The Role of Labor Market Institutions in Shaping Euro Area Monetary Policy Transmission*

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First Version: July 2024
This Version: January 2025

Abstract

We examine the role of labor market institutions (union density, benefit replacement rates and employment protection legislation) in shaping the transmission mechanism of monetary policy. Our theoretical framework posits that higher union density flattens the Phillips curve, thereby magnifying the effect on output while attenuating the impact of monetary policy shocks on inflation. This is empirically confirmed by means of an interacted panel VAR model estimated for euro area countries. Conversely, benefit replacement rates and employment protection legislation exhibit limited influence. Our findings identify a structural rather than a cyclical phenomenon for shaping monetary policy effectiveness. They underscore the importance of labor market characteristics in this context, particularly within a monetary union where heterogeneous labor markets can lead to inefficient inflation and output differentials.

Keywords: Monetary policy; Labor market institutions; Euro area; Interacted panel VAR.
JEL Codes: C32, C33, E52, J21, J38 .

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1. Introduction

The unexpected rise in inflation in the euro area (and the US) has ignited a lively debate not only about its exact causes but also about the appropriate policy responses from a central bank perspective. In the euro area, inflation has risen sharply in 2021 and 2022, peaking at around 10 percent (CPI, year-over-year) in late 2022. After the re-opening of the economy following the Covid-19 pandemic, supply disruptions associated with pandemic-induced sectoral reallocations stoked inflation at first (Bernanke and Blanchard, 2024; Dao et al., 2024). However, strong fiscal relief programs across advanced economies and the tightening of labor markets shifted the focus more towards demand factors (Benigno and Eggertsson, 2023; Bergholt et al., 2023). For the fulfillment of their mandate towards price stability, central banks monitor, *inter alia*, developments on the labor market carefully.[1] While the labor market is shaped by cyclical developments, structural factors are equally important.[2] In the euro area, labor markets are characterized by relatively strong rigidities, such as a longer duration of unemployment spells, inflexible wages, strong unionization, or higher benefit replacement rates as compared to the US. As highlighted in the theoretical work by Christoffel, Kuester and Linzert (2009), labor market characteristics potentially alter the monetary policy transmission mechanism.

This paper thus investigates theoretically and empirically the efficacy of monetary policy conditional on structural impediments that arise from the labor market. In detail, the analysis focuses on three labor market institutions (LMIs): union density (UD), benefit replacement rates (BRR), and employment protection legislation (EPL). The paper develops a dynamic stochastic general equilibrium (DSGE) model to qualitatively assess the role of distinct LMIs for the transmission of monetary policy. This theoretical analysis motivates the empirical model. While the LMIs are modeled as deep structural parameters affecting the equilibrium steady state in the DSGE model, we use slow-moving variables to characterize these institutions as interaction terms in a macroeconomic model. For this purpose, we estimate a Bayesian interacted panel vector autoregressive (IPVAR) model for a set of euro area countries in which the LMI variables serve as an exogenous interaction term. The value of the interaction terms affects the moving average representation as an approximation of the steady state. We empirically identify a structural monetary policy shock and examine how the transmission of monetary policy is affected given specific structural arrangements on the labor market.

[1] See, for instance, a speech by Lane (2024) in which he discussed that "wage growth is expected [...], reflecting the ongoing gradual correction of the real wage gap."

[2] See, for instance, Nickell (1997), Snower and La Dehesa (1997), or Layard and Nickell (1999).

The paper proceeds in two parts. In the first part, we develop a relatively standard theoretical model to assess how LMIs affect the transmission of monetary policy. The DSGE model incorporates three main frictions: (i) search frictions in the labor market, (ii) nominal frictions through Calvo pricing, and (iii) firing costs. Structural elements of the labor market can affect inflation dynamics and the transmission of monetary policy via price or quantity rigidities. Price rigidities affect the real wage through the level of the household's reservation wage (via BRR) or through workers' bargaining power in the Nash wage bargaining (proxied through UD). Quantity rigidities are modeled via firing costs (proxied through EPL) that affect labor demand. We show that this leads to a flattening of the Phillips curve, which is particularly pronounced for UD. Contractionary monetary policy then causes a stronger output reaction while the inflation response flattens. This trade-off between output and inflation is virtually similar for all three indicators used in the theoretical exercise.

In the second part of the paper, we corroborate the theoretical predictions in an empirical exercise. We estimate an IPVAR model for a sample of euro area countries for the time period 1999-2023. We measure union density by the percentage of workers affiliated with a union, the benefit replacement rates measure the generosity of the unemployment benefit system, and employment protection legislation is given by a synthetic strictness indicator. Given more flexible or rigid labor market characteristics, we retrieve different approximations of the steady state. This is possible by using the LMI indicators as exogenous interaction variables in the IPVAR. Once this is established, we identify euro area monetary policy shocks using high-frequency surprises provided by Altavilla et al. (2019). We use the framework of Jarociński and Karadi (2020) to disentangle monetary policy shocks from information effects. Using the identified euro area monetary policy shocks on the euro area as a whole, we pursue a block exogenous structure for individual member countries. This is justified by the assumption that each member country is a small open economy relative to the euro area block. This allows us to trace an aggregate euro area shock to individual member countries. We then use the mean-group estimator of Pesaran and Smith (1995) to compute expressions for a typical euro area country.

Similar to the theoretical analysis, we vary the degree of labor market rigidities in a typical euro area economy. We examine the two corner cases of a high ("more rigid") and a low ("more flexible") value for each LMI separately by analyzing impulse response functions (IRFs) and forecast error variance decompositions (FEVDs). In the IRF analysis, we investigate the responses to a 25 basis points (bps) contractionary monetary policy shock. This shock acts as a demand shifter in the euro area aggregate and causes both a decline in output and price / wage measures. However, for a typical euro area country our estimates suggest that price and wage responses are much more pronounced in flexible labor markets than in rigid labor markets. On the contrary, we observe a stronger output reaction in the model with more rigid labor markets. This effect holds only for UD, while BRR and EPL show no significant differences in the price, wage, or output response. However, median effects also point to stronger output effects in more rigid labor markets for BRR and EPL.

The remainder of the paper is structured as follows. In the next section, we discuss the related literature and our contribution. Section 3 provides an overview on the data and a descriptive analysis of the structural labor market indicators. Section 4 develops the theoretical model and discusses the extent to which changes in labor market institutions shape the transmission mechanism of monetary policy. Section 5 introduces the econometric model and provides an empirical assessment of the theoretical model's implications. Finally, Section 6 concludes.

2. Related literature

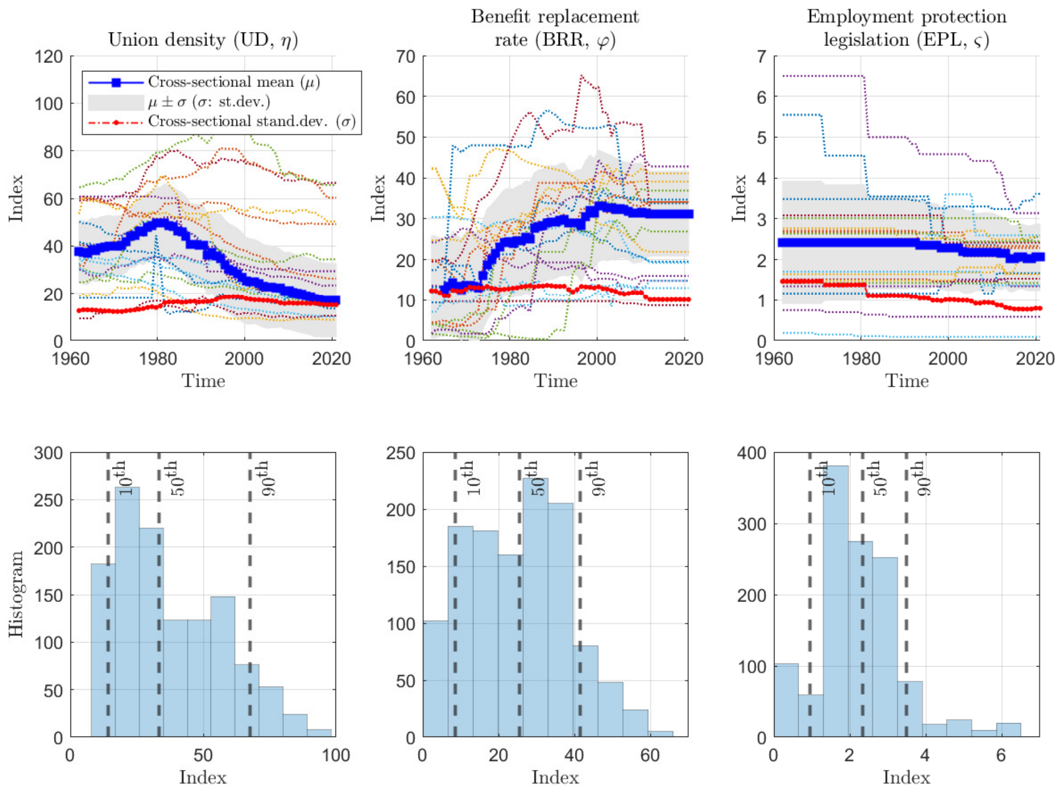
Our contribution is related to two important streams of the literature. The first concerns elements that shape the effectiveness of monetary policy, while the second studies the role of labor market institutions for business cycle fluctuations.

Much of the discussion of the potentially diminished traction of monetary policy has focused on the reduced policy space as a result of a fall in the equilibrium real rate of interest (see, for instance, Gust et al., 2017). Another focal point of the debate has been the smaller impact of monetary policy on inflation due to a flattening of the Phillips curve (see, for instance, Negro et al., 2020; Glocker and Piribauer, 2021). However, it is also possible that aggregate demand itself has become less responsive to monetary policy. In other words, the IS curve may have steepened – a hypothesis laid out by Borio and Hofmann (2017). Gornemann, Kuester and Nakajima (2016) highlight the distributional consequences transmitting strongly through frictional labor markets where wage earners face higher cyclical unemployment risk. The existing empirical and theoretical literature has focused primarily on cyclical labor market characteristics that shape the transmission of monetary policy (see Galí, 2022, among others) and on the effects of monetary policy on the labor market in general (see Cantore, Ferroni and León-Ledesma, 2021; Zens, Böck and Zörner, 2020). We instead take the opposite view, that is, we assess the extent to which structural labor market characteristics shape the transmission channel of monetary policy. This has received considerably less attention.

Gnocchi, Lagerborg and Pappa (2015) provide stylized facts about the relationship between labor market institutions and business cycle fluctuations. They highlight that more flexible institutions are associated with lower business cycle volatility. Abbritti and Weber (2018) find that LMIs have a large and significant effect on both unemployment and inflation dynamics. Stricter employment protection legislation (EPL) and higher union density (UD) mute the reaction of unemployment but increase the response of inflation to external shocks. The results of Gnocchi, Lagerborg and Pappa (2015) imply that countries with very rigid or very flexible labor markets can have similar inflation and unemployment dynamics. Zanetti (2009) and Zanetti (2011), for instance, provide theoretical evidence highlighting the importance of employment protection legislation (EPL) in shaping aggregate dynamics. In contrast to that, more generous unemployment benefits have the opposite effect (see Thomas and Zanetti, 2009; Campolmi and Faia, 2011, for instance): as the higher benefits improve workers' outside option, this in turn reduces the responsiveness of wages and inflation to aggregate fluctuations, and triggers an increase in unemployment volatility. In an empirical evaluation, Lastauskas and Stakėnas (2024) show that active labor market policies are inflationary but reduce unemployment but these outcomes depend on the monetary policy stance. We instead take the opposite approach and examine the transmission of policy shocks conditional on the labor market policies in place.

The literature examining how structural characteristics of the labor market shape the transmission of cyclical shocks is relatively scarce. Abbritti and Weber (2018) focus on external shocks, while other papers examine the alteration of fiscal policy shocks (Brückner and Pappa, 2012; Cacciatore et al., 2021; Boeck, Crespo Cuaresma and Glocker, 2022). The effects of monetary policy and how the transmission mechanism is altered due to labor market institutions has so far not been investigated through an empirical model. This is surprising given that central banks put emphasis on monitoring the labor market. In addition to this, little consensus emerges from the literature on structural labor market characteristics and their implication for inflation dynamics. For instance, Christoffel, Kuester and Linzert (2009) highlight in a theoretical model that labor market institutions affect the extent of nominal wage stickiness, which has direct implications for the efficacy of monetary policy. Moreover, in the case of stricter employment protection legislation, some studies find no effect on inflation volatility (Merkl and Schmitz, 2011) or output volatility (Rumler and Scharler, 2011), while others find a negative effect on unemployment or output volatility (Faccini and Rosazza Bondibene, 2012) or an inverted U-shaped effect on the relative unemployment to output volatility (Lochner, 2024). The available evidence is similarly inconclusive in the case of benefit replacement rates or for the impact of unions on business cycle dynamics.

Figure 1: Labor market institutions (LMIs) overview



Notes: This figure shows the time trajectory (upper panels) and the histograms across time (lower panels) of three labor market indicators for the countries of the euro area. The Greek letters attached to each LMI indicator refer to their parametric counterpart in the theoretical model.

3. Labor market indicators in the euro area

Figure 1 shows the time path and distribution of three labor market indicators for countries of the euro area currency block: union density (UD, left-hand side panels), benefit replacement rate (BRR, middle panels), and employment protection legislation (EPL, right-hand side panels). The top panels display the time trajectory for each country, extended for the cross-sectional mean (blue line) and standard deviation (red line). The histograms for each labor market indicator across time and countries are displayed in the bottom panels.

The data for the LMIs are taken from the CEP-OECD institutions database. The original data are arranged on an annual frequency. We extend the data until 2022 using information from the OECD; the ICTWSS Database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts (Visser, 2013); and calculations as described in Abbritti and Weber (2018).

The measurement of union density (UD) is predominantly derived from survey data, supplemented by administrative data adjusted for non-active and self-employed members when necessary. UD is calculated as the proportion of wage and salary earners who are affiliated with trade unions relative to the total number of wage and salary earners. Elevated values of UD signify a higher proportion of trade union membership and, consequently, a heightened influence of trade unions in wage negotiation processes.[3] Unemployment benefit replacement rates (BRR) gauge the extent to which income is sustained following a specified duration of unemployment. This metric is determined by the ratio of net household income during a designated month of unemployment to the net household income preceding job loss. Higher BRR values denote more comprehensive unemployment benefit systems. These two metrics, UD and BRR, exert influence on employment, job vacancies, and unemployment dynamics through their impact on labor costs. In contrast, employment protection legislation (EPL) is anticipated to directly impact the quantity of labor, with wage adjustments typically occurring as secondary effects. The EPL index serves as a composite measure, encapsulating the stringency of regulations concerning dismissals and the utilization of temporary employment contracts. The employment protection indicator for each year references the regulatory framework in effect as of January 1st. Elevated values of the indicator signify a greater degree of employment protection within the given jurisdiction.

The upper panels of Figure 1 show significant variation in each LMI across time and countries. The mean of UD has decreased across countries since 1980, while cross-country heterogeneity has increased over the same time period. The decline in UD suggests a shift towards decentralized wage setting in advanced economies. Bhuller et al. (2022) note in this context that this measure overlooks the significant coverage of collective bargaining agreements, which often extends beyond union membership. Particularly in continental European and some Scandinavian countries, collective bargaining covers a larger share of the workforce, including non-union positions (see also Flanagan, 1999). Although the OECD provides data on the extent of collective wage bargaining coverage, it is not feasible in our setting due to its limited country coverage and short time-span.

A similar divergent pattern arises for BRR. While the average rates have increased across countries, the extent of cross-country heterogeneity has steadily decreased. In the context of EPL, the average value across countries has steadily declined, coinciding with a decrease in country heterogeneity.

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

The particular form of time variation of the LMIs holds significant importance for the structure of the subsequent theoretical and empirical models. Although all LMIs exhibit some time variation in all countries, the fluctuations therein extend beyond business cycle fluctuations. Therefore, changes in the LMIs can be regarded as structural changes – while they vary over time, in most cases, changes occur solely once in a decade. Hence, regarding the theoretical model, the LMIs are introduced as structural parameters in a standard model for the business cycle (such as the New Keynesian model or the real business cycle model). As for the empirical model, the LMIs are considered weakly exogenous variables as they are not affected by short-term monetary policy.

The histograms for each LMI over time and countries are shown in the lower panels in Figure 1. Although the distribution in each case resembles a bell shape, the significant degree of skewness stands out. The distribution is most skewed in the case of UD, followed by EPL and BRR. To account for this, in the empirical model we use the values indicated by the 10th and 90th percentiles to operationalize the interaction term. Although this is not an exhaustive list of indicators to describe the overall design of the labor market, these three dimensions cover the most commonly used aspects in studies of the impact of labor market institutions on macroeconomic developments.

4. The theoretical model

In our theoretical framework, we integrate the structure of a Diamond-Mortensen-Pissarides (DMP) model with a conventional New Keynesian (NK) framework, drawing from the framework proposed by Merz (1995), Andolfatto (1996), Krause and Lubik (2007), and Monacelli, Perotti and Trigari (2010).[4] This amalgamation aims for parsimony while directing attention to the role of labor market institutions.

The model incorporates three frictions: (i) search frictions, central to the DMP model, which render vacancy posting costly and reflect labor market inefficiencies; (ii) nominal frictions through sticky price adjustments, which allow for the integration of monetary policy, giving rise to inertial price adjustments in response to shocks; and (iii) firing costs, which constrain labor reallocation and allow to introduce employment protection legislation.

We assume representative households and firms. Among firms, there are intermediate goods producers and final goods producers. Each intermediate goods producing firm employs n_t workers and posts v_t

[4] We aim for a closed-economy model setting. An alternative is to use an small-open economy framework with an exogenously set interest rate mirroring a member country of the euro area and monetary policy set by the ECB. However, we approximate the euro area as a whole to simplify the model framework.

vacancies. These firms incur a cost of κ per posted vacancy and of b_t^s per laid-off worker. The total number of unemployed workers searching for jobs is $u_t = 1 - n_t$. New hires (m_t) are determined by 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 of a firm filling a vacancy is given by $q_t = m_t / v_t = \bar{m}\theta_t^{-\gamma}$, where $\theta_t = v_t/u_t$ represents the labor market tightness (LM-tightness). Similarly, the probability of an unemployed worker finding a job is $p_t = m_t/u_t = \bar{m}\theta_t^{1-\gamma}$. Firms and workers regard q_t and p_t as exogenous. Furthermore, each firm separates from a fraction $\varrho(\tilde{a}_t)$ of its existing workers each period. This fraction consists of an exogenous component ($\bar{\varrho}$) and an endogenous one. Following Krause and Lubik (2007), job destruction probabilities a_t are sampled each period from a distribution $F(a_t)$ with positive support and density $f(a_t)$. If $a_t < \tilde{a}_t$, where $\tilde{a}_t$ is an endogenously determined threshold value, a job is terminated. This process leads to an endogenous job separation rate $F(\tilde{a}_t)$. The total separation rate is expressed as $\varrho(\tilde{a}_t) = \bar{\varrho} + (1 - \bar{\varrho})F(\tilde{a}_t)$.

4.1 Intermediate-good producers

The (representative, intermediate-goods producing) firm uses labor to produce output y_t according to $y_t = \bar{A} n_t A(\tilde{a}_t)$, where $\bar{A} > 0$ is a common productivity factor and $A(\tilde{a}_t) = E[a | a \geq \tilde{a}_t] = \frac{1}{1-F(\tilde{a}_t)} \int_{\tilde{a}_t}^{\infty} a dF(a)$ is the conditional expectation of productivity being larger than the endogenously determined critical threshold. 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

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

Profits are given by $\pi_t^F = \mu_t y_t - w_t n_t - \kappa v_t - F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b_t^s$, where the output price is normalized to unity, $1/\mu_t$ is the price markup ($1/\mu_t$ represents the real value of a unit of output, which is directly related to the real marginal cost for the representative intermediate-goods producing firm) and $w_t = \int_{\tilde{a}_t}^{\infty} \frac{\tilde{w}_t(a)}{1-F(\tilde{a}_t)} dF(a)$ is the (average) real wage weighted according to the idiosyncratic job productivity. The last term captures firing costs (Cacciatore et al., 2021) of which $(n_{t-1} + m_{t-1})(1 - \bar{\varrho})F(\tilde{a}_t)$ 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. 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 (4.1). E_t is the (conditional) expectation operator; and $\Lambda_{t,t+k}$ the stochastic discount factor, defined below. The first order conditions give rise to[5]

$$
\begin{aligned}
F_t^n &= \mu_t mpl_t - w_t + E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right] && (4.2)\\
\frac{\kappa}{q_t} &= E_t\Lambda_{t,t+1}\left[(1-\varrho(\tilde{a}_{t+1}))F_{t+1}^n - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right] && (4.3)\\
\mu_t mpl_t &= w_t - b_t^s - \frac{\kappa}{q_t} && (4.4)
\end{aligned}
$$

where $mpl_t = y_t/n_t = \bar{A}\cdot A(\tilde{a}_t)$ is the marginal product of labor. In equation (4.2), F_t^n captures the (shadow) value to the firm of employing one additional worker which is made up of: (i) the marginal product of a worker, (ii) the 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 productivity ($\tilde{a}_t$) 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 or when (iii) wages decline.

4.2 Final-good producers

Final-goods producers buy intermediate goods and sell them to the households. We assume that there is a continuum of final goods producers indexed by $i \in [0,1]$. They are perfectly competitive in their input markets and monopolistically competitive in their output market. Their price setting is subject to nominal rigidities à la Calvo (1983).

The probability that a firm cannot re-optimize its price for k periods is given by ξ^k. For final-goods producer re-optimizing its price at time t, profit maximization implies that he chooses a target price P_t^* that maximizes the following stream of future profits $E_t\sum_{k\geq 0}\xi^k\Lambda_{t,t+k}\int_0^{y_{t+k}(j)}(P_t^* - mc_{t+k}P_{t+k})dq$ subject to the demand constraint $y_{t+k}(j) = (P_t^*/P_{t+k})^{-\varepsilon}y_{t+k}$. Marginal costs are given by $mc_t = \mu_t$. The first order condition with respect to the price P_t^* implies that the following condition has to hold

[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 (4.1)-(4.3), this equation can be further simplified to: $(1-\bar{\varrho})(1-F(\tilde{a}_t))\left(\frac{\partial A(\tilde{a}_t)}{\partial \tilde{a}_t} - \frac{\partial w_t}{\partial \tilde{a}_t}\right) = b_t^s(1-\bar{\varrho})f(\tilde{a}_t) + \left(\mu_t mpl_t - w_t + \frac{\kappa}{q_t}\right)\frac{\partial \varrho(\tilde{a}_t)}{\partial \tilde{a}_t}$. Using the derivatives of $\frac{\partial A(\tilde{a}_t)}{\partial \tilde{a}_t}$, $\frac{\partial w_t}{\partial \tilde{a}_t}$ and $\frac{\partial \varrho(\tilde{a}_t)}{\partial \tilde{a}_t}$ yields the following expression: $\tilde{w}_t(a) = b_t^s + \frac{\kappa}{q_t} + \bar{A}a$. Finally, operating on both sides with $\int_{\tilde{a}_t}^{\infty}\frac{dF(a)}{1-F(\tilde{a}_t)}$ and using the definition of the production function gives Equation (4.4).

$E_t \sum_{k\geq 0} \xi^k \Lambda_{t,t+k} y_{t+k}(j) \left(P_t^* - \frac{\varepsilon}{\varepsilon-1} mc_{t+k} P_{t+k}\right) = 0$. Finally, the definition of the price index P_t yields the following law of motion: $P_t = \left(\xi P_{t-1})^{1-\varepsilon} + (1-\xi)(P_t^*)^{1-\varepsilon}\right)^{\frac{1}{1-\varepsilon}}$.

4.3 Households

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

The budget constraint is given by

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

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 member and lump-sum subsidies. 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 and the employment flow. Optimization leads to the following conditions

$$1 = E_t \left[\Lambda_{t,t+1} \frac{R_t}{1+\pi_t}\right], \tag{4.7}$$

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

where $1 + \pi_t = P_t / P_{t-1}$, λ_t is the Lagrange multiplier attached to equation (4.5) and $\lambda_t H_t^n$ the one attached to equation (4.6). Furthermore, $\tilde{w}_t^b = w_t - b_t^u$, $mrs_t = -u_{n,t}/\lambda_t$ and $u_{n,t} < 0$ is the marginal dis-utility of working.[6] 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.

In equation (4.8), H_t^n captures the household's (shadow) value of having one additional employed member. It has three components: (i) the increase in utility from higher income when an additional member is employed, (ii) the decrease in utility from less leisure, captured by the marginal dis-utility of work, and (iii) the continuation utility value, given by the contribution of a current match to a household's employment in the next period.

4.4 Nash wage bargaining

Wages are set each period based on Nash-bargaining of the (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:

$$\eta F_t^n = (1-\eta) H_t^n. \tag{4.9}$$

The equilibrium wage per worker can also be viewed from the reservation wage of the firm and the worker. The reservation wage of a worker (firm) is given by the minimum (maximum) wage acceptable. Since H_t^n (F_t^n) describes the marginal value to the worker (firm) of having one further worker employed, the reservation wages of a worker and a firm are hence determined by $H_t^n = 0$ and $F_t^n = 0$. In this situation, the worker and the firm are not willing to increase or to decrease labor supply and demand. Using equations (4.2) and (4.3), the reservation wages are given by

$$\overline{w}_t^F = \mu_t mpl_t + E_t \Lambda_{t,t+1} \left[(1 - \varrho(\tilde{a}_{t+1})) F_{t+1}^n - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right], \tag{4.10}$$

$$\underline{w}_t^H = mrs_t + b_t^u - (1 - \varrho(\tilde{a}_{t+1}) - p_{t+1}) E_t \Lambda_{t,t+1} H_{t+1}^n. \tag{4.11}$$

The equilibrium wage is then given by $w_t = (1-\eta)\overline{w}_t^F + \eta \underline{w}_t^H$, or equivalently by

$$w_t = (1-\eta)(mrs_t + b_t^u) + \eta \left(\mu_t mpl_t + E_t \Lambda_{t,t+1} \left[\kappa \theta_{t+1} - b_{t+1}^s (1 - \bar{\varrho}) F(\tilde{a}_{t+1}) \right] \right). \tag{4.12}$$

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)

[6] 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.

on the other. Higher unemployment benefits (b_t^u) 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.

4.5 Policy, aggregate resource and the government budget constraints

The government budget constraint satisfies $B_t = R_{t-1}B_{t-1} + b_t^u u_t + T_t^s$. Fiscal policy is governed by 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, and a specification for firing costs according to $b_t^s = \bar{\varsigma} + \varsigma w_{t-1}$, and 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.

Monetary policy is governed by a Taylor rule according to

$$i_t = \rho_i i_{t-1} + (1 - \rho_i)(\phi_\pi \pi_t + \phi_y \hat{y}_t) + \varepsilon_t, \tag{4.13}$$

where ε_t is the monetary policy shock, $i_t = \ln(R_t)$ and the hat-notation refers to the log-deviation from the steady state.

Finally, using the household and government budget constraints, and the expression for firms' profits, we obtain the aggregate resource constraint

$$y_t = c_t + \kappa v_t + F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1}v_{t-1})b_t^s. \tag{4.14}$$

This equation closes the model.

4.6 Embedding LMIs in the model

The LMIs discussed in Section 3 are reflected in the model through three key structural parameters: ς, η, and φ. The parameter η signifies workers' bargaining power, representing their advantage in wage negotiation processes. This can also be interpreted as a measure of union strength or the level of centralization in wage bargaining, with higher centralization typically favoring workers' position in wage bargaining. φ denotes the ratio of unemployment benefit payments to pre-dismissal wages, directly determined by governments. While comparatively less ambiguous than η and ς, φ exhibits significant variation across countries and over time, with some nations adjusting benefits based on crisis severity. Lastly, ς reflects the government's influence on employment protection, encompassing various mechanisms such as stringent layoff regulations, short-time work schemes, and potential severance payments.[7]

Mapping empirical LMIs to their theoretical counterparts in DSGE models remains qualitative. Quantitative data on firing costs, particularly regarding severance payments and notice periods, are often lacking at the country level. Furthermore, measures for union density and unemployment benefit replacement rates face similar challenges, as they may not fully capture non-monetizable elements of labor market institutions such as administrative and judicial procedures.

4.7 Equilibrium, model solution, and dynamic simulations

We collect the LMI parameters of interest in the vector $\boldsymbol{\vartheta} = [\eta, \varphi, \varsigma]$ and consider a log-linearized solution of the rational expectations model around its steady state,

$$\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})z_t = \boldsymbol{\Psi}_1(\boldsymbol{\vartheta})z_{t-1} + \varepsilon_t \tag{4.15}$$

where the vector z_t contains the endogenous variables and the scalar ε_t is the monetary policy shock as of equation (4.13). The matrix $\boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$ governs the dynamics among the dependent variables and the vector $\boldsymbol{\Psi}_0(\boldsymbol{\vartheta})$ determines the contemporaneous impact of the monetary policy shock on the endogenous variables. We assess the extent to which each of the three parameters in $\boldsymbol{\vartheta}$ shapes the response of the endogenous variables to a monetary policy shock by computing impulse response functions (IRFs) based on a calibration of the model's parameters as outlined in Table A1 in Appendix A.2.

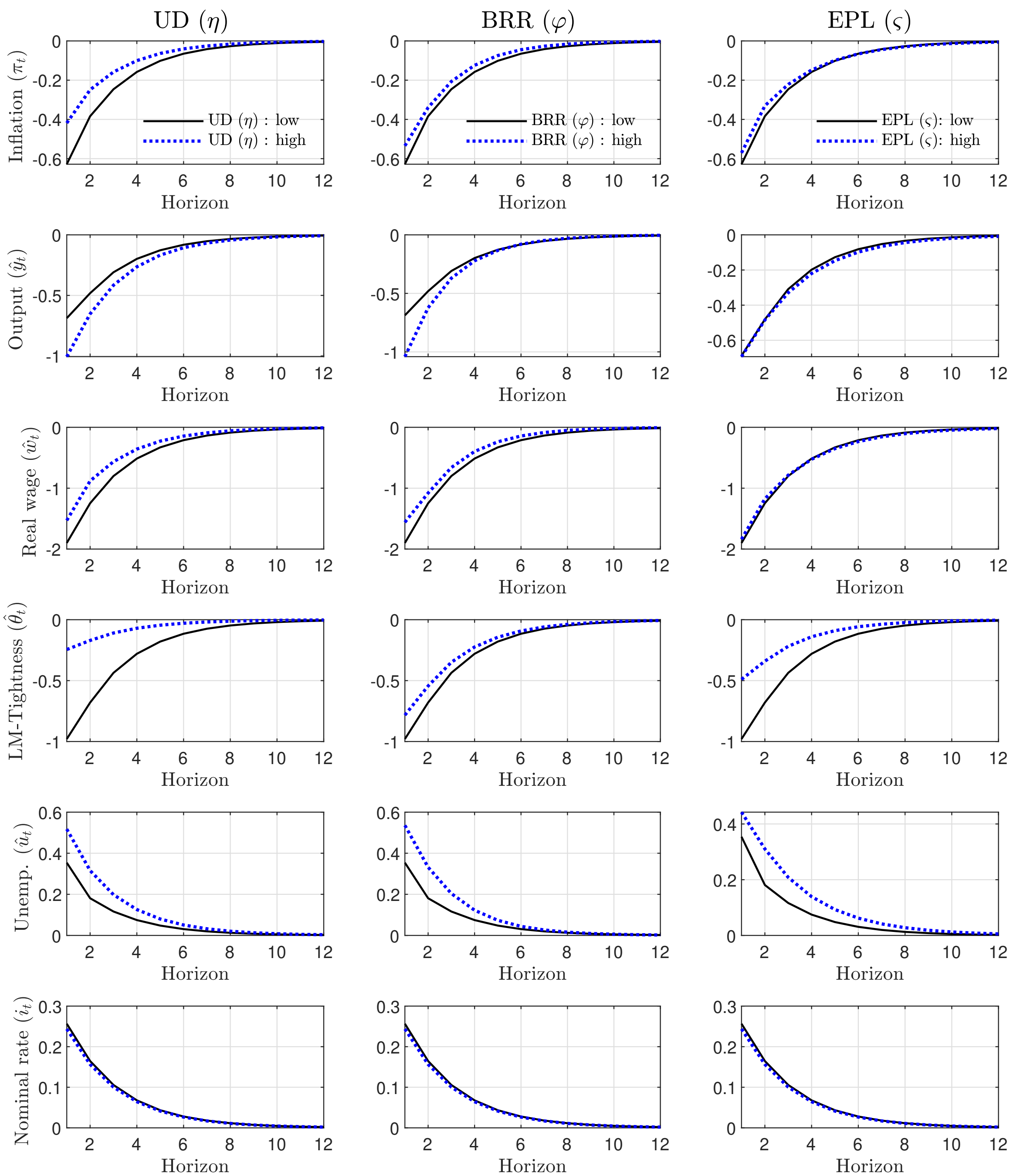
As the IRFs are continuous functions of $\boldsymbol{\vartheta}$, we can display them over a whole range of values of $\boldsymbol{\vartheta}$. This is shown in Figures 2 and 4. The former shows the whole path of the IRFs for two distinct values of the LMIs (high vs. low value). The latter, in turn, shows the value of the IRFs for two distinct horizons solely, however, for wide range of the LMIs.

As highlighted in the figures, the monetary tightening causes a drop in inflation, output, and the real wage: The increase in the nominal rate raises the real rate, which induces households to substitute present consumption for future consumption. The resulting drop in current consumption reduces output. This is accompanied by a drop in the demand for labor, which exerts downward pressure on the real wage.

The responses of the low values of the LMIs (black solid lines) use the values as depicted in Table A1. The high values for the LMIs add 30 basis points to each LMI's value of the "low" state. Importantly, the LMIs

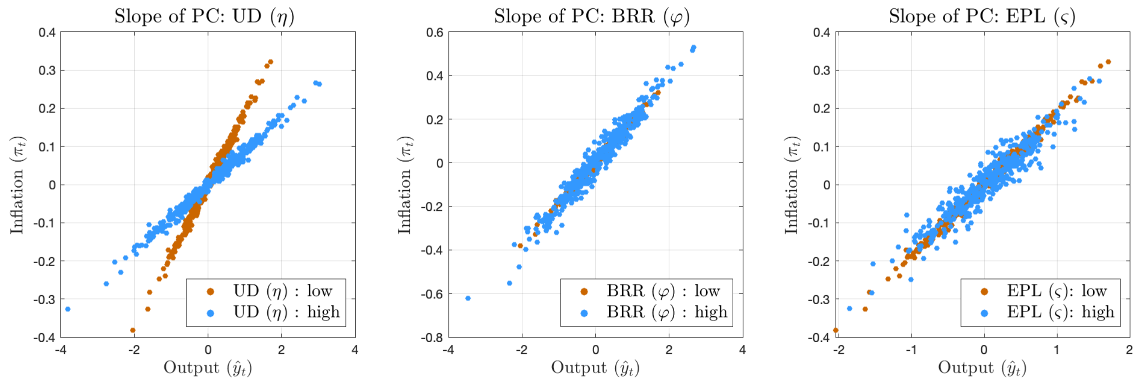
[7] This parameter also includes one-time payments to the social security system, commonly observed in many countries to offset unemployment insurance costs resulting from dismissals.

Figure 2: Impulse response functions of the baseline model.



Notes: Each subplot illustrates the impulse response functions following a contractionary monetary policy shock, contingent upon two levels of the LMIs.

Figure 3: Slope of the Phillips curve.



Notes: Dynamics in the baseline model with less (brown) and more (blue) stringent LMIs in response to a monetary policy shock. The x-axis depicts output ($\hat{y}_t$) and the y-axis inflation (π_t) both as percentage point deviation from their steady states.

do not qualitatively alter the path of the impulse response functions, though there are sizeable quantitative differences.

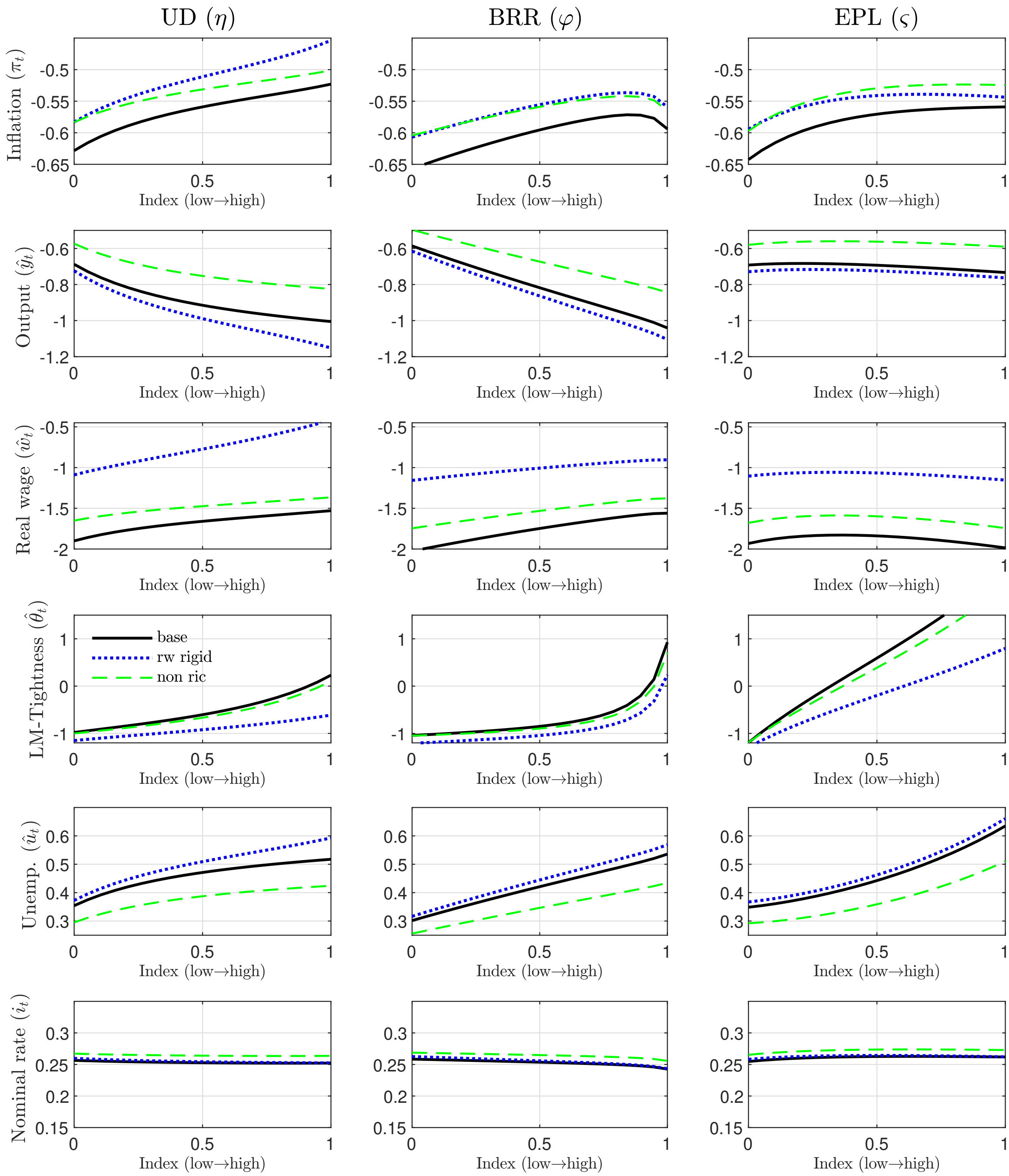
For each LMI in $\boldsymbol{\vartheta}$, we have that the parameter space is bounded between zero and unity. Hence, the differences in the IRFs from low and high values of the LMIs can be compared directly across the LMIs.

4.8 Implications of the LMIs for monetary policy efficacy

The IRFs across low and high values of the LMIs highlight a particular trade-off: For low values of the LMIs, the reaction of inflation is pronounced while the one of output is moderate. The opposite applies for high values of the LMIs, where output reacts strongly, while inflation only reacts moderately. This suggests that the LMIs crucially affect the slope of the Phillips curve, which is illustrated in Figure 3. Higher values of the LMIs lead to a flattening of the Phillips curve, which shapes the quantitative effect of the demand contraction induced by an interest rate hike. It renders the output reaction more pronounced, while the inflation response is flattened at the same time. As can be seen in Figures 3 and 2, this turns out to be pronounced in the case of UD, while only to a minor extent in the case of the BRR and EPL. As Figure 4 highlights, changes in the LMIs have unique effects on inflation, output and the real wage in case of UD only, while in case of BRR and EPL a concave pattern emerges for some variables.

To better understand how the LMIs affect the monetary policy transmission mechanism, it is useful to consider the reservation wages of households and firms ($\underline{w}_t^H$ and $\overline{w}_t^F$). With respect to the household, the drop in current consumption (hence increase in marginal utility of consumption) and the rise in leisure (or the drop in employment; hence decrease in marginal utility of leisure) leads to a drop in the marginal

Figure 4: Sensitivity of the impact responses to the LMIs.



Notes: Each subplot illustrates the impulse response function's value one month after a contractionary monetary policy shock, contingent upon various levels of the LMIs. The solid black line (*base*) represents the model delineated in Section 4, while the dotted blue lines (*rw rigid*) introduce real wage rigidity, and the dashed green lines (*non ric*) incorporate non-Ricardian households which are characterized by limited asset market participation (both extensions are detailed in the Appendix).

rate of substitution between consumption and leisure ($mrs_t = -u_{n,t}/u_{c,t}$), which reduces the household's reservation wage ($\underline{\mathrm{w}}_t^H$). With a view to the (intermediate-goods producing) firm, the increase in the interest rate reduces the present-value of employing an additional worker, which hence reduces the firm's reservation wage ($\underline{\mathrm{w}}_t^F$). The real wage hence declines. The decline therein is stronger the higher is the firm's bargaining power $(1-\eta)$. As a consequence, when η takes on a small value, the drop in the real wage is stronger. This in turn exerts a more positive effect on the value of employment (F_t^n), which mitigates the drop in employment and hence in output. On the other hand, the stronger drop in the real wage when η is low exacerbates the decline in marginal costs causing a more pronounced reduction in the inflation rate. This explains the differences in the impulse response functions in the left-hand side panels in Figures 2 and 4.

Changes in the BRR and EPL lead to similar effects on the transmission mechanism of monetary policy shocks. However, in the case of EPL, the quantitative impact is negligibly small, while at the same time the overall influence is uncertain. In the Appendix, we present several extensions to the theoretical analysis, including an analysis based on a more general calibration and model extensions that include real wage rigidity and non-Ricardian households. The implication of all these exercises is that the main results presented here prevail as can be seen in Figure 4.

5. The empirical model

We empirically validate our theoretical model by examining the conditional response to monetary policy shocks using an interacted panel vector autoregressive (IPVAR) model. The IPVAR model is used to estimate how the matrices $\mathbf{\Psi}_1(\boldsymbol{\vartheta})$ and $\mathbf{\Psi}_0(\boldsymbol{\vartheta})$ of the system given by equation (4.15) depend on the structural parameter of interest. We consider a first-order approximation of these matrix functions around the sample mean of each interaction variable $\vartheta_l \in \boldsymbol{\vartheta}$, given by $\bar{\vartheta}_l$

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

In the IPVAR specification the regressors are not only covariates at different lags, but also the interaction terms formed by them. The response coefficients are thus allowed to change deterministically with the interaction terms, which in our case capture structural characteristics of the labor market. This implies that the impulse response functions can be evaluated for different constellations of LMIs in order to empirically validate the theoretical results presented earlier.

There are two assumptions involved in the empirical approach used here. First, the interaction variables are assumed to be exogenous with respect to business cycle shocks. We defend this assumption as follows. While the LMIs show a reasonable variation over time, this is limited to changes over a long-term horizon, implying that the LMIs are unlikely to respond to monetary policy shocks that affect only the short run.[8]

The second assumption is the usual linearity assumption embedded in the IPVAR specification used to mimic the approximation in equation (5.1). In principle, the linearity assumption could be relaxed by considering various nonlinear combinations of the LMIs in the IPVAR model. However, depending on the number of observations and the parameters of interest in the estimation, overfitting the model becomes a problem in our setting, so we stick to linear specifications with interactions specified in this form rather than evaluating more complex nonlinear parameterizations of the model.

Before we describe the econometric framework in detail, we discuss the data. We use a monthly data set for the time period 1999M1 to 2023M6 both at the aggregate euro area level and at the individual country level for nine selected member countries. Euro area refers to the (current) twenty member countries. At the euro area level, we have a $n_e \times 1$ vector of endogenous variables, $\boldsymbol{e}_t$, consisting of high-frequency interest and stock market surprises (defined further below), core CPI (log-level), the real wage (log-level), real GDP (log-level), the shadow rate (Wu and Xia, 2016) (level), and the term structure (difference between the yields on short-term (3 months) and long-term (10 years) government bonds of (aggregated) euro area countries, level). For a set of $N = 9$ individual countries, we have a $n_y \times 1$ vector of endogenous variables, $\boldsymbol{y}_{it}$ ($i = 1, \ldots, N$), including core CPI (log-level), the real wage (log-level), and real GDP (log-level). The real wage is calculated as compensation of employees per employee and deflated by the GDP deflator.[9] The countries in our sample are Austria, Belgium, Finland, France, Germany, Italy, the Netherlands, Spain, and Portugal.

5.1 Econometric framework

In the econometric framework, we pursue a block exogenous structure in which the euro area variables are exogenous for the individual member countries. This is motivated by the small open economy characteristic

[8] An empirical evaluation confirms this. The result is that, using the LMIs as additional endogenous variable, a tightening of monetary policy causes negligible changes in the LMIs (as measured by the IRFs), with the IRF of each LMI not statistically different from zero. We interpret this evidence as supporting the assumption that the LMIs are exogenous to monetary policy shocks.

[9] Compensation of employees, GDP and the GDP deflator are available at a quarterly frequency. We have disaggregated compensation of employees to a monthly frequency using collective wage indices; for GDP we have used industrial production and retail sales indices and for the GDP deflator we have used the core CPI and the producer price index (PPI).

of each member country relative to the euro area block. This yields the following system of equations for the euro area, $\boldsymbol{e}_t$, and individual country, $\boldsymbol{y}_{it}$, variables:

$$\begin{pmatrix} \boldsymbol{A}_0^{11} & \boldsymbol{0} \\ \boldsymbol{A}_{i0}^{21} & \boldsymbol{A}_{i0}^{22} \end{pmatrix} \begin{pmatrix} \boldsymbol{e}_t \\ \boldsymbol{y}_{it} \end{pmatrix} = \sum_{j=1}^{P} \begin{pmatrix} \boldsymbol{A}_j^{11} & \boldsymbol{0} \\ \boldsymbol{A}_{ij}^{21} & \boldsymbol{A}_{ij}^{22}(\boldsymbol{\vartheta}) \end{pmatrix} \begin{pmatrix} \boldsymbol{e}_{t-j} \\ \boldsymbol{y}_{it-j} \end{pmatrix} + \begin{pmatrix} \boldsymbol{\mu}^e \\ \boldsymbol{\mu}_i^y \end{pmatrix} + \begin{pmatrix} \boldsymbol{\nu}_t^e \\ \boldsymbol{\nu}_{it} \end{pmatrix}, \tag{5.2}$$

where $\boldsymbol{A}_j^{11}$ are $n_e \times n_e$ and $\boldsymbol{A}_{ij}^{22}$ are $n_y \times n_y$ coefficient matrices ($j = 1, \ldots, P$) for the euro area and the individual country block, respectively. Lastly, the $n_y \times n_e$ matrix of coefficients models the transmission of the euro area aggregate on the individual country without allowing for any dynamic feedback to the euro area block. The joint error vector $\boldsymbol{\nu}_t = (\boldsymbol{\nu}_t^{e\top}, \boldsymbol{\nu}_{it}^{\top})^\top$ satisfies $\boldsymbol{\nu}_t \sim N(0, \boldsymbol{\Sigma})$, where $\boldsymbol{\Sigma}$ is a block diagonal matrix. This corresponds to $\boldsymbol{\nu}_t^e \sim \mathcal{N}(\boldsymbol{0}, \Sigma^e)$ and $\boldsymbol{\nu}_{it} \sim \mathcal{N}(\boldsymbol{0}, \boldsymbol{\Sigma}_i)$, which renders the two error vectors independent of each other. Note also that both $\boldsymbol{\Sigma}^e$ and $\boldsymbol{\Sigma}_i$ $(\forall i)$ are diagonal matrices containing only the volatilities of the respective variables since the model is written in recursive form. Hence, $\boldsymbol{e}_t$ is neither contemporaneously related nor Granger caused by $\boldsymbol{y}_{it}$.[10] We specify deterministics in the vector $\boldsymbol{\mu} = (\boldsymbol{\mu}^{e\top}, \boldsymbol{\mu}_i^{y\top})^\top$, which consists of a constant, a linear time trend, measures of global industrial production and commodity prices to control for global conditions, and a dummy variable to account for the turmoil surrounding the Covid-19 pandemic in early 2020 (Cascaldi-Garcia, 2022).

We analyze the role of the LMIs in mediating the effect of (monetary policy) shocks by allowing the coefficients of the $\boldsymbol{A}_{ij}^{22}(\boldsymbol{\vartheta})$ matrices to vary as follows

$$\boldsymbol{A}_{ij}^{22}(\boldsymbol{\vartheta}) = \boldsymbol{\Phi}_{ij}^{22} + \sum_{l=1}^{L} \boldsymbol{\Lambda}_{ij,l} \vartheta_{it,l}, \qquad \forall\, j \geq 1 \tag{5.3}$$

where ϑ_{it} refers to LMI l of country i at time t. $\boldsymbol{\Lambda}_{ij}$ is a coefficient matrix describing the impact of the LMI on the reduced form coefficients $\boldsymbol{A}_{ij}^{22}(\vartheta)$ in the model. To preserve computational tractability, we always consider only one out of the three interaction variables ($\boldsymbol{\vartheta}$) in the IPVAR model.[11] The model, consisting of the equations (5.2) and (5.3), is estimated using Bayesian techniques. We use a natural conjugate prior and draw all parameters jointly from the posterior.

We report the mean-group estimator of Pesaran and Smith (1995), which was used in an IPVAR setting by Towbin and Weber (2013) and Sá, Towbin and Wieladek (2014). The mean-group estimator of the country

[10] We assess this joint hypothesis by means of the Schwartz information criterion. The results indeed favor the assumption of the block exogeneity structure in the IPVAR model.

[11] In principle, all three interaction variables could be used simultaneously, but this quickly exhausts computing power.

block in the IPVAR is given by

$$\overline{A}_j^{2k}(\boldsymbol{\vartheta}) = \sum_{i=1}^{N} A_{ij}^{2k}(\boldsymbol{\vartheta})/N \qquad \forall k = 1, 2, \quad \forall j \geq 1. \tag{5.4}$$

We have equivalent expressions for $\overline{\boldsymbol{\Phi}}_j^{22}$ and $\overline{\boldsymbol{\Lambda}}_{j,l}$. We compute impulse response functions using these average coefficient estimates and interpret them as responses in a typical euro area country.

The elements describing the structural characteristics of the LMIs in the DSGE model shape the parameter matrices of the theoretical model ($\boldsymbol{\Psi}_j(\boldsymbol{\vartheta})$) and thus the transmission channel of the endogenous variables to exogenous shocks. The empirical IPVAR model treats these structural features in exactly the same way as they affect the parameter matrices of the theoretical model. It is instructive to examine the case of one lag ($P = 1$) a bit closer. The IPVAR representation corresponding to equation (5.1) is given by $\overline{A}_1^{22}(\boldsymbol{\vartheta}) = \boldsymbol{\Psi}_1(\boldsymbol{\vartheta})$, while the average effect is given by $\overline{\boldsymbol{\Phi}}_1^{22} = \boldsymbol{\Psi}_1(\overline{\boldsymbol{\vartheta}})$ and the partial derivative by $\overline{\boldsymbol{\Lambda}}_{1,l} = \partial \boldsymbol{\Psi}_1(\boldsymbol{\vartheta})/\partial \vartheta_l$. Finally, we have that $\overline{A}_0^{22} = \boldsymbol{\Psi}_0(\boldsymbol{\vartheta})$, which implies that the IPVAR coefficients capturing the contemporaneous relationship among the endogenous variables does not depend on the interaction variables.[12]

5.2 Identification

We measure unanticipated euro area monetary policy surprises via high-frequency proxies.[13] These high-frequency surprises are price changes of interest rate derivatives in a narrow window surrounding monetary policy announcements of the ECB and labeled as *surprises*.[14] These changes then reflect the exogenous changes in expectations solely due to announcement of the new monetary policy stance. Following Jarociński and Karadi (2020), we identify a *monetary policy shock* via a negative co-movement of surprises in interest rate derivatives and a stock market index.[15] We provide a brief exposition of the framework on how we

[12] We considered an extension in the form of allowing the A_{i0}^{22} coefficients to also depend on the interaction variables. However, every element in $\boldsymbol{\Lambda}_{i0,l}$ turned out to be close to zero and statistically indistinguishable from zero; and moreover, this extension did not change the results of the model with A_{i0}^{22} independent of the interaction variables. For these reasons, we have opted for the simpler and more parsimonious specification.

[13] See, inter alia, Kuttner (2001), Gürkaynak, Sack and Swanson (2005), Gertler and Karadi (2015), Nakamura and Steinsson (2018), Altavilla et al. (2019), Jarociński and Karadi (2020), Miranda-Agrippino and Ricco (2021), Bauer and Swanson (2023), and Badinger and Schiman (2023).

[14] Following Altavilla et al. (2019), we use the surprises in the *monetary event window* as the union in the surprises of the press release and the press conference window. Surprises in the press release/conference window are the difference between a pre- and post-release/conference window. In each window, they use the median price in a 10min window to cleanse the data of any misquotes. An exact timeline can be found in Altavilla et al. (2019).

[15] We follow the methodology in Jarociński and Karadi (2020) and Jarociński (2022) to construct the surprises in the interest rate derivatives as the first principal component of the surprises in interest rate derivatives with maturities up to 1 year. We use the Overnight Index Swaps (OIS) with maturities 1-, 3- and 6-months and 1-year. OIS rates capture market expectations of the future level of the Euro Overnight Index Average (EONIA). We rescale each interest rate surprise so that it has the standard deviation of the respective 1-year instrument. We do so for five high-frequency surprises: the current-month OIS rate, the 3-months OIS rate, and the eurodollar futures at the horizons of two, three, and four quarters, respectively. The advantage of this procedure is

implement this in the aggregate euro area model, in which we modify the aggregate element in equation (5.2). Let $\boldsymbol{e}_t = (\boldsymbol{m}_t^\top, \boldsymbol{w}_t^\top)^\top$ denote an $(n_m + n_w) \times 1$ vector of euro area aggregate variables. $\boldsymbol{m}_t$ is a vector of exogenous high-frequency instruments aggregated to a monthly frequency, while $\boldsymbol{w}_t$ consists of endogenous macroeconomic and financial variables on a monthly frequency. In our specific empirical setting, we have $n_w = 5$ and $n_m = 2$, which yields

$$\begin{pmatrix} \boldsymbol{m}_t \\ \boldsymbol{w}_t \end{pmatrix} = \underbrace{\begin{pmatrix} \mathbf{0} \\ \boldsymbol{c}^w \end{pmatrix}}_{=\boldsymbol{c}^e=(\boldsymbol{A}_0^{11})^{-1}\boldsymbol{\mu}^e} + \sum_{j=1}^{P} \underbrace{\begin{pmatrix} \mathbf{0} & \mathbf{0} \\ \boldsymbol{B}_j^{mw} & \boldsymbol{B}_j^{ww} \end{pmatrix}}_{=\boldsymbol{B}_j=(\boldsymbol{A}_0^{11})^{-1}\boldsymbol{A}_j^{11}} \begin{pmatrix} \boldsymbol{m}_{t-j} \\ \boldsymbol{w}_{t-j} \end{pmatrix} + \underbrace{\begin{pmatrix} \boldsymbol{\varepsilon}_t^m \\ \boldsymbol{\varepsilon}_t^w \end{pmatrix}}_{=\boldsymbol{\varepsilon}_t=(\boldsymbol{A}_0^{11})^{-1}\boldsymbol{\nu}_t^e}, \qquad \boldsymbol{\varepsilon}_t \sim \mathcal{N}\left(\mathbf{0}, \boldsymbol{\Omega}^e\right), \tag{5.5}$$

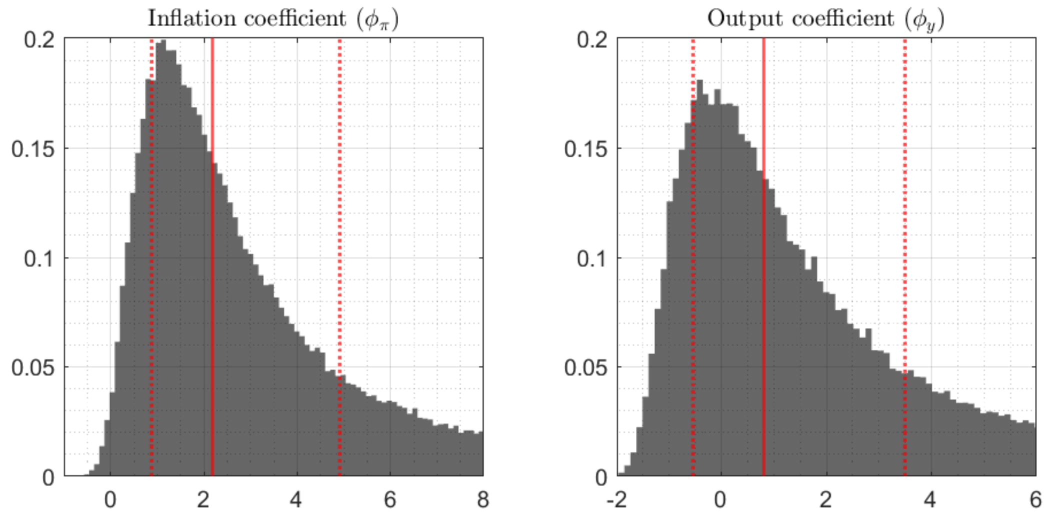
where $\boldsymbol{c}^w$ denotes the $n_y \times 1$ vector of constants, $\boldsymbol{B}_j^{ww}$ is the $n_y \times n_y$ matrix of coefficients for lag j, $\boldsymbol{B}_j^{mw}$ denotes an $n_m \times n_y$ matrix of coefficients, and $\boldsymbol{\Omega}^e = (\boldsymbol{A}_0^{11})^{-1}\boldsymbol{\Sigma}^e(\boldsymbol{A}_0^{11})^{-1\top}$ is an $n_e \times n_e$ covariance matrix. We assume that the surprises in $\boldsymbol{m}_t$ have zero mean and do not depend on the lags of either $\boldsymbol{m}_t$ or $\boldsymbol{w}_t$. These restrictions are plausible as long as financial market surprises are unpredictable.

For the identification of the monetary policy shock, we use sign restrictions on the proxies. The monetary policy shock is identified through a negative co-movement between the two proxies. Sign restrictions are implemented using the algorithm outlined in Rubio-Ramirez, Waggoner and Zha (2010). For that, we assume a uniform prior on the space of rotations conditionally on satisfying the imposed sign restrictions. This yields then a set-identified semi-structural model.

We estimate equation (5.5) using a two-step approach. First, we estimate only the upper part of the equation involving $\boldsymbol{m}_t$, followed by the lower part, involving $\boldsymbol{w}_t$. Although this may seem to introduce problems regarding the use of estimated regressors in the second step, we address this by incorporating the full posterior distribution of the upper part estimation into the lower part estimation. We do so by estimating the lower part for each draw of the posterior distribution of the monetary policy shock obtained from the upper part. This ensures that the uncertainty inherent in the estimation and identification of the monetary policy shock is fully accounted for in the estimation of the lower part. This reduces computational complexity quite *a lot*. The reason is that we do not have to impose sign-restrictions for the identification of monetary policy in each country-specific model given by (5.2) but only have to do this once.

that it partly also captures unconventional monetary measures, for instance, forward guidance. As stock market surprise, we use the surprises in the Euro Stoxx 50 index, which is a market capitalization-weighted stock market index including 50 blue-chip companies from 11 euro area countries. We obtain the data from the Euro Area Monetary Policy Event Study Database (Altavilla et al., 2019).

Figure 5: Estimated Taylor rule coefficients.



Notes: The sub-plots show the histograms of the estimated coefficients for inflation (ϕ_π) and output (ϕ_y) in the interest equation of the BVAR model. The red solid lines indicate the median estimate and the red dotted lines refer to the 16th and 84th percentile of the posterior distribution.

5.3 Empirical results

To examine whether the interactions with the LMIs affect the dynamics of the endogenous variables in the IPVAR model, we assess the posterior distribution of the impulse response functions (IRFs) at different levels of the interacting covariates and compute IRFs and forecast error variance decompositions (FEVDs). Since we always use one of the three LMIs as the interaction variable at a time, the impulse response functions are obtained after varying the interaction variable between low and high values, which we set to the values of the 10th and 90th percentiles of the sample distribution, as shown in the lower subpanels of Figure 1. Specifically, the coefficient matrices for a typical country with a high share of a particular LMI are $\overline{\boldsymbol{A}}_j^{22}(\boldsymbol{\vartheta})\Big|_{\vartheta_l^{\text{High}}} = \overline{\boldsymbol{\Phi}}_j^{22} + \overline{\boldsymbol{\Lambda}}_{j,l}\vartheta_l^{\text{High}}$ for LMI l and at lag j. Similarly, the coefficient matrices for a typical country with a low value of a particular LMI are $\overline{\boldsymbol{A}}_j^{22}\Big|_{\vartheta_l^{\text{Low}}} = \overline{\boldsymbol{\Phi}}_k^{22} + \overline{\boldsymbol{\Lambda}}_{j,l}\vartheta_l^{\text{Low}}$. We proceed in the same manner when evaluating any of the three LMIs in shaping the impulse response functions and in decomposing the forecast error variance. All estimations are performed with a lag length of one, as suggested by the BIC. We analyze 8,000 posterior draws after discarding the first 2,000 as burn-ins.

Before we turn to the analysis of the IRFs, we inspect the outcomes of the global model in more detail. Our empirical setup allows us to estimate values for the structural parameters of inflation and output in the interest rate equation, commonly referred to as the Taylor rule coefficients. To this end, we follow Baumeister and

Hamilton (2024) and estimate structural coefficients using external instruments.[16] The results are presented in Figure 5. The coefficients in the interest rate equation are normalized such that the coefficient for the shadow rate is set to unity, while those for the inflation rate (ϕ_π) and output (ϕ_y) are adjusted for the sign. The figure illustrates the posterior density, along with the 50th percentile (solid red line) and the 16th and 84th percentiles (dotted red lines).

As shown in Figure 5, the median estimate for the inflation coefficient (ϕ_π) is 2.2, and it is statistically significantly different from zero, based on a 68 percent confidence interval. Moreover, the magnitude of this estimate is substantial, indicating alignment with the Taylor principle. In contrast, the estimate of the coefficient for output exhibits quite large uncertainty. Its median estimate is low (0.8) and it is not statistically significantly different from zero, given a 68 percent confidence interval. However, both estimates align roughly to conventional parameter choices in DSGE models for the interest rate reaction function. We also note that the uncertainty in those coefficients arise from two sources: parameter uncertainty of the estimation and also uncertainty regarding the identification of monetary policy.

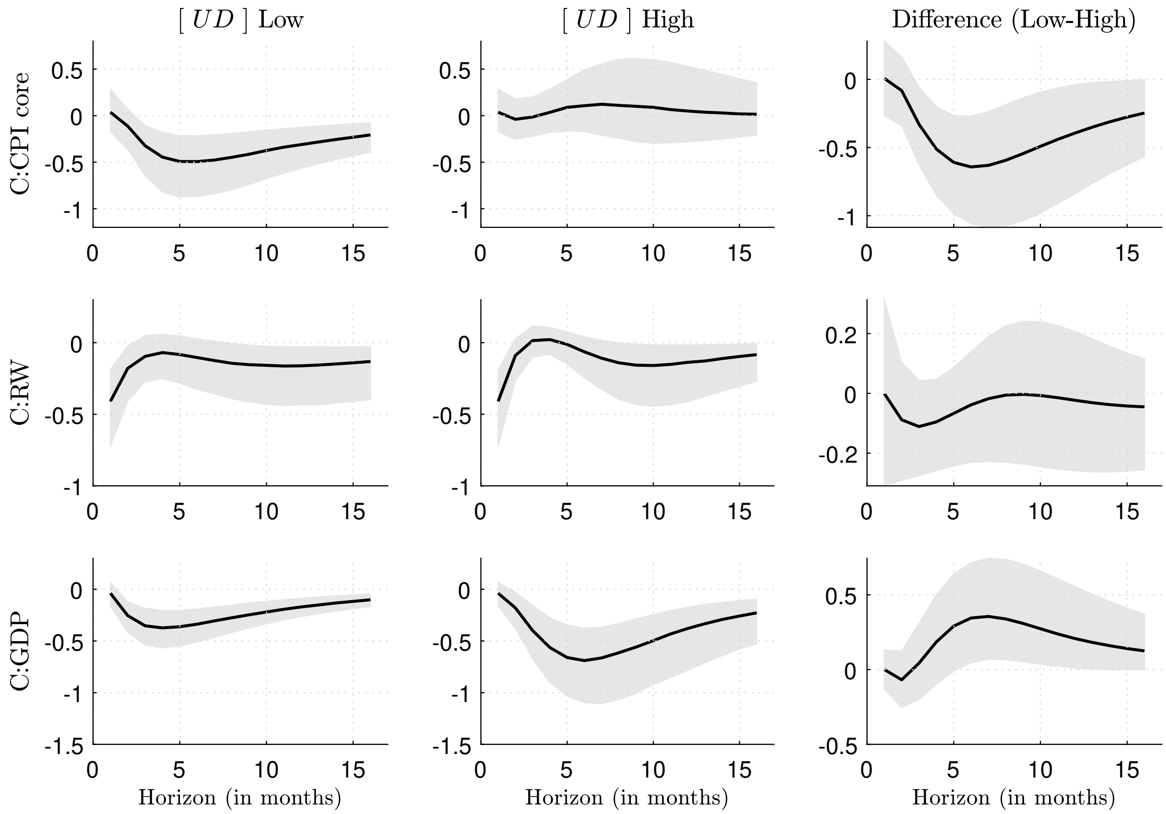
Figures 6-8 show the IRFs of real GDP, real wage, and core CPI of the individual countries to the monetary policy shock for different levels of the LMIs. Each figure contains three columns: the first column shows the IRFs and 68 percent credible intervals from the posterior when the relevant interaction variable is at a low level (10th percentile); the second column shows the IRFs when the interaction variable is at a high level (90th percentile); and the third column shows the difference between the two previous IRFs.[17] This allows us to assess whether variations in the interaction variables have a significant impact on the dynamic adjustment to the shock.

We investigate the outcomes of a monetary policy shock, normalized to a 25 bps increase in the shadow rate on impact. We report the group-mean estimator of the individual countries, while the IRFs of the euro area aggregates can be found in the appendix, see Figure C1. To summarize the results on the euro area

[16] We follow their procedure as follows. Given that the vector $\hat{\boldsymbol{h}}_1$ is a single column of the structural impact matrix, and defined by $\boldsymbol{\varepsilon}_t = \boldsymbol{H}\boldsymbol{u}_t$ where $\boldsymbol{u}_t$ is a vector of structural shocks while $\boldsymbol{\varepsilon}$ is the vector of reduced-form innovations, Baumeister and Hamilton (2024) show that we can estimate the coefficients of the structural equation (up to an unknown constant g and sign convention) by defining $\hat{\boldsymbol{v}}_1 = (\boldsymbol{I} - \sum_{j=1}^{P} \hat{\boldsymbol{B}}_j)^{-1}\hat{\boldsymbol{\Omega}}^e \hat{\boldsymbol{h}}_1 g$. We add the first term to allow for interest rate smoothing in the structural equation. We arrive at an estimate of the structural coefficients by using $\hat{\boldsymbol{a}}_1^\top = \hat{\boldsymbol{v}}_1^\top / (\hat{\boldsymbol{v}}_1^\top \boldsymbol{e}_i)$, where $\boldsymbol{e}_i$ is the column i of the identity matrix of dimension n_e with i being the position of the shadow rate in the vector of endogenous variables.

[17] The IRFs for low and high values of the interaction variable are correlated. We follow Abbritti and Weber (2018) and compute a test statistic using the impulse responses from draws of the posterior parameters. For each of the 8,000 posterior draws, we compute the differences between the response of each variable to a monetary policy shock under different values of the interaction variable of interest. This yields a distribution of the difference in responses, which can then be used to compute credible intervals. The IRFs for low and high values of a given interaction variable can be considered statistically different from each other if the credible interval of their difference (shown in the third column) is above or below zero.

Figure 6: Impulse response functions: union density (UD).

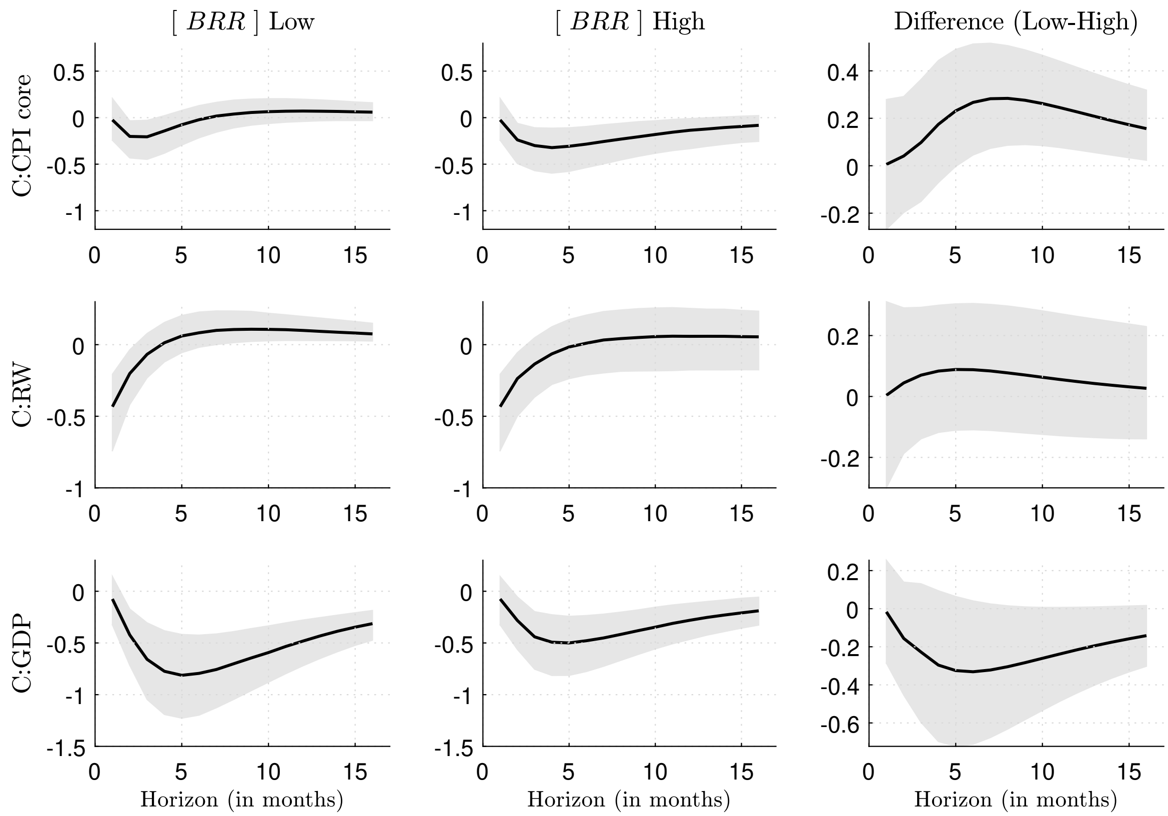


Notes: The sub-plots in the first column depict impulse response functions when the UD has a low level (10th percentile), the second column when the UD has a high level (90th percentile), and the third column reports the difference between a low and high value. Black solid lines denote the median response, while the gray dashed areas are the 68 percent posterior credible intervals. We report the group-mean estimator of the individual countries ("C:"). The horizontal axis measures time in months. The vertical axis measures deviation from pre-shock level in percent.

level, the monetary policy shock elicits an increase in the shadow rate causing a decline in core CPI, the real wage, and real GDP. The risk premium given by the term premium increases. Across all LMIs, we find that a contractionary monetary policy shock at the euro area aggregate transmits to an average euro area country expected according to the theoretical model. We observe a decline in output (real GDP), prices (core CPI), and the real wage. From a quantitative point of view, the responses show significant variation across different values of the LMIs.

Figure 6 reports the results for the sensitivity of the monetary policy transmission mechanism with respect to the union density (UD). When union density is high, a 25 bps increase in the shadow rate induces a drop in real GDP of approximately 0.7 percent after five months and a strong decline by 0.4 percent on impact of the real wage. Core CPI barely reacts when we assume a high value of UD. These values are in line with

Figure 7: Impulse response functions: benefit replacement rate (BRR).

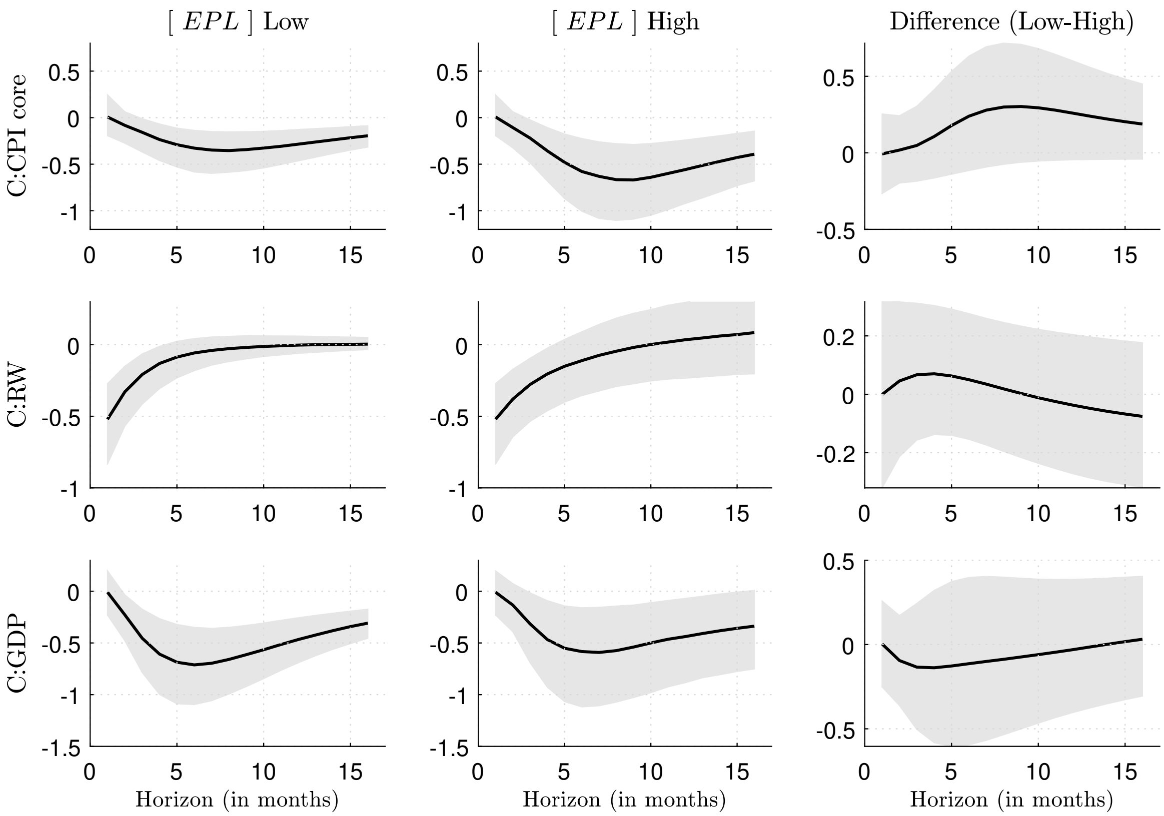


Notes: The sub-plots in the first column depict impulse response functions when the BRR has a low level (10th percentile), the second column when the BRR has a high level (90th percentile), and the third column reports the difference between a low and high value. Black solid lines denote the median response, while the gray dashed areas are the 68 percent posterior credible intervals. We report the group-mean estimator of the individual countries ("C:"). The horizontal axis measures time in months. The vertical axis measures deviation from pre-shock level in percent.

estimates presented in other studies, such as Cantore, Ferroni and León-Ledesma (2021), who document an output contraction of approximately 0.5 percent in the euro area in response to a 20 bps rise in the nominal rate. Similarly, Glocker and Piribauer (2021) provide an estimate of 0.6 percent in this context.

These estimates exhibit noticeable variations when UD assumes a low value (left-hand side panels). Output displays a much more muted reaction, peaking only at −0.4 percent after three months. The price and wage responses are, however, more pronounced when imposing a low value of UD. Core CPI shows a 0.5 percent decline before gradually returning towards the pre-shock level. The real wage shows a somewhat more sluggish returning to the pre-shock level. Crucially, the disparities in the IRFs between low and high values of UD are statistically significant for both prices and output, as illustrated in the right-hand side panels. We interpret this as compelling evidence supporting the notion that union density influences the transmission

Figure 8: Impulse response functions: employment protection legislation (EPL).



Notes: The sub-plots in the first column depict impulse response functions when the EPL has a low level (10th percentile), the second column when the EPL has a high level (90th percentile), and the third column reports the difference between a low and high value. Black solid lines denote the median response, while the gray dashed areas are the 68 percent posterior credible intervals. We report the group-mean estimator of the individual countries ("C:"). The horizontal axis measures time in months. The vertical axis measures deviation from pre-shock level in percent.

channel of monetary policy shocks. Specifically, when union density is high, alterations in the monetary policy stance yield strong effects on prices and wages and provoke substantial effects on output. Conversely, when union density is low, output responds weakly while price effects are more pronounced. This empirical evidence aligns with the implications of the theoretical model.

Figure 7 illustrates the findings regarding the sensitivity of the transmission mechanism of monetary policy concerning the benefit replacement rate (BRR). Once more, the IRFs exhibit qualitative alignment with those predicted by the theoretical model: output and prices decline, as does the real wage. The point estimates closely mirror those of the IRFs depicted in Figure 6 for union density. Similarly to before, Figure 7 shows a trade-off between output and prices. While the IRF of the differences is only statistically significant for prices, the one for output shows high uncertainty. Still, the median clearly points to stronger output effects

Table 1: Forecast error variance decomposition

	CPI (core)		*Real wage*		*GDP*	
Horizon	Low	High	Low	High	Low	High
Union density (UD, η)						
2	0.1	0.1	0.1	0.1	0.1	0.1
4	1.3	1.3	0.2	0.2	0.3	0.3
8	3.0	1.0	0.4	0.4	1.2	2.9
12	4.5	1.2	0.6	0.7	1.4	4.4
Benefit replacement rate (BRR, φ)						
2	0.0	0.0	0.0	0.0	0.0	0.0
4	0.3	0.3	0.2	0.1	0.1	0.1
8	1.0	1.0	0.4	0.3	1.8	2.0
12	1.5	1.4	0.5	0.5	2.3	2.3
Employment protection legislation (EPL, ς)						
2	0.0	0.0	0.0	0.0	0.0	0.0
4	0.3	0.3	0.0	0.0	0.0	0.0
8	1.0	0.9	0.0	0.0	1.0	0.9
12	1.5	1.4	0.1	0.1	2.0	1.8

Notes: The values refer to the median of the corresponding posterior distribution of the IPVAR model. *Horizon* is measured in *months*. The columns *Low* and *High* refer to values of the interaction terms below and above the 50th percentile of the cross-country distribution in the interaction variables (UD, BRR, EPL) used for the computation of the forecast error variance decomposition.

in the model with low BRR. Again, we do not find strong differences for the real wage. Consequently, based on these estimates, it can be inferred that the benefit replacement rate hardly exerts a relevant influence on the transmission mechanism of monetary policy.

Similar conclusions are drawn in the analysis of employment protection legislation (EPL). The estimates are presented in Figure 8. As was observed with the benefit replacement rate (BRR), the IRFs in the case of EPL exhibit qualitative alignment with those anticipated by the theoretical model: both output and prices decline, alongside a decrease in the real wage. Furthermore, the point estimates closely resemble those observed in the IRFs depicted in Figure 6 for union density. However, when using EPL as the interaction variable, median responses of the differences are in line with the theoretical exercise but the IRFs are not statistically indistinguishable across low and high values of the interaction variable. This lack of difference is underscored in the right-hand side panels of Figure 8. Consequently, we interpret this as evidence supporting the absence of an effect of EPL on the transmission mechanism of monetary policy.

Table 1 reports the results of the forecast error variance decomposition (FEVD) of the IPVAR model along two dimensions: (i) for different horizons (two, four, eight and twelve months) and (ii) for two different values of the LMIs. The results are again presented only for the country block variables (core CPI, real wage, and GDP) of the IPVAR model.

Monetary policy shocks account for a relatively small share of the variation in (country-specific) GDP, explaining less than two percent when the horizon considered is short and UD is low (upper part of the table). However, when UD is high, more than four percent is explained. While this is still a seemingly low figure for the variation in GDP explained by monetary policy shocks, it is significantly larger than in the case of low UD. This confirms the results presented before, which indicate that the impact of monetary policy shocks on GDP increases with the level of UD. Conversely, the core CPI shows a contrasting pattern, with more than four percent of its variation attributable to monetary policy shocks when UD is low. This share declines significantly as UD increases. Finally, the amount of variation in the real wage explained by monetary policy shocks is largely unaffected by UD.

Looking at the results for the BRR and EPL (middle and bottom segments of the table, respectively), there is little difference in the variance decomposition of the forecast error for low and high values of these labor market indicators (LMIs). This confirms the results presented earlier, suggesting that these variables have a more negligible impact on the transmission channel of monetary policy.

In Appendix C, we report a number of robustness checks to the empirical model. We investigate cross-correlations to a number of possible omitted variables and conduct sub-sample stability checks. The outcomes prevail when only investigating the second half of the sample. Lastly, we conduct checks by replacing and adding endogenous variables, exogenous variables, or using different measures for prices. Results are robust to these choices.

6. Conclusion

In this paper, we provide evidence that structural features of the labor market affect the transmission mechanism of monetary policy based on the analysis of a theoretical and the results of an empirical model. The structural labor market indicators that we examine in the analysis are limited to (i) union density, (ii) the benefit replacement rate and (iii) employment protection legislation. All three labor market indicators exhibit significant variation across countries and time. However, temporal variations are long-term oriented and hence go beyond the length of typical business cycles. As a consequence, if these labor market indicators were to affect the monetary policy transmission mechanism, then this constitutes another structural feature in shaping the former.

To sum up, our theoretical model shows that the structural labor market indicators affect the slope of the Phillips curve. In particular, higher levels of union density flatten the Phillips curve. This, in turn, implies that when union density is high, a monetary contraction has a large impact on output, but only a muted one on inflation. The opposite is true when union density is low. No clear-cut pattern emerges for the remaining two labor market indicators (benefit replacement rates and employment protection legislation). While they both tend to affect the monetary transmission channel too, U and inverted-U turn patterns emerge.

Our empirical analysis largely corroborates these theoretical propositions. Employing a block exogenous vector autoregressive (VAR) model, labor market indicators are incorporated as interaction terms, enabling the assessment of impulse response functions across various labor market conditions. Our results indicate that monetary tightening exerts downward pressure on prices, output, and the real wage. Notably, union density emerges as a significant determinant in shaping the transmission mechanism of monetary policy. Specifically, heightened union density accentuates the impact on output while tempering the effect on prices, while, diminished union density exhibits the opposite effect. Conversely, the benefit replacement rate and employment protection legislation exhibit negligible influence on the transmission mechanism.

These findings underscore the pivotal role of labor market characteristics as structural elements in shaping the transmission mechanism of monetary policy. They hold noteworthy implications, particularly in the context of a monetary union. When member states possess heterogeneous labor markets, differential transmission of shocks ensues, resulting in asymmetric effects of symmetric shocks, potentially leading to inefficient inflation and output differentials.

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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.

Production, costs and prices

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

Households

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

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

Budget and resource constraints

- $B_t = R_{t-1} B_{t-1} + b^u_t u_t + \bar{T}^s_t$
- $y_t = c_t + \kappa v_t + F(\tilde{a}_t)(1 - \bar{\varrho})(n_{t-1} + q_{t-1} v_{t-1}) b^s_t$

Policy

Table A1: Calibration of the model.

Parameter	Description	Value	Range
β	Discount factor	0.992	[0.95 – 0.999]
γ	Elasticity of matching of unemployed persons	0.68	[0.05 – 0.95]
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]
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]
ϕ_π	Inflation sensitivity in the Taylor rule	1.5	[1.0 – 2.5]
ϕ_y	Output sensitivity in the Taylor rule	0.5	[0.0 – 1.0]
ρ_i	Nominal interest rate smoothing	0.6	[0.0 – 0.9]
ξ	Calvo price stickiness	0.7	[0.45 – 0.95]
η	Bargaining power of workers (UD)	0.2	—
φ	Unemployment benefit replacement rate (BRR)	0.0	—
ς	Firing costs in relation to last wage (EPL)	0.0	—

- $i_t = \rho_i i_{t-1} + (1-\rho_i)(\phi_\pi \pi_t + \phi_y \hat{y}_t) + \varepsilon_t$ with $\varepsilon_t \sim N(0, \sigma_\varepsilon^2)$

- $b_t^s = \bar{\varsigma} + \varsigma w_{t-1}$, $b_t^u = \varphi_{b^s} w_{t-1}$ and $\bar{T}_t^s = \bar{T}^s + \varphi_{T^s} B_t$

where $mpl_t = y_t/n_t$ is the marginal product of labor, and a is log-normally distributed of which F is the cumulative density function (c.d.f.). A hat over a variable signifies the log deviation from its steady state value. The particular functional form of the instantaneous utility function is given by: $u(c,n) = \frac{c^{1-\sigma}(1+(\sigma-1)\phi n)^\sigma - 1}{1-\sigma}$.

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.

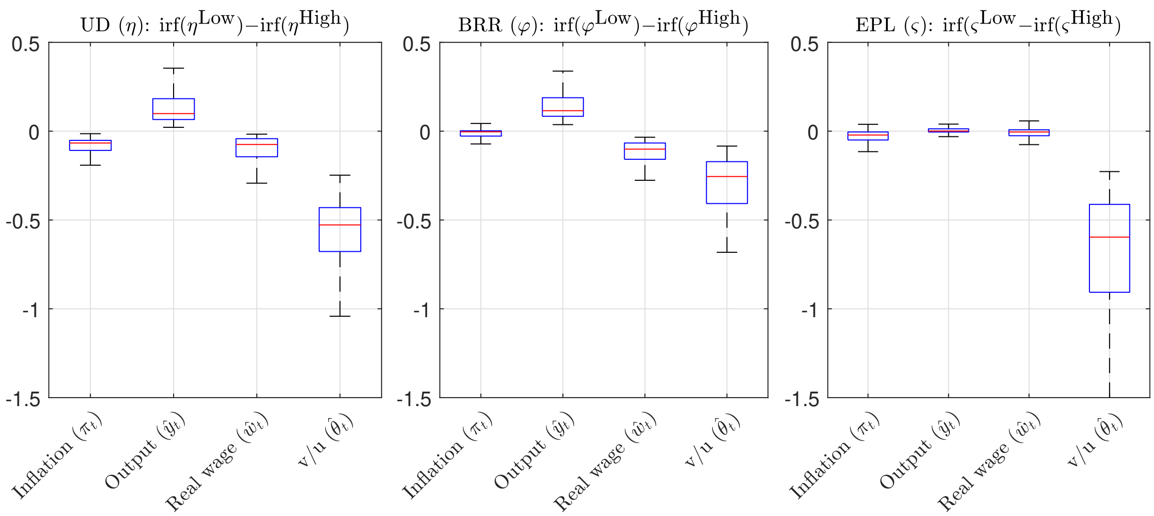
In particular, from $\bar{m} = p/\theta^{1-\gamma}$ we get the probability of a vacancy being filled $q = \bar{m}\theta^{-\gamma}$, the number of employed and unemployed persons $n = (1 - \varrho(\tilde{a}))p/((1 - \varrho(\tilde{a}))p + \varrho(\tilde{a}))$ and $u = 1 - n$, the number of vacancies posted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

Using equations (4.2), (4.3) and (4.12) 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 $(\varphi, \eta, \beta, \bar{\varrho}, \zeta, ...)$. 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 $c = y - \kappa v$. 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 $\bar{T}^{s} = -\varphi_{b^{s}} w(1 - n)$. $\bar{T}^{s} < 0$, it can be interpreted as lump-sum taxes.

Our benchmark calibration is summarized in Table A1. Given that our focus is on the role of the LMIs in the transmission of fiscal spending shocks, we do not calibrate our model to a particular economy. We closely follow Christoffel, Kuester and Linzert (2009) for the choice of the values of the structural parameters. They estimate a DSGE model with an extended labor market structure in their model based on the data for the euro area. Since the countries in our sample are part of the euro area, we rely on the estimates in Christoffel, Kuester and Linzert (2009). The complementarity coefficient σ in the households' instantaneous utility function $u(c, n)$ is set to one, which corresponds to the separable utility case. As a final remark, it is important to emphasize that the detailed description of the procedure for computing the steady state should reveal a key insight: changes in the LMIs not only impact the short-run equilibrium, as captured by the impulse response functions to shocks, but also influence the long-run equilibrium, that is, the steady state.

Figure A1: Impulse responses from the theoretical model: sensitivity assessment.



Notes: The figure shows the difference in the impulse response functions at low and high values of the LMIs of inflation, output, real wage, and labor market tightness to a monetary policy shock for different calibrations of the underlying structural parameters.

A.3 LMIs and the monetary transmission channel – a more general calibration

The impulse response functions (IRFs) of the theoretical model proposed in Section 4 are shaped by the specific calibration chosen. Consequently, an alternative calibration could yield markedly different outcomes. We leverage this notion as the foundation for investigating the robustness of our theoretical model's principal findings across a broader spectrum of calibrations. Throughout this inquiry, we center our attention on discerning the sensitivity in variables' reactions to a monetary policy shock as the LMIs vary. To this end, we proceed as follows.

We consider a continuum of values for all structural model parameters other than the LMIs (η, φ and ς). We simulate the model over a wide range of different values for these 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 LMIs (η, φ and ς). For instance, in the case for η: for a particular draw of the structural parameters (except the three LMIs), we solve the model for $\eta = 0.2$ and compute impulse response functions. For the same draw of the structural parameters (except the three LMIs) we then also solve the model using $\eta = 0.7$. 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.

We focus on the impact responses. The three scenarios (UD, η; BRR, φ; and EPL, ς) are depicted in the sub-panels in Figure A1. Each box-plot shows the difference in the impact response for each of the three cases for the following variables: output, inflation, 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 a LMI of interest relative to a high value of the same LMI.

The sub-panel on the left confirms the results presented in Section 4 within a more general calibration of the model. Changes in union density (UD) underscore a discernible trade-off in the inflation-output dynamic. In response to monetary tightening, the decline in inflation is particularly pronounced under high levels of UD, while it remains more moderate when UD is low. Conversely, the degree of output contraction increases as the UD level rises. Moreover, this sub-panel highlights two key aspects. First, the striking deviation of the impact responses from zero underscores the distinctiveness of the results explained in Section 4 under a more general calibration. Second, the considerable breadth of the box plots highlights the notable influence of specific UD values on the shape of the IRFs.

The middle sub-panel shows the results for the benefit replacement rate (BRR). While the impact of changes in the BRR on the output response shows clear characteristics - in particular, a marked escalation of the output contraction corresponding to higher levels of the BRR - the impact on the inflation response remains unclear. This ambiguity is illustrated by the box plot for inflation, which is predominantly centered around zero. This observation is consistent with the results shown in Figure 4, which highlights the concave nature of the inflation response. This confirms the non-uniqueness of the BRR in shaping the adjustment of inflation following a monetary contraction. Nevertheless, as the box plots for all variables are widely scattered, the level of the BRR generally tends to leave a pronounced effect on the shape of the variable IRFs.

Lastly, the sub-panel on the right-hand side illustrates the responsiveness of the IRFs to variations in the level of employment protection legislation (EPL). Notably, the box plots for both inflation and output are predominantly centered around zero, suggesting that changes in EPL do not distinctly influence the sensitivity of these variables in response to a monetary policy shock. Furthermore, the narrow width of the box plots in each case, particularly when contrasted with their counterparts in the left-hand side and middle sub-panels, underscores the minimal quantitative impact of alterations in EPL on the adjustment of inflation and output to a monetary policy shock. This observation once again corroborates the findings presented in Figure 4.

The exercise conducted to establish Figure A1 considers variations in all structural parameters of the model listed in Table A1 jointly (except for the LMIs). This undoubtedly obscures the relative importance of individual parameters in shaping the steady state and the impulse response functions. To address this, we performed an additional exercise in which only one parameter at a time, from those listed in Table A1, is varied. The purpose of this exercise is to assess the quantitative relevance of changes in specific parameters for the steady state, the impulse response functions, and thus for our overall results. In view of the results of this additional exercise, our findings indicate that the parameters most relevant from a purely quantitative perspective are (i) the output sensitivity in the Taylor rule (ϕ_y), (ii) the elasticity of matching of unemployed persons (γ), and (iii) the consumption-leisure complementarity coefficient (σ), which are among the most influential. While these parameters have a sizable quantitative effect on the impulse response functions, they do not alter the qualitative outcomes of the model across a reasonable parameter space.

A.4 Extensions

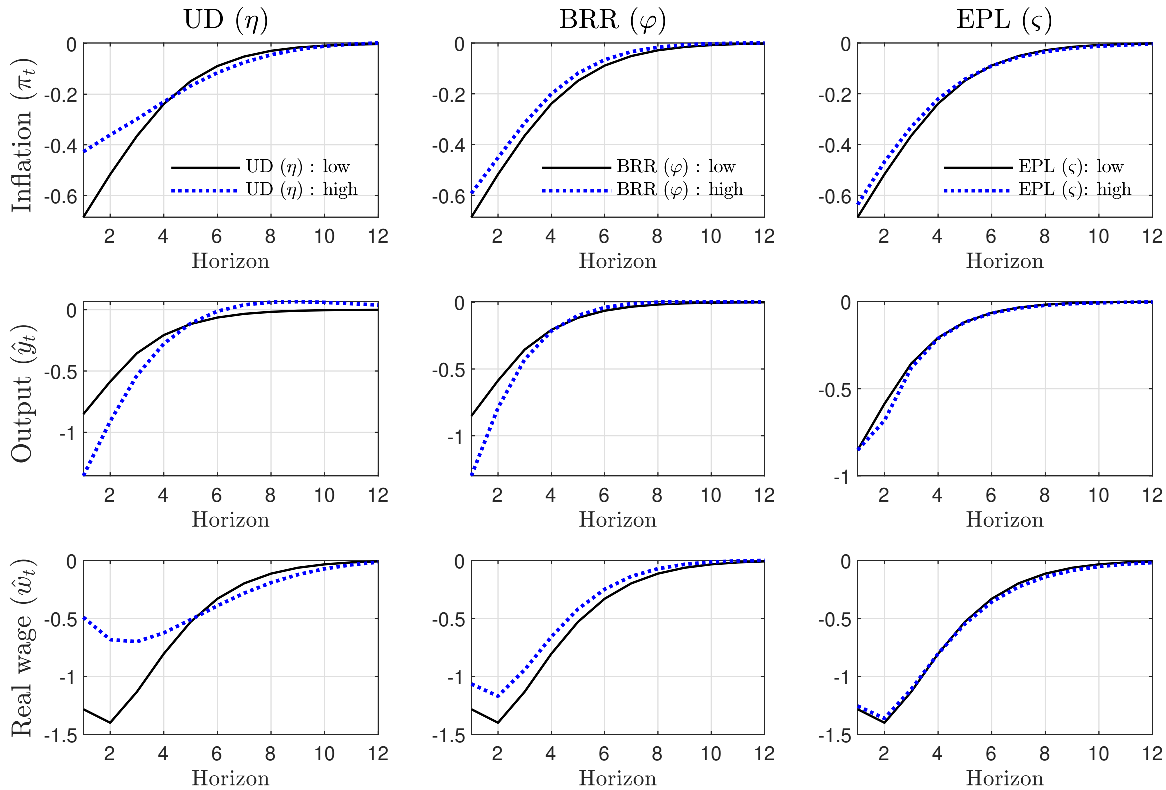
We have shown that a more general calibration of the model confirms the primary results based on the baseline calibration of the model. Another aspect that could also influence the main results relates to the specific structure of the model, including the frictions and model components considered. To address this aspect, the following sections examine various model extensions and their role in shaping our main (theoretical) results.

Real wage rigidity

Blanchard and Galí (2007) considered a particular extension of the standard New Keynesian model to address two important critiques: (i) a lack of a source of sufficient intrinsic inflation inertia and (ii) a lack of a meaningful trade-of-between stabilization of inflation and the output (even in the presence of supply shocks).

In this context, they introduced real wage rigidity, which they consider as a relevant and important real imperfection in the standard New Keynesian model. This emerges from the slow adjustment of real wages to underlying labor market conditions. The existence of such real wage rigidities has been pointed to by many authors as a feature needed to account for a number of labor market facts (for example Hall, 2005). Blanchard and Galí (2007) show that, once the model is extended to take account of this imperfection, it naturally delivers inflation inertia, a meaningful trade-off between inflation and output stabilization and fits the data better.

Figure A2: Impulse responses from the theoretical model with real wage rigidity.



Notes: Each subplot shows the impulse response functions (IRFs) following the contractionary monetary policy shock, conditional on two different levels of the LMI. The IRFs are from the baseline model extended with real wage rigidity.

We take this as a starting point for examining the sensitivity of our model's variables to a monetary policy shock under varying LMIs. 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 monetary policy 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 is 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 the real wage rate (w_t) responds 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 (4.12) to the following

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

where $\check{w}_t = (1-\eta)(mrs_t + b_t^u) + \eta\left(\mu_t mpl_t + E_t\Lambda_{t,t+1}\left[\kappa\theta_{t+1} - b_{t+1}^s(1-\bar{\varrho})F(\tilde{a}_{t+1})\right]\right)$. The parameter ϱ_w captures the extent of real wage rigidity. Equation (A.1) 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 Zanetti (2007) and Gertler, Huckfeldt and Trigari (2020) and the papers cited therein.

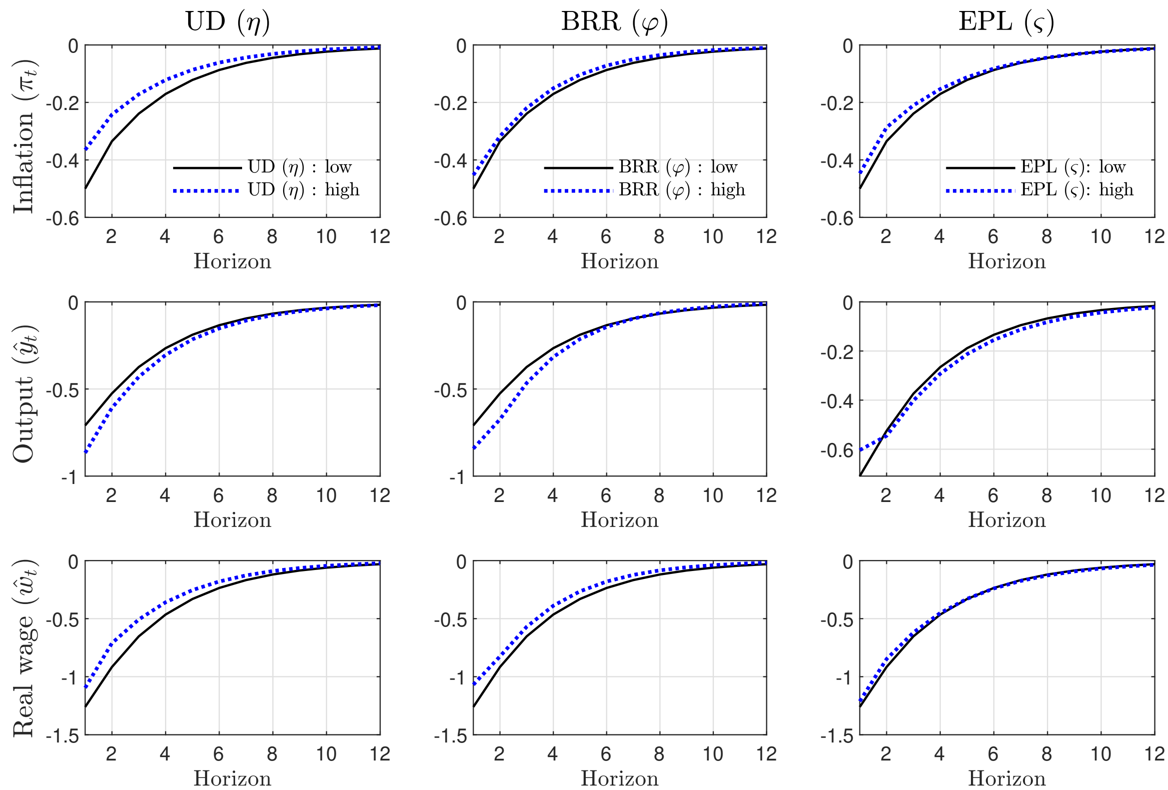
The results of this model extension are shown in Figure A2, which contrasts the impulse response functions (IRFs) to a monetary policy contraction at low and high values of the LMI. Figure 4, on the other hand, contrasts the effects of this extension with those of the baseline model across the whole range of different values for the LMIs. The simulations are based on a moderate level of real wage inertia ($\varphi_w = 0.3$). As can be seen from these figures, the introduction of rigid real wages does not change the qualitative course of the IRFs, although it does lead to quantitative adjustments. With rigid wages, the intensity of output contraction becomes more pronounced, while that of real wages (and hence inflation) is reduced due to their stickiness. Importantly for our setting, we find that the influence of LMIs in shaping the monetary policy transmission mechanism remains qualitatively unchanged. This observation confirms the robustness of our baseline results.

Limited asset market participation

Bilbiie and Straub (2013) highlighted the important role of limited asset market participation for the efficacy of monetary policy. According to these authors, in a conventional sticky-price model, standard aggregate demand logic is inverted at low enough asset market participation: interest rate increases become expansionary, and only passive monetary policy ensures equilibrium determinacy in this situation. Put differently, if asset market participation is low, aggregate demand is positively related to real interest rates.

According to Bilbiie and Straub (2013) limited asset market participation introduces a distributional dimension which shapes the inverted aggregate demand schedule. Non-Ricardian households (that is, non-asset holders) introduce an opposing effect on firms' profits: both marginal cost (wage) and sales (output and hours) fall.

Figure A3: Impulse responses from the theoretical model with limited asset market participation.



Notes: Each subplot shows the impulse response functions (IRFs) following the contractionary monetary policy shock, conditional on two different levels of the LMI. The IRFs are from the baseline model extended with limited asset market participation (i.e. non-Ricardian consumers/households).

The relative size of these reductions (and the final effect on profits) depends on the relative size of the share of non-Ricardian households (and the labor supply elasticity). If the share is high, an increase in profits would occur in response to a monetary contraction that would generate a positive income effect on Ricardian households. This could give rise to an increase in output despite an initial interest rate hike.

We follow Galí, López-Salido and Valles (2007) and add the second consumer type of households (non-Ricardian households) into the baseline model. The households outlined in the baseline model are now referred to as *Ricardian* households and their consumption is henceforth referred to as c_t^r (same for their labor supply n_t^r). *Non-Ricardian* 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 $c_t^{nr} = w_t n_t^{nr} + b_t^u(1 - n_t^{nr}) + T_t^{s,nr}$. Aggregate consumption and employment are given by a weighted average of the corresponding variables for each household 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 A3 shows the effects of the LMIs on the IRFs for output, employment, etc. in the extended model for different values of the LMIs. Figure 4 in turn compares the effects of this extension to the results of the baseline model across the whole range of different values for the LMIs. The simulations are based on a share of one-quarter of non-Ricardian households ($\lambda = 0.25$).

As can be seen from these figures, the introduction of limited asset market participation does not change the qualitative course of the IRFs, although it does lead to quantitative adjustments. With a reasonable share of non-Ricardian households, the intensity of output contraction becomes less pronounced and the same applies to real wages. Importantly for our setting, we find that the influence of LMIs in shaping the monetary policy transmission mechanism remains qualitatively unchanged. This observation confirms the robustness of our baseline results.

Habit formation

While habit formation may not change the results qualitatively, it significantly affects the shape of the IRFs and, consequently, the timing of the transmission channel for monetary policy shocks. To explore this, we extend the model by extending the utility function $u(c_t, n_t)$ to the following:

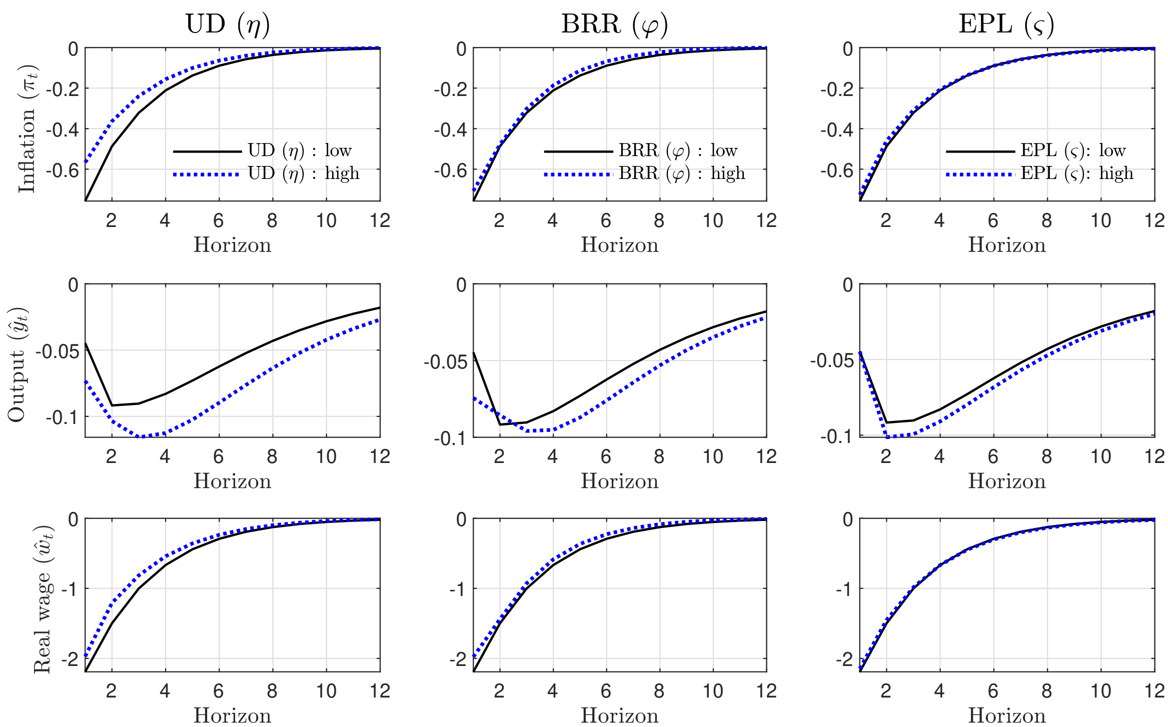
$$u(c_t, c_{t-1}, n_t) = \frac{(c_t - hc_{t-1})^{1-\sigma}(1 + (\sigma - 1)\phi n_t)^{\sigma} - 1}{1 - \sigma} \tag{A.2}$$

where the parameter h captures the degree of habit formation by households. This extension allows us to analyze how this friction leads to a hump-shaped adjustment path in response to a monetary policy shock. Our goal is to better align the IRFs of the DSGE model with those observed in the BVAR model.

Figure A4 provides the results. It shows the IRFs of inflation, output and the real wage to a contractionary monetary policy shock. The IRFs are again subjected to different levels of UD to examine the role of this structural element in shaping the transmission mechanism of monetary policy. As can be seen, the introduction of habit formation leads to a lagged response in consumption and hence output. The degree of inertia in output adjustment increases with the degree of habit formation.

The most important aspect for our purposes is the effect of UD in this respect. While the output response has changed qualitatively, the impact of the UD on the output response is still consistent with the results presented in Section 4. The contraction in output increases with the level of union density, while the opposite

Figure A4: Impulse responses from the theoretical model with habit formation.



N*Notes:* Each subplot shows the impulse response functions (IRFs) following the contractionary monetary policy shock, conditional on two different levels of the LMI. The IRFs are from the baseline model extended with habit formation.

is true for the inflation rate. This is also the case for the remaining two parameters that capture BRR and EPL. We conclude that while this extension changes the shape of the IRFs, it does not alter the main findings on how the LMIs shape the transmission channel of monetary policy.

B. Data appendix

Data come from various sources. The data for the total and the core consumer price index (HICP, All-Items, 2015=100 and Core HICP, Overall Index Excluding Energy & Unprocessed Food, 2015=100), real GDP (Gross Domestic Product, Constant Prices, SA, Chained, EUR), employment (Total Employment, Employees, Persons, SA, in thousands), unemployment (Harmonised Unemployment, according to ILO definition), industrial production (Production in Industry (NACE B-D), 2015=100), the wage bill (Wages & Salaries, Current Prices, SA, EUR), the 10-year government bond rate (Government Bond, 10 Year, Yield, benchmark) and the 3-months government bond rate are from Eurostat (and accessed via Macrobond). The short term interest rate (EONIA Rate, Historical Close, Average of Period) and the Shadow rate (European Central Bank Shadow Rate, Wu-Xia) are taken from the ECB data portal.

As we consider a monthly frequency, quarterly real GDP data are extended to a monthly frequency by using the Chow-Lin method and industrial production and retail sales as within-quarter indicator. We do the same for employment where unemployment is used a within-quarter indicator. The real wage is the (gross) wage rate per worker computed as the wage bill relative to the number of employed people, deflated with the CPI and disaggregated with the collective wage indices. All time series cover the period 1999M1 to 2023M6. All series are seasonally adjusted. The data consist of $N = 9$ countries (Austria, Belgium, Finland, France, Germany, Italy, the Netherlands, Spain and Portugal) and the euro area aggregate.

Labor market indicators (trade union density–UD, average gross unemployment benefit replacement rate–BRR, employment protection legislation–EPL) are taken from the CEP-OECD institutions database. The original data set contains annual observations until 2004. We extend the data until 2023 using information from the OECD; the ICTWSS Database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts; and by relying on the calculations as described in Abbritti and Weber (2018). The sample includes $N = 9$ countries and euro are aggregates, corresponding to the ones described previously. Data on labor market indicators are complete.

C. Further details on the empirical model

The results shown in Section 5 only involve those of the country part of the IPVAR model. For completeness, we provide the impulse response functions for the variables of the euro area in Figure C1. These IRFs are not shaped by the interaction variables, which only affect the dynamics on the country level. Hence, no distinction has to be made between the IRFs as of a low or a high value of the interaction variable.

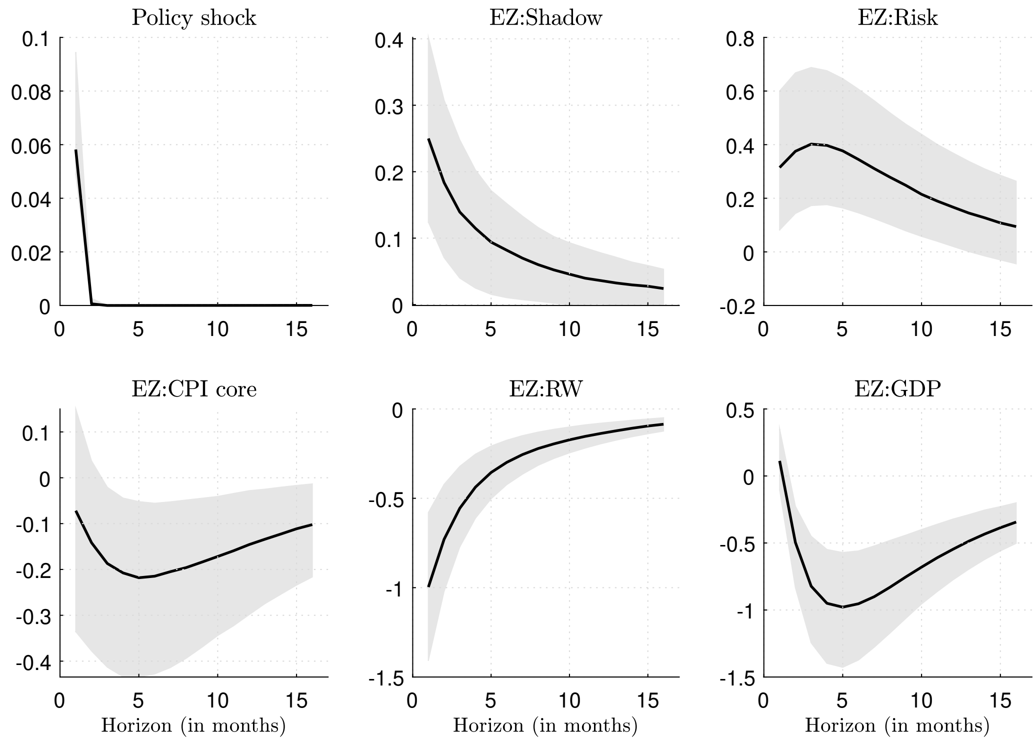
The IRFs in Figure C1 highlight that a 25 basis point increase in the monetary policy rate (shadow rate) exerts downward pressure on prices, the real wage rate and output, while upward pressure on the risk measure used (term structure of government bonds). The reaction in prices and output is significantly delayed highlighting the extent of inertia in their response to a monetary tightening, with a statistically significant response occurring after half year only.

From a quantitative point of view, the 25 basis point increase in the shadow rate causes prices (core CPI) to decline by around 0.2 percent after half a year and output (GDP) to drop by around 0.8 percent. While the price response is quantitatively in line with the estimate provided in Cantore, Ferroni and León-Ledesma (2021), the output response reported here is a bit lower in size, though more persistent. However, in contrast to Cantore, Ferroni and León-Ledesma (2021), in our case, the IRF for the real wage displays a statistically significant reaction, with a qualitative path in line with the predictions of the theoretical model put forth in Section 4.

C.1 Extensions to the empirical model

In this section we consider robustness checks and further extensions to the empirical model. Separate Ljung-Box tests on each residual time series cannot reject the null hypothesis that they follow processes which are uncorrelated over time. However, it is still possible that omitted variables matter for the results. To check whether the identified monetary policy shock is correlated with other (omitted) variables, we first follow Glocker and Towbin (2012) and compute correlations of the estimated structural disturbance with variables that a large class of general equilibrium models suggests as being jointly generated by various shocks. We compute correlations up to six leads and lags between the shock and the growth rate of local stock market indices, the stock market index of the euro area (EURO STOXX 50), the implied volatility index of the EURO

Figure C1: Impulse response functions of the euro area variables.



Notes: Impulse response functions of the euro area variables to a monetary policy shock. Black solid lines denote the median response, while the gray dashed areas are the 68 percent posterior credible intervals. The horizontal axis measures time in months. The vertical axis measures deviation from pre-shock level in percent.

STOXX 50 (VSTOXX), the Brent oil price[18], and employment. The cross-correlations indicate that none of the omitted variables correlates significantly with the structural shock.[19]

Secondly, we conducted a sub-sample check by splitting the sample in half. In this analysis, we focus exclusively on the UD, as it exhibits sufficient variation in both sub-samples across all countries. Comparing the results to those in Figure 6[20], we observe the following: for the early sub-sample, the responses are qualitatively similar but quantitatively more pronounced. As a result, the differences in the impulse response functions are larger in magnitude, with both inflation and output responses remaining statistically significantly different from zero. In the second sub-sample, we find a weaker response in real wages, while the responses for inflation and output are quantitatively consistent with those in Figure 6. However, for output, the difference

[18] We use the cyclical component of the oil price obtained after applying the Christiano-Fitzgerald filter on the logarithm of the oil price.

[19] The statistical importance of the cross-correlations has been judged by means of the upper and lower limits of an asymptotic 95 percent confidence tunnel for the null hypothesis of no cross-correlation.

[20] The results are available upon request.

in the impulse response functions between the low and high UD values is no longer statistically significant, whereas it still is for inflation.

Thirdly, we assess the changes in the results when additional variables are used in the BPVAR model which we address in detail below.

Additional endogenous variables

The theoretical model identifies the measure of labor market tightness ($\theta_t = v_t/u_t$) as a key variable for the wage-setting process, the labor market equilibrium, and thus for the overall dynamics of the model. We collected data on the number of vacancies and the number of unemployed (both obtained from the Eurostat database). Due to data availability constraints, two countries were dropped from the sample in this case. This was the case for Italy, as no vacancy data were available, and for Spain, as the time span covered by the series available to us was too short to be used in the econometric model.

We estimate the same model as presented in Section 5, but now we extende the country part to include labor market tightness ($\theta_t = v_t/u_t$) as an additional endogenous variable. The results for the extended model are shown in Figure C2. We only provide the results for the case when union density (UD) is used as interaction variable as for the other two cases (BRR and EPL) no statistically significant difference of the IRFs was found for low and high values of the interaction variables.

The figure shows that the baseline results hold in the extended setup. A tightening of monetary policy exerts downward pressure on prices, output and the real wage. Most importantly, the price response is stronger when UD is low, while the opposite is true for the output response. Regarding the new variable, labor market tightness (θ_t), we find that monetary tightening leads to a decrease in this variable, which is in line with the theoretical model (see Figure 2). However, we find that a statistically significant response only occurs when UD takes a high value. When UD takes a low value, the IRF of the tightness measure has a quantitatively negligible reaction, which is also not statistically different from zero.

We interpret these results as an additional confirmation of our baseline results. It should be noted, however, that these results are based on a limited sample, as two countries were omitted. Nevertheless, the basic elements of the baseline results hold in this extension.

Next a measure for the labor market tightness (θ_t), we also considered unemployment as an additional endogenous variable. Starting from the baseline specification as put forth in Section 5, we added unemployment in the country part of the Bayesian panel VAR model for each country and estimated and simulated the model in the same way as before. The results for this extension are provided in Figure C3. Again, we confine the analysis to the case when UD is used as interaction variable, as in the case of the other two (BRR and EPL), the difference of the IRFs across low and high values of the interaction variables is never statistically different from zero.

The results highlight that the baseline results for prices, the real wage rate and output still apply. All of them decline in response to a monetary tightening. The drop in prices is pronounced when UD takes on a low value, while the opposite applies for output. As regards the new variable, unemployment, we observe that its reaction is positive. The rise in unemployment in response to a monetary tightening is negligibly small though when UD takes on a low value, however, the size of its reaction rises with UD and, most importantly, it then also displays a statistically significant reaction. As a consequence, we observe a difference of its IRF across a low and a high value of UD. We again interpret these results as an additional confirmation of our baseline results.

Additional (truly) exogenous variables

We conducted a comprehensive examination of the baseline results with a focus on the inclusion of additional exogenous variables in the model. This analysis entails several key enhancements, including the specification of trend components (both linear and quadratic), the adoption of a lag structure comprising two lags based on AIC criteria (in contrast to the single lag chosen in the baseline results following BIC selection), and the incorporation of further exogenous variables within the Bayesian panel VAR framework. Notably, these additional variables encompass various influential factors such as the price of natural gas (specifically, the Dutch Title Transfer Facility (TTF) index), the logarithmic representation of industrial production in the United States, the Global Supply Chain Pressure Index as established by the US Federal Reserve, and lastly, the Baltic Dry Index, which serves as a gauge for the cost of maritime freight transportation.

The results for the output response are provided in Figures C4 and C5. Our analysis reveals that despite the introduction of these enhancements, the qualitative nature of the baseline results put forth in Section 5 remains largely unchanged, underscoring the robustness of our findings to various model extensions. However, upon closer examination, it becomes evident that several of the additional exogenous variables exhibit estimated coefficients that are statistically indistinguishable from zero, indicative of their overall irrelevance within our model framework. This observation holds true even when considering different lag structures for these variables, further emphasizing their lack of substantive impact on the outcomes of interest.

Using different price measures

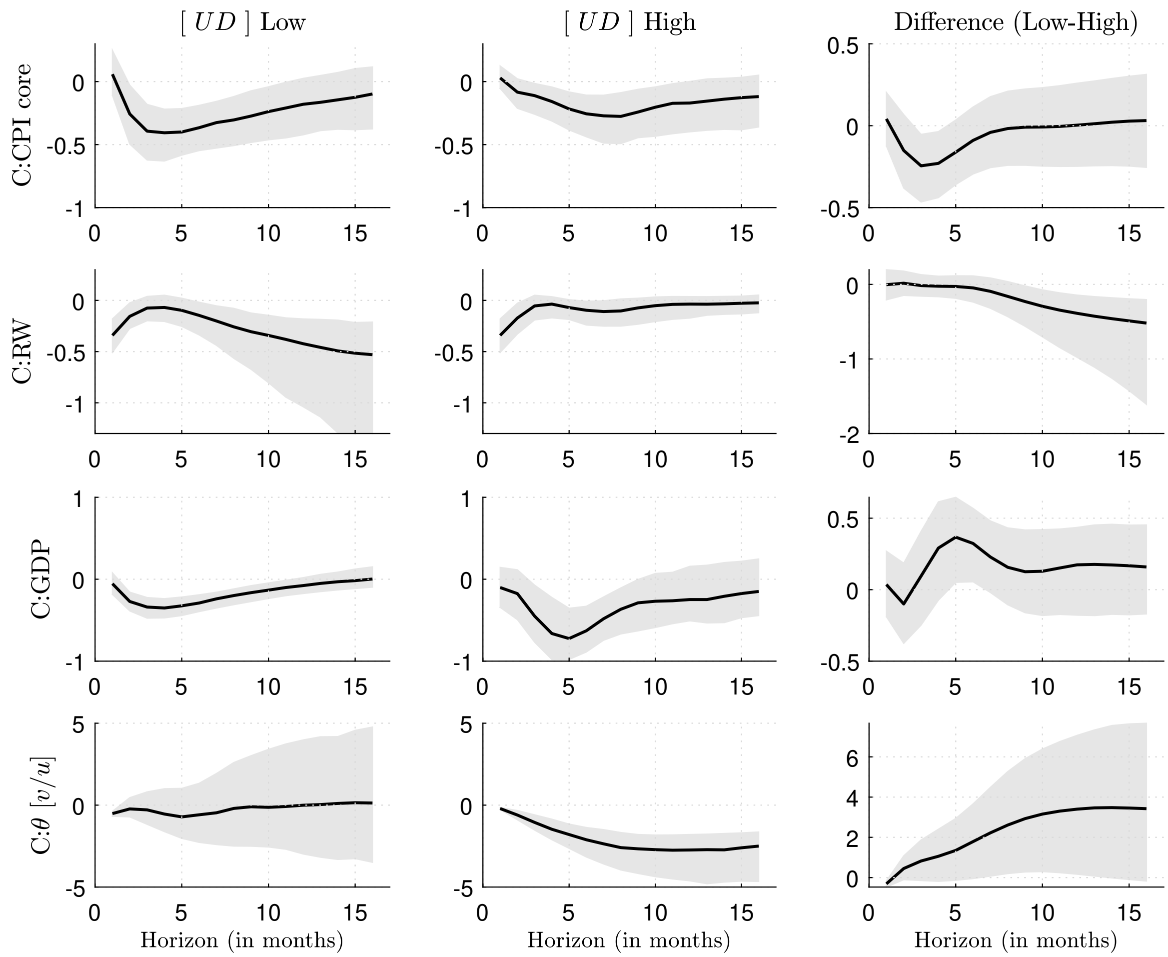
A final robustness check concerns the price measure used in the analysis. So far, we have considered the CPI core as the measure in this context, motivated by the fact that it generally contains a higher share of domestically produced goods and services than the CPI. While many other researchers use the CPI instead, we examine the sensitivity of our results to the choice of price measure used. We consider three alternative measures, namely (i) the CPI, (ii) the GDP deflator[21], and (iii) the consumption deflator.[22]

We perform the same exercise as in Section 5 and examine the sensitivity of the price response to a monetary contraction. We present the results only for UD as an interaction variable, since for the others (BRR and EPL) no significant effect of these variables in the IRFs was found. The results are shown in Figure C6. We find that in all cases prices tend to fall in response to monetary contraction. This is true regardless of the price measure used. However, in some cases the responses are not statistically significantly different from zero. This aspect already leads to the second important finding here: the size of the price reaction depends on the level of UD. In general, the price reaction is statistically significantly different from zero and quantitatively large when UD is low, but the opposite is true when UD is high. Thus, the difference between the responses at low and high UD levels is statistically different from zero. This confirms the robustness of our baseline results to extensions with different price measures.

[21] We used the CPI and the PPI (producer price index) and the Chow-Lin method of disaggregating the GDP deflator from a quarterly to a monthly frequency.

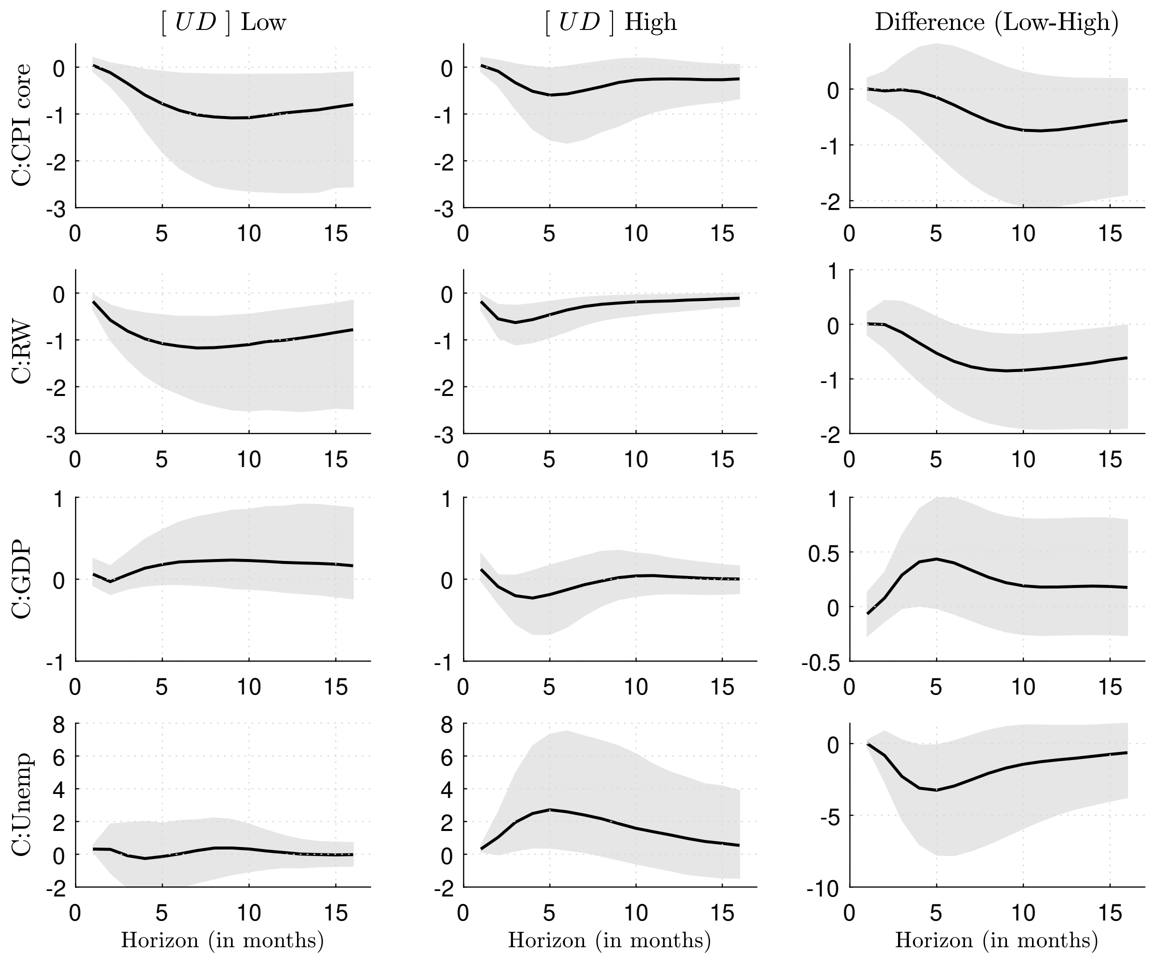
[22] We used the CPI and the Chow-Lin method of disaggregating the consumption deflator from a quarterly to a monthly frequency.

Figure C2: Impulse response functions: UD with labor market tightness ($\theta = V/U$) as additional variable.



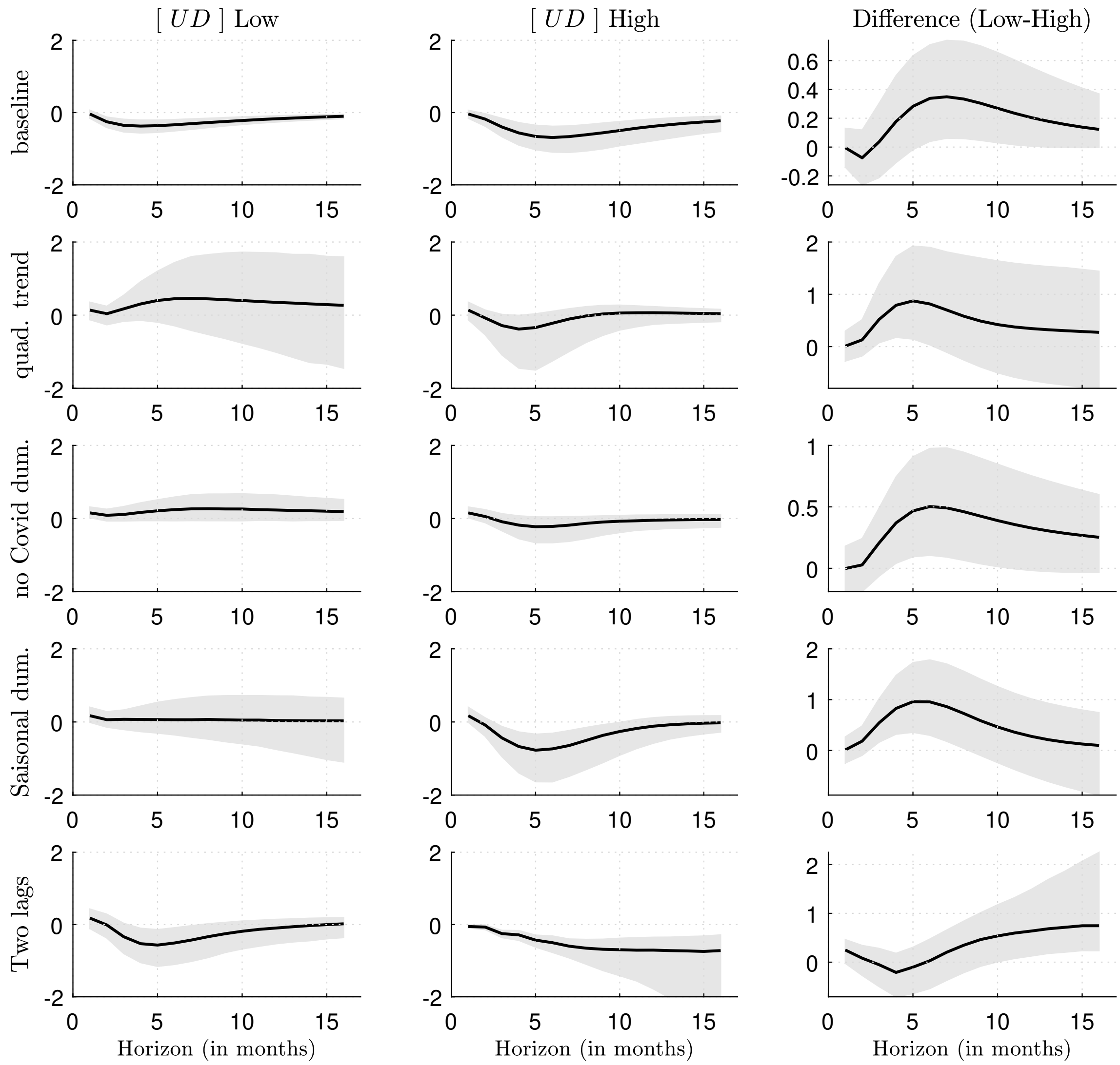
Notes: The sub-plots in the first column depict IRFs and corresponding 68 percent posterior credible intervals, when the UD has a low level (10th percentile). The sub-plots in the second column depict IRFs and 68 percent posterior credible intervals, when the UD has a high level (90th percentile). The third column shows the difference of the two IRFs and the 68 percent posterior credible interval. θ_t is given by the number of vacancies relative to the number of unemployed persons.

Figure C3: Impulse response functions: UD with unemployment as additional variable.



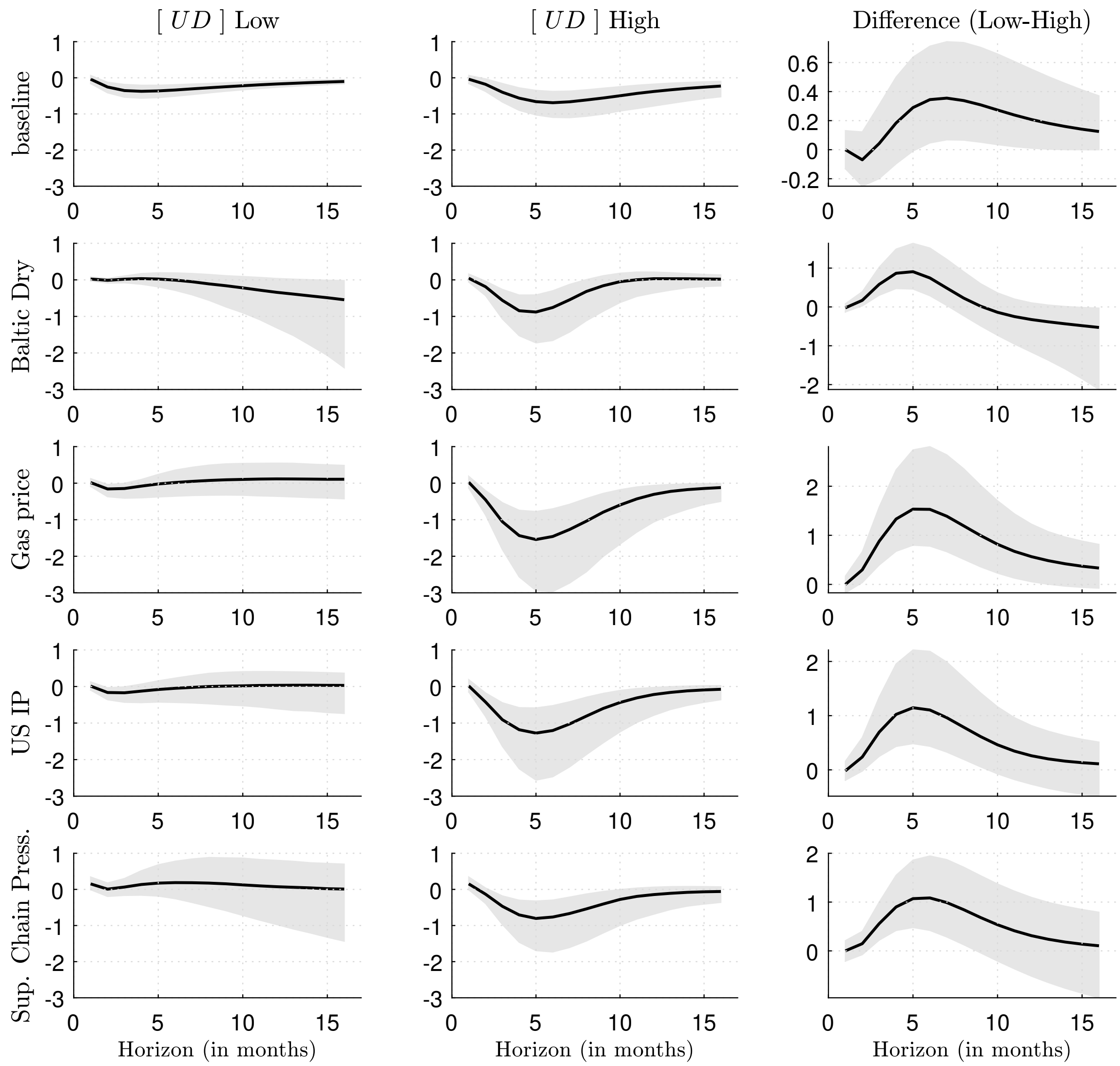
Notes: The sub-plots in the first column depict IRFs and corresponding 68 percent posterior credible intervals, when the UD has a low level (10th percentile). The sub-plots in the second column depict IRFs and 68 percent posterior credible intervals, when the UD has a high level (90th percentile). The third column shows the difference of the two IRFs and the 68 percent posterior credible interval. Unemployment (i.e. number of unemployed persons) enters in logarithmic terms.

Figure C4: Impulse response functions for output response with UD as interaction variable.



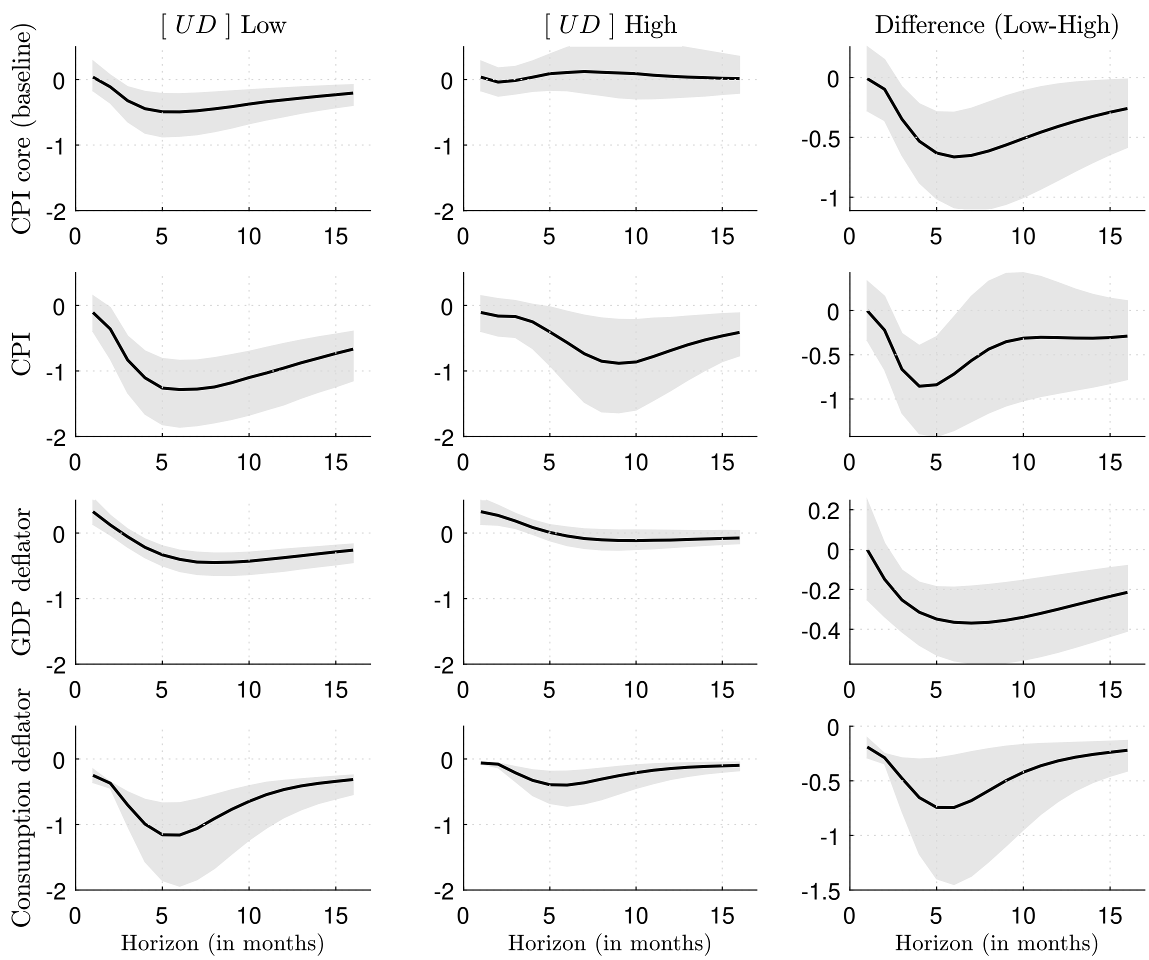
Notes: The sub-plots in the first column depict IRFs and corresponding 68 percent posterior credible intervals, when the UD has a low level (10th percentile). The sub-plots in the second column depict IRFs and 68 percent posterior credible intervals, when the UD has a high level (90th percentile). The third column shows the difference of the two IRFs and the 68 percent posterior credible interval. The first line shows the output response for the baseline specification (see Section 5), the second line shows the output response when a quadratic trend is added to the baseline model, the third line shows the output response when the dummies for the Covid-19 period are omitted, the fourth line shows the output response when the seasonal dummies are omitted, and the fifth line shows the output response when two lags are considered.

Figure C5: Impulse response functions for output response with UD as interaction variable



Notes: The sub-plots in the first column depict IRFs and corresponding 68 percent posterior credible intervals, when the UD has a low level (10th percentile). The sub-plots in the second column depict IRFs and 68 percent posterior credible intervals, when the UD has a high level (90th percentile). The third column shows the difference of the two IRFs and the 68 percent posterior credible interval. The first line shows the output response for the baseline specification (see Section 5), the second line shows the output response when the Baltic dry index is added to the baseline model, the third line shows the output response when the natural gas price (according to Dutch TTF) is added to the baseline model, the fourth line shows the output response when the US industrial production index is added to the baseline model, and the fifth line shows the output response when the global supply chain pressure index (from the US Fed) is added to the baseline model.

Figure C6: IRFs for price response with UD as an interaction variable.



Notes: The sub-plots in the first column depict IRFs and corresponding 68 percent posterior credible intervals, when the UD has a low level (10th percentile). The sub-plots in the second column depict IRFs and 68 percent posterior credible intervals, when the UD has a high level (90th percentile). The third column shows the difference of the two IRFs and the 68 percent posterior credible interval. The first line shows the price response for the baseline specification (see Section 5), the second line shows the price response when the CPI is used instead of the CPI core in the otherwise baseline model, the third line shows the price response when the GDP deflator is used instead of the CPI core, and the fourth line shows the price response when the consumption deflator is used instead of the CPI core.