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Is Unemployment on Steroids in Advanced Economies?

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Abstract

Despite conventional macroeconomic theory is based on the idea that demand shocks can only have temporary effects on unemployment, several European economies display highly persistent unemployment dynamics. The theory of hysteresis challenges this view and points out that, under certain conditions, demand disturbances can have permanent effects. In this paper, we find strong empirical evidence of unemployment hysteresis in advanced economies since the 1990s. Relying on an identification scheme instigated by an insider/outsider model, we study the effects of demand shocks allowing for cross-country heterogeneous dynamics, and exploit such heterogeneity to investigate what institutional settings have the potential to soften or amplify the effects of demand shocks. Our results indicate that strengthening labor market institutions that promote a faster adjustment of real wages, removing disincentives for firms to hire and for workers to be employed, and improving the matching between labor supply and labor demand can lessen the effects of adverse demand shocks and lead to a faster reversion of unemployment rates to pre-shock levels.

Keywords: Advanced economies, hysteresis, panel VAR, persistence, unemployment, unit root.

JEL Codes: E24, E31, E32.

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"Any reasonable reader of the data has to recognize that this financial crisis has confirmed the doctrine of hysteresis more strongly than anyone could have anticipated."

— Larry Summers, April 2, 2014. Speech at Center for Budget and Policy Priorities

1 Introduction

Conventional macroeconomic analysis is based on the idea that demand shocks only have temporary effects on unemployment. Since Friedman (1968) articulated the natural rate hypothesis, it became standard to think about the unemployment rate as a stationary variable reverting to its natural rate, without any possibility for actual unemployment to affect its natural rate. In reality, however, unemployment rates proved very persistent. In a review of unemployment dynamics in 20 developed countries, Ball (2009) finds that large increases in the non-accelerating inflation rate of unemployment (NAIRU) are associated with monetary tightenings. These findings suggest that demand shocks can lead to persistent effects in the NAIRU and, therefore, in actual unemployment rates. While most standard models incorporating nominal rigidities and matching frictions predict some persistence in the deviations of unemployment from its natural rate (see, for example, Christiano et al., 2016), in many cases output losses appear somewhat closer to permanent rather than persistent (Cerra and Saxena, 2008 and Cerra and Saxena, 2017).

In a seminal paper, Blanchard and Summers (1987) reconcile the empirical evidence of path-dependent unemployment rates with the theory by arguing that wage bargaining in a context of strong unions reflect the interest of the insiders and can result in permanent effects of demands shocks, a phenomenon they label as *hysteresis*. They themselves recognize that this argument is too strong, but other mechanisms can generate hysteresis.[1] Layard and Nickell (1987), for example, point out that a longer unemployment duration may lead to disenfranchising through skill loss, less interest of firms in hiring, and discouragement, effectively reducing the relevance of long-term unemployed workers in the wage formation process. This in turn, would exacerbate the persistence in unemployment rates and, in the words of Blanchard and Wolfers (2000), "possibly even halting" the return to lower unemployment rates. Also, Blanchard (2018) notes that if a slowdown in economic activity is associated with lower spending on research and development, total factor productivity may be permanently lower as a result of a smaller "stock" of past research and development efforts.

But how relevant is the theory of unemployment hysteresis? Research on hysteresis has been somewhat neglected since the Great Moderation, but persistent unemployment rates in many European countries and the permanent output losses in the aftermath of the GFC revamped the discussion (Ball et al., 2014 and Cœuré, 2017). Galí (2015) tests the relevance of the hysteresis hypothesis by taking "seriously" the non-stationarity of the unemployment rate in the euro area, and investigating its sources and empirical plausibility through the lens of a New-Keynesian framework. He finds that while the natural rate hypothesis cannot explain the patterns observed in the data, the hysteresis hypothesis can account for the stability of wage inflation together with nonstationary

[1] Blanchard (2018) notes that even if the insiders do not care about the outsiders, they still risk to become unemployed if economic conditions deteriorate, hence they might accept lower wages, especially if unemployment is high. Also, high levels of unemployment strengthen the bargaining position of firms at the moment of hiring, as they can draw from a larger pool of candidates.

movements in the unemployment rate during the post-1994 period.[2]

The literature also emphasized that labor market institutions play a major role in determining unemployment duration, including by softening or amplifying the effect of demand shocks. For instance, better coordination in bargaining could be reflected in a faster adjustment of real wages to shocks, thereby reducing hysteresis. On the other hand, if bargaining mostly reflects the interests of a specific category of workers (e.g., prime-age workers), it would lead to little sensitivity of wages of any other category (e.g., young workers) to their unemployment. Ljungqvist and Sargent (1995) show that unemployment duration can be longer when unemployment benefits are generous and labor taxes increase, while Mortensen and Pissarides (1999) analyze the role of unemployment benefits and employment protection policies when demand shocks hit, reaching similar conclusions. At the empirical level, there is no clear-cut evidence. Some key contributions to the literature include Blanchard and Wolfers (2000), that conduct an empirical analysis of the interactions of common shocks to unemployment and labor market institutions and find that they are crucial to explain the heterogeneity in unemployment dynamics and their persistence across Europe; and Nunziata et al. (2002) and Nickell et al. (2005), that do not find robust evidence for the role of the interactions between shocks and institutions.[3]

In this paper, we first test and argue that unemployment dynamics can be approximated by unit root processes in most of the 23 advanced economies in the sample since the 1990s. We then rely on this finding to identify aggregate demand shocks to unemployment through a modified version of the insider/outsider model of Balmaseda et al. (2000) and Amisano and Serati (2003), which, in our version, does not restrict demand disturbances to have zero long-run effects on unemployment. Finally, we estimate the impact of these shocks allowing for cross-country heterogeneous dynamics in the context of a panel structural vector autoregressive (PSVAR) model, and in a second-stage we exploit such heterogeneity to investigate what institutional settings have the potential to soften or amplify the effects of demand shocks. We find strong evidence of unemployment hysteresis, raising skepticism about the natural rate hypothesis.[4] Our results also provide suggestive yet preliminary evidence that the generosity of unemployment benefits, labor taxation, union density, and more coordinated wage setting bargaining amplify the impact of demand shocks; while incentives for specific categories of workers to look for employment, such as the diffusion of part-time employment, the length of maternity leave, higher statutory retirement age, more generous pension systems, more migrant-friendly policies, as well as higher spending on active labor market programs (ALMP) curb these effects.

The rest of the paper is structured as follows. Section 2 discusses the theoretical framework used for the shock identification. Section 3 illustrates a list of stylized facts suggesting that unemployment rates in advanced economies tend to be very persistent. Section 4 presents a two-stage empirical strategy which relies on the restrictions implied by the theoretical model and uncovers the interaction effects between shocks and institutions. Section 5 discusses the empirical results. Section 6 concludes.

[2] Galí (2015) also notes that the long-run trade-off hypothesis can account for the secular rise in unemployment during the 1970s and 1980s.

[3] A full literature review is beyond the scope of the paper.

[4] Galí (ibid.) and Blanchard (2018), among others, come to similar conclusions.

2 Theoretical Setup

The framework here adopted draws from Blanchard and Summers (1987), Balmaseda et al. (2000), and Amisano and Serati (2003). It consists of an aggregate demand equation:

$$y_t = \phi(d_t - p_t) + a\theta_t \tag{1}$$

where real GDP, y_t, depends on real aggregate demand, $(d_t - p_t)$, and productivity, θ. It also features an aggregate supply equation of the following form:

$$y_t = n_t + \theta_t \tag{2}$$

where real GDP depends on labor, n_t, and productivity.

Prices are set in line with wages, w_t, adjusted for productivity:

$$p_t = w_t - \theta_t \tag{3}$$

In turn, wages are set to obtain a certain level of expected employment, n_t^e, that is a function of the previous level of participation, l_{t-1}, the previous level of employment, n_{t-1}, as well as a wage push factor, z_t^w (as in ibid.), and a wage pull factor, s_t^w (our addition to the model):

$$w_t = arg\{w_t : n_t^e = \lambda l_{t-1} + (1-\lambda)n_{t-1} - z_t^w + s_t^w\} \tag{4}$$

where the wage push factor is a function of the institutional settings likely to increase wage sensitivity to demand shocks, such as, but not limited to, unemployment benefits and labor taxes, and the wage pull factor is a function of the labor market institutions that operate in the opposite direction, such as, but not limited to, part-time employment and ALMP:

$$z_t^w = \sum_Z^N \xi^Z Z_t \tag{5}$$

$$s_t^w = \sum_S^M \psi^S S_t \tag{6}$$

where the terms Z_t and S_t are vectors of N and M variables, respectively, proxying institutional settings.

In the model of Blanchard and Summers (1987), the parameterization of λ depends on the union density in the labor market. In settings with strong unions, wage decisions are "likely to give little weight to the interests of unemployed members and less to the interests of non-members", effectively assigning a smaller weight to the outsiders (i.e., low λ). In contrast, in non-unionized settings, incumbent workers have some bargaining power owing to low fixed costs and worse re-employment prospects, which is reflected in a higher weight to outsiders (i.e., high λ).[5] Hence, if unions are sufficiently strong, wages are set unilaterally to make expected employment equal to current employment, $n_t^e = n_t$, generating *full* hysteresis, as any shock to employment will have permanent effects with no tendency to return to the pre-shock employment level. When outsiders have some bargaining power, λ increases producing *partial* hysteresis.

Including a wage push factor in equation (4) can amplify (or reduce) the extent of hysteresis. The mechanism at work can be clarified with an example. In a country with strong unions and high unemployment, unions would negotiate for higher wages to protect insiders and the authorities may decide to increase unemployment benefits. This would strengthen the unions' efforts to protect the insiders by increasing the wage push factor z_t^w, which would result in an outcome approaching full hysteresis. As noted by Amisano and Serati (2003), introducing a wage push factor is akin to have a time-varying λ. The same mechanism operates for the wage pull factor. In a country with low union density and high unemployment, increases in public spending on ALMP would lead to a higher wage pull factor s_t^w. This, in turn, would be reflected in a higher level of expected employment and lower wages. The outcome would then be characterized by partial hysteresis. More generally, one could easily think of other possible scenarios in which the decisions of the authorities about the wage push and pull factors lead to full or partial hysteresis.

Finally, we further specify the labor supply by linking current labor force participation to expected real wages, past unemployment, a wage push factor, and also a wage pull factor:

$$l_t = \alpha(w_t - p_l^e) - bu_{t-1} + z_t^l - s_t^l + \tau_t \tag{7}$$

where b is the discouragement effect and τ_t is a stochastic disturbance.[6] Having specified real wages and labor force participation in equation (4) and equation (7) as functions of the wage push and pull factors has the advantage of allowing for interactions among the institutions of equation (5) and equation (6), in the spirit of Blanchard and Wolfers (2000).

As a last step, we specify the disturbances to supply, demand, and labor force participation as random walk processes:

$$\begin{aligned} \Delta\theta_t &= \varepsilon_t^s \\ \Delta d_t &= \varepsilon_t^d \\ \Delta\tau_t &= \varepsilon_t^l \end{aligned} \tag{8}$$

where the shocks ε_t^s, ε_t^d, and ε_t^l, are effectively productivity, aggregate demand, and labor supply shocks.[7]

[5] If unemployment is low and fixed costs to hire outsiders are also low, new firms may hire outsiders generating competition in the goods markets and lowering wages. If unemployment is high, re-employment prospects are gloomier for workers if laid off, forcing insiders to accept lower wages. In this case, however, insiders may harass the outsiders as replacing the whole labor force is not likely to be cost effective (Lindbeck and Snower, 1986).

[6] As in Amisano and Serati (2003), we model current labor force as a function of expected real wages rather than actual real wages and past unemployment rather than contemporaneous unemployment, providing a better characterization of how long-term unemployment strengthens the bargaining position of the insiders.

[7] In this framework, oil price shocks are captured as negative productivity shocks, which increase prices (thereby lowering real wages) and raise unemployment.

We solve the model for real wages, real output, and unemployment, which yields the following equations:

$$\Delta(w_t - p_t) = \varepsilon_t^s \tag{9}$$

$$\begin{aligned}\Delta y_t =& \big[\phi\varepsilon_t^d + (\phi+\alpha)\varepsilon_t^s - z_t^w + s_t^w\big] \\ &+ \lambda\frac{\big[z_{t-1}^w - s_{t-1}^w + \Delta z_{t-1}^l - \Delta s_{t-1}^l\big](1-\phi-\alpha)\varepsilon_{t-1}^s + \varepsilon_{t-1}^l - \phi\varepsilon_{t-1}^d + \alpha\varepsilon_{t-2}^s\big]}{\big[1-(1-b-\lambda)L - bL^2\big]}\end{aligned} \tag{10}$$

$$u_t = \frac{z_t^w - s_t^w + \Delta z_t^l - \Delta s_t^l + (1-\phi-\alpha)\varepsilon_t^s + \varepsilon_t^l - \phi\varepsilon_t^d + \alpha\varepsilon_{t-1}^s}{\big[1-(1-b-\lambda)L - bL^2\big]} \tag{11}$$

Equation (11) suggests that the unemployment dynamics depend positively on the discouragement effect, the impact of past employment on wages, and the wage push factor; and negatively on the wage pull factor. Compared to the framework of Balmaseda et al. (2000), for any given value of b, full hysteresis is not limited to the case in which $\lambda = 0$. In fact, the presence of the wage push and pull factors can re-produce a persistent behavior of the unemployment rate.

While this setup suggests that real wages depend solely on productivity shocks, an assumption on the role of aggregate demand and labor supply shocks has to be made by analyzing the stationarity properties of the unemployment rates. On the one hand, if the unemployment rate is $I(0)$, we can assume that aggregate demand shocks do not have any long-run effect on real output, which is consistent with the natural rate hypothesis.[8] On the other hand, if the unemployment rate is $I(1)$, aggregate demand shocks do have an impact on real output in the long run, and labor supply shocks are the one with no long-run effect on real output. It is therefore critical to assess the persistence of the unemployment rate to opt for the appropriate identification strategy.

3 Stylized Facts

3.1 Unemployment Persistence

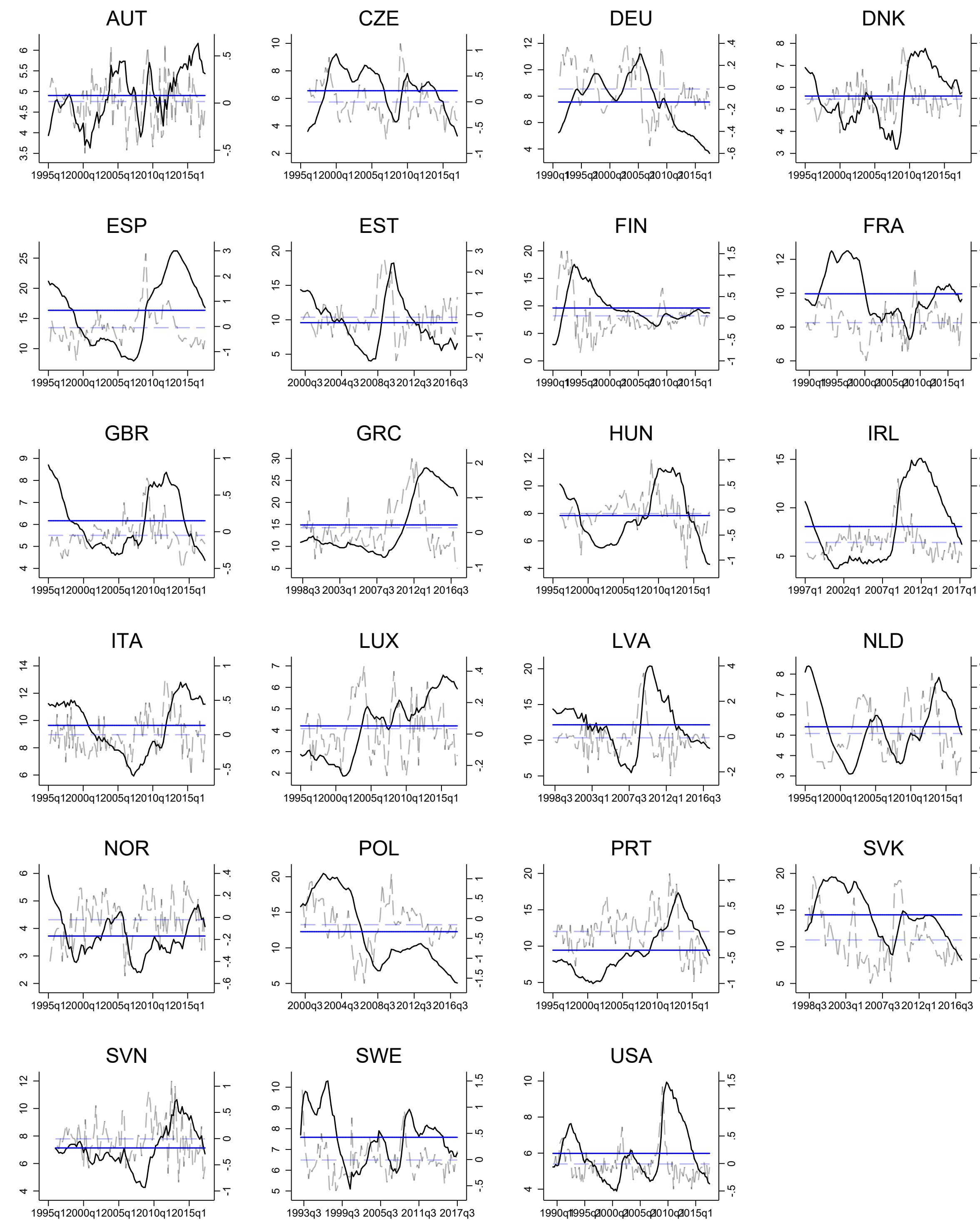
A visual inspection of unemployment rates in 23 advanced economies since 1990s reveals a great deal of persistence.[9] As shown in Figure 1, most of the economies in the sample show persistent swings in the series with limited mean reversion. Over a number of quarters ranging between 61 and 114 depending on data availability across countries, the lines denoting the unemployment rates cross the lines representing the country-specific means less than four times on average. On the other hand, changes in unemployment rates appear stationary around their mean. To further appreciate the strong persistence of unemployment rates, we plot the autocorrelation coefficients in Figure 2. The coefficient remains statistically different from zero for a minimum of one and a half years up to about two and a half years, indicating the random walk nature of the unemployment series.

[8] Galí (2015) shows that the natural rate hypothesis cannot account for the patterns of unemployment and wage inflation observed in the euro area between 1970 and 2014. Specifically, the strong persistence in unemployment cannot be reconciled with a mean reverting behavior implied by the theory.

[9] See Appendix A for the list of countries in the sample.

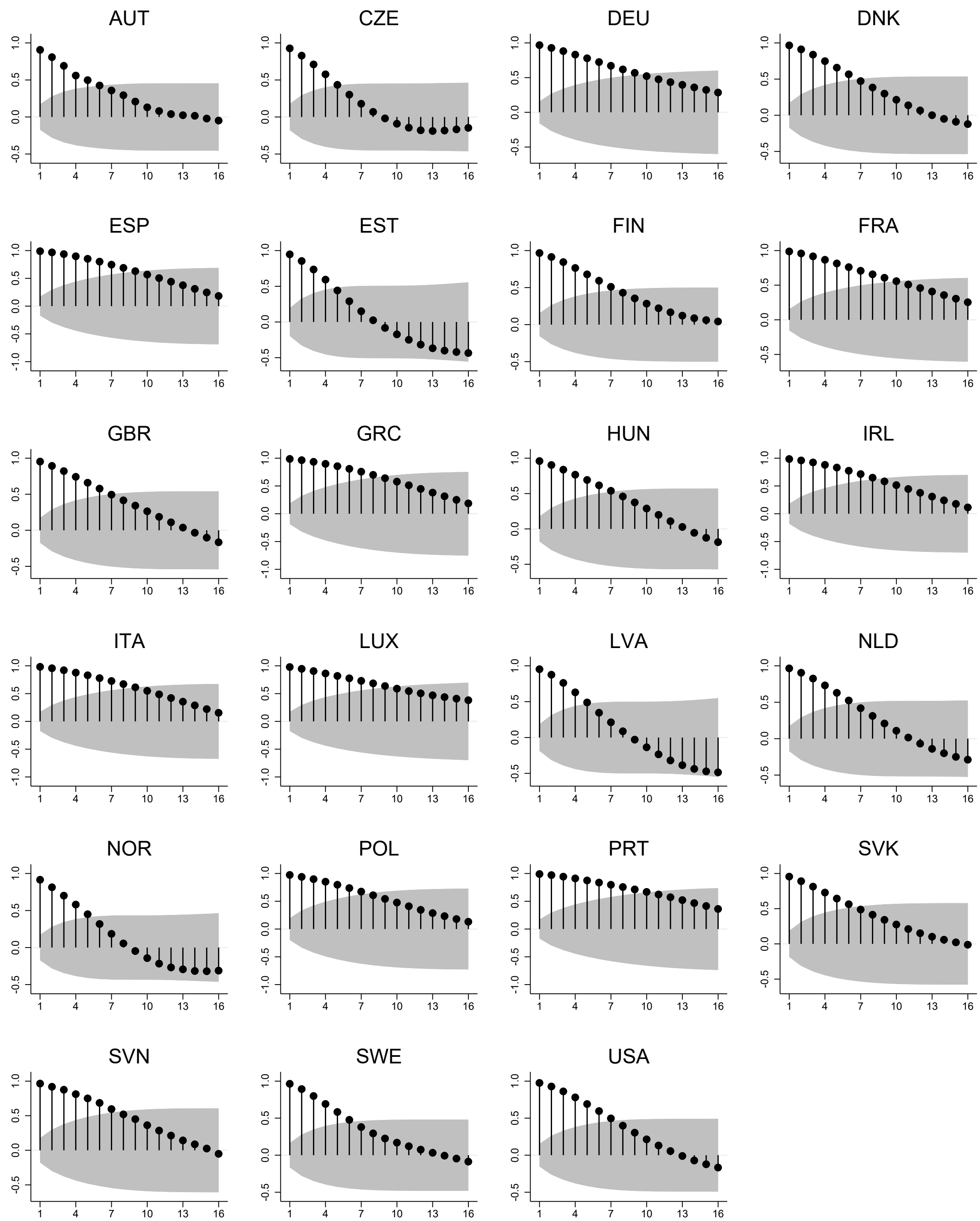
Figure 1: Unemployment Rate Levels and Changes

(Left Y-axis: unemployment rate, percent; right Y-axis: changes in unemployment rate, pp)



Source: Authors' calculations.
Notes: The (dashed) solid black lines denote the (change in the) unemployment rates, and the (dashed) solid blue lines denote the mean of the (change in the) unemployment rates. The change in the unemployment rates and its mean are measured on the right axis.

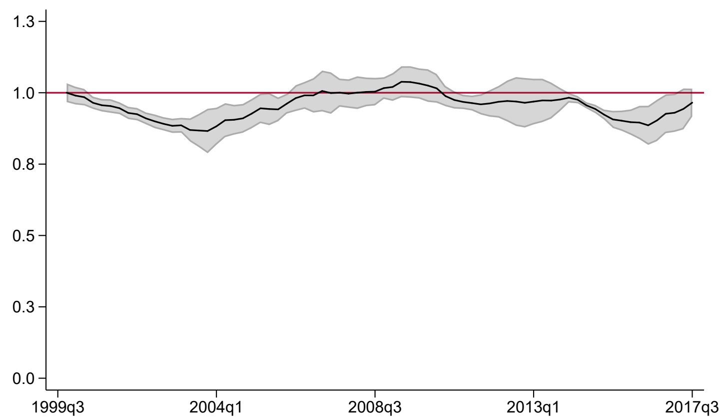
Figure 2: Autocorrelation of the Unemployment Rate
(Percent)



Source: Authors' calculations.
Notes: The black lines denote the autocorrelation coefficients and the shaded are denote the 90 percent confidence interval.

The persistence, however, might have varied over time. For instance, changes in institutional settings could have favored a more or less dynamic adjustment of unemployment to shocks. To explore this possibility, we run a rolling panel regression of unemployment on its lag (as well as on country- and time-fixed effects) and plot the coefficient over time. Figure 3 shows that the coefficient hovers around one for the length of the sample. In the few periods for which the coefficient is statistically different from one, the point estimate does not fall below 0.9, confirming that the average country displays a unit root process for the unemployment rate.

Figure 3: Coefficient on Lag Unemployment rate of an $AR(1)$ Regression
(Percent)



Source: Authors' calculations.
Notes: The label on the x-axis shows the end date of the 5-year rolling sample. The black line denotes the coefficient on the lagged unemployment rate, and the shaded area denotes the 95 percent confidence interval calculated with HAC standard errors. The regression specification includes country- and time-fixed effects.

We now move to perform formal unit root tests for unemployment. Table 1 reports the Augmented Dickey-Fuller (ADF) test statistic including the intercept as well as including the intercept and a trend component, both for the level of the unemployment rates and the changes.[10] For the great majority of the advanced economies in the sample, the null hypothesis of the presence of a unit root process cannot be rejected. When including the trend, only in the case of Finland, Netherlands, and Sweden the results suggest that the unemployment rate is stationary. In the case of Greece, Ireland, and Latvia, the null hypothesis cannot be rejected at the 10 percent significance level for the changes in the unemployment rate, indicating some level of persistence even for that. These results are not new in the literature. Balmaseda et al. (2000) and Galí (2015), for example, come to similar conclusions, but note the power of the conventional unit root test is impaired when samples are finite.

As a way to circumvent the limited power of the ADF test, we perform a second set of unit root test relying on the Johansen (1991) framework for cointegration testing. In line with Balmaseda et al. (2000), we specify a vector autoregression in levels including real wages, real output, and unemployment plus an unrestricted linear trend. If we reject the null hypothesis that there are no cointegrating vectors, and cannot reject both that the number of cointegrating vectors is one and that the cointegrating vector has the form $\big[0,0,1\big]$ for $\big[\big[(w-p),y,u\big]\big]$, then we can conclude that u is $I(0)$ while the other variables are $I(1)$ processes, with no cointegration among the variables. On the other hand, if we cannot reject that the form of the cointegrating vector is $\big[(w-p),y,u\big]$,

[10] Real wages and real output are unambiguously integrated of order one. The results of the unit root tests for these series are available upon request.

Table 1: Unit Root Tests for Unemployment Rates Based on ADF Test

Country	Obs	Intercept		Intercept and trend	
		Levels	Changes	Levels	Changes
AUT	91	-1.289	-6.782***	-2.112	-6.755***
CZE	85	-2.064	-3.606***	-2.593	-3.931**
DEU	107	-0.713	-3.685***	-2.040	-4.150***
DNK	91	-2.245	-5.953***	-3.073	-5.956***
ESP	91	-1.678	-3.141**	-1.888	-3.090*
EST	70	-2.320	-4.174***	-2.329	-4.152***
FIN	111	-3.287	-3.194**	-4.503**	-3.341*
FRA	114	-1.944**	-4.457***	-2.608	-4.440***
GBR	90	-1.774	-4.249***	-1.763	-4.214***
GRC	77	-2.239	-2.811*	-2.903	-2.745
HUN	86	-1.183	-4.653***	-1.125	-4.646***
IRL	82	-1.849	-2.536	-1.912	-2.493
ITA	91	-1.520	-3.520***	-1.564	-3.614**
LUX	90	-1.255	-4.371***	-2.897	-4.331***
LVA	77	-2.953	-2.977**	-2.971	-2.962
NLD	90	-3.454**	-3.077**	-3.899**	-3.179*
NOR	91	-2.716**	-5.668***	-2.640	-5.702***
POL	70	-0.868*	-3.468**	-2.170	-3.441*
PRT	91	-1.335	-4.104***	-1.290	-4.134***
SVK	78	-1.481	-3.467**	-3.232*	-3.682**
SVN	86	-1.946	-3.320**	-1.962	-3.288*
SWE	99	-3.070	-6.449***	-2.920	-6.393***
USA	114	-2.791**	-4.323***	-2.742	-4.353***

Source: Authors' calculations.
Notes: The lag structure is based on the Schwartz information criterion. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

we conclude that u is $I(1)$. The results reported in Table 2 confirm the presence of a unit root for most advanced economies. In the case of Denmark, Estonia, the United Kingdom, Luxembourg, and the United States, however, the $I(0)$ nature of the unemployment series cannot be rejected. Interestingly, there is no match between the results across the two unit root test approaches.

Table 2: Unit Root Tests for Unemployment Rates Based on Johansen Trace Test

Country	Obs	H_0 : $r = 0$	H_0' : $r = 1$	H_0'' : $\beta = [0, 0, 1]$	Conclusion
AUT	82	45.112	16.069	12.710	$I(1)$
CZE	79	53.719	15.807	17.647	$I(1)$
DEU	105	41.887	21.215	6.119	$I(1)$
DNK	83	41.355	15.462	4.381	$I(0)$
ESP	86	53.924	19.836	7.962	$I(1)$
EST	63	62.243	31.429	2.050	$I(0)$
FIN	102	67.584	22.276	27.197	$I(1)$
FRA	111	27.596	8.955	10.945	$I(1)$
GBR	88	23.597	11.792	4.455	$I(0)$
GRC	69	29.152	9.695	9.928	$I(1)$
HUN	81	44.357	22.92	8.939	$I(1)$
IRL	74	56.942	19.251	18.457	$I(1)$
ITA	89	40.21	20.31	15.74	$I(1)$
LUX	88	22.501	10.585	2.776	$I(0)$
LVA	68	38.843	15.216	11.245	$I(1)$
NLD	88	48.497	20.555	9.838	$I(1)$
NOR	89	31.847	13.754	6.034	$I(1)$
POL	61	59.913	18.48	37.157	$I(1)$
PRT	89	27.298	8.685	15.81	$I(1)$
SVK	73	27.221	9.165	5.363	$I(1)$
SVN	81	32.768	7.702	17.673	$I(1)$
SWE	95	24.49	6.729	9.483	$I(1)$
USA	111	33.353	13.648	4.611	$I(0)$

Source: Authors' calculations.
Notes: The specification include an unrestricted linear trend. The lag structure is based on the Schwartz information criterion. Critical values for $H_0 : r = 0$ and $H_1' : r = 1$ are from Table 2 (case 2, Trace) in Osterwald-Lenum (1992).

As a way to reconcile these results, we perform panel unit root tests. Given that the degree of persistence can change by country in line with the institutional settings, we employ tests that allow for heterogeneous parameters, i.e. Im et al. (2003) and Maddala and Wu (1999).[11] The results in Table 3 confirm that unemployment rates are not stationary, while changes in unemployment rates are. Based on this evidence, we conclude that some shocks seem to trigger permanent effects in unemployment rates.

3.2 Unemployment and Wage Inflation

After establishing the non-stationarity of unemployment rates, we look into possible reasons for that. The main mechanism generating hysteresis in the framework of Blanchard and Summers (1987) is related to the nature of the wage setting institutions, which might be conducive to re-

[11] These tests differ in that Im et al. (2003) treat the parameter of interest as varying across countries and focus on the between-country dimension, while Maddala and Wu (1999) treat all the parameters as potentially varying across countries and test by pooling significance values across members of the panels.

Table 3: Panel Unit Root Tests for Unemployment Rates

	Intercept		Intercept and trend	
	Levels	Changes	Levels	Changes
Im et al. (2003)	-1.492*	-11.993***	-0.382	-9.896***
Maddala and Wu (1999)	51.991	242.325***	49.622	188.241***

Source: Authors' calculations.
Notes: The lag structure is based on the Schwartz information criterion. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

duced sensitivity of wages to changes in unemployment. If that is the case, unemployment and wage inflation should not be as negatively correlated as one would presume in a conventional Phillips curve relationship. Simple plots of the unemployment rate and wage inflation in Figure 4 suggest a variety of experiences across advanced economies. Overall, this evidence is consistent with a negatively-sloped Phillips curve in most of the countries as shown in Figure 5, but the relationship is weak and at times even flat or positively sloped for some countries.

The panels in Figure 5 are also useful to detect whether there is an evident time pattern to the flattening or steepening of the Phillips curve in correspondence of the Global Financial Crisis. At the country level, there are no clear changes in slope over the sample period. The degree of the steepness is difficult to assess from these panels, as the variance of wage inflation changes significantly across countries. Thus, we estimate the following country-specific reduced form of the wage Phillips curve equation:

$$\pi_t^w = \alpha + \beta \bar{\pi}^p_{[t-4,t-1]} + \gamma u_t + \epsilon_t \tag{12}$$

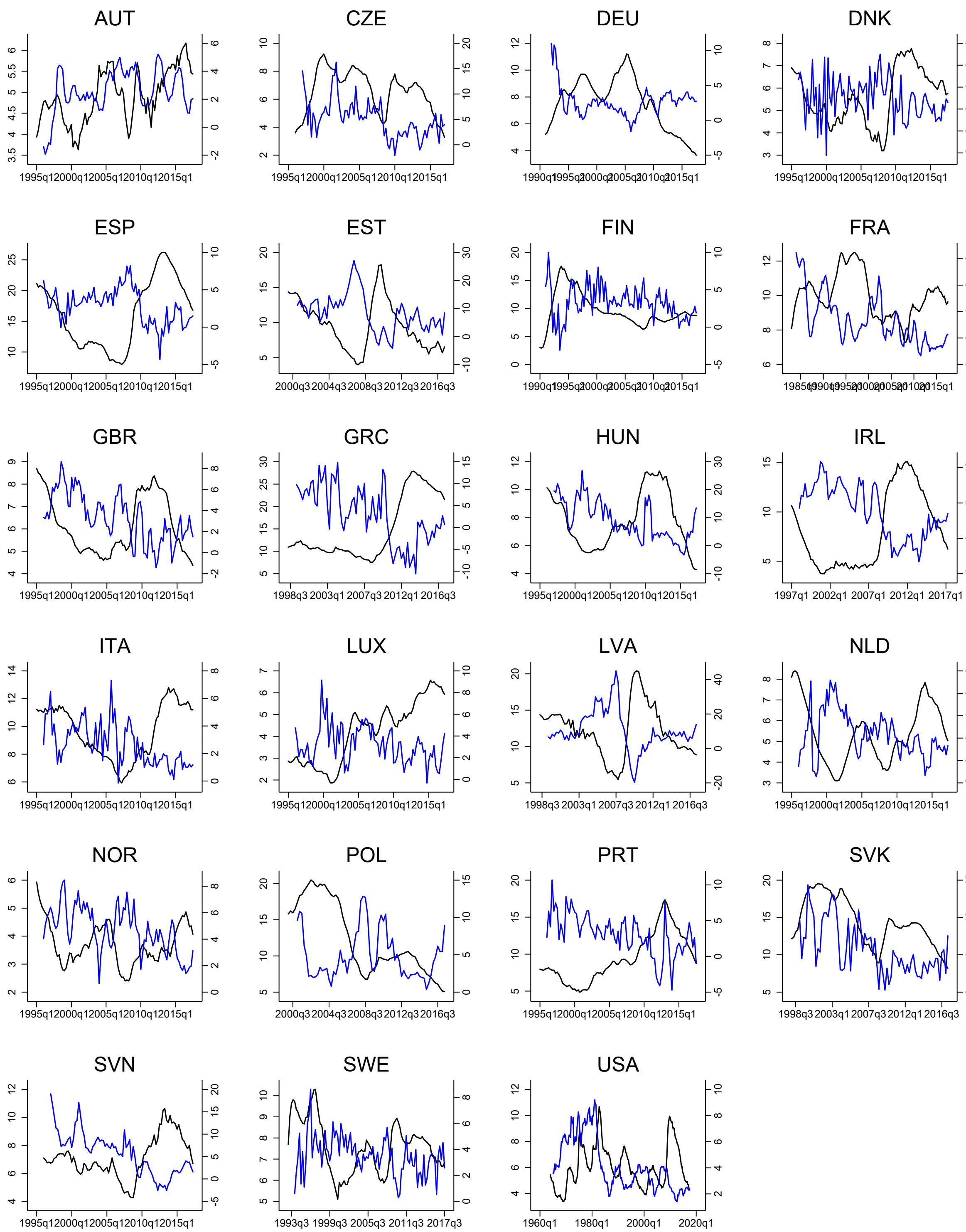
where π_t^w is (y-o-y) wage inflation at time t, $\bar{\pi}^p_{[t-4,t-1]}$ is the average (y-o-y) price inflation calculated over the last four quarters to capture possible indexation, and u_t is the unemployment rate.[12] In addition, we also estimate a panel regression adding country- and time-fixed effects. We find a significant negative relationship for all countries except Austria, Czech Republic, France, the United Kingdom, Slovak Republic, and Sweden, for which the coefficient on the unemployment rate is either negative and not significant or positive. When the relationship is negative and significant, the magnitude of the slope of the Phillips curve shows a lot of dispersion, ranging from -2.5 to -0.2. A panel estimation returns a coefficient of 0.4.

The existence of a negative relationship between the unemployment rate and wage inflation can be driven by several other factors that are not contemplated in the reduced-form specification of equation (12). What matters, however, is whether unemployment and wage inflation share a long-run relationship. We test for that by employing the Engle-Granger test for cointegration. The results reported in Table 5 indicate that over the sample period, there is little evidence that employment rates and wage inflation co-move in the long run. Only in a handful of cases (Germany, Finland, Hungary, Italy, Luxembourg, and Portugal) we cannot reject the null hypothesis of cointegration.[13]

To complete the cointegration analysis, we also perform panel cointegration tests. As the evidence

[12] Galí (2011) provides the theoretical underpinnings of the reduced form of the New-Keynesian Wage Phillips curve. In this paper, we estimate a similar specification, also estimated in Galí (2015) for the United States and the euro area. Adding the change in the unemployment rate to the regressors and substituting the unemployment rate with a measure of the unemployment gap obtained by filtering the original series do not alter the conclusions. Estimating a hybrid version of the wage Phillips curve at the country-level is complicated by data availability for

Figure 4: Unemployment Rate and Wage Inflation
(Left Y-axis: unemployment rate, percent; right Y-axis: wage inflation, percent)



Source: Authors' calculations.
Notes: The black lines denote the unemployment rates, and the blue lines denote wage inflation. Wage inflation is measured on the right axis.

Table 4: Reduced-Form Phillips Curve Estimations
(Dependent variable: wage inflation)

	(1) AUT	(2) CZE	(3) DEU	(4) DNK	(5) ESP	(6) EST	(7) FIN	(8) FRA
Past price inflation	0.629***	-0.153	0.527***	0.357**	0.343**	0.691**	-0.200	0.365***
	(0.143)	(0.111)	(0.165)	(0.176)	(0.146)	(0.270)	(0.154)	(0.132)
Unemployment rate	0.155	0.717***	-0.488***	-0.532***	-0.254***	-1.217***	-0.175**	-0.084
	(0.160)	(0.231)	(0.046)	(0.142)	(0.040)	(0.215)	(0.079)	(0.073)
Constant	0.959	0.378	5.183***	5.245***	6.054***	16.969***	5.153***	2.969***
	(0.814)	(1.472)	(0.327)	(0.922)	(0.854)	(2.273)	(0.838)	(0.815)
Observations	82	76	98	82	82	61	102	105
R-squared	0.160	0.086	0.520	0.175	0.551	0.421	0.075	0.061

	(9) GBR	(10) GRC	(11) HUN	(12) IRL	(13) ITA	(14) LUX	(15) LVA	(16) NLD
Past price inflation	-0.414	-1.264***	0.817***	0.264***	0.440***	0.148	-0.065	0.398***
	(0.441)	(0.261)	(0.149)	(0.084)	(0.112)	(0.198)	(0.252)	(0.126)
Unemployment rate	-0.390	-0.851***	-1.028***	-0.733***	-0.223***	-0.411***	-2.526***	-0.996***
	(0.348)	(0.082)	(0.345)	(0.063)	(0.079)	(0.140)	(0.274)	(0.134)
Constant	6.473***	18.330***	11.548***	8.906***	3.805***	4.731***	40.998***	6.760***
	(1.377)	(1.891)	(2.567)	(0.688)	(0.927)	(0.938)	(3.796)	(0.812)
Observations	81	68	77	73	82	81	68	81
R-squared	0.125	0.532	0.374	0.804	0.240	0.131	0.598	0.429

	(17) NOR	(18) POL	(19) PRT	(20) SVK	(21) SVN	(22) SWE	(23) USA	(24) Panel data
Past price inflation	-0.012	0.424**	0.114	0.042	0.700***	0.336**	0.207***	0.315**
	(0.204)	(0.207)	(0.257)	(0.171)	(0.100)	(0.143)	(0.059)	(0.149)
Unemployment rate	-1.764***	-0.312***	-0.506***	0.414**	-1.207***	-0.207	-0.267***	-0.448***
	(0.218)	(0.065)	(0.137)	(0.194)	(0.191)	(0.173)	(0.044)	(0.147)
Constant	11.005***	7.274***	7.481***	-0.484	11.048***	4.001***	4.245***	
	(0.874)	(0.923)	(1.702)	(2.171)	(1.564)	(1.285)	(0.280)	
Observations	82	61	82	69	77	90	105	1,865
R-squared	0.400	0.229	0.419	0.203	0.623	0.095	0.410	0.345
Countries								23

Source: Authors' calculations.
Notes: The specification in column (1) includes country- and quarter-fixed effects. HAC standard errors are reported in parentheses. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

Table 5: Cointegration Tests between Unemployment Rate and Wage Inflation

Country	Obs	Engle-Granger statistic	Number of coint. vectors from Johansen test
AUT	82	-3.987**	0
CZE	76	-2.982	0
DEU	98	-4.618***	1
DNK	82	-4.258***	0
ESP	82	-2.690	0
EST	61	-1.656	0
FIN	102	-3.144*	1
FRA	105	-1.922	0
GBR	81	-1.564	0
GRC	68	-2.660	0
HUN	77	-2.389	1
IRL	73	-2.998	0
ITA	82	-3.652**	1
LUX	81	-3.051	1
LVA	68	-1.791	0
NLD	81	-3.301*	0
NOR	82	-2.112	0
POL	61	-1.925	0
PRT	82	-4.649***	1
SVK	69	-3.007	0
SVN	77	-2.271	0
SWE	90	-2.752	0
USA	105	-1.872	0

Source: Authors' calculations.
Notes: The Engle-Granger and Johansen tests include four lags. In the case of the Johansen test, the model includes an unrestricted constant, and the number of observations is the one reported plus one. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

presented so far highlighted some cross-country differences, we employ the tests by Pedroni (1999) and Pedroni (2004) that allow for heterogeneous parameters across countries and include within- and between-country dimension tests, both parametric and non-parametric. Compared to tests such as Kao (1999), these tests have the advantage of not imposing homogeneous dynamics, thus avoiding the risk of being mispecified if heterogeneous dynamics exist. The great majority of the results of the panel cointegration tests in Table 6 suggest that unemployment rates and wage inflation do not share a long-run relationship. In sum, the findings indicating a unit root process for unemployment rates, together with the absence of a long-run relationship between unemployment rates and wage inflation, constitute a *prima facie* evidence supporting the hysteresis hypothesis and raising skepticism about the natural rate hypothesis.

Table 6: Panel Cointegration Tests between Unemployment Rate and Wage Inflation

	Intercept	Intercept and trend
Pedroni (1999) and Pedroni (2004)		
Panel v-Statistic	0.157	5.739***
Panel ρ-Statistic	-1.073	0.481
Panel PP-Statistic	-1.040	0.132
Panel ADF-Statistic	1.145	2.012
Group ρ-Statistic	-2.042**	1.529
Group PP-Statistic	-2.509***	1.447
Group ADF-Statistic	-0.941	2.640

Source: Authors' calculations.
Notes: The table reports the versions of the statistics that are not weighted by the member specific long-run conditional variances. The lag structure is based on the Schwartz information criterion. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

4 Empirical Strategy and Shock Identification

The fact that institutional rigidities causing unemployment hysteresis can be more pervasive in some countries than in others suggests that there is no reason why one should expect to observe the same degree of persistence in the response of unemployment to aggregate demand shocks. Rather, cross-country responses are likely to display different dynamics. To deal with this complexity, in the first stage of the analysis we employ the heterogeneous PSVAR model of Pedroni (2013), which relaxes the usual assumption of homogeneous dynamics among the members of the panel.[14] Beyond unveiling the heterogeneous dynamics, we are also interested in identifying which institutional features determine the severity of the shock. To do that, we take advantage of the distribution of the unconditional responses of the unemployment rates (with respect to the possible factors that determine hysteresis) to shocks in aggregate demand estimated with the heterogeneous PSVAR, and run a second-stage regression of these estimated responses on a wide set of institutional variables.[15]

We now discuss more formally the first stage of the analysis. We model the dynamic relationship

inflation expectations.

[13]In the case of Luxembourg, the cointegration results are not reliable given that we reject the presence of a unit root in Table 2.

[14]Failing to account for such heterogeneity by simply pooling data as in conventional dynamic panel methods may result in inconsistent estimation and inference of the relationships (see Pesaran and Smith, 1995).

[15]See Appendix A for the list of countries in the sample at each stage of the analysis.

among these variables with a heterogeneous PSVAR of the following form:

$$B_i \Delta z_{i,t} = A_i(L)\Delta z_{i,t-1} + v_{i,t} \tag{13}$$

where $z_{i,t}$ is a vector of (demeaned) endogenous variables: the log of real wages, $(w_{i,t} - p_{i,t}) - (\bar{w}_i - \bar{p}_i)$; the log of real output, $y_{i,t} - \bar{y}_i$; and the unemployment rate, $u_{i,t} - \bar{u}_i$. $A_i(L)$ = is a lag polynomial allowing for country-specific lag lengths according to the usual information criteria. The subscripts $i = 1, ..., N_t$ and $t = 1, ..., T_i$ on the time and cross-section dimensions take into account that the panel may be unbalanced. Compared to Amisano and Serati (2003), we do not intentionally include any institutional variable.[16] In this respect, our empirical strategy is more similar to the one of Balmaseda et al. (2000) as our aim is to estimate *unconditional* responses to the shocks and recover the correlations with the wage push and pull factors in a second stage.

To obtain the country-specific structural residuals and responses, we estimate a set of N reduced-form VARs, one for each country i:

$$\begin{aligned} B_1 \Delta z_{1,t} &= A_1(L)\Delta z_{1,t-1} + v_{1,t} \\ &\vdots \\ B_1 \Delta z_{N,t} &= A_N(L)\Delta z_{N,t-1} + v_{N,t} \end{aligned} \tag{14}$$

and we recover the structural shocks from the reduced-form residuals, $\varepsilon_{i,t} = B_i^{-1} v_{i,t}$. As noted in Blanchard and Summers (1987), when the unemployment rate is $I(1)$, equation (4) becomes $w_t = arg\{n_t^e = n_{t-1}\}$. Thus, Balmaseda et al. (2000) show that in this case $\Delta z_{i,t} = R_i(L)\varepsilon_{i,t}$ translates into the following expression with restrictions on the long-run structural coefficient matrix $R_i(1)$:

$$\begin{bmatrix} \Delta(w-p) \\ \Delta y \\ \Delta u \end{bmatrix} = \begin{bmatrix} r_{1,1}(1) & 0 & 0 \\ r_{2,1}(1) & r_{2,2}(1) & 0 \\ r_{3,1}(1) & r_{3,2}(1) & r_{3,3}(1) \end{bmatrix} \begin{bmatrix} \varepsilon^s \\ \varepsilon^d \\ \varepsilon^l \end{bmatrix} \tag{15}$$

Under this identification scheme, aggregate demand shocks are allowed to have a permanent effect on both real output and unemployment, consistent with the $I(1)$ properties discussed in Section 3.[17] Real wages, in turn, are only determined by productivity shocks in the long run, in line with a constant returns to scale assumption for the production function. Finally, labor supply shocks have no permanent effects on output, as it remains solely determined by productivity and aggregate demand shocks.

By unconditionally estimating the heterogeneous PSVAR with respect to the institutional characteristics, the impulse response function (IRF) coefficients will contain information about the role played by the same institutional factors in determining the impact of aggregate demand shocks on unemployment. Hence, we take advantage of the heterogeneous nature of the methodology in a couple of ways. First, we compute descriptive statistics for the cross-sectional distribution of the structural responses, $R_i(L)$, including the median, mean, and interquartile ranges, to present representative impulse responses for the heterogeneous dynamics in our sample. Second, we run a second stage regression to identify which institutional settings (or wage push and wage pull factors) are more likely to be associated with large responses of the unemployment rate to shocks in

[16] Amisano and Serati (2003) include labor taxes and unemployment benefits in the exogenous block of the VAR. However, this choice could be questioned as these variable may not be really exogenous to any of the variables in the endogenous bloc.

[17] This is in contrast with the identification scheme used by Balmaseda et al. (2000) and Amisano and Serati (2003), in which the partial hysteresis hypothesis supported by the evidence of stationary unemployment rates leads to impose the restriction for which aggregate demand shocks have no permanent effects on output.

aggregate demand. Formally, we estimate the following cross-section equation:

$$Y_{i,h} = \alpha + \beta \bar{X}_i + \pi_L + \epsilon_{i,L} \tag{16}$$

where $Y_{i,h}$ is the response of unemployment to aggregate demand shocks for country i at horizon h. π_h denotes the horizon-fixed effects. X is a vector of proxies for: institutional settings commonly associated with unemployment hysteresis, including union density and coordination of wage setting; policies that could change the incentives for firms to hire and for workers to be unemployed, such as labor taxation and unemployment benefits; programs that improve the matching between labor supply and labor demand, measured as public spending on ALMP (e.g., training programs and job-search assistance); and other institutional settings that could make it easier for specific categories of workers to look for employment or remain employed, such as the friendliness of migration policies, part-time employment, the length of maternity leave, the statutory retirement age, and the generosity of the pension system. While it is difficult to argue that institutional characteristics are exogenous to the size of the unemployment fluctuations and therefore to interpret the results in a casual manner, we still believe that this approach can provide preliminary insights about the interactions between shocks and institutional settings.

5 Results

5.1 Heterogeneous PSVAR

Figure 6 presents the median, the average, and the 25^{th} and 75^{th} percentile cumulative responses of the endogenous variables to productivity, aggregate demand, and labor supply shocks. Showing the distribution of the IRFs is particularly informative as it provides information about the extent to which countries with different institutional settings display evidence of hysteresis.

We start the discussion of the results from the real wage responses. Productivity shocks increase real wages both in the short and long run, however the dispersion around the mean is sizable, with the effect for the 75^{th} percentile being about 2.5 times the one for the 25^{th} percentile. In the case of an aggregate demand shock, the average response of real wages is counter-cyclical, however the magnitude of the median effect is only one third of it, and the response for the 25^{th} percentile reveals that for a quarter of the countries in the sample the effect is positive.[18] Finally, labor supply shocks tend to have a positive effect on real wages, but similar to the case of aggregate demand shocks, the effect is dispersed with a quarter of the sample showing a negative effect.

With respect to the real output responses, the short-run effect is positive for all shocks. Consistent with the restrictions imposed, the effect of a labor supply shock converges towards zero over the long run. In the case of a productivity shock and an aggregate demand shock, the effect is persistent, even though it displays a much larger cross-country variation for the productivity shock.

Finally, we look at the unemployment responses. Differently from Balmaseda et al. (2000) and Amisano and Serati (2003) that employ an identification scheme based on the partial hysteresis hypothesis, the median responses show persistence for all shocks. In the case of a productivity

[18] Balmaseda et al. (2000) finds a positive effect for the United States and a negative one for the OECD countries in the sample. They note that the popularity of Real Business Cycle theories in the United States and the extensive use of models based on sticky wages in OECD countries could be associated to this finding.

shock, however, the effect is dispersed across countries, with over one quarter of the sample displaying a zero effect. The response to an aggregate demand shock—our focus in this paper—is large, negative, and persistent, suggesting strong hysteretic effects. A one standard deviation increase in aggregate demand is associated with a fall in the unemployment rate by about 0.6 percentage points after two years for the median country. It should be noted, however, that the effect is dispersed, with half of the countries in the sample exhibiting a response ranging between about 0.4 and 1 percentage points.

We now assess the statistical significance of the median responses. Figure 7 shows the median responses together with the bootstrapped 95 percent confidence intervals. All the median responses discussed so far are statistically significant.

To complete the first stage of the analysis, we evaluate the relevance of each shock in explaining the variability of the unemployment rates. To do so, we rely on the forecast error variance decomposition (FEVD) shown in Figure 8. The left panel shows that, on average, 70 percent of the variability in unemployment rates is due to labor supply shocks in the short run, followed by aggregate demand shocks with 20 percent, and productivity shocks with 10 percent. However, as time goes by, the proportion of unemployment variation explained by aggregate demand shocks increases and reaches about 40 percent one year and a half after the shock. This increase is mostly offset by a fall in the variation associated with labor supply shocks. Once again, we analyze the cross-country distribution of these results. In the mid panel, we plot the median and the interquartile range of the proportion of unemployment rates' variability due to aggregate demand shocks. The results suggest that there is little difference between the mean and the median, and that for half of the sample aggregate demand shocks can explain between 30 and 50 percent of total variability in unemployment rates in the long run. Finally, in the right panel we show the bootstrapped 95 percent confidence interval for the median contribution of aggregate demand shocks to the unemployment variance, which confirms that it is statistically different from zero.

5.2 The Role of Labor Market Institutions

There are many institutional characteristics that can possibly determine the extent to which unemployment rates react to aggregate demand shocks, or, in other words, prevent unemployment rates to revert to pre-shock levels. We assess the role of these institutional factors by using the IRF coefficients as a dependent variable and regressing them on the variables proxying institutional settings.[19] IRF coefficients, however, vary period by period, implying that no single horizon provides an unambiguous measure of the size of the responses. Using a cross-horizon average is complicated by the fact that we have only 23 countries in the sample (i.e., 23 IRF coefficients at any given horizon), and only for a subset of these we have data for the institutional proxies, making statistical inference challenging. To address this, we pool all IRF coefficients together and run a panel regression using the ordinary least squares (OLS) estimator with autocorrelation and heteroskedasticity consistent (HAC) standard errors, as well as the weighted least squares (WLS) estimator with the inverse of the IRFs' squared standard error as weights.[20]

Table 7 reports the regression results. The estimated coefficients present signs that generally go in the expected direction. We find that higher labor taxation, generous unemployment benefits, restrictive migration policies, and strong unions seem to amplify the effects of aggregate demand shocks; while higher spending on ALMP curbs the effects of aggregate demand shocks on unem-

[19] A negative IRF coefficient describes a fall in unemployment in response to an aggregate demand shock.

[20] Lewis and Linzer (2005) show that if the White (1980) correction for heteroscedasticity is used, simple unweighted OLS to estimate the second stage generates conservative inferences of the second-stage parameters.

ployment. Also, incentives for specific categories of workers, such as the diffusion of part-time employment, the length of maternity leave, higher statutory retirement age, and more generous pension systems help limiting the impact of aggregate demand shocks on unemployment. The only counter-intuitive result is the one for coordination of wage setting, which seems to amplify the effect of the shocks, possibly because of inefficiencies associated with a coordinated approach.

Table 7: Regressions of Unemployment Responses
(Dependent variable: responses of unemployment rate to an aggregate demand shock)

	(1) OLS	(2) WLS
Tax wedge	-0.004***	-0.010***
	(0.000)	(0.002)
Unemployment replacement ratio	-0.028***	-0.006***
	(0.002)	(0.002)
Public spending on ALMP	0.007***	0.004***
	(0.002)	(0.001)
Restrictiveness of migrant integr. policies	-0.033***	-0.009***
	(0.002)	(0.003)
Union density	-0.003***	-0.005***
	(0.000)	(0.001)
Coordination of wage setting	-0.159***	-0.083***
	(0.0006)	(0.015)
Share of part-time employment	0.025***	0.017***
	(0.003)	(0.003)
Job-protected maternity leave	0.003***	0.002***
	(0.000)	(0.000)
Statutory retirement age	0.136***	0.033***
	(0.007)	(0.012)
Public spending on old-age pensions	0.180***	0.179***
	(0.010)	(0.014)
Observations	432	432
R-squared	0.754	0.715

Source: Authors' calculations.
Notes: The specifications include horizon-fixed effects. HAC standard errors are reported in parentheses. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

5.3 Robustness

We perform a battery of robustness checks. For the countries where the null hypothesis of nonstationary unemployment rate is rejected in Table 2, we run VARs using an identification scheme consistent with the natural rate hypothesis (i.e., assuming no long-run effects of aggregate demand shocks on unemployment rates and real output). As shown in the left panel of Figure 9, the average and median dynamics are very similar to the ones presented in Figure 6. In the mid panel of Figure 9, we report the IRFs of the countries for which we reject the hypothesis of unit root in Table 2. Despite the different identification strategy, Denmark, Estonia, Luxembourg, and the United States show very persistent effects of aggregate demand shocks on unemployment, lasting about three and a half to over five years and eventually converging to zero due to the identification restriction. In the case of the United Kingdom, the effect turns surprisingly positive.[21] Also,

[21]When demand shocks are allowed to have a long-run effect on unemployment as in our baseline specification, the impact turns negative also for the United Kingdom, further raising concerns about an identification strategy that restricts the long-run effect to zero.

we make the extreme-and-against-evidence assumption that all countries have stationary unemployment rates and impose restrictions consistent with no permanent effects of aggregate demands shocks on unemployment. Even under such restrictions, the results suggest that the median IRFs take four years to die out, as shown in the right panel of Figure 9.

Finally, we present the results of the second-stage regressions of specific time horizons in Table 8. Unsurprisingly, the coefficients are imprecisely estimated due to the reduced number of observations. However, the signs of the coefficients are the same as in Table 7, and the magnitudes are broadly consistent. In some cases, even with such reduced number of observations, some coefficients turn out significant. This is the case of unemployment benefits, restrictiveness of migration policies, share of part-time employment, job-protected maternity leave, statutory retirement age, and public spending on old-pensions.

Table 8: Regressions of Unemployment Responses at Specific Horizons
(Dependent variable: response of unemployment rate to an aggregate demand shock)

	(1) $h = 0$	(2) $h = 4$	(3) $h = 8$	(4) $h = 12$	(5) $h = 16$	(6) $h = 20$	(7) $h = 23$
Tax wedge	-0.001	-0.004	-0.008	-0.006	-0.003	-0.003	-0.004
	(0.007)	(0.011)	(0.013)	(0.014)	(0.015)	(0.015)	(0.015)
Unemployment replacement ratio	-0.007*	-0.017**	-0.025**	-0.032**	-0.035**	-0.037**	-0.038**
	(0.004)	(0.005)	(0.008)	(0.010)	(0.012)	(0.013)	(0.014)
Public spending on ALMP	0.003	-0.001	0.006	0.006	0.015	0.017	0.015
	(0.004)	(0.006)	(0.007)	(0.009)	(0.011)	(0.012)	(0.012)
Restrictiveness of migrant integr. policies	-0.005	-0.025**	-0.038*	-0.039	-0.038	-0.039	-0.040
	(0.006)	(0.008)	(0.018)	(0.022)	(0.025)	(0.028)	(0.029)
Union density	-0.002	-0.003	-0.005	-0.004	-0.002	-0.001	-0.001
	(0.001)	(0.003)	(0.004)	(0.005)	(0.006)	(0.006)	(0.006)
Coordination of wage setting	-0.064	-0.155**	-0.157	-0.175	-0.182	-0.172	-0.162
	(0.046)	(0.057)	(0.095)	(0.119)	(0.144)	(0.153)	(0.155)
Share of part-time employment	0.003	0.025*	0.047**	0.034*	0.019	0.017	0.020
	(0.005)	(0.011)	(0.014)	(0.016)	(0.017)	(0.015)	(0.014)
Job-protected maternity leave	0.002**	0.003**	0.004**	0.004*	0.003	0.003*	0.004*
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Statutory retirement age	0.044*	0.109**	0.150**	0.166**	0.155*	0.146*	0.142*
	(0.021)	(0.045)	(0.060)	(0.067)	(0.070)	(0.070)	(0.070)
Public spending on old-age pensions	0.061	0.162*	0.241**	0.222*	0.184*	0.168	0.167
	(0.032)	(0.069)	(0.091)	(0.095)	(0.096)	(0.094)	(0.093)
Observations	18	18	18	18	18	18	18 2
R-squared	0.700	0.826	0.843	0.802	0.746	0.739	0.748

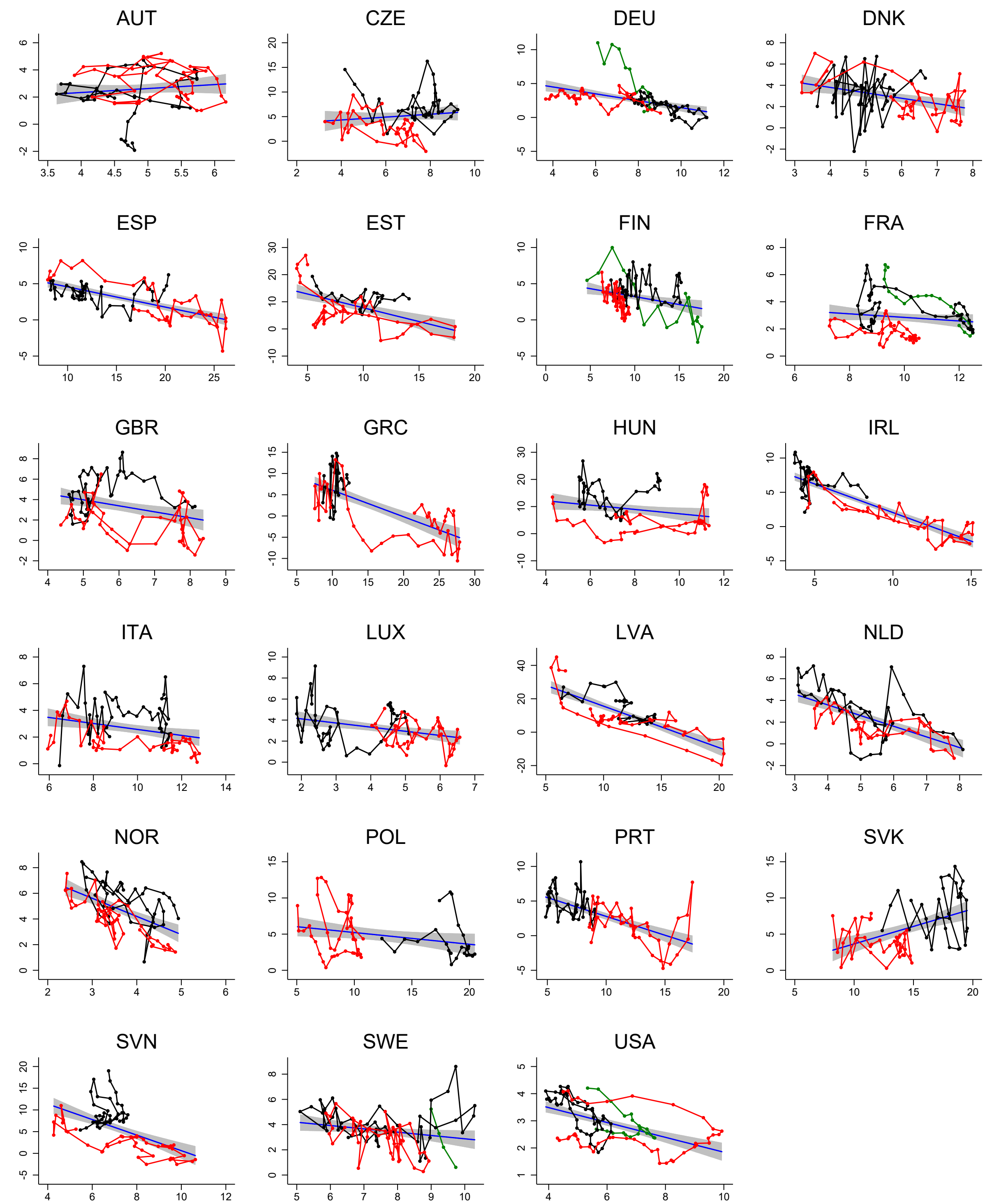
Source: Authors' calculations.
Notes: HAC standard errors are reported in parentheses. ***, **, and * indicate statistical significance at 1, 5, and 10 percent, respectively.

6 Conclusions

How persistent are unemployment rates? Can policymakers act to make demand shocks less severe and hence favor a faster reversion to pre-shock levels of unemployment? In this paper, we address these questions in the context of an insider/outsider model. Specifically, we adapt a modified version of the model of Blanchard and Summers (1987) and similar to the one employed by Balmaseda et al. (2000) and Amisano and Serati (2003) to obtain identification restrictions consistent with persistent unemployment rates. We then estimate the impact of demand shocks on unemployment rates and use the cross-country heterogeneity to explore what institutions help soften or amplify the effect of demand disturbances.

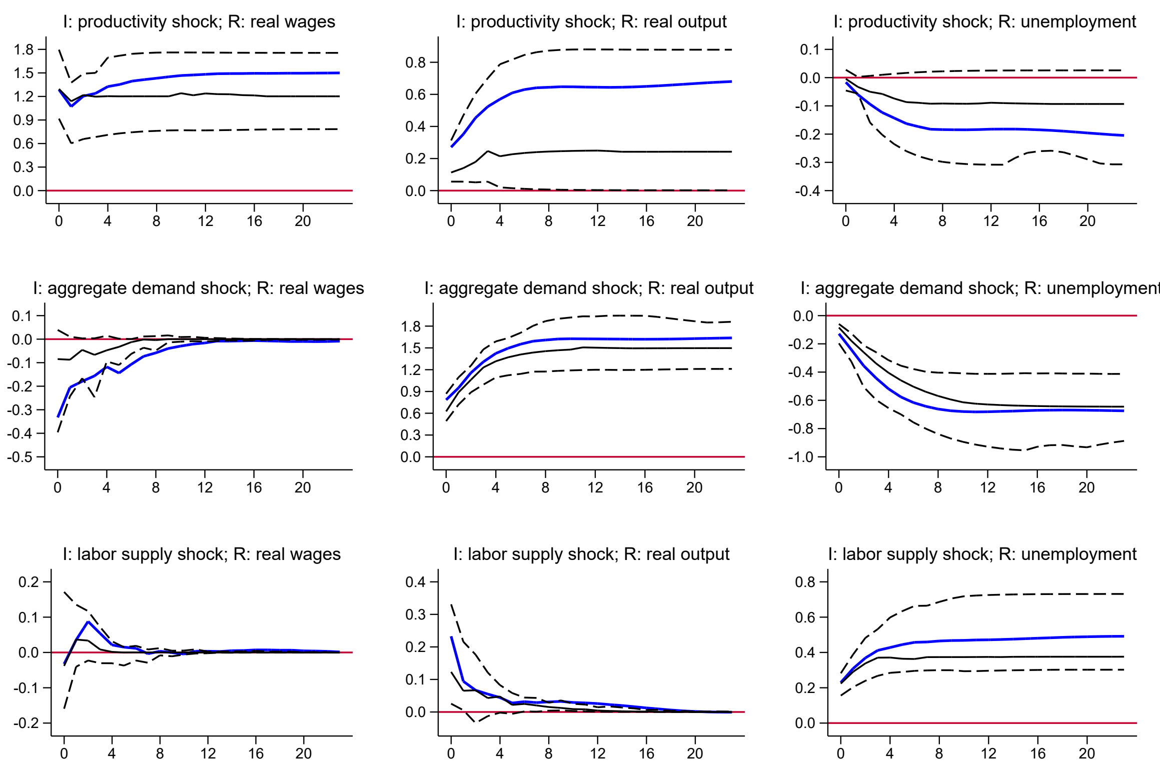
We find strong evidence of unemployment hysteresis, with demand shocks showing persistent effects on unemployment. The analysis of the cross-country unemployment responses provides suggestive yet preliminary evidence that disincentives for firms to hire and for workers to be employed (such as labor taxation and unemployment benefits), and impediments to a quick adjustment of real wages (proxied by union density) amplify the effects of demand shocks. At the same time, programs that improve the matching between labor supply and labor demand (proxied by public spending on ALMP) and incentives for specific categories of workers to look for employment or remain employed (such as the diffusion of part-time employment, the length of maternity leave, higher statutory retirement age, more generous pension systems, and more migrant-friendly policies) curb the effects of demand shocks on unemployment. In sum, our results suggest that strengthening labor market institutions can lead to less severe effects of demand shocks on unemployment and therefore a faster reversion to pre-crisis unemployment levels.

Figure 5: Phillips Curve
(Y-axis: wage inflation, percent; X-axis: unemployment rate, percent)



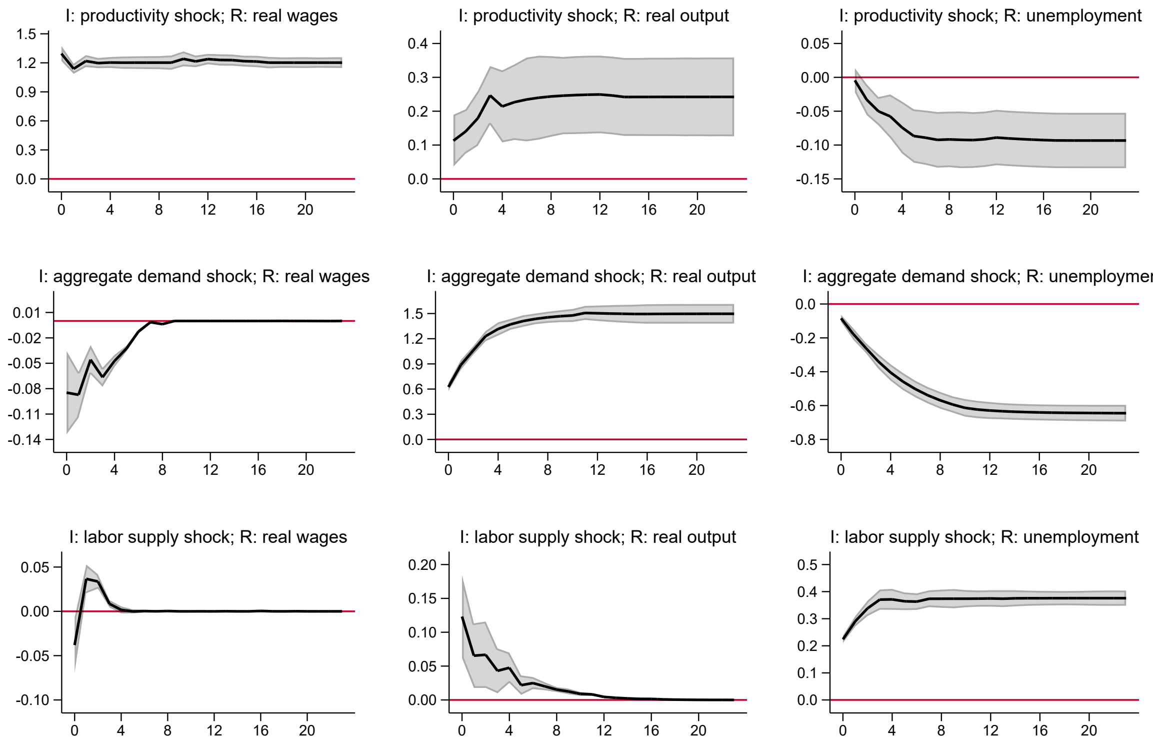
Source: Authors' calculations.
Notes: The green (black) [red] dots and lines denote the combinations of unemployment rates and wage inflation before 1995 (between 1995 and 2006) [since 2007]. The blue lines denotes the linear predictions and the shaded areas denote the 95 percent confidence intervals.

Figure 6: IRF Distribution from Heterogeneous PSVAR
(Percent)



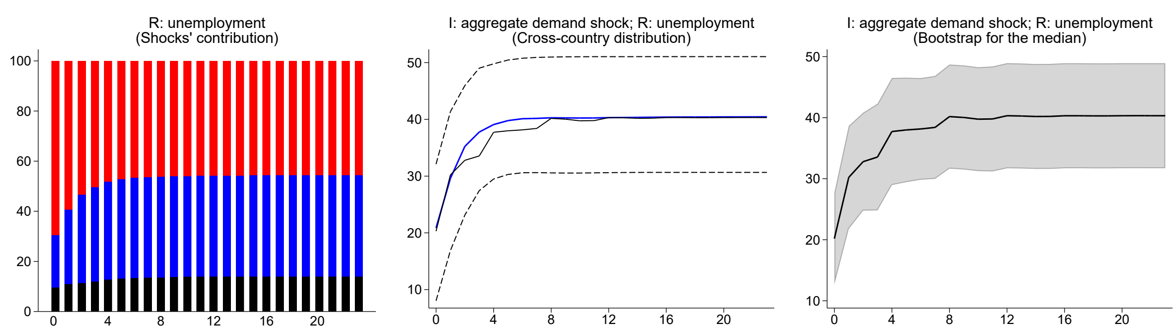
Source: Authors' calculations.
Notes: The black (blue) solid lines represent the median (average) effects, the dashed black lines represent the 25th and 75th percentiles.

Figure 7: Bootstrap of Heterogeneous PSVAR
(Percent)



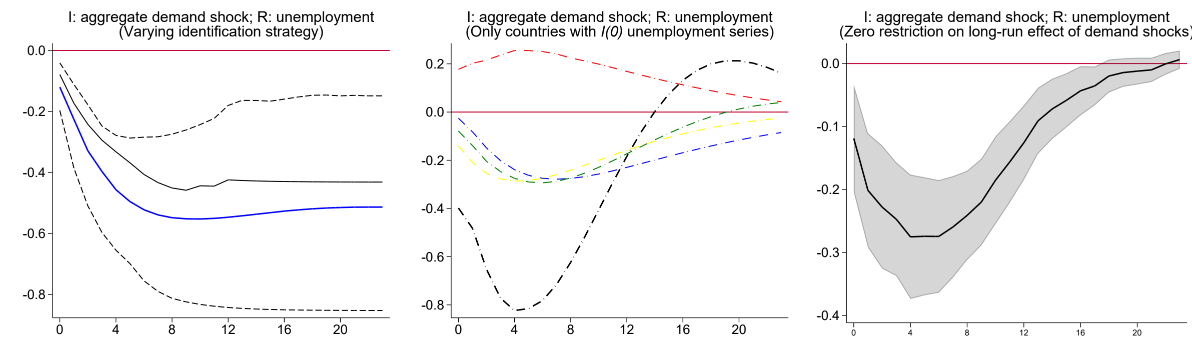
Source: Authors' calculations.
Notes: The black lines represent the median effect, the shaded areas represent the 95 percent confidence interval calculated from a resampling simulation with 100 repetitions.

Figure 8: FEVD
(Percent)



Source: Authors' calculations.
Notes: In the left panel, the black (blue) [red] bars denote the contributions of productivity (aggregate demand) [labor supply] shocks for the average unemployment response. In the mid panel, the black (blue) solid lines represent the median (average) contribution of aggregate demand shocks to the variance in unemployment responses, and the dashed black lines represent the 25^{th} and 75^{th} percentiles. In the right panel, the black lines represent the median contribution of aggregate demand shocks to the variance in unemployment responses, and the shaded areas represent the 95 percent confidence interval calculated from a resampling simulation with 100 repetitions.

Figure 9: IRFs with Alternative Identification Strategies
(Percent)



Source: Authors' calculations.
Notes: In the left panel, the black (blue) solid lines represent the median (average) effects, and the dashed black lines represent the 25th and 75th percentiles. In the mid panel, the black (blue) [red] /green/ \yellow\ dot-dashed line denotes Denmark (Estonia) [United Kingdom] /Luxembourg/ \United States\. In the right panel, the black line represents the median effect and the shaded areas represent the 95 percent confidence interval calculated from a resampling simulation with 100 repetitions.

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Appendix A. Sample

The sample for the first stage of the analysis (and the stylized facts) is composed by the following countries: Austria, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Luxembourg, Netherlands, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, United Kingdom, and the United States.

The sample for the second stage of the analysis excludes Estonia, Hungary, Latvia, Poland, and Slovenia due to missing data for some explanatory variables.

Appendix B. Data Sources

In the following, we report how we construct the variables entering the second stage of the empirical analysis and the original data sources:

- Unemployment rates are defined as the seasonally-adjusted share of unemployed to the labor force. Source: OECD, Employment database; and FRED, Economic Data.
- Real wages are calculated by deflating the seasonally-adjusted hourly nominal wages in nominal currency by the GDP deflator. Source: OECD, Employment database and Benefits and Wages database; and FRED, Economic Data.
- Real GDP is measured in constant local currency. Source: OECD, National Accounts Statistics; and Fred, Economic Data.
- The labor tax wedge is defined as the ratio between the average tax paid by a single-earner family (one parent at 100 percent of average earnings with two children) and the corresponding total labor cost for the employer. The labor tax wedge is available from the OECD for 2000 to 2016, and was extended back to 1980 using Bassanini and Duval (2006) and IMF (2016). The latter series is available only in uneven years; the value of the labor tax wedge in even years is obtained by linear interpolation. Sources: OECD, Tax database; Bassanini and Duval (2006); and IMF (2016).
- The generosity of the unemployment benefits system is measured as the gross replacement rate, which, in turn, is computed as the gross unemployment benefit levels as a percentage of previous gross earnings. The summary measure with the best coverage is the average of the gross unemployment benefit replacement rates for two earnings levels, three family situations, and three durations of unemployment. Such measures are available in uneven years, and are interpolated to obtain their values for even years. The reported values are for the average worker from 2001 to 2011, and average production worker from 1961 to 2005. The two series are spliced. Source: OECD, Benefits and Wages database.
- Public expenditure on ALMP is calculated as ALMP spending per unemployed person in percent of GDP per capita, following Gal and Theising (2015). Source: OECD, Employment database.
- Restrictiveness of migration policy is an index with information about all changes to the existing legal framework relevant for migration (see also De Resende, 2014). We focus on major changes in policies guiding the post-entry rights or other aspects of migrants' integration. Source: International Migration Institute, DEMIG POLICY database.
- Union density is measured as net union membership as a proportion of wage earners in employment. Source: OECD, Employment database.
- Coordination of wage setting is an index of the centralization of bargaining. The index runs from 1 to 5 with values defined as (1) Fragmented wage bargaining, confined largely to individual firms or plants, (2) mixed industry and firm-level bargaining, weak government coordination through minimum wage setting or wage indexation, (3) negotiation guidelines based on centralized bargaining, (4) wage norms based on centralized bargaining by peak association with or without government involvement, and (5) maximum or minimum wage rates/increases based on centralized bargaining. Source: Amsterdam Institute for Advanced Labour Studies, Database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention, and Social Pacts.

- Part-time employment is measured as the proportion of employees with a part-time contract to total employees. Source: OECD, Employment database.
- Job-protected maternity leave is defined as the total number of weeks of job-protected maternity, parental, and extended leave available to mothers, regardless of income support. Source: OECD, Family database.
- The statutory retirement age is defined as the population-weighted average at which workers can retire. Source: Social Security Programs throughout the World.
- To capture the generosity of pension schemes, we rely on the measure with the best country and time coverage, which is public spending on old-age pensions as a percent of GDP. Source: OECD, Social protection database.