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The Inflation-Unemployment Trade-Off: Empirical Considerations and a Simple US-Euro Area Comparison

A Relação entre Inflação e Desemprego: Considerações Empíricas e uma Comparação Simples entre EUA e Zona Euro

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ABSTRACT
This paper uses recently developed robust estimation methods to empirically reassess the long-standing inflation-unemployment trade-off debate. Indeed, we study to what extent unemployment-based New Keynesian Phillips Curves are informative about the relationship between inflation dynamics and labor market conditions. In particular, we attempt to quantify the 'elasticities' of inflation with respect to unemployment in two economies, the US and the Euro Area, whose labor market characteristics are admittedly very different. We find that the relevance of the inflation-unemployment trade-off and its empirical adequacy is greatly enhanced once the informational content of key labor market variables is explored in our estimations.
Keywords: Phillips curve; unemployment; model averaging.

JEL Classification: E24; E31; C26

1. Introduction

The inflation-unemployment trade-off has played a key role in the development of modern macroeconomics and policymaking. Central banks regularly monitor labour market conditions, as these are important sources of business cycle fluctuations (e.g. Zanetti, 2011), but also because they affect firms' pricing decisions through their impact on marginal costs. Thus, if there is a significant link between the structural features of labour markets and inflation, this will have significant consequences for monetary policy and how it is transmitted to the economy. Given the potential importance of labour market conditions for inflation dynamics and the recent developments in the literature, there is an opportunity to reassess the long-standing debate about the inflation-unemployment relationship.

The original Phillips curve essentially stemmed from the observation of a historical inverse relation between inflation and unemployment rates. This ad-hoc relationship has more recently been superseded by micro-founded structural versions, based on the behaviour of optimising forward-looking agents. The so-called New Keynesian Phillips curve (NKPC) describes inflation as being driven by inflationary pressures, either in the form of an output gap (Fuhrer and Moore, 1995) or conveyed by firms' marginal costs, as in Galí and Gertler (1999).

A distinct, but related, approach that has attracted a great deal of interest is one that incorporates labour market frictions into the NK theory of inflation. Blanchard and Galí (2007 and 2010), Ravenna and Walsh (2008), and Krause, et al. (2008), for example, show how standard measures of marginal cost are incorrect in the presence of these rigidities. Accounting for such frictions provides an additional and broader source of inflation persistence, giving rise to a NKPC that explicitly depends on unemployment – at first resembling a "traditional" Phillips curve, albeit stemming from a properly micro-founded framework.

The contribution of this paper is to empirically re-examine the inflation-unemployment relationship by making use of the NKPC with labour market frictions of Blanchard and Galí (2010, BG henceforth) and Ravenna and Walsh (2008, RW henceforth). In particular, we are interested in quantifying the elasticities of inflation with respect to unemployment of two economies, the US and the Euro Area (EA), which are known to have markedly distinct labour market characteristics. These specifications provide a theoretically sound, dynamic description of the inflation-unemployment trade-off, in contrast with traditional approaches investigating the empirical connection between unemployment and inflation with little formal theory. In turn, this allows us to relate the empirical relationships with the stylised facts of these two economies.

Indeed, we first explore a simple implication of the baseline BG formulation: while it is impossible to identify and estimate all the structural parameters of the model, we note that a reduced-form approach is still able to convey interesting information, as we can infer the level of labour market sclerosis and its relationship with inflation from the relative magnitude of the coefficients associated with unemployment and the change in unemployment. This, in itself, provides a check on the empirical adequacy and relevance of the inflation-unemployment trade-off. We then consider an extended specification discussed in both BG and RW, which offers a more complete description of the dynamic relationship between inflation and unemployment. To achieve this, we employ a recently developed model averaging approach for instrumental variables estimation that allows us to circumvent some of the difficulties typically associated with inference on NKPCs. This is very convenient in our case, given that we are mainly interested in 'composite' unemployment elasticities.[1]

Our results suggest that a *stricto sensu* implementation of the theoretical models, although broadly adequate, provides mild empirical support only. However, once we consider several extensions that relax some of the limiting assumptions implied by the theoretical framework in BG and RW, we are able to provide a richer empirical description of the inflation-unemployment trade-off and the empirical adequacy of these models is strengthened. In particular, using additional information about key labour market variables (such as measures of the NAIRU, labour market tightness and separation rates) improves the fit and the precision of our estimations.

Our paper is related to, and complements, the work of Ravenna and Walsh (2008), which provides empirical tests on the connection between the structural features of the US labor market and inflation, and that of Krause et al. (2008), which focus on a structural approach to analyse the relevance of a specific form of labour market rigidity for the US economy (search-match frictions, albeit without allowing for real wage rigidities). Both these papers emphasise 'labour market augmented' definitions of real marginal costs as the driving variable for inflation. In contrast, we focus on the unemployment elasticities of inflation by studying versions of the same model in which inflation is explicitly written in terms of unemployment.[2] This then allows us to carry out feasible and meaningful comparisons between the US and the EA economies, given the lack of detailed data on relevant labour market variables for the latter.

While we acknowledge that the EA comprises countries with different labour market characteristics, for the purposes of our study it makes sense to treat them as a bloc. First, they share a common monetary policy framework with the European Central Bank targeting an EA wide measure of inflation. Our study aims precisely to understand to what extent inflation in the EA bloc is driven by aggregate labour market pressures. Second, recent cross-country evidence produced by Jolivet, et al. (2006) and Hobijn and Sahin (2009), for example, support the view that the US labour market is much more fluid – higher job finding rates and lower unemployment durations – than any of the EA countries, so it is interesting to understand how these differences are translated in terms of the inflation-unemployment relationship.

It is also important to recognise the inherent difficulties of our single-equation estimation approach. An alternative approach to investigate the empirical merits of these NKPC specifications would be to employ system Bayesian methods on a DSGE model (Galí et al., 2011), for example, who add unemployment as an observable to a DSGE model for the US). While a full-information approach is in principle more efficient, a misspecification in a bloc of the model may spillover to the whole model (Ruge-Murcia, 2007). Thus, our single-equation method may be seen as a complement to this approach, by allowing us to focus solely on the specification of the NKPC. Also, a method-of-moments framework is, in principle, more

[1] Although not strictly 'elasticities', we follow RW in designating the coefficients associated with the right-hand side variables of the unemployment-based NKPC as elasticities.

[2] Both BG and RW derive a linearized version of the NKPC with inflation as a function of unemployment, but do not study this empirically.

robust to statistical misspecifications, as it only requires minimal distributional assumptions. Moreover, our approach allows for simple comparisons between different economies to be carried out, which might otherwise prove too cumbersome with fully specified DSGE models.

The paper proceeds as follows. The next section describes briefly the theoretical framework that shows explicitly the role of unemployment in determining inflation dynamics. In Section 3, we analyse the baseline specifications using GMM and the Model Averaging procedure of Kuersteiner and Okui (2010), while in Section 4 we extend the empirical framework to allow for a more flexible description of the inflation-unemployment relationship. Section 5 concludes.

2. Nkpc Models with Unemployment

A number of recent papers has attempted to modify the New Keynesian setup by introducing labour market frictions in addition to standard nominal rigidities. Here, we focus on two contributions that lead to very similar specifications of the NKPC with inflation written as a function of unemployment fluctuations.

2.1. A simplified nkpc with unemployment

BG construct a model with staggered price and nominal wage rigidities combined with Diamond-Mortensen-Pissarides-type search and match frictions, with the addition of real wage rigidities. This setup gives rise to explicit interactions amongst productivity shocks, unemployment fluctuations and inflation. By making some simplifying assumptions (hiring costs small relative to output and small separation rates; see BG for details), these authors first express the inflation rate π_t as a function of labour market tightness $\hat{x}_t$ (with $x \in [0,1]$ defined as the steady-state ratio of hires to unemployment) and log labor productivity $\hat{\alpha}_t$ (assumed to follow an AR(1) process)

$$\pi_t = \eta\hat{x}_t' - \Psi\gamma\hat{\alpha}_t \tag{1}$$

with “ . ” denoting variables in deviation-from-steady-state form, where $\gamma \in [0,1]$ reflects real wage rigidities, while η and Ψ are composite parameters that depend on frictions such as hiring costs, firms gross markup and the degree of price stickiness.

Noting that the relation between labour market tightness and unemployment $\hat{u}_t$ can be rewritten as

$$(1-u)\delta\hat{x}_t = -\hat{u}_t + (1-\delta)(1-x)\,\hat{u}_{t-1} \tag{2}$$

where $\delta \in (0,1)$ is the (exogenous) separation rate, BG then derive a simplified NKPC

$$\pi_t = \kappa\hat{u}_t + \kappa(1-\delta)(1-x)\,\hat{u}_{t-1} - \Psi\gamma\hat{\alpha}_t \tag{3}$$

with $\kappa = \eta/\delta(1 - u)$, or equivalently

$$\pi_t = -\kappa(1 - (1 - \delta)(1 - x))\,\hat{u}_t - \kappa(1 - \delta)(1 - x)\Delta\hat{u}_t - \Psi\gamma\hat{\alpha}_t. \tag{4}$$

These simple equations are convenient in that they allow us to draw conclusions on the inflation-unemployment trade-off and its relationship with labour market rigidities. In particular, as BG (p. 16) comment "[t]he more sclerotic the labour market, the weaker the effect of the level of unemployment, and the stronger the effect of the change in unemployment."

Moreover, using $\zeta = (1 - \delta)(1 - x)$ as an overall measure of labour market rigidities, we expect the US economy to display high fluidity and low unemployment (i.e., large δ and x). In contrast, the Euro-area labour market is generally considered more sclerotic, i.e., with a larger ζ. If these effects are found in the data, this implies that the BG formulation not only captures the inflation-unemployment trade-off, but it also correctly reflects the nature of labour market rigidities in the economies under study.

As alluded to above, this seemingly simple formulation involves several deep parameters describing the model economy. However, notice that it suffices to estimate (3) and (4) in reduced form, i.e.

$$\pi_t = -\kappa_1\hat{u}_t + \kappa_2\hat{u}_{t-1} + \kappa_3\hat{\alpha}_t + \varepsilon_t \tag{5}$$

$$\pi_t = -\kappa_1^*\hat{u}_t + \kappa_2^*\hat{u}_{t-1} + \kappa_3^*\hat{\alpha}_t + \varepsilon_t^* \tag{6}$$

where $\kappa_1 = -\kappa$, $\kappa_2 = -\kappa\zeta$, $\kappa_1^* = -\kappa(1 - \zeta)$, $\kappa_2^* = -\kappa\zeta$, and $\kappa_3 = \kappa_3^* = -\Psi\gamma$ and ε_t, ε_t^* are uncorrelated disturbances. Estimating (5) allows us to identify separately κ and ζ, while (6) is used to compare the relative magnitude of the effects of the level of unemployment and the change in unemployment for the two economies.[3]

2.2. An extended NKPC with unemployment

The simplicity of the model above is quite appealing from an empirical perspective. Nevertheless, it is also interesting to consider the richer specifications of the unemployment-based NKPC of both BG and RW.[4] The latter also incorporates a theory of unemployment into the NKPC setup with search frictions. Interestingly, both papers lead to a very similar extended NKPC of the form

$$\pi_t = -\beta E_t\pi_{t+1} - \kappa_0\hat{u}_t + \kappa_L\hat{u}_{t-1} + \kappa_F E_t\hat{u}_{t+1} - \kappa_p\hat{\alpha}_t \tag{7}$$

[3] It should be noted that κ itself is a nonlinear function of the separation rate δ and labour market tightness x, which makes identification and subsequent estimation of these parameters difficult without strong assumptions (and calibration) about the remaining structural parameters, something we wish to avoid here.

[4] The BG framework can also be viewed as an employment adjustment cost model and therefore slightly distinct from RW, but for the purpose of our analysis, these differences are not crucial.

where, again, the composite elasticities κ_0, κ_L, κ_F and κ_p are complicated functions of deep parameters describing labour market frictions and nominal rigidities (see BG and RW for details). One of the distinctive features of RW is the inclusion of a 'cost channel' effect, whereby the real interest rate has a direct impact on inflation, thus adding another channel for monetary policy to affect inflation. We will consider this effect in the empirical section, by estimating an extended version of (7), namely

$$\pi_t = -\beta E_t\pi_{t+1} - \kappa_0\hat{u}_t + \kappa_L\hat{u}_{t-1} + \kappa_F E_t\hat{u}_{t+1} - \kappa_p\hat{\alpha}_t + \kappa_p\hat{r}_t + \varepsilon_t \tag{8}$$

where $\hat{r}_t$ is the real interest rate ($i_t - E_t\pi_{t+1}$).

Equations (7)-(8) have interesting implications. First, under sensible values for the underlying structural parameters, we expect the magnitude of κ_0 to dominate κ_L and κ_F. Also, the model predicts that inflation is a lot more responsive to unemployment dynamics if the labour markets are rigid, i.e., we would expect larger (in absolute terms) κ_0, κ_L, and κ_F for the EA when compared to the US.[5] This is consistent with RW's calibration exercise, which suggests that the higher the separation and vacancy rates (higher fluidity), the lower the κ_0, while the higher the labour share of surplus (e.g. in an economy where workers have higher bargaining power), the larger the inflation elasticity with respect to unemployment. Furthermore, RW demonstrate that the magnitude of the cost channel effect depends positively on the rigidity of labour markets, so that we expect a larger κ_R for the EA.

RW estimate an equation in which inflation depends on the probability of filling a posted vacancy, denoted as $\hat{q}_t$ in their paper, rather than on unemployment. While data for $\hat{q}_t$ (the ratio between the job finding probability and labour market tightness) is available for the US, that is not the case for the Euro Area, hence the convenience of using (8) for our comparison. However, the effects go in the same direction: the higher the (contemporaneous) unemployment rate, the higher the probability of filling a posted vacancy and therefore the impact on inflation should be negative, the converse being true for expected values of $\hat{u}_{t-1}$ and $\hat{q}_{t+1}$, as predicted by the models.

Given that we are mainly interested in obtaining empirical estimates of these elasticities for the US and the EA, this reduced-form specification entails little loss of information, as discussed next.

3. **Baseline Estimations**

3.1. Data and methods

In order to empirically quantify the predictions of the models discussed in the previous section, we use quarterly data for the sample period 1970-2007. The start of the sample period is determined by the availability of reliable data for the Euro Area, while we restrict the study up to 2007 because of (public) data availability for synthetic measures of some

[5] Using BG's calibration, we would have κ0 $\simeq$ 0.14, κL $\simeq$ 0.06, and κF $\simeq$ 0.08 for the EA, while for the US the values would be $\kappa_0 \simeq 0.09$, $\kappa_L \simeq 0.02$, and $\kappa_F \simeq 0.06$.

important labour market variables. The sampling period is similar to RW, though, thus ensuring some degree of comparability.

US data is taken from the FRED database at the Federal Reserve Bank of St Louis, namely US inflation measured by the GDP deflator (other measures such as CPI or PCE expenditures produce similar results), the (demeaned) unemployment rate and labour productivity, measured as the log ratio of GDP and total non-farm employment.[6] The corresponding data for the EA comes from the updated Area-Wide Model database of Fagan et al. (2001), with similar transformations applied.

Endogeneity issues usually plague NKPC formulations, namely due to the presence of expectations in these models (which lead to forecast errors picked up by the error term), errors-in-variables due to the use of proxies (for marginal costs, productivity, etc.) and because shocks affecting inflation are likely to be correlated with the driving variables as well. To deal with these problems, estimation is usually carried out by GMM, although this itself entails several problems (see Martins and Gabriel, 2009; and Kleibergen and Mavroeidis, 2009). Indeed, inference concerning the NKPC is plagued with the 'many weak instruments curse', partly because any information these instruments may contain on future inflation fluctuations would have already been exploited by the central banks to contain inflation. In practice, this leads to a very limited number of (arbitrarily chosen) instruments being used, resulting in a loss in efficiency.

To attenuate these difficulties, we suggest using the procedure proposed by Kuersteiner and Okui (2010), which allows us to construct "optimal instruments" by applying a model averaging (MA) approach to the first stage of two-stage least squares (2SLS). This procedure has several advantages: i) it delivers a more favourable trade-off between bias and efficiency relative to estimators that rely on a single set of instruments; ii) no ad-hoc choice of instruments or weak instruments pre-testing must be entertained, and iii) it possesses good finite sample properties even when there are many weak instruments available, which is likely to be our case.

The weights for first-stage model averaging are chosen to minimise the asymptotic mean squared error and are found numerically as the solution of a quadratic programming problem. We use the MA-2SLS$_P$ version, where the weights are constrained to be positive, in the interval [0,1] (see Kuersteiner and Okui, 2010, for details). For comparison, however, we also present results from GMM estimation with a heteroskedasticity and autocorrelation consistent Newey-West weight matrix.

For conciseness, we focus on a baseline instrument set containing two lags of the variables appearing explicitly in each estimated model (inflation rate, unemployment, productivity and real interest rates in the case of (8)), plus the labour share, commodities price inflation, wage inflation, interest rate spread and HP-filtered real GDP, thus following the previous literature on the NKPC (e.g. Galí and Gertler, 1999; and Ravenna and Walsh, 2008). Although results vary little when more lags are included (or some variables excluded), we found that the Stock and Yogo (2005) 2SLS-bias based test usually failed to reject the null

[6] We use productivity in levels, as in RW. These authors show that it is relatively straightforward to allow for long-run productivity growth in their setup. Detrending productivity, regardless of the detrending/filtering method, results in the corresponding estimated elasticity being positive, which is not consistent with the theoretical predictions. Also, using productivity defined as output per hour (only possible for the US) delivers almost identical results.

of weak instruments (though, strictly speaking, this test is valid for IV estimation only). For the MA-2SLS estimation, in particular, changes in the instrument set had very little impact on final results, mainly due to the first-step averaging, which underlines the robustness of this procedure.

3.2. Results

Table 1 presents the estimation results for equations (5)-(6) under this baseline setting. A general assessment reveals that the MA-2SLS estimator tends to deliver slightly smaller, but more precisely estimated, elasticities (a general feature throughout the paper). Also, parameter estimates are broadly similar across the different estimators, which suggests that results are relatively robust. Under column Sargan-*J*, we see that tests of over-identifying restrictions (the usual J-test for GMM and a Sargan-type test for MA-2SLS) were also satisfied.

Table 1: Reduced-form Phillips Curve, equations (5)-(6), 1970:1-2007:4

	From eq. (5)						
	κ_1^*	κ_2^*	κ_3^*	κ	ζ	Sargan-*J*	Stock-Yogo
US							
GMM	0.184 (0.066)	0.710 (0.296)	– 0.014 (0.005)	– 0.894 (0.321)	0.794 (0.077)	0.729	8.819
MA-2SLS$_p$	0.252 (0.032)	0.460 (0.151)	– 0.003 (0.001)	– 0.511 (0.148)	0.899 (0.002)	1.000	§
Euro-area							
GMM	0.134 (0.021)	1.242 (0.334)	– 0.036 (0.004)	– 1.376 (0.337)	0.903 (0.026)	0.614	12.861*
MA-2SLS$_p$	0.191 (0.021)	0.950 (0.246)	– 0.029 (0.003)	– 1.018 (0.242)	0.933 (0.033)	1.000	§

Notes: Standard errors (s.e.) in parentheses; nonlinear Wald-type s.e. in the case of ζ; the Sargan-*J* column refers to *p*-values of tests for over-identifying restrictions; Stock-Yogo refers to the Stock and Yogo (2005) statistic (based on 2SLS bias) for the null hypothesis of weak instruments; * denotes rejection at the 10% significance level (critical value: 10.47); § denotes the same statistic.

As explained above, we are interested in comparing the magnitudes of ζ and the κ^*'s for the two economies, which should reveal the relative degree of labour market sclerosis. The results lend some support to the view that the US labour market appears to be more fluid than the Euro-area one and, therefore, the nature of the inflation-unemployment trade-off is distinct for these two economies. The point estimate for the composite parameter ζ that we derive from the estimations is lower for the US, but not substantially so. Indeed, although we cannot test whether or not the coefficients are the same for the two economies, we observe that the implicit confidence intervals overlap, so we cannot confidently say that this measure of overall labour market rigidity is significantly different for the US and the Euro-area.

However, we find the impact of current unemployment on inflation (κ_1^*) to be stronger in the US case relative to $\Delta\hat{u}_t$, while the effect of changes in unemployment (κ_2^*) is relatively larger for the Euro-area: taking κ_1^* as a proportion of κ_2^*, the lowest value for the US is 0.26, while the highest for the Euro-area is 0.20. Interestingly, a direct comparison of the coefficients of the two economies agrees with this conclusion. In fact, κ_1^* tends to be larger for the US, while κ_2^* dominates in the Euro-area case.

Note also that the coefficient κ_3^* on labour productivity delivers the expected sign for both economies, indicating a negative relationship with inflation. This reflects in-sample underlying trends for these variables, with inflation displaying a long decline, while productivity has steadily grown. Noticeably, our estimates indicate that this variable has a smaller effect on US inflation than on European inflation (although broadly in line with the results in RW). Recall that the (composite) labour productivity elasticity is an increasing function of the degree of real wage rigidit ies, so these results reinforce the view that labour market frictions play a more significant role in European inflation dynamics. In the next section, we will attempt to appraise the role of productivity in more detail.

Table 2: BG-RW extended model, baseline case

	Model: $\pi_t = \beta E_t\pi_{t+1} + \kappa_0^*\hat{u}_t + \kappa_L^*\hat{u}_{t-1} + \kappa_F^* E_t\hat{u}_{t+1} + \kappa_p^*\hat{\alpha}_t + \kappa_R^*\hat{r}_t$						
	β	κ_0^*	κ_L^*	κ_F^*	κ_p^*	κ_R^*	Sargan-J
US							
GMM	0.877 (0.038)	– 0.982 (0.412)	0.336 (0.195)	0.694 (0.231)	– 0.001 (0.0004)		0.589
MA-2SLS$_p$	0.804 (0.058)	– 0.010# (0.120)	0.132 (0.080)	0.171 (0.073)	– 0.004# (0.003)		0.835
GMM	0.903 (0.057)	– 1.150 (0.382)	0.529 (0.174)	0.688 (0.240)	– 0.008 (0.003)	0.028 (0.014)	0.438
MA-2SLS$_p$	0.806 (0.058)	– 0.013# (0.126)	0.127# (0.085)	0.139 (0.081)	– 0.003# (0.003)	0.014 (0.012)	0.985
Euro Area							
GMM	0.953 (0.089)	– 0.974# (1.453)	0.131# (0.663)	– 0.861# (0.817)	– 0.000# (0.003)		0.277
MA-2SLS$_p$	0.616 (0.110)	– 0.470# (0.442)	– 0.166# (0.238)	0.623 (0.261)	– 0.011 (0.004)		0.361
GMM	0.843 (0.071)	– 3.440 (1.025)	– 1.819 (0.502)	– 1.660# (0.551)	0.001# (0.001)	0.033 (0.015)	0.507
MA-2SLS$_p$	0.841 (0.220)	– 3.876# (3.085)	– 2.003# (1.480)	– 1.941# (1.645)	0.006# (0.007)	0.057# (0.036)	0.367

Notes: See notes to Table 1; # denotes not significant at the 10% significance level; $\hat{r}_t$ is measured as the short term (3-month) interest rates minus the inflation rates in period t+1; Stock-Yogo critical values are only available for models with up to three endogenous regressors.

Turning to the extended BG-RW specification discussed in section 2.2, estimates of equations (7)-(8) are presented in Table 2.[7] The most noticeable result is the fact that, although the relative magnitude of the coefficients is the anticipated one (with κ_0^* dominating κ_L^* and κ_F^* and these elasticities being larger for the EA than the US), the empirical adherence of this extended specification is somewhat weak. Indeed, several of the estimated coefficients are statistically insignificant and in some cases the signs of the coefficients are not in accordance with the model. This is particularly more severe in the case of the EA, where most parameters are not significant and κ_L^* and κ_F^* display the wrong (negative) sign. Interestingly, the productivity elasticity κ_p^* is estimated to be smaller, for both economies, than in the simpler version estimated in Table 1, while the cost channel given by κ_R^* is non-negligible, although smaller than that reported in RW for the US (ranging between 0.07 and 0.15).

[7] We follow the typical practice of parodying expected values of inflation (and unemployment here as well) by realised values, which naturally introduces an additional measurement disturbance in the error term. See Dufour et al. (2006), who consider using survey-based measures of expectations, as well as Kleibergen and Mavroeidis (2009) and discussions therein.

Also, the results differ little across GMM and the MA estimator, although the latter tends to produce more insignificant coefficients.

These difficulties are not wholly unexpected, as the strong persistence in unemployment rates, especially for the EA (autocorrelation coefficient in excess of 0.95), means that it will be difficult to distinguish between the effects of current unemployment and lags/leads of this series on inflation, thus leading to imprecision in the estimations. This is not inconsistent with the relative success of the baseline specification estimated in Table 1, as it only depends on current unemployment. In the next section, we consider ways of improving the empirical fit of these specifications.

4. Extensions

The baseline results discussed above focus on a simple implementation of BG, which nevertheless produces interesting insights, broadly consistent with the stylised facts about labour market rigidities in the US and the Euro-area and that are robust across different estimation methods. Nevertheless, the strict adherence to the theoretical framework of BG and RW is somewhat limiting from an empirical perspective. Thus, we next consider a few extensions that add realism to our empirical exercise, analysing their individual impact on the baseline estimations.[8]

4.1. Time-varying nairu

Albeit convenient from a theoretical point of view, the assumption of a constant steady-state level of unemployment ($\bar{u}_t$) in the specification above is restrictive and runs counter the evidence of fluctuations in estimated NAIRUs, both for the US and the Euro Area (see Staiger et al., 1997; and Fabiani and Mestre, 2000 and 2004). Thus, we explore how the baseline results are affected by allowing for this source of variation in $\bar{u}_t$.

Several methods for estimating the NAIRU are possible and there is considerable uncertainty about its true level. We considered "official" estimates of the NAIRU, as well as estimates based on our sample. In the case of the US, the Congressional Budget Office publishes a NAIRU based on their potential output estimates. In addition, the OECD also publishes annual estimates of the NAIRU for both the US and the Euro Area. In the latter case, estimates are only available from 1991 onwards, so we use estimates from Fabiani and Mestre (2004) for the earlier part of the sample. Given that these NAIRU measures only display annual variation, the resulting quarterly stepwise deviations are HP-filtered to create a smoother series. It turns out that the resulting measures of $\bar{u}_t$ are highly correlated with those obtained from simple HP-filtering or quadratic detrending of the corresponding unemployment rate series (correlations in excess of 0.8). To save space, while ensuring consistency for both economies, we report results using the OECD-based NAIRU deviations (results vary little when other measures are employed).

[8] In what follows, the instruments are naturally based on the lags of the new variables considered under each specification.

Table 3: Equations (5)-(6) with new variables

Using NAIRU as $\bar{u}_t$ (with baseline $\hat{\alpha}_t$)					From eq. (3)			
		κ_1^*	κ_2^*	κ_3^*	κ	ζ	Sargan-J	Stock-Yogo
US	GMM	0.222 (0.072)	0.683 (0.297)	– 0.014 (0.005)	– 0.905 (0.320)	0.755 (0.049)	0.736	13.988*
	MA-2SLS$_p$	0.267 (0.034)	0.485 (0.151)	– 0.004 (0.0002)	– 0.551 (0.147)	0.879 (0.002)	1.000	§
Euro Area	GMM	0.196 (0.031)	1.566 (0.314)	– 0.049 (0.002)	– 1.762 (0.324)	0.889 (0.021)	0.553	13.656*
	MA-2SLS$_p$	0.194 (0.021)	1.429 (0.178)	– 0.037 (0.003)	– 1.573 (0.531)	0.909 (0.029)	1.000	§
US only								
Using Fernald TFP as $\hat{\alpha}_t$ (with baseline $\bar{u}_t$)								
	GMM	0.203 (0.069)	0.751 (0.315)	– 0.006 (0.002)	– 0.954 (0.345)	0.787 (0.073)	0.699	13.991*
	MA-2SLS$_p$	0.257 (0.026)	0.401 (0.155)	– 0.0001 (0.0001)	– 0.559 (0.155)	0.719 (0.001)	0.999	§
NAIRU as $\bar{u}_t$ and Fernald TFP as $\hat{\alpha}_t$								
	GMM	0.237 (0.076)	0.667 (0.312)	– 0.005 (0.002)	– 0.904 (0.339)	0.738 (0.097)	0.714	14.132*
	MA-2SLS$_p$	0.194 (0.036)	0.410 (0.174)	– 0.001 (0.000)	– 0.504 (0.171)	0.813 (0.001)	1.000	§
NAIRU as $\bar{u}_t$ and Fernald TFP as $\hat{\alpha}_t$, plus separation rate and tightness								
US	κ_1^*	κ_2^*	κ_3^*	κ	ζ	tightness	sep. rate	Sargan-J
GMM	0.324 (0.161)	0.802 (0.280)	– 0.003 (0.001)	– 1.113 (0.376)	0.721 (0.080)	0.021 (0.002)	0.810 (0.132)	0.802
MA-2SLS$_p$	0.308 (0.034)	0.432 (0.164)	– 0.004 (0.001)	– 0.540 (0.103)	0.801 (0.001)	0.011 (0.003)	1.033 (0.107)	1.000
NAIRU as $\bar{u}_t$ and tightness								
Euro-area								
GMM	0.215 (0.078)	1.478 (0.259)	– 0.052 (0.005)	– 1.792 (0.256)	0.825 (0.083)	0.088 (0.028)	–	0.705
MA-2SLS$_p$	0.224 (0.088)	1.388 (0.431)	– 0.041 (0.005)	– 1.612 (0.406)	0.861 (0.072)	0.011 (0.033)	–	1.000

Notes: See notes to Table 1. The first panel reports estimates using the NAIRU as $\bar{u}_t$ and labour productivity as $\hat{\alpha}_t$ (baseline case); the middle panel reports estimates for the US using the utilisation-adjusted TFP measure, first with the unemployment rate as $\bar{u}_t$ (baseline case); the bottom panel includes labour market tightness (and separation rates for the US) as additional regressors.

The results are shown in the first panel of Table 3, with no noticeable differences when compared to the baseline results of Table 1. Indeed, the magnitude of the κ_i^*, κ and ζ coefficients is similar. This suggests that controlling for a time-varying steady-state unemployment level does not change the nature of the baseline results discussed above. However, given that it offers a more realistic interpretation of movements in steady-state levels of unemployment, we will also consider this measure for the extended BG-RW specification further below.

4.2. The role of productivity measures

The theoretical framework outlined above does not include capital and defines a production function that is linear in labour, which implies that labour productivity coincides with total factor productivity (TFP). Ideally, one would use measures of "exogenous TFP", but it is well known that TFP as conventionally calculated may be mismeasured, due to variable input utilisation, non-constant returns to scale, reallocation effects, etc. (Basu, Fernald and Kimball, 2006). These effects make measured TFP endogenous and may lead to biased coefficients in estimated Phillips curves, as we found in the baseline estimation results. Suppose cycles are primarily caused by non-technology shocks, such as demand shocks. A positive demand shock raising marginal costs induces an increase in inflation, but it also increases observed productivity due to an increase in utilisation.

To attenuate these potential biases, we re-estimate the baseline BG model using the utilisation-adjusted quarterly TFP series constructed by Fernald (2012). This variable still omits some of the corrections that may be done with annual data, but it is, as far as we know, the best proxy available for true "exogenous" TFP.

Unfortunately, neither standard nor utilisation-adjusted TFP measures are available for the Euro Area at a quarterly frequency and for our sample period. Existing measures are annual from 1980 onwards and do not include all Euro Area countries.[9] Thus, our analysis focuses on the US case only. We conjecture that the estimation results, relative to those using labour productivity, would be probably similar had a TFP measure been available for the Euro Area. Indeed, the correlation between the available annual TFP growth measure and labour productivity growth is 0.71, while the correlation between the corresponding US variables at the quarterly frequency is 0.70.

The middle panel of Table 3 contains results for estimations based on the Fernald (2012) measure of US total factor productivity. The first set reports results using unemployment rates, while the second set considers the NAIRU deviations as $\bar{u}_t$, as discussed above. At a first glance, estimates are largely similar to the baseline results of Table 1 and the top panel of Table 3, regardless of the measure for $\bar{u}_t$. However, the coefficient κ_3^* associated with $\hat{\alpha}_t$ is now estimated to be considerably smaller, thus suggesting that "purified" technology shocks, while relevant, appear to play a minor role in explaining the dynamics of US inflation. Also, note that ζ is now estimated to be smaller than in the baseline estimations, indicating that the degree of labour market rigidities is smaller once corrected measures of technology shocks are employed. Again, this is consistent with the idea that labour productivity will

[9] Data available from the EU KLEMS database.

include effects on the intensive margin that may lead to higher decreases in marginal costs and therefore have a more negative impact on inflation, but in a way that it also biases the coefficients associated with unemployment. This suggests that using a "purified" TFP measure is important in an empirical setup, to ensure consistency with the theory and to avoid misspecifications.

The presence of productivity in the NKPC, although desirable from a theoretical point of view, raises some empirical difficulties, given the absence of precise measures of productivity. For robustness, we also considered a restricted version of the baseline specification that allows us to obtain estimates of some parameters of interest, namely ζ, without the use of $\hat{\alpha}_t$. Indeed, assuming that $\hat{\alpha}_t$ follows an AR(1) process with autoregressive coefficient ρ, then (3) can be rewritten as

$$\pi_t = \rho\pi_{t-1} + \phi_1\hat{u}_t + \phi_2\hat{u}_{t-1} + \phi_3\hat{u}_{t-2} \tag{9}$$

where $\phi_1 = -\kappa$, $\phi_2 = \kappa(\zeta + \rho)$ and $\phi_3 = -\rho\kappa\zeta$ so that π_t is now a function of current and two lags of unemployment, with a more traditional 'intrinsic' persistence component in π_{t-1}.

Table 4: Restricted version, Eq. (9), GMM estimation

	ρ	ϕ_1	ϕ_2	ϕ_3	ζ	Sargan-J
Baseline						
US	0.839 (0.058)	– 0.865 (0.214)	1.449 (0.326)	– 0.607 (0.131)	0.836 (0.142)	0.244
Euro-area	0.934 (0.022)	– 0.841 (0.388)	1.465 (0.709)	– 0.635 (0.328)	0.814 (0.128)	0.237
New variables						
US	0.852 (0.037)	– 0.582 (0.221)	0.863 (0.344)	– 0.343 (0.131)	0.743 (0.114)	0.616
Euro-area	0.909 (0.021)	– 1.432 (0.368)	1.583 (0.651)	– 1.164 (0.299)	0.894 (0.093)	0.338

Notes: See notes to Table 1; MA-2SLS and Stock-Yogo tests are not available in the case of nonlinear estimation.

Table 4 reports estimation results when we directly recover the parameters using their nonlinear relationships, both for the baseline variables of Section 2 (first panel) and when we use the NAIRU and TFP measures. Again, the results, although slightly less precise, are remarkably consistent with previous estimations, namely for the relative magnitudes of κ and ζ for the two economies. Note that using the variables introduced in this section improves the fit and precision in these estimations. Moreover, it is interesting to observe that the persistence of productivity, as captured by ρ, is higher for the EA, which helps to explain the larger (negative) impact of productivity on inflation noted previously, as one could anticipate from the BG setup. Thus, it seems sensible to include measures of productivity in estimations of the NKPC, as the biases stemming from the use of proxies for $\hat{\alpha}_t$ appear to be relatively modest.

4.3. Labour market variables

As discussed above, a potential shortcoming of our baseline exercise is the fact that we focus on (constant) steady-state levels of some relevant variables. Note that from (2) there is a direct relationship between tightness, separation rates and unemployment rates, which suggests that they could be used interchangeably to study the effects on inflation dynamics. However, as explained above, there can be significant discrepancies between the theoretical framework and its empirical implementation. Indeed, it is likely that cyclical variations in labour market tightness and separation rates, for example, in addition to their influence in unemployment dynamics, may have additional explanatory power in determining the inflation-unemployment trade-off.

Indeed, as discussed in Shimer (2005), the cyclical component of labour market tightness (the vacancy-unemployment ratio) displays a great deal more volatility than predicted by standard models (and than most of the variables studies so far). On the other hand, while there has been some debate about the cyclicality of separation rates (Hall, 2005), recent literature (Fujita and Ramey, 2009 and 2012; Elsby et al., 2009) emphasises the contribution of changes in separation rates to unemployment fluctuations. Thus, it is an interesting exercise to assess how variations in these key labour market variables affect inflation dynamics, even after we control for variation in unemployment and productivity.[10]

Measures of labour market tightness and time varying separation rates are readily available for the US (see Shimer, 2005; Shimer, 2012; and the author's webpage).[11] For the Euro Area, data availability is an issue for all variables that are not present in the AWM database. Nevertheless, we were able to construct a time series for labour market tightness (the vacancies to unemployment ratio) in the Euro Area for our sample period, by constructing a vacancies index (based on data available for approximately two thirds of the EA countries, as in Christoffel et al., 2009) and, using the same methodology to construct EA-wide unemployment levels.[12] Following Shimer (2005), we extract the cyclical component of this measure by employing the HP filter with smoothing parameter 10^5 (a similar transformation is used for the US separation rates).

We then gauge the contribution of these variables in two concurrent ways. First, we add these variables to the baseline BG equation, therefore directly controlling for variation in the key labour market variables. Second, indirectly, by adding lags of labour market tightness (and separation rates in the case of the US) to the instrument set, thus utilising correlations of these variables with unemployment and productivity measures to obtain potentially better estimates of the inflation-unemployment trade-off.

Results for this exercise are displayed in the bottom panel of Table 3. Including these new variables generally improves the fit and estimation precision, but in a way that reinforces the initial conclusions. Indeed, *i*) the composite parameter ζ measuring the degree of

[10] Note that this echoes Ravenna and Walsh (2008) NKPC specification that is explicitly written in terms of the probability of filling a vacancy, itself a function of labour market tightness.

[11] For labour market tightness, US vacancies were constructed by splicing the Help Wanted Advertising Index used in Shimer (2005) and data from the Job Openings and Labor Force Turnover Survey (JOLTS).

[12] The composition of the EA is updated as data for both vacancies and unemployment levels become available for each country (source: OECD Statistics).

rigidities is lower for the US than the EA, but more so than in the baseline estimation, while the relative magnitude of the κ_i^*'s support the premise of the BG model that the more sclerotic the labour market, the stronger the elasticity of $\Delta\hat{u}_t$ relative to the effect of $\hat{u}_t$. The lowest ratio of estimated κ_1^* to κ_2^* is now 0.40 for the US, while for the EA the highest ratio is 0.16.

It is also interesting to check the direct impact of labour market tightness on inflation. The tightness coefficient, although estimated to be relatively small (but larger than productivity's in absolute terms for the US), is highly significant and positive, as expected.[13] This indicates that there may be further cyclical effects in labour market tightness influencing inflation, but that are not captured solely by deviations from the NAIRU.

This is further supported by the inclusion of time-varying separation rates in the case of the US economy. Allowing for a time-varying separation rate uncovers additional responses of inflation to labour market conditions, with a sizeable and significant estimated coefficient. This can be explained by the fact that both the US inflation and separation rates appear to share a secular decline (even after HP filtering the latter variable), particularly after the early 1980's. This decline in separation rates was noted by Shimer (2012), but not previously related to inflation dynamics.

4.4. The extended RW-BG specification with new variables

We now assess how the empirical modifications discussed above impact on estimates of the extended RW-BG specification. In Table 5 we present results with NAIRU as $\hat{u}_t$ for both the US and EA, Fernald's (2012) TFP as $\hat{\alpha}_t$ for the US and adding labour market tightness (and separation rates for the US) to the instrument set. Doing so delivers a significant improvement in terms of fit compared with Table 2, particularly for the US case, with all coefficients now statistically significant, with the correct signs and with sensible magnitudes, and similar for both the GMM and MA estimators. Indeed, we find a one-to-one relationship between inflation and current deviations from the NAIRU, with $\kappa_0 > \kappa_F > \kappa_L$, as expected. As before, the effects of productivity and real interest rate fluctuations are mostly significant, but relatively small. The results are also consistent with the RW estimations based on $\hat{q}_t$, as discussed in section 2.2.

[13] This is consistent with the finding of Ravenna and Walsh (2008), who consider the (negative) effects on inflation of the probability of filling a posted vacancy, itself inversely proportional to labour market tightness.

Table 5: BG-RW extended model, new variables

	Model: $\pi_t = \beta E_t \pi_{t+1} + \kappa_0^* \hat{u}_t + \kappa_L^* \hat{u}_{t-1} + \kappa_F^* E_t \hat{u}_{t+1} + \kappa_p^* \hat{\alpha}_t + \kappa_R^* \hat{r}_t$						
	β	κ_0^*	κ_L^*	κ_F^*	κ_p^*	κ_R^*	Sargan-J
US							
GMM	0.931 (0.030)	– 0.943 (0.364)	0.350 (0.180)	0.481 (0.203)	– 0.001 (0.0001)	–	0.820
MA-2SLS$_p$	0.787 (0.053)	– 1.015 (0.117)	0.252 (0.081)	0.415 (0.070)	– 0.002 (0.001)	–	0.986
GMM	0.861 (0.024)	– 1.089 (0.262)	0.424 (0.121)	0.681 (0.145)	– 0.001 (0.0001)	0.007 (0.004)	0.716
MA-2SLS$_p$	0.785 (0.052)	– 1.008 (0.121)	0.411 (0.085)	0.512 (0.078)	– 0.002 (0.001)	0.017 (0.012)	0.975
Euro Area							
GMM	0.913 (0.076)	– 2.865 (1.148)	1.10 (0.541)	1.803 (0.636)	– 0.002# (0.003)	–	0.404
MA-2SLS$_p$	0.693 (0.115)	– 0.899 (0.266)	0.114# (0.174)	0.749 (0.144)	– 0.011 (0.006)	–	0.862
GMM	0.759 (0.088)	– 3.846 (1.301)	1.563 (0.653)	2.305 (0.678)	– 0.010 (0.004)	0.007# (0.011)	0.389
MA-2SLS$_p$	0.753 (0.111)	– 0.709 (0.325)	0.103# (0.202)	0.502 (0.161)	– 0.013 (0.005)	0.055 (0.017)	1.000

Note: See notes to Table 2.

In the case of the EA, the empirical adequacy of this model is also improved, although some coefficients are occasionally insignificant (in particular κ_L with the MA estimator). The unemployment elasticities are estimated to be larger than those of the US, as expected, given that the κ's are proportional to the degree of labour market rigidities. Also, both their relative magnitudes and their 'net effect' on impact (i.e. $\kappa_0 - \kappa_F - \kappa_L$) are in accordance with the calibration exercises in BG and RW. The improvements in fit appear to be due to the use of additional information concerning labour markets and the fact that deviations from the NAIRU are slightly less persistent than unemployment rates, thus helping to discriminate the dynamic effects of $\hat{u}_t$ on inflation.

5. **Conclusion**

Although the inflation-unemployment trade-off is at the heart of modern macroeconomics, most empirical studies consider alternative drivers of inflation dynamics in competing specifications, namely measures of real activity such as (proxies for) marginal cost or output gaps. However, empirical support for these relationships is mixed, at best. This paper attempts to shed light on the empirical usefulness of unemployment-based NKPCs stemming from a fully micro-founded framework, more particularly in explaining observed differences between the US and the Euro Area economies. Although our study focuses on a simple implementation of Ravenna and Walsh (2008) and Blanchard and Galí (2010), it nevertheless produces interesting insights.

We find that these specifications are broadly consistent with the stylised facts about inflation persistence and labour market rigidities in the US and the Euro-area. More precisely, we found that, once appropriate adjustments are made and the informational content of relevant labour market variables (such as the NAIRU, time-varying separation rates and labour market tightness) is explored, an unemployment-based NKPC, both in its simple and extended form, produces results that are in line with the theoretical predictions. Indeed, our results are robust across different estimation methods and show that unemployment and productivity elasticities are larger for the EA compared to the US. Given that these elasticities reflect underlying labour markets characteristics, this, in turn, is consistent with the view that the US labour market is considerably more fluid that the Euro-Area one, as discussed by Jolivet et al. (2006), Hobijn and Sahin (2009), and Elsby et al. (2013).

It is important to recognise that several of the limitations identified in empirical studies of the NKPC are also present here. First, there are clear identification difficulties, with the specifications studied here depending on a considerable number of deep parameters that cannot be recovered in a single-estimation setup. The fact that our interest is on 'reduced-form' elasticities mitigates, but cannot obviate, this fact. Second, the models studied here depend on expectations for inflation and unemployment and also measures of productivity, all of these difficult to measure or observe accurately.

Nevertheless, we suggest that exploring the additional information provided by labour market variables may help to understand inflation rate dynamics and, therefore, to better inform economic policy. This is, in fact, consistent with the current practice of some central banks, such as the "forward guidance" principle, which puts labour market conditions at the centre of monetary policy decisions.

An interesting challenge would be to study the inflation-unemployment relationship in the aftermath of the Great Recession. Indeed, inflation has remained relatively stable, despite the increase in unemployment, though this has also been accompanied by a decline in productivity. Given that the models studied here hinge on the relationship among inflation, unemployment and productivity, they will provide a useful tool in disentangling the contribution of these effects to the recent modest and protracted decline of inflation rates. Thus, once more data becomes available, further empirical research on this topic would clearly be a worthwhile pursuit.

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