Intrinsic persistence of wage inflation in New Keynesian models of the business cycles*

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Abstract

Our paper derives and estimates an NK wage Phillips curve that accounts for intrinsic inertia. Our approach considers a model of wage-setting featuring an upward-sloping hazard function, based on the idea that the probability of resetting a wage depends on the time elapsed since the last reset. According to our specification, we obtain a wage Phillips curve including also backward-looking terms, which account for persistence. With GMM estimation we test the slope of the hazard function. Then, placing our equation in a small-scale NK model, we investigate its dynamic properties by Bayesian estimation. Model comparison shows that our model outperforms commonly used alternative methods to introduce persistence.

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Keywords: time-dependent wage adjustments, intrinsic inflation persistence, DSGE model, hybrid Phillips curves, model comparison.

1 Introduction

Micro empirical evidence suggests that nominal wages are sticky and wage inflation is persistent (Barattieri *et al.*, 2014). In aggregate models, these imperfections play a crucial role in the transmission of monetary policy and in our understanding of business cycle fluctuations (Christiano *et al.*, 2005; Rabanal and Rubio-Ramirez, 2005; Olivei and Tenreyro, 2010). If wage stickiness is neglected, macroeconomic models are actually unable to mimic the inertial dynamics of output observed in the data, unless implausibly high degrees of price stickiness are assumed (Christiano *et al.*, 2005). Moreover, the inertial structural component of wage inflation also plays a role in

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affecting price inflation persistence, which is present in the data (Fuhrer and Moore, 1995; Fuhrer, 2011). Everything equal, in fact, price inflation depends on expected future nominal marginal cost, which in turn depends on wages.

In macroeconomic models, stickiness and wage inflation persistence are usually captured by assuming wage adjustment processes *á la* Calvo and some forms of indexation to past price inflation (e.g., Erceg *et al.*, 2000; Christiano *et al.*, 2005; Rabanal and Rubio-Ramirez, 2005; Smets and Wouters, 2007). The former accounts for wage stickiness, whereas the latter introduces intrinsic inflation persistence.[1] The success of these assumptions is mainly related to their simplicity and tractability (Tsoukis *et al.*, 2009). However, both assumptions seem to be rejected by the micro evidence on wage-setting process.

In a recent study for the U.S., using quantitative data from the Survey of Income and Program Participation, Barattieri *et al.* (2014) find that wages are sticky, but the hazard function of a nominal wage change is not constant. This is in contrast to the Calvo mechanism, whereby the probability of changing a price is independent of the time elapsed from the last price reset (i.e., the hazard function is flat). In their sample, the hazard function is initially increasing, with a peak at twelve months, signaling that the probability of observing a wage reset positively depends on the time elapsed from the last wage adjustment, i.e., newer wages are stickier than older. Moreover, the probability of a wage change does not vary across quarters, i.e., the wage reset process is not affected by seasonality.

Barattieri *et al.* (2014) also reject wage indexation. They find that only a fraction of wages are reset in every period, while the remaining wages are left unchanged. This is at odds with the assumption of wage indexation, which entails that all the wages are updated in every period. Moreover, the degree of wage indexation largely varies across time (Holland, 1988) and seems to be endogenously determined by the business cycle fluctuations or other factors (Hofmann *et al.*, 2010; Di Bartolomeo *et al.*, 2015). As a result, the degree of wage indexation might not be a structural parameter in the sense of Lucas (1976). In general, inflation indexation can be considered an *ad hoc* assumption to introduce persistence because it is not supported by the survey evidence (see, e.g., Dhyne *et al.*, 2005; Fabiani *et al.*, 2005).

In light of the above facts, we propose a different approach for modelling the wage adjustment process and introduce wage inflation persistence. Our starting point is Sheedy (2007, 2010),[2] who shows that, by assuming a positive hazard function in price setting,[3] a Phillips curve with any number of lags in past inflation rates can be obtained. We borrow Sheedy's mechanism and extend it to wage setting.

We assume that the length of a wage spell directly influences the reset probability. If the slope of the hazard is positive (negative), the probability of posting a new wage increases (decreases) for the wages left fixed for many periods. Considering a non-constant hazard function, we derive

[1] Intrinsic persistence is the structural inertia that is proper for the inflation process, i.e., generated by past inflation. See, e.g., Angeloni *et al.* (2006), Rudd and Whelan (2006), Fuhrer (2011).

[2] Price adjustments with non-constant hazard functions are considered by many other papers— including Taylor (1980), Goodfriend and King (1997), Dotsey *et al.* (1999), Wolman (1999), Guerrieri (2001, 2002), and Mash (2004). These models are based on state or time-dependent assumptions and focus on price dynamics.

[3] Micro evidence on price setting about the slope of the hazard function is mixed. Results depend on the sample, countries, periods considered and methodologies used (see, e.g., Cecchetti, 1986; Nakamura and Steinsson, 2013). For a detailed discussion on the positive hazard function based on macroeconomic evidence, see instead Sheedy (2010) and Yao (2011).

a wage equation that accounts for stickiness and intrinsic wage inflation inertia, without assuming that all the wages are adjusted in every period (as for indexation)—in line with Barattieri *et al.* (2014). After deriving the wage Phillips curve, we estimate our equation by a single equation using generalized method of moments (GMM) and in a small-scale DSGE model by Bayesian techniques.[4]

Our paper makes both theoretical and empirical contributions to the literature on wage-setting behavior and inflation persistence. To our knowledge, this is the first attempt to capture wage intrinsic persistence by assuming a positive hazard function and to estimate the resulting wage Phillips curve by using macro data. Our main results can be summarized as follows.

We extend the price adjustment mechanism proposed by Sheedy (2007, 2010) to the wage-setting process. Assuming time-dependent reset probabilities, we analytically derive a wage forward-looking Phillips curve that also embeds past terms for wage inflation rates. In doing this, we provide microfoundations and a theoretical justification for intrinsic wage inflation persistence that is not at odds with the micro evidence.

We estimate our wage Phillips curve as a single equation using GMM. We find that the estimated hazard parameters are positive and statistically significant. This provides evidence that a positive hazard function emerges for wage changes at a the macro aggregate level. Moreover, our estimation also provides evidence for intrinsic (versus extrinsic) wage inflation persistence.

Including our equation in a DSGE macro model, we generalize Erceg *et al.* (2000) to take account of possible time-dependent wage adjustments. By estimating our DSGE macro model, we find that hazard gradients are positive for both prices and wages—confirming our GMM results and those of Sheedy (2010). Then, following Benati (2008, 2009), we test the robustness of time-dependent adjustments to policy regime shifts. By considering sub-samples, we find that parameters encoding intrinsic persistence remain significantly different from zero also during the Great Moderation—so they are "deep" in the sense of Lucas (1976).

Finally, we find that our model outperforms alternative specifications for price and wage adjustments. Following Rabanal and Rubio-Ramirez (2005), we evaluate the empirical performance of different models by comparing marginal likelihoods (via Bayes factor). As alternatives, we consider flat hazard functions (price and wage Phillips curves *á la* Calvo) and past or dynamic indexation mechanisms (see Galí and Gertler, 1999; Christiano *et al.*, 2005).[5]

The rest of the paper is organized as follows. In the next section, we introduce the hazard function and show the derivation and estimation of our time-dependent wage Phillips curve. Section 3 presents a DSGE simple small-scale model characterized by price and wage Phillips curve able to account for intrinsic inflation persistence. Section 4 provides ours model estimations and compares them to commonly used alternatives based on different types of inflation indexation. A final section concludes and provides some future lines of research.

[4] Regarding price adjustment in macro models with rational expectations, there is a long debate about single equation estimation and full model specification. See, e.g., Galí *et al.* (2005), Lindé (2005), Mavroeidis (2005), Rudd and Whelan (2005). We consider both approach for the sake of robustness.

[5] Similar results are found by Laforte (2007) for price setting. He finds that, in terms of the Bayes factor, the predictive ability of a model with positive hazard functions (Wolman, 1999) is strongly higher than that of models with indexation (Smets and Wouters, 2007) and sticky information (Mankiw and Reis, 2002).

2 Wage Phillips curve and intrinsic persistence

This section illustrates the main characteristics of a hazard function and shows how to derive a wage Phillips curve under the assumption that wages are reset following a vintage-dependent mechanism. Moreover, we perform the GMM estimation of our Phillips curve to test if the sign of the hazard slope is positive as well as the statistical significance of the Phillips curve coefficients.

2.1 Hazard function and time dependent adjustment

According to Sheedy (2007),[6] the probability of a wage adjustment is not random as in the Calvo specification, but depends on the time elapsed since the last wage reset. This means that the probability of posting a new wage is not equal among households, but it is a positive function of time. Formally, wage adjustments are defined by using a hazard function, which expresses the relationship between the probability of updating a wage and the duration of wage stickiness. The hazard function is defined by the sequence of probabilities $\{\alpha_{w,l}\}_{l=1}^{\infty}$, where $\alpha_{w,l}$ represents the probability of resetting a wage that remained unchanged for l periods. The hazard function is specified as follows:[7]

$$\alpha_{w,l} = \alpha_w + \varphi_w \left(1 - \alpha_{w,l-1}\right)^{-1}, \qquad \text{for } l > 1 \tag{1}$$

where α_w is the initial value of the hazard function (for $l = 1$) and φ_w is its slope. In what follows we assume that only one parameter controls the slope of the hazard,[8] as described below:

$$\begin{cases} \varphi_w = 0, & \longrightarrow \text{the hazard is flat (Calvo case);} \\ \varphi_w > 0, & \longrightarrow \text{the hazard is upward-sloping;} \\ \varphi_w < 0, & \longrightarrow \text{the hazard is downward-sloping.} \end{cases} \tag{2}$$

Thus, the hazard is positive if $\varphi_w > 0$. As mentioned above, a positive hazard function translates into a higher probability of updating a wage reset from many periods ago. Equation (1) helps us to grasp the intuition about this point: Whenever $\varphi_w > 0$, $\alpha_{w,l}$ shifts upward, implying $\alpha_{w,l+1} > \alpha_{w,l}$, this means that older wages will more likely be reset than newer wages.

Each hazard function is related to a survival function, which expresses the probability that a wage remains fixed for l periods. As for the hazard, the survival function is defined by a sequence of probabilities: $\{\varsigma_{w,l}\}_{l=0}^{\infty}$, where $\varsigma_{w,l}$ denotes the probability that a wage fixed at time t will still be in use at time $t + l$. Formally, the survival function is defined by:

$$\varsigma_{w,l} = \prod_{h=1}^{l} \left(1 - \alpha_{w,h}\right) \tag{3}$$

with $\varsigma_{w,0} = 1$.

By making use of (3), we can rewrite the non-linear recursion (1) for the wage adjustment

[6] Note that the hazard function modeled in Sheedy (2010) is equivalent to Sheedy (2007). Differences are based only on parameterization choice. Both hazard functions lead to the same Phillips curve specification.

[7] Further details about the hazard function properties are in Appendix A.

[8] In Appendix A, we provide the general case in which the sequence $\{\varphi_{w,l}\}_{l=1}^{n}$ affects the hazard gradient.

probabilities as a linear recursion for the corresponding survival function:

$$\varsigma_{w,l} = (1-\alpha_w)\varsigma_{w,l-1} - \varphi_w \varsigma_{w,l-2}, \quad \text{for } l > 1 \tag{4}$$

with $\varsigma_{w,1}$ =$(1-\alpha_w)$ for $l = 1$.[9] Let $\theta_{w,lt}$ denote the proportion of households earning at time t a wage posted at period $t-l$. The sequence $\{\theta_{w,lt}\}_{l=0}^{\infty}$ indicates the distribution of the duration of wage stickiness at time t. If both the hazard function and the evolution over the time of the distribution of wage duration satisfy certain conditions,[10] the following three relations are obtained:

$$\begin{cases} \theta_{w,l} = (\alpha_w + \varphi_w)\,\varsigma_{w,l} \\ \alpha_w^e = \alpha_w + \varphi_w \\ D_w^e = \frac{1-\varphi_w}{\alpha_w+\varphi_w} \end{cases} \tag{5}$$

where $\theta_{w,l}$ represents the unique stationary distribution to which the economy always converges, α_w^e indicates the unconditional probability of a wage reset and D_w^e denotes the expected duration of wage stickiness.

2.2 Time-dependent wage Phillips curve derivation

The supply side of the economy we consider is fairly standard (see EHL, 2000). It is composed of a continuum of monopolistically competitive firms indexed on the unit interval $\Omega \equiv [0,1]$. The production function of the representative firm $i \in \Omega$ is described by a Cobb-Douglas without capital:

$$Y_t(i) = A_t N_t(i)^{1-\phi}, \tag{6}$$

where $Y_t(i)$ is the output of good i at time t, A_t represents the state of technology, $N_t(i)$ is the quantity of labor employed by $i-$firm and $1-\phi$ is the labor share. The quantity of labor used by firm i is defined by:

$$N_t(i) = \left[\int_{\Omega} N_t(i,j)^{\frac{\varepsilon_w - 1}{\varepsilon_w}} dj \right]^{\frac{\varepsilon_w}{\varepsilon_w - 1}} \tag{7}$$

where $N_t(i,j)$ is the quantity of j-type labor employed by firm i in period t and ε_w denotes the elasticity of substitution between workers. Cost minimization with respect to the quantity of labor employed yields to the labor demand schedule:

$$N_t(i,j) = \left(\frac{W_t(j)}{W_t} \right)^{-\varepsilon_w} N_t(i) \tag{8}$$

where $W_t(j)$ is the nominal wage paid to $j-$type worker and W_t is the aggregate wage index defined in the following way:

$$W_t = \left[\int_0^1 W_t(j)^{1-\varepsilon_w} dj \right]^{\frac{1}{1-\varepsilon_w}} \tag{9}$$

[9] It derives from (3).
[10] See Appendix A for all the restrictions that must be satisfied.

We consider a continuum of monopolistically competitive households indexed on the unit interval $\Theta \equiv [0,1]$. Each household supplies a different type of labor $N_t(j) = \int_\Omega N_t(i,j)di$ to all the firms. The representative household $j \in \Theta$ chooses the quantity of labor $N_t(j)$ to supply, to maximize the following separable utility:

$$U(C_t(j), N_t(j)) = E_0 \left\{ \sum_{t=0}^{\infty} \beta^t \left[G_t \frac{(C_t(j) - hC_{t-1}(j))^{1-\sigma}}{1-\sigma} - \frac{N_t(j)^{1+\gamma}}{1+\gamma} \right] \right\} \tag{10}$$

where E_0 is the expectation operator conditional on time $t = 0$ information, β is the stochastic discount factor, σ denotes the relative risk aversion coefficient, γ is the inverse of the Frisch labor supply elasticity and h is an external habit on consumption. Finally, G_t is a preference shock that affects the marginal utility of consumption and is assumed to follow an $AR(1)$ stationary process. The household faces a standard budget constraint specified in nominal terms:

$$P_t(j) C_t(j) + E_t [Q_{t+1,t} B_t(j)] \leq B_{t-1}(j) + W_t(j) N_t(j) + T_t(j) \tag{11}$$

where $P_t(j)$ is the price of good j, $B_t(j)$ denotes holdings of one-period bonds, Q_t is the bond price, T_t represents a lump-sum government nominal transfer. Finally, $C_t(j)$ represents the consumption of household j and it is described by a CES aggregator: $C_t(j) = \left(\int_\Theta C_t(i,j)^{\frac{\varepsilon_p - 1}{\varepsilon_p}} di \right)^{\frac{\varepsilon_p}{\varepsilon_p - 1}}$, where $C_t(i,j)$ denotes the quantity of i-type good consumed by household j and ε_p is the elasticity of substitution between goods.

In our framework, households are wage setters. In setting wages, each maximizes (10) internalizing the effects of labor demand (8) and taking account of (11). Households are subject to a random probability of updating their wage, but, according to our time-dependent mechanism, a wage change will be more likely to be observed when the last wage reset happened far in the past. Formally, suppose that at time t a household sets a new wage, denoted by W_t^*,[11] if the household still earns this wage at time $\tau \geq t$, its relative wage will be W_t^*/W_τ, and the household utility can be written as $U\left[W_t^*/W_\tau; C_{\tau|t}; N_{\tau|t}\right]$;[12] by considering the survival function, the household will then choose its optimal reset wage by solving:

$$\max_{W_t^*} \sum_{\tau=t}^{\infty} \varsigma_{w,\tau-t} E_t \left\{ \left(\prod_{s=t+1}^{\tau} \frac{\pi_s^p}{I_s} \right) U \left[\frac{W_t^*}{W_\tau}; C_{\tau|t}; N_{\tau|t} \right] \right\} \tag{12}$$

where $\pi_s^p = P_s/P_{s-1}$ is the gross price inflation rate and $I_s = i_s/i_{s-1}$ is the gross nominal interest rate. This maximization is subject to the budget constraint (11) and the labor demand schedule (8). Equation (12) yields the following first-order condition:

$$\sum_{\tau=t}^{\infty} \varsigma_{w,\tau-t} E_t \left(\frac{W_t^*}{W_\tau} \right)^{-\varepsilon_w} \left(\prod_{s=t+1}^{\tau} \frac{\pi_s}{I_s} \right) \left[U_c(C_{\tau|t}, N_{\tau|t}) \frac{N_{\tau|t}}{P_\tau} (1-\varepsilon_w) - \varepsilon_w U_n(C_{\tau|t}, N_{\tau|t}) \frac{N_{\tau|t}}{W_t^*} \right] = 0 \tag{13}$$

[11] Because each household solves the same optimization problem, henceforth, index j is omitted.

[12] $C_{\tau|t}$ and $N_{\tau|t}$ denote the level of consumption and the labor supply at time τ of a household, respectively, for which the last wage reset was in period t.

where $U_c(C_{\tau|t}, N_{\tau|t})$ is the marginal utility of consumption and $-U_n(C_{\tau|t}, N_{\tau|t})$ is the marginal disutility of labor. Considering that the marginal rate of substitution between consumption and leisure is $MRS_{\tau|t} = -\frac{U_n(C_{\tau|t}, N_{\tau|t})}{U_c(C_{\tau|t}, N_{\tau|t})}$, and the steady state wage mark-up is $\mu_w = \frac{\varepsilon_w}{\varepsilon_w - 1}$, equation (13) can be rearranged and expressed in terms of the optimal wage reset as:

$$W_t^* = \left[\frac{\mu_w \left(\sum_{\tau=t}^{\infty} \varsigma_{w,\tau-t} \beta^{\tau-t} MRS_{\tau|t} P_\tau\right)}{\sum_{\tau=t}^{\infty} \varsigma_{w,\tau-t} \beta^{\tau-t}}\right] \tag{14}$$

Assuming that the economy has converged to $\{\theta_{w,l}\}_{l=0}^{\infty}$, the wage level (9) can be expressed as a weighted-average of the past reset wages:

$$W_t = \left(\sum_{l=0}^{\infty} \theta_{w,l} W_{t-l}^{*\,1-\varepsilon_w}\right)^{\frac{1}{1-\varepsilon_w}} \tag{15}$$

As common practice in DSGE models, we log-linearize (14) and (15) around a deterministic steady state. Specifically, there is no trend inflation, i.e., $\Pi^w = 1$ and $\Pi^p = 1$, where Π^w and Π^p represent the steady state of the wage and price inflations, respectively. This assumption implies that the steady state value for the relative reset wage is 1 and that the steady state of the real interest rate is equal to β^{-1}. Thus, we obtain:[13]

$$w_t^* = \sum_{\tau=t}^{\infty} \left(\frac{\beta^{\tau-t} \varsigma_{w,\tau-t}}{\sum_{j=0}^{\infty} \beta^j \varsigma_{w,j}}\right) [w_\tau - \Xi_w \mu_\tau^w] \tag{16}$$

where $\Xi_w = \frac{1}{1+\varepsilon_w \gamma}$ and

$$w_t = \sum_{l=0}^{\infty} \theta_{w,l} w_{t-l}^* \tag{17}$$

Equations (16) and (17) describe the wage adjustment mechanism. The time-dependent wage Phillips curve is derived by combining them with (4) and (5).

Specifically, inserting (4) in (16), we obtain:

$$w_t^* = \beta(1-\alpha_w) E_t w_{t+1}^* - \beta^2 \varphi_w E_t w_{t+2}^* + \left[1 - \beta(1-\alpha_w) + \beta^2 \varphi_w\right] (w_t - \Xi_w \mu_t^w) \tag{18}$$

By making use of (5), equation (17) can be recast as follows:

$$w_t = (1-\alpha_w) w_{t-1} - \varphi_w w_{t-2} + (\alpha_w + \varphi_w) w_t^* \tag{19}$$

where we have used the fact that the stationary distribution of the wage duration (5) can be rewritten in a recursive way as:

$$\theta_{w,l} = (1-\alpha_w)\theta_{w,l-1} - \varphi_w \theta_{w,l-2} \quad \text{for } l > 1 \tag{20}$$

with $\theta_{w,0} = \alpha_w + \varphi_w$ and $\theta_{w,1} = (1-\alpha_w)(\alpha_w + \varphi_w)$ because of (5) and (3).

[13] Small-caps letters denote log-deviations from the steady state.

The general expression for the wage Phillips curve is obtained from (18) and (19):

$$\pi_t^w = \psi_w \pi_{t-1}^w + \beta \left[1 + (1-\beta)\psi_w\right] E_t \pi_{t+1}^w - \beta^2 \psi_w E_t \pi_{t+2}^w - k_w \mu_t^w, \tag{21}$$

where π_t^w is the wage inflation and μ_t^w is the wage mark-up. Moreover:

$$\begin{cases} \psi_w = \frac{\varphi_w}{(1-\alpha_w)-\varphi_w[1-\beta(1-\alpha_w)]} \\ k_w = \frac{(\alpha_w+\varphi_w)\left[1-\beta(1-\alpha_w)+\beta^2\varphi_w\right]}{(1-\alpha_w)-\varphi_w[1-\beta(1-\alpha_w)]}\Xi_w \end{cases}, \tag{22}$$

where ψ_w and k_w are coefficients depending on the hazard parameters.[14] In particular, φ_w and α_w control the slope and the initial level of the hazard function, respectively.

As highlighted previously, our wage Phillips curve establishes that current wage inflation is determined not only by the expectations over future wage inflation[15] but now also has a "history dependent" dimension as an endogenous lagged term affects it. Unlike the case with indexation, here the past term is on wage inflation and not on price inflation, indicating a "purely" intrinsic inertia. Moreover, the presence of this term does not derive any more from an *ad hoc* assumption, but it now has a clear theoretical foundation.

The backward term in wage inflation derives from a "selection effect" stemming from our pricing mechanism. In all sticky wage models, after a temporary shock has vanished, there are still incentive to change prices because the average price level changes and not all wage setters are able to adjust in every period. In a Calvo framework, the incentive to raise wages exactly compensate the incentive to reduce them. By contrast, in our framework, after a positive shock, the former dominates the latter, generating inflation persistence. The intuition of our result is explained more in detail as follows.

Suppose that at time t a temporary positive cost-push shock hits the economy, as nominal wages are sticky, a part of them does not change, while the remaining fraction increases. Therefore, the average wage level increases. At time $t+1$, after the shock has vanished, the wage setters who did adjust at t find their relative wage too low and want to raise their nominal wage (the "catch-up" effect). By contrast, the wage setters who did adjust at t now find their relative wage too high and, hence, want to lower it (the "roll-back" effect).

In the Calvo context, the two effects exactly countervail one another and inflation stabilizes right away. When non-constant hazard functions are considered, one effect dominates over the other. In the case of the positive hazard function, the "catch-up" effect prevails over the "roll-back" effect because wages that remained unvaried at t are more likely to be changed than newly set wages. As a result, aggregate wage inflation is also still positive after the shock has dissipated, generating inflation persistence. In terms of (21), this is captured by the presence of the past wage inflation rate. In our framework, on average, the "selection effect" is positive because the relative gains of adjusting a wage are more likely to be higher for wages that have remained unchanged for longer periods.

[14] The evolution of the wage Phllips curve coefficients for the general case $n > 1$ is reported in Appendix A.

[15] Both inflation at time $t+1$ and $t+2$ are relevant. Although the coefficient associated with the latter is negative, the overall effect of expected inflation is positive on its current rate. The second-order term in the difference equation captures the dynamics of the adjustment process. See Sheedy (2007) for a discussion.

A positive slope for the hazard also implies that the coefficient on the backward term is always positive, which is consistent with observed wage inflation dynamics. Finally, our Phillips curve encompasses the Calvo purely forward-looking specification as the particular case $\varphi_w = 0$ (a flat hazard) implies that ψ_w drops to zero and, consequently, we return to the standard textbook case.

2.3 Hazard function estimation

As in Sheedy (2007, 2010), we estimate our wage Phillips curve via GMM, to inspect the shape of the hazard function and the statistical significance of the coefficients attached to the lead and lag components of the wage equation. For the sake of simplicity we show only the estimation of (21) when $n = 1$.[16] Because it is not easy to find an observable proxy for the wage mark-up, the latter can be expressed as a function of unemployment, as in Galí *et al.* (2011):

$$\mu_t^w = \gamma u_t \tag{23}$$

where u_t represents the unemployment gap. Therefore, (21) becomes:

$$\pi_t^w = \psi_w \pi_{t-1}^w + \beta \left[1 + (1-\beta)\,\psi_w\right] E_t \pi_{t+1}^w - \beta^2 \psi_w E_t \pi_{t+2}^w - k_w \gamma u_t \tag{24}$$

To perform a GMM estimation of (24), we need to use a set of instruments, to correctly identify all the coefficients. Let z_{t-1} represent a vector of observable variables known at time $t-1$: Under rational expectations the error forecast of π_t^w is uncorrelated with information contained in z_{t-1}, and thus, the following orthogonality condition holds:

$$E_t \left\{ \left[\pi_t^w - \psi_w \pi_{t-1}^w - \beta(1 + (1-\beta)\,\psi_w) E_t \pi_{t+1}^w + \beta^2 \psi_w E_t \pi_{t+2}^w + k_w \gamma u_t\right] z_{t-1} \right\} = 0 \tag{25}$$

Following Galí and Gertler (1999), because (25) is non-linear in the structural parameters, we normalize the orthogonality condition in the following way:

$$E_t \left\{ \chi_w \left[\pi_t^w - \psi_w \pi_{t-1}^w - \beta(1 + (1-\beta)\,\psi_w) E_t \pi_{t+1}^w + \beta^2 \psi_w E_t \pi_{t+2}^w + k_w \gamma u_t\right] z_{t-1} \right\} = 0 \tag{26}$$

where $\chi_w = (1-\alpha_w) - \varphi_w \left[1 - \beta\,(1-\alpha_w)\right]$.

Our estimation is made using quarterly U.S. data ranging from 1960:1 to 2011:4: All the data are from the FRED database. Wage inflation is measured by the *compensation per hour*, whereas for the unemployment rate, we use the *civilian unemployment rate.* The set of instruments is composed of the lags of the following observable variables: wage inflation, unemployment, price inflation, consumer price index, output gap, labor share, and the spread between ten-year Treasury Bond and three-month Treasury Bill yields. In particular, six lags in price inflation, wage inflation and CPI, four lags for the output gap and two lags for the remaining instruments are used.[17]

[16] For $n > 1$, we find that the extra leads and lags deriving from this specification are not statistically significant.

[17] The sample range, the data and the wage Phillips curve specification used for GMM estimation differ from those that will be used for the Bayesian estimation. We made this choice to avoid the possibility that the Bayesian comparison might unduly favor our model with respect to the alternatives considered. However, we chose to perform a "non-informative" estimation for the hazard slope parameters to test the robustness of our comparison (see Section 4).

The structural form of (26) is estimated by imposing $\beta = 0.99$, $\varepsilon_w = 8.85$ and $\gamma = 2$; wage elasticity is derived as in Galí (2011) by using $\varepsilon_w = [1 - \exp(-\gamma u^n)]^{-1} = 8.85$, where we assume $\gamma = 2$ and a natural unemployment rate u^n equal to 6%, as the average rate of the period considered. The reduced form coefficients (see (22)) are a convolution of the structural parameters estimated and they are obtained by substituting these parameters into them; the standard errors are computed using the delta method.[18]

The results for the structural form estimation are reported in Table 1. We show the estimation for the structural parameters φ_w (hazard slope) and α_w (hazard initial value); moreover, we also report D_w^e and α_w^e (computed as in (5)) and the $J - stat$.

Table 1 – Wage Phillips curve estimation (structural form)[19]

α_w	φ_w	D_w^e	α_w^e	$J - stat$
0.318*	0.126*	1.964*	0.444*	19.527
(0.050)	(0.030)	(0.146)	(0.033)	[0.813]

Notes: a 6-lag Newey-West estimate of the covariance matrix is used.

Standard errors are shown in parentheses.

For the *J-stat* the p-value is shown in brackets.

* denotes statistical significance at 5% level.

All the coefficients estimated are statistically significant, and the hazard function is estimated to be upward-sloping. Wages are estimated to be rigid, as an adjustment comes every two quarters. The $J - stat$ is a test of over-identifying moment condition: In our case, we accept the null hypothesis that the over-identifying restrictions are satisfied (the model is "valid").

We now report the reduced form of (25), obtained by substituting the estimated values of α_w and φ_w into (22).

$$\pi_t^w = \underset{(0.038)}{0.197\pi_{t-1}^w} + \underset{(0.000)}{0.991E_t\pi_{t+1}^w} - \underset{(0.037)}{0.193E_t\pi_{t+2}^w} - \underset{(0.006)}{0.03u_t} \qquad (27)$$

Also under this specification, all the coefficients are statistically significant at the 5% level (standard errors computed using delta method are reported in parentheses). Our wage Phillips curve, in line with the underlying theory, is able to capture the well-known negative relation between the unemployment gap and the wage inflation, as highlighted by the negative coefficient measuring the slope of the curve. In Figure 1, we provide a graphical representation for the hazard and survival functions derived from our estimation and computed by using (1) and (4), respectively. The hazard clearly shows a positive slope, meaning that a time-dependent mechanism for wage adjustment emerges.

[18] See Papke and Wooldridge (2005).

[19] The estimation has been performed by using Cliff's (2003) GMM package for MATLAB available from https://sites.google.com/site/mcliffweb/programs.

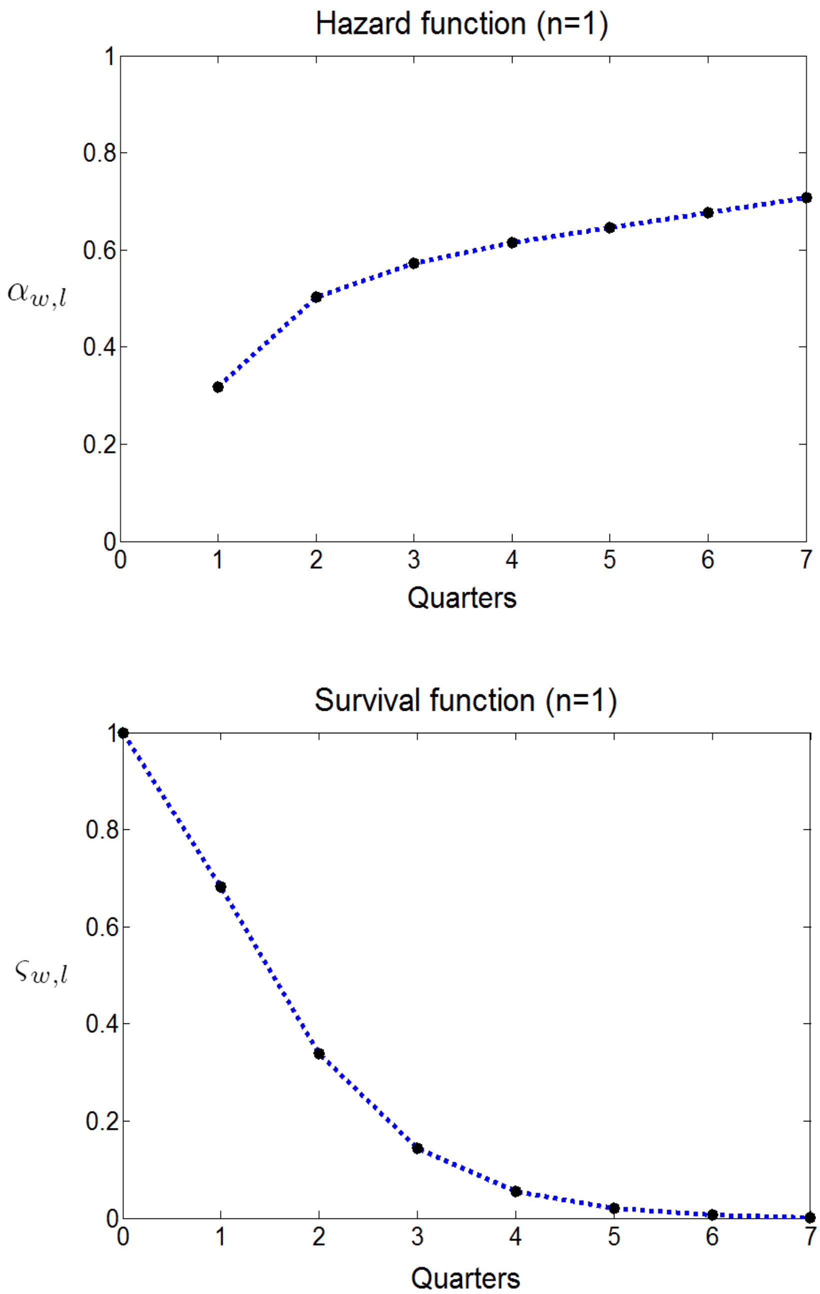

Figure 1 - Hazard and survival function deriving from GMM estimation.

3 The DSGE model

Our model generalizes EHL (2000) by assuming that price and wage adjustments are governed by a time-dependent mechanism. Moreover, to improve its empirical realism, we also consider habit formation in consumption, which implies persistence in the IS curve. As the main difference with EHL (2000) is the derivation of the Phillips curves, as explained in Section 2, the description of the model derivation is not detailed. For a full derivation, we refer to EHL (2000). All the equation reported herein are expressed as log-linear deviations from the steady state.

3.1 The log-linearized economy

The demand side of the economy is described by a simple IS curve, which is obtained by log-linearizing the Euler equation around the steady state. As usual, the Euler equation is derived by maximizing the utility function (10) subject to the budget constraint (11). Formally,

$$y_t = \frac{1}{1+h} E_t y_{t+1} + \frac{h}{1+h} y_{t-1} - \frac{1-h}{\sigma(1+h)} \left(i_t - E_t \pi^p_{t+1} + E_t g_{t+1} - g_t \right), \tag{28}$$

where y_t is the output, π^p_t is the price inflation rate, i_t is the nominal interest rate set by the central bank, and g_t is a preference shock. The lagged term on output is due to the presence of external consumption habits.

Price adjustment is described by a Phillips curve with a time-dependent mechanism similar to that used for wages. As shown by Sheedy (2007), by using the same structure of (1) for prices, the price Phillips curve is:[20]

$$\pi^p_t = \psi_p \pi^p_{t-1} + \beta \left[1 + (1-\beta) \psi_p \right] E_t \pi^p_{t+1} - \beta^2 \psi_p E_t \pi^p_{t+2} + k_p \left(mc_t + \zeta_t \right), \tag{29}$$

where mc_t is the real marginal cost and ζ_t is a price mark-up shock. The coefficients ψ_p and k_p are a function of the parameters characterizing the hazard function for prices φ_p and α_p:

$$\begin{cases} \psi_p = \frac{\varphi_p}{(1-\alpha_p) - \varphi_p [1-\beta(1-\alpha_p)]} \\ k_p = \frac{(\alpha_p + \varphi_p)\left[1-\beta(1-\alpha_p)+\beta^2\varphi_p\right]}{(1-\alpha_p)-\varphi_p[1-\beta(1-\alpha_p)]} \eta_{cx} \end{cases}$$

The parameter φ_p controls the gradient of the hazard function, and α_p is its starting level; the coefficient $\eta_{cx} = \frac{1-\phi}{1-\phi+\phi\varepsilon_p}$ is the elasticity of a firm's marginal cost with respect to average real marginal cost and depends on the labor share $(1-\phi)$ and the elasticity of substitution between goods (ε_p).

The log-linearized real marginal cost is:

$$mc_t = \omega_t + n_t - y_t, \tag{30}$$

where ω_t denotes the real wage and n_t is the amount of hours worked. Equation (30) is derived from cost minimization subject to the production function (6), which can be written in a log-

[20] The curve is derived in a similar way as explained in Section 2. For a detailed derivation of it, see Sheedy (2007).

linearized form as:

$$y_t = a_t + (1 - \phi)\, n_t, \tag{31}$$

where a_t is the technology shock.

The wage adjustment is described by the wage Phillips curve (21) previously derived, i.e.:

$$\pi_t^w = \psi_w \pi_{t-1}^w + \beta\,[1 + (1 - \beta)\,\psi_w]\, E_t \pi_{t+1}^w - \beta^2 \psi_w E_t \pi_{t+2}^w - k_w(\omega_t - mrs_t), \tag{32}$$

where the wage mark-up is written as the difference between the real wage (ω_t) and the marginal rate of substitution (mrs_t) because the labor market is characterized by imperfect competition.

The real wage, by assumption, follows:

$$\omega_t = \pi_t^w - \pi_t^p + \omega_{t-1}. \tag{33}$$

The marginal rate of substitution between consumption and hours worked is obtained from the wage setter's problem, and it equals the ratio between the marginal utility of leisure and consumption. Formally, in log-linear terms, it is given by:

$$mrs_t = \frac{\sigma}{1 - h}\,(y_t - h y_{t-1}) + \gamma n_t - g_t, \tag{34}$$

Finally, monetary policy is assumed to follow a simple Taylor rule:

$$i_t = \rho_r i_{t-1} + (1 - \rho_r)\,(\delta_\pi \pi_t^p + \delta_x y_t) + \upsilon_t, \tag{35}$$

where ρ_r captures the degree of interest rate smoothing, δ_π and δ_x measure the response of the monetary authority to the deviation of inflation and output from their steady state values; υ_t is a monetary policy shock.

Aside from the monetary disturbance,[21] all the shocks considered in the model follow an $AR(1)$ process:

$$\begin{cases} a_t = \rho_a a_{t-1} + \varepsilon_t^a, \\ g_t = \rho_g g_{t-1} + \varepsilon_t^g, \\ \zeta_t = \rho_\zeta \zeta_{t-1} + \varepsilon_t^\zeta, \\ \upsilon_t = \varepsilon_t^\upsilon, \end{cases} \tag{36}$$

where $\varepsilon_t^j \sim N\left(0, \sigma_j^2\right)$ are white noise shocks uncorrelated among them and ρ_j are the parameters measuring the degree of autocorrelation, for $j = \{a, g, \zeta\}$.

Summarizing, our model consists of six equations, describing the following: the dynamic IS (28); the price Phillips curve (29); the real marginal cost (30); the production function (31); the wage Phillips curve (32); the real wage dynamics (33); the marginal rate of substitution (34); and the Taylor rule (35). Shock dynamics are described by (36).

[21] Monetary policy persistence is already captured by the lagged term in (35). We have, however, successfully checked the robustness of our result with respect to alternative assumptions. Specifically, we have considered an $AR(2)$ process for the interest rate in equation (35). Results are available upon request.

4 Empirical analysis

We estimate our model (28)-(36) for the U.S. economy by Bayesian techniques. Our choice is motivated by the fact that Bayesian methods outperform GMM and maximum likelihood in small samples.[22] After writing the model in state-space form, the likelihood function is evaluated using the Kalman filter, whereas prior distributions are used to introduce additional non-sample information into the parameters estimation: Once a prior distribution is elicited, posterior density for the structural parameters can be obtained by reweighting the likelihood by a prior. The posterior is computed via numerical integration by making use of the Metropolis-Hastings algorithm for Monte Carlo integration; for the sake of simplicity, all structural parameters are assumed to be independent from each other.

We use four observable macroeconomic variables: real GDP, price inflation, real wage, and nominal interest rate. The dynamics are driven by four orthogonal shocks, including monetary policy, productivity, preference and price mark-up; because the number of observable variables is equal to the number of exogenous shocks, the estimation does not present problems deriving from stochastic singularity.[23] The estimation of the model is performed by using informative priors and, as a robustness check, non-informative priors for the parameters characterizing the slope of the hazard function.

We aim to test whether the model exhibits a positive hazard function, i.e., the time-dependent price/wage adjustments holds. Following Benati (2008, 2009), we also test the robustness of our pricing mechanism to policy regime shifts. By considering only price rigidity and flexible wages, Benati (2009) analyses several models to build inflation persistence including Sheedy (2007).[24] He finds evidence of positive-sloping hazard functions, but, by considering the Great Moderation subsample, he also finds that the parameters encoding the hazard slope have dropped to zero in the last thirty years. He concludes that these parameters depend on the monetary regime referring to the switch in the way monetary policy is conducted as discussed in Clarida *et al.* (2000). However, he only focuses on price inflation: We generalize his setup by considering staggered wages with a possible time-dependent adjustment process in the labor markets. As mentioned above, nominal rigidity and persistence in wages may have important implications for both inflation persistence and monetary policy effects.

After estimating our model for the full sample (1960:1-2008:4), we also consider a smaller one (1982:1-2008:4), representative of the Great Moderation, to investigate if a positive hazard function still holds in a period characterized by small volatility in shocks and more aggressive tactics by central bankers in the fight against inflation.

Finally, we evaluate the empirical performance of our time-dependent Phillips curves in relation to alternative specifications commonly used in the literature. We consider the traditional forward-looking Phillips curves derived in EHL (2000) extended with price and wage indexation, which is often one main assumption to account for inflation persistence. Model comparison is based on

[22] For an exhaustive analysis of Bayesian estimation methods, see Geweke (1999), An and Schorfheide (2007) and Fernández-Villaverde (2010).

[23] Problems deriving from misspecification are widely discussed in Lubik and Schorfheide (2006) and Fernández-Villaverde (2010).

[24] Specifically, Benati (2009) analyzed Fuhrer and Moore (1995), Galí and Gertler (1999), Blanchard and Galí (2007), Sheedy (2007), Ascari and Ropele (2009).

log-marginal likelihood. To apply this methodology, we will show how the models compared here are nested.

The next subsection presents the data used and prior distributions. Subsection 4.2 provides the estimation for the baseline model. Subsection 4.3 evaluates our time-dependent model against alternative specifications.

4.1 Data and prior distributions

In our estimations, we use U.S. quarterly data. All the time series used are from the FRED database maintained by the Federal Reserve Bank of St. Louis. The *real gross domestic product* is used as measure of the output; the *effective Fed funds rate* is used for the nominal interest rate. Price inflation is measured using the *GDP implicit price deflator* taken in log-difference. Real wage is obtained dividing the nominal wage, measured by the *compensation per hour in nonfarm business sector,* by the *GDP implicit price deflator*. All the variables have been demeaned; output and real wage are detrended using Baxter and King's bandpass filter.

Our choices about prior beliefs are as follows. The coefficients of the Taylor rule are centered on a prior mean of 1.5 for inflation and 0.125 for the output gap, which are Taylor's (1999) estimates, and follow a Normal distribution. These values are quite standard in the literature. The smoothing parameter is assumed to follow a Beta distribution, with a mean of 0.6 and a standard deviation equal to 0.2. The same choice has been made for the consumption habit. We assume that the inverse of Frisch elasticity is based on a Gamma distribution, with a mean of 2 and a standard deviation 0.375. These priors are fairly diffuse and broadly in line with those adopted in previous studies, such as Del Negro *et al.* (2007), Smets and Wouters (2007), Justiniano and Primicieri (2008), and Justiniano *et al.* (2013).

For the hazard function coefficients, we perform an "informative estimation" using as priors coefficients estimated from a single equation GMM;[25] we assign a Normal distribution to φ_p and φ_w with a standard deviation equal to 0.2, whereas α_p and α_w follow a Beta distribution with a standard deviation of 0.1. As a robustness check, following Benati (2009), we also estimate the model using non-informative priors for the parameters affecting the slope of the hazard function, instead of those derived from the GMM estimations. Unlike his approach, we use a Uniform distribution with support $[-1, 1]$: The choice of such a large interval is motivated by the fact that we want to investigate if the hazard slope is positive, negative or zero.

We need to calibrate some parameters to avoid identification problems.[26] Because we consider a production function without capital, it is difficult to estimate β and ϕ, which are set to 0.99 and 0.33, respectively. Similarly, we fix $\varepsilon_p = 6$ and $\varepsilon_w = 8.85$, implying a price and wage mark-up equal to 1.20 and 1.12. Price elasticity is calibrated following Sheedy (2007), to be coherent with the hazard priors derived from his GMM estimation. As explained in Section 2, wage elasticity is set equal to 8.85. Finally, all the autoregressive coefficients of the shocks follow a Beta distribution,

[25] The values of φ_w and α_w estimated in Section 2 are used as priors. For the hazard characterizing price adjustment, we directly use as priors the GMM estimates of Sheedy (2007).

[26] The identification procedure has been performed by using the Identification toolbox for Dynare, which implements the identification condition proposed by Iskrev (2010a, 2010b). For a review of identification issues arising in DSGE models, see Canova and Sala (2009).

with a mean of 0.5 and a standard deviation equal to 0.2. The prior for the shock standard deviations is an Inverse Gamma, with a mean of 0.01 and 2 degrees of freedom.

4.2 Estimation results

Our estimations are reported in Table 2, which also summarizes the 90% probability intervals and our beliefs about the priors. The table describes the results for the full sample and the Great Moderation. We report posterior estimation of the shocks and structural parameters, obtained by the Metropolis-Hastings algorithm, when informative priors for the hazard slope are used.

Table 2 – Prior and posterior distributions[27]

	Prior distribution			Posterior distribution (full sample)			Posterior distribution (Great Moderation)		
	Density	Mean	St. Dev.[28]	Mean	5%	95%	Mean	5%	95%
σ	Gamma	1.0	0.375	1.324	0.673	1.955	1.227	0.581	1.820
γ	Gamma	2.0	0.375	2.515	2.041	2.997	2.249	1.732	2.748
h	Beta	0.6	0.2	0.906	0.866	0.946	0.908	0.863	0.955
δ_π	Normal	1.5	0.25	1.423	1.197	1.650	1.851	1.524	2.158
δ_x	Normal	0.125	0.05	0.215	0.152	0.279	0.164	0.096	0.235
ρ_r	Beta	0.6	0.2	0.818	0.787	0.850	0.850	0.819	0.883
α_p	Beta	0.132	0.1	0.020	0.001	0.042	0.063	0.001	0.124
φ_p	Normal	0.222	0.2	0.195	0.157	0.233	0.128	0.048	0.213
α_w	Beta	0.318	0.1	0.126	0.073	0.179	0.151	0.073	0.228
φ_w	Normal	0.126	0.2	0.242	0.210	0.277	0.250	0.203	0.297
ρ_a	Beta	0.5	0.2	0.781	0.706	0.854	0.850	0.819	0.883
ρ_g	Beta	0.5	0.2	0.768	0.717	0.817	0.802	0.738	0.867
ρ_ζ	Beta	0.5	0.2	0.825	0.762	0.889	0.822	0.732	0.910
σ_a	Inv. Gamma	0.01	2	0.019	0.013	0.025	0.014	0.008	0.019
σ_g	Inv. Gamma	0.01	2	0.053	0.038	0.068	0.044	0.028	0.059
σ_υ	Inv. Gamma	0.01	2	0.002	0.002	0.002	0.001	0.001	0.002
σ_ζ	Inv. Gamma	0.01	2	0.020	0.013	0.028	0.030	0.012	0.047

In the full sample case, the estimated hazard function is upward-sloping, because φ_p and φ_w are both positive. Thus, the time-dependent mechanism seems to be able to account for inflation inertia for both prices and wages. The duration of a price spell is 3.7 quarters, whereas wages appear to be less sticky, because their duration is 2.05 quarters.[29] These durations are similar to those obtained in the literature, in which it is also common to find higher duration in price setting compared to wages (e.g., Rabanal and Rubio-Ramirez, 2005; Galí *et al.*, 2011).

[27] The posterior distributions are obtained using the Metropolis-Hastings algorithm; the procedure is implemented using the Matlab-based Dynare package. Mean and posterior percentiles are from two chains of 250,000 draws each from the Metropolis-Hastings algorithm, for which we discarded the initial 30% of draws.

[28] For the Inverse Gamma distribution the degrees of freedom are indicated.

[29] The durations (D_i^e) of price and wage stickiness are computed by using the following relation: $D_i^e = \frac{1-\varphi_i}{\alpha_i+\varphi_i}$ for $i = \{p, w\}$ (see (5)).

The estimations for the parameters characterizing the utility function (i.e., habit, relative risk aversion and inverse of Frisch elasticity) are coherent with the standard findings in the literature (see, e.g., Del Negro *et al.* 2007; Smets and Wouters, 2007; Justiniano and Primicieri, 2008; Justiniano *et al.*, 2013).

The response of monetary authority to inflation and the output gap is in line with the Taylor principle; estimated coefficients of the monetary rule are in line with the literature. The estimated degree of interest rate smoothing is 0.82. All the shocks exhibit a high degree of autocorrelation, of approximately 0.8.

By considering the Great Moderation period, as expected, we find a more aggressive monetary policy stance (Clarida *et al.*, 2000). Differently from Benati (2009), we find that the hazard functions still exhibit positive slopes also in this sub-sample. This result gives us evidence that a pricing mechanism based on hazard function still holds also in a period characterized by a central bank more concerned in fighting inflation, as highlighted by the higher estimated coefficient for δ_π. As a result, intrinsic persistence also holds for the Great Moderation period.

The price duration increases to 4.5 quarters. This is highlighted by the fact that the hazard function sloping is still positive, but smaller. This fact is in line with macroeconomic theory: During the Great Moderation, inflation has dropped, the cost of not adjusting a price is smaller compared to the previous period and this translates into a longer price spell.

By contrast, computed wage stickiness is rather stable, reflecting the fact that wage bargaining is more influenced by institutional factors related to the labor market than by the monetary policy.[30] The stability of wage duration over time is also found by Rabanal and Rubio-Ramirez (2005).

In Figure 2, we plot prior distribution, posterior distribution and posterior mode of the estimated parameters.

[30] Considering a Walrasian labor market, as in Benati (2009), may force the estimated price Phillips curve to capture also wage stickiness present in the data. This leads to an overestimation of price duration, which, in the Great Moderation sub-sample, drops the hazard coefficients to zero, implying a quite flat hazard function and no intrinsic persistence in price inflation.

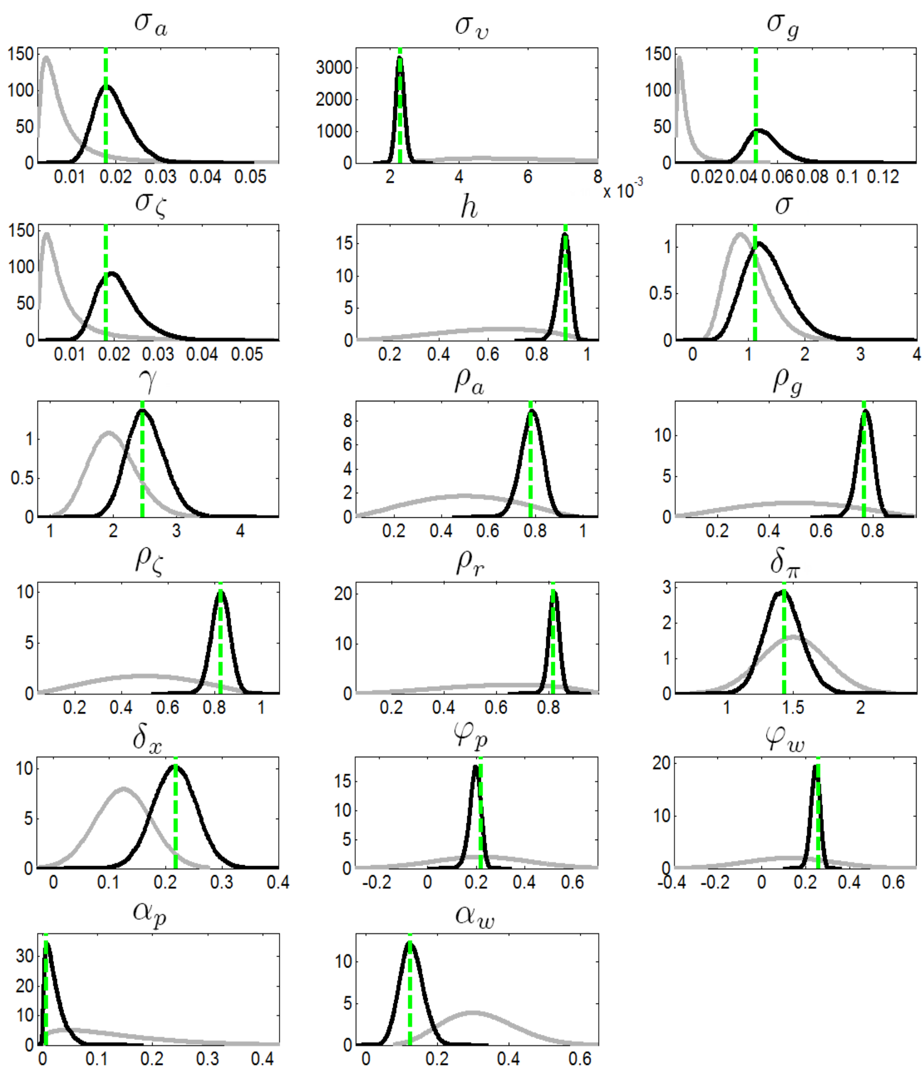

Figure 2 - Prior distribution (grey curve), Posterior distribution (green curve) and Posterior mode (dotted line) of the estimated parameters.

Bayesian estimations of DSGE models can be quite sensitive to the choice of priors for model-specific parameters and other assumptions regarding, e.g., measures of variables used and shock specifications. Thus, we have checked the robustness of our analysis by considering also uniform priors for the parameters φ_p and φ_w with support $[-1; 1]$,[31] whereas the prior distributions for the remaining parameters are the same as those used previously. The results are provided in Table 3.

[31] Choosing this large support we can test if the hazard slope is negative, positive or flat. The prior mean is centered on 0.

The "non-informative" estimation confirms our results about the hazard function, which is still characterized by positive slope, both in the full sample and during the Great Moderation; the estimated parameters for the hazard slope are very similar to the ones estimated under "informative" priors. This result shows as the hazard function mechanism is robust to a change of policy.

Table 3 - Prior and posterior distributions under non-informative priors

	Prior distribution			Posterior distribution (full sample)			Posterior distribution (Great Moderation)		
	Density	Mean	St. Dev.	Mean	5%	95%	Mean	5%	95%
σ	Gamma	1.0	0.375	1.321	0.670	1.933	1.227	0.595	1.834
γ	Gamma	2.0	0.375	2.511	2.021	2.974	2.251	1.738	2.753
h	Beta	0.6	0.2	0.906	0.868	0.948	0.909	0.865	0.957
δ_π	Normal	1.5	0.25	1.428	1.203	1.661	1.855	1.545	2.171
δ_x	Normal	0.125	0.05	0.215	0.151	0.277	0.165	0.096	0.234
ρ_r	Beta	0.6	0.2	0.818	0.787	0.850	0.851	0.820	0.882
α_p	Beta	0.132	0.1	0.020	0.001	0.041	0.067	0.001	0.133
φ_p	Uniform	0	0.57	0.195	0.158	0.236	0.125	0.042	0.213
α_w	Beta	0.318	0.1	0.126	0.073	0.177	0.151	0.072	0.225
φ_w	Uniform	0	0.57	0.243	0.209	0.276	0.252	0.207	0.298
ρ_a	Beta	0.5	0.2	0.780	0.704	0.854	0.832	0.755	0.910
ρ_g	Beta	0.5	0.2	0.768	0.719	0.817	0.800	0.738	0.866
ρ_ζ	Beta	0.5	0.2	0.824	0.760	0.887	0.824	0.737	0.914
σ_a	Inv. Gamma	0.01	2	0.019	0.013	0.025	0.014	0.008	0.019
σ_g	Inv. Gamma	0.01	2	0.052	0.038	0.067	0.044	0.029	0.061
σ_υ	Inv. Gamma	0.01	2	0.002	0.002	0.002	0.001	0.001	0.002
σ_ζ	Inv. Gamma	0.01	2	0.020	0.013	0.028	0.029	0.013	0.044

We have also successfully checked the robustness of our results by considering different model specifications (i.e., a model without habit), various specifications for the Taylor rule (as already mentioned) and alternative series for observable variables.[32] Results are available upon request.

In the figures below, we plot the dynamic behavior of the model variables, described by the Bayesian impulse response functions, conditional to price mark-up and monetary policy shocks. The solid line represents the estimated response, with the shaded area capturing the corresponding 95 percent confidence interval. The horizontal axis measures the quarters after the initial shock. The monetary shock is illustrated in Figure 3, whereas the cost-push shock is described by Figure 4.

The monetary shock affects both real and nominal variables and has persistent effects on output. As expected, in response to the monetary tightening, GDP declines with a characteristic

[32] In particular, we have considered a different measure for prices by using the *nonfarm business sector implicit price deflator*. With regard to wages, we have considered alternatives measures given by: *average hourly earnings of production*; *business sector compensation per hour*; *hourly earnings for manufacturing sector*.

hump-shaped pattern. It reaches a trough after four quarters, and then, it slowly reverts back to its initial level. Both price inflation and real wage also exhibit a hump-shaped behavior. Similarly, a positive cost-push shock affects all the variables and has persistent effects on output. In both cases, the pattern of interest rate, real output, hours, real wage and inflation are in line with the literature (see, e.g., Smets and Wouters, 2003; Christiano *et al.*, 2005; Güneş and Millard, 2012).

In regard to wage dynamic adjustments, which are the core of our investigation, wage inflation has a pattern similar to that of price inflation. In Figure 3, following a monetary shock, wage inflation has a moderate hump shape and progressively returns to its steady state with a little overshooting. In the case of a cost-push shock, described in Figure 4, the dynamic response of wage inflation is strongly hump shaped with a trough after approximately ten quarters, which denotes a high level of wage persistence. We will discuss in more detail the response of nominal wage inflation in the next section, when we compare the responses of our model to alternatives in which inertia is introduced by wage indexation to past prices.

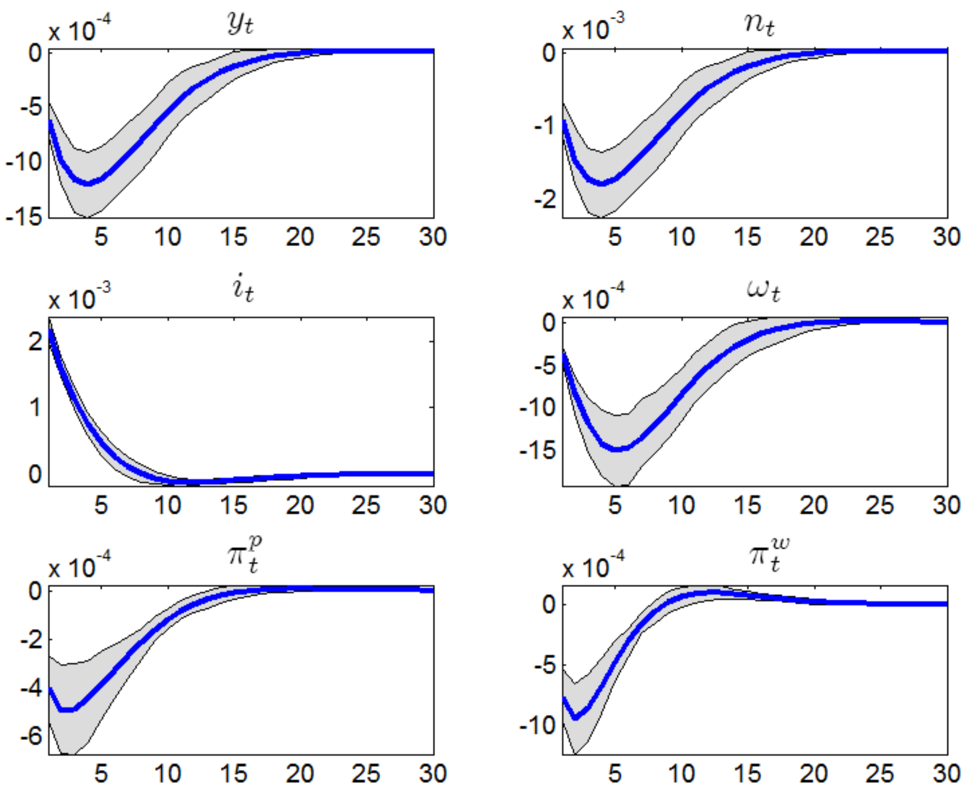

Figure 3 - Bayesian IRFs conditional to a monetary shock.

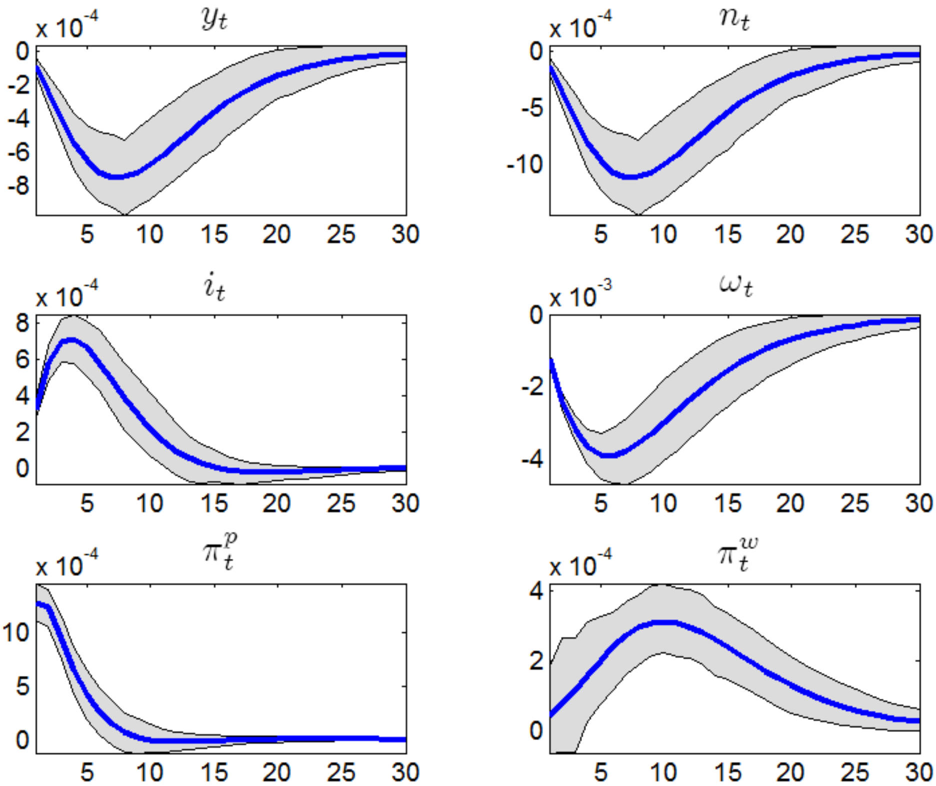

Figure 4 - Bayesian IRFs conditional to a price mark-up shock.

4.3 Time-dependent Phillips curves vs. alternatives

In this section, we aim to compare the empirical performance of our time-dependent Phillips curves to different specifications accounting for price and wage inflation inertia. Specifically, we focus on EHL (2000), augmenting it by indexation because, as discussed, this is a common way to introduce price and wage persistence in New Keyensian DSGE models. We consider two different widely used forms of indexation and we compare these alternatives to our baseline model in terms of log-marginal density.

4.3.1 Alternative price-setting mechanisms

Simply by setting $\varphi_p = 0$ and $\varphi_w = 0$ in (29) and (32), we obtain flat hazard functions and, therefore, price and wage Phillips curves *á la* Calvo as in EHL (2000), which is nested in our model. Formally, these two purely forward-looking curves are

$$\pi_t^p = \beta E_t \pi_{t+1}^p + \lambda_p (mc_t + \zeta_t) \tag{37}$$

$$\pi_t^w = \beta E_t \pi_{t+1}^w - \lambda_w \mu_t^w \tag{38}$$

where $\lambda_p = \frac{\alpha_p[1-\beta(1-\alpha_p)]}{1-\alpha_p}\eta_{cx}$ and $\lambda_w = \frac{\alpha_w[1-\beta(1-\alpha_w)]}{1-\alpha_w}\Xi_w$.

Equations (37) and (38) clearly do not exhibit any persistence. However, price and wage inertia can be easily introduced by considering indexation to past price inflation. Two popular ways to do it have been proposed by Galí and Gertler (1999) and Christiano *et al.* (2005). The former obtains a lagged term in the aggregate Phillips curve by assuming the existence of a fraction of agents who set prices/wages according to a backward-looking rule of thumb. The latter, instead, considers dynamic indexation, i.e., all agents not able to reset their price/wage adjust according to past price inflation, and consequently, the past inflation rate appears in the Phillips curve.

According to Galí and Gertler (1999), the Phillips curves can be rewritten as follows:

$$\pi_t^p = \frac{\xi_p}{\Lambda_p}\pi_{t-1}^p + \frac{\beta(1-\alpha_p)}{\Lambda_p}E_t\pi_{t+1}^p + \lambda_p^\xi(mc_t + \zeta_t) \tag{39}$$

$$\pi_t^w = \frac{\xi_w}{\Lambda_w}\pi_{t-1}^p + \frac{\beta(1-\alpha_w)}{\Lambda_w}E_t\pi_{t+1}^w - \lambda_w^\xi\mu_t^w \tag{40}$$

where ξ_p (ξ_w) measures the fraction of rule-of-thumb agents, i.e., the degree of price (wage) indexation to past price inflation; $\Lambda_p = 1-\alpha_p+\xi_p[\alpha_p + (1-\alpha_p)\beta]$, $\Lambda_w = 1-\alpha_w+\xi_w[\alpha_w + (1-\alpha_w)\beta]$, $\lambda_p^\xi = \frac{(1-\xi_p)(1-\alpha_p)\lambda_p}{\Lambda_p}$, and $\lambda_w^\xi = \frac{(1-\xi_w)(1-\alpha_w)\lambda_w}{\Lambda_w}$.

Instead, assuming dynamic indexation as in Christiano *et al.* (2005), equation (37) and (38) can be written as:

$$\pi_t^p = \frac{\iota_p}{(1+\iota_p\beta)}\pi_{t-1}^p + \frac{\beta}{1+\iota_p\beta}E_t\pi_{t+1}^p + \lambda_p^\iota(mc_t + \zeta_t) \tag{41}$$

$$\pi_t^w = \iota_w\pi_{t-1}^p - \iota_w\beta\pi_t^p + \beta E_t\pi_{t+1}^w - \lambda_w\mu_t^w \tag{42}$$

where ι_p (ι_w) denotes the degree of price (wage) indexation to last period's price inflation and $\lambda_p^\iota = \frac{\lambda_p}{(1+\iota_p\beta)}$.

4.3.2 Model comparison

Our formalization nests different models of price and wage adjustment. Differences only depend on the Phillips curve parameterization. With different assumptions on φ_p, φ_w, ι_p, ι_w, ξ_p, ξ_w, we can consider positive or flat hazard functions augmented by the two different types of indexation aforementioned. We compare our baseline (BASE) to three alternative scenarios:[33]

1. EHL model with dynamic indexation (DYNind), by considering (41) and (42);

2. EHL model with rule of thumb indexation *á la* Galí-Gertler (GG), by considering (39) and (40).

3. EHL model with inertia only in wage equation (MIXED), i.e., Phillips curves are (37) and (32).

[33] We omit the comparison with a model characterized by simple forward-looking Phillips curves *á la* Calvo because this model has not intrinsic persistence. However, Rabanal and Rubio-Ramirez (2005) showed that this model exhibits quite the same performance as a model with indexation.

The measure used to compare the models is the log-marginal likelihood, which is a measure of the fit of a model in explaining the data.[34] The aim is to evaluate if the way in which price and wage are modeled adjustment affects the fit of a model. The model with the highest log-marginal likelihood better explains the data.[35] Table 4 reports our results.

Table 4 - Log-marginal data densities and Bayes factors for different models[36]

Model	Log-marginal data density	Bayes factor vs. BASE
BASE	3615.6	
BASE (non-info)	3613.6	$\exp[-2.0]$
MIXED	3598.1	$\exp[-17.5]$
DYNind	3569.7	$\exp[-45.9]$
GG	3564.8	$\exp[-50.8]$

The difference, in terms of marginal likelihood, between Galí-Gertler specification and dynamic indexation is minimal. According to Jeffreys' scale of evidence,[37] this difference must be considered as "slight" evidence in favor of DYNind with respect to GG. However, our model clearly outperforms both the alternatives considered: In particular, the Bayes factor gives "very strong" evidence in favor of our specification. As the models considered differ only for the Phillips curves, this result signals that the data prefer a pricing method based on positive hazard functions; this is not surprising as the micro evidence rejects both constant hazard and indexation. Under "non-informative" priors, we observe a slight decrease of the marginal likelihood: This happens because under diffuse priors there is an increase in model complexity, and this penalizes the marginal data density (this effect dominates the improvement in model fit).

In the comparison between our time-dependent adjustment model and those with dynamic indexation or rule-of-thumb agents, the fit of the different models is judged by the log-marginal likelihoods. The differences are significant (as evidenced by the Bayes factors). The models only differ in the Phillips curves. Moreover, all price equations are observable equivalent for the different specifications considered (all include backward- and forward-looking terms for the main dependent variables and imply similar reduced forms). Therefore, the great improvement in the fit of our model should be necessarily explained by differences in the wage equations.

Comparing the dynamics of our model with those generated by models with indexation, we observe large differences for the nominal wage dynamics. These differences are plotted in Figure 5. Unsurprisingly, the dynamics of all the other variables are qualitatively similar among the models considered here; thus, we do not report them (the dynamics of other variables in case of indexation mechanisms are similar to those reported in Figures 3 and 4).

[34] The estimated reduced forms of the Phillips curves considered are reported in Appendix B.

[35] For details on model comparison technique, see Fernández-Villaverde and Rubio-Ramirez (2004), Rabanal and Rubio-Ramirez (2005), Lubik and Schorfheide (2006), Riggi and Tancioni (2010).

[36] For the computation of the marginal likelihood for different model specifications, we used the modified harmonic mean estimator, based on Geweke (1999). The Bayes factor is the ratio of posterior odds to prior odds (see Kass and Raftery, 1995).

[37] Jeffreys (1961) provided a scale for the evaluation of the Bayes factor indication. Odds ranging from 1:1 to 3:1 give "very slight evidence"; odds ranging from 3:1 to 10:1 constitute "slight evidence"; odds ranging from 10:1 to 100:1 constitute "strong to very strong evidence"; odds greater than 100:1 give "decisive evidence."

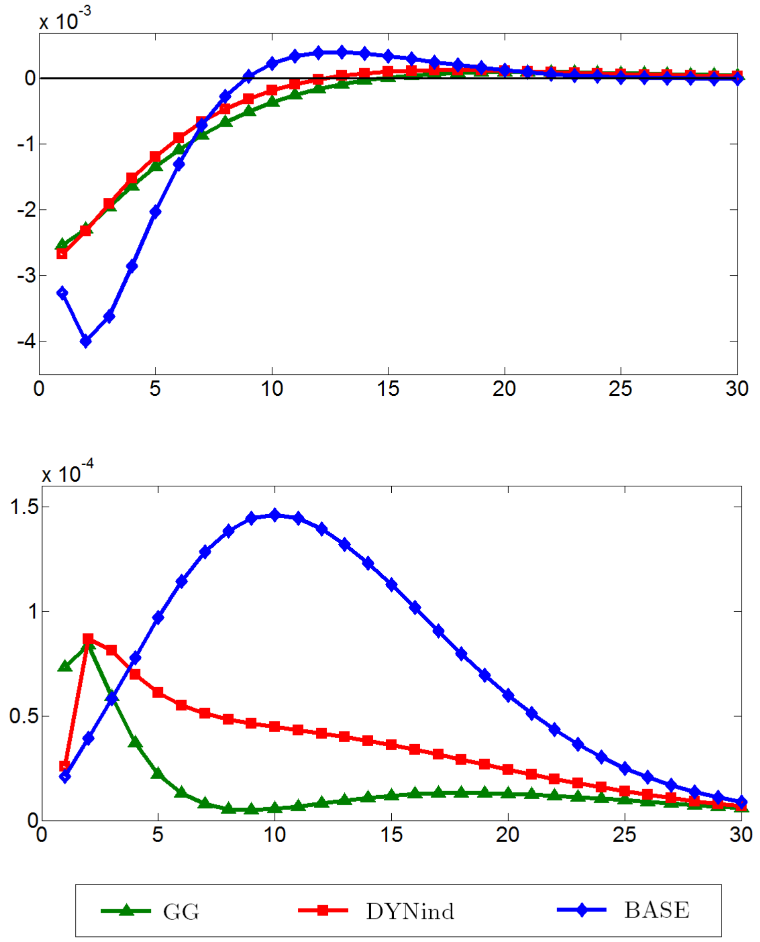

Figure 5 - IRFs of the wage inflation to 1% monetary policy (upper panel) and price mark-up (lower panel) shocks for several model specifications.

Considering first the monetary shock, in both cases of indexation to past price, wage inflation follows a smooth path, while in our framework, beyond exhibiting a little hump immediately after the shock, it is more volatile. Under a cost-push shock, for both the models with indexation, wage inflation responds with a spike in the early periods and then slowly returns to its stationary value. A model with a vintage-dependent wage adjustment instead generates a dynamic quite different with respect to indexation models.

The higher predictive ability associated to our model by the Bayesian comparison comes out from the different dynamic behavior of nominal wages generated by our wage Phillips curve. This occurs because the reduced forms of the wage Phillips curves are different. Our mechanism embeds the wage Phillips curve of a lagged term for wage inflation (intrinsic inertia). In this respect, it differs from the wage equations with indexation, for which the backward term is on past price inflation (inherited inertia). As the marginal likelihood is a measure of the ability of a model to fit the data, we can conclude that our wage-setting mechanism better captures the information contained in the data with respect to models with some form of indexation.

In line with the above observations, the crucial role of wage adjustment is also highlighted by the fact that the log-marginal likelihood of MIXED is significantly higher than that of indexation models (see Table 4). It is worth recalling that MIXED refers to the model with time-dependent adjustment only in wages, while prices follow the common Calvo's scheme. This result confirms the key role of nominal wage rigidities to explain the data—as claimed by, e.g., Christiano *et al.* (2005), Rabanal and Rubio-Ramirez (2005), Olivei and Tenreyro (2010).

5 Conclusions

Our paper proposed a new approach to model wage inflation persistence and evaluated its empirical relevance. By assuming that wage adjustments are governed by a vintage-dependent mechanism, we showed how to derive a New Keynesian wage Phillips curve that also embeds backward terms for past wage inflation (intrinsic persistence). In our specification, the presence of endogenous-lagged terms does not rest on the unrealistic assumption of indexation to past price inflation rates, but it has a theoretical reason justified by the presence of a positive selection effect. Because of wage stickiness, after the shock has vanished, wage setters continue to adjust wages. In the standard Calvo model, upward and downward adjustments compensate (no selection effect). In our case, due to a positive hazard function, a wage setter is more likely to adjust wages in the same direction of the past shock, inducing wage persistence.

We successfully tested the relevance of intrinsic wage persistence by using both GMM and Bayesian methods. Lagged terms for wage inflation are significantly different from zero in single equation GMM estimation. Placing our equation in a small-scale DSGE model, by Bayesian estimations, we confirmed that an upward-sloping hazard function emerges for both prices and wages. By comparing log-marginal likelihoods, we found that our model outperforms alternatives based on popular mechanisms for modeling inflation persistence. We showed that the key rationale of this result is in our wage adjustment mechanism. Our result confirms the crucial role of nominal wage rigidities to understand the economic fluctuations (see, e.g., Christiano *et al.*, 2005). Finally, we test that the hazard function slope does not change with the policy regime.

It would be interesting to observe other features of the different modeling choices for price-setting, e.g., implications for welfare and optimal monetary policy. However, this is beyond the scope of the current paper, and we leave it for future research.

Appendix A – Hazard function properties

This appendix provides further details about the hazard function properties. We refer to Sheedy (2007) for the proofs relative to the hazard function mentioned here. See, in particular his Appendix A.2 and A.5. Moreover, we show the evolution of the hazard and the resulting wage Phillips curve for the general case $n > 1$.

Assuming that $\Gamma_t \subset \Theta$ denotes the set of households that post a new wage at time t, the length of wage stickiness can be defined as:

$$D_t(j) \equiv \min\{l \geq 0 \mid j \in \Gamma_{t-l}\} \tag{43}$$

where $D_t(j)$ is the duration of a wage spell for household j for which the last reset was l periods ago.

As explained in the text, the hazard function is defined by a sequence of probabilities: $\{\alpha_{w,l}\}_{l=1}^{\infty}$, where $\alpha_{w,l}$ represents the probability to reset a wage that remained unchanged for l periods. This probability is defined as: $\alpha_{w,l} \equiv \Pr(\Gamma_t \mid D_{t-1} = l-1)$.

Each hazard function is related to a survival function, which expresses the probability that a wage remains fixed for l periods. As for the hazard, the survival function is defined by a sequence of probabilities: $\{\varsigma_{w,l}\}_{l=0}^{\infty}$, where $\varsigma_{w,l}$ denotes the probability that a wage fixed at time t will still be in use at time $t+l$.

The hazard function can be reparameterized by making use of a set of $n+1$ parameters and rewritten as (1) if $n=1$ and in the following way for the general case $n>1$:

$$\alpha_{w,l} = \alpha_w + \sum_{j=1}^{\min(l-1,n)} \varphi_{w,j} \left[\prod_{k=l-j}^{l-1} (1-\alpha_{w,k})\right]^{-1}, \tag{44}$$

where α_w is the initial value of the hazard function and $\varphi_{w,j}$ is its slope; n is the number of parameters that control the slope. The sequence of parameters $\left\{\varphi_{w,l}\right\}_{l=1}^{n}$ affect the gradient of the hazard function in the following way:

$$\begin{cases} \varphi_{w,l} = 0, \, \forall \; l = 1, ..., n \longrightarrow \text{the hazard is flat (Calvo case);} \\ \varphi_{w,l} \geq 0, \, \forall \; l = 1, ..., n \longrightarrow \text{the hazard is upward-sloping;} \\ \varphi_{w,l} \leq 0, \, \forall \; l = 1, ..., n \longrightarrow \text{the hazard is downward-sloping.} \end{cases} \tag{45}$$

The survival function (4) in the general case is rewritten as:

$$\varsigma_{w,l} = (1-\alpha_w)\varsigma_{w,l-1} - \sum_{h=1}^{\min(l-1,n)} \varphi_{w,h}\varsigma_{w,l-1-h} \tag{46}$$

Following Sheedy (2007), we assume that the hazard function satisfies two weak restrictions:

$$\begin{cases} \alpha_{w,1} < 1, \text{ meaning that is allowed a degree of wage stickiness;} \\ \alpha_{w,\infty} > 0, \text{ with } \alpha_{w,\infty} = \lim_{l\to\infty} \alpha_{w,l}. \end{cases} \tag{47}$$

We now introduce $\theta_{w,lt} \equiv \Pr(D_t = l)$ which denotes the proportion of households earning at time t a wage posted at period $t-l$. The sequence $\{\theta_{w,lt}\}_{l=0}^{\infty}$ indicates the distribution of the duration of wage stickiness at time t. This distribution evolves over time according to:

$$\begin{cases} \theta_{w,0t} = \sum_{l=1}^{\infty} \alpha_{w,l}\theta_{w,l-1,t-1} \\ \theta_{w,lt} = (1-\alpha_{w,l})\,\theta_{w,l-1,t-1} \end{cases} \tag{48}$$

If the hazard function satisfies the restrictions (47) and the evolution over the time of the distribution of wage length evolves as in (48), then *a)* from whatever starting point, the economy always converges to a unique stationary distribution $\{\theta_{w,l}\}_{l=0}^{\infty}$. Hence $\theta_{w,lt} = \theta_{w,l} = \Pr(D_t = l)$, $\forall t$; *b)* let's consider (1) and assume that the economy has converged to $\{\theta_{w,l}\}_{l=0}^{\infty}$; the relations expressed in (5) are obtained. For $n > 1$, conditions in (5) become:

$$\begin{cases} \theta_{w,l} = \left(\alpha_w + \sum\limits_{h=1}^{n} \varphi_{w,h}\right) \varsigma_{w,l} \\ \alpha_w^e = \alpha_w + \sum\limits_{l=1}^{n} \varphi_{w,l} \\ D_w^e = \frac{1-\sum_{l=1}^{n} l\varphi_{w,l}}{\alpha_w + \sum_{l=1}^{n} \varphi_{w,l}} \end{cases} \tag{49}$$

Then, inserting (46) in (16), we obtain:

$$w_t^* = \beta(1-\alpha_w)E_t w_{t+1}^* - \sum_{l=1}^{n} \beta^{l+1}\varphi_{w,l}E_t w_{t+l+1}^* + \left[1 - \beta(1-\alpha_w) + \sum_{l=1}^{n} \beta^{l+1}\varphi_{w,l}\right](w_t - \Xi_w \mu_t^w) \tag{50}$$

By making use of (49), equation (17) can be recast as follows:

$$w_t = (1-\alpha_w)\, w_{t-1} - \sum_{l=1}^{n} \varphi_{w,l} w_{t-1-l} + \left(\alpha_w + \sum_{h=1}^{n} \varphi_{w,h}\right) w_t^* \tag{51}$$

Finally, the expression for the wage Phillips curve in the case $n > 1$ is obtained by mixing (50) and (51):

$$\pi_t^w = \sum_{l=1}^{n} \psi_{w,l}\pi_{t-l}^w + \sum_{l=1}^{n+1} \delta_{w,l} E_t \pi_{t+l}^w - k_w \mu_t^w \tag{52}$$

where the coefficients $\psi_{w,l}$, $\delta_{w,l}$ and k_w have the following parameterization:

$$\begin{aligned}
\psi_{w,l} &= \frac{\varphi_{w,l} + \sum_{h=l+1}^{n} \varphi_{w,h}\left[1 - \beta(1-\alpha_w) + \sum_{k=1}^{h-1} \beta^{k+1}\varphi_{w,k}\right]}{\chi_w} && \text{for } l = 1, ..., n \\
\delta_{w,1} &= \frac{\beta\left[(1-\alpha_w) - \sum_{h=1}^{n} \beta^h \varphi_{w,h}\left(\alpha_w + \sum_{k=1}^{h-1} \varphi_{w,k}\right)\right]}{\chi_w} \\
\delta_{w,l+1} &= -\frac{\beta^{l+1}\left[\varphi_{w,l} + \sum_{h=l+1}^{n} \beta^{h-1}\varphi_{w,h}\left(\alpha_w + \sum_{k=1}^{h-1} \varphi_{w,k}\right)\right]}{\chi_w} && \text{for } l = 1, ..., n \\
k_w &= \frac{\Xi_w\left[\left(\alpha_w + \sum_{h=1}^{n} \varphi_{w,h}\right)\left[1 - \beta(1-\alpha_w) + \sum_{h=1}^{n} \beta^{h+1}\varphi_{w,h}\right]\right]}{\chi_w}
\end{aligned}$$

where $\chi_w = (1-\alpha_w) - \sum_{h=1}^{n} \varphi_{w,h}\left[1 - \beta(1-\alpha_w) + \sum_{k=1}^{h-1} \beta^{k+1}\varphi_{w,k}\right]$.

It is easy to check that if we assume that only one parameter controls the slope of the hazard function (i.e., $n = 1$), the wage Phillips curve (52) becomes that reported in the paper, i.e., (21).

Appendix B – Reduced form Phillips curves from the Bayesian estimation

The reduced form from the Bayesian estimation for the price and wage Phillips curves in the EHL model with dynamic indexation (DYNind) are:

$$\pi_t^p = 0.138\pi_{t-1}^p + 0.853E_t\pi_{t+1}^p + 0.01mc_t \tag{53}$$
$$\pi_t^w = 0.153\pi_{t-1}^p - 0.151\pi_t^p + 0.99E_t\pi_{t+1}^w - 0.01\mu_t^w \tag{54}$$

The EHL model with rule-of-thumb indexation *á la* Galí-Gertler (GG) implies:

$$\pi_t^p = 0.144\pi_{t-1}^p + 0.847E_t\pi_{t+1}^p + 0.009mc_t \tag{55}$$
$$\pi_t^w = 0.106\pi_{t-1}^p + 0.884E_t\pi_{t+1}^w - 0.01\mu_t^w \tag{56}$$

Finally, our time-dependent specification is associated with:

$$\pi_t^p = 0.2\pi_{t-1}^p + 0.99E_t\pi_{t+1}^p - 0.196E_t\pi_{t+2}^p + 0.012mc_t \tag{57}$$
$$\pi_t^w = 0.288\pi_{t-1}^w + 0.99E_t\pi_{t+1}^w - 0.283E_t\pi_{t+2}^w - 0.007\mu_t^w \tag{58}$$

As shown in the paper, the above reduced forms imply that models without time-dependent adjustment capture persistence in the wage equation by past price inflation (i.e., inherited inertia), whereas (57)-(58) include a backward term for wage inflation. Both for price and wage inflation equations, our estimated Phillips curves capture a higher degree of persistence, as highlighted by the coefficients attached to backward inflation, with respect to models based on indexation. This is not surprising as, since the Great Moderation, indexation to past inflation has progressively vanished, and thus, the parameter encoding it has progressively dropped to zero (see Benati, 2008).

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