Journal of Macroeconomics 63 (2020) 103166

Contents lists available at ScienceDirect

Journal of Macroeconomics

journal homepage: www.elsevier.com/locate/jmacro

Wage indexation and the monetary policy regime☆

Selien De Schryder*,[a], Gert Peersman[a], Joris Wauters[a,b]

[a] Department of Economics, Ghent University, Sint Pietersplein 5, Gent B-9000, Belgium
[b] National Bank of Belgium, de Berlaimontlaan 14, Brussels 1000, Belgium

ARTICLE INFO

Keywords:
Wage indexation
Monetary policy regimes
Cross-country panel
Phillips curve

JEL classification:
C23
E42
J30

ABSTRACT

We estimate a New Keynesian wage Phillips curve for a panel of 24 OECD countries and allow the degree of wage indexation to past inflation to vary according to structural characteristics. We find that the degree of wage indexation is significantly lower for countries with an inflation target. However, this effect vanishes when we control for the degree of goods market competition. By contrast, more goods market competition is consistently associated with lower wage indexation. This robust finding puts into question whether embedding a constant degree of wage indexation in standard DSGE models is truly structural.

1. Introduction

New Keynesian dynamic stochastic general equilibrium (DSGE) models typically assume sticky wages and partial wage indexation to past inflation. Notably, the degree to which wages are indexed to past inflation is hard-wired as a fixed and policy invariant parameter (e.g., Christiano et al., 2005; Smets and Wouters, 2007).

The assumption of a constant degree of wage indexation has, however, been rejected by institutional and empirical evidence for the United States (US). In particular, the coverage of private-sector workers by cost-of-living adjustment (COLA) clauses rose substantially between the late 1960s and mid-1980s, after which it declined again (Holland, 1986). Hofmann et al. (2012) estimate the extent of wage indexation in the US over time and find a considerably higher degree of indexation during the “Great Inflation” of the 1970s compared to the earlier and later periods. They infer that the reduction in indexation from the mid-1970s to 2000 led to a decline in the long-run impact of a supply and demand shock on prices by respectively 44% and 39%. Since changes in wage indexation practices have clearly had significant macroeconomic consequences, it is essential to understand why the degree of indexation changed over time.

Hofmann et al. (2012) argue that the rise and fall of wage indexation can be explained by a weaker reaction of the Federal Reserve to inflation during the “Great Inflation” and more aggressive inflation stabilization before and after this period. Intuitively, a regime

☆ We thank the editor David VanHoose and three anonymous referees for their constructive comments and suggestions. We further thank Julio Carrillo, Zulal Denaux, Andrej Drygalla, Markus Eberhardt, Gerdie Everaert, Simon Price, Arnoud Stevens, Rasmus Thönnessen, and conference participants of the 2014 Ghent-Lille workshop on economics, the 2015 RGS Doctoral Conference in Economics, the 2015 IAES conference, the 2015 SNDE conference and the 2015 SMYE conference for their comments and suggestions. We are grateful to Joshua Chan for sharing his replication files and to Elena Rusticelli for sharing her NAIRU data. We acknowledge financial support from the Research Foundation Flanders (FWO) and the Special Research Fund (BOF) of Ghent University. All remaining errors are ours.

* Corresponding author.
E-mail addresses: selien.deschryder@ugent.be (S. De Schryder), gert.peersman@ugent.be (G. Peersman), joris.wauters@nbb.be (J. Wauters).

https://doi.org/10.1016/j.jmacro.2019.103166
Received 15 November 2018; Received in revised form 30 October 2019; Accepted 30 October 2019
Available online 06 November 2019

of high and volatile inflation fosters the use of wage indexation clauses as protection against inflation uncertainty.[1] This conjecture is also supported by economic theory that shows that the optimal proportion of wage contracts indexed to inflation increases with the variance of monetary disturbances (Gray, 1978) and that the likelihood of indexation rises when inflation uncertainty is higher (Ehrenberg et al., 1983).

However, other explanations for changes in wage indexation have also been put forward. Carrillo et al. (2017) demonstrate that utility-maximizing workers base their indexation choice on the relative importance of shocks in the economy. More specifically, workers want to index wages to past inflation when technology and permanent shocks to the inflation target dominate output fluctuations, but not when aggregate demand and temporary inflation target shocks dominate. Based on a counterfactual calibration exercise for the US, they attribute the high degree of wage indexation in the 1970s primarily to very volatile supply-side shocks relative to demand-side shocks. Changes in the monetary policy rule and the stability of the inflation target, in contrast, had probably only a minor influence on wage indexation in the US. Furthermore, Duca and VanHoose (1998) find empirical evidence supporting their theory of an inverse relationship between the degree of product market competition and wage indexation. Intuitively, increased goods market competition raises the price elasticity of demand, which makes equilibrium employment less sensitive to demand shocks and thereby reduces the incentive to index. Finally, the degree of wage indexation might also be linked to labor market characteristics. For example, Messina and Sanz-de Galdeano (2014) find that a decline in union coverage and a more decentralized wage bargaining reduces wage indexation. In sum, it remains unclear to what extent the degree of wage indexation in the economy, and its variation over time, can be linked to the monetary policy regime in place. Our paper aims to fill this gap.

This study is the first to use cross-country data to examine whether wage indexation varies systematically across monetary policy regimes. We estimate variants of the reduced-form empirical New Keynesian wage Phillips curve of Galí (2011) on a panel dataset covering 24 OECD countries between 1975Q1 and 2016Q4. We measure structural differences in wage indexation across countries, as well as changes in wage indexation over time. We innovate relative to the literature in two ways. First, we take into account several possible drivers behind changes in wage indexation. To measure the effects of the monetary policy regime, we distinguish between the presence of an inflation, money growth, and exchange rate target. This distinction is motivated by the differences between the underlying dynamics of these strategies and their impact on the formation of inflation expectations. For example, inflation-targeting central banks typically try to stabilize inflation in the short to medium term, whereas money growth targeting is more a commitment to low inflation in the long run.[2] In addition, we take into account the degree of product market competition and wage bargaining characteristics as alternative drivers for wage indexation changes.

Second, we use a panel dataset to exploit the information across the cross section and to increase the number of observations significantly. This is crucial because monetary policy targets and wage bargaining characteristics can show limited variation over specific periods, yet differ importantly between countries. By contrast, existing studies typically focus on the US alone and thereby capture less variation.

The estimation results provide important considerations for macroeconomic analysis and policymakers. First, our benchmark regressions without interaction effects show that wage indexation is significant and economically relevant for the sample under analysis. Second, when allowing wage indexation to vary depending on the monetary target, we find an economically and statistically significant reduction of wage indexation for countries with an inflation target. Exchange rate or monetary targets, by contrast, do not show systematic effects. This result resonates with Benati (2008), who estimates the price Phillips curve on historical data for a set of countries and finds the price indexation parameter to be very low or zero under stable monetary policy regimes with clearly defined nominal anchors. Third, when we include structural characteristics on the degree of product market competition and wage bargaining, the effect of inflation targeting essentially vanishes. More precisely, we find evidence that a higher degree of product market competition lowers the degree of wage indexation in an economically and statistically significant way. This effect is found to withstand a broad set of robustness checks. Therefore, we generalize the results from Duca and VanHoose (1998) for a broad set of countries and a longer time span.

Since product market competition is an outcome of economic policy, we conclude that the constant indexation assumption embedded in standard DSGE models is susceptible to the Lucas (1976) critique, i.e., it is not intrinsic to the deep structure of the economy and not a policy invariant parameter. This is relevant for policymakers, given the macroeconomic importance of wage indexation on output and inflation. Furthermore, our results suggest that monetary policy is not a crucial driver of wage indexation, which is consistent with Carrillo et al. (2017). But we caution against interpreting the statistically insignificant effects from the monetary policy targets as definitive evidence that monetary policy is irrelevant for wage indexation. Our results are consistent with insignificant effects of monetary policy, but they are also consistent with the view that it is challenging to disentangle monetary policy effects from product market competition effects. Indeed, we show that the shift towards more product market competition is strongly correlated with the adoption of inflation targets.

Our work relates to literature that links changing wage or price dynamics to shifts in macroeconomic policy. For instance, Messina and Sanz-de Galdeano (2014) use micro-level data to document how Brazil's and Uruguay's disinflation policies changed the nature of wage rigidities. Alogoskoufis and Smith (1991) study wage and price inflation series from 1892 to 1987 for the US and the UK; they report coinciding shifts in the wage Phillips curve and price inflation persistence, which they link to departures from international fixed exchange rate regimes. Levin et al. (2004) find that inflation expectations appear to be more forward-looking, and inflation less persistent, in inflation-targeting countries.

[1] Holland (1986) and Ragan and Bratsberg (2000) attribute the rise of indexation practices in the 1970s to much higher inflation uncertainty.

[2] The inflationary outcomes of the three different types of nominal anchors also turn out to be different (Fatás et al., 2007).

A related study is also Fatás et al. (2007), who find that having an explicit quantitative target for monetary policy, in particular, an inflation target, is systematically related to a lower average level of price inflation. Moretti (2014) furthermore finds that inflation targeting and product market deregulation have both lowered the level of price inflation in OECD economies. More recently, López-Villavicencio and Saglio (2017) provide cross-country evidence that wage indexation to past inflation has decreased when countries shifted to a low inflation environment.[3] Finally, our study is related to the literature that analyzes the role of monetary policy institutions for inflation outcomes and economic growth, such as central bank independence (Alesina and Summers, 1993) and transparency (Sterne et al., 2002; Eijffinger and P., 2006; Dincer and Eichengreen, 2014).

The remainder of the paper is organized as follows: In the next section, we present the estimation results for a benchmark wage Phillips curve model with a constant degree of wage indexation. In Section 3, we extend the benchmark model to analyze the influence of the monetary policy regime on the extent of backward-looking wage indexation, while controlling for the degree of goods market competition and labor market characteristics. Finally, Section 4 concludes.

2. Wage Phillips curve with constant indexation

We first derive a benchmark empirical New Keynesian wage Phillips curve in Section 2.1. Section 2.2 presents the data for these main variables and discusses some econometric issues. The estimation results are shown in Section 2.3.

2.1. Model specification

Our theoretical framework is based on Galí (2011), who derives the empirical wage Phillips curve from a New Keynesian model that includes the unemployment rate. He derives both a theoretical foundation for the empirical wage Phillips curve relation and a structural interpretation of the reduced form coefficients. The model assumes staggered wage setting as in Erceg et al. (2000), which means that a worker's wage cannot be re-optimized in every period. When the wage cannot be reset, it is assumed to be indexed to a measure of price inflation ($\bar{\pi}_{t-1}^{p}$), the central bank's inflation target (π^*) and trend productivity growth (g): $\gamma\bar{\pi}_{t-1}^{p} + (1-\gamma)\pi^* + g$, with parameter $\gamma \in (0, 1)$ determining the weights.

When we denote wage inflation by π_t^w and the difference between unemployment and the natural rate by $\hat{u}_t \equiv u_t - u^n$, the model's reduced form representation is given by

$$\pi_t^w = \alpha + \gamma\bar{\pi}_{t-1}^{p} + \psi_0\hat{u}_t + \psi_1\hat{u}_{t-1}, \tag{1}$$

with $\alpha \equiv (1-\gamma)\pi^* + g$ (Galí, 2011, eq. 19).

Bringing (1) to a panel data setting results in the econometric *benchmark wage inflation model:*

$$\pi_{i,t}^w = \alpha_i + \gamma\bar{\pi}_{i,t-1}^{p} + \psi_0\hat{u}_{i,t} + \psi_1\hat{u}_{i,t-1} + \eta_{i,t}, \tag{2}$$

where subscripts i and t indicate the country and time period, and α_i represents $(1-\gamma)\pi_i^* + g_i$ plus additional country-specific and time-invariant effects. Our main interest is the degree of indexation to past price inflation (γ), which is expected to lie between 0 and 1. Galí considers both past inflation π_{t-1}^{p} and smoothed past inflation $1/4(\pi_{t-1}^{p} + \pi_{t-2}^{p} + \pi_{t-3}^{p} + \pi_{t-4}^{p})$ as indexation measure $\bar{\pi}_{t-1}^{p}$.[4] The central bank's inflation target and trend productivity growth are assumed to be constant. The benchmark wage inflation model can be extended with a time-varying measure of trend inflation as discussed in detail in Appendix B.

2.2. Panel dataset and econometric considerations

Our dataset consists of an unbalanced panel covering quarterly data for 24 OECD economies.[5] We use the information from a group of countries to broaden the information set and to increase the power of the tests. The sample period runs from 1975Q1 to 2016Q4, where we keep the estimation sample fixed over the different model specifications to compare the coefficient estimates.

We consider the OECD average hourly earnings of employees in the manufacturing sector as our measure of wages. The earnings data are comparable to wage rate series that proxy for the basic wages or cost-of-living allowances, but they provide a more complete measure of the overall wage income because they also include premium pay for overtime and bonuses.[6] One caveat of using the

[3] López-Villavicencio and Saglio (2017) estimate wage Phillips curves which allow the degree of wage indexation to vary according to a high, medium, and low inflation regime. The results reveal that wage indexation has been lower in low inflation environments. We are not aware of theories predicting a link between wage indexation and the level of inflation. A possible explanation is that high-inflation regimes are typically also characterized by more volatile inflation and hence more inflation uncertainty, which encourages indexation.

[4] The empirical wage Phillips curve specification in (2) is based on the assumption that the unemployment rate follows an AR(2) process (see Galí, 2011, equation (18) p.447, where an AR(2) process is assumed for US data). This AR(2) assumption fits the data optimally for 9 of the 24 countries in our sample based on the standard Bayesian information criterion, the largest number of countries in total for which the lag length is optimal when we let the lag length vary from 1 to 5. Based on these observations, we deem the AR(2) assumption on the unemployment rate process to be a valid choice for the reduced form wage Phillips curve specification (2) in our panel data setting.

[5] The 24 countries are: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Estonia, Finland, France, Germany, Hungary, Ireland, Italy, South Korea, the Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, the United Kingdom and the United States.

[6] Compensation rates, which also include employer contributions to social security or social insurance schemes, are a widely used alternative measure. However, this series is not available for 10 of the 24 countries in our sample.

Table 1
Results benchmark model.

Regressand:	Wage inflation ($\pi^w_{i,t}$)			
	(1) POLS	(2) FE	(3) MG	(4) IV
Unemployment gap ($\hat{u}_t$)	−0.401*** (0.095)	−0.369*** (0.092)	−0.412*** (0.104)	−0.739*** (0.179)
Lagged unemployment gap ($\hat{u}_{t-1}$)	0.352*** (0.094)	0.315*** (0.092)	0.318*** (0.118)	0.677*** (0.175)
Lagged price inflation ($\bar{\pi}^p_{i,t-1}$)	0.863*** (0.050)	0.822*** (0.059)	0.619*** (0.072)	0.841*** (0.030)

Note: Results are shown for pooled OLS (POLS), fixed effects (FE), mean group (MG), and instrumental variable (IV) estimations. Symbols *, **, and *** denote significance at 10%, 5%, and 1% levels, respectively. Standard errors are in brackets (robustified to heteroskedasticity and serial correlation within cross sections for the pooled estimators). Instrumental variables: second lag of the unemployment gap for the current unemployment gap and the second lag of smoothed price inflation for price inflation. The first stage IV F-statistic (Kleibergen-Paap RK Wald weak identification test statistic) equals 93.4. Sample: $\overline{T}=123.4$, max $T=168$, min $T=60$ and N = 24.

earnings data for the manufacturing sector, instead of the entire private sector, is that we might pick up industry-specific wage dynamics related to differential changes in the labor share of manufacturing and the relative importance of the manufacturing sector in the aggregate economy. The data availability of entire private sector data is, however, highly restricted in time.

Our price measure is the quarterly all-items consumer price index. We construct quarter-on-quarter wage and price inflation series (respectively π_t^w and π_t^p) as 100 times the log differences of the wage and price indices. For our main results we take smoothed past inflation $\bar{\pi}^p_{t-1} \equiv 1/4(\pi^p_{t-1} + \pi^p_{t-2} + \pi^p_{t-3} + \pi^p_{t-4})$ as the measure of price inflation to which wages are indexed. The unemployment gap is the difference between the unemployment rate and the NAIRU. Before 1990Q1, annual NAIRU series have been interpolated to a quarterly frequency with a quadratic polynomial. Appendix A provides further details on the coverage and data definitions.

To estimate the benchmark wage inflation model, we need to take into account two econometric issues. First, Eq. (2) postulates homogeneous regression parameters. When the regression does not contain lagged dependent variables, as in Eq. (2), and the regressors are strictly exogenous, both homogeneous and heterogeneous estimators deliver consistent average estimates when the coefficients differ randomly (Pesaran and Smith, 1995). To check to which extent the coefficient homogeneity assumption affects our small sample panel coefficient estimates, we apply a dispersion type test based on the work of Swamy (1970) and test the equality of slope coefficients across individual regressions.

Second, we consider the possibility of the joint determination of wage inflation and the unemployment gap as well as potential endogeneity between current wage inflation and the price indexation measure. To account for these forms of endogeneity and verify the robustness of the results, we employ an instrumental variable (IV) estimator where we instrument both the price inflation variable and the current unemployment gap with their respective second lag.

2.3. Results

Table 1 depicts the estimation results for the benchmark wage inflation model (2) for the homogeneous Pooled OLS (POLS), the Fixed Effects (FE) and the heterogeneous Mean Group (MG) panel estimators in the first three columns. All estimates indicate a substantial degree of wage indexation to past inflation.

The point estimates of γ range from 0.62 to 0.86, which is in line with the estimated range of 0.52 and 0.83 obtained by Galí (2011) for the US. The negative contemporaneous and positive lagged effects of the unemployment gap are in line with theoretical expectations (Galí, 2011). On impact, a 1 percentage point decline in the unemployment gap leads to a respective 0.37% to 0.41% increase in nominal wages in the FE and MG models. All parameters are statistically significant at the 1% level.[7] Tests on coefficient homogeneity however point to a significant cross-sectional heterogeneity in the coefficients.[8]

The last column of Table 1 shows the FE estimates[9] of the IV model. The results point to a similar indexation coefficient and larger unemployment coefficients, although the latter are less precisely estimated. The wage indexation coefficients appear to be robust to accounting for the potential endogeneity concerns between wage inflation and respectively lagged price inflation and the unemployment gap.

[7] We perform the empirical analysis in STATA 15 with the user-written xtmg routine of Eberhardt (2012) for the MG estimates. We estimate cluster robust standard errors which are robust for heteroskedasticity and first order autocorrelation in the residuals. In the fixed effects regressions, robust standard errors were computed with the xtreg robust option, as advocated by Cameron and Miller (2015).

[8] Swamy's test of homogeneity of the pooled ordinary least square (POLS) estimates generates a significant chi-square test statistic of 597.21. Also the homogeneity assumption of the intercept term in the POLS estimator can be rejected. Wald tests on the individual-specific OLS slope coefficients (MG estimation) reject the equality across individuals for each regressor based on chi-square test statistics of 171.13, 90.82, and 94.31 for the respective regressors $\bar{\pi}^p_{i,t-1}$, $\hat{u}_{i,t}$ and $\hat{u}_{i,t-1}$.

[9] Country-specific IV regressions point to weak identification. We consider the FE estimator as the baseline pooled estimator since the homogeneity assumption of the intercept term in the POLS estimator can be rejected.

3. Wage Phillips curve with variable wage indexation

3.1. Linking wage indexation to structural drivers

Monetary policy A number of theoretical studies conclude that the degree of wage indexation to past inflation may depend on inflation uncertainty and the conduct of monetary policy. Specifically, Gray (1978) presents a neoclassical model with short-term wage rigidities and uncertainty, and shows that the degree of wage indexation that minimizes the deviation of output from full-information output increases with the variability of monetary disturbances. Ehrenberg et al. (1983) demonstrate in an efficient contract model that a rise in inflation uncertainty may lead to greater use of wage indexation because wage indexation helps to insulate the worker's real wage from the effects of unanticipated inflation, whereas it reduces the impact of lower-than-anticipated inflation on the real cost of labor inputs for firms. Carrillo et al. (2017) show that utility maximizing wage setters raise the extent of wage indexation to past inflation when the variability of permanent shocks to the inflation target of the central bank increases, whereas the amount of indexation declines when there is a rise in the volatility of temporary inflation target shocks.

A possible link between wage indexation, inflation uncertainty and monetary policy is also supported by institutional and empirical evidence for the US.[10] Holland (1986) demonstrates that the proportion of cost-of-living adjustment clauses in major collective bargaining agreements was much higher in the 1970s and first half of the 1980s than the preceding and subsequent periods. He finds that this pattern can be explained by a sizable increase of inflation uncertainty in the 1970s, measured by the mean squared forecast error of inflation surveys. Hofmann et al. (2012) estimate the evolution of US wage dynamics over time, and find a degree of wage indexation to past inflation of 0.91 during the "Great Inflation" of the 1970s, compared to 0.30 and 0.17 before and after this period.[11] Hofmann et al. (2012) argue that this evolution can be explained by a shift in the monetary policy reaction function of the Federal Reserve. More specifically, a weakly inflation stabilizing conduct of monetary policy in the 1970s resulted in high and volatile inflation, which encouraged the use of indexation clauses in wage contracts as a protection against inflation uncertainty. Conversely, the credible establishment of price stability after the disinflation of the early 1980s reduced the need for protection against unforeseen inflation, thus mitigating wage indexation.[12] Given the suggestions of an important role of monetary policy for wage indexation practices, we formally examine the influence of the monetary policy regime on the wage indexation parameter in the following sections for the panel of OECD countries while accounting for other possible drivers.

Product market competition The degree of product (or goods) market competition is considered to affect the attractiveness of wage indexation (Duca and VanHoose, 1998; Duca, 1998). More specifically, the conjecture is made that wage indexation becomes more attractive when the ex-ante variance of aggregate demand shocks increases relative to that of aggregate supply shocks. In the face of aggregate demand shocks, indexation keeps wages close to their market-clearing level, whereas the opposite holds in case of aggregate supply shocks (Gray, 1976). The benefit of indexation in the light of aggregate demand shocks - less variation in real wages -, however, lowers with the degree of product market competition because firms become less sensitive to aggregate demand shocks. By contrast, the costs of indexation following aggregate supply shocks - more variation in employment - increase with more competition. Combining higher costs of wage indexation following supply shocks with smaller benefits following demand shocks leads to the prediction that more product market competition negatively impacts the degree of wage indexation. Duca and VanHoose (1998) document empirical support for this theoretical prediction. They find that greater goods-market competition lowers the extent of wage indexation in the US for the period 1956–1995.

Wage bargaining characteristics Changes in the wage bargaining process have been shown to influence the extent of wage indexation practices. Cecchetti (1987), for instance, documents that policy interventions in the bargaining process in the US during the 1960s and early 1970s altered the effective degree of indexation. More recently, Messina and Sanz-de Galdeano (2014) relate the different evolution of wage indexation in Uruguay and Brazil to the dynamics in the centralization level of wage bargaining and changes in the union coverage. The authors find that a decline in union coverage and a more decentralized wage bargaining reduces wage indexation. Carrillo et al. (2017) further show that the economy's equilibrium degree of wage indexation can differ depending on whether the labor market coordination is centralized or decentralized.

We focus on the coordination level of wage bargaining and the union coverage rate to measure the effects of changes in the wage bargaining process on wage indexation. Both reflect the wage-setting power of labor unions in wage bargaining negotiations. Indices of the coordination level of wage bargaining capture the degree to which major institutional players' decisions extend to lower-level institutions. Union coverage instead reflects the degree to which collective agreements are applied to non-unionized workers.

[10] The link between monetary policy and wage setting more generally has been the subject of theoretical work. Ball and Cecchetti (1991) and Waller and VanHoose (1992) study the interaction between wage indexation and monetary policy, while Fethke and Policano (1987) determine the optimal pattern of negotiation and the pattern of monetary policy intervention.

[11] Hofmann et al. (2012) first estimate a time-varying parameters Bayesian structural vector autoregressive (TVP-BVAR) model, and document considerable time variation in the impulse responses of wages and prices to aggregate supply and demand shocks. In a second step, the parameters of a standard DSGE model containing a wage Phillips curve are estimated for respectively 1960Q1, 1974Q1 and 2000Q1, by matching the impulse responses from the TVP-BVAR for these periods with the impulse responses of the DSGE model using a Bayesian impulse response matching procedure.

[12] Similar to Hofmann et al. (2012), Muto and Shintani (2014) perform an empirical evaluation of the New Keynesian Wage Phillips Curve for Japan and the US. They show with rolling window regressions that the importance of wage indexation has declined over time for both countries, which they explain via lower and more stable inflation. Ascari et al. (2011) also find a similar pattern of time-variation in wage indexation for the US by means of rolling window estimates of a reduced form wage equation.

3.2. Econometric model specification

We explore whether the degree of wage indexation varies according to the aforementioned structural drivers by extending the benchmark wage Phillips curve of Section 2 as follows:

$$\pi_{i,t}^{w} = \alpha_i + \tilde{\alpha}'D_{i,t} + (\gamma + \tilde{\gamma}'D_{i,t})\bar{\pi}_{i,t-1}^{p} + \psi_{0,i}\hat{u}_{i,t} + \psi_{1,i}\hat{u}_{i,t-1} + \eta_{i,t}, \tag{3}$$

We allow for cross-country heterogeneity in the fixed effects and unemployment gap coefficients given the rejection of coefficient homogeneity for the benchmark model. We nevertheless use common parameters for the interaction variables to measure the drivers of changes in wage indexation. Despite the fact that the interaction variables are country- and time-specific, many of the regime indicators under examination display substantial persistence over time for multiple countries (infra Section 3.3) which prevents the estimation of Eq. (3) at the individual country level.

$D_{i,t}$ is a $k \times 1$ vector of variables that are interacted with past inflation and the intercept, whereas $\tilde{\alpha}$ and $\tilde{\gamma}$ are $k \times 1$ vectors with the corresponding interaction coefficients.[13] In the first instance, we solely focus on the most popular candidate driver of wage indexation, the monetary policy regime. Following Fatás et al. (2007), we asses whether having a monetary target matters in the first place, and distinguish between the type of target in a second step. In this case, $d_{i,t} \equiv (mp_{i,t})'$. The policy target interaction term (MP) therefore indicates the presence of an explicit quantitative monetary target (TARGET), or respectively an inflation target (IT), a money growth target (MON) or an exchange rate target (ER). Next, we additionally control for the effects of product market competition (ETCR) and union power (COORD and COV) in the wage bargaining process, implying that $d_{i,t} \equiv (mp_{i,t}, etcr_{i,t}, coord_{i,t}, cov_{i,t})'$.

3.3. Measuring the structural drivers

Monetary policy regimes We identify the monetary policy regime by the presence of an explicit quantitative monetary target following Fatás et al. (2007), see Appendix A.2. It is commonly accepted that a policy regime that clearly commits to a nominal anchor can help promote price stability and stabilize inflation expectations (Mishkin, 2007). A quantitative monetary target should help to lower inflation uncertainty. Accordingly, the degree of wage indexation is also expected to be lower in a monetary policy regime with an explicit nominal anchor. The advantage of defining a policy regime by the presence of a quantitative target is that it can easily be measured and verified in an objective and mechanistic way (Fatás et al., 2007).

There exist different types of nominal anchors, and in this work, we consider three different monetary targets: the inflation rate, the exchange rate, and the money supply. These different targeting strategies have distinct characteristics. With an inflation target, policy decisions are based on conditional medium-run inflation forecasts, which allows for close monitoring by the private sector. This is especially the case for an inflation targeting strategy, given its high degree of transparency and accountability (Svensson, 1999).

A fixed exchange rate target is also easily communicated to the public, but monetary policy cannot react to domestic shocks independently from the anchor country. This restriction makes the fixed exchange rate vulnerable to speculative attacks, which increases inflation uncertainty. Money growth targeting also offers a transparent anchor to the public, but this strategy should be seen primarily as a way to communicate a commitment to low and stable inflation in the long run (Issing, 1996). Targeting a money aggregate can lead to more inflation uncertainty in the short to medium term due to the unstable relationship between inflation and money aggregates and the presence of money demand shocks.

This distinction is also validated by the empirical literature analyzing the macroeconomic outcomes of different monetary regimes. Fatás et al. (2007), for instance, find that an inflation targeting regime lowers inflation significantly more than an exchange rate or a monetary targeting regime. Furthermore, inflation targeting has anchored inflation expectations (Walsh, 2009) and lowered inflation persistence (Mishkin, 2007).[14] Fixed exchange rate regimes have, in some cases, been linked to higher output volatility and lower and more stable inflation (Ghosh et al., 1997; Levy-Yeyati and Sturzenegger, 2001). Monetary target regimes have been found to keep inflation under control in the longer run by means of a flexible approach towards the target rule and an active and elaborate communication of the monetary policy strategy to the public (Mishkin, 1999).

Fig. 1 summarizes the evolution of the monetary regimes for the countries in our sample over 1960Q1-2016Q4. The overall pattern is consistent with the shifts in policy regimes documented in the literature (Fatás et al., 2007). Exchange rate targeting was the dominant regime during the 1960s and early 1970s while in the mid-1970s some countries shifted to monetary targeting.[15] Since the early 1990s, many countries switched to an inflation target, which becomes the dominant targeting strategy at the end of the sample. We classify eurozone members as having both an inflation and a monetary target since the emergence of the European Central Bank in 1999, which is why the series jump at this point. Notice that the total number of regimes increases over time and always exceeds the number of countries, which is due to countries combining different targets. Table 2 shows that there are 58 regime

[13] Interacting with the intercept prevents the estimation of spurious interaction effects (Brambor et al., 2006).

[14] For evidence that inflation targeting leads to better-anchored inflation expectations, see e.g., Gürkaynak et al. (2007, 2010) and Levin et al. (2004).

[15] Exchange rate frameworks characterized as a managed or free float are not considered as an exchange rate targeting regime.

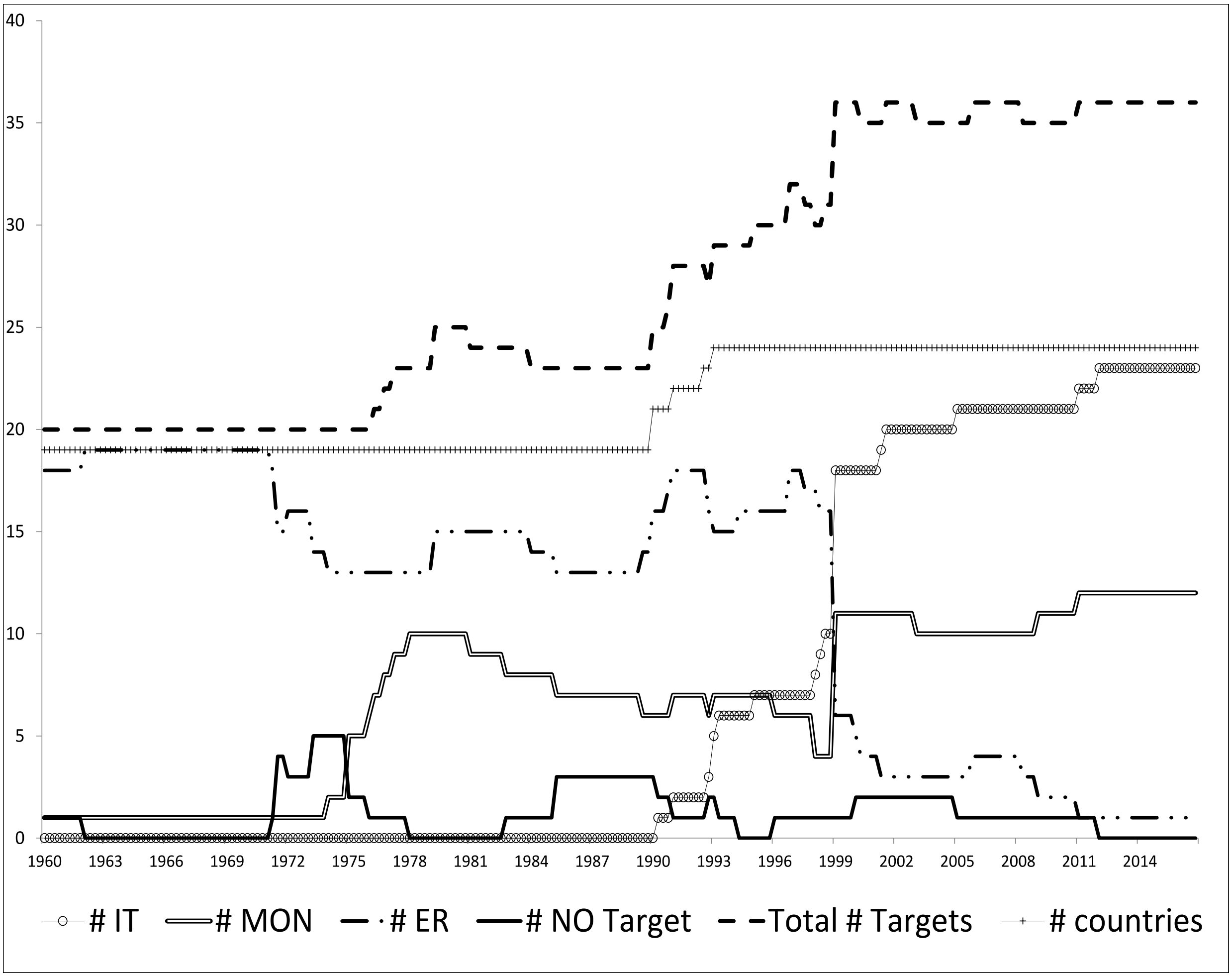

Fig. 1. Monetary targets. **Note:** The figure shows the evolution of the amount of countries with an Inflation, Monetary, and/or Exchange rate target over time, as well as countries without a formal target. We also show the amount of countries in the sample, and the total number of targets at each point in time.

Table 2
Summary statistics: Target switches.

From To	ER	MON	IT	No target
ER	-	9	13	13
MON	1	-	2	3
IT	0	0	-	0
No Target	5	6	6	-

Note: The numbers indicate the total amount of events where a target was replaced by another. ER,MON & IT stand for respectively exchange rate targeting, monetary targeting and inflation targeting.

switches in total.[16] We conclude that there is quite some variation in the data in terms of policy regime shifts in the panel of countries under analysis. At the same time, the persistence of the regimes at the individual country level is also apparent, which emphasizes the need for a cross-country dimension to examine the effects of policy regimes.

Product market competition The degree of product market competition is measured by the OECD indicators of regulation in non-manufacturing sectors (NMR), more specifically the aggregate indicator of regulation in energy, transport and communications

[16] Regime switches occur when a monetary target is abandoned and replaced by another target. Of these 58 cases, only 7 are "double counted" because a country with one target switched to a regime with 2 other targets, and vice versa. If we also include cases where there is a *change* in the policy regime without abandoning the regime already in place (including cases where, for instance, a monetary targeter adds an inflation target as second objective), we obtain a total of 76 regime changes.

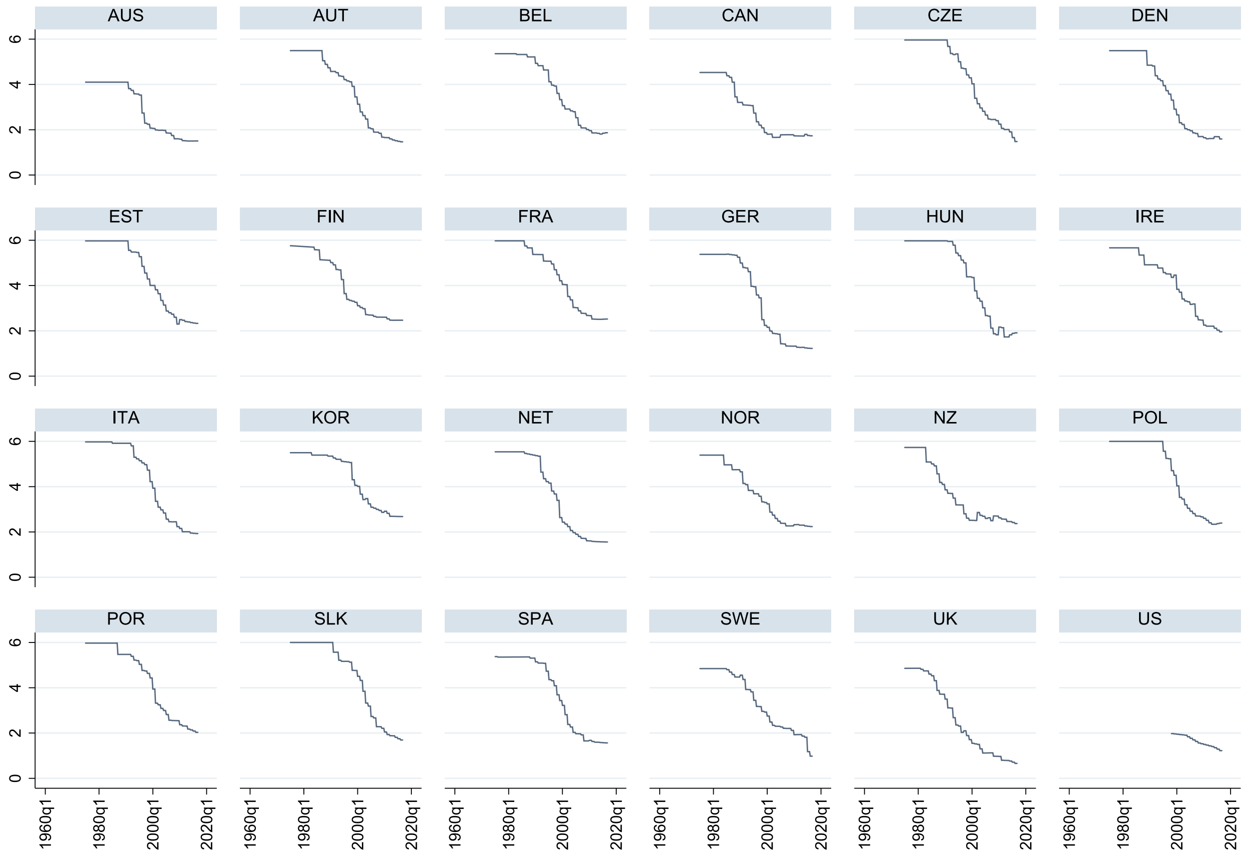

Fig. 2. ETCR indicator for product market competition (index: 0–6). **Note:** The figure shows the evolution of the OECD ETCR product market regulation indicator - a measure of goods market competition. A decrease in the ETCR indicates more competition.

(ETCR). This aggregate indicator is based on the average of the regulation indicators for 7 network sectors (electricity, gas, telecom, post, rail, airlines, and road). The indicators combine restrictiveness measures of regulation for 4 categories (entry, public ownership, vertical integration, and market structure) which scale from 0 to 6 (from least to most restrictive). Barker (2010, chapter 5), compares the ETCR against other proxies for product market competition and argues in its favor. This indicator of product market regulation is available on an annual frequency from 1975 to 2013. We extended this series to 2016 using the highly correlated Fraser Institute's "starting a business" series, after which we interpolated it to a quarterly frequency. See Appendix A for more details on the data series.

An increase in the indicator hence reflects more restrictive regulation and therefore a decrease in competitiveness. The indicator shows a similar evolution over time across countries and reflects a general decrease in product market regulation since the mid-1980s (Fig. 2). Although there is some variation in the level of the regulation indicator across countries, the indicator is more than halved by the end of the sample in each country relative to the 1980s.

Wage bargaining institutions The wage bargaining coordination variable (COORD) is an indicator obtained from the ICTWSS database (Visser, 2019) and ranges from 1 (decentralized) to 5 (highly centralized). Fig. 3 shows the evolution over time in the coordination measure for the countries in our sample. The indicator varies strongly across countries, reflecting substantial cross-sectional variation. At the same time, there is a difference in the extent of time variation across countries, with about half of the countries displaying sizable variation over time while the remainder of the countries show a large degree of persistence.

The union coverage rate (COV) is sourced from the same dataset and gives the percentage of employees covered by collective bargaining agreements of all employed wage and salary earners where an adjustment is made for the possibility that some sectors or occupations are excluded from the right to bargain. The union coverage rate (Fig. 4) in general moves gradually over time with both up- and downward trends, although the coverage rate remains stable at high levels for some countries. Taken together, the indicators of union power in wage bargaining depict sizable cross-sectional variation while the extent of time variation differs across the countries under analysis.

3.4. Results

Table 3 shows the estimates of Eq. (3). The upper part of the table shows the coefficients belonging to interactions between past inflation and a range of structural characteristics. In the lower part of the table, we report the average degree of indexation in each

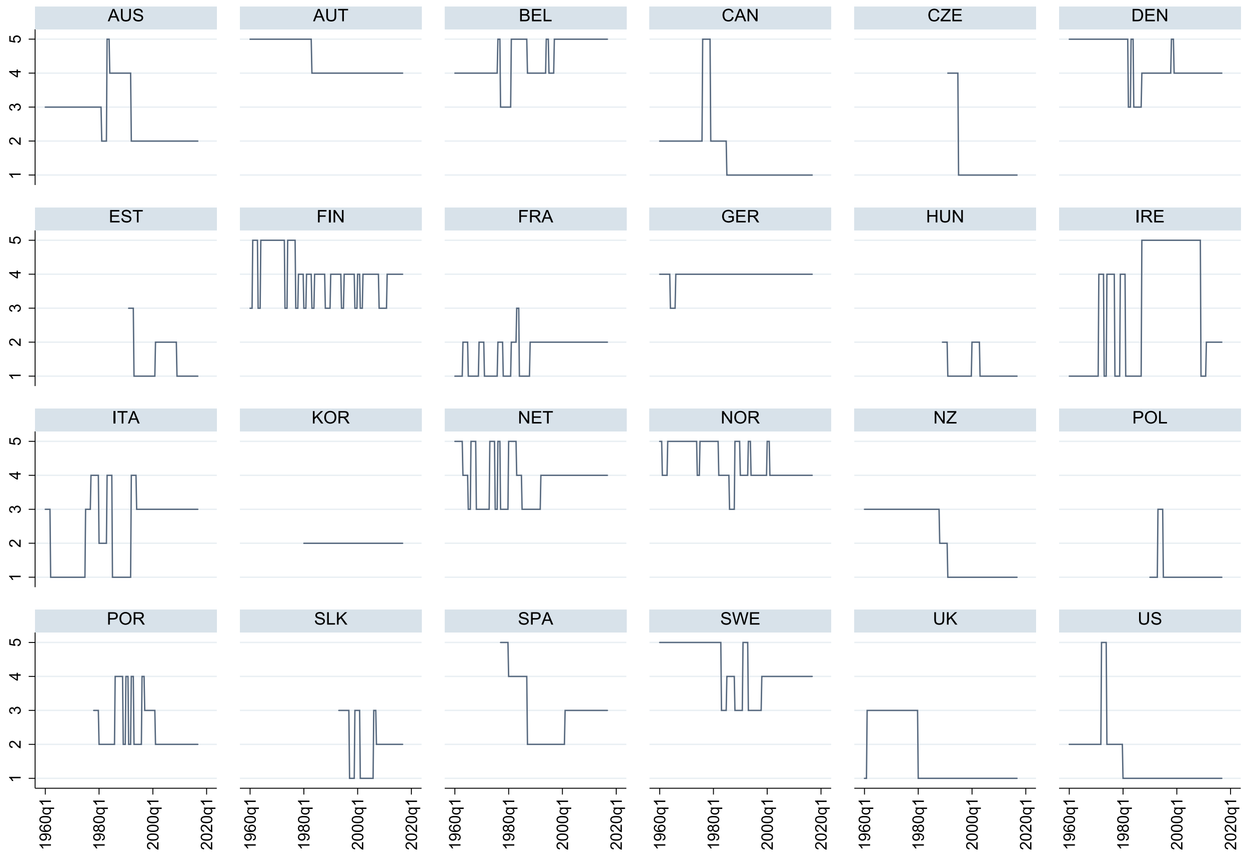

Fig. 3. Degree of wage bargaining coordination (index: 1–5).

monetary target regime. For columns 3 and 4, where the interaction effects are not confined to the monetary policy regime dummies, the average degree of indexation is calculated while the product market competition (ETCR) and wage bargaining indicators are set to their sample means.[17] The first column indicates that having a monetary target, irrespective of the type of target, does not significantly affect the degree of wage indexation. This result remains unchanged when controlling for other characteristics (column 2).

However, a closer look at the effect of an inflation, money growth, or exchange rate target, provided in column 3, shows that the type of target matters. Specifically, monetary targeting regimes with an explicit inflation target have a strong effect on the degree of wage indexation, as the interaction effect implies a reduction of wage indexation of 0.57 for an inflation targeting regime compared to a regime without a monetary policy target. Wage indexation to past inflation remains statistically significant under an inflation targeting regime but at a low value of 0.16. A money growth targeting strategy, by contrast, has a significant upward effect on the wage indexation coefficient.[18] The wage indexation parameter increases in this case to 0.97. For an exchange rate targeting strategy, we observe a substantially more muted reduction in indexation of 0.02, which is not statistically significant.

Columns 2 and 4 indicate that, in line with the theoretical predictions, an increase in product market competition (ETCR) significantly lowers wage indexation. Wage bargaining characteristics, by contrast, do not significantly alter the degree of indexation. Over both specifications, the product market regulation effect is around 0.22, which means that a decrease in competition (rising ETCR) has large positive effects on the degree of wage indexation. For the sample means of the regulation indicator (3.37), union coverage (0.66), and coordination (3), the average indexation coefficient becomes 0.33.

More important, however, is that the effect of the inflation target dummy, and the effects of the monetary policy target dummies more generally, vanish with the control for the additional drivers of wage indexation. A closer inspection of the model shows that this change is solely due to the inclusion of the product market regulation variable.

Table 4 shows that taking account of potential endogeneity between wage inflation and, respectively, lagged price inflation and the unemployment gap leads to very similar estimation results.

Discussion

Our finding of the importance of goods market competition for wage indexation corroborates the results from Duca and

[17] Total indexation equals $\partial E[\pi^{w}_{i,t}|\bar{\pi}^{p}_{i,t-1}, ...]/\partial \bar{\pi}^{p}_{i,t-1} = \gamma + \tilde{\gamma}' D_{i,t}$. The standard errors are computed with the Delta method.

[18] This effect is sensitive to the time period under analysis (see Appendix B).

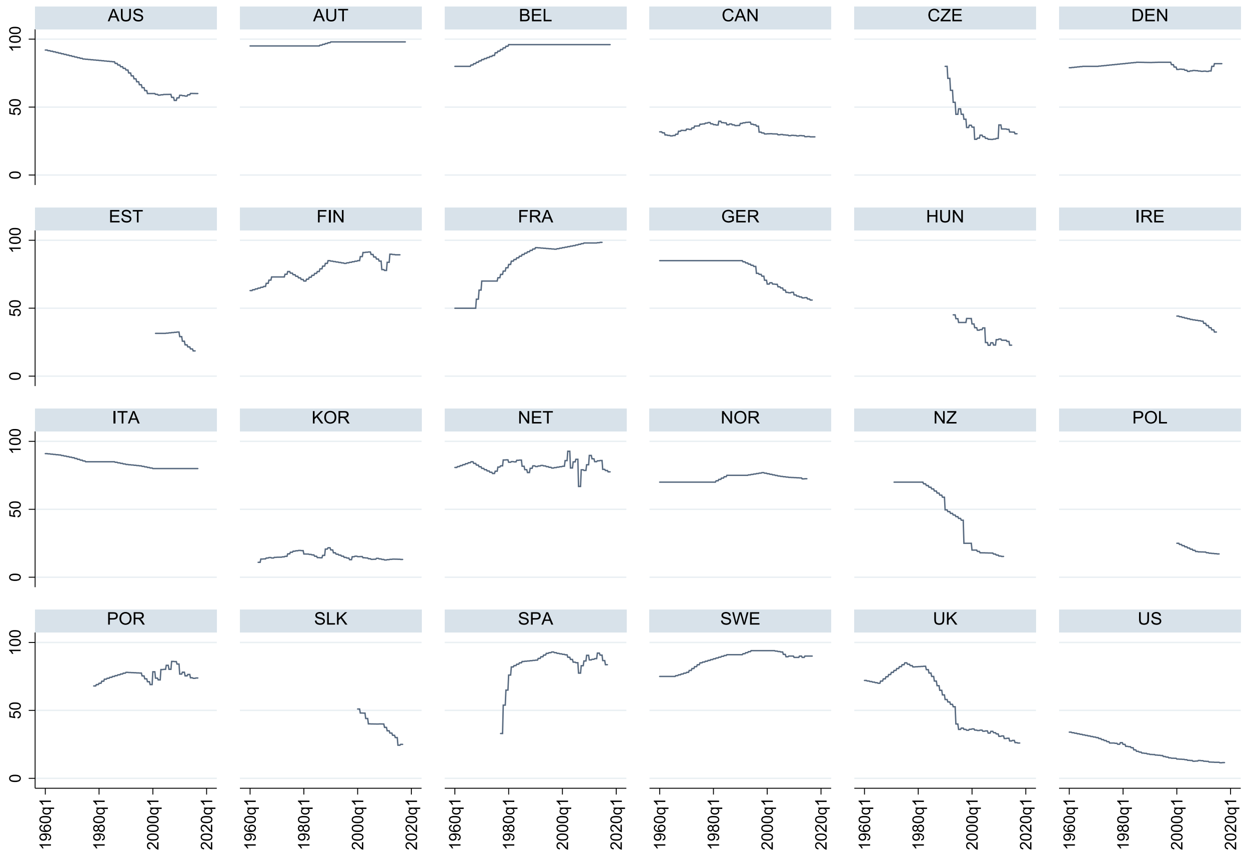

Fig. 4. Union coverage rate (percentage). **Note:** Figs. 3 and 4 respectively show the evolution of the ICTWSS wage bargaining coordination and union coverage indicators.

VanHoose (1998) for a broad set of countries and a longer time span that includes inflation-targeting central banks. We show in Appendix B that this competition effect withstands a broad set of robustness checks, each time remaining economically and statistically significant. Our results suggest that monetary policy is not a crucial driver of wage indexation, which is consistent with what Carrillo et al. (2017) found for the US.

However, one needs to be careful in the interpretation of these results given the similar time variation in product market competition and the shift to inflation targets. Fig. 5 shows that the shift towards more product market competition - a decline in ETCR - is indeed closely associated with the adoption of inflation targets.[19] Although the competition indicators generally started to decline before the first adoption of an inflation target in 1990, their descent picked up pace in the 1990s when more countries adopted an inflation target. Our results are hence consistent with insignificant effects of monetary target regimes, but they are also consistent with the view that it is challenging to disentangle monetary policy effects from product market competition effects.

4. Conclusions

We have examined the standard assumption in New Keynesian DSGE models that wage indexation to past price inflation is invariant to policy regimes. In particular, we have estimated the reduced form empirical wage Phillips curve of Galí (2011) with a panel model for 24 advanced economies and allowed the degree of backward-looking wage indexation to vary according to the monetary policy regime while controlling for the degree of product market competition and wage bargaining characteristics

When we allow wage indexation to vary depending on the monetary target, we find an economically and statistically significant reduction of wage indexation for countries with an inflation target. Specifically, the degree of wage indexation is substantially lower in an inflation target regime compared to a regime without a target or with a monetary or exchange rate target. However, when we include structural characteristics on wage bargaining and the degree of product market competition, the effect of inflation targeting essentially vanishes. More precisely, we find robust evidence that a higher degree of product market competition lowers the degree of

[19] Fig. 5 is a binscatter plot (Stepner, 2013) between ETCR and the inflation targeting dummy, which works as follows. All ETCR observations, sorted from small to large, are divided into 30 bins. Next, the average ETCR and corresponding IT values are calculated for each bin and, finally, the figure shows the scatterplot of these averages and the fitted regression line.

Table 3
Results interaction model.

Regressand:	Wage inflation ($\pi_{i,t}^{w}$)			
	(1) FE	(2) FE	(3) FE	(4) FE
Interactions with $\bar{\pi}_{i,t-1}^{p}$:				
Target	0.105 (0.175)	−0.005 (0.129)		
IT			−0.568*** (0.150)	−0.066 (0.207)
MON			0.239*** (0.081)	0.131 (0.080)
ER			−0.021 (0.115)	−0.120 (0.097)
ETCR		0.220*** (0.037)		0.227*** (0.062)
COORD		−0.021 (0.027)		0.017 (0.025)
COV		0.088 (0.196)		−0.133 (0.131)
Total indexation effect [ETCR, COORD and COV at means]:				
Absence of quantitative TARGET	0.731*** (0.161)	0.297*** (0.104)	0.729*** (0.112)	0.327*** (0.117)
Presence of quantitative TARGET	0.835*** (0.062)	0.292*** (0.058)		
Inflation targeting regime (IT)			0.160** (0.067)	0.261** (0.132)
Monetary targeting regime (MON)			0.968*** (0.162)	0.458*** (0.163)
Exchange rate targeting regime (ER)			0.708*** (0.051)	0.207* (0.116)

Note: Symbols *, **, and *** denote significance at 10%, 5%, and 1% levels, respectively. Standard errors are in brackets (robustified to heteroskedasticity and serial correlation within cross sections for the pooled estimators). Sample: $\overline{T}=123.4$, max $T=168$, min $T=60$ and N = 24.

Table 4
Results IV regression with interaction effects.

Regressand:	Wage inflation ($\pi_{i,t}^{w}$)			
	(1) IV	(2) IV	(3) FE_{homo}	(4) FE_{homo}
Interactions with $\bar{\pi}_{i,t-1}^{p}$:				
IT	−0.422*** (0.135)	−0.073 (0.169)	−0.571*** (0.150)	−0.250 (0.164)
MON	0.273*** (0.063)	0.204*** (0.077)	0.257*** (0.076)	0.171** (0.080)
ER	0.067 (0.090)	−0.015*** (0.100)	0.006 (0.119)	−0.087 (0.096)
ETCR		0.191*** (0.052)		0.171*** (0.056)
COORD		0.039 (0.031)		0.020 (0.029)
COV		−0.255 (0.235)		−0.170 (0.179)
Total indexation effect [ETCR, COORD and COV at means]:				
Absence of quantitative TARGET	0.658*** (0.093)	0.321*** (0.118)	0.717*** (0.113)	0.427*** (0.111)
Inflation targeting regime (IT)	0.236** (0.106)	0.248** (0.117)	0.146** (0.059)	0.176* (0.101)
Monetary targeting regime (MON)	0.931*** (0.084)	0.525*** (0.115)	0.974*** (0.167)	0.598*** (0.165)
Exchange rate targeting regime (ER)	0.725*** (0.044)	0.306*** (0.113)	0.724*** (0.047)	0.340*** (0.097)

Note: Symbols *, **, and *** denote significance at 10%, 5%, and 1% levels, respectively. Standard errors are in brackets (robustified to heteroskedasticity and serial correlation within cross sections for the pooled estimators). Instrumental variables: second lag of the unemployment gap for the current unemployment gap and the second lag of smoothed price inflation for price inflation. The first stage IV F-statistics (Kleibergen-Paap RK Wald weak identification test statistic) equal 23.24 and 12.65 for columns (1) and (2). "FE_{homo}" refers to the FE estimates with homogeneous coefficients on the unemployment variables which serve as the relevant benchmark for the IV estimators for which homogeneous coefficients are imposed. Sample: $\overline{T}=123.4$, max $T=168$, min $T=60$ and N = 24.

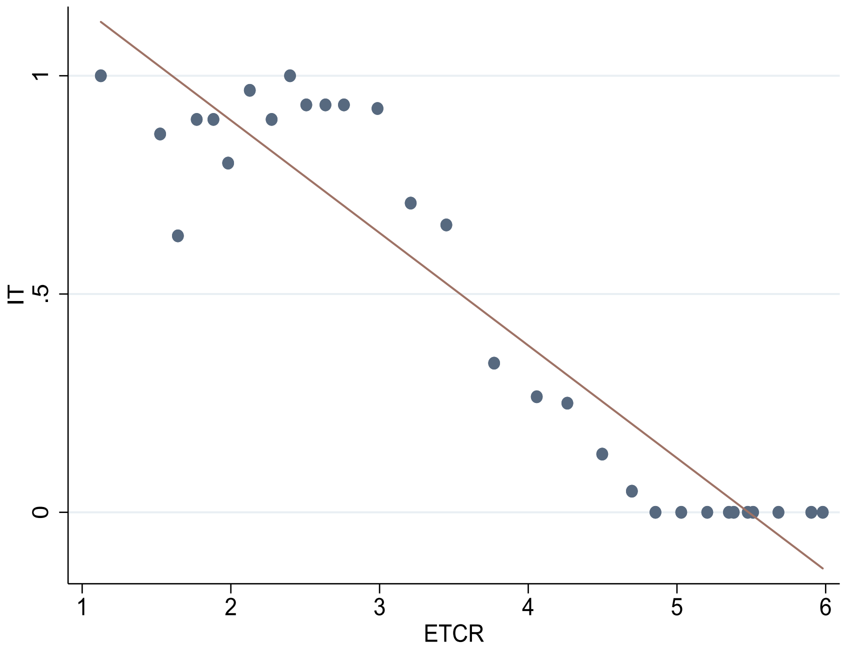

Fig. 5. The association between product market competition and inflation targeting. **Note:** The figure shows a binscatter plot (see Stepner, 2013) of the degree of goods market competition (ETCR) against the inflation targeting dummy.

wage indexation in an economically and statistically significant way. Therefore, we generalize the results from Duca and VanHoose (1998) for a broad set of countries and a longer time span.

Since product market competition is an outcome of economic policy, we conclude that the constant indexation assumption embedded in standard DSGE models is susceptible to the Lucas (1976) critique, i.e., it is not intrinsic to the deep structure of the economy and not a policy invariant parameter. This is relevant for policymakers, given the macroeconomic importance of wage indexation on output and inflation.

Our findings suggest that monetary policy is not a crucial driver of wage indexation. Although this is consistent with the result of Carrillo et al. (2017) for the US, we caution against interpreting this result as conclusive evidence that monetary policy is irrelevant for wage indexation. Since the evolution of product market competition correlates with the adoption of inflation targeting, it could be that it is challenging to disentangle monetary policy effects from product market competition effects. One avenue for future research is to find ways to resolve this issue.

Appendix A. Data sources and construction of variables.

A1. Macroeconomic series

Hourly earnings: average total earnings in manufacturing paid per employee per hour, seasonally adjusted [OECD, Main Economic Indicators database, quarterly frequency]. We compute quarter-on-quarter wage inflation $\pi^w_{i,t}$ as $100LN\,(P^w_{i,t}/P^w_{i,t-1})$, where $LN(.)$ stands for the natural logarithm and $P^w_{i,t}$ for the nominal wage index of country i in quarter t.[20]
Note: The earnings series constitutes wage rates plus overtime payments, bonuses and gratuities regularly and irregularly paid, remuneration for time not worked, and payments in kind. Not included are employer contributions to social security or insurance schemes and unfunded employee social benefits paid by employers. Adding these components to the earnings series delivers compensation rates.
Unemployment rate: the unemployment rate is defined as the ratio of the number of unemployed workers to the working population [OECD, Economic Outlook No.104 database, quarterly frequency].
Note: The German series before 1992 was extended backward based on the growth rates of unemployment rate as percentage of the civilian labor force for West-Germany (source: Bundesbank, series BBK01.USCY01).
Unemployment gap: The unemployment gap is the difference between the unemployment rate and the NAIRU, expressed as percentages. The former is described above. Concerning the latter, we combined several OECD sources to obtain long NAIRU series for each country. The NAIRU series from 1990Q1 onwards is obtained from the Economic Outlook No.104 database. We are

[20] We initially started from a data sample of 29 OECD countries. Due to data availability of the labor market indicators, three countries were dropped (Iceland, Israel and Mexico). We further eliminated Japan and Luxembourg from the sample as these countries are characterized by severe outliers concerning the coefficients on the unemployment rate variables.

Table 5
Overview number of observations per country.

country	time span		#	country	time span		#
Australia	1984Q1	2016Q4	132	Italy	1975Q1	2016Q4	168
Austria	1975Q1	2016Q4	168	South Korea	1992Q2	2016Q4	99
Belgium	1975Q1	2016Q4	168	Netherlands	1975Q1	2016Q4	168
Canada	1975Q1	2016Q4	168	New Zealand	1989Q2	2011Q4	91
Czech Republic	1995Q2	2016Q4	87	Norway	1975Q1	2014Q4	160
Denmark	1975Q1	2016Q4	168	Poland	2000Q1	2015Q4	64
Estonia	2001Q1	2015Q4	60	Portugal	2000Q2	2016Q4	67
Finland	1975Q1	2015Q4	164	Slovak Republic	2000Q1	2016Q4	68
France	1975Q1	2014Q4	160	Spain	1981Q2	2016Q4	143
Germany	1990Q2	2016Q4	107	Sweden	1975Q1	2016Q4	168
Hungary	1995Q2	2014Q4	79	United Kingdom	1975Q1	2016Q4	168
Ireland	2000Q1	2014Q4	60	United States	1998Q1	2016Q4	76

Note: We have an unbalanced panel with different starting and ending dates for some countries. The benchmark sample is 1975Q1-2016Q4, i.e. the longest possible sample while including the product market competition indicator.

grateful to Elena Rusticelli for providing us with these series at a quarterly frequency. To extend this series back in time, we backdated the NAIRU series from the 104 Economic Outlook with the NAIRU series from the Economic Outlook No. 98 database. This historical NAIRU data was interpolated from an annual to a quarterly frequency with Eviews by using a quadratic polynomial (with the constraint that the average of the quarterly values in a year equals the annual value). Finally, the most recent NAIRU series was backdated using the historical NAIRU series using growth rates.

Price inflation: consumer price index, all items, non-seasonally adjusted [OECD, Main Economic Indicators database, quarterly frequency]. The price indices were seasonally adjusted using the X13-Arima procedure in Jdemetra+. Quarter-on-quarter price inflation $\pi_{i,t}^{p}$ is defined as $100LN\,(P_{i,t}/P_{i,t-1})$, where $P_{i,t}$ denotes the price index of country i in quarter t. Our main results are based on smoothed (year-on-year) past inflation $\pi_{i,t-1}^{p} = \frac{1}{4}(\pi_{i,t-1}^{p} + \pi_{i,t-2}^{p} + \pi_{i,t-3}^{p} + \pi_{i,t-4}^{p})$.

Trend inflation: To proxy long-term expectations for each country, we estimate trend inflation rates $\pi_{i,t}^{*}$ with the AR-Trend-bound model of Chan et al. (2013). The details of the estimation procedure are given in Appendix C.

A2. Monetary targets

Monetary policy quantitative target dummies: 0–1 dummy variables that indicate whether a country has a formal inflation, exchange rate, or monetary target in the respective period. If an exchange rate regime is classified as having a managed or freely floating exchange rate, we consider it to have no formal target [quarterly frequency].

Our main sources for the classification of quantitative monetary targets are Fatás et al. (2007) and Houben (2000). An additional source for inflation targeting regimes is Rose (2007); for the exchange rate and monetary targeting regimes, it is Borio and White (2003). Table 6 presents the country-specific overview of the different monetary regimes.

Note: Conflicting dates were examined, and remaining gaps were filled based on central banks' websites, individual central bank reports and, where necessary, additional publications.

A3. Product market competition indicator

Our measure of product market competition is the OECD indicator of regulation in non-manufacturing sectors (NMR), more specifically the aggregate indicator of regulation in energy, transport and communications (ETCR) [OECD, Dataset: Regulation in energy, transport and communications 2013, annual frequency 1975–2013]. This aggregate indicator is based on the average of the regulation indicators for 7 network sectors (electricity, gas, telecom, post, rail, airlines, and road). The indicators combine restrictiveness measures of regulation for 4 categories (entry, public ownership, vertical integration, and market structure) which scale from 0 to 6 (from least to most restrictive). This series is the only OECD product market competitiveness indicator from this widely used database that is available on an annual frequency and for a relatively long period (1975–2013). For instance, Barker (2010, chapter 5), compares the ETCR against other proxies for product market competition and argues in its favour.

We use a regression model to extend the ETCR series to 2016. More specifically, we found that the ETCR indicator is highly correlated (-0.72 on average) with the Fraser Institute's "starting a business" indicator, which is available for 2000–2016. Therefore, as a first step, we regressed each country's ETCR indicator on its first lag and the (contemporaneous) "starting a business" indicator for the period 2000–2013.[21] As a second step, we predicted the ETCR values for 2014–2016 and, finally, we converted the annual series to a quarterly frequency by keeping the annual value constant over the four quarters within the year.

[21] In some cases, we added additional right-hand-side variables to the regression equation in order to obtain well-behaved residuals. Based on the BIC criterion, we included either more lags of the ETCR, lags of the "starting a business" indicator, a time trend, or a combination of them.

Table 6
Overview of periods when a monetary policy target was in place.

Country	Type of target			
	Exchange rate	Money growth	Inflation	None
Australia	/	1984Q1-1985Q1	1993Q2-2016Q4	1985Q2-1993Q1
Austria	1975Q1-1998Q4	1999Q1-2016Q4	1999Q1-2016Q4	/
Belgium	1975Q1-1998Q4	1999Q1-2016Q4	1999Q1-2016Q4	/
Canada	/	1975Q1-1982Q3	1991Q1-2016Q4	1982Q4-1990Q4
Czech Republic	1995Q2-1997Q2	1995Q2-1997Q4	1998Q1-2016Q4	/
Denmark	1975Q1-2016Q4	/	/	/
Estonia	2001Q1-2010Q4	2011Q1-2015Q4	2011Q1-2015Q4	/
Finland	1975Q1-1992Q3	1999Q1-2015Q4	1993Q1-2015Q4	1992Q4
	1996Q4-1998Q4			
France	1975Q1-1998Q4	1976Q4-2014Q4	1999Q1-2014Q4	/
Germany	1990Q2-1998Q4	1990Q2-2016Q4	1999Q1-2016Q4	/
Hungary	1995Q2-2008Q1	/	2001Q3-2014Q4	/
Ireland	/	2000Q1-2014Q4	2000Q1-2014Q4	/
Italy	1979Q2-1992Q2	1975Q1-2016Q4	1999Q1-2016Q4	/
	1996Q4-1998Q4			
South Korea	1992Q2-1998Q1	1992Q2-2002Q4	1998Q2-2016Q4	/
Netherlands	1975Q1-1998Q4	1977Q2-1980Q4		
		1999Q1-2016Q4	1999Q1-2016Q4	/
New Zealand	/	/	1990Q2-2011Q4	1989Q2-1990Q1
Norway	1975Q1-1992Q4	/	2001Q2-2014Q4	1993Q1-1994Q1
	1994Q2-2001Q1			
Poland	2000Q1	/	2000Q1-2015Q4	/
Portugal	/	2000Q2-2016Q4	2000Q2-2016Q4	/
Slovak Republic	2005Q4-2008Q4	2009Q1-2016Q4	2005Q1-2016Q4	2000Q1-2004Q4
Spain	1989Q3-1998Q4	1981Q2-1989Q2	1995Q1-2016Q4	/
		1999Q1-2016Q4		
Sweden	1975Q1-1992Q4	/	1993Q1-2016Q4	/
United Kingdom	1990Q4-1992Q3	1976Q1-1992Q3	1992Q4-2016Q4	1975Q1-1975Q4
United States	/	/	2012Q1-2016Q4	1998Q1-2011Q4

Notes: 1) The table shows for each country in which periods a certain monetary policy target was in place. We consider exchange rate target, money growth target and inflation targets, and we define a period without a formal target as one where none of the forementioned three are present. 2) Finland, Italy and Norway have two separate periods with an exchange rate; The Netherlands and Spain have two periods with a money growth target. 3) We have an unbalanced panel with different starting and ending dates for some countries. The benchmark sample is 1975Q1-2016Q4, i.e. the longest possible sample while including the product market competition indicator.

A4. Wage bargaining indicators

Our two wage bargaining indicators are sourced from the ICTWSS version 6.0. Database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts, 1960–2017 (Visser, 2019).

Coordination: an indicator of the degree of wage coordination [ICTWSS codename: Coord] which ranges from 1 to 5. The ICTWSS codebook describes the values as follows:

5 = Binding norms regarding maximum or minimum wage rates or wage increases issued as a result of a) centralized bargaining by the central union and employers' associations, with or without government involvement, or b) unilateral government imposition of wage schedule/freeze, with or without prior consultation and negotiations with unions and/or employers' associations.

4 = Non-binding norms and/or guidelines (recommendations on maximum or minimum wage rates or wage increases) issued by a) the government or government agency, and/or the central union and employers' associations (together or alone), or b) resulting from an extensive, regularized pattern setting coupled with high degree of union concentration and authority.

3 = Procedural negotiation guidelines (recommendations on, for instance, wage demand formula relating to productivity or inflation) issued by a) the government or government agency, and/or the central union and employers' associations (together or alone), or b) resulting from an extensive, regularized pattern setting coupled with high degree of union concentration and authority.

2 = Some coordination of wage setting, based on pattern setting by major companies, sectors, government wage policies in the public sector, judicial awards, or minimum wage policies.

1 = fragmented wage bargaining, confined largely to individual firms or plants, no coordination

[Refers to the private sector, across branches or sectors; based on Kenworthy (2001b), Kenworthy (2001a). Note that this is an indicator of the "degree, rather than the type, of coordination" (Kenworthy, 2001b:78), (...) "based on a set of expectations about which institutional features of wage setting arrangements are likely to generate more or less coordination" (2001a:80).]

We converted this annual series to a quarterly frequency by keeping the annual value constant over the four quarters within the year.

Coverage: Adjusted bargaining (or union) coverage rate [ICTWSS codename: AdjCov]. This variable ranges from 0 to 100 and measures "employees covered by valid collective (wage) bargaining agreements as a proportion of all wage and salary earners in employment with the right to bargaining, expressed as percentage, adjusted for the possibility that some sectors or occupations are excluded from the right to bargain." For this series, we first imputed the missing values between observations through linear interpolation. Then, we converted the annual series to a quarterly frequency by keeping the annual value constant over the four quarters within the year.

Appendix B. Robustness checks

Section 3 shows that having an explicit inflation target seems to reduce the extent of wage indexation in OECD countries unless one explicitly controls for the influence of product market competition. In this section, we show that these results withstand a battery of robustness checks.

First, we find that our benchmark results are robust to *i)* replacing the unemployment gap measure $\hat{u}_{i,t}$ with the unemployment rate $u_{i,t}$, and *ii)* using past inflation as the indexation measure ($\bar{\pi}^p_{t-1} \equiv \pi^p_{t-1}$). The interaction model results are further robust to *i)* the inclusion of time fixed effects, as well as *ii)* controlling for the level of inflation, the duration of the regimes, the combination of different monetary targets, and *iii)* different classifications of IT regimes (i.e. discriminating between loose and strict inflation targets as in Samarina et al. (2014) and between range and point inflation targets based on the classification in Apel and Claussen (2017)). In what follows, we discuss other checks in more detail.

Controlling for trend inflation

One robustness check involves the extension of the benchmark wage inflation model (2) with a time-varying measure of trend inflation, $\pi^*_{i,t}$. A time-varying trend inflation rate is commonly interpreted as the inflation target of the central bank and a reflection of long-term inflation expectations. Its inclusion enables us to check whether the degree of wage indexation depends on the dynamics in trend inflation. Cogley and Sbordone (2008), for example, find that the estimated degree of price indexation becomes small and statistically insignificant when they extend a standard DSGE model with a time-varying trend inflation rate.[22] What is more, the inclusion allows us to account for the forward-looking behavior of economic agents.

In contrast to the benchmark wage inflation model (2), where the intercept contains a constant trend inflation, the modified model specification has trend inflation as a right hand side variable:

$$\tilde{\pi}^w_{i,t} = \alpha_i + \gamma\tilde{\pi}^p_{i,t-1} + \psi_0\hat{u}_{i,t} + \psi_1\hat{u}_{i,t-1} + \eta_{i,t}, \tag{4}$$

where $\tilde{\pi}^w_{i,t} \equiv \pi^w_{i,t} - \pi^*_{i,t}$ and $\tilde{\pi}^p_{i,t-1} \equiv \bar{\pi}^p_{i,t-1} - \pi^*_{i,t}$.[23] We transform the interaction effects model with heterogeneous unemployment gap coefficients (Eq. (3)) in the same way. We employ a smoothed trend based on the work of Chan et al. (2013), henceforth CKP, as a measure of trend inflation $\pi^*_{i,t}$. Appendix C outlines the estimation procedures of this trend measure.[24]

The estimated wage indexation effects are similar to those of Table 3 although somewhat smaller in size. The effects of regimes with an inflation or monetary target remain statistically significant (Table 7, column 2) and again vanish once product market regulation and the wage bargaining characteristics are taken into account (column 3). From this, we conclude that the results are robust to the inclusion of time-varying trend inflation and hence to the inclusion of long-term forward-looking inflation expectations.

Longer sample period While we keep the estimation sample fixed over the different model specifications to facilitate a correct comparison of the coefficients, the time period is confined to 1975Q1-2016Q4 due to the availability of the product market competition indicator. Given the data availability for the variables in the benchmark NKWPC and the monetary regime indicators, we can however enlarge the time period to 1960Q1-2016Q4 for the interaction specifications without the inclusion of the wage bargaining and product market competition proxies. When we include the observations starting in 1960 (Table 7), the benchmark results are not materially affected. For the interaction model, the monetary target effect however becomes much smaller and is insignificantly different from zero. This points to the importance of picking up the increased use of monetary targets since the mid-1970s relative to the years prior to 1975 in measuring the effect of a monetary targeting strategy. The effect of the inflation target interaction is instead highly similar.

Time effects Inflation targets were adopted since 1990 and they make up for a large part of the regime shifts in our sample (see Fig. 1). Similarly, product markets have witnessed a general deregulation trend over time. While the results are robust to the

[22] Cogley and Sbordone (2008) extend the Calvo pricing model in a standard DSGE model to incorporate drifts in trend inflation and derive a reduced-form New Keynesian Phillips Curve (NKPC) with time-varying coefficients. This reduced form NKPC is estimated on US data using a Bayesian time-varying VAR. Under the assumption that non-optimized prices are fully indexed to a mixture of current trend inflation and one-period lagged inflation, the reduced form NKPC collapses to a more traditional NKPC based on the inflation gap with constant coefficients and without additional forward-looking terms. The latter specification is similar to our *benchmark wage gap model* (4).

[23] Eq. (4) is obtained by assuming that the coefficients on past and trend inflation (γ and γ^*) in $\pi^w_{i,t} = \alpha_i + \gamma\bar{\pi}^p_{i,t-1} + \gamma^*\pi^*_{i,t} + \psi_0\hat{u}_{i,t} + \psi_1\hat{u}_{i,t-1} + \eta_{i,t}$ sum to one. This hypothesis cannot be rejected at the 5% level for each of the 3 panel estimators. Note that we implicitly assume in (4) that $\pi^w_{i,t}$ has no immediate effect on $\pi^*_{i,t}$. We find this assumption to be reasonable because our measure of trend inflation captures inflation expectations at an infinite horizon.

[24] We apply robust standard errors which are consistent in the presence of unknown forms of autocorrelation and heteroskedasticity. Note however that we treat trend inflation as a known variable such that the potential generated regressor bias of the standard errors (Pagan, 1984) is not taken into account.

Table 7
Robustness checks for trend inflation and full sample.

Regressand:	Wage inflation ($\pi_{i,t}^{w}$)				
	Trend inflation			Full sample	
	(1)	(2)	(3)	(4)	(5)
Unemployment gap ($\hat{u}_t$)	−0.344*** (0.094)			−0.407*** (0.090)	
Lagged unemployment gap ($\hat{u}_{t-1}$)	0.290*** (0.092)			0.307*** (0.085)	
Lagged price inflation ($\bar{\pi}_{i,t-1}^{p}$)				0.892*** (0.044)	
Lagged price inflation gap ($\tilde{\pi}_{i,t-1}^{p}$)	0.684*** (0.147)				
Interactions with $\bar{\pi}_{i,t-1}^{p}$:					
IT		−0.611*** (0.206)	0.253 (0.255)		−0.544*** (0.185)
MON		0.389** (0.174)	0.208 (0.153)		0.036 (0.089)
ER		−0.034 (0.142)	−0.171 (0.103)		−0.084 (0.116)
ETCR			0.310*** (0.073)		
COORD			0.034 (0.041)		
COV			0.248 (0.328)		
Total indexation effect [ETCR, COORD and COV at means]:					
Absence of quantitative TARGET		0.578*** (0.146)	0.085 (0.135)		0.882*** (0.121)
Inflation targeting regime (IT)		−0.033 (0.130)	0.338* (0.174)		0.338*** (0.130)
Monetary targeting regime (MON)		0.967*** (0.233)	0.293 (0.193)		0.918*** (0.126)
Exchange rate targeting regime (ER)		0.543*** (0.124)	−0.086 (0.147)		0.798*** (0.062)

Note: Columns 1–3 show the results for the model that takes trend inflation into account. Columns 4–5 show results on the full sample which also includes pre-1975 data (ETCR starts in 1975). Symbols *, **, and *** denote significance at 10%, 5%, and 1% levels, respectively. Standard errors are in brackets (robustified to heteroskedasticity and serial correlation within cross sections for the pooled estimators).

Table 8
Robustness checks for time effects.

Regressand:	Wage inflation ($\pi_{i,t}^{w}$)			
	Great Moderation dummy		Pre-2007 sample	
	(1)	(2)	(3)	(4)
Interactions with $\bar{\pi}_{i,t-1}^{p}$:				
Great moderation dummy	−0.157*** (0.048)	−0.130*** (0.043)		
IT	−0.575*** (0.151)	−0.062 (0.207)	−0.511*** (0.171)	−0.278 (0.192)
MON	0.208** (0.078)	0.097 (0.076)	0.237** (0.100)	0.147 (0.122)
ER	0.009 (0.124)	−0.106 (0.105)	0.039 (0.124)	−0.070 (0.099)
ETCR		0.236*** (0.062)		0.186** (0.084)
COORD		0.010 (0.026)		0.025 (0.033)
COV		−0.149 (0.154)		−0.222 (0.146)
Total indexation effect [ETCR, COORD and COV at means; GM dummy = 1]:				
Absence of quantitative TARGET	0.627*** (0.117)	0.247** (0.113)	0.706*** (0.112)	0.419*** (0.116)
Inflation targeting regime (IT)	0.052*** (0.076)	0.186 (0.139)	0.195* (0.112)	0.141 (0.148)
Monetary targeting regime (MON)	0.835*** (0.172)	0.344** (0.159)	0.943*** (0.177)	0.566*** (0.217)
Exchange rate targeting regime (ER)	0.636*** (0.037)	0.142 (0.111)	0.745*** (0.045)	0.350** (0.158)

Note: Columns 1–2 show results for the model with a Great Moderation dummy. Columns 3–4 show results for the pre-Great Financial Crisis sample. Symbols *, **, and *** denote significance at 10%, 5%, and 1% levels, respectively. Standard errors are in brackets (robustified to heteroskedasticity and serial correlation within cross sections for the pooled estimators). Sample: $\overline{T} = 123.4$, max $T = 168$, min $T = 60$ and N = 24.

Table 9
Robustness checks for nonlinearities.

Regressand:	Wage inflation ($\pi^w_{i,t}$)			
	Spline function		Thresholds	
	(1)	(2)	(3)	(4)
Interactions with $\hat{u}_t$:				
I(Unemployment gap ($\hat{u}_t$) > 0)	0.229 (0.156)	0.192 (0.138)		
Lower threshold			−0.167 (0.187)	−0.164 (0.182)
Upper threshold			0.095** (0.039)	0.099** (0.040)
Interactions with $\hat{u}_{t-1}$:				
Lower threshold			0.163 (0.191)	0.106 (0.172)
Upper threshold			−0.078 (0.049)	−0.053 (0.051)
Interactions with $\bar{\pi}^p_{i,t-1}$:				
IT	−0.591*** (0.149)	−0.264 (0.159)	−0.585*** (0.154)	−0.257 (0.162)
MON	0.259*** (0.078)	0.165** (0.078)	0.256*** (0.074)	0.162** (0.077)
ER	0.006 (0.115)	−0.100 (0.088)	0.006 (0.124)	−0.097 (0.101)
ETCR		0.181*** (0.055)		0.176*** (0.055)
COORD		0.016 (0.028)		0.016 (0.029)
COV		−0.147 (0.177)		−0.141 (0.181)
Total indexation effect [ETCR, COORD and COV at means]:				
Absence of quantitative TARGET	0.721*** (0.108)	0.417*** (0.102)	0.722*** (0.117)	0.421*** (0.114)
Inflation targeting regime (IT)	0.129** (0.063)	0.153 (0.101)	0.138** (0.059)	0.164* (0.097)
Monetary targeting regime (MON)	0.980*** (0.165)	0.581*** (0.157)	0.979*** (0.170)	0.583*** (0.167)
Exchange rate targeting regime (ER)	0.726*** (0.045)	0.317*** (0.094)	0.729*** (0.045)	0.324*** (0.095)

Note: Columns 1–2 show results for the specification with interaction terms on the sign of the unemployment gap. Columns 3–4 show results for the models with unemployment threshold effects. Symbols *, **, and *** denote significance at 10%, 5%, and 1% levels, respectively. Standard errors are in brackets (robustified to heteroskedasticity and serial correlation within cross sections for the pooled estimators). Sample: $\overline{T} = 123.4$, max $T = 168$, min $T = 60$ and N = 24.

inclusion of time fixed effects, we perform two additional tests to verify that we are not confounding the interaction effects with more specific breaks in either the effectiveness of monetary policy conduct or a shift in the economic environment.

First, we consider whether the results are robust to accounting for the Great Moderation, an era characterized by global declines in inflation volatility. We defined a time dummy changing value (from zero to one) between 1984Q1 and 2007Q4 to mimic the Great Moderation period and added it to the set of interaction variables $D_{i,t}$ in the estimation of (3). This dummy has a statistically significant effect of roughly -0.15 on the indexation parameter (columns 1 and 2 of Table 8) but leaves the interaction coefficients more or less unchanged.

Second, we estimated the interaction effects model on a sample with only pre-2007Q1 data to accommodate any concerns related to potential time variation in the coefficients resulting from the post Great Financial Crisis (GFC) period. Columns 3 and 4 of Table 8 show that the restriction to the pre-GFC sample also delivers similar results. Therefore, the interaction effects do not seem to confound with a time effect.

Nonlinearities to unemployment gap Some recent studies (Donayre and Panovska, 2016; Kumar and Orrenius, 2016) document the existence of nonlinearities in the relationship between wage growth and unemployment. To analyze the effects of such potential nonlinearities on the interaction effects, we take two approaches. First, we estimate a spline function which allows the effects of unemployment to vary according to whether the unemployment gap is positive or negative, similar to Kumar and Orrenius (2016). More specifically, we add an interaction term that multiplies the unemployment gap variables with an indicator variable being 1 if unemployment is above the natural rate and 0 otherwise.

Second, we estimate a three-regime threshold model following Donayre and Panovska (2016) but with exogenously determined threshold values being equal to the country-specific 15th and 85th percentiles of lagged unemployment. The (homogenous) slope of the NKWPC[25] is allowed to vary when the unemployment gap is below the lower threshold value or above the upper value. In both

[25] Homogeneity of the unemployment coefficients is imposed to avoid a very large amount of coefficient estimates. The same conclusion however holds with heterogeneous coefficients on the unemployment gap variables.

cases, the effects of the monetary regimes are not affected by the allowance for these nonlinearities (see Table 9).[26]

Appendix C. The AR-trend-bound model for trend inflation

For each country in our sample, we estimate trend inflation $\pi^*_{i,t}$ from quarter-on-quarter price inflation $\pi^p_{i,t} \equiv 100ln(P_{i,t}/P_{i,t-1})$ using the model of Chan et al. (2013). For details on the model, we refer the reader to their paper. We next discuss how our priors differ from theirs.

We impose the same set of generally uninformative priors for all countries. Relative to the setup of Chan et al. (2013), we make initial conditions more diffuse by setting $\omega^2_{\pi^*} = \omega^2_h = 25$ and set π^*_0 equal to the mean of the first four observations of the inflation series, if this figure is positive, and otherwise we set it to zero. Unlike Chan et al. (2013), we impose no bounds for trend inflation. Concerning the inverse gamma priors for the error variances, we follow Cogley and Sargent (2005) in placing the most weight on the sample information by using an IG prior with a single degree of freedom for all three variances: $\sigma^2_{\pi^*}$, σ^2_ρ and $\sigma^2_h \sim IG(\frac{1}{2}, \frac{0.01^2}{2})$.

We estimate the model using Bayesian MCMC methods, as detailed in Chan et al. (2013), and we draw 70,000 samples and discard the first 20,000 as burn-in. Of the remaining 50,000 draws, we keep every 25th draw in order to break the autocorrelation of the chains, which implies that we finally end up with 2000 posterior draws. The results are available upon request. For each country, we keep the median draws from the distribution of trend inflation in order to detrend wage and price inflation, as given in (4).

Supplementary material

Supplementary material associated with this article can be found, in the online version, at 10.1016/j.jmacro.2019.103166

References

Alesina, A., Summers, L., 1993. Central bank independence and macroeconomic performance: Some comparative evidence. J. Money Credit Bank. 25, 151–162.
Alogoskoufis, G.S., Smith, R., 1991. The Phillips curve, the persistence of inflation, and the Lucas critique: evidence from exchange-rate regimes. Am. Econ. Rev. 81, 1254–1275.
Apel, M., Claussen, C., 2017. Inflation targets and intervals' an overview of the issues. Sveriges Riksbank Econ. Rev. 2017:1, 83–103.
Ascari, G., Branzoli, N., Castelnuovo, E., 2011. Trend Inflation, Wage Indexation, and Determinacy in the U.S. Quaderni di Dipartimento 153. University of Pavia, Department of Economics and Quantitative Methods.
Ball, L., Cecchetti, S., 1991. Wage indexation and discretionary monetary policy. Am. Econ. Rev. 81 (5), 1310–1319.
Barker, R., 2010. Corporate Governance, Competition, and Political Parties: Explaining Corporate Governance Change in Europe. Oxford University Press.
Benati, L., 2008. Investigating inflation persistence across monetary regimes. Q J Econ 123, 1005–1060.
Borio, C., White, W., 2003. Whither Monetary and Financial Stability: The Implications of Evolving Policy Regimes. Proceedings - Economic Policy Symposium - Jackson Hole. pp. 131–211.
Brambor, T., Clark, W.R., Golder, M., 2006. Understanding interaction models: improving empirical analyses. Polit. Anal. 14, 63–82.
Walsh, C.W., 2009. Inflation targeting: what have we learned? Int. Finance 12, 195–233.
Cameron, C., Miller, L., 2015. A practitioner's guide to cluster-robust inference. J. Hum. Res. 50 (2), 317–372.
Carrillo, J.A., Peersman, G., Wauters, J., 2017. Endogenous wage indexation and aggregate shocks. BIS Working Papers 604. Bank for International Settlements.
Cecchetti, S., 1987. Indexation and incomes policy: a study of wage adjustment in unionized manufacturing. J. Labor Econ. 5, 391–412.
Chan, J.C.C., Koop, G., Potter, S., 2013. A new model of trend inflatio. J. Bus. Econ. Stat. 31, 94–106.
Christiano, L.J., Eichenbaum, M., Evans, C., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. J. Polit. Econ. 113, 1–45.
Cogley, T., Sargent, T., 2005. Drifts and volatilities: monetary policies and outcomes in the post WWII US. Rev. Econ. Dyn. 8 (2), 262–302.
Cogley, T., Sbordone, A., 2008. Trend inflation, indexation, and inflation persistence in the new Keynesian Phillips curve. Am. Econ. Rev. 98, 2101–2126.
Dincer, N.N., Eichengreen, B., 2014. Central bank transparency and independence: updates and new measures. Int. J. Cent. Bank. 10, 189–259.
Donayre, L., 2014. Estimated thresholds in the response of output to monetary policy: are large policy changes less effective? Macroecon. Dyn. 18 (1), 41–64.
Donayre, L., Panovska, I., 2016. Nonlinearities in the U.S. wage Phillips curve. J. Macroecon. 48(C), 19–43.
Duca, J.V., 1998. How increased product market competition may be reshaping America's labor markets. Economic and Financial Policy Review, Federal Reserve Bank of Dallas (IV), 2–1698.
Duca, J.V., VanHoose, D., 1998. The rise of goods-market competition and the decline in wage indexation: a macroeconomic approach. J Macroecon 20 (3), 579–598.
Eberhardt, M., 2012. Estimating panel time-series models with heterogeneous slopes. Stata J. 12, 61–71.
Ehrenberg, R.G., Danziger, L., San, G., 1983. Cost-of-living adjustment clauses in union contracts: a summary of results. J Labor Econ 1, 215–245.
Eijffinger, S.C., P., G., 2006. How transparent are central banks? Eur. J. Polit. Econ. 22, 1–21.
Erceg, C.J., Henderson, D.W., Levin, A., 2000. Optimal monetary policy with staggered wage and price contracts. J. Monet. Econ. 46, 281–313.
Fatás, A., Mihov, I., Rose, A., 2007. Quantitative goals for monetary policy. J. Money Credit Bank. 39, 1163–1176.
Fethke, G., Policano, A., 1987. Monetary policy and the timing of wage negotiations. J. Monet. Econ. 19 (1), 89–105.
Galí, J., 2011. The return of the wage Phillips curve. J. Eur. Econ. Assoc. 9, 436–461.
Ghosh, A.R., Gulde, A.M., Ostry, J.D., Wolf, H., 1997. Does the Nominal Exchange Rate Regime Matter? Working Paper 5874. National Bureau of Economic Research.
Gray, J., 1976. Wage indexation: a macroeconomic approach. J. Monet. Econ. 2, 221–235.
Gray, J., 1978. On indexation and contract length. J. Polit. Econ. 86, 1–18.
Gürkaynak, R.S., Levin, A.T., Marder, A.N., Swanson, E., 2007. Inflation targeting and the anchoring of inflation expectations in the western hemisphere. Econ. Rev. 25–47.
Gürkaynak, R.S., Levin, A.T., Swanson, E., 2010. Does inflation targeting anchor long-run inflation expectations? Evidence from the US, UK, and sweden. J. Eur. Econ. Assoc. 8, 1208–1242.
Hofmann, B., Peersman, G., R., S., 2012. Time variation in U.S. Wage Dynamics. J. Monet. Econ. 59, 769–783.
Holland, A., 1986. Wage indexation and the effect of inflation uncertainty on employment: an empirical analysis. Am. Econ. Rev. 76, 235–243.
Houben, A., 2000. The evolution of monetary policy strategies in Europe. Financial and Monetary Policy Studies, Vol.34. Kluwer Academic Publishers, Dordrecht/

[26] A caveat of imposing exogenous thresholds, rather than estimating them endogenously, is that the choice of the particular threshold may taint the significance of the estimated nonlinearities (Donayre, 2014).

Boston.
Issing, O., 1996. Is monetary targeting in germany still adequate? In: Issing, O. (Ed.), Monetary Policy in an Integrated World Economy: Symposium 1995Tübingen, Mohr.
Kenworthy, L., 2001. Wage-Setting Coordination Scores. Department of Sociology, Emory University. Unpublished manuscript
Kenworthy, L., 2001. Wage-setting institutions: a survey and assessment. World Polit. 54, 57–98.
Kumar, A., Orrenius, P., 2016. A closer look at the Phillips curve using state-level data. J. Macroecon. 47(A), 84–102.
Levin, A.T., Natalucci, F.M., Piger, J., 2004. The macroeconomic effects of inflation targeting. Fed. Reser. Bank St. Louis Rev. 86, 51–80.
Levy-Yeyati, E.L., Sturzenegger, F., 2001. Exchange rate regimes and economic performance. IMF staff papers 47. National Bureau of Economic Research, Inc.
López-Villavicencio, A., Saglio, S., 2017. The wage inflation-unemployment curve at the macroeconomic level. Oxf Bull Econ Stat 79 (1), 55–78.
Lucas, R.E., 1976. Econometric policy evaluation: a critique. Carnegie-Rochester Conference Series on Public Policy 1, 19–46.
Messina, J., Sanz-de Galdeano, A., 2014. Wage rigidity and disinflation in emerging countries. American Economic Journal: Macroeconomics 6 (1), 102–133.
Mishkin, F., 1999. International experiences with different monetary policy regimes. J Monet Econ 43, 579–605.
Mishkin, F., 2007. Will Monetary Policy Become More of a Science? NBER Working Papers 13566. National Bureau of Economic Research, Inc.
Moretti, L., 2014. Inflation targeting and product market deregulation. J Macroecon 40, 372–386.
Muto, I., Shintani, K., 2014. An Empirical Study on the New Keynesian Wage Phillips Curve: Japan and the US. MPRA Paper 53934. University Library of Munich, Germany.
Pagan, A., 1984. Econometric issues in the analysis of regressions with generated regressors. Int Econ Rev (Philadelphia) 25, 221–247.
Pesaran, M.H., Smith, R., 1995. Estimating long-run relationships from dynamic heterogeneous panels. J Econom 68, 79–113.
Ragan, J.F., Bratsberg, B., 2000. Un-COLA: why have cost-of-living clauses disappeared from union contracts and will they return? South Econ J 67, 304–324.
Rose, A., 2007. A stable international monetary system emerges: inflation targeting is bretton woods, reversed. J Int Money Finance 26, 663–681.
Samarina, A., Terpstra, M., De Haan, J., 2014. Inflation targeting and inflation performance: a comparative analysis. Appl Econ 46 (1), 41–56.
Smets, F., Wouters, R., 2007. Shocks and frictions in US business cycles: a Bayesian DSGE approach. Am. Econ. Rev. 97, 586–606.
Stepner, M., 2013. BINSCATTER: Stata Module to Generate Binned Scatterplots. Statistical Software Components S457709. Boston College Department of Economics. Revised 24 Nov 2013
Sterne, G., Stasavage, D., Chortareas, G., 2002. Does it pay to be transparent? International evidence from central bank forecasts. Fed. Reser. Bank St. Louis Rev. 84 (4), 99–118.
Svensson, L., 1999. Inflation targeting as a monetary policy rule. J. Monet. Econ. 43, 607–654.
Swamy, P., 1970. Efficient inference in a random coefficient regression model. Econometrica 38 (2), 311–323.
Visser, J., 2019. ICTWSS database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts. Amsterdam Institute for Advanced Labour Studies, University of Amsterdam, Amsterdam. Version 6.0.
Waller, C.J., VanHoose, D., 1992. Discretionary monetary policy and socially efficient wage indexation. Q. J. Econ. 107 (4), 1451–1460.