The Phillips Curve and Long-Term Unemployment

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Abstract

This paper studies the impact of long-term unemployment on the dynamics of inflation. Labor market theories such as insider-outsider models predict that this type of unemployed are less relevant in the price formation process than the newly unemployed. In this paper I investigate whether evidence of this behavior is present in a set of 19 OECD countries. For this purpose, I propose a new way to specify the Phillips Curve that allows for different unemployment lengths to enter the model in a time-varying NAIRU setting. This is done by constructing an index of unemployment that assigns different weights to the unemployed based on the length of their unemployment spell. The results show that unemployment duration does matter in the determination of inflation, and that a smaller weight ought to be given to the long-term unemployed. Cross-country differences in the results can be explained by a number of labor market institutions. This modified model has important implications for the policy maker: It produces more accurate forecasts of inflation and more precise estimates of the NAIRU.

Key Words: Long-term unemployment, Phillips curve, NAIRU, Kalman filter.

JEL Classification: C22, E31, E50, J64

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1 Introduction

Over the last two decades, one of the most important labor market phenomena in many developed countries has been the emergence and persistence of long-term unemployment.[1] Starting in the early 1980s, the number of long-term unemployed in many OECD countries soared in relation to the already growing number of unemployed. As a result, considerable research has been devoted to study issues such as the process leading to long-term unemployment, its effects, and possible solutions.[2]

The objective of this paper is to study the impact of long-term unemployment on the dynamics of inflation. This is an important issue because the inverse short-run relationship between price changes and unemployment, as captured by the Phillips Curve and the NAIRU (the Non-Accelerating Inflation Rate of Unemployment), is widely used by policymaking institutions to assess the desired stance of monetary policy and to forecast inflation (Boone *et al*, 2002). However, in the presence of long-term unemployment, the aggregate rate of unemployment may provide a distorted measure of the true demand pressures exerted on inflation. On this subject, the OECD argues that when long-term unemployment is high "...unemployment becomes a poor indicator of effective labor supply, and macroeconomic adjustment mechanisms- such as downward pressure on wages and inflation when unemployment is high- will then not operate effectively..." (OECD, 2002, p.189). The argument rests on the assumption that the long-term unemployed play an unimportant role in the setting of prices and wages. This has a number of important implications for the policy maker: If the long-term unemployed become less relevant to price formation, then the downward pressure of unemployment on prices decreases and unemployment becomes more persistent (Blanchard and Wolfers, 2000). Furthermore, if long-term unemployment is high, a given reduction in inflation may require extra contractionary measures as the pool of long-term unemployed will not contribute much to bringing inflation down.

In this paper I provide evidence of the role that unemployment duration plays in the determination of inflation using a set of nineteen OECD countries. This is the first paper that undertakes such a systematic, multi-country study. In the spirit of Nickell (1987) and

[1] Following the preferred OECD terminology, I will define as long-term unemployed those individuals in the labor force who have been out of work for one year or longer. Short-term unemployed will be those out of work for less than one year.

[2] For a more comprehensive analysis of the trends, incidence and composition of long-term unemployment see OECD (1983, 1987, 2002) and Layard *et al* (1991). Machin and Manning (1999) survey the literature on long-term unemployment.

Manning (1994), I propose a modified version of an otherwise standard Phillips Curve model that allows for different unemployment lengths to enter the estimation. This is done by constructing an index of unemployment that assigns different weights to the unemployed based on the length of their unemployment spell. These weights are a measure of the impact that the unemployed have on inflation. This deviates from the standard practice of using the aggregate unemployment rate.[3] Optimal weights are determined by the estimation of the model by maximum likelihood using the Kalman filter. The use of the Kalman filter enables the estimation of a time-varying NAIRU. This is an important point of departure from Nickell (1987) and Manning (1994), who assume a constant NAIRU.

The results obtained show that unemployment duration is important for the understanding of inflation dynamics, and that a smaller weight ought to be given to the long-term unemployed. The results also show that in those countries where long-term unemployment is high (namely, some Western European countries), the long-term unemployed play little role in the setting of prices and wages. This contrasts with non-European OECD countries, where all the unemployed have similar impact, regardless of the length of their spell. These cross-country variations can be explained by some of the institutions that characterize labor markets in the OECD, such as union coverage levels and employment protection.

Insofar as the monetary authority employs Phillips Curve models and the NAIRU to asses inflationary pressures and to forecast inflation, the results in this paper are relevant to the policy maker. That is, by looking at a break down of unemployment in terms of duration, the policy maker receives more accurate information concerning inflationary developments. As the results will further show, this modified version of the Phillips Curve produces more accurate forecasts of inflation at both the one-year and two-year horizons, and generates more precise estimates of the NAIRU, with an average improvement of around 20 percent.

The paper is organized as follows. Section 2 reviews the evolution of unemployment in the OECD and possible explanations. Section 3 presents the baseline and modified econometric models and discusses a number of estimation issues. Section 4 lays out the main empirical results of both models. Section 5 relates the results to a number of labor market institutions. Section 6 checks for robustness of the results. Section 7 concludes.

[3] The standard unemployment rate gives equal weight to all the unemployed, regardless of the length of their spell

2 Evolution and Studies of Unemployment in the OECD

Figure 1 shows the evolution of unemployment in the OECD over the last two decades. There are three periods of pronounced growth in the aggregate unemployment rate (bold line). These are related to the global recessions of the 1970s and to the slowdown of the early 1990s. The individual country data show some important differences across countries. While countries outside Europe have been able to maintain relatively low and stable levels of unemployment, Western European countries have, for the most part,[4] suffered from persistently high and volatile levels of unemployment. Unemployment in Europe quickly jumped from 2.9 percent in 1974 to a peak of nearly 10.5 percent in 1985, remaining at high levels for the rest of the decade. On the other hand, growth in unemployment outside Europe was much less pronounced and by the end of the 1980s it was back to its original level. The global slowdown of the early 1990s also had some important and interesting effects: While it caused another big increase in unemployment in Europe, it was short-lived and relatively painless outside Europe.

A large number of studies have attempted to explain the behavior of unemployment across OECD countries (see Bean, 1994; Nickell, 1997; Siebert, 1997; Blanchard and Wolfers, 2000; Ljungqvist and Sargent, 1998). These studies claim that the emergence of long-term unemployment provides an insight into these unemployment experiences.[5] From Figure 1, it is easy to see that most of the unemployment growth in the OECD can be attributed to an increase in long-term unemployment. Its rate quickly jumped from less than 1 percent in 1976 to almost 3 percent in 1985, remaining at high levels ever since. On the other hand, the short-term unemployment rate moved around its long-run trend. These studies also argue that the incidence of long-term unemployment was much higher in the European countries.[6]

2.1 Studies on Long-Term Unemployment

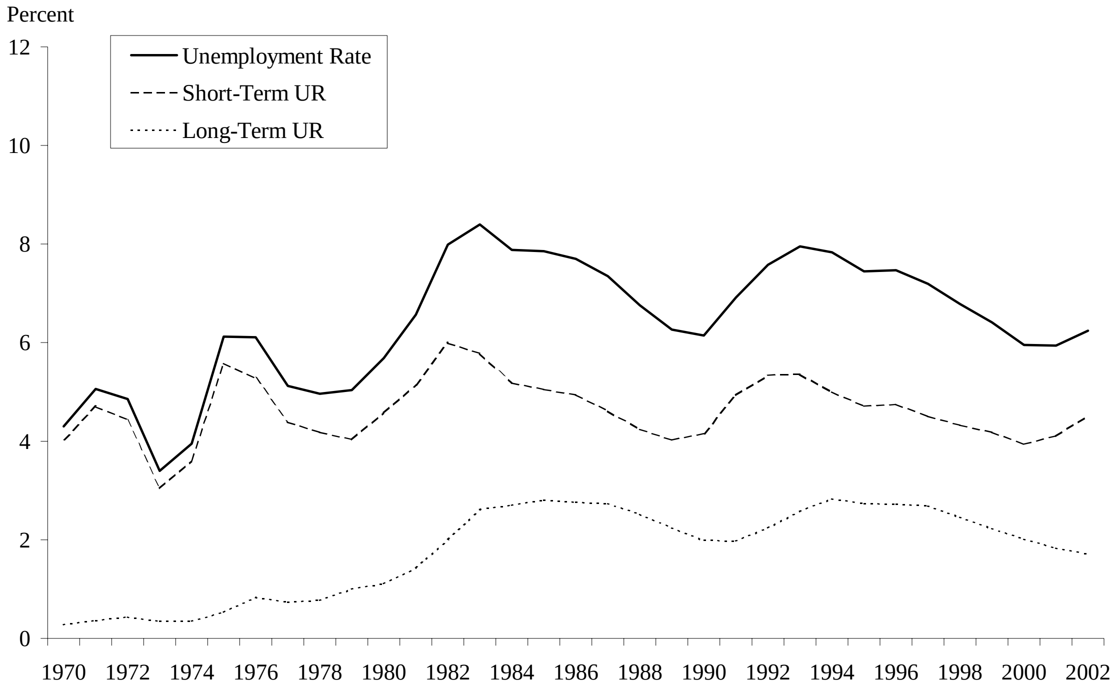
The transition from unemployment to long-term unemployment has spawned an abundant literature in labor economics seeking to provide microeconomic foundations to the problem.

[4] Even within the group of European nations, the behavior of unemployment has displayed very little homogeneity across countries. Nickell (1997) warns against this lumping but claims that it is convenient for analytical purposes.

[5] This is related to the concept of hysteresis introduced by Blanchard and Summers (1986): The existence of long-term unemployed will result in unemployment becoming more persistent. This deviation of unemployment from its equilibrium value will cause the equilibrium value itself to change over time.

[6] The problem of long-term unemployment continues to this day. The OECD (2002) reports that in 2000, over 50% of the unemployed in Italy, Greece, Belgium, Ireland, and Germany were long-term unemployed.

Figure 1: The Evolution of Unemployment in the OECD



Source: OECD and author's calculations

One argument is that as the unemployment spell lengthens, workers lose some of their human capital. An immediate consequence is that they become less employable. Theoretical studies by Pissarides (1992) and Ljungqvist and Sargent (1998) use this loss of skills assumption to explain why some individuals become long-term unemployed after a temporary negative shock to unemployment. Similarly, after some time unemployed, individuals become discouraged and diminish their job search intensity, lowering their probability of finding employment (see Devine and Kiefer, 1991; Schmitt and Wadsworth, 1993). Another strand of the literature focuses on the firm's behavior in relation to the long-term unemployed. Blanchard and Diamond (1994), Lockwood (1991), and Acemoglu (1995) conclude that firms prefer to hire newly unemployed individuals over those individuals with longer unemployment spells. In a process they call "ranking", Blanchard and Diamond (1994) assume that a firm receiving multiple job applications always picks the applicant with the shortest unemployment spell. This implies that the exit rate from unemployment becomes a negative function of duration[7] and the overall state of the labor market.

A crucial implication of the literature presented above is that those individuals who have been unemployed short-term will have the greatest impact on wage setting. On the wage formation effects of long-term unemployment, Blanchard and Diamond (1994) point out that "...one implication is that long-term unemployment, per se, has little effect on wages." The argument is that wages depend on the labor market prospects of the employed or newly unemployed, rather than on the prospects of the average unemployed. Efficiency wage models (Akerlof and Yellen, 1986) give support to this idea: The equilibrium or "efficiency wage" is set by firms as an incentive to the currently employed. The literature on insider-outsider models[8] arrives at similar conclusions: The long-term unemployed, as outsiders, have little influence on the wage bargaining process, while the insiders, the employed or newly unemployed, have the ability to impose their wage aspirations.

While most of the micro literature reviewed above takes a theoretical approach, there is only a small number of empirical studies that look for evidence of the effects discussed, largely for the UK. Studies by Nickell (1987) and Manning (1994) use UK data to claim that the long-term unemployed fail to exert downward pressure on earnings, or equivalently, that there is no significant association between this type of unemployment and wages (Manning, 1994). Franz (1987) arrives at similar conclusions using data for West Germany. Nevertheless, the results in these studies are not very conclusive (Blanchflower and Oswald, 1984) and should be interpreted with caution because of two important shortcomings: They concentrate on one country for a small time period, and they do not allow for a time-varying NAIRU. Both of these shortcomings are addressed in this paper.

3 Econometric Model: The Phillips Curve and the NAIRU

The short-run trade-off between inflation and unemployment has become one of the most important tools in the design and implementation of monetary policy (Gordon, 1997). Closely associated with this trade-off is the concept of the NAIRU, or that level of unemployment consistent with stable inflation.

[7] Lockwood (1991), and Acemoglu (1995) arrive to a similar conclusion. They claim that firms use unemployment duration as a signal of the individual's productivity level on which to base their hiring decisions.

[8] Lindbeck and Snower (1989) survey the literature on insider-outsider theories.

The NAIRU can be inferred from an expectations-augmented Phillips Curve of the following general form:[9]

$$\pi_t = \pi_t^e + \gamma \left(u_t - u_t^N\right) + \varepsilon_t \tag{1}$$

where π_t and π_t^e denote realized and expected inflation, and u_t^N is the natural rate or NAIRU at time t. The disturbance ε_t is assumed to be i.i.d. normal with zero mean and variance σ_ε^2. ε accounts for supply shocks that shift the inflation-unemployment trade-off, such as import prices or changes in the exchange rate.[10]

Equation (1) is similar to that used by Ball and Mankiw (2002), and Ball and Moffit (2001). It facilitates the analysis of annual data, as is the case in this paper. It also allows the contemporaneous unemployment gap to enter as a regressor. This assumes that there is no contemporaneous feedback from inflation to unemployment.[11]

There are two key issues concerning the estimation of equation (1). The first one is the specification of the inflation expectations. The second one is the modelling of the unobserved NAIRU. In relation to the former, it has become practice in much of the literature (see Staiger *et al*, 1997) to assume that expectations are adaptive, that is, $\pi_t^e = \pi_{t-1}$. Under adaptive expectations, (1) becomes:

$$\Delta\pi_t = \gamma \left(u_t - u_t^N\right) + \varepsilon_t$$

In regards to the modelling of the NAIRU, it is now widely accepted that it varies over time (see King and Watson, 1994; Steiger *et al*, 2001; Gordon 1997). On this subject, most of the recent literature assumes that the NAIRU follows a random walk.

3.1 Unemployment Duration Version of the Phillips Curve

This section introduces a modified version of the standard Phillips Curve model that accounts for different lengths in the duration of unemployment.[12] As previously discussed, the standard Phillips Curve uses the aggregate unemployment rate to measure economic activity and demand pressures on inflation. However, this may not be the most accurate indicator of inflationary

[9] Staiger et all (1997, 2001), Greenslade *et all* (2003), and Fabiani and Mestre (2001) are a few of the numerous studies on the Phillips Curve and the NAIRU.

[10] Section 6 on robustness will explicitly take into account the effect of supply shocks.

[11] Appendix B in Gruen *et al* (1999) explains the exogeneity assumptions relevant to the estimation of Phillips Curves.

[12] The idea of modifying the Phillips Curve by including other measures of unemployment is not new. Duca (1996) adds data on duration of unemployment, Roed (2002) uses job vacancy rates, and Ball and Moffitt (2001) considers productivity growth.

pressures, given that all the unemployed are entered with equal weights, regardless of the length of their spell. As an alternative, this paper proposes an index of unemployment that gives different weight to individuals based on the length of their unemployment spell. This index would indeed become a truer measure of inflationary pressures. The index takes the following form:

$$\tilde{U} = \alpha U_s + (1 - \alpha) U_l \tag{2}$$

where α is the weight assigned to the short-term unemployed, U_s is the short-term unemployment rate and U_l is the long-term unemployment rate. The value of α will be determined by the estimation. α can be interpreted as the impact that the short-term unemployed have on changes in inflation. Therefore, $\alpha = 0.5$ implies that both short-term and long-term unemployed have equal effect on inflation.

For the purpose of this paper, the duration version of the Phillips Curve will now be expressed as:

$$\Delta\pi_t = \gamma \left(\tilde{U}_t - \tilde{U}_t^N \right) + \varepsilon_t. \tag{1'}$$

where $\tilde{U}_t^N$ is the equilibrium value of the index of unemployment previously described.

This paper also modifies the standard Phillips Curve framework by modelling the NAIRU as a random walk with an stochastic drift. This is done to better capture the upward trend in unemployment observed in most European countries during the 1980s and 1990s (Laubach, 2001, and Fabiani and Mestre, 2001). Accordingly, the NAIRU follows the following process:

$$\tilde{U}_t^N = \tilde{U}_{t-1}^N + \mu_{t-1} + \nu_t \tag{3}$$

where

$$\mu_t = \mu_{t-1} + \eta_t \tag{4}$$

where ν_t and η_t are assumed i.i.d. normal $\left(0, \sigma_\nu^2\right)$ and $\left(0, \sigma_\eta^2\right)$ respectively, and uncorrelated with ε_t and with each other. Equations (1'), (3), and (4) can be expressed in state-space form and estimated by maximum likelihood using the Kalman filter (Harvey, 1989).

3.2 Estimation Issues

The estimation of (1'), (3), and (4) requires a number of assumptions in terms of the behavior of some of the variables and the treatment of some the parameters.[13]

Modelling the NAIRU as a random walk with a drift implies that the NAIRU is an I(2) process (given that the drift is I(1) itself). This paper will assume the unemployment gap to be I(0), which implies that the change in inflation must be I(0) as well. Table 12 in the appendix shows results from augmented Dickey-Fuller unit root tests for $\Delta\pi$. The table contains the *t*-tests results for the null hypothesis that the data contain a unit root. Given the corresponding critical values, the null hypothesis is soundly rejected for all the countries in the sample except for Denmark (rejected at the 5% level). Therefore, the results confirm that the change in inflation is I(0).

Another important issue is the estimation of the two parameters affecting the time variation of the NAIRU: $\sigma_\nu^2/\sigma_\varepsilon^2$ for high frequency variations and σ_η for low frequency variations. This is a problem akin to the selection of the smoothness parameter in the Hodrick-Prescott filter (Gordon, 1997).

The short-term volatility of the NAIRU is determined by the signal-to-noise ratio: $\sigma_\nu^2/\sigma_\varepsilon^2$. The larger the ratio, the more volatile the NAIRU is, whereas a ratio of zero implies a constant NAIRU. In principle, median-unbiased estimates of the signal-to-noise ratio can be obtained using the procedure in Stock and Watson (1998). However, as reported by Laubach (2001) and OECD (2000), the estimation of the signal-to-noise can be problematic.[14] In this paper I will follow Steiger *et al* (1997), Laubach (2001), and OECD (2000), and will fix the signal-to-noise ratio at the same value for every country. I tested alternative values based on the range of values obtained when the parameters are freely estimated. The value chosen is $\sigma_\nu^2/\sigma_\varepsilon^2 = 0.04$. This is similar to Laubach's 0.049.

Similarly, the value of σ_η, the long-term variation of the NAIRU, will be fixed at 0.02.

It is important to note that the main results in the paper are robust to the choice of these parameters. Section 6 further elaborates on this issue.

[13] Before the algorithm can be started, the vector of parameters needs to be initialized. Initial values for the coefficient on the unemployment gap are obtained from an OLS estimation of equation (1'). This procedure, suggested by Hamilton (1994), is similar to that employed by Fabiani and Mestre (2001). The initial value for the state variable will be the first observation of the HP-filtered unemployment rate, that is, $\tilde{U}_0^N = U_0^{hp}$. Results are not affected by the use of alternative starting values.

[14] This is related to so-called pile-up problem: The MLE of the signal-to-noise ratio of a nonstationay state variable such as the NAIRU, is downward biased towards zero.

4 Empirical Results

This section presents the estimation results. For every country in the sample, I am estimating a baseline Phillips curve model using two different specifications. The first one employs the standard unemployment rate, while the second one employs the unemployment index previously described. This facilitates the assessment of the performance of the modified model with respect to the standard model. The sample consists of 19 OECD countries: 14 European and 5 non-European. Long-term unemployment is defined as those individuals in the labor force unemployed for one year or longer. Short-term unemployed are those individuals out of work for less than one year.

4.1 Main Model Results

Results from estimating the Phillips Curve models for the countries in the sample are reported in Table 1 and Table 2. Table 1 displays results for the European OECD countries whereas Table 2 does it for the non-European countries. Each table contains results for both the standard and the modified models. For each of the specifications, the coefficient on the unemployment gap and standard errors are reported. Additionally, for the duration model, the value of the estimated weight on short-term unemployment, α, and its standard error are reported as well.

Focusing first on Table 1, columns three and four show that the γ coefficients on the unemployment gap have the expected negative sign, and are quite precisely estimated. All the coefficients are significant at the 10% level or better. This is consistent with results obtained by the OECD (2000) that find the contemporaneous unemployment gap to be quite indicative of changes in inflation in all the OECD countries in their sample. Column five contains the value of α, the weight on short-term unemployment. There is a good deal of cross-country variation in the estimates. For countries like Spain, Portugal, Ireland, and Greece, the value of α is around 0.9 or higher. This implies that the short-term unemployed alone have most of the ability to affect prices. In other countries such as Holland, Germany, and Sweden, this ability is more evenly distributed between both groups of unemployed (α values closer to 0.5). These results are consistent with the argument that the long-term unemployed have a diminished ability to influence prices. The precision with which these coefficients are estimated also varies. In some cases. they are estimated quite precisely, while in others (Finland, Portugal, and the UK), there is greater uncertainty around the estimate.

Table 1. Estimation Results (OECD Europe)

Country	Sample	Standard γ	Modified γ	Modified α	LR
Belgium	1973-02	-0.643	-1.015	**0.733**	7.962
		(0.124)	(0.182)	(0.060)	0.000
Denmark	1983-02	-0.268	-1.381	**0.741**	5.092
		(0.112)	(0.552)	(0.065)	0.035
Finland	1978-02	-1.168	-0.743	**0.804**	12.449
		(0.307)	(0.148)	(0.163)	0.000
France	1969-02	-0.232	-0.620	**0.768**	8.136
		(0.051)	(0.116)	(0.108)	0.000
Germany	1973-02	-0.350	-0.592	**0.630**	9.471
		(0.129)	(0.173)	(0.035)	0.000
Greece	1983-02	-0.739	-2.074	**0.947**	10.947
		(0.321)	(0.629)	(0.134)	0.000
Ireland	1979-02	-0.225	-1.299	**0.967**	11.759
		(0.087)	(0.401)	(0.043)	0.000
Italy	1979-02	-0.728	-1.922	**0.860**	14.390
		(0.347)	(0.801)	(0.191)	0.000
Netherlands	1973-02	-0.518	-0.937	**0.672**	6.838
		(0.096)	(0.148)	(0.028)	0.006
Norway	1979-02	-1.105	-1.633	**0.729**	4.993
		(0.467)	(0.671)	(0.100)	0.038
Portugal	1986-02	-0.765	-1.728	**0.881**	9.275
		(0.340)	(0.683)	(0.140)	0.000
Spain	1977-02	-0.243	-0.847	**0.942**	17.880
		(0.053)	(0.167)	(0.013)	0.000
Sweden	1971-02	-0.475	-0.653	**0.659**	3.160
		(0.079)	(0.104)	(0.084)	0.085
UK	1973-02	-1.045	-2.587	**0.839**	12.683
		(0.342)	(0.772)	(0.183)	0.000
			Average:	**0.798**	
				(0.084)	

Note: White robust standard errors in parenthesis.
p values reported for LR test.

Table 2. Estimation Results (OECD Non-Europe)

		Standard	Modified		LR
Country	Sample	γ	γ	α	
Australia	1978-02	-0.749	-0.827	**0.639**	3.372
		(0.312)	(0.337)	(0.221)	0.068
Canada	1976-02	-0.682	-1.268	**0.556**	3.609
		(0.175)	(0.318)	(0.085)	0.053
Japan	1977-02	-1.612	-0.772	**0.583**	2.838
		(0.715)	(0.324)	(0.127)	0.094
N. Zealand	1986-02	-0.899	-1.392	**0.698**	7.296
		(0.381)	(0.561)	(0.168)	0.000
US	1968-02	-1.348	-2.161	**0.538**	3.074
		(0.263)	(0.403)	(0.040)	0.089
			Average:	**0.603**	
				(0.127)	

Note: White robust standard errors in parenthesis.
p values reported for LR test.

The standard model is equivalent to the modified model when $\alpha = 0.5$ (they are nested). Given two nested models, the likelihood ratio test can be used to compare the two models correcting for the number of restrictions. The last column in Table 1 reports the likelihood ratio for the hypothesis that $\alpha = 0.5$. Given the number of restrictions, the test statistic follows a $\chi^2_{(1)}$. The test results show that the null hypothesis is always rejected at the 10% level or better. This confirms that the modified model outperforms the standard model in explaining changes in inflation.

Table 2 reports the same set of results for the non-Europe OECD countries in the sample. As in the previous table, the coefficients on the unemployment gap have the correct negative sign and are statistically significant. The weight α also indicates that for this group of countries the short-term unemployed have greater impact on prices than the long-term unemployed. Finally, the likelihood ratio test validates the use of the modified model.

Comparing results across the two groups of countries, the most interesting difference lies in the estimated value of α. This value tends to be larger in the European group of countries: The average α for the European countries is 0.798, whereas the average for the non-European

Table 3a. Variability of the NAIRU

	Standard	Modified
All	1.645	1.295
Europe	1.911	1.464
Non-Europe	0.897	0.823

is 0.603. As the next section will show, these differences across countries can be explained by some of the institutions that characterize labor markets in the OECD, such as union coverage levels and employment protection.

These results show that the incidence of long-term unemployment is key to understanding the true demand pressures on prices, and that this effect is not equal across countries.

4.2 Time Path of the NAIRU

One of the key features of the Phillips Curve is that it provides estimates of the NAIRU, a concept widely used by policy makers. Figure 2 in the appendix contains NAIRU estimates with 95% confidence intervals (CI) and the unemployment rate. For each country, the solid line represents the standard model NAIRU, with its shaded 95% CI. The modified model NAIRU and CI are shown in dashed lines. NAIRU estimates for the modified model have been mean adjusted to make them comparable to the standard model estimates. The time profiles are consistent with prior beliefs on the time behavior of the NAIRU.[15] In most European countries, the NAIRU's upward trend is followed by a gradual decline starting in the mid to late 1990s. Outside this group of countries, the NAIRU displays a less volatile behavior. These results are similar to those obtained by Laubach (2001), and OECD (2000).

The use of the modified model has an important implication for the time path of the NAIRU: It reduces its variability. Table 3a shows this decrease in variability (measured by the standard deviation of the NAIRU). For a number of European countries, this translates into NAIRUs that rose by less than what the actual variation in unemployment would have suggested. Correspondingly, for these countries, the modified NAIRU was lower than the standard NAIRU during the periods of high unemployment growth. This implies that output expansions to reduce unemployment would not have necessarily been as inflationary as expected. Ireland

[15] Gordon (1997) imposes some limitations on the low and high frequency variations of the NAIRU.

Table 3b. Confidence Intervals

	Standard	Modified	% Change
All	4.159	3.473	-0.198
Europe	4.254	3.426	-0.242
Non-Europe	3.895	3.603	-0.0810

presents a good example of this. Ireland's tame inflation of the late 1980s and early 90s is considered puzzling given the strong output growth and declining unemployment of the time. One suggested explanation is based on strong productivity growth leading to a decline in the NAIRU (Ball, 1999). The results in this paper suggest an alternative explanation: The usual estimation of the NAIRU is misspecified because it does not consider the effects of long-term unemployment. Properly accounting for these effects results in a lower profile for the NAIRU and a plausible explanation for the Irish puzzle. At its peak in 1989, the modified model implies a NAIRU over 15% lower than the standard model (12.3% NAIRU versus 14.5% for the standard model). A similar case is found in Sweden and Finland during the 1990s. In both these countries, unemployment shot up dramatically, with a large proportion of this growth coming from the long-term unemployed. Under the modified model, this translates into a flatter NAIRU than what the standard model would have implied (14% and 16% lower at their peaks in 2002.and 1994 respectively).

4.3 Confidence Intervals

The use of the Kalman filter has the advantage that it provides an estimate of the uncertainty around the NAIRU. This estimate is calculated from the error variance for the unobserved state. However, the uncertainty around the NAIRU is also affected by the fact that the true parameters in the model are unknown. I will use the Monte Carlo methods suggested by Hamilton (1994) to obtain confidence bands around the NAIRU that take into account both sources of uncertainty.[16]

As reflected in Figure 2, there is a good amount of uncertainty around the estimates of the NAIRU. This is a well documented problem of the NAIRU literature. The 95% CI

[16] These methods consist on obtaining simulated parameters based on the distribution of the set of parameters initially estimated. From each different draw of parameters, a new NAIRU series can be derived.

tends to be considerably large, and in two cases, Japan and Norway, it completely includes the unemployment rate. The US NAIRU is the most precisely estimated.

This uncertainty problem is solved to some extent by the modified model. Table 3b reports the unweighted mean across countries and across years of the width of the 95% confidence bands for both models, and the corresponding percentage change. The numbers in the table show a considerable reduction in the uncertainty around the NAIRU (19.8 percent reduction in the overall mean width of the NAIRU). The reduced uncertainty can also be observed in the graphs in Figure 2. The dashed CIs are considerably narrower, allowing for a better identification of the NAIRU with respect the unemployment rate.

The estimation of more precise NAIRUs is a major improvement of the modified model over the standard model of the NAIRU, and of great importance to the policy maker.

4.4 Euro Area Analysis

The previous analysis can be extended to investigate the unemployment-inflation trade-off in the euro area as a whole. For this purpose, I am constructing area-wide aggregate variables from individual country data.[17] Unemployment series are summed across countries. To obtain the area-wide consumer price index series I am using the "Index method" described in Fagan and Henry (1998) and Fabiani *et al* (2001). The aggregate index is constructed as the weighted sum of the individual country indices, with fixed weights based on each country's output.

The synthetic euro area data are used to estimate the standard and modified models of the Phillips Curve. The main results are presented in Tables 4 and 5, and Figure 3. Estimation values show that the coefficient on the unemployment gap is highly statistically significant regardless of the model used. The value for α is 0.734, which is lower than the straight average of 0.798 for the set of European OECD countries. Nevertheless, this value of α for the euro area seems to be consistent with the individual country results. In terms of the NAIRU, the modified model produces a more accurate estimate of the euro area NAIRU, with an 18% reduction in the mean width of the 95% CI. Euro area results are largely driven by two countries, Germany and France, that account for almost 50% of the labor force. As Figure

[17] Euro area aggregate series contain data for all 12 countries excluding Austria and Luxembourg, as no consistent series on unemployment duration is available for these two countries. Given the small size of their labor force, this exclusion is innocuous.

Table 4. Estimation Results (Euro Area)

		Standard	Modified		LR
Country	Sample	γ	γ	α	
Euro area	1973-02	-0.399	-0.827	**0.734**	9.327
		(0.093)	(0.177)	(0.128)	

Note: White robust standard errors in parenthesis.

Table 5. Changes in the NAIRU (Euro Area)

Confidence Interval Width			Nairu Variation	
Standard	Modified	%Change	Standard	Modified
3.442	2.925	-0.177	2.415	1.670

Note: Variation measured by the standard deviation of the NAIRU.

3 shows, the shape of the euro area NAIRU resembles the equivalent series for Germany and France.

Overall, the results show that the individual country results hold at the euro area level.

4.5 Implications for Forecasting

The Phillips Curve has become one of the most popular forecasting tools for inflation. In this section I will follow Stock and Watson (1999) to compare the forecasting performance of the standard and modified models of the Phillips Curve. Similar to their baseline model, I will be estimating equations of the form

$$\pi_{t+h}^{h} - \pi_t = \phi U_{gap,t} + \varepsilon_{t+h} \tag{5}$$

where $\pi_t^h = \ln(P_t/P_{t-h})$ is the h-year inflation rate, and π_t is inflation over the past year. U_{gap} takes two possible values: The first value is the gap between the unemployment rate and the standard NAIRU. The second is the gap between $\tilde{U}$ and $\tilde{U}^N$. Equation (5) will be estimated recursively using OLS to obtain out of sample forecasts of the change in inflation. That is,

Table 6. Forecast Results

	1-yr ahead forecast RMSE						
	Real Time		Ex-post		Constant		AR(1)
	Std.	Mod.	Std.	Mod.	Std.	Mod.	
All	0.0891	0.0740	0.0854	0.0682	0.125	0.121	0.258
Europe	0.0963	0.0754	0.0885	0.0716	0.130	0.126	0.281
Non-Eu.	0.0764	0.0698	0.0725	0.0678	0.119	0.117	0.229
	2-yr ahead forecast RMSE						
	Real Time		Ex-post		Constant		AR(1)
	Std.	Mod.	Std.	Mod.	Std.	Mod.	
All	0.283	0.268	0.294	0.271	0.351	0.349	0.384
Europe	0.286	0.270	0.299	0.273	0.356	0.353	0.419
Non-Eu.	0.267	0.262	0.271	0.266	0.344	0.341	0.371

I will estimate the model using only data available before the forecast period. For example, to forecast the change in inflation from period t to period $t+h$ I will estimate (5) using data up to and including period t. For the next forecast period, I will add one more observation to the data, and so on. This way, for each country and for each measure of the gap, I will obtain a forecast series for the change in inflation for the period 1995-2002. Given the annual nature of the data, I will calculate one-year and two-year ahead forecasts of inflation. Stock and Watson (1999) assume a constant NAIRU in their estimations. I will be assuming instead three different specifications for the NAIRU. The first one is a constant NAIRU. The second one is a real time NAIRU: A full model is estimated and a new NAIRU is calculated recursively every period using only data available at the time of the estimation. Finally, I will be using an *ex-post* NAIRU, uniquely calculated using the full sample. In addition, I am also comparing the results from these estimations with the forecast obtained from a univariate, AR(1), model of the change in inflation.

The results obtained are summarized in Table 6. The table displays the average across countries of the root mean squared error (RMSE) of the forecast for each of the possible estimations. A number of conclusions can be drawn from these results. First, the modified model outperforms the standard model at both, the one-year and two-year horizons, and for all three specifications of the Phillips Curve. This is reflected on the lower RMSE values (improvement over the standard model is 20-25%). Second, the improvement from using the modified model is more pronounced in the European than in the non-European countries. This is to expect, given that the modified model affects primarily, but not exclusively, the European countries. Third, the ability to forecast changes in inflation decreases dramatically as we move from the one-year to the two-year horizon; the RMSEs at the two-year horizon are considerably larger. The first of these results has very important implications for the use of the Phillips curve as a forecasting device: Forecasting using the Phillips Curve can be improved on by disaggregating unemployment in terms of duration.

In regards to the univariate forecasts, a number of recent studies (Atkenson and Ohanian, 2001) question the usefulness of Phillips Curves as forecasting tools, and claim that simple univariate models are as good, if not better, predictors of inflation. The results in this paper do not support these arguments, as shown on the last column in Table 4. The RMSEs of the AR(1) forecasts are larger than those obtained using time-varying NAIRU Phillips Curves. Fabiani and Mestre (2000) perform a similar forecasting exercise using Euro Area data and conclude that "the Phillips Curve outperforms a simple AR model of inflation, sometimes by a wide margin."

4.5.1 Evaluating the Forecasts

For a proper assessment and comparison of the various forecasts, it is important to have a statistical measure of their differences. One of these measures is the forecast combining regression used by Stock and Watson (2001) and others. The procedure entails the estimation of equations with the form

$$\Delta\pi_t = \omega\Delta\pi_t^{f,M} + (1-\omega)\,\Delta\pi_t^{f,ST} + \varepsilon_t \tag{6}$$

where $\Delta\pi_t^{f,M}$ is the forecast change in inflation obtained using the modified model of the Phillips Curve, and $\Delta\pi_t^{f,ST}$ is the equivalent forecast obtained with the standard model. The value of ω determines the relative performance of the two competing models. The higher the value of ω, the better the performance of the modified model over the standard model. Ideally, equation (6) ought to be estimated individually for each country. Unfortunately, the small number of forecast observations hinders the ability to produce such estimates at the country level. To get around this problem, I am specifying (6) as a system of equations, one for each country. Estimation of the system by GLS will use Seemingly Unrelated Regressions (SUR).

Table 7. Forecast Evaluation Results

	1 yr. ahead		
	Real Time	Ex-post	Constant
All	0.848	0.872	0.571
	(0.356)	(0.282)	(0.472)
Europe	0.985	1.024	0.627
	(0.269)	(0.140)	(0.417)
Non Europe	0.728	0.683	0.445
	(0.296)	(0.327)	(0.349)

Note: Estimation by GLS using SUR.
Standard errors in parenthesis.

The use of SUR lies on the assumption that the residuals are contemporaneously correlated across equations. This assumption seems plausible for the forecast errors in (6), since they are all generated from equivalent models. To increase the precision of the estimation, I am also assuming that the estimated coefficients are equal across countries.

The results obtained are reported in Table 7. It displays the value of ω for each of the possible forecast estimations and for the three country groups. Focusing on the forecasts at the one year horizon, the modified version of the Phillips Curve outperforms the standard version when the NAIRU is allowed to vary. This is implied by the values of ω estimated to be greater than 0 and close to 1. When the NAIRU is assumed to be constant, the modified model is not significantly different from the standard model, with ω values estimated quite imprecisely.

5 The Role of Labor Market Institutions

It is often argued that the different unemployment experiences in the OECD are the result of institutions that shape labor markets behavior. Nickell (1997) finds that labor market institutions can explain a great deal of the variation in unemployment and long-term unemployment across OECD countries. Similarly, Siebert (1997) claims that "...institutional differences between Europe and the United States can explain their different unemployment pictures."

This section seeks to relate the α values previously obtained to some of the institutions

Table 8. Weights and Labor Market Institutions

	alpha	*alpha*	*alpha*	*alpha*	*alpha*
Constant	0.596	0.635	0.682	0.724	0.519
	(13.817)	(9.629)	(10.127)	(11.131)	(11.031)
Emp. Protection	0.014				0.016
	(4.180)				(4.892)
Coverage		0.002			0.001
		(1.769)			(1.671)
Duration			0.027		0.024
			(1.942)		(1.567)
Labor Market Policies				-0.006	-0.008
				(-2.138)	(-1.943)
Adj. R^2	0.332	0.132	0.184	0.249	0.444

Note: t-stat. in parenthesis based on White heteroskedasticity-robust s.e.
Source: Labor market data from Nickell and Layard (1999)

known to affect labor markets. The finding of a link between these two would imply that the same variables used to explain variations in employment performance across countries can also be used to explain the relative importance of the long-term unemployed in the determination of price changes. I will be focusing on the five institutions most widely mentioned in the literature. These are: the employment protection index (refers to the legal regulation of the hiring and firing of workers), union coverage index (proportion of workers actually covered by union bargaining), active labor market policies (government expenditures to help unemployed get back to work), and finally, unemployment benefits duration (in years) and replacement rate.

Table 8 looks at the relationship between these labor market variables and α. I am regressing α on the set of labor market institutions. I tested different specifications, from more parsimonious to less parsimonious, to compare the individual and combined effects of these variables. Only the most interesting results are reported. All the variables, except the replacement rate of benefits, are robust predictors of α, with coefficients significant at the 10% or better. All the coefficients have the expected sign: positive for those institutions that contribute to a smaller impact of the long-term unemployed on inflation, and negative for those institutions (labor market policies) that make the long-term unemployed more likely to influence prices.

Given high levels of employment protection that limit firing and hiring by firms, employers will hire first those unemployed short-term (considered more productive and less risky), making the long-term unemployed (less productive and more costly) less likely to compete for jobs and therefore, affect prices. In regards to unions, to the extent that negotiations are based on the wage aspirations of the employed or the short-term unemployed, there is little room for the long-term unemployed to impact wages. Insider-outsider models of wage determination have similar implications. On the other hand, labor market policies such as assistance with job search and training, will make the long-term unemployed more attractive to employers and more likely to compete for jobs, increasing their influence on wages. Finally, a longer duration of unemployment benefits is likely to increase the time spent unemployed and to reduce the unemployed's impact on wages, as job search efficiency and human capital will deteriorate.

These results seem to indicate that in some countries (those where regulations are more worker friendly) these institutions have the effect of moving the unemployed (and particularly the long-term unemployed) to the fringe of the labor market, causing them to lose their ability to affect prices and wages.

6 Robustness to Alternative Specifications

This section tests the robustness of the results to three alternative specifications of the model. First, I will define the NAIRU in terms of wage inflation instead of price inflation. This is a sensible modification, given the relationship between unemployment and wages. Second, I will allow for measures of supply shocks to enter the estimation. In particular, I will consider the effect of the real price of oil. Finally, I will test the sensitivity of the results to the choice of the signal-to-noise ratio.

6.1 The Wage Phillips Curve

Although the standard practice is to estimate the NAIRU in terms of price inflation, there are some studies (Gordon, 1998, Gruen *et al*, 1999) that use some measure of wage inflation instead. I will follow these studies and use the growth rate of trend unit labor costs as the dependent variable. This variable is defined as the growth rate of nominal wages minus the rate of growth in trend labor productivity. The Phillips Curve now takes the form

$$\Delta w = \gamma \left(U_t - U_t^N \right) + \varepsilon_t \tag{7}$$

Table 9. Estimation Results (Wage Inflation)

	Standard	Modified		LR		Standard	Modified		LR
	γ	γ	α			γ	γ	α	
Aus.	-0.753	-0.773	**0.613**	3.114	Japan	-1.503	-0.629	**0.594**	2.820
	(0.302)	(0.319)	(0.165)	0.080		(0.785)	(0.297)	(0.138)	0.096
Belg.	-0.617	-1.000	**0.756**	7.409	Neth.	-0.572	-1.048	**0.732**	6.713
	(0.126)	(0.191)	(0.072)	0.000		(0.115)	(0.183)	(0.031)	0.007
Can.	-0.553	-1.293	**0.571**	3.570	N. Z.	-0.937	-1.480	**0.681**	6.732
	(0.172)	(0.389)	(0.099)	0.058		(0.413)	(0.611)	(0.199)	0.005
Denk.	-0.216	-1.514	**0.784**	5.240	Norw	-1.132	-1.478	**0.704**	4.640
	(0.087)	(0.579)	(0.105)	0.030		(0.461)	(0.611)	(0.091)	0.042
Fin.	-1.200	-0.732	**0.831**	10.74	Port.	-0.788	-1.931	**0.910**	9.388
	(0.341)	(0.159)	(0.138)	0.000		(0.362)	(0.786)	(0.117)	0.000
Fran.	-0.299	-0.671	**0.743**	8.014	Spain	-0.250	-0.918	**0.939**	17.27
	(0.070)	(0.132)	(0.169)	0.000		(0.057)	(0.187)	(0.025)	0.000
Germ.	-0.396	-0.547	**0.645**	10.26	Swed	-0.412	-0.778	**0.691**	3.075
	(0.127)	(0.136)	(0.030)	0.000		(0.068)	(0.125)	(0.127)	0.086
Gree.	-0.694	-2.151	**0.925**	9.743	U.K.	-1.267	-2.747	**0.883**	11.95
	(0.280)	(0.714)	(0.174)	0.000		(0.427)	(0.886)	(0.142)	0.000
Irel.	-0.208	-1.138	**0.962**	11.91	U.S.	-1.365	-2.220	**0.541**	2.969
	(0.089)	(0.386)	(0.047)	0.000		(0.269)	(0.427)	(0.042)	0.092
Italy	-0.784	-1.825	**0.892**	15.68					
	(0.364)	(0.691)	(0.147)	0.000					

Note: White robust standard errors in parenthesis.

where Δw is change in the growth rate of trend unit labor costs. Equation (7) is equivalent to equation 5 in Gordon (1998). It assumes that the same factors that affect price inflation also affect wage inflation. The equation will be estimated using both, the standard and modified models of the NAIRU. Tables 9 contains the results from the estimation. A comparison of these results with those obtained when price inflation is used as the dependent variable shows that the choice of inflation measure does not represent a significant change in the estimates. This follows from the fact that both variables, wage inflation and price inflation, are quite similar and track each other very closely. The average correlation between the two variables is 0.861, with some countries where the correlation is around 0.950. The different values of α remain mostly unchanged across estimations, with the average difference being less than 5

Table 10. Estimated α Weights (Shocks Augmented Model)

Australia	0.655	Belgium	0.718	Canada	0.539
	(0.197)		(0.053)		(0.069)
Denmark	0.753	Finland	0.816	France	0.732
	(0.094)		(0.119)		(0.093)
Germany	0.658	Greece	0.919	Ireland	0.952
	(0.056)		(0.102)		(0.058)
Italy	0.877	Japan	0.578	Netherlands	0.663
	(0.207)		(0.093)		(0.022)
N. Zealand	0.684	Norway	0.736	Portugal	0.890
	(0.145)		(0.104)		(0.156)
Spain	0.921	Sweden	0.675	UK	0.853
	(0.020)		(0.077)		(0.216)
US	0.526				
	(0.028)				

percent. The results in the tables are corroborated by a graphical analysis of the the resulting NAIRUs.[18] The time profiles of the NAIRUs are similar for both measures of inflation. In general, the results do not seem to be affected by the choice of inflation variable.

6.2 The Effect of Supply Shocks

As described in equation (1), Phillips Curve equations typically include a vector of variables that capture the impact of supply shocks. The most common of these variables are commodity prices such as oil, the exchange rate, and relative import prices. In this subsection I will be estimating the same set of equations augmented by the contemporaneous change in the real price of oil.[19]

Table 10 contains the results from these augmented Phillips Curves. It focuses on the different weights α resulting from the estimations.[20] Once again, the results are consistent with those obtained by the baseline model. The value of α does not deviate significantly from its original value. The coefficient on the unemployment gap (not reported) remains significant

[18] Given the similarities with the price inflation NAIRUs, the graphical results are not reported but are available upon request.

[19] Using other measures of supply shocks such as an import price index produced similar results.

[20] Complete tables and graphical results are not included but are avaible from the author.

Table 11. Estimated α Weights (Alternative $\sigma_\nu^2/\sigma_\varepsilon^2$ values)

	0.02	0.04	0.06		0.02	0.04	0.06
Australia	0.641	0.639	.0648	Japan	0.568	0.583	0.601
	(0.216)	(0.221)	(0.209)		(0.128)	(0.121)	(0.115)
Belgium	0.737	0.733	0.745	Netherlands	0.659	0.672	0.663
	(0.066)	(0.060)	(0.069)		(0.033)	(0.028)	(0.030)
Canada	0.545	0.556	0.568	N. Zealand	0.705	0.698	0.703
	(0.082)	(0.085)	(0.091)		(0.174)	(0.168)	(0.171)
Denmark	0.744	0.741	0.733	Norway	0.720	0.729	0.736
	(0.068)	(0.065)	(0.076)		(0.107)	(0100)	(0.105)
Finland	0.808	0.804	0.811	Portugal	0.885	0.881	0.886
	(0.167)	(0.163)	(0.168)		(0.138)	(0.140)	(0.141)
France	0.741	0.768	0.799	Spain	0.940	0.942	0.944
	(0.113)	(0.108)	(0.101)		(0.013)	(0.013)	(0.015)
Germany	0.625	0.630	0.621	Sweden	0.665	0.659	0.655
	(0.033)	(0.035)	(0.029)		(0.089)	(0.084)	(0.077)
Greece	0.940	0.947	0.943	UK	0.826	0.839	0.824
	(0.128)	(0.121)	(0.125)		(0.190)	(0.183)	(0.193)
Ireland	0.968	0.967	0.965	US	0.541	0.538	0.535
	(0.043)	(0.043)	(0.041)		(0.036)	(0.040)	(0.035)
Italy	0.855	0.860	0.867				
	(0.186)	(0.191)	(0.194)				

in all the countries in the study. Overall, these results tend to corroborate the main findings of the paper.

6.3 Changes to the Signal-to-Noise Ratio

The choice of the signal-to-noise ratio, $\sigma_\nu^2/\sigma_\varepsilon^2$, determines the high-frequency variation in the NAIRU. In the analysis thus far I have followed the literature in imposing a fixed, arbitrary value for this ratio. This value was chosen to produce time estimates of the NAIRU in line with prior expectations of its shape and in line with estimates in other studies. In this section I will test the sensitivity of α, the weight on short-term unemployment, to alternative values of the signal-to-noise-ratio. Table 11 contains the value of α for three different $\sigma_\nu^2/\sigma_\varepsilon^2$: the baseline value of 0.04, a high value of 0.06 and a low value of 0.02. These alternative values will affect the high-frequency but not the long-run variation of the NAIRU. As the table shows, the value of α is not very sensitive to changes in the signal-to-noise ratio. Variations of the estimated values fall within a relatively narrow range.

This result shows that the choice of the signal-to-noise ratio matters for the time-path of the NAIRU but not for the estimated value of α.

7 Conclusions

The emergence of long-term unemployment has shaped the unemployment experiences of many developed (OECD) countries over the last two decades. Labor market theories predict that the long-term unemployed are less relevant in the price formation process than the newly unemployed. This paper has investigated the implications of these predictions for the short-run trade-off between inflation and unemployment implied by the Phillips Curve. Using a new way to specify the Phillips Curve that allows different unemployment lengths to enter the model, this paper finds that unemployment duration matters for inflation dynamics, and that the long-term unemployed have a smaller effect on inflation. Moreover, the impact of the long-term unemployed is not found to be uniform across countries. In some countries, in particular some Western European countries, the long-term unemployed have a negligible effect on changes in prices. This variation across countries can be explained by some of the institutions that characterize labor markets in the OECD, such as employment protection and unionization levels. These are the same variables that are used to explain the incidence of long-term unemployment. Therefore, changes in the labor market geared to promote employment among the long-term unemployed should also have an impact on their ability to influence prices.

The modified model of the Phillips Curve proposed in this paper has important implications for the policy maker. By looking at the distribution of unemployment in terms of duration, a better measure of inflationary pressures can be developed. This paper finds that this improved measure produces more accurate forecasts of inflation at both, the one-year and two-year horizons. There are also implications for the estimation of the NAIRU. The modified model generates more precise estimates of the NAIRU, with an average reduction in the mean width of the confidence bands of close to 20 percent.

The results in this paper suggest a number of future research avenues. Similar to Schweitzer (2003), it would be interesting to study the relationship between α and the data on re-employment probabilities, and to combine the analysis with alternative measures of labor market slack. This paper has also shown that information on unemployment duration can help improve the policy maker's assessment of the dynamics of inflation. Additionally, a model can be developed linking the policy maker's actions to changes in unemployment and how they translate into short-term and long-term unemployment.

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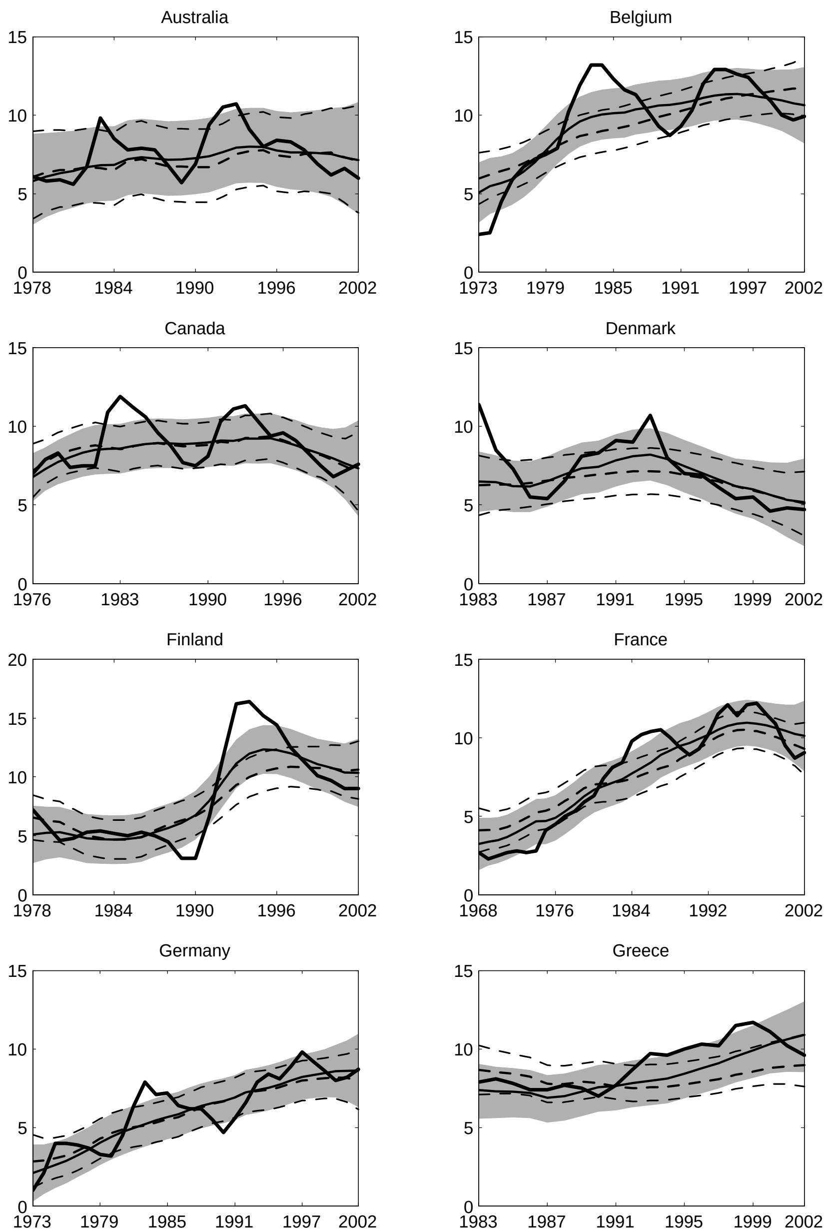
Table 12. Stationarity Analysis on $\Delta\pi$
(ADF Unit Root Test Results)

	t-stat.	Lags		t-stat.	Lags		t-stat.	Lags
Australia	-4.889	1	Belgium	-4.289	1	Canada	-4.5482	0
Denmark	-3.167	3	Finland	-4.321	1	France	-4.798	0
Germany	-5.125	0	Greece	-5.300	1	Ireland	-5.024	2
Italy	-5.496	0	Japan	-5.818	1	Netherl.	-6.736	0
N. Zealand	-6.168	0	Norway	-5.467	3	Portugal	-6.428	3
Spain	-7.205	0	Sweden	-6.989	0	UK	-6.128	0
US	-6.284	1						

Critical Values: 1%, -3.58; 5%, -2.93; 10%, -2.60.
t-stat. based on model with constant and no trend.
Optimal lag length based on BIC.

Figure 2: Unemployment and the NAIRU (CPI Inflation)



Bold- UR; *Solid*- Standard NAIRU; *Dash*- Modified NAIRU

Figure 2: Unemployment and the NAIRU (cont.)

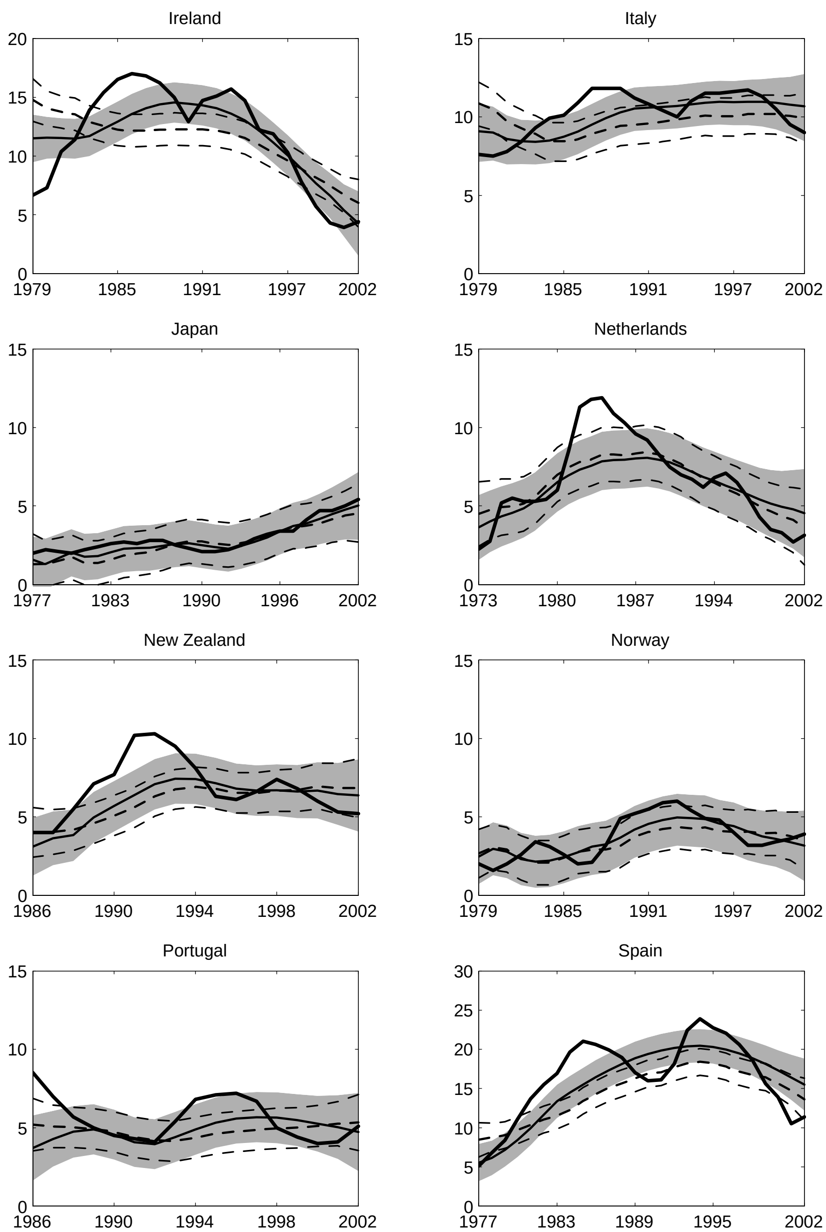


Figure 2: Unemployment and the NAIRU (cont.)

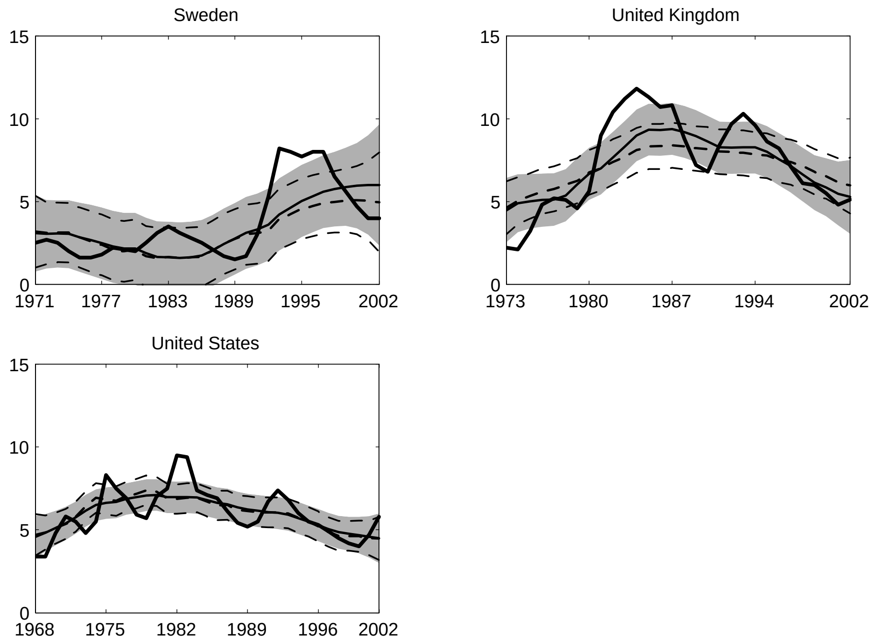
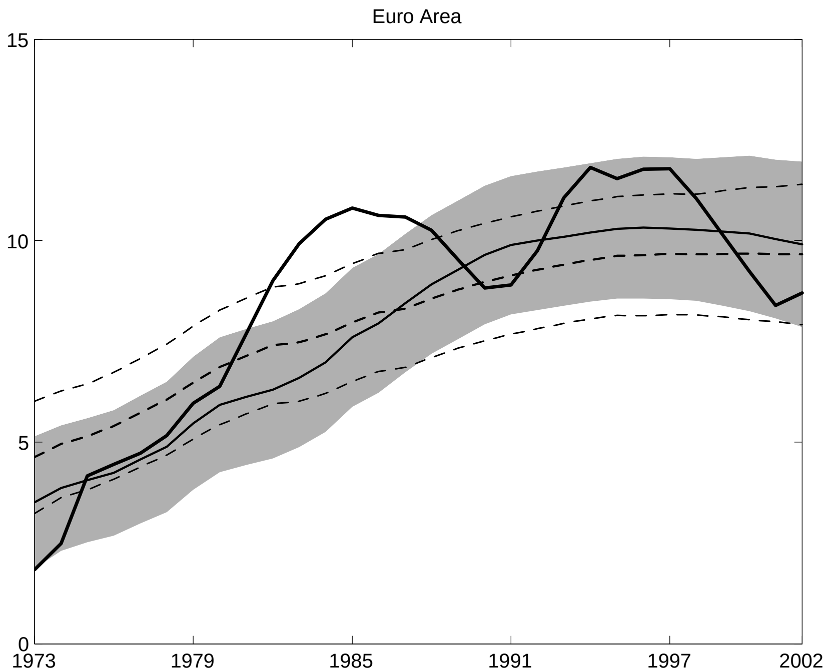


Figure 3. Unemployment and the NAIRU, Euro Area



Bold- UR; *Solid*- Standard NAIRU; *Dash*- Modified NAIRU