Begging thy coworker – Labor market dualization and the wage growth slowdown in Europe

Lukas Lehner[1], Paul Ramskogler[2] and Aleksandra Riedl[3]

Abstract

As temporary employment has become a pervasive feature of modern labor markets, reasons for wage growth have become less well understood. To determine whether these two phenomena are related, we investigate whether the dualized structure of labor markets affects macroeconomic developments. Specifically, we incorporate involuntary temporary workers into the standard wage Phillips curve to examine wage growth in 30 European countries for the period 2004-2017. To aggregate wages by job contract type and control for changes in the workforce composition, we rely on individual-level data. This approach allows us to show, for the first time, that there is a strong negative effect of the incidence of involuntary temporary workers on aggregate wage growth. This effect, which we name the competition effect, is particularly pronounced in countries where wage bargaining institutions are weak. Our findings shed further light on the reasons for the secular slowdown of wage growth after the global financial crisis.

Keywords: wages, segmentation, unemployment, labor market institutions, Europe

JEL classification: J31 Wage Level and Structure; Wage Differentials, J42 Monopsony; Segmented Labor Markets, J82 Labor Force Composition

[1]University of Oxford, E-Mail: lukas.lehner@spi.ox.ac.uk.

[2]Oesterreichische Nationalbank. paul.ramskogler@oenb.at. The views expressed in this paper are exclusively those of the authors and do not necessarily reflect those of the Oesterreichische Nationalbank (OeNB) or the Eurosystem.

[3]Oesterreichische Nationalbank.aleksandra.riedl@oenb.at.

1 Introduction

Temporary employment has become a pervasive feature in European labor markets. The reason is simple: Temporary employment is more flexible in terms of labor cost adjustments than permanent employment. Further, it is well established that temporary employment comes with a wage penalty (Kahn 2016; Pavlopoulos 2013). For many employers, temporary employment is, thus, a cheaper source of labor. Yet, the flexibility of temporary contracts often disproportionately benefits employers instead of workers (Hyman 2018), resulting in substantial levels of *involuntary* temporary workers. Involuntary temporary employment has behavioral implications central to our analysis. Notably, *involuntary* temporary employment, by definition, forms part of the labor supply for permanent employment and thus may foster competition between different segments of workers.

We propose a mechanism through which the presence of involuntary temporary employees dampens the bargaining power and wage growth of permanent workers. This "competition effect" is not captured by the unemployment rate. To empirically identify a potential competition effect, we draw on a sample of 30 European countries in the period 2004-2017. Given the large country heterogeneity with respect to the incidence of involuntary temporary employment in Europe, the effect is expected to be mainly driven by differences *across countries*. To what extent temporary employment can influence the bargaining power of workers likely depends on a country's institutional framework. Therefore, we expect the institutional framework to reinforce or dampen the impact of temporary employment on wage growth.

While cross-country variation of temporary employment might predominate, the within-country variation linked to business cycle dynamics should not be neglected. The incidence of temporary workers increases in the early stages of a recovery (i.e., when the unemployment rate starts to decline) and falls swiftly in the downturn. Hence, temporary employees can create an additional source of slack in labor markets, which necessitates examining potential competition between temporary and permanent workers that occur *over the time dimension*. At the same time, fluctuations in temporary employment also affect aggregate wages by changing the share of workers who incur a wage penalty due to their temporary contract.

For instance, temporary workers are typically the first to be laid off during a recession, which mechanically increases average wages through a pure *composition effect*. To adjust for a changing composition of employment, we use worker-level data to construct an adjusted wage growth variable that allows us to focus on competition effects only.

Building on the insights of a relatively recent strand of literature (Bellani and Bosio 2019; Damiani et al. 2018), we thus explore how labor market dualities have affected wage growth in Europe since 2004. We do so by investigating the impact of involuntary temporary employment on wage growth in a wage Phillips curve model and by exploring the role of institutions. We then examine whether involuntary temporary employment helps to understand the Phillips curve flattening in the recovery period after the 2008 Global Financial Crisis (GFC).

2 Labor market dualization and the competition effect

Involuntary temporary employment is pervasive in European economies (ILO 2016) with an increasing share of workers experiencing temporary employment (Latner 2022). Driving this trend of labor market dualization are structural changes, such as the service sector's growth (Marx 2011) and deregulation policies (Polavieja 2006; Emmenegger et al. 2012; Thelen 2014; Biegert 2014). Shares vary across countries (Figure 1), averaging 5.5% in 2017 compared to a 7.4% average unemployment rate. Does its prevalence impact wage-setting?

2.1 The relationship between temporary and permanent workers

We introduce the *competition effect* to understand the relationship between temporary and permanent workers and investigate its macroeconomic consequences. Drawing on industrial relations scholarship, we hypothesize that the presence of involuntary temporary workers restrains wage growth due to elevated job insecurity for permanent workers, which weakens their bargaining position. Indeed, temporary employment contributes to a rise in perceived job insecurity (Kuroki 2012, p. 564), which has been suggested to explain wage restraint

(Katz et al. 1999). Our hypothesis is supported by Damiani et al. (2018), who show that reductions in employment protection for temporary workers can reduce overall wage shares. Bellani and Bosio (2019) find that, at the occupational level, wages of permanent employees are negatively affected by the incidence of overall temporary employees (i.e., voluntary and involuntary). Ramskogler (2021) indicates that overall temporary employment has a negative effect on aggregate (unadjusted) wage growth in Europe.

Exacerbating the *competition effect*, employers may foster discord between workers to prevent the emergence of a unified labor bloc (Bellani and Bosio 2019). For instance, temporary agency workers are used to mitigate wage pressures (Houseman et al. 2003; Drenik et al. 2023), and reforms in temporary employment have worsened conditions for permanent employees (Dolado et al. 2002), in particular for those with lower and middle incomes (Weisstanner 2020). Empirically, evidence of competition between permanent and temporary employees has been found (Voinea 2018). In line with earlier work (Piore 1979; Western and Healy 1999), we suggest that different segments of workers and their interaction with labor market institutions affect the wage-setting process over and above standard macroeconomic factors.

The competition between temporary and permanent workers could invert the established insider-outsider logic (Lindbeck and Snower 1988, 2002), which in its early work on dual labor markets challenged human capital theory (Doeringer and Piore 1971; Rosen 1972; Reich et al. 1973; Piore 1983; Dickens and Lang 1985). Applied to temporary work contracts, the insider-outsider theory suggests that larger hiring and firing transaction costs for insiders create two labor market segments. Insiders enjoy relatively higher economic security than outsiders and can extract rents to the detriment of outsiders by securing higher wages. This results in the wage penalty that is well established: under equal conditions, temporary workers receive smaller paychecks than permanent workers (Kahn 2016; Pavlopoulos 2013). Employers have the incentive to replace permanent employees with temporary ones if the transaction costs associated with hiring and firing permanent employees are lower than the wage penalty (Koutentakis 2008).

As such, the wage penalty can obscure the empirical analysis of the *competition effect*. Changes in the share of temporary workers who suffer from the wage penalty mechanically affect aggregate wages. This results in a *composition effect* on wage growth, which we correct for.

2.2 The role of labor market institutions

The competition effect hypothesized interacts with labor institutions. Olson (1971) proposed that significant but non-encompassing collective interests are detrimental at the societal level. The implications for wage determination are widely discussed (Calmfors et al. 1988; Soskice 1990). In the insider-outsider model, it is easier for insiders to protect their rents at moderate levels of worker organization. At high levels, there are fewer outsiders to bear externalization costs, while at low levels, insiders face stronger competitive pressures.

The simplest measure of the inclusiveness of trade unions is their membership density (Lange 1984). According to Olson's theory, we expect a negligible competition effect in countries with moderately encompassing membership. We expect a large effect in countries with low membership, as insiders lack sufficient power to protect themselves. In contrast, countries with encompassing membership are expected to experience no competition effect, as outsider interests are likely to be internalized.[1] Only the Nordics, Belgium, Malta, and Cyprus still have high membership rates above 40% of the labor force (Appendix Figure A.1). Determining the cut-off between low and medium trade union density (TUD) is less obvious. The distribution of membership rates suggests a cut-off at 20%, which we follow but test for robustness using different cut-off values.

While TUD seems a suitable measure to capture how encompassing unions are, we also use collective bargaining coverage (CBC) for robustness. We assign countries again into three groups: The high group with above 85% coverage comprises mainly countries with an automatic extension of CBC including France, Spain, Belgium, Austria, Finland, and

[1] We demonstrate in the empirical analysis that the competition effect is stronger in countries with weak institutions and that this relationship does not result from weak trade unions causing higher shares of temporary employment.

Sweden, while the low group with less than 35% coverage comprises exclusively of Central Eastern European countries and the UK.

Another measure of robustness is wage bargaining coordination. The high group comprises countries with established norms about wage bargaining, predominantly Nordic and Continental European countries. The middle group includes countries that rely on procedural guidelines for coordination with no regularized patterns, such as Spain, Ireland, and Switzerland. The lowest group again contains traditionally Central Eastern European countries, the UK, and some Mediterranean countries.

We also analyze the impact of employment protection legislation (EPL), which refers to the level of protection provided to permanent workers against individual and collective dismissals. We expect the effect of EPL to be ambiguous: Stricter employment protection may enhance workers' bargaining power but also incentivize hiring temporary workers to avoid dismissal costs, thus intensifying competition between temporary and permanent workers. Furthermore, workers subject to more stringent EPL have more to lose if they were to be replaced with temporary employees, putting permanent workers in a *golden cage*. We categorize countries into high and low EPL groups. Appendix Table E.2 displays the descriptive statistics and country groupings.

2.3 The post-crisis Phillips curve flattening

After the GFC, wage growth fell short of expectations based on the established relationship between unemployment and wages (Kahn 1980; Blanchflower and Oswald 1994). Early signs of a weakened relationship between wages and inflation emerged around 2012 (Anderton and Boele 2015; Ciccarelli and Osbat 2017; Moretti et al. 2019). To explain the wage growth slowdown, conventional measures of labor market slack were adjusted for hidden slack, including discouraged workers and involuntary working time reductions, which had increased during the recession (Hurley and Partini 2017; IMF 2017; Hong et al. 2018; Nickel et al. 2019). Bell and Blanchflower (2019) construct a labor under-utilization index, which enhances the fit of the Phillips curve for Europe (Bell and Blanchflower 2021) and the U.S. (Blanchflower et al.

2022).

Involuntary temporary employment constitutes a specific form of labor under-utilization, commonly used by firms to absorb labor market fluctuations (Draeger and Marx 2017; Hijzen et al. 2017). Temporary workers have lower job stability (Hirsch 2016; Autor and Houseman 2010; Gebel and Giesecke 2011), which suggests that they are more likely to be laid off during unfavorable economic conditions (Costain et al. 2010). Consequently, temporary employment declines faster than permanent employment during economic downturns. An increase in temporary jobs does not provide the same employment opportunities for job seekers and may result in elevated uncertainty among permanent workers as no functional equivalent jobs, and thus outside options for job switchers, open up. If newly created jobs predominantly offer temporary contracts, the reduction in the unemployment rate may thus have a smaller effect on wage growth.

Given the negative correlation between temporary employment and unemployment, fluctuations in temporary employment may balance out the effect of unemployment changes (Appendix Figure A.2). Notably, the incidence of involuntary temporary employment reached a historical peak during the 2013-2017 recovery. This warrants investigating whether the high prevalence of temporary workers has contributed to the Phillips curve flattening. Involuntary temporary workers, desiring permanent contracts, represent part of the labor supply for permanent employment and constitute a form of hidden slack. Therefore, any job created based on temporary contracts should reduce the impact of unemployment rate reductions on wage growth, resulting in a flatter Phillips curve.

3 The empirical approach

Macroeconomic research as discussed in our theory section has relied primarily on country-level data. As available aggregated wage data do not distinguish between permanent and temporary employees, macroeconomic approaches were unable to disentangle competition from composition effects. The distinction is fundamental since variation in our main independent variable – the incidence of involuntary temporary employment – mechanically affects wages as an inherent component of the employment composition. Hence, an observed negative relationship between temporary workers and wage growth may be the result of changes in the composition of workers with different wage levels. We therefore correct for this composition effect to identify the *competition effect*. In contrast to macroeconomic approaches, industrial relations research has distinguished between wages of temporary and permanent employees by using individual-level data to investigate heterogeneous effects on employment and wages at the meso- or micro-level. We contribute by bringing both strands together. This allows us to account for changes in the share of temporary employees to assess whether competition effects influence macroeconomic outcomes.

More specifically, we rely on worker-level data to construct a country-year panel for wage growth of only permanent employees. If a competition effect exists, temporary workers have a negative impact on the wage growth of permanent employees. As permanent workers make up around 90% of Europe's labor force, wage growth of permanent employees is very likely to be close to overall wage developments. Nevertheless, the use of worker-level data allows us to construct a wage growth series for *all* employees in a country, while netting out potential composition effects. Hence, we also estimate the sensitivity of *overall* aggregate wage growth with respect to the prevalence of temporary work in Europe.

3.1 Adjusting wage growth for a changing employment composition

To construct our dependent variable, wage growth, we rely on EU-SILC, a representative population survey containing the longest-running cross-national dataset available with annual information on employment and wages.[2] It allows us to distinguish employees on temporary contracts from permanent ones, which is crucial for our research question. Although the primary focus of EU-SILC lies in collecting representative data on income rather than on the labor market status, the share of temporary employees in total employees in EU-SILC

[2]The EU-LFS does not contain the wage level. The EU-SES is only conducted every four years.

(11.7%) is quite comparable to the respective figure in the Labor Force Survey (13.9%).[3] We discuss the data and aggregation of country-level time series in Appendix B. To confirm the validity of our aggregation, we compare our time series to Eurostat's officially published EU-SILC country-level data (Figure B.1-B.2) and to the OECD's time series on wages based on national accounts as well as on survey and admin data (Figure B.3).

In addition to the *unadjusted* aggregate wage growth variable of all employees, we calculate the wage growth of permanent workers based on the information about employees' contract type. Figure 2 illustrates wage dynamics in Europe for both contract groups separately. What stands out immediately is that wage growth has slowed down since the onset of the GFC, a stylized fact discussed in our theory section. Interestingly, this applies to both groups but seems to be more pronounced in the group of temporary workers. It might be related to the strong *relative* demand for temporary employees before the onset of the crisis (Appendix Figure A.2), which could have accelerated wage growth for temporary workers compared to permanent workers. Likewise, the weakened relative demand during 2008-2014 might explain the observed slower wage growth of temporary workers, while the economic recovery gaining traction from 2015 has fuelled demand for temporary labor, thereby lifting their wages.

To obtain the wage growth of a pseudo-workforce with a constant employment composition over time, we employ *inverse probability weighting (IPW)* (Rosenbaum and Rubin 1983; DiNardo et al. 1996; Fortin et al. 2011). First, we use a logit model to predict the probability of each observation of being in temporary employment per year and country, pairing the base year t (2004 or earliest available) with each of the following years $t+n$: $\ln \frac{p}{1-p} = \beta_0 + \sum_{t=1}^{m} \beta_t x_t$ where x_t is employment contract that we control for.[4] We estimate the re-weighting factors for each year and country separately. Second, we adjust the weights for each observation so that the re-weighted sample has the employment composition with regard

[3] The reported figures represent the weighted average of the share of temporary employees in total employees of all European countries in our sample in the observation period 2004-2017. Note also that the share of temporary employment according to EU-SILC seems to follow a quite similar pattern over time compared to the respective series from the EU-LFS. This is reflected by a relatively high correlation coefficient of 0.88 between both series.

[4] To adjust for changes in employment shares based on other observable characteristics, we repeat the same procedure using gender, migration background, educational attainment, and work experience as additional controls in the logit model (Figure B.1).

to the first year available. For the base year, we keep the original weight $g_1 = g$, whereas for control individuals, we use the predicted probability $p(x)$ to receive the adjusted survey weights $g_{1+n} = g\frac{p(x)}{1-p(x)}$.[5] Finally, we aggregate the worker-level data at the country-year level to obtain our adjusted measure for aggregate wage growth that is based on a counterfactual employment composition constant over time with regard to employment contracts.

Figure 3 presents the adjusted wage growth variable for each country (averaged over the whole sampling period) and the employment composition effect. The latter purely represents a mechanical effect from the changing share of temporary workers over time. Since temporary workers suffer a wage penalty compared to permanent workers, an increasing share of temporary workers lowers the aggregate average wage given a constant penalty. Adjusted wage growth represents the counterfactual rate of wage growth if the share of temporary workers would have remained constant over time. The size of the composition effect is very heterogeneous at the country-level and sizeable in some countries, in particular Denmark, Serbia and the UK. However, interestingly, it does not play a large role for Europe as a whole. Some countries are characterized by substantial wage differences between temporary and permanent workers and have experienced a strong increase in temporary work. However, even in those cases, temporary workers as a share of all employees have only changed by a few percentage points over several years, resulting in a minor impact of employment composition changes on wages. For robustness, we adjust wage growth additionally for employment composition changes by gender, migration background, educational attainment, and work experience, which warrants slightly larger effects (Appendix Figure C.1).

Although the difference between adjusted and unadjusted wage growth is quite small overall, it must be stressed that only by adjusting can we identify the underlying mechanism that impacts wage growth. Without the adjustment for employment composition, we would not know whether composition or competition is driving our results.

[5] It is not possible to fully rebalance continuous covariates with a semiparametric method, but we can eliminate a large part of the variation in individual worker characteristics over time with our re-weighting procedure.

3.2 Estimating factors of wage growth

The most widely used empirical model to study the determinants of wage growth is the wage Phillips curve. The traditional wage Phillips curve relates nominal wage growth to labor market slack. Additional determinants typically considered are (expected) inflation and labor productivity growth (Nickel et al. 2019). We use such an augmented Phillips curve model to study the impact of dualization on nominal wage growth in Europe. We estimate a standard reduced form equation in a panel data framework of the form:

$$\dot{W}_{i,t} = \alpha_1 + \alpha_2 \mathrm{U}_{i,t} + \alpha_3 \mathrm{Prod.}_{i,t} + \alpha_4 \mathrm{Infl.}_{i,t} + \alpha_5 \mathrm{Invol.\ Temp}_{i,t} + \mu_i + \tau_t + \epsilon_{i,t} \tag{1}$$

As outlined in the section above, our dependent variable is nominal wage growth obtained from EU-SILC. As a benchmark, we first study the dynamics of the unadjusted aggregate wage growth to represent the workhorse Phillips curve model. In the second step, we analyze the nominal wage growth of permanent workers only, and we finally implement our main dependent variable, which is nominal wage growth net of composition effects ($\dot{W}_{i,t}$). While most studies estimating wage Phillips curves use quarterly data[6], we have to stick to an annual frequency (as in the original contribution by Phillips (1958) or more recently by Kiss and Van Herck (2019)) as the computation of our dependent variable is only feasible based on yearly data. Our sample includes 30 European countries (i) and ranges from $t = 2004, ..., 2017$, which leaves us with roughly 340 observations.[7] We intentionally choose a static representation as we do not observe any persistence in wage dynamics (likely due to the annual frequency of our sample). Moreover, as the time-invariant country effects (μ_i) are correlated with the regression variables, we employ the fixed-effects estimator (FE), where unobservable country effects are assumed to be fixed (and not random). We compute standard errors clustered at the country-level in all specifications to control for potential serial correlation in the error term within each country.

[6]Examples are Bonam et al. (2021), Nickel et al. (2019) and Bulligan and Viviano (2017)

[7]Data for some countries are only available after 2004. We drop 19 observations when there was a break in the time series of wages according to Eurostat (due to a change of source or survey methodology).

As a baseline, we use the conventional labor market slack indicator, which is the headline unemployment rate $U_{i,t}$, but we also consider several other measures of slack for robustness. Further, we control for the impact of labor productivity ($\mathrm{Prod}_{i,t}$) on wages, which we measure as the growth rate of real output per employment, as well as for inflation ($\pi_{i,t}$). Studies using quarterly wage growth data often employ (one quarter) lagged inflation implying backward-looking expectations (Ramskogler 2021; Nickel et al. 2019; IMF 2017). Given the annual frequency of our data, we assume a contemporaneous effect from inflation (measured as the annual change in the harmonized index of consumer prices) on nominal wage growth.[8]

Finally, and most importantly, we add to our Phillips curve specification a variable to identify the competition effect. So far, studies exploring the impact of dualized labor markets on wages have considered *overall* temporary employment (Ramskogler 2021; Bellani and Bosio 2019). However, the limitation is that not all temporary workers look for a permanent contract. We identify the competition effect, by focusing on *involuntary* temporary employees as a share of the active working-age population ($\mathrm{Invol.\ Temp}_{i,t}$). This segment of disadvantaged workers prefers a permanent contract over their temporary one, which we expect to cause the competition effect. A detailed description of the measurement of all variables and their sources is included in Appendix Table E.1.

4 Results

In the first part of this section, we present estimation results concerning the identification of the competition effect and its macroeconomic consequences (Tables 1 and 2). In the second part, we test whether the magnitude of the competition effect depends on labor market institutions (Table 3). Finally, we explore whether the rise in involuntary temporary employment can explain the flattening of the Phillips curve in Europe during the post-GFC recovery period (Table 4).

[8] We also consider a survey-based measure capturing forward-looking inflation expectations provided by the European Commission (expected inflation). As this variable is not available for two of our countries (Switzerland and Norway) and does not improve the explanatory power, we stick to realized consumer price inflation.

4.1 Identifying the competition effect at the macroeconomic level

We present the workhorse Phillips curve specification augmented by involuntary temporary employment in column (1) of Table 1. The coefficient estimates have the expected signs and are statistically significant. Labor productivity has a positive effect on wages. We also observe that inflation drives up wages with a regression coefficient of around 1.[9] By contrast and as expected, an increase in the unemployment rate reduces nominal wage growth. Yet, $Invol.\,Temp_t$ – our variable of main interest – is also negatively associated with nominal wage growth and is statistically significant. A rise in the share of involuntary temporary employees by 1 percentage point leads to a decrease in nominal wage growth by almost 1 percentage point. As we have considered the *unadjusted* growth rate of wages so far, the coefficient estimate captures both potential composition *and* competition effects. However, before we alter the dependent variable to isolate the competition effect, we include time dummies in our model to control for common shocks that might have affected wage dynamics equally across countries over time, such as the GFC. In fact, a test of joint significance shows that the time dummies have high explanatory power. Their inclusion also reduces coefficient estimates of all variables except $Invol.\,Temp_t$, as we can see in column (2). This is particularly true for inflation, which becomes statistically non-significant.[10] As time dummies are significant in all the following model specifications, we included them to avoid biased estimates (Baltagi 2005).

We now alter our dependent variable in column (3) by considering the nominal wage growth of employees with permanent contracts only. This allows us to estimate the competition effect, as we isolate the part of wage growth that cannot be affected by changes in relative weights between temporary and permanent workers. Compared to column (2), all coefficient estimates remain broadly the same. This result has two main implications. First, it strongly supports our thesis that the incidence of a dualized labor market has negative spillover effects

[9] This finding likely is linked to the annual context of our estimations (Kiss and Van Herck 2019). While it is not very common in the literature to use quarterly data, Rusinova et al. (2015) show that if four lags of inflation are considered in quarterly estimations, the aggregate effect again accumulates to close to 0.9.

[10] Obviously, price dynamics across countries have followed a very similar pattern over time. Certainly, a common driver of inflation across countries could be oil, which is known to significantly affect consumer price dynamics.

on the dynamics of wages of employees with permanent contracts. This is consistent with Bellani and Bosio (2019) who find that the density of temporary contracts within occupation- and age-specific groups negatively affects average wages for permanent workers belonging to the same group. In addition to their findings, our results show that competition effects are also relevant in a macroeconomic context, where other important wage growth determinants like the unemployment rate are accounted for. The second important implication is that *composition* effects seem to be negligible in Europe in the period 2004-2017. A first indication of these rather small composition effects is the relatively low contribution of employment composition to adjusted wage growth across countries presented in Figure 3. In column (4) we show that the sensitivity of wage growth with respect to the unemployment rate decreases when the Phillips curve is specified without controlling for temporary employment.

Unlike in previous studies (Ramskogler 2021; Bellani and Bosio 2019), our empirical setting allows us to focus on *involuntary* (rather than on *overall*) temporary employment to measure the degree of labor market dualization. To reveal whether this is indeed the relevant measure in our context, we add the share of *voluntary* temporary employees and report the results in column (5). Comparing the coefficients of both indicators reveals *involuntary* temporary employees drive wage growth of permanent workers, while the impact of workers, who have voluntarily chosen to have a temporary contract ($Vol.\ Temp_t$) is non-significant. Hence, when measuring dualization, it is crucial to quantify those employees who would prefer to be employed on a permanent basis. Note also that the magnitude of the coefficient estimate (and its statistical significance) would drop substantially if we considered overall temporary employment (instead of $Invol.\ Temp_t$). To proxy labor market dualization, our results, thus, strongly suggest that – whenever feasible – *involuntary* rather than *overall* temporary employment should be considered.

To investigate the impact of involuntary temporary employment on overall aggregate wage growth, we re-estimate specification (2) by employing *adjusted* wage growth, i.e., wage growth net of composition effects. The results are depicted in column (1) of Table 2 and show almost unchanged coefficient estimates (compared to model (2) in Table 1). This is consistent with our previous observation, namely that composition effects are empirically only of minor importance. Further, our results resemble Ramskogler (2021), who finds a significant negative effect from temporary employment on *unadjusted* wage growth in Europe. Additionally, we confirm that the underlying mechanism behind the observed negative relationship arises from a *competition* rather than a *composition* effect.[11]

Before turning to the issue of reverse causality, we provide two interesting extensions of this result. First, we assess the *economic* significance of dualized labor markets by re-estimating specification (1) based on standardized variables. As reported in column (2), involuntary temporary employment turns out to be the most relevant determinant for wage growth followed by unemployment.[12] However, taking into account the uncertainty surrounding the parameter estimates, both variables are equally meaningful in explaining nominal wage growth.[13] Hence, involuntary temporary employment has been at least as important as unemployment in shaping nominal wage dynamics in Europe.

While in specification (1) we control for composition effects with respect to the type of employment contract, other possible aspects of composition could affect wage growth as well. Some demographic groups may have more bargaining power (e.g., prime-age native-born males) or a higher marginal productivity than others. The composition of who is selected into temporary jobs may change over the business cycle. Hence, in addition to contract type, we adjust wage growth for changes in the employment composition by gender, migration background, educational attainment, and work experience. As highlighted in column (3), this new measure does not alter the observed results with respect to the competition effect. However, the impact of productivity on wage growth becomes non-significant. This is most probably because productivity is largely captured at the worker-level by having netted out changes in the share of skilled and experienced workers.

We now turn to the issue of a potential simultaneity bias arising from reverse causality be-

[11] As a robustness check, we exclude one country at a time from the sample to show that the result does not depend on one particular country (Appendix Table D.1).

[12] An increase in the rate of involuntary temporary employment by one standard deviation leads to a drop in nominal (composition adjusted) wage growth by half a standard deviation.

[13] A test on parameter equality is not rejected.

tween nominal wage growth and labor market slack. Usually, reverse causality is approached by inserting the slack variable in its one-period lagged form into the Phillips curve model (Ramskogler 2021; Byrne and Zekaite 2020; Nickel et al. 2019). While this is certainly a valid approach when using quarterly data, it is not feasible in our case given the annual frequency of the data. Fortunately, in the case of reverse causality, the fixed-effect estimate of the impact of unemployment on wage growth would be downward biased rather than upward, as higher wage growth should cause higher labor market slack (IMF 2017; Wooldridge 2009). The same logic applies to the dualization measure. If wage growth accelerates, it is presumable that employers increasingly demand temporary employees as they are cheaper and associated with lower firing costs. Hence, our findings concerning the importance of temporary employment for wage growth are likely not mistaken even in the presence of reverse causality.

An alternative approach to account for a potential simultaneity bias is to use instrumental variable techniques. As exogenous instruments are not at hand, neither for unemployment nor for temporary employment, we use internal instruments, i.e., time lags of the variables in the model. In particular, we employ the difference GMM estimator[14] (Arellano and Bond 1991; Blundell and Bond 1998) and treat both variables as endogenous (by using the lagged levels of the variables as instruments). As displayed in column (4), involuntary temporary employment and unemployment have the expected negative signs and are statistically significant. However, compared to the fixed effect estimation in column (1), we observe an increase in the coefficient estimate for both variables. Obviously, controlling for simultaneity has an effect on the estimates in the direction that we expected. A very similar result can be found in Bellani and Bosio (2019). Finally, in column (5) we add inflation to the set of endogenous variables and show that this alteration does not have any significant influence on the estimation outcome.

[14] Our dependent variable is not persistent. For this reason, we choose a static representation and the difference rather than the system GMM estimator.

4.2 The effect of institutions

As highlighted in the theory section, the magnitude of the competition effect may depend on a country's labor market institutions. We take account of these considerations in Table 3, where we interact involuntary temporary employment with different institutional variables. We consider the wage growth of permanent workers as the dependent variable in all specifications. In addition to interaction effects, we control for the direct effect that institutions might exert on wage growth. This rules out a potential omitted variable bias that might arise if weak trade unions were to cause higher shares of temporary employment due to their inability to prevent the substitution of good jobs with bad jobs.

Following our country grouping based on trade union density, we report group-specific differences in the competition effect in column (1). As we anticipated, the competition effect is significant and large only in countries with low trade union density. Moreover, we can conclude that the competition effect does not arise because weak trade unions cause higher shares of involuntary temporary employees, as we have controlled for the direct impact of institutions. Rather, the observed effect in countries with weak institutions results from the fact that permanent employees cannot use union power to protect themselves from negative wage pressures caused by temporary employment. Trade union density (TUD_t) affects wage growth positively, as expected (Kahn 1979; Stansbury and Summers 2020).

Analogously, our findings are confirmed when investigating different proxies for how wage bargaining encompasses workers. These proxies include collective bargaining coverage (CBC) as in specification (2) or wage bargaining coordination (Coord) as in specification (3). In all these cases, the competition effect is most pronounced at the lowest institutional level.[15] Our results, thus, support the hypothesis that competition effects from labor market dualization heavily depend on domestic labor market institutions.

Further, we stress that the competition effect is higher in countries with more stringent

[15] Moreover, we find a negative impact from CBC and Coord on wage growth, which corresponds to common expectations (Soskice 1990; Hancke and Soskice 2003). This happens as higher levels of centralization and coordination help to prevent wages from taking an inflationary turn (i.e., internalizing negative externalities). Coordination is expected to be the most relevant factor (Soskice 1990), which is confirmed by our findings.

employment protection legislation (EPL) as shown in specification (4). This likely is the result of the golden cage effect: Permanent workers in countries with more stringent EPL, have more to lose if they were to be laid off. The monetary value of a job can be understood as a function of the wage received per period and the probability of future job loss, which amounts to the discounted income stream expected. In this logic, workers trade job security, corresponding to the likelihood of wage receipt in the future, against higher wages in the present. More stringent EPL increases the chance of continued employment and thus can correspond to lower wages.

In the last specification of Table 3, we want to highlight that the competition effect, although it varies in magnitude across country groups, is significant in determining *weighted* aggregate wage growth in Europe. In fact, the competition effect remains highly significant and increases in magnitude when we put more weight on countries that are larger, as we demonstrate in column (5).[16] This observation is consistent with the fact that low TUD countries expose a particularly strong competition effect, as the countries belonging to this group (comprising 14 countries) make up more than 75% of overall employment in Europe and drive the aggregate weighted effect.

Finally, empirically analyzing differences across country groups always involves choosing the "right" threshold that divides countries into the respective groups. In the case of TUD and Coord, a disproportionately high share of countries are clustered in the first group (i.e., countries with the weakest institutions) as rationalized in the theory section. This categorization might work against finding a statistically significant effect for the medium and high groups. As a robustness check, we consider an alternative clustering procedure that groups the countries more evenly by using terciles of the institutional variables. We report our results in Appendix Table D.2.

In columns (1) to (4), we show that our results hold when the number of countries does not vary across groups. In column (5) of Table C.2, we base our country grouping on a joint set of institutional variables to form three equally sized clusters. Concretely, we perform

[16] The relative weight of each country is based on the number of employed persons in 2005.

a principal component analysis (PCA) to understand the correlation of the institutional variables, as the PCA reduces their multi-dimensional character by identifying their common grounds (components). We consider the first resulting component, which explains 75% of the overall variation of TUD, Coord and CBC.[17] Using terciles of the constructed index to form the grouping again confirms our main finding. The magnitude of the competition effect decreases with the strength of institutions. Lastly, in column (6) we employ an index for union strength developed by Metten (2021) who uses a more sophisticated theoretically informed PCA to identify determinants of trade union strength. Employing this index again supports our hypothesis that the competition effect is large in countries where insiders do not have sufficient power to shelter themselves from competitive pressure.[18]

4.3 The Phillips curve flattening

So far, we have shown that competition effects play a statistically significant role in explaining aggregate nominal wage growth and that they interact with the institutional dimension. Does this help to understand the observed flattening of the Phillips curve in Europe after the GFC? To tackle this question, we extend the Phillips curve framework by allowing for a different unemployment parameter after the crisis. This allows us to test the Phillips curve flattening and study possible interaction effects with temporary employment. We employ adjusted wage growth in all specifications to rule out possible composition effects and to obtain results that reflect the wage dynamics of both temporary and permanent employees. To investigate post-crisis differences, we first interact unemployment as well as involuntary temporary employment with a post-crisis dummy that equals 1 for the period 2013-2017 and 0 for the preceding period.[19] Second, following the hidden slack literature, we construct a labor market slack measure by summing up unemployed and involuntary

[17]We have not considered EPL in constructing the index, as EPL does not load the first component of the PCA in the same direction as the remaining three institutional variables. Hence, EPL does not seem to capture the inclusiveness of wage bargaining but other underlying institutional factors. Recall that the competition effect is more pronounced for countries with high EPL.

[18]The index is not available for Malta. Moreover, due to several missing values at the beginning and end of the individual time series, we have not imputed missing values and therefore did not include the index as an additional regressor into the model.

[19]We have tested other thresholds as well. It turned out that the break in slope parameters is most pronounced when the post-crisis period is defined from 2013 onward.

temporary employees ($Slack_t$) to study their joint impact on wage growth before and after the crises.[20]

The corresponding results are summarized in Table 4. The first column shows a model that allows for a crisis interaction term on the unemployment rate without considering temporary employment. The slope parameter of the unemployment rate is statistically different across the two time periods and points to a decreased sensitivity of wage dynamics to unemployment of more than 50% since the post-crisis period. While a decrease in the unemployment rate boosted wage growth by 0.56 percentage points before 2013, this sensitivity declined to 0.18 percentage points[21] in the post-crisis period. Our results, thus, support the empirical findings in the literature that indicate a lower explanatory power of labor market slack measures in the post-crisis period (Byrne and Zekaite 2020).

In column (2), we add temporary employment into the model and allow for different slope parameters on this variable as well. Two things stand out. First, the sensitivity of nominal wage growth with respect to involuntary temporary employment remains largely unchanged. Even though we observe an increase in the impact of temporary employment after 2012, it is not statistically significant. Second, adding labor market dualization into the model does not help to understand the flattening of the Phillips curve as the slope parameter of $U_t * \textit{post-crisis}$ remains unchanged. Interestingly though, employing the variable that summarizes unemployment and temporary employment (i.e., $Slack_t$) leads to a different conclusion, as can be seen in column (3). The flattening is still observable since the slope parameter on $Slack_t * \textit{post-crisis}$ is positive, but it is smaller and statistically not significant.

Given the thus far inconclusive results concerning the role of temporary employment for the Phillips curve flattening, we follow the literature on hidden slack by considering a broader measure of the unemployment rate as well as involuntary part-time employment (as another source of potential hidden slack) and re-estimate the first three specifications accordingly. We extend the headline unemployment rate by additionally considering discouraged as well

[20]Both Unemp_t and Invol. Temp_t, representing $Slack_t$, are measured as a share of the active working-age population.

[21]The slope parameter of unemployment after 2012 is obtained as follows: -0.56 + 0.38 = -0.18.

as marginally attached workers (U-5). Moreover, we account for employees who work part-time but do so involuntarily. Note however, that unlike by Bell and Blanchflower (2019), we are not able to account for labor under-utilization based on desired hours of work due to data availability. Instead, we have to rely on headcounts to capture the degree of underemployment. The flattening disappears when adding involuntary temporary employment next to the broader unemployment measure (columns (4) to (6)). The same holds true when investigating overall slack as defined above.[22]

Finally, in the remaining three specifications, we employ the cyclical components of our independent labor market variables. In specifications (7) and (8), we rely on the concept of the non-accelerating wage rate of unemployment (NAWRU) and consider the unemployment gap arising between the headline unemployment rate (U-3) and the NAWRU. In model (8), we add the cyclical component of involuntary temporary employment, which we compute by applying an HP filter.[23] For the final model, we use the same filtering technique to de-trend the time series of labor market slack ($Slack_t$). The results of the last three models are quite similar to the ones obtained when considering the narrow definition of the unemployment rate (U-3). Adding labor market dualization does not explain the flattening of the Phillips curve as can be seen from a comparison of columns (7) and (8), but dampens the flattening in specification (9), where unemployment is considered jointly with involuntary temporary employment.

Overall, the presented results concerning the interaction between dual labor markets and the flattening of the Phillips curve are not robust. In our view, though, the findings point to a potential role of involuntary temporary employment in the hidden slack debate. One reason for the inconclusive results might be the relatively short time period of our analysis. Adding more observations might eventually result in more robust findings, especially if temporary employment were to increase further. Moreover, becoming more granular concerning the

[22]That is, summing up the broader unemployment measure U-5 and involuntary temporary employment.

[23]To avoid the end-point problem of the HP filter (Orphanides and van Norden 2002), we consider the most recent data, which are available until 2021 for all countries in our sample. Following Ravn and Uhlig (2002), we set the smoothing parameter to 6.25. As the filtering technique does not allow gaps within the time series, we impute four observations in AT (2004,2005), MT (2004) and ES (2005). However, we do not use these observations for estimating our models.

slack variable (e.g., by considering the variable created by Bell and Blanchflower (2019)) could help to improve estimation efficiency for future research.

5 Conclusion

In this paper, we have demonstrated that competition between involuntary temporary and permanent workers has suppressed wage growth in Europe. This means that the higher the incidence of temporary workers who are involuntarily on a temporary contract, the lower the growth rate of wages. The effect is clearly present when investigating (i) the rate of wage growth of permanent employees alone and when employing (ii) adjusted aggregate wage growth that nets out potential composition effects caused by fluctuations in the share of temporary employment. Moreover, we have illustrated that involuntary temporary employment has been at least as important as the unemployment rate in shaping wage dynamics in Europe. Hence, the competition effect is not only statistically but also economically significant.

On top, the cross-country nature of our analysis also allowed us to investigate the role of institutions. We have shown that the competition effect is more pronounced when wage bargaining institutions are weak, which is consistent with industrial relations scholarship. Crucially, our findings are robust when we put more weight on larger countries, thus ruling out the possibility that only small countries drive the results.

Finally, we have presented some tentative evidence that the competition effect might help to understand the strange flattening of the Phillips curve in Europe during the post-GFC recovery period. In fact, we have shown that accounting for the incidence of involuntary temporary employees in defining labor market slack explains the flattening of the Phillips curve to some extent. However, our findings in this regard are the least (statistically) significant and thus leave ample room for further research.

Overall, our analysis shows the important macroeconomic consequences of the dualized structure of labor markets. Despite the recent uptick in wage growth, the entrenchment of temporary employment calls for macroeconomic policies that are cognizant of the labor market structure.

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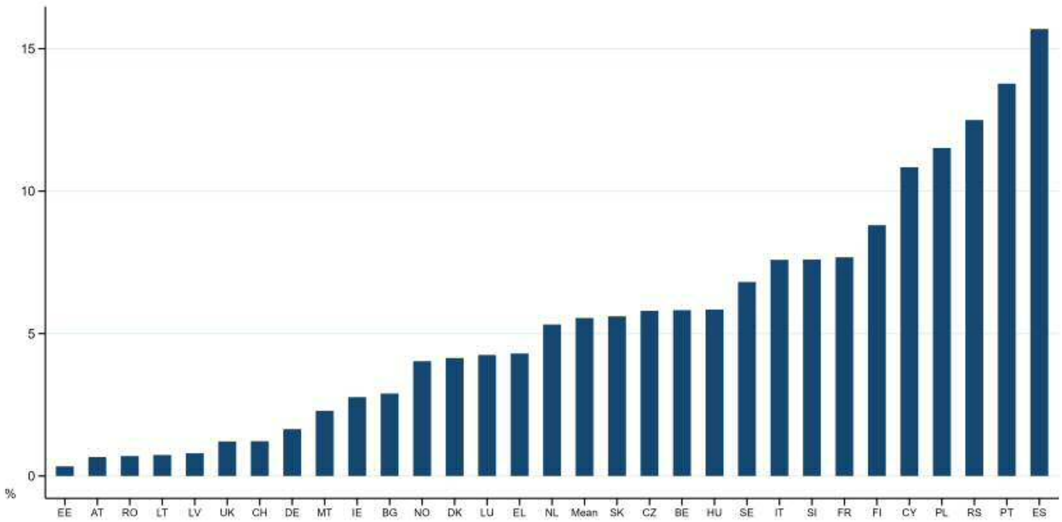
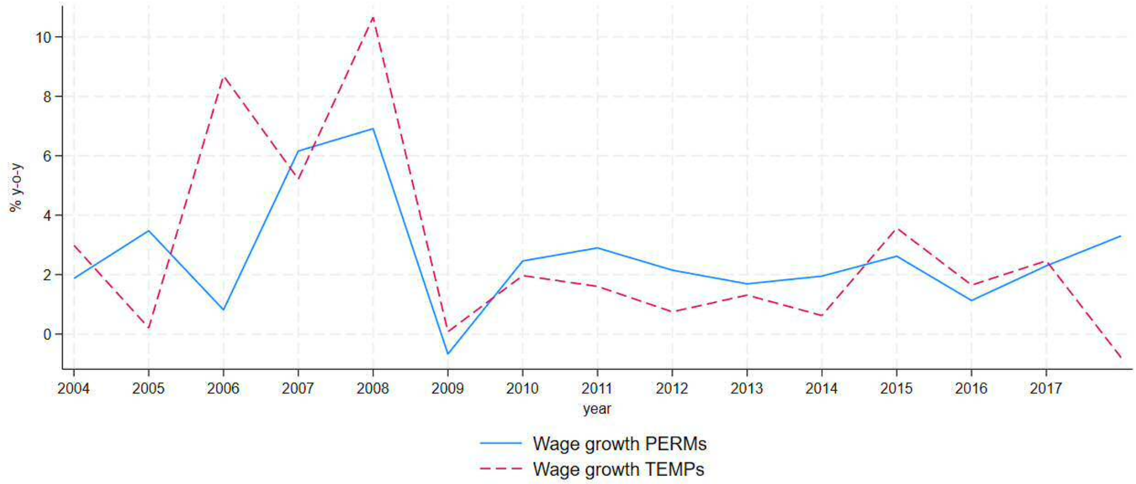
Figures and Tables

Figure 1: Involuntary temporary employment in Europe, 2017



Note: Mean is an unweighted average of all countries shown. Involuntary temporary workers are shown as a share of the labor force aged 15 to 74.
Source: Eurostat/EU-LFS: lfsa_etgar and lfsa_agan.

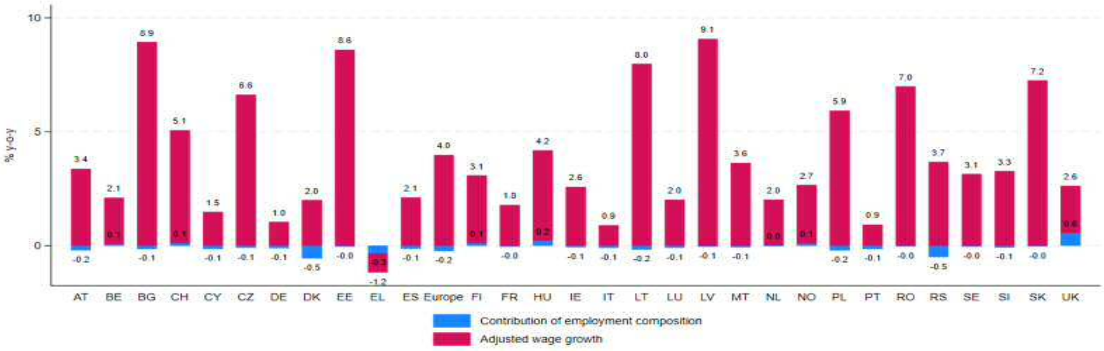
Figure 2: Wage growth permanent and temporary workers



Note: The average annual change in nominal wages is shown as a weighted average for European countries.

Source: EU-SILC.

Figure 3: Wage growth 2004-2017 divided in contribution of employment composition and adjusted wage growth



Note: Wage growth is adjusted for a changing employment composition by contract. Wage growth refers to the average annual change in nominal wages 2003-2017. Europe refers to the simple average of all countries shown.

Source: Authors' computations based on EU-SILC.

Table 1: Identifying the competition effect

	all workers, unadjusted		permanent contract workers		
Dep. var.: wage growth	work-horse PC (1)	incl. time dummies (2)	competition effect (3)	excl. Temp (4)	incl. vol. Temp (5)
$Prod_t$	0.57***	0.33**	0.34**	0.33**	0.34**
	(3.43)	(2.21)	(2.46)	(2.24)	(2.46)
$Infl_t$	0.90***	0.55	0.49	0.54	0.46
	(3.00)	(1.18)	(1.06)	(1.20)	(0.97)
U_t	-0.66***	-0.51***	-0.49**	-0.44**	-0.51**
	(-3.91)	(-2.90)	(-2.64)	(-2.09)	(-2.61)
$Invol. Temp_t$	-0.96**	-0.92***	-0.98***		-1.04***
	(-2.66)	(-2.82)	(-3.21)		(-3.17)
$Vol. Temp_t$					-0.34
					(-1.18)
Cons	11.88***	11.19***	11.35***	5.56**	13.28***
	(5.31)	(3.82)	(3.91)	(2.58)	(3.13)
Model	FE	FE	FE	FE	FE
TimeD	excl.	incl.	incl.	incl.	incl.
N	344	344	344	344	344

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis and are based on cluster-robust standard errors by country.

Table 2: The impact of the competition effect on *adjusted* wage growth

Dep. var.: adjusted wage growth	adjusted by contract (1)	standardized coefficients (2)	adjusted by all controls (3)	reverse causality (4)	(5)
$Prod_t$	0.33**	0.14**	0.29	0.35**	0.35***
	(2.26)	(2.26)	(1.58)	(2.51)	(2.61)
$Infl_t$	0.54	0.16	0.53	0.47	0.44
	(1.16)	(1.16)	(1.13)	(1.08)	(0.85)
U_t	-0.50***	-0.33***	-0.58***	-0.57***	-0.65***
	(-2.78)	(-2.78)	(-3.47)	(-3.59)	(-4.18)
$Invol.\,Temp_t$	-0.95***	-0.54***	-0.89**	-1.67**	-1.87**
	(-2.92)	(-2.92)	(-2.74)	(-2.06)	(-2.29)
Cons	11.34***	-0.01	11.68***	15.88***	17.73***
	(3.91)	(-0.09)	(3.94)	(3.24)	(3.59)
Model	FE	FE	FE	GMM	GMM
Ar1				-2.81	-2.78
Ar2				-0.72	-0.71
Hansen				12.83	14.74
Hansen p-val				0.80	0.97
TimeD	incl.	incl.	incl.	incl.	incl.
N	344	344	343	344	344

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis and are based on cluster-robust standard errors by country. Specification (3) includes wage growth adjusted by contract type, gender, migration, education, and work experience. Specifications (4) and (5) are estimated by first difference GMM (using orthogonal deviations). We use the Stata command xtabond2 and employ the second level lag (up to 11 lags) of the endogenous variables as instruments. As the cross-section dimension is rather small (i.e., 30 countries), we use standard IV instruments rather than GMM-type instruments to limit the instrument count (by using the collapse option (Roodman 2009)). Specification (4) treats only U_t and $Invol.\,Temp_t$ as endogenous, while specification (5) assumes that all variables are endogenous except $Prod_t$.

Table 3: The competition effect and the role of institutions

Dep. var.: wage growth of perm. workers	Inst: TUD (1)	Inst: CBC (2)	Inst: Coord. (3)	Inst: EPL (4)	weighted sample (5)
$Prod_t$	0.17	0.29	0.28*	0.04	0.04
	(0.89)	(1.46)	(1.83)	(0.25)	(0.15)
$Infl_t$	0.59	0.44	0.50	-0.20	-0.25
	(1.53)	(0.98)	(1.17)	(-0.55)	(-0.38)
U_t	-0.60**	-0.61***	-0.58***	-0.31**	-0.45**
	(-2.44)	(-3.17)	(-3.06)	(-2.09)	(-2.08)
$Invol.\,Temp_t$					-1.13***
					(-3.39)
...*low Inst*	-1.40***	-1.39**	-1.47***		
	(-4.89)	(-2.44)	(-3.64)		
...*med. Inst*	0.36	-0.20	-0.76		
	(0.32)	(-0.50)	(-1.25)		
...*high Inst*	-0.11	-0.81	-0.29		
	(-0.35)	(-1.40)	(-1.05)		
$Invol.\,Temp_t$					
...*low EPL*				-0.74**	
				(-2.52)	
...*high EPL*				-1.15**	
				(-2.11)	
TUD_t	0.43***				
	(3.49)				
CBC_t		-0.09*			
		(-1.95)			
$Coord_t$			-1.24*		
			(-1.86)		
EPL_t				-0.75	
				(-0.53)	
Cons	-2.33	15.64***	14.80***	12.99***	14.33***
	(-0.55)	(4.54)	(3.67)	(3.50)	(3.91)
Model	FE	FE	FE	FE	FE
TimeD	incl.	incl.	incl.	incl.	incl.
N	300	302	344	278	344

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis and are based on cluster-robust standard errors by country. The institutional variables are employment protection legislation (EPL), trade union density (TUD), collective bargaining coverage (CBC), and coordination of wage setting (Coord). As the CBC time series has a lot of gaps, we impute missing values with lagged available values. Column (5) represents estimates from a weighted regression. The relative weight of each country is based on its number of employed persons in 2005.

Table 4: Phillips curve flattening: the role of dualization; Dep. variable: adjusted wage growth

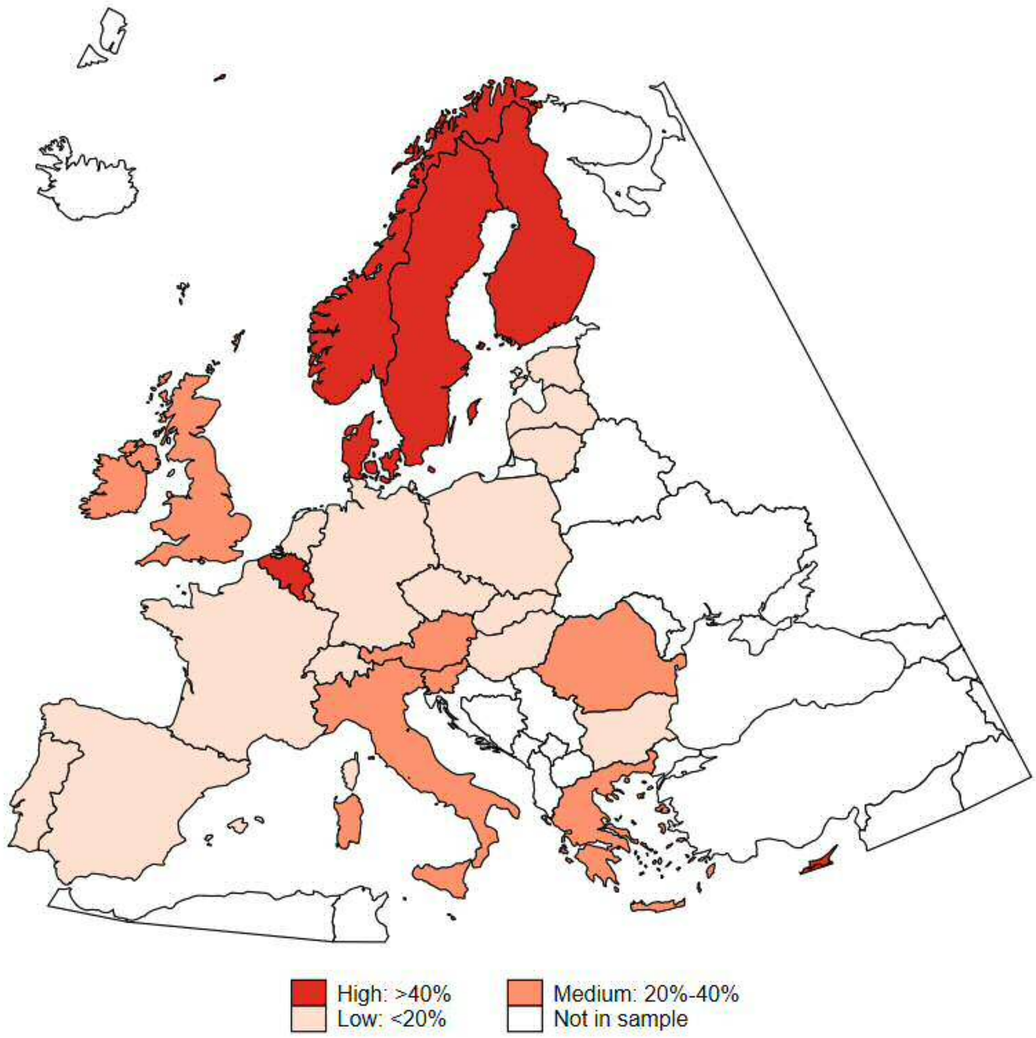
	headline U			broad U & invol. part-time			de-trended labor market var.		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$Prod_t$	0.34	0.34	0.35	0.36	0.34	0.35	0.29	0.26	0.33*
	(1.53)	(1.66)	(1.67)	(1.56)	(1.58)	(1.58)	(1.50)	(1.46)	(1.77)
$Infl_t$	0.99**	0.98**	0.92**	1.06**	1.04**	0.99**	0.81**	0.70*	0.44
	(2.65)	(2.51)	(2.52)	(2.70)	(2.62)	(2.66)	(2.22)	(1.93)	(1.16)
U_t	-0.56**	-0.58**		-0.44*	-0.50***		-1.04***	-1.06***	
	(-2.17)	(-2.63)		(-2.03)	(-2.79)		(-2.81)	(-3.07)	
$U_t * post\text{-}crisis$	0.38*	0.38**		0.30*	0.26		1.06***	1.05***	
	(1.80)	(2.14)		(1.76)	(1.61)		(2.87)	(3.04)	
$Invol.\ Temp_t$		-0.92**			-0.95**			-1.80***	
		(-2.34)			(-2.29)			(-3.51)	
$Invol.\ Temp_t * post\text{-}crisis$		-0.10			-0.08			1.19	
		(-0.90)			(-0.59)			(1.28)	
$Slack_t$			-0.56**			-0.51**			-2.28***
			(-2.46)			(-2.49)			(-4.08)
$Slack_t * post\text{-}crisis$			0.17			0.14			1.49*
			(1.47)			(1.21)			(1.78)
$Invol.\ Part._t$				-0.08	0.40	0.36			
				(-0.19)	(0.78)	(0.72)			
Cons	2.38	8.55***	6.38**	2.76	8.42***	6.17**	1.39	1.72	1.85
	(1.59)	(3.16)	(2.75)	(1.58)	(3.17)	(2.49)	(1.24)	(1.50)	(1.55)
Model	FE	FE	FE	FE	FE	FE	FE	FE	FE
TimeD	incl.	incl.	incl.	incl.	incl.	incl.	incl.	incl.	incl.
N	285	285	285	285	285	285	285	285	285

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis and are based on cluster-robust standard errors by country. $Slack_t$ is measured as the sum of the respective unemployment rate and involuntary temporary employment in models (3) and (6). Models (7) to (9) use the cyclical components of the employed labor market variables, i.e., U_t is based on the NAWRU (OECD) and $Invol.\ Temp_t$ and $Slack_t$ are based on an HP filter, where $Slack_t$ is defined as the sum of the headline unemployment rate and involuntary temporary employment.

Online appendix for: Begging thy coworker – Labor market dualization and the slow-down of wage growth in Europe

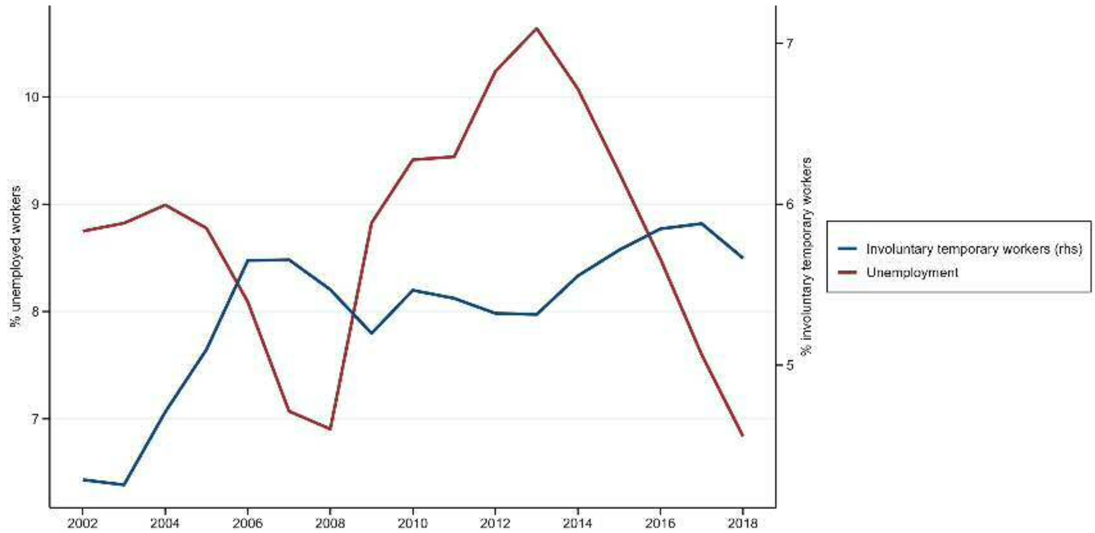
A Appendix: Stylized facts

Figure A.1: Trade union density in Europe



Note: Trade union density refers to dependent employees that are trade union members as a share of the total number of dependent employees. Numbers are reported as an annual average 2003-2018.
Source: OECD/AIAS ICTWSS database.

Figure A.2: Labor market dualization in Europe



Note: Weighted averages of 28 European countries in our sample excluding CH and RS due to data availability for the entire time series. Values for AT 2004, 2005, ES 2005, MT 2004 and FR 2002 are imputed for the graph only. Unemployment and involuntary temporary workers as a share of the labor force aged 15 to 74.
Source: Eurostat/EU-LFS: lfsa_etgar, lfsa_eegais, lfsa_agan.

B Appendix: EU-SILC data

B.1 Wage growth aggregated vs. published (net)

The EU-SILC data has been the established standard for cross-country income comparisons in Europe. The survey combines demographic variables from the current year with wages from the previous year (except for Ireland and the UK) (Eurostat 2018). Since we focus on wage growth, we use the year of the reported wage, i.e., one year prior to the other data collected. We use all waves 2004-2018 and hence yield an (unbalanced) macro-panel of wage data spanning the period 2003-2017.[1] We use *gross employee cash or near cash income (PY010G)* for dependent workers as our main variable for wages since we are interested in their pre-tax wages. We rely on the *number of hours usually worked per week in their main job (PL060)* to compute hourly wages at the individual level. We compute aggregate measures at the country level using the *personal cross-sectional weight (PB040)* and compute the average annual change in nominal hourly wages.

A large effort is put into the harmonization of definitions and variables across countries, although some caveats apply due to national differences in data collection. The income reference period for most countries is the calendar year previous to the survey year with two exceptions: Ireland and the UK. In Ireland, the income reference period is the last twelve months. In the United Kingdom, the current income is annualized and aims to refer to the current calendar year, i.e., weekly estimates are multiplied by 52, monthly by 12 (Eurostat 2018). Since this data started being collected for 2004, an increasing number of countries have shifted to rely on national registries to construct or correct the wage variables strengthening accuracy and reliability (for a detailed overview see Goedeme and Trindade (2020) and Lohmann (2011)). We carefully examine national particularities in our data cleaning and aggregation procedure following Trindade and Goedeme (2019) on the income variables and GESIS (2021) in addition to the EU-SILC methodological guidelines and national quality reports to ensure maximum cross-country comparability. However,

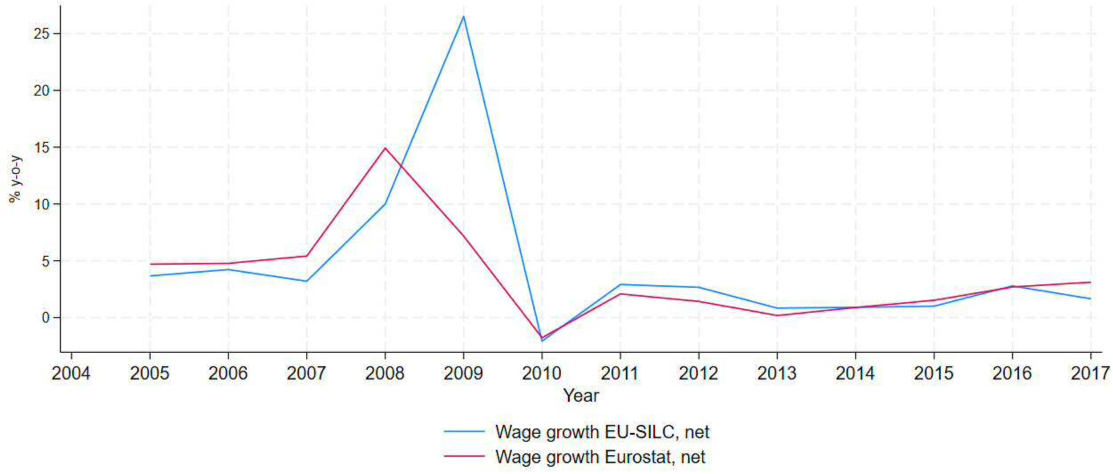
[1] The effective sample starts only in 2004 because our dependent variable is wage growth.

processing and aggregating individual-level data always entails a series of small decisions that can affect the outcome. To be transparent, we document our aggregation procedure in detail in our annotated Stata code. Eurostat does not publish sufficient details on their procedure for data processing and aggregation that we could follow. For a detailed discussion of EU-SILC representativeness, in particular regarding sampling design, see Goedeme (2013); Zardo Trindade and Goedeme (2016).

To assess the validity of our aggregation, we compare the published aggregate of wages by Eurostat based on EU-SILC to our country aggregation of the individual-level data. Since Eurostat does not publish an aggregate series for *gross* wages from EU-SILC but only for *net* wages, we use *net* wages (*net employee cash or near cash income (PY010N)* in EU-SILC) for comparison. Our aggregated series aligns closely with the officially published time series across Europe (Appendix Figure B.1 and Figure B.2), although with two limitations. First, in 2009, several countries changed from survey to register data for documenting wages in EU-SILC, resulting in some differences prior to the adjustment, most notably in the year of change (2009). As a result, the alignment of the two series is substantially improved from 2010 onwards. Second, wage growth for Cyprus has an unreliable profile in *net* terms, although our series for *gross* wages in Cyprus is smoother (Figure B.2).[2] Excluding both the year 2009 and Cyprus, we obtain a correlation coefficient of 0.91 between our aggregated series and the officially published Eurostat data.

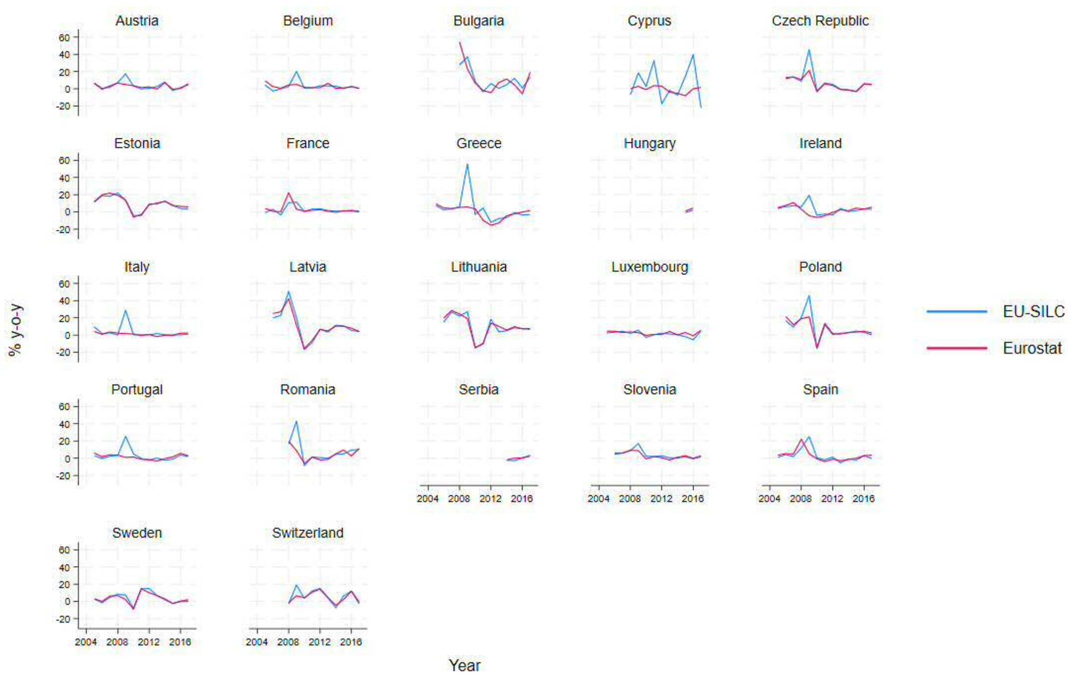
[2]Goedeme and Trindade (2020) indicate that Cyprus relies on surveys to collect income data but matches it with register data to correct for apparent mistakes and keeps extreme or outlying values in the data if they have been verified.

Figure B.1: Net wage growth EU-SILC aggregated vs published by Eurostat



Note: Wage growth refers to average annual change in nominal wages 2003-2017 and is computed as a weighted average for European countries in our sample with available data for net wages.
Source: Authors' computations based on EU-SILC and Eurostat.

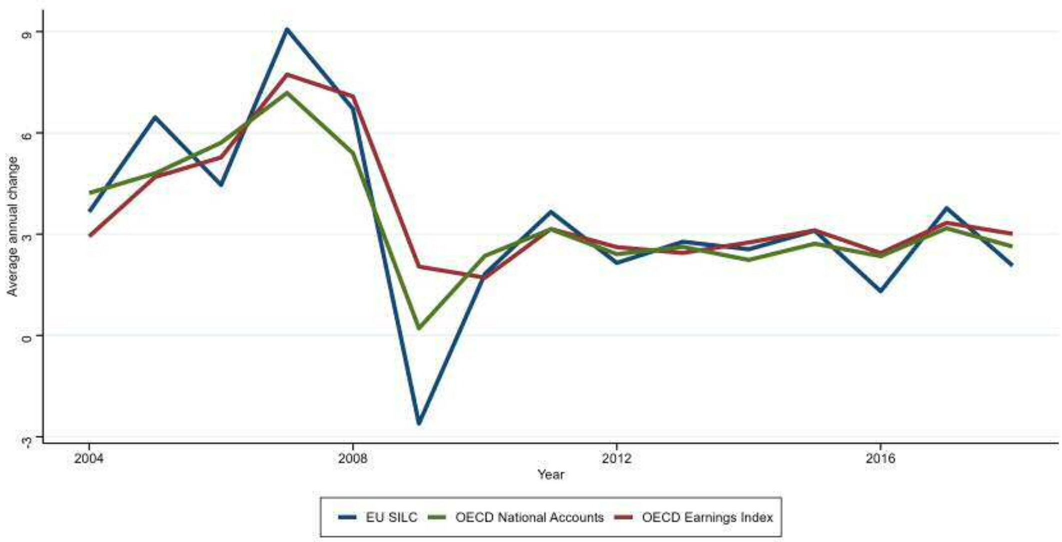
Figure B.2: Net wage growth EU-SILC aggregated vs published by Eurostat
for countries with available data for net wages



Note: Wage growth refers to average annual change in nominal wages. Results are shown for all countries with available data for net wages.
Source: Authors' computations based on EU-SILC.

Figure B.3 compares our baseline wage growth measure from EU-SILC (PY010G) to established wage data from national accounts as well as surveys and registries. While our measure aligns closely with both series for the 2010-2015 period, some differences occur in the earlier and later years. However, differences between wage measures of different sources are rather common: a correlation analysis reveals that EU-SILC data still aligns closer to each of the two OECD series than the two series align between each other. Our measure aligns closest with the national accounts measure for average annual wages per full-time equivalent dependent employee (CPNCU). It is computed by dividing the total wage bill by the number of employees multiplied by the ratio of average usual weekly hours per full-time employee to average usually weekly hours for all employees. This approach is rougher compared to ours since we compute hourly wage growth based on respective hours worked on an individual basis. Our measure also conforms the OECD Earnings Index (MEI) that aggregates wage developments (LCEAPR IXOBSA). Differences are likely because the MEI only includes private sector employees based on survey and administrative data.

Figure B.3: Wage growth EU SILC (gross) vs OECD National Accounts vs OECD Earnings Index

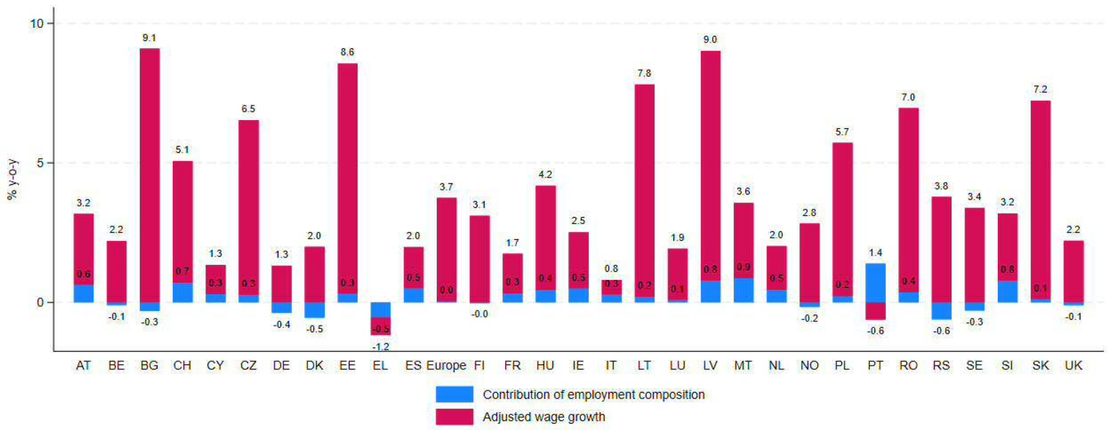


Note: Europe simple average includes all country-year observations with data available for both measures: AUT, BEL, DEU, DNK ESP, FIN, FRA, GBR, ISL, ITA, LTU, LUX, LVA, POL, PRT, SVK.
Source: EU-SILC and OECD.

C Appendix: Micro Adjustment

C.1 Adjusted vs unadjusted wage growth

Figure C.1: Wage growth adjusted with all controls for European countries



Note: Wage growth is adjusted for a changing employment composition by contract, gender, migration background, educational attainment, and work experience. Wage growth refers to the average annual change in nominal wages 2003-2017. Europe refers to the simple average of all countries shown. Periods may be limited depending on countries' data availability.

Source: Authors' computations based on EU-SILC.

D Appendix: Robustness

In Table D.1 we check the sensitivity of the competition effect when countries are excluded one at a time from the sample. The estimation is based on the model presented in column (1) of Table 1. In Table D.2 we report group-specific differences in the competition effect with respect to countries' institutional characteristics based on an alternative clustering procedure. The extensions are based on our baseline results presented in Table 2.

Table D.1: Sensitivity of labor market dualization to country exclusion

Country	Invol. Temp.	T-stat.	Country	Invol. Temp.	T-stat.
AT	-0.96***	-2.96	IT	-0.98***	-2.95
BE	-0.97***	-2.84	LT	-1.06***	-3.30
BG	-1.00***	-3.11	LU	-1.00***	-3.10
CH	-0.96***	-3.01	LV	-0.87**	-2.70
CY	-1.04***	-3.09	MT	-0.97***	-2.92
CZ	-0.94***	-2.81	NL	-0.97***	-2.95
DE	-0.91***	-2.80	NO	-1.01***	-3.07
DK	-0.94***	-2.87	PL	-0.81**	-2.26
EE	-1.00***	-3.00	PT	-0.92***	-2.78
EL	-0.86**	-2.43	RO	-0.83**	-2.67
ES	-0.93**	-2.52	RS	-0.91**	-2.55
FI	-0.90**	-2.75	SE	-0.98***	-2.85
FR	-0.94***	-2.85	SI	-0.94***	-2.88
HU	-1.08***	-3.31	SK	-0.81**	-2.62
IE	-1.05***	-3.09	UK	-0.95***	-2.89

Note: Two-tailed significance levels: $*$: 10% $**$: 5% $***$: 1%. T-statistics are based on cluster-robust standard errors by country. Dependent variable: adjusted wage growth, i.e., counterfactual overall aggregate wage growth assuming a constant share of temporary employees in total employees over time (base year: 2004).

Table D.2: The competition effect and the role of institutions

Dep. var.: wage growth perm. workers	Inst: TUD (1)	Inst: CBC (2)	Inst: Coord. (3)	Inst: EPL (4)	Inst: PCA (5)	Inst: Metten I. (6)
$Prod_t$	0.21	0.32*	0.29*	0.01	0.25	0.40**
	(1.23)	(1.79)	(1.84)	(0.06)	(1.43)	(2.53)
$Infl_t$	0.62	0.47	0.50	-0.15	0.44	0.46
	(1.61)	(1.03)	(1.15)	(-0.42)	(0.99)	(1.00)
U_t	-0.57**	-0.57***	-0.58***	-0.35**	-0.64***	-0.51**
	(-2.43)	(-2.83)	(-3.09)	(-2.42)	(-3.07)	(-2.71)
$Invol.\,Temp_t$						
...low Inst	-1.28***	-1.21**	-1.50***	-0.30	-1.42***	-1.72***
	(-3.27)	(-2.30)	(-3.57)	(-0.84)	(-3.04)	(-5.10)
...med. Inst	-1.12	-0.50	-0.80	-0.75**	-1.51**	-0.60
	(-1.59)	(-1.11)	(-1.41)	(-2.12)	(-2.12)	(-1.42)
...high Inst	0.05	-0.57	-0.22	-2.07***	-0.04	-0.13
	(0.15)	(-1.09)	(-0.76)	(-3.73)	(-0.17)	(-0.38)
TUD_t	0.40***					
	(3.36)					
CBC_t		-0.08*				
		(-1.89)				
$Coord_t$			-1.26*			
			(-1.90)			
EPL_t				-1.10		
				(-0.92)		
PCA_t					-2.26*	
					(-2.00)	
Cons	-1.41	15.07***	14.34***	14.62***	11.75***	10.24***
	(-0.32)	(4.23)	(3.46)	(4.33)	(5.79)	(3.58)
Model	FE	FE	FE	FE	FE	FE
TimeD	incl.	incl.	incl.	incl.	incl.	incl.
N	300	302	344	278	302	340

Note: Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis and are based on cluster-robust standard errors by country. The grouping into low, medium and high countries corresponds to the first, second and third tercile of the respective institutional variable. PCA is an index generated by principal component analysis based on TUD, CBC, and Coord (first component). Note, that for running the PCA analysis we imputed some missing observations in TUD and CBC to increase sample size. Metten I. is a composite index for union strength developed by Metten (2021).

E Appendix: Variable measurement and sources

Table E.1: Variable measurement and sources

Abbrev.	Variable	Measurement	Source
Wage growth	Annual average change in nominal wages	Gross employee cash or near cash income (PY010G)/months employed (PL073+PL074)/hours worked (PL060) (separated for full time and part time) aggregated with personal cross-sectional weights (PB040)	EU-SILC
Prod.	labor productivity growth	Nominal GDP/employment*100, annual change	Eurostat (naida_10_pe, naida_10_gdp)
Infl.	HICP Inflation	Annual average change of HICP	Eurostat (prc_hicp_aind)
Exp. Infl.	Inflation expectations	Monthly consumer survey asking for price trends over the next months, yearly average over 12 months	European Commission
U	Unemployment rate (U-3)	Unemployed (ILO definition) in % of active working age population (aged 15-74)	Eurostat (lfsa_urgan)
Invol. Temp	Involuntary temporary employment	Employees with a temporary contract who could not find a permanent job, in % of active working age population (aged 15-74)	Eurostat (lfsa_etgar, lfsa_eegais, lfsa_agan)
Temp	Temporary employment	Employees with a temporary contract, in % of active working age population (aged 15-74)	Eurostat (lfsa_etgadc, lfsa_agan)
Vol. Temp	Voluntary temporary employment	Temp − Invol. Temp	
U-5	U-5 Unemployment rate	Unemployed incl. discouraged (not seeking, but available) and marginally attached workers (available, but not seeking)	Eurostat (lfsa_urgan, lfsa_sup_age)
NAWRU	Non-accelarating wage rate of unemployment	Estimates from a model-based approach, European Union, 2017.	European Commission
Invol. Part	Involuntary part-time employment	Share of involuntary part-time employees in labor force, in %	OECD Statistics
EPL	Employment protection legislation	Strictness of employment protection – individual and collective dismissals (regular contracts)	OECD Statistics
TUD	Trade union density	Union members in % of employees (administrative and survey data)	OECD, ICTWSS
CBC	Collective bargaining coverage	Percentage of employees with the right to bargain	OECD Statistics
Coord	Coordination of wage setting	degree of coordination in wage bargaining on an ordinal 5-point scale	OECD Statistics

Table E.2: Country grouping according to strength of labor market institutions

	TUD			CBC			Coord			EPL			PCA		Metten I.	
Country	Mean	Base	Robust	Mean	Base	Robust	Mean	Base	Robust	Mean	Base	Robust	Mean	Robust	Mean	Robust
AT	29	2	2	98	3	3	4.0	3	3	2.3	1	2	1.3	3	8.4	3
BE	54	3	3	96	3	3	5.0	3	3	1.8	1	1	2.3	3	9.1	3
BG	18	1	1	.	.	.	1.4	1	1	.	.	.	.	.	5.7	1
CH	17	1	1	52	2	2	3.0	2	3	1.4	1	1	-0.4	2	4.9	1
CY	54	3	3	.	.	.	2.3	2	2	.	.	.	.	.	6.6	2
CZ	15	1	1	30	1	1	1.0	1	1	3.4	2	3	-1.8	1	5.7	1
DE	19	1	2	60	2	2	4.0	3	3	2.6	2	3	0.2	2	7.6	3
DK	68	3	3	79	2	2	4.0	3	3	1.5	1	1	1.9	3	7.3	2
EE	7	1	1	26	1	1	1.8	1	1	2.0	1	1	-1.8	1	5.3	1
EL	22	2	2	67	2	2	1.8	1	1	2.8	2	3	-0.6	2	7.6	2
ES	16	1	1	86	3	3	2.9	2	2	2.2	1	2	0.2	2	7.5	2
FI	68	3	3	87	3	3	3.6	3	3	2.1	1	1	1.9	3	8.8	3
FR	9	1	1	98	3	3	2.0	1	2	2.6	2	2	-0.2	2	7.9	3
HU	11	1	1	25	1	1	1.0	1	1	1.8	1	1	-2.0	1	6.8	2
IE	29	2	2	38	2	1	2.6	2	2	1.2	1	1	-0.6	1	6.4	2
IT	35	2	3	80	2	3	3.2	3	3	2.9	2	3	0.7	3	8.9	3
LT	9	1	1	10	1	1	1.0	1	1	2.5	2	2	-2.4	1	5.4	1
LU	36	2	3	57	2	2	2.2	2	2	2.1	1	2	-0.2	2	7.9	3
LV	14	1	1	17	1	1	1.0	1	1	3.0	2	3	-2.1	1	4.9	1
MT	55	3	3	.	.	.	1.0	1	1	.	.	.	.	.	.	.
NL	19	1	2	81	2	3	4.1	3	3	3.3	2	3	0.7	3	7.3	2
NO	50	3	3	73	2	2	4.0	3	3	2.3	2	2	1.3	3	8.2	3
PL	18	1	2	19	1	1	1.0	1	1	2.3	1	2	-2.0	1	5.6	1
PT	19	1	2	78	2	2	2.0	1	2	3.8	2	3	-0.3	2	7.3	2
RO	32	2	2	.	.	.	2.5	2	2	.	.	.	.	.	6.6	2
RS	.	.	.	.	.	.	1.0	1	1	1.7	1	1	.	.	5.9	1
SE	69	3	3	90	3	3	4.0	3	3	2.5	2	2	2.1	3	9.0	3
SI	29	2	2	75	2	2	2.7	2	2	2.4	2	2	0.2	2	7.3	2
SK	16	1	1	34	1	1	1.8	1	1	2.7	2	3	-1.4	1	5.9	1
UK	26	2	2	30	1	1	2.0	1	2	1.4	1	1	-1.1	1	5.5	1

Note: Countries are classified either into two (1 - low, 2 - high) or three groups (1 - low, 2 - medium, 3 - high). Base and Robust refer to the groupings considered in the models of Tables 2 and D.2 respectively.

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