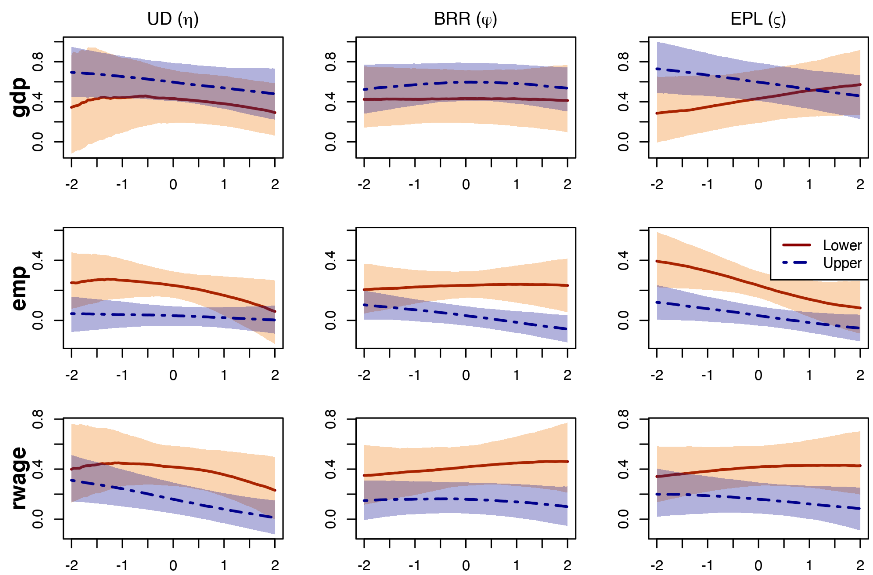
In Figure D8, we examine the fiscal multipliers both on a within- and between-country variation basis. In line with the theoretical results and the analysis conducted in subsection 5.4, we use Equation 4.13 to compute multipliers for the effects after four quarters ("one-year multiplier", horizon $P = 4$) 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; 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 Figure D10 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 5.4. In the "upper" group of countries, cyclical policies do not have a strong effect on labor market

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

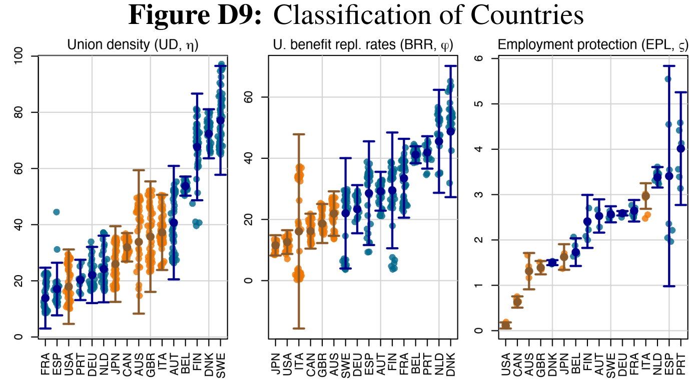
variables. Cyclical policies still affect the fiscal multipliers in the "lower" group of countries – with a clear downward-sloping effect along the within-country variation.

Figure D8: 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.

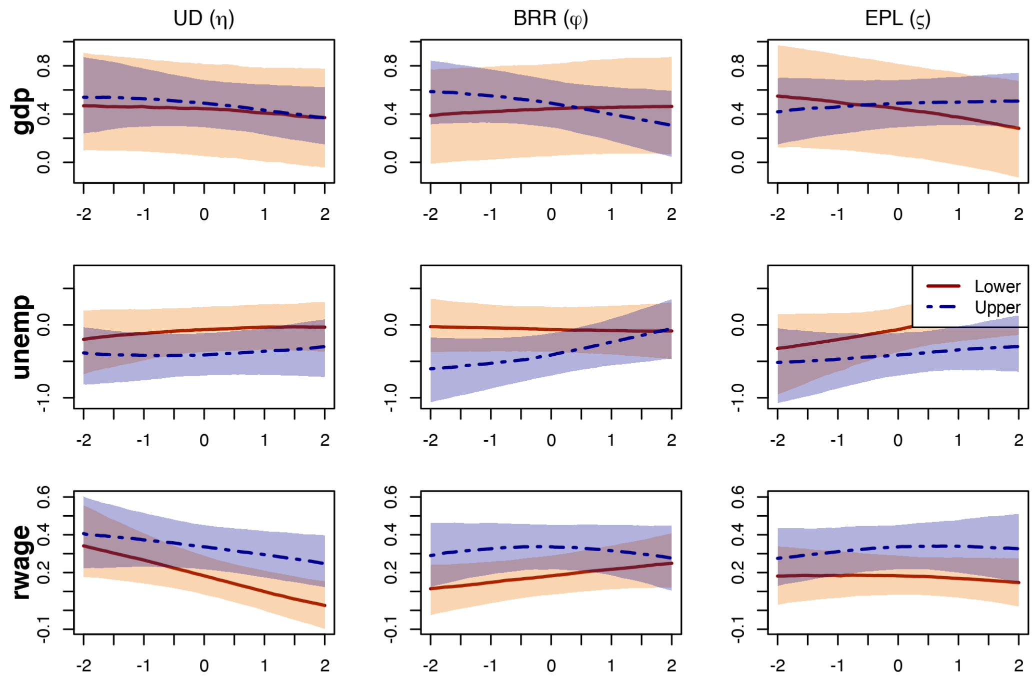
Figure D9: Classification of Countries



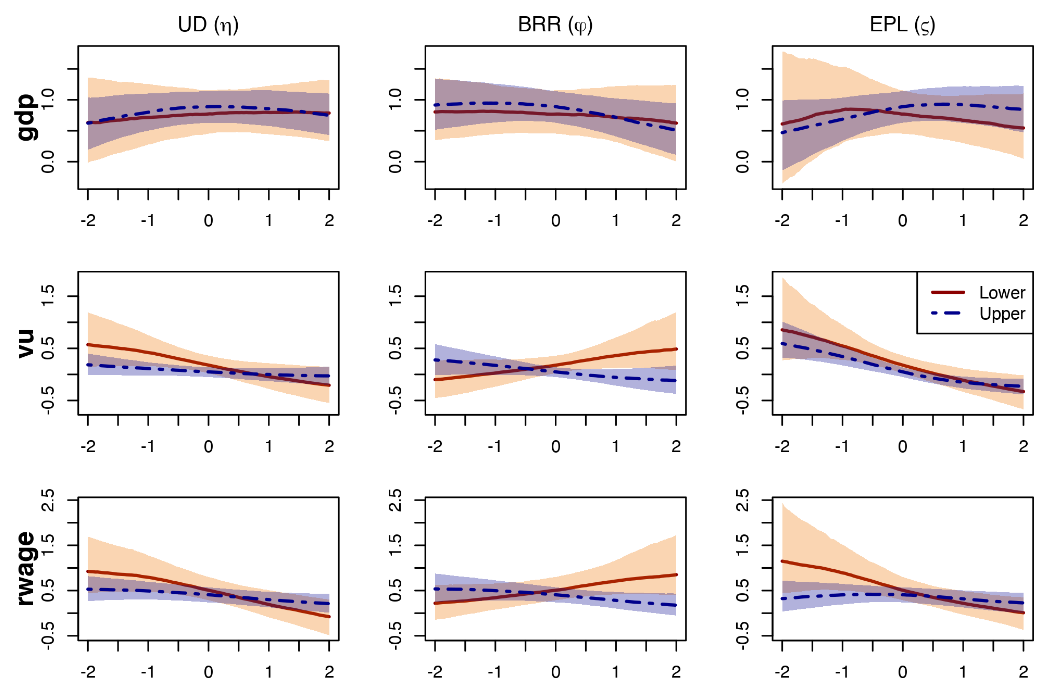
Notes: Each sub-plot shows the mean of each LMI for each country, together with two standard deviation in each direction. The points are observed data for the respective country. Color shadings differentiate countries belonging to the *Upper Group* (blue) and *Lower Group* (orange).

Figure D10: 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.