Bulligan and Viviano *IZA Journal of Labor Policy* (2017) 6:9
DOI 10.1186/s40173-017-0087-z

IZA Journal of Labor Policy

ORIGINAL ARTICLE Open Access

Has the wage Phillips curve changed in the euro area?

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Abstract

Increasing evidence shows that in the aftermath of the global financial crisis, in the euro area, the relationship between price inflation and economic slack became stronger. Instead, there is no clear evidence of a strong(er) relationship between wage inflation and unemployment. In this paper, we estimate a Phillips curve with time-varying coefficients separately for Italy, Spain, Germany and France and we find that, with the exception of Germany, after the global financial crisis, the sensitivity of hourly wage changes to labour market slack increased. Second, by the use of administrative microdata, available only for Italy, we relate daily wage changes to the local unemployment rate. The results confirm the steepening of the Phillips curve after 2008, also when controlling for composition effects.

Keywords: Wage growth, Phillips curve, Parameter instability

JEL Classification: E24, E31, E58

1 Introduction

After the global financial crisis, the debate about the short-run determinants of inflation has gained momentum. In the USA, where empirical evidence supports (if any) the hypothesis of a flattening of the Phillips curve (PC; see Blanchard 2016 for a critique[1]), the debate started immediately after the global financial crisis, with the so-called missing disinflation puzzle (see, e.g. IMF 2013) and continued after the subsequent recovery (the "missing inflation puzzle"). Various explanations have been proposed. According to some studies, the flattening of the PC is caused by firms' and households' expectations (e.g. Coibion and Gorodnichenko 2015; Davig 2016, who proposes a model which explains PC parameter instability). For others, it depends on downward nominal wage rigidities (e.g. Daly and Hobijn 2014), which in Europe are mainly determined by institutional factors (typically wage bargaining institutions; see Izquierdo et al. 2017). In the euro area, the empirical evidence is rather mixed. On the one hand, some papers show that in recent years, the negative relationship between price inflation and economic slack has become stronger (see e.g. Giannone et al. 2014; Riggi and Venditti 2015). Yet the twin puzzles observed in the USA occurred also in the Euro area (see Ciccarelli and Osbat 2017). Furthermore it is not clear whether also the correlation between nominal wage growth and unemployment, indeed quite weak before the crisis, has changed too. Actually, non-linearities in the price inflation

Phillips curve do not necessarily imply non-linearities also in the wage Phillips curve, as other factors (for instance, the frequency of price adjustments, see e.g. Fabiani and Porqueddu 2017) may affect prices when slack increases. It is then important to complement the recent findings of the literature about price inflation with an analysis of the relationship between wage inflation and economic slack.

In this paper, we present both macro- and micro evidence supporting the hypothesis of a steepening of the wage Phillips curve in the largest euro area countries after 2008. First, we focus on Germany, France, Italy and Spain, and, by estimating country-specific Phillips curves, we test the hypothesis of parameter instability. Overall, we find support for the hypothesis that the slope of the PC changed after 2008. Second, we carry out time-varying regressions for the same set of countries and two standard PC specifications. We use the kernel-based non-parametric estimators proposed by Giraitis et al. (2013), based on the simple idea that in the presence of structural breaks, the estimator of the parameter of interest (the slope of the Phillips curve in our case) should put lower weight to older observations than to more recent ones. We find evidence of a steepening of the PC with respect to unemployment after 2008 in France, Italy and Spain, but not in Germany, where the PC is still steeper than in the other euro area countries but became flatter after the global financial crisis.

We then focus on Italian microdata. First, by using simple OLS models with firm fixed effects, we investigate the potential impact on the Phillips curve of changes in firm composition. In a setting which controls for firms' entry and exit, as well as firm unobservable characteristics, we estimate how average wage changes react to the local unemployment rate. We find evidence of an increase in wage flexibility at the firm level after 2008 and no evidence that changes in the composition of firms explain the steepening of the PC.

Last, we focus on workers. Our data do not allow us to separately analyse the dynamics of base and variable wage components, the last being one of the most important sources of wage flexibility in Europe (see Babecký et al. 2009). So we focus on workers' composition effects. We estimate the correlation between individual wage changes and the local unemployment rate in different samples of incumbent workers and fixed-term workers. In all specifications, we find a steepening of the PC. We then conclude that the higher elasticity of wages to labour market slack found in aggregate data after 2008 is not (only) driven by composition effects.

Analyses on the wage PC in the euro area are rather scant because it is generally believed that a wage PC in the euro area is rather weak and that downward nominal wage rigidities do not allow wages to react to slack.[2] All in all, our paper contributes to the literature on the wage PC showing that in some of the largest euro area countries, the relationship holds also for wages and it is likely to be non-linear.

In the literature there is no consensus on the possible causes of non-linearities, so that it is not clear whether the Phillips curve (both the price and the wage version of the PC) is steeper in a boom or in a recession.[3] Indeed, Anderton and Bonthuis (2015) find evidence of a flattening of the PC, using a panel of euro area countries. Their specification however is rather rigid as the models used do not allow to take into account different country-specific trends. On the opposite side, it is possible that in deep and prolonged recessions, as the one started in 2008, wages become more reactive to slack, as workers and unions are more willing to accept a wage cut in order to preserve employment.

Among structural changes, it is worth mentioning that the steepening/flattening of the PC can also be caused by changes in the composition of workers, due to hiring and firing decisions of firms. For instance, a steepening of the PC may occur if workers with more flexible wages are more likely to remain employed in recessions. On the opposite side, a flattening occurs if during a recession, firms prefer to hoard more productive (and highly paid) labour. It also follows that reforms increasing the flexibility on the use of labour can affect the speed of workforce re-composition and consequently the shape of the PC. On the same ground, also, firms' entry and exit can affect aggregate wage dynamics. This may happen for instance if more productive firms are also more likely to remain in the market during downturns. Conversely, it could happen that market forces tend to select firms with a more flexible wage structure.

Adamopoulou et al. (2016a), using microdata for Italy, find that after 2008, Italian firms used fixed-term workers to slow down the dynamics of wages. Adamopoulou et al. (2016b) show also that in Italy a non-negligible share of the evolution of wages can be attributed to changes in the composition of firms. The authors, however, do not provide estimates of the PC. Font et al. (2015) show that in Spain, real wages of newly hired workers and temporary workers are more sensitive to the business cycle, suggesting indirectly that changes in workers' composition can affect the dynamics of aggregate wages and its reaction to unemployment. With respect to these papers, our contribution is not only to use very flexible tests and techniques for the case of a time-varying PC but also to show that composition effects may certainly help to explain part the steepening of the PC. Composition effects, however, are just part of the possible explanation as some non-linearities still survive when changes in composition are accounted for.

This paper is organized as follows. In Section 2, we present evidence, based on aggregate data, confirming that in the euro area, in spite of existing institutional rigidities, wage growth negatively correlates with labour market slack. This is a preliminary step for formally testing the hypothesis of changes in the slope of the PC in the remaining part of the section. In Section 3, we present micro estimates for Italy, which allow us to control for a very wide sources of composition effects. Section 4 briefly concludes.

2 Macro evidence

2.1 Has the relationship changed? Preliminary evidence

We first check for the existence of a negative correlation between wage growth and unemployment. We start the analysis by estimating a reduced form statistical relationship between the unemployment rate and wage inflation. Figure 1 shows the scatterplots of the unemployment rate in the euro area and annual nominal wage growth between 1999Q1 and 2015Q4. The corresponding OLS estimates confirm that the coefficient on the unemployment rate is equal to −0.25 and statistically significant.

We then come back to aggregate macro variables and focus on the following basic model:

$$\Delta^4 w_t = c + \beta \Delta^4 w_{t-1} + \gamma U_t + \varepsilon_t \quad (1)$$

where $\Delta^4 w_t$ is the y-o-y nominal hourly wage growth in the private sector and U_t, as in Fig. 1, is the economy-wide unemployment rate. The lagged dependent variable $\Delta^4 w_{t-1}$ is aimed at capturing persistence in wage dynamics, which for instance can

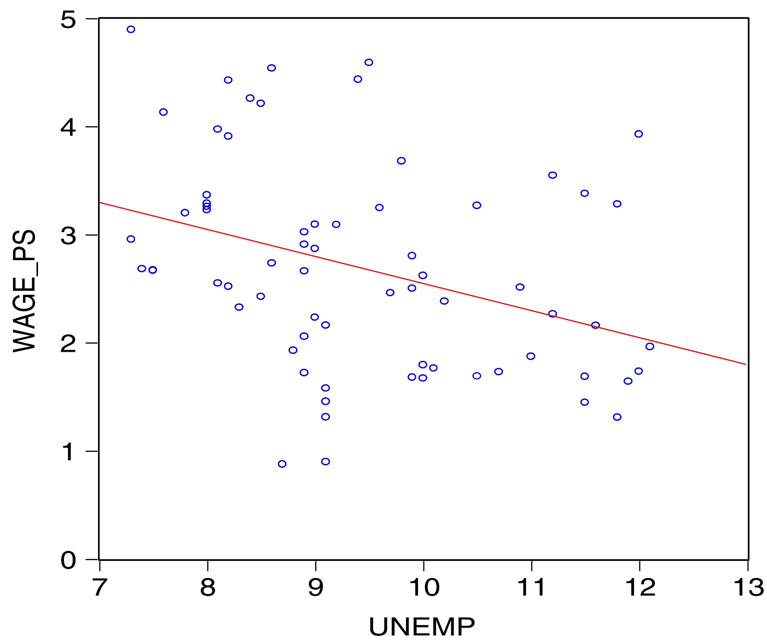

Fig. 1 Euro area, 19 countries. Nominal wage growth and unemployment

arise in case of wage rigidities (see also Schmieder and von Wachter 2010, for micro evidence). The unemployment rate proxies for labour market slack.

We estimate Eq. 1 recursively on an expanding window where the first sample is 1999Q1–2007Q4 which goes from the start of stage three of European Monetary Union to the global financial crisis. Then, we re-estimate the same equation adding one observation at a time. The last sample covers the period 1999Q1–2015Q4. This sub-period allows us to further neutralize possible remaining differences due to monetary policy before 1999. Figures 2 and 3 report the estimated coefficient on the unemployment rate for the euro area and for its four biggest economies. The figures suggest the following:

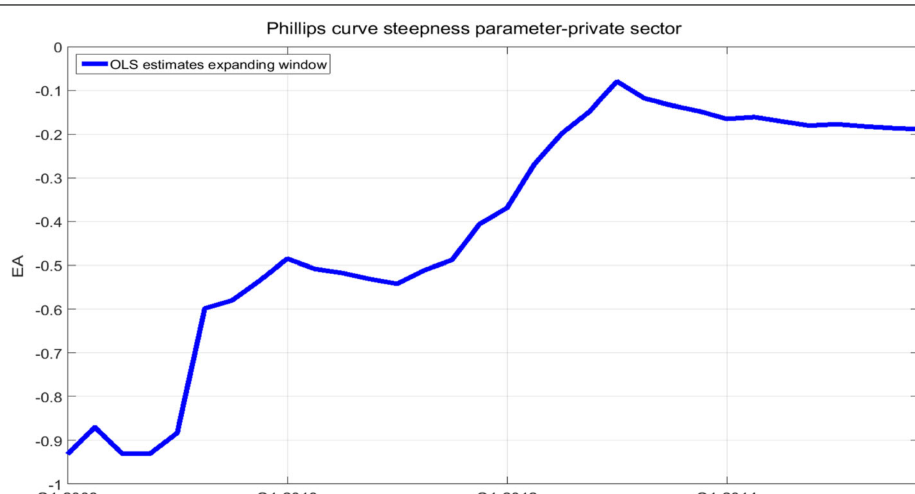

Fig. 2 Phillips curve expanding window regressions in the euro area. The *blue lines* show the OLS estimates of the parameter of unemployment rate for different samples: the first point refers to the sample 1999Q1–2008Q1 and the last point refers to the sample 1999Q1–2015Q4

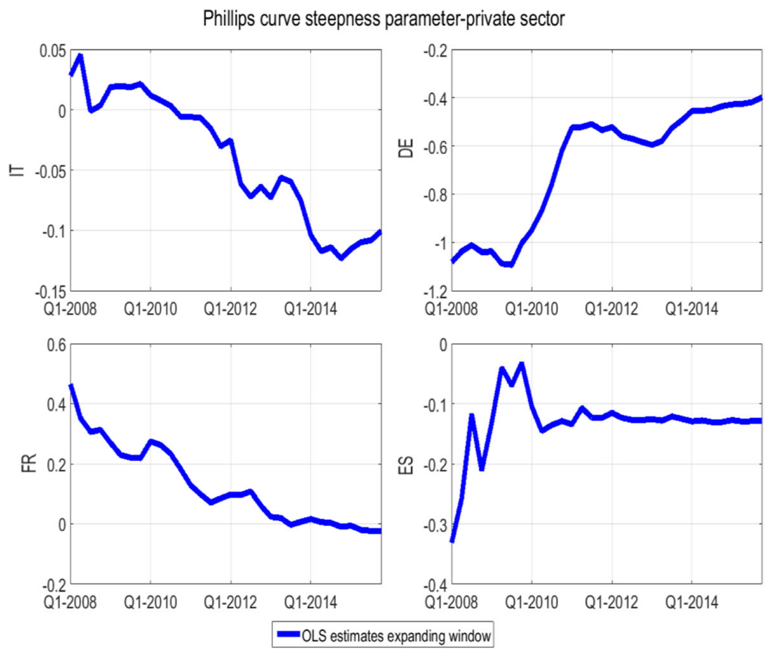

Fig. 3 Phillips curve expanding window regressions in the euro area countries. The *blue lines* show the OLS estimates of the parameter of unemployment rate for different samples: the first point refers to the sample 1999Q1–2008Q1 and the last point refers to the sample 1999Q1–2015Q4

(i) the correlation between unemployment and wage inflation in the euro area is rather unstable over time. While remaining negative, its absolute value has decreased monotonically from the beginning of 2008 to 2012 but it has increased again—albeit marginally—after 2012; (ii) its pattern over time reflects closely the pattern observed in Germany until 2012 (Fig. 3); (iii) the euro area aggregates mask heterogeneous results across its four biggest economies. The relationship between wages and labour market slack seems to have become more negative in Italy and less negative in Germany (but still higher than in the other countries). In Spain, after a maximum around 2010, it has become more negative again, while in France has moved from being positively signed to being negligible. While Eq. 1 represents a very naïve descriptive model, it provides prima facie evidence in favour of parameter instabilities in the four major euro area countries.

Last, it is important to mention that the choice of the unemployment rate as a measure of slack, instead of the unemployment gap is intended to avoid that our estimates are affected by mis-measurement of the unemployment gap. From a theoretical point of view, the use of the unemployment rate only implies that the natural rate of unemployment is assumed to be constant (our estimates include a constant term). This hypothesis can be particularly restrictive for the period that we analyse, as many euro area countries carried out labour market reforms to increase labour market flexibility. In Section 2.4, however, we directly deal with the potential impact of this assumption.

2.2 Testing for the instability of the Phillips curve

In this section, we adopt a time series approach to evaluate the case of time-varying sensitivity of wages to labour market conditions. We use country-by-country time series, and we check whether there is evidence supporting the case for parameter instability after controlling for other determinants of wage inflation like past and expected inflation.

First, we estimate a specification where wage inflation at time t is a function of its own lags, the current or past unemployment rate and expected inflation, measured by the average inflation rate expected 6-quarter ahead by a sample of financial analysts surveyed by Consensus Economics,

$$\Delta^4 w_t = c + \sum_{i=1}^{k} \beta_i \Delta^4 w_{t-i} + \delta E_t \Delta p_{t+1} + \gamma U_{t-h} + \varepsilon_t \tag{2}$$

where lags between 0 and 4 are considered for the unemployment rate. As Eq. (1), Eq. 2 is consistent with theoretical models where wage inflation has some degree of persistence (staggered-contract models), like the one proposed by Galì (2011). On the one hand, Galì (2011) derives a wage Phillips within the New Keynesian framework. On the other, he includes lagged terms of wage inflation to capture the serial correlation found in the data. This formulation also allows for wage inflation to respond to current and expected business cycle conditions.

The first type of evidence is based on tests of parameter instability over the period 1999Q1–2015Q4. Under the null hypothesis, the parameters of the PC are stable, while under the alternative hypothesis, they evolve as driftless random walks. As conventional break-point tests show low power when the break occurs towards the end of the sample, we resort to the tests suggested by Busetti (2012). According to the author, the locally most powerful test can be modified so as to achieve higher power.

The locally most powerful (LMP) test has the following form:

$$L = \hat{\sigma}^{-2} T^{-2} \sum_{t=1}^{T} S_t' V^{-1} S_t \tag{3}$$

where $\hat{\sigma}^2 = (T-K)^{-1} \sum_{t=1}^{T} \widehat{u_t^2}$, $S_t = \sum_{j=t}^{t} \widehat{u_j} \widehat{x_{1j}}$, $V = T^{-1} \sum_{t=1}^{T} x_{1t} x_{1t}'$ and $\hat{u}_t = y_t - x_t' \hat{\beta}$ are the OLS residuals from regressing y_t on a set of x_t variables for a subset of which we are interested in testing parameter instability (x_{1t} which might include all x_t variables). Busetti (2012) suggests to use the following variations to focus on breaks that occur only in the last fraction π of the sample $L(\pi) = \hat{\sigma}^{-2} (T-\pi T)^{-2} \sum_{t=\pi T+1}^{T} S_t' V^{-1} S_t$ and to test parameter instability with:

$$\text{Sup–}L = \text{Sup}_{\pi \in \Pi} L(\pi) \tag{4}$$

$$\text{Exp–}L = \log \int_{\pi \in \Pi} \exp(L(\pi)) \pi d(\pi) \tag{5}$$

The main idea behind these tests is to increase the power of original LMP tests by focusing on the latest part of the sample and additionally by giving increasing weight to observations close to the end of the sample (Exp-L). We test for a break in the last 25%, the last 10% of the sample and the central interval covering 98% of the sample.

The tests are run for five specifications of Eq. 2 where we let the parameter h, which controls the lag relationship between wage dynamics and the unemployment rate, range from 0 to 4 (for robustness) and keep k, the parameter controlling the number of lags of wage growth, fixed to 1. Table 1 collects the average (across values of h) results for tests conducted on all coefficients (panel a) and on the coefficient on the

Table 1 End of sample instability test (Eq. 2)

	Italy	France	Germany	Spain
	All parameters			
Sup-L(.75)	1	0	1	.6
Sup-L(.90)	1	0	1	.6
Exp-L(.75)	1	0	1	.4
Exp-L(.90)	1	0	1	.6
Sup-L(.01–.99)	1	0	1	.6
Exp-L(.01–.99)	1	0	1	.6
	Single parameter			
Sup-L(.75)	1	0	1	.6
Sup-L(.90)	1	0	1	.6
Exp-L(.75)	1	0	1	.6
Exp-L(.90)	1	0	1	.6
Sup-L(.01–.99)	1	0	1	.6
Exp-L(.01–.99)	1	.2	1	.6

Values reported are the average across $h = 0,\ldots,4$ where the result of the test is coded with 1 if the test rejects at the 10% the null hypothesis of constant parameters in favour of time-varying parameters, with 0 if the null hypothesis of constant parameters cannot be rejected at the 10%

unemployment rate alone (panel b). Since values close to 1 are supportive of parameter instability (see note to the table), we find clear evidence of parameter instability for Italy and Germany, mixed evidence for Spain (which is the results of stronger evidence when lags of the unemployment rate are used) and no evidence for France.

We run the same tests also on a slightly different specification of the PC where we substitute past wage dynamics for past inflation dynamics. This formulation is still compatible with the Galì's model (2011), under the reasonable assumption that in euro area countries, wages are to some extent indexed to past inflation.

$$\Delta^4 w_t = c + \sum_{i=1}^{k} \beta_i \Delta^4 p_{t-i} + \delta E_t \Delta p_{t+1} + \gamma U_{t-h} + \varepsilon_t \tag{6}$$

Also, under this specification, there is clear evidence in favour of parameter instability for Italy and Germany and mixed evidence for Spain. However, there is also evidence in favour of instability for France (see Table 2).

2.3 Time-varying coefficients

The second type of evidence is based on time-varying parameter regressions. More specifically, we explicitly allow the parameters in Eqs. 2 and 6 to vary over time (before doing that we set $h = 0$ and $k = 1$):

$$\Delta^4 w_t = c_t + \beta_t \Delta^4 w_{t-1} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t \tag{2bis}$$

or alternatively,

$$\Delta^4 w_t = c_t + \beta_t \Delta^4 p_{t-1} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t \tag{6bis}$$

Following Riggi and Venditti (2015), we resort to the kernel-based non-parametric estimators proposed by Giraitis et al. (2013). This type of estimators is based on the idea

Table 2 End of sample instability test (Eq. 6)

	Italy	France	Germany	Spain
	All parameters			
Sup-L(.75)	1	1	1	.4
Sup-L(.90)	1	1	1	.4
Exp-L(.75)	1	1	1	.4
Exp-L(.90)	1	1	1	.4
Sup-L(.01–.99)	1	1	1	.4
Exp-L(.01–.99)	1	1	1	.4
	Single parameter			
Sup-L(.75)	1	1	1	.4
Sup-L(.90)	1	1	1	.4
Exp-L(.75)	1	1	1	.4
Exp-L(.90)	1	1	1	.4
Sup-L(.01–.99)	1	1	1	.4
Exp-L(.01–.99)	1	1	1	.4

Values reported are the average across $h = 0,\ldots,4$ where the result of the test is coded with 1 if the test rejects at the 10% the null hypothesis of constant parameters in favour of time-varying parameters, with 0 if the null hypothesis of constant parameters cannot be rejected at the 10%

that in the presence of structural breaks, older observations should be discounted more than more recent ones. Giraitis et al. (2013) find that such methods perform well in forecasting several US macroeconomic series. In a subsequent paper, Giraitis et al. (2014) show that they also are consistent and asymptotically normally distributed and, most important for our purposes, have very good small sample properties. As suggested by the authors, we choose a Gaussian kernel and set the discounting parameter to $H = T^r$ where $r = 0.5$. However, unlike the authors, we use a one-sided kernel which at each time t only considers current and past observations. It follows that the estimated parameter instability between consecutive periods t and $t + 1$ is due both to a new data point entering the estimation sample as well as changes in the weighting structure. Figure 4a, b shows the time-varying 50th, 16th and 84th percentiles of the distribution for the parameter of the unemployment rate, the so-called steepness of the PC for the two specifications tested.

Both figures confirm that the responsiveness of wage dynamics to the unemployment rate has changed since the pre-crisis period. In particular for Germany, the data support the existence of a Phillips curve and suggest a flattening of the slope. On the contrary for Spain and Italy, even if in these countries before the crisis, the Phillips curve is not supported by the data (the coefficient on the unemployment rate is either zero or positive[4]), after the crisis, we find a negative relationship. Finally for France, we find mixed results: the data are not consistent with the implication of the Phillips curve if specified as in Eq. 1 (the coefficient on the unemployment rate is positive throughout the whole sample), but they are more in line with a Phillips curve after 2011 according to the alternative specification.

According to the (median) results for the first specification (Eq. 2bis), at the end of 2015, an increase by one percentage point (pp) in the Italian unemployment rate is associated with a decrease by 0.15 pp in annual wage inflation on impact and with a

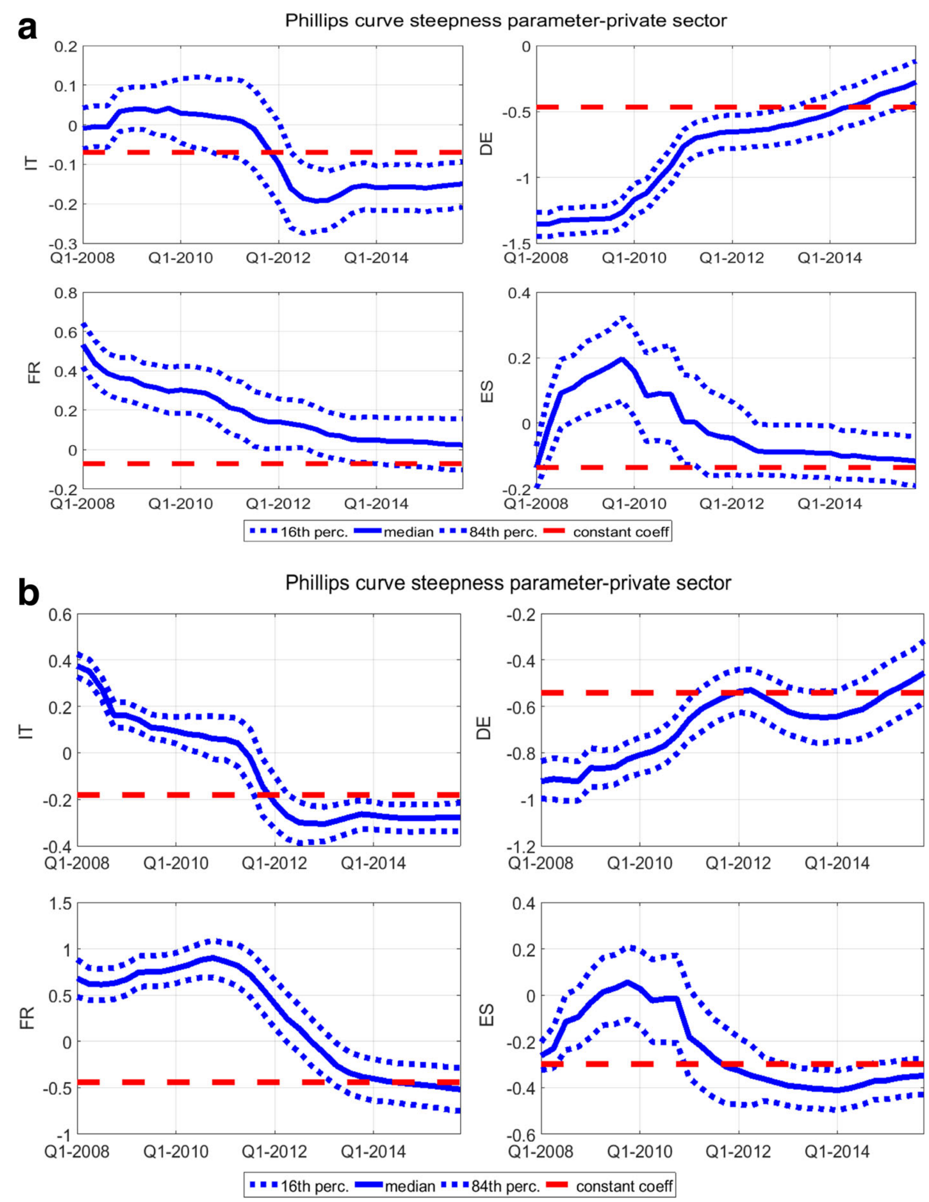

Fig. 4 Country-specific evidence of instability of the unemployment parameter in wage Phillips curve. **a** $\Delta^4 w_t = c_t + \sum_{i=1}^{h} \beta_{it} \Delta^4 w_{t-i} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t$. **b** $\Delta^4 w_t = c_t + \sum_{i=1}^{h} \beta_{it} \Delta^4 p_{t-i} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t$. The *blue lines* are 16th, 50th and 84th percentile from time-varying parameter model. The *red line* is OLS estimate from constant parameter model estimated over period 1999Q1–2015Q4

decline by 0.3 pp, after accounting for the lag structure. Very similar results are found for Spain. In Germany, despite the recent flattening, wage inflation changes are slightly stronger (0.4 and 0.6 pp respectively), while in France, the short-run change of wage inflation is modest (0.02 pp) and not significant (it is close to −0.5 pp accounting for lags, although not significant). The alternative specification (Eq. 6bis) does not distinguish between short-run and long-run correlation, and as far as the latter are concerned, these are in line with the estimates obtained from the first specification.

2.4 Changes in sectoral composition and robustness

So far, the results refer to the whole private sector. However, it is interesting to look at sector-specific wage changes as results could differ substantially between labour vs capital intensive sectors or depend on the type of worker and contract prevailing in specific sectors. Moreover, results for the whole private sector could be driven by changes in the relative weight of the manufacturing sector, typically considered as more rigid and more unionized, and the service sector. So, carrying out different estimates for each sector allows us to control for changes in the simplest source of composition effects, i.e. sectoral composition occurring after 2008. We re-run the analysis looking at hourly wages for industry and services separately (according to the NACE classification), while, absent sector-specific measures of labour market slack, we keep using the total economy unemployment rate. Again, results are shown for both specifications and reported in Fig. 5a, b.

For Italy, there is evidence of increased responsiveness of wage growth to the unemployment rate in both industry and services. For Spain, the increased steepness seems to reflect changes in the service sector, whereas in the industrial sector since the crisis, the responsiveness of wages has stopped declining. For Germany, the decline reflects both industry and service developments, while in France, a non-significant relationship between wage growth and the unemployment rate is confirmed by sector-specific results.

If we look at the second specification, the results are broadly similar to those obtained under the first specification with the exception of France where the evidence of increased responsiveness found for the whole private sector reflects developments in both industry and services.

So far, we have performed the analysis for given parameters regarding the amount of discounting of past data and the lag relationship between wage growth and the unemployment rate. In order to shed light on the role of the discounting factor in shaping results, Fig. 6 shows how the latter changes for different values of the parameter H (which in our setting controls the amount of discounting), while in Fig. 7, we show the results obtained under different lag structures. Regarding the discount factor, starting from the benchmark case ($r = 0.5$ and therefore $H = T^{0.5}$) as we reduce the amount of discounting applied to past data (r increases), the degree of parameter instability declines but does not vanish. Figure 7 shows instead the effect of changing the lag relationship between unemployment and wage dynamics. For Italy, France and Germany, the results do not seem to depend from the choice of this parameter, while for Spain, results are more sensitive to it.

Last, our results could be potentially affected by the assumption of a constant natural rate of unemployment. This hypothesis could be very restrictive for our analysis as structural reforms, as the one undertaken in many euro area countries after 2008, can have both reduced wage inflation and the natural rate on unemployment, causing the change in the slope that we find in the PC. To check also analytically that our results remain unaffected, we estimate the correlation between wage inflation and slack for the periods before and after 2008. We carry out separate exercises by using the unemployment rate or the unemployment gap (OECD estimates from 1995 to 2014). Our results are fully confirmed (details available upon request): the correlation between wage inflation and the unemployment gap is non-significant in Italy, Spain and France before

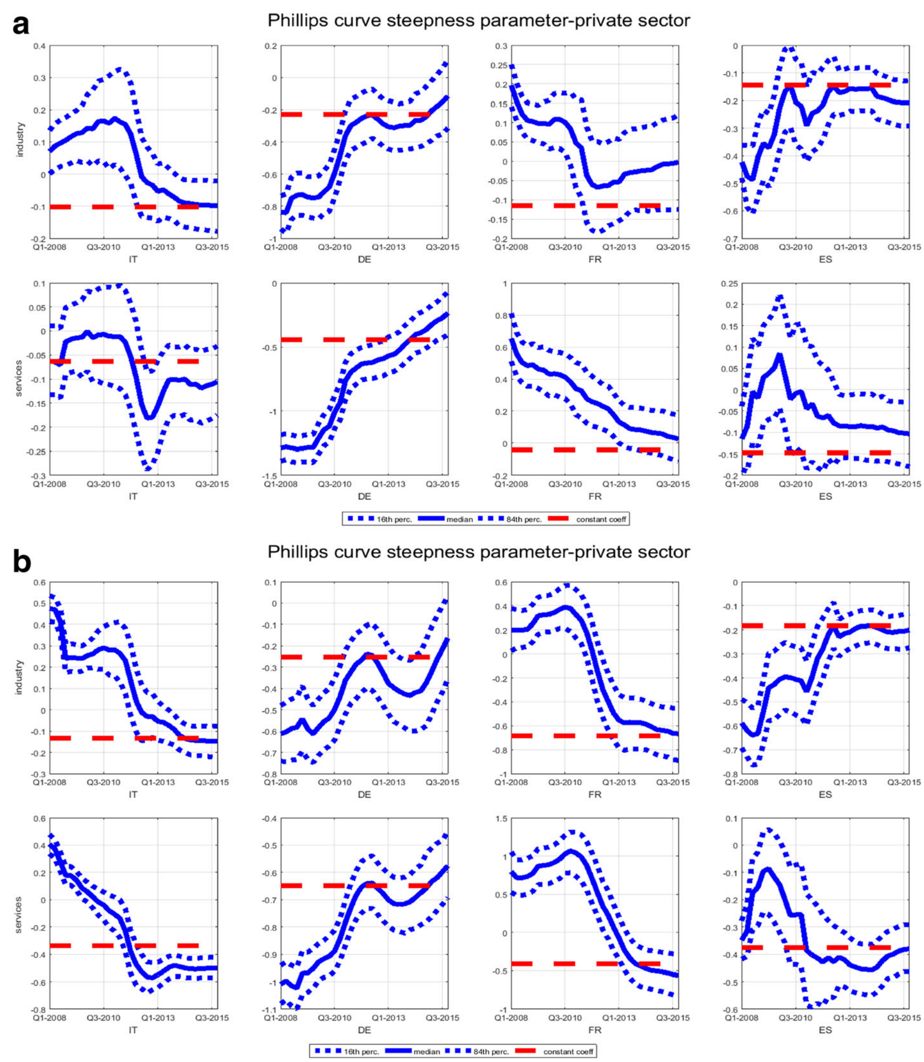

Fig. 5 Sector-specific evidence of instability of the unemployment parameter in wage Phillips curve. **a** $\Delta^4 w_t = c_t + \sum_{i=1}^{h}\beta_{it}\Delta^4 w_{t-i} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t$. **b** $\Delta^4 w_t = c_t + \sum_{i=1}^{h}\beta_{it}\Delta^4 p_{t-i} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t$. *Blue lines* are 16th, 50th and 84th percentile from time-varying parameter model. The *red line* is OLS estimate from constant parameter model estimated over period 1999Q1–2015Q4

2008, while it is negative and significant in Germany. Afterwards, it becomes negative and significant in Italy, France and Spain (as described in Section 2.3)[5]; it is still negative, but small and not significant in Germany.

3 What is (not) behind the steepening? Evidence from Italy

3.1 The data

To check for the robustness of our results, we rely on a single country, Italy, and on microdata on firms and workers. In particular, we focus on how wage growth responds to the local province-level unemployment rate.[6] Data on wages consist of social security payments made by all private-sector firms with at least one employee to the Italian National Social Security Institute (INPS). From this master data, INPS extracts two datasets. The first consists of the universe of firms, and for each firm, it includes the

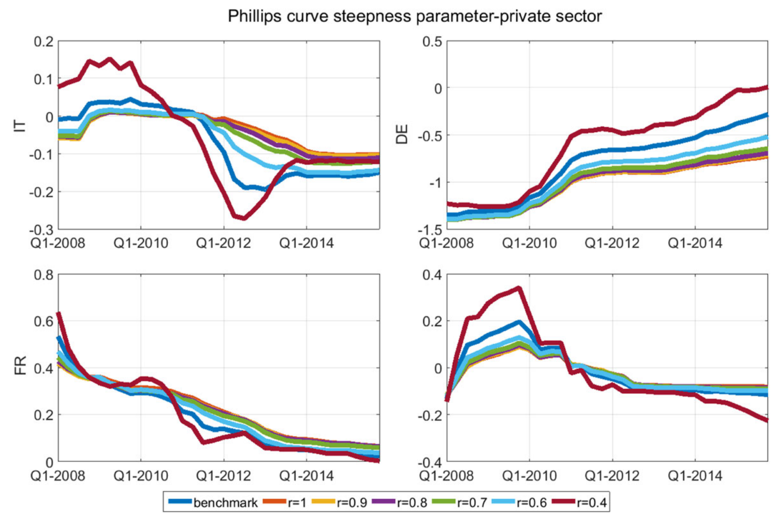

Fig. 6 Robustness to discounting factor $\Delta^4 w_t = c_t + \sum_{i=1}^{h} \beta_{it} \Delta^4 w_{t-i} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t$

average number of employees and firms' wage bill, together with a firm unique identifier and the province where the firm is located. It is then possible to calculate average wages. The last available data refer to year 2013. The second consists of the employment histories of all workers born on the 1st or the 9th day of each month (6.5% of total workforce in the Italian private sector). The worker extraction, updated to 2014, provides information on demographics, the annual gross wage, the number of days worked, the main characteristics of the contract and maternity and sick leaves, as well as whether the worker benefits from the Italian wage supplementation scheme, *Cassa integrazione guadagni* (CIG). The province-level unemployment rate is released by the Italian National Statistical Institute (Istat) since 2004.

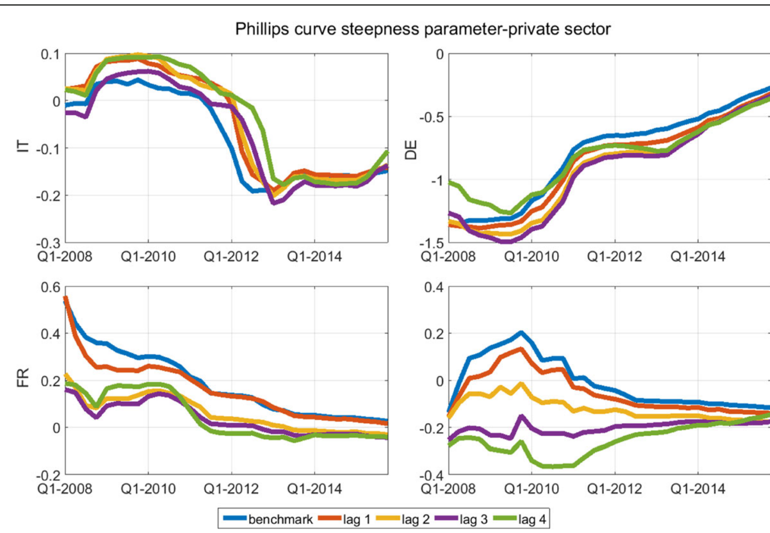

Fig. 7 Robustness to lag relationship between unemployment and wage dynamics $\Delta^4 w_t = c_t + \sum_{i=1}^{h} \beta_{it} \Delta^4 w_{t-i} + \delta_t E_t \Delta p_{t+1} + \gamma_t U_t + \varepsilon_t$

3.2 Firm composition

We first look at firms' adjustments, and we estimate the following equation using data from 2004 to 2013:

$$\Delta w_{tfp} = \delta_t + \delta_f + \beta \Delta w_{t-1,pf} + \delta_1 U_{t,p} + \delta_2\, U_{t,p} * D_{\text{post2008}} + \epsilon_{tfp} \tag{7}$$

where all data have an annual frequency; Δw_{tfp} is the y-o-y change in the average gross salary paid to employees by firm fth, located in province p; δ_t and δ_f are the time and firm fixed effects; and $U_{t,p}$ is the unemployment rate in province p at time t. Since the number of firms which change province each year is negligible, the term δ_f captures also province-level characteristics. The dummy D_{post2008} is equal to 1 if the observation refers to a year between 2008 to 2013 (0 otherwise), and δ_2 captures changes in the elasticity of the PC after 2008. Equation 7 allows us to strictly control for confounding factors affecting firms' labour cost evolution. The identification of the slope of the PC and its steepening is then based on the variability of the unemployment rate across provinces, whereas the impact of inflation expectations, which (reasonably) do not vary at the local level, is captured by time fixed effects.

The coefficients δ_1 and δ_2 should however partly capture also the reaction of wages to the aggregate unemployment rate. We then estimate the following equation:

$$\Delta w_{tfp} = \delta_f + \beta \Delta w_{t-1,pf} + \delta_1 U_{t,p} + \delta_2\, U_{t,p} * D_{\text{post2008}} + \delta_3 E_t \Delta p_{t+1} + \epsilon_{tfp} \tag{8}$$

where time dummies are skipped and aggregate shocks, which include also wage increases set by nation-wide collective agreements, are captured by inflation expectations $E_t \Delta p_{t+1}$. The coefficients δ_1 and δ_2 represent then the correlation between the unemployment rate and wage growth not captured by the correlation between wage growth and inflation expectations. Since both Eqs. 7 and 8 include the lagged value of the dependent variables, estimates can be biased (and since we are using panel data, this bias does not vanish as the time dimension of our dataset increases). Thus, we estimate Eq. 8 also by the use of a GMM estimator. The inclusion of the lagged dependent variable in 8 not only allows for a higher correspondence with the specifications used in Section 2 but also provides a further check that changes in the slope of the PC do not depend on the assumption underlying the different specifications.

Preliminarily, consider Fig. 8, which reports the y-o-y elasticity of wage growth to the unemployment rate, obtained by estimating Eq. 7 on the interactions among $U_{t,p}$ and time fixed effects.[7] Overall, we find a negative relationship between wage changes and unemployment, entirely determined by the increase of elasticity from 2009 onwards. (The positive spike in year 2008 has been probably due by the growth of the base wage component determined by collective pay agreements, mostly renewed in 2008, i.e. before the reform of collective bargaining undertaken by social partners in 2009.)

In Table 3, instead, we report the results obtained by estimating Eq. 7 (column 1) and Eq. 8 (columns 2 and 3, OLS and GMM estimators respectively).

Table 3 is also split into three parts. The upper one refers to the whole population of firms. In column 1, which is based on Eq. 7, the effect of province level differences in the unemployment rate on wage growth is rather small and significant only after 2008. In the second column, referring to Eq. 8 (OLS estimator), the correlation between the unemployment rate and wage growth is larger, being δ_1 equal to –.225 and highly significant. Interestingly after 2008, the correlation increases and the new slope of the

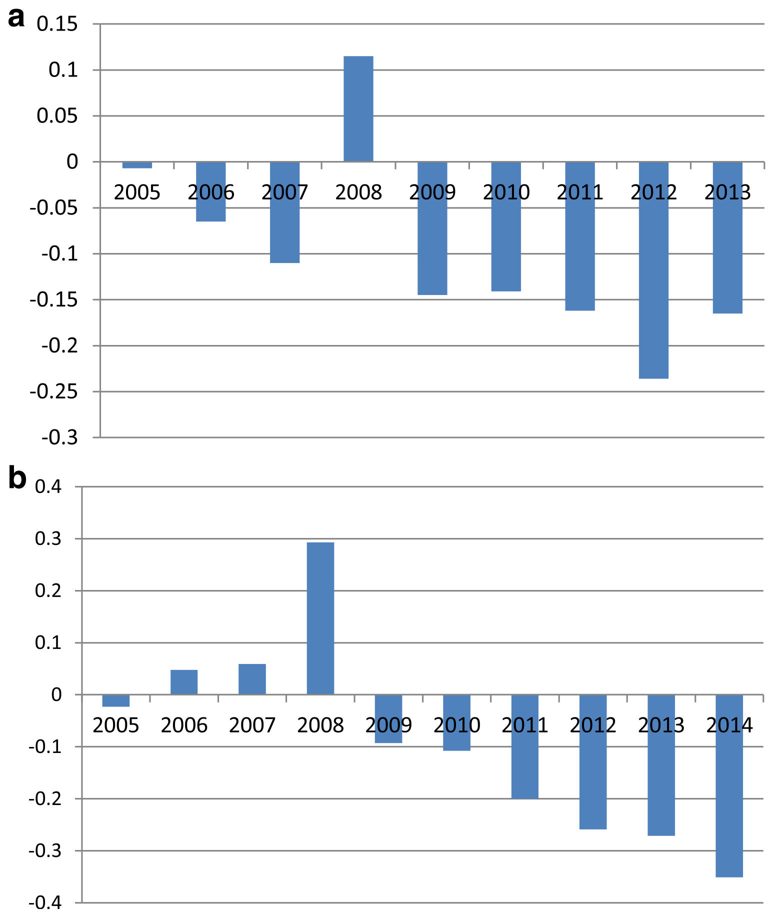

Fig. 8 Phillips curve in Italy: coefficients of year dummies interacted with province-level unemployment rates. Effect on wages calculated after removing composition effects and turnover. **a** Firms. **b** Workers. INPS data: sample of firms and workers. Robust standard errors indicate that the coefficients are negative and significantly different from zero after 2009 (at standard significance levels)

PC is −.291 (−.225 plus −.066), a value not so far from to the one obtained by macro estimates. The last column reports the GMM estimates which confirm the main findings. In columns 2 and 3, interestingly, the coefficient on inflation expectation is positive, even if rather small.

We then control for possible selection bias induced by firm entry and exit. For instance, if during the global financial crisis, firms with higher wage growth are more likely to exit the market, then our estimates are biased upward. The same happens if after 2008, only firms with lower wage growth enter the market.

To address this potential source of bias, in the central part of the table, we select all firms which were already operating before 2007, i.e. before the crisis, and we re-estimate Eqs. 7 and 8 on this subsample. Again, we find an increase in the elasticity of average wage growth to local unemployment (column 1) which rises further when we estimate Eq. 8. Also, when we consider a sample composed only of firms which did not exit the market during the period 2008–2013 (bottom part of Table 3), the results confirm a significant steepening of the PC after 2008. We then conclude that the steepening of the PC does not depend on changes in the composition of firms.

Table 3 Wage Phillips curve in Italy: firm-level estimates (p values in brackets)

	OLS	OLS	GMM
	All firms		
Δw_{t-1}	−0.00006	−0.00006	0.274
	(0.152)	(0.155)	(0.000)***
U_t	−0.00015	−0.225	−0.169
	(0.972)	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.006	−0.066	−0.083
	(0.014)**	(0.000)***	(0.000)***
$E_t\Delta p_{t+1}$		0.004	0.006
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Firm fixed effects	Yes	Yes	Yes
Observations	1.21E + 07	1.21E + 07	1.15E + 07
Number of firms	2,474,469	2,474,469	2,381,249
	All firms (already operating before 2008)		
Δw_{t-1}	−0.0004	−0.0004	0.334
	(0.002)***	(0.002)***	(0.000)***
U_t	0.001	−0.223	−0.18
	(0.775)	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.006	−0.065	−0.082
	(0.016)**	(0.000)***	(0.000)***
$E_t\Delta p_{t+1}$		0.006	0.008
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Firm fixed effects	Yes	Yes	Yes
Observations	1.04E + 07	1.04E + 07	9987063
Number of firms	1760753	1760753	1729694
	All firms (already operating before 2008 + present until 2013)		
Δw_{t-1}	−0.002	−0.002	0.200
	(0.045)**	(0.045)**	(0.000)***
U_t	0.004	−0.301	−0.230
	(0.432)	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.009	−0.053	−0.082
	(0.000)***	(0.000)***	(0.000)***
$E_t\Delta p_{t+1}$		0.005	0.006
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Firm fixed effects	Yes	Yes	Yes
Observations	6,414,023	6,414,023	6,354,583
Number of firms	750,235	750,235	750,235

Robust p values in parentheses
Significant at *10%, **5% and ***1%
Dependent variable: percentage change in the annual average wage per worker

3.3 Workers' composition and turnover

By using the random sample of employees described in Section 3.1, we now focus on workers, by replicating what already done for firms. Preliminarily, since compensation through CIG is only partial and workers in CIG face a decrease in nominal wages, we skip them from the analysis.

We estimate the same model as in 7, which is now equal to

$$\Delta w_{ipt} = \delta_t + \delta_i + \beta \Delta w_{t-1,\ p,i,} + \delta_1 U_{t,p} + \delta_2\ U_{t,p} * D_{\text{post2008}} + \epsilon_{ipt} \tag{7bis}$$

where i denotes the ith individual, p the province where the ith individual works, t is the time and δ_i and δ_t are the corresponding fixed effects. As before, the parameter δ_1 captures the correlation between the local unemployment rate and wage growth net of aggregate time trends, while δ_2 captures changes in the elasticity after 2008. As before, we do not need to include province fixed effects as they are taken into account by individual fixed effects. As in Section 3.2, we also estimate:

$$\Delta w_{ipt} = \delta_i + \beta \Delta w_{t-1,pi} + \delta_1 U_{t,p} + \delta_2\ U_{t,p} * D_{\text{post2008}} + \delta_3 E_t \Delta p_{t+1} + \epsilon_{ipt} \tag{8bis}$$

where we refer to individual i instead of firm f.

The bottom part of Fig. 8 refers to Eq. 7bis and reports the estimated elasticities of δ_1 interacted with time dummies (and estimated by excluding δ_t and $\delta_2\ U_{t,p} * D_{\text{post2008}}$, as for firms). Once again, we find a negative relationship between wage changes and unemployment, entirely determined by the increase in the elasticity from 2009 onwards.

Table 4 is divided into three parts, corresponding to different samples and three columns. In the upper part of the table, the sample is composed of job stayers, i.e. workers who worked for no less than 12 months in both time t and $t + 1$ and who are aged no more than 54 (54+ workers are excluded to avoid mis-measurement due to retirement). We also exclude workers who change firm between two consecutive years, and/or job position within the same firm, so that y-o-y wage changes are not affected by changes in the characteristic of the job. With this sample selection, we focus on quite stable workers: 95% are employed with a permanent job contract and 5% with a fixed-term ones and with a job duration longer than 1 year.

The central part of the table refers instead to a sample which is further selected, to control for the potential source of (upward) bias, which could arise if, during the crisis, firms hire only those workers who are expected to have lower expected wage growth. We then focus only on workers already employed in the firm before 2007.[8]

Last, it is possible that firms during the crisis preferred to retain only workers with lower expected wage growth and fired the other ones. To control also for this possible source of selection, in the bottom part of the table, we further select our sample by considering only workers who were continuously employed during the period 2008–2014.

As in Table 3, the first column reports the estimates of Eq. 7bis with time fixed effects. The parameters δ_1 and δ_2 are then identified thanks to the province-level variability. The estimates reported in the other two columns instead capture time trends by the use of inflation expectations and use two different estimators: OLS and GMM. In all the specifications and subsamples, we find evidence of a steepening of the PC.

In Table 5, we further split the sample of job stayers into two: blue and white collars. If variable wage components allow firm to compress wage growth when unemployment

Table 4 Wage Phillips curve in Italy: individual estimates; job stayers (*p* values in brackets)

	OLS	OLS	GMM
	All workers (employed at both time t and $t+1$)		
Δw_{t-1}	−0.211	−0.205	0.077
	(0.000)***	(0.000)***	(0.000)***
U_t	0.005	−0.417	−0.178
	(0.461)	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.008	−0.063	−0.072
	(0.053)*	(0.000)***	(0.000)***
$E_t\Delta p_{t+1}$		0.010	0.009
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Firm fixed effects	Yes	Yes	Yes
Observations	3,588,786	3,588,786	3,173,763
Number of firms	683,357	683,357	648,836
	All workers (already employed before 2008; same firm, not changing position from t to $t+1$)		
Δw_{t-1}	−0.186	−0.181	0.082
	(0.000)***	(0.000)***	(0.000)***
U_t	−0.003	−0.318	−0.178
	(0.595)	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.007	−0.061	−0.078
	(0.082)*	(0.000)***	(0.000)***
$E_t\Delta p_{t+1}$		0.010	0.010
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Firm fixed effects	Yes	Yes	Yes
Observations	2,585,290	2,585,290	2,460,382
Number of firms	507,951	507,951	497,330
	All workers (already employed before 2008 and working until 2014; same firm, not changing position from t to $t+1$)		
Δw_{t-1}	−0.189	−0.18	0.109
	(0.000)***	(0.000)***	(0.000)***
U_t	−0.023	−0.367	−0.204
	(0.004)***	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	0.002	−0.039	−0.055
	(0.780)	(0.000)***	(0.000)***
$E_t\Delta p_{t+1}$		0.009	0.009
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Firm fixed effects	Yes	Yes	Yes
Observations	1,063,247	1,063,247	1,057,311
Number of firms	132,092	132,092	132,015

Robust *p* values in parentheses
Significant at *10%, **5%, ***1%
Dependent variable: percentage change in the daily wage

Table 5 Wage Phillips curve in Italy: individual estimates; job stayers—blue and white collars (*p* values in brackets)

	White collars	Blue collars
Δw_{t-1}	0.111	0.065
	(0.000)***	(0.000)***
U_t	−0.15	−0.243
	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.121	−0.003
	(0.000)***	(0.767)
$E_t \Delta p_{t+1}$	0.009	0.008
	(0.000)***	(0.000)***
Year dummies	Yes	No
Individual fixed effects	Yes	Yes
Observations	101,185	62,486
Number of workers	687,536	496,136

Robust *p* values in parentheses
Significant at *10%, **5%, ***1%
GMM estimates. Dependent variable: percentage change in the daily wage

is high, the PC relationship could vary also by the type of worker, since typically the incidence of variable wages is higher for white collars than for blue collars (see Bryson et al. 2012). Here, we find a negative relationship with no steepening for blue collars and a large steepening for white collars, confirming indirectly that the increased shape of the PC can be influenced by adjustments in the variable component of wages.

In Table 6, we re-run the estimates on job movers, defined as those who changed firm or job position or were not employed in the firm before 2008 or lost their job after 2008. All the results are confirmed.

We now consider a specific type of turnover, determined by job contract duration and we focus on fixed-term workers (independently on the number of months worked

Table 6 Wage Phillips curve in Italy: individual estimates; job movers (*p* values in brackets)

	OLS	OLS	GMM
Δw_{t-1}	−0.218	−0.212	0.065
	(0.000)***	(0.000)***	(0.000)***
U_t	0.019	−0.412	−0.152
	(0.017)**	(0.000)***	(0.000)***
$U_t * D_{\text{post2008}}$	−0.012	−0.087	−0.073
	(0.042)**	(0.000)***	(0.000)***
$E_t \Delta p_{t+1}$		0.009	0.010
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Individual fixed effects	Yes	Yes	Yes
Observations	2,525,539	2,525,539	2,116,452
Number of workers	610,333	610,333	563,192

Robust *p* values in parentheses
Significant at *10%, **5% and ***1%
Dependent variable: percentage change in the daily wage

in a year). Since each fixed-term worker can have more than one job spell within a year, for each temporary worker, we calculate the percentage change in the average yearly daily wage. The estimates are reported in Table 7 and once again the estimates confirm the steepening of the PC after 2008, especially in the GMM specification, which supports the hypothesis of a strong increase in the slope of the PC in this segment (also when one considers the long-run effect). So, the increase of the slope of the PC can depend on composition effects, i.e. on and increase in the share of fixed-term contracts in total workforce. However, since the steepening is found also for job stayers, we can conclude that composition effects are definitely not the only cause behind the steepening of the PC in Italy.

Last, for robustness, we have also carried out some regressions (not reported) to check for further alternative specifications which could determine non-linearities, like for instance higher elasticity of wages when unemployment increases than when it decreases. We have also checked whether the PC reacts to short-run changes in unemployment (ΔU_t). The results do not support the existence of asymmetric response to unemployment nor of a short-run reaction of wages to changes in the unemployment rate.

As for firms, changes in workers' compositions are unlikely to be the (only) cause of the steepening of the PC in Italy after 2008.

4 Conclusions

The behaviour of inflation during and after the great financial crisis has sparked renewed interest in the Phillips curve. The so-called twin puzzle of missing disinflation in the aftermath of the great financial crisis and persistently low inflation in spite of the ongoing recovery in the euro area following the sovereign debt crisis have led many to rethink about the Phillips curve. While most of the literature has focused on the relationship between consumer' and/or producer' price inflation and economic slack, in this paper, we look at wage inflation for two important reasons. First, there is evidence

Table 7 Wage Phillips curve in Italy: individual estimates; fixed-term workers (*p* values in brackets)

	OLS	OLS	GMM
Δw_{t-1}		−0.048	0.260
		(0.000)***	(0.000)***
U_t	−0.088	−0.067	−0.014
	(0.000)***	(0.000)***	(0.858)
$U_t * D_{\text{post2008}}$	0.008	−0.058	−0.304
	(0.560)	(0.001)***	(0.000)***
$E_t\Delta p_{t+1}$		0.005	0.009
		(0.000)***	(0.000)***
Year dummies	Yes	No	No
Individual fixed effects	Yes	Yes	Yes
Observations	791,638	419,922	419,922
Number of workers	349,803	192,702	192,702

Robust *p* values in parentheses. Since fixed-term workers can have more than one job spell within a year, the dependent variable is the percentage difference between the average daily wages in two consecutive years
Significant at *10%, **5% and ***1%
Dependent variable: percentage change in the average daily wage

that in the euro area, domestic factors played an important role besides global disinflationary forces. Second, in several euro area countries, labour markets have gone through some modification following the globalization process, the monetary union and more recently the double dip recession. In particular, the interaction of a deep and prolonged crisis with policy reforms raises the question whether parameter instability and/or non-linearities might shape the response of wages to the unemployment rate. We tackle these issues from a variety of angles, using both macro time series models and micro panel data (for Italy only). We find evidence that the wage Phillips curve has changed since the great financial crisis. In particular, we find evidence of an increased correlation between wage inflation and the unemployment rate in Italy, France and Spain while such correlation has diminished in Germany. For Italy, this macro evidence is supported also by microeconometric evidence. We then try to uncover some of the factors behind it and suggest two explanations. One refers to the increased pro-cyclicality of incumbent wages, through the increasing use of flexible wage schemes, and the other refers to the increased use of fixed-term contracts which by their own nature leads to higher frequency of re-setting wages, a phenomenon that mirrors the higher frequency of price adjustment recorded among Italian firms after the great financial crisis (Fabiani and Porqueddu 2017) and put forward as an explanation of the increased responsiveness of price inflation to economic slack. Our results are robust to a variety of controls and in particular to changes over the business cycle in the composition of the universe of firms (due to entry and exit phenomena) and of the workforce.

Endnotes

[1]Indeed, the debate about a flattening of the PC in advanced economies started before the global financial crisis. Borio and Filardo (2007) relate it to a structural change due to globalization and the stringer competitive pressure of Asian economies. Gaiotti (2010), using firm-level data, confirms the existence of a flattening but finds that it is not related to globalization but is probably related to the moderate dynamics of inflation expectations.

[2]An exception is Rosolia (2015) finds that in Italy, negotiated wage growth is negatively affected by labour market slack. The correlation of aggregate actual wages and unemployment is however quite weak because of the long duration of centrally bargained job contracts (3 years in Italy). Izquierdo et al. (2017), using the data of a survey conducted by the European central banks within the "Wage dynamics network" project (WDN), show that, in spite of the presence of institutional rigidities preventing nominal and real wage cuts, during the period 2010–2013, more than 20% of French and Spanish firms hit by a shock could freeze or cut wages. This share was equal to 17% in Italy and 11% in Germany.

[3]Different explanations based on the relationship between firms' pricing decisions and capacity utilization support the hypothesis of (short-run) non-linearities when market slack is very low (e.g. Clark et al. 1996, for the US case).

[4]In Italy, with the exception of the years following 2008, robust evidence of a PC can be found only for the 1970s (see Visco 1984).

[5]It is respectively equal to −.11 in Italy, −.13 in France, −.18 in Spain and −.05 in Germany.

[6]While there is a huge literature on the responsiveness of the PC at the local unemployment rate (see e.g. Bodo and Sestito, 1994, for Italy; more recently Gregg et al., 2014, for the UK), empirical studies on the wage PC at the local level are rather scant.

[7]Of course, also, the term $\delta_2\ U_{t,p} * D_{\text{post2008}}$ is excluded.

[8]The opposite case, i.e. that during the crisis, firms prefer to hire workers with higher growth potential, would lead to a downward bias of our estimates, without significant affect in our conclusions.

Acknowledgements
We thank Stefano Marucci for the valuable research assistance; Effrosyni Adamopoulou, Emmanuele Bobbio, Marta De Philippis and Federico Giorgi for the helpful discussions and Matteo Bugamelli, Fabio Busetti, Claudio Lucifora, Paolo Sestito, the editor and an anonymous referee for the comments. The views expressed here are solely ours; in particular, they do not necessarily reflect those of the Bank of Italy.
Responsible editor: Denis Fougère

Competing interests
The IZA Journal of Labor Policy is committed to the IZA Guiding Principles of Research Integrity. The authors declare that they have observed these principles.

Publisher's Note

Received: 8 May 2017 Accepted: 13 July 2017
Published online: 29 August 2017

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