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Abstract

This paper studies the role of long-term unemployment in the determination of prices and wages. Labor market theories such as insider-outsider models predict that this type of unemployed are less relevant in the wage formation process than the newly unemployed. This paper looks for evidence of this behavior in a set of OECD countries. For this purpose, I propose a new specification of the Phillips Curve that contains different unemployment lengths 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 spell. The results show that unemployment duration matters in the determination of prices and wages, and that a smaller weight ought to be given to the long-term unemployed. This modified model has important implications for the policy maker: It produces more accurate forecasts of inflation and more precise estimates of the NAIRU.

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

JEL classification: C22, E31, E50, J64

Non-technical summary

The emergence of long-term unemployment has shaped the unemployment experiences of many developed (OECD) countries over the last two decades. Two key issues concerning this type of unemployment are of particular research interest. First, longer unemployment spells can be related to lower transition probabilities out of unemployment and into employment. Second, the long-term unemployed are less relevant to wage and price formation than the newly unemployed. This paper investigates the importance of the second of these issues for the short-run trade-off between inflation and unemployment implied by the Phillips Curve and the NAIRU (the Non-Accelerating Inflation Rate of Unemployment). This is a relevant question, given that the inverse short-run relationship between prices and unemployment is widely used by policymaking institutions to assess the desired stance of monetary policy. Yet in the presence of long-term unemployment, the aggregate rate of unemployment may provide a distorted measure of the true demand pressures exerted on prices and wages. This argument rests on the assumption that the long-term unemployed play a marginal role in the wage formation process. In this paper, I investigate whether evidence of this behavior is present in a set of 19 OECD countries. It is the first paper that undertakes such a systematic, multi-country study. The analysis uses 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. This deviates from the standard practice of using the aggregate unemployment rate. Optimal weights are determined by the estimation of the model by maximum likelihood using the Kalman filter.

The results obtained show that unemployment duration does matter in the determination of prices and wages as concluded by the Phillips Curve estimations, and that a smaller weight ought to be given to the long-term unemployed, confirming theoretical arguments presented in the paper. 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 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. Insofar as the monetary authority employs Phillips Curve models and the corresponding NAIRUs derived 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. 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 of the Phillips Curve generates more precise estimates of the NAIRU, with an average reduction in the mean width of the confidence bands of close to 20 percent.

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.[2] 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.[3]

The objective of this paper is to study the implications of long-term unemployment in the determination of prices and wages. This is an important issue because the inverse short-run relationship between prices and unemployment, as captured by the Phillips Curve and the NAIRU (the Non-Accelerating Inflation Rate of Unemployment), is widely used by policy-making 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 prices and wages. 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

[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] The OECD (1983, 1987) mentions 1982 as a year with particularly sharp increases in long-term unemployment in several countries.

[3] 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.

determination of prices and wages 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 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 prices. This deviates from the standard practice of using the aggregate unemployment rate.[4] 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 does matter in the determination of prices and wages, 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.

[4] 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

The unemployment experience in the OECD countries over the last two decades shows remarkable contrasts, with large disparities in its evolution across member 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,[5] suffered from persistently high and fairly volatile levels of unemployment. However, this has not always been the case. The upper panel of Figure 1 shows the unemployment rates for three different groups of countries: OECD Europe, OECD non-Europe, and OECD non-Europe excluding the US. For the greater part of the 1970s unemployment in Europe remained at low levels, comparable to those in other countries (and lower than in the US). Only at the end of the 1970s and early 80s, after the second oil shock and the subsequent disinflationary policies, did unemployment in Europe start to sharply rise in relation to the non-European countries. It quickly jumped from a rate of 2.9 percent in 1974 to a peak of nearly 10.5 percent in 1985. It remained at high levels for the rest of the decade. On the other hand, growth in unemployment outside Europe was much less pronounced, it reversed trend earlier, and by the end of the 1990s it was back to its pre-shock levels. The global slowdown of the early 1990s also had some important and interesting implications for unemployment: While it caused another big increase in unemployment in Europe, it was short-lived and relatively painless outside.

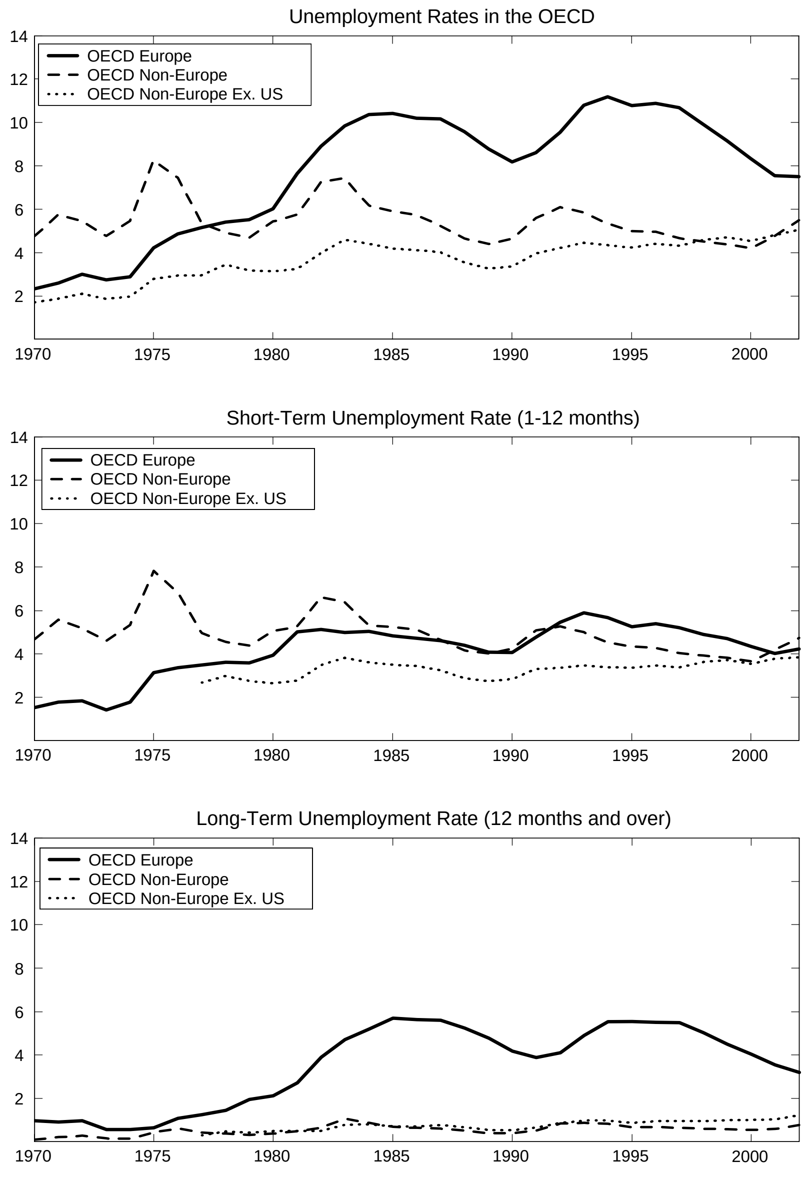
A large number of studies have attempted to explain these differences in the behavior of unemployment (see Nickell, 1997; Siebert, 1997; Blanchard and Wolfers, 2000; Ljungqvist and Sargent, 1998). These studies argue that the emergence of long-term unemployment provides an insight into the unemployment experiences in many OECD countries from the early 80s into 90s.[6] The middle panel of Figure 1 depicts short-term unemployment rates while the lower panel shows long-term unemployment rates. It is easy to see that most of the unemployment growth in Europe can be attributed to the striking growth in long-term unemployment. Its rate quickly jumped from about 1 percent in 1976 to almost 6 percent in 1985, remaining at high levels ever since.[7] On the other hand the behavior of short-term unemployment was

[5] 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.

[6] 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.

[7] 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

similar to that in other countries; short-term unemployment in OECD Europe averaged 4.9 percent during the1980s and 1990s, versus 4.8 percent in non-European OECD countries (3.3 percent if excluding the US).

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. 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[8] 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: If firms prefer to hire the newly unemployed because they are assumed to be more productive and less costly, the equilibrium or "efficiency wage" is determined by the wage demands of this preferred group. The literature

[8]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.

on insider-outsider models[9] 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.

Similarly, a number of studies use microdata to assess the impact of local unemployment on individual wages (see Pekkarinen (2001) for Finland, Blackaby and Hunt (1992) for the UK, and Winter-Ebmer (1996) for Austria). These studies find a positive relationship between long-term unemployment and wages.

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.

The NAIRU can be inferred from an expectations-augmented Phillips Curve of the following general form[10]:

$$\pi_t - \pi_t^e = \beta(L)\left(\pi_{t-1} - \pi_{t-1}^e\right) + \gamma(L)\left(u_{t-1} - u_{t-1}^N\right) + \delta(L) X_t + \varepsilon_t \tag{1}$$

where π_t and π_t^e denote realized and expected inflation, $\beta(L)$, $\gamma(L)$, and $\delta(L)$ are polynomials in the lag operator, u_t^N is the NAIRU at time t, and X_t is a vector of possible supply shocks

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

[10] 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.

(typically commodity prices or import prices). The disturbance ε_t is assumed to be i.i.d. normal with mean zero 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.[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 all* (1997)) to assume that expectations follow a random walk, that is, $\pi_t^e = \pi_{t-1}$, so $\pi_t - \pi_t^e = \Delta\pi_t$. In regards to the modelling of the NAIRU, it is now widely accepted that it varies over time[12] (see King and Watson (1994), Steiger *et all* (2001), Gordon (1997)). On this subject, most of the recent literature assumes that the NAIRU follows a random walk, and equation (1) is augmented with the following process for the NAIRU:

$$u_t^N = u_{t-1}^N + \nu_t \tag{2}$$

where ν_t is assumed to be i.i.d. normal with mean zero and variance σ_ν^2 and uncorrelated with ε_t at all leads and lags. The system formed by equations (1) and (2) can be expressed in its state-space form and can be estimated by maximum likelihood using the Kalman filter. A key advantage of the Kalman filter is that it can generate standard errors for the estimates of the NAIRU.

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[13]. 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 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 wage and price pressures. The index takes the

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

[12] In many initial studies, especially for the US, the NAIRU was assumed to be constant.

[13] 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.

following form:

$$\tilde{U} = \alpha U_s + (1-\alpha) U_l \tag{3}$$

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

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

This specification of the Phillips Curve is similar to that used by Gordon (1997, 1998), Ball and Mankiw (2002),the OECD (2000), and others in that it allows the contemporaneous unemployment gap to enter as a regressor. This assumes that there is no contemporaneous feedback from inflation to unemployment[14]. This specification also implies that inflation expectations follow a random walk, so the model can be estimated in first differences of inflation.

This paper also modifies the standard Phillips Curve framework by modeling the NAIRU as a random walk with an stochastic drift. This is done to better capture the movements in unemployment observed in most European countries (Laubach, 2001, and Fabiani and Mestre, 2001). Accordingly, equation (2) is now replaced by

$$\tilde{U}_t^N = \tilde{U}_{t-1}^N + \mu_{t-1} + \nu_t \tag{2'}$$

where

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

where η_t is assumed i.i.d. normal $\left(0, \sigma_\eta^2\right)$, and uncorrelated with ε_t and ν_t. Equations (1'), (2'), and (4) can be expressed in state-space form and estimated using the Kalman filter. Note that the modified version of the Phillips Curve is parsimonious. It omits supply shock variables or lag values of the unemployment index. This is mostly the result of data limitations. Nevertheless, section 6 checks for robustness of the results to alternative specifications of the model.

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

3.2 Estimation Issues

The system formed by equations (1'), (2'), and (4) can be estimated by maximum likelihood as described in Harvey (1989) and Hamilton (1994, ch. 13). However, before proceeding with the estimation of the parameters, a number of assumptions are required in terms of the behavior of some of the variables and the treatment of some the parameters.

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 contains 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).

Before the Kalman filter algorithm can be started, the vector of parameters needs to be initialized, including the state variable (the NAIRU). Initial values for the coefficient on the unemployment gap are obtained from an OLS estimation of equation (1')[15]. This procedure, suggested by Hamilton (1994), is similar to the one employed by Fabiani and Mestre (2001). The initial guess for the state variable will be the first observation of the HP-filtered unemployment rate, that is, $\tilde{U}_0^N = U_0^{hp}$. It is important to note that the results obtained are robust to the use of alternative starting values.

The final issue concerning the use of the Kalman filter deals with the smoothness of the NAIRU. This is a problem akin to the selection of the smoothness parameter in the Hodrick-Prescott filter (Gordon, 1997). The 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, both components of the signal-to-noise ratio can be estimated by the maximum likelihood procedure. However, as reported by Laubach (2001), OECD (2000), and others, the estimation of the signal-to-noise ratio leads to very flat NAIRUs[16]. In this paper, I will follow the approach of Steiger *et all* (1997), Laubach (2001), and others, and will fix the signal-to-noise ratio at values in line with the existing literature.

[15] The OLS estimation is done using the standard unemployment rate and its HP-filtered values. The use of the unemployment rate assumes that the initial value of α is .5.

[16] This is related to so-called pile-up problem: The ML estimate of the variance of a nonstationay state variable with small true variance, such as the NAIRU, is downward biased towards zero.

An alternative procedure to estimate median-unbiased estimates of the signal-to-noise ratio suggested by Stock and Watson (1998) was initially tested, but the results were not very satisfactory[17]. For the same arguments just explained, I will also fix the value of σ_η.

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. As discussed in the previous section, some assumptions are needed in terms of the underlying parameters of the model. In particular, the values 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) need to be determined. As in Laubach (2001), I will fix $\sigma_\nu^2/\sigma_\varepsilon^2$ and σ_η at the same value for every country. I tested alternative values for both parameters based on the range of values obtained when I let the parameters be freely determined by the estimation. The values chosen were $\sigma_\eta = 0.02$ and $\sigma_\nu^2/\sigma_\varepsilon^2 = 0.04$. These are relatively close to Laubach's 0.015 and 0.049 respectively, and result in time profiles of the NAIRU that fall in line with those in other studies (OECD, 2000).

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

[17] The estimation of the parameters in the signal-to-noise ratio led to very imprecise estimates, with a great deal of variation across countries.

Table 1. Estimation Results (OECD Europe)

Country	Sample	UR	LTU	Standard γ	Modified γ	Modified α	LR
Belgium	1973-02	11.06	6.89	-0.643	-1.015	**0.733**	7.962
				(0.124)	(0.182)	(0.060)	0.000
Denmark	1983-02	7.14	2.09	-0.268	-1.381	**0.741**	5.092
				(0.112)	(0.552)	(0.065)	0.035
Finland	1978-02	8.40	2.23	-1.168	-0.743	**0.804**	12.449
				(0.307)	(0.148)	(0.163)	0.000
France	1969-02	9.91	4.48	-0.232	-0.620	**0.768**	8.136
				(0.051)	(0.116)	(0.108)	0.000
Germany	1973-02	7.10	3.18	-0.350	-0.592	**0.630**	9.471
				(0.129)	(0.173)	(0.035)	0.000
Greece	1983-02	9.06	4.50	-0.739	-2.074	**0.947**	10.947
				(0.321)	(0.629)	(0.134)	0.000
Ireland	1979-02	11.89	6.79	-0.225	-1.299	**0.967**	11.759
				(0.087)	(0.401)	(0.043)	0.000
Italy	1979-02	10.40	6.55	-0.728	-1.922	**0.860**	14.390
				(0.347)	(0.801)	(0.191)	0.000
Netherlands	1973-02	7.31	3.55	-0.518	-0.937	**0.672**	6.838
				(0.096)	(0.148)	(0.028)	0.006
Norway	1979-02	3.76	0.54	-1.105	-1.633	**0.729**	4.993
				(0.467)	(0.671)	(0.100)	0.038
Portugal	1986-02	5.58	2.64	-0.765	-1.728	**0.881**	9.275
				(0.340)	(0.683)	(0.140)	0.000
Spain	1977-02	17.65	9.45	-0.243	-0.847	**0.942**	17.880
				(0.053)	(0.167)	(0.013)	0.000
Sweden	1971-02	4.33	0.93	-0.475	-0.653	**0.659**	3.160
				(0.079)	(0.104)	(0.084)	0.085
UK	1973-02	8.39	3.32	-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)

Country	Sample	UR	LTU	Standard γ	Modified γ	Modified α	LR
Australia	1978-02	7.71	2.20	-0.749	-0.827	**0.639**	3.372
				(0.312)	(0.337)	(0.221)	0.068
Canada	1976-02	9.12	1.28	-0.682	-1.268	**0.556**	3.609
				(0.175)	(0.318)	(0.085)	0.053
Japan	1977-02	3.07	0.67	-1.612	-0.772	**0.583**	2.838
				(0.715)	(0.324)	(0.127)	0.094
N. Zealand	1986-02	6.83	2.11	-0.899	-1.392	**0.698**	7.296
				(0.381)	(0.561)	(0.168)	0.000
US	1968-02	6.21	0.54	-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.

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.

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 countries is 0.603. This difference in α can be related to the presence of long-term unemployment in the respective countries: The average long-term unemployment rate (column 4) is 4.08% in the European countries[18] and 1.35% in the non-European. Portugal and the US provide an interesting example of this: As Blanchard and Portugal (2001) note they both have quite low unemployment rates (5.58% and 6.21% unemployment rate respectively). However, as reported in the last column on Tables 1 and 2, the long-term unemployed in Portugal have very little impact on prices ($\alpha = 0.881$) while those in the US have a considerable effect ($\alpha = 0.538$). This translates into much higher long-term unemployment in Portugal (2.64%) than in the US (0.54%). The result follows from the fact that a higher α represents less downward pressure on wages, and therefore, more long-term unemployment.

The values of α obtained can also be related to the dynamics of unemployment. As in Bean (1994), and OECD (1995) one can look at data on flows out of unemployment (estimated as the difference between the average monthly level of inflows and the monthly average change in unemployment over one year) across countries as a proxy for the probability of finding a job . These data can be compared to the values of α to see if there is a relationship between α and the probability of re-employment. Columns 2 and 3 in Table 12 show that there is an inverse relationship between the value of α and the data on flows out of unemployment. The correlation between the two variables is -0.67. Therefore, higher α are associated with

[18] The low rates of long-term unemployment in countries such as Sweden and Finland may reflect the fact that many individuals who would otherwise be counted as long-term unemployed are in subsidized employment or training. The effect of this group on inflation is hard to quantify, but it could influence the results for these countries.

a smaller probability of getting out of unemployment. This result can be motivated by the rigidities and institutions that affect unemployment dynamics in the different countries. Similar inverse relationship can be found between the rate of job offers and α, with a negative correlation of -0.53 (See Column 4). Finally, for a smaller sample of European countries, one can analyze the re-employment probabilities of the long-term unemployed. Column 5 shows the proportion of re-employed individuals who were long-term unemployed using a small sample of the unemployed. For this reduced group of countries, a negative relationship between α and the re-employment probabilities of the long-term unemployed is found. This result gives support to the argument that the long-term unemployed have a smaller ability to compete for jobs and to therefore affect prices and wages. Given the exploratory nature of this exercise, a more in-depth analysis of the relationship between α and the re-employment probabilities of the long-term unemployed is left for future research.

Overall, the results in this section show that the incidence of long-term unemployment is key to understanding the true pressures on prices, and that high long-term unemployment is associated with the long-term unemployed having little effect on prices. As Section 5 will show, this latter result can be related to the nature of the institutions that characterize labor markets in the OECD countries under study.

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.[19] 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

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

Table 3a. Variability of the NAIRU

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

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 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

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

sources of uncertainty.[20]

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 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.[21] 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

[20] 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.

[21] 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.

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 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, 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: Every period, a full model is estimated and a new NAIRU is calculated recursively 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[22]. 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

[22] Individual country results are available from the author upon request.

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
Euro area	0.0971	0.0758	0.0891	0.0719	.0133	0.129	0.287
Non-Europe	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
Euro area	0.291	0.273	0.304	0.279	0.360	0.356	0.426
Non-Europe	0.267	0.262	0.271	0.266	0.344	0.341	0.371

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 6. 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."

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)
Euro area	1.026	0.925	0.562
	(0.212)	(0.108)	(0.387)
Non Europe	0.728	0.683	0.445
	(0.296)	(0.327)	(0.349)

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

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 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 pooling the forecast data and am specifying (6) as a system of equations, one for each country. Estimation of the system by GLS will use Seemingly Unrelated Regressions (SUR). 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 four 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 a 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 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 8a 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

Table 8a. 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)

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.

An alternative explanation to the results in Table 8a is that the estimation is only capturing the indirect impact of institutions on weights through their impact on long-term unemployment. That is, the institutions affect long-term unemployment but have no direct impact on the weights. To disentangle this direct effect from the effect via long-term unemployment one can

Table 8b. Weights, Unemployment, and Labor Market Institutions

	alpha	*alpha*	*alpha*	*alpha*	*alpha*
Constant	0.588	0.555	0.634	0.617	0.528
	(15.329)	(11.943)	(11.981)	(11.617)	(12.492)
Emp. Protection	0.007				0.012
	(2.326)				(1.904)
Coverage		0.002			0.002
		(1.730)			(1.679)
Duration			-0.006		0.013
			(−0.419)		(0.769)
Labor Market Policies				-0.001	-0.005
				(−0.158)	(−1.153)
LTU	0.025	0.033	0.036	0.035	0.012
	(2.135)	(4.382)	(5.525)	(4.588)	(1.863)
Adj. R^2	0.433	0.454	0.381	0.378	0.414

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

add measures of unemployment as regressors to control for the indirect effect. Table 8b shows the results from these estimations. Once again, only the most significant results are reported. The new set of estimations seems to partially confirm this alternative explanation. When the long-term unemployment rate is added as a regressor, only the employment protection index and the level of union coverage are found to be significant. The other two variables, unemployment benefits duration and labor market policies, are no longer significant. This result implies that the employment protection index and the union coverage are the only variables that have a direct impact on the weights beyond their indirect effect via long-term unemployment. On the other hand, unemployment benefits duration and labor market policies only affect the weights via their impact on long-term unemployment.

The results in this section seem to indicate that in some countries (those where regulations are more worker friendly) these institutions, especially the employment protection index and the level of union coverage, cause the unemployed (and particularly the long-term unemployed) to lose their ability to affect prices and wages. Nevertheless, these results should be interpreted with caution and as a exploratory exercise given the small number of observations (19) in the sample.

Table 9. Estimation Results (Wage Inflation)

	Standard	Modified			Standard	Modified	
	γ	γ	α		γ	γ	α
Australia	-0.753	-0.773	0.613	Italy	-0.784	-1.825	0.892
	0.075	0.068	(0.165)		0.088	0.067	(0.147)
Belgium	-0.617	-1.000	0.756	Japan	-1.503	-0.629	0.594
	0.000	0.000	(0.072)		0.107	0.093	(0.138)
Canada	-0.553	-1.293	0.571	Netherlands	-0.572	-1.048	0.732
	0.021	0.017	(0.099)		0.000	0.000	(0.031)
Denmark	-0.216	-1.514	0.784	N. Zealand	-0.937	-1.480	0.681
	0.098	0.084	(0.105)		0.092	0.077	(0.199)
Finland	-1.200	-0.732	0.831	Norway	-1.132	-1.478	0.704
	0.051	0.032	(0.138)		0.081	0.077	(0.091)
France	-0.299	-0.671	0.743	Portugal	-0.788	-1.931	0.910
	0.064	0.048	(0.169)		0.079	0.068	(0.117)
Germany	-0.396	-0.547	0.645	Spain	-0.250	-0.918	0.939
	0.056	0.046	(0.030)		0.013	0.000	(0.025)
Greece	-0.694	-2.151	0.925	Sweden	-0.412	-0.778	0.691
	0.089	0.070	(0.174)		0.041	0.026	(0.127)
Ireland	-0.208	-1.138	0.962	UK	-1.267	-2.747	0.883
	0.076	0.050	(0.047)		0.059	0.052	(0.142)
				US	-1.365	-2.220	0.541
					0.000	0.000	(0.042)

Note: p values reported for γ.
Standard errors for α in parenthesis

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 all, 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

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

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. Table 9 contains the results from the estimation. A comparison of these results with those obtained when consumer 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 percent. The results in the tables are corroborated by a graphical analysis of the the resulting NAIRUs[23]. 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 real price of oil in dollar terms[24].

Table 10 contains the results from these augmented Phillips Curves. It focuses on the

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

[24] Using other measures of supply shocks such as an import price index produced similar results and are not reported.

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)				

different weights α resulting from the estimations[25]. 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 inclusion of a supply shock in the form of the change in the real price of oil should have an effect on the overall price level but should not have a big impact on the distribution of unemployment. As previously shown, labor market variables tend to determine it. The coefficient on the unemployment gap (not reported) remains significat in all the countries in the study. There is also no clear pattern on how the inclusion of this variable affects the value of α. 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 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

[25] 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)				

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. On the other hand, time estimates of the NAIRU do seem to be more sensitive to changes in the signal-to-noise ratio.

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						

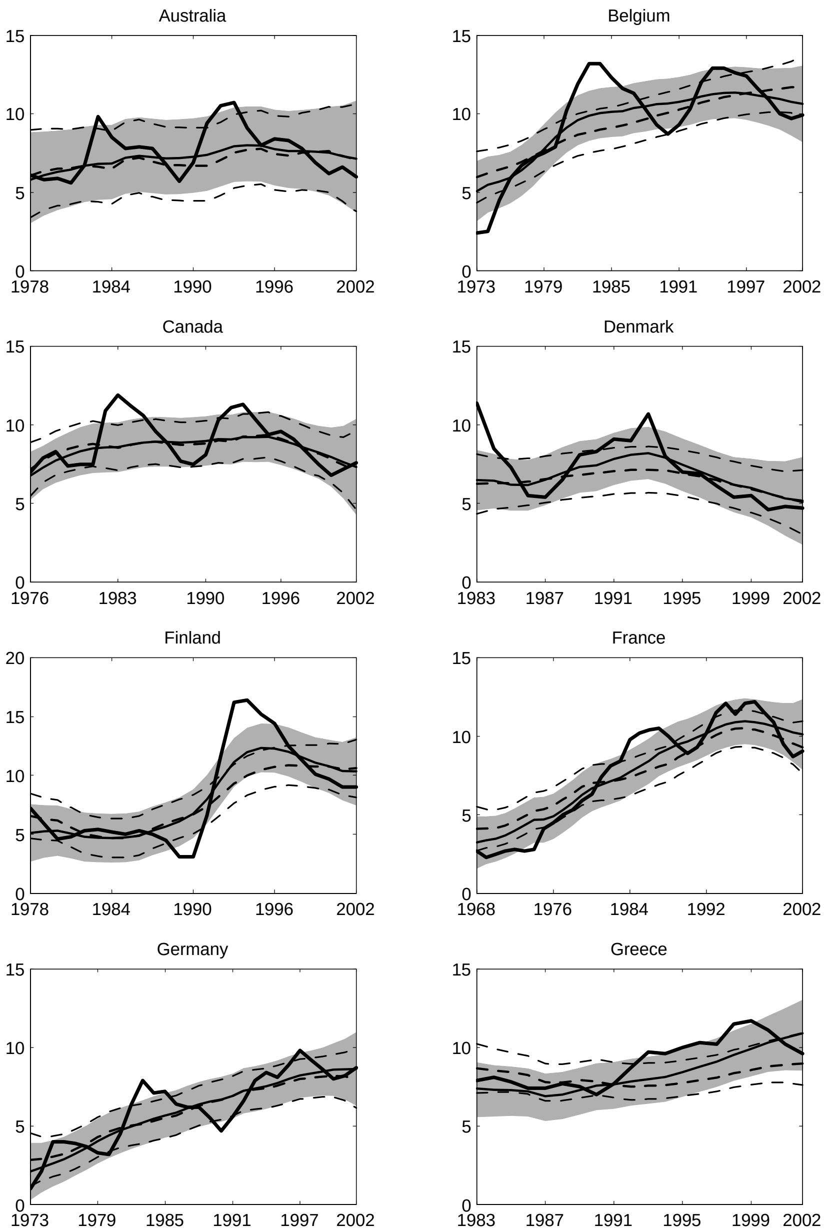
Critical Values: 1%, -3.58; 5%, -2.93; 10%, -2.60.
t-stat. based on model with constant and no trend.
Optimal lag length based on BIC.

Table 13. Weights and Re-employment probabilities

Country	*alpha*	Outflow	Offer Rate	% LTU
Australia	0.639	14.4	n.a.	n.a.
Belgium	0.733	8.6	0.083	30.51
Canada	0.556	27.5	n.a.	n.a.
Denmark	0.741	21.4	0.103	28.92
Finland	0.804	13.9	0.087	n.a.
France	0.768	3	0.067	28.23
Germany	0.63	9	0.098	31.43
Greece	0.947	4.7	0.045	22.82
Ireland	0.967	3.8	0.058	33.70
Italy	0.86	9.5	0.064	31.42
Japan	0.583	14.4	n.a.	n.a.
Netherlands	0.672	6.4	0.06	n.a.
Norway	0.729	21.6	n.a.	n.a.
NZ	0.698	17.7	n.a.	n.a.
Portugal	0.881	15.3	0.03	32.41
Spain	0.942	2.7	0.082	24.48
Sweden	0.659	18.4	n.a.	n.a.
UK	0.839	9.3	0.076	27.27
US	0.538	37.6	n.a.	n.a.

Sources: OECD (1995), Table 1.9 for data on flows. Addison et al(2004) for data on offer rates, and OECD (2002), Table 4.6 for data on LTU re-employment probabilities.

Figure 2: Unemployment and the NAIRU (CPI Inflation)



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

Figure 2: Unemployment and the NAIRU (cont.)

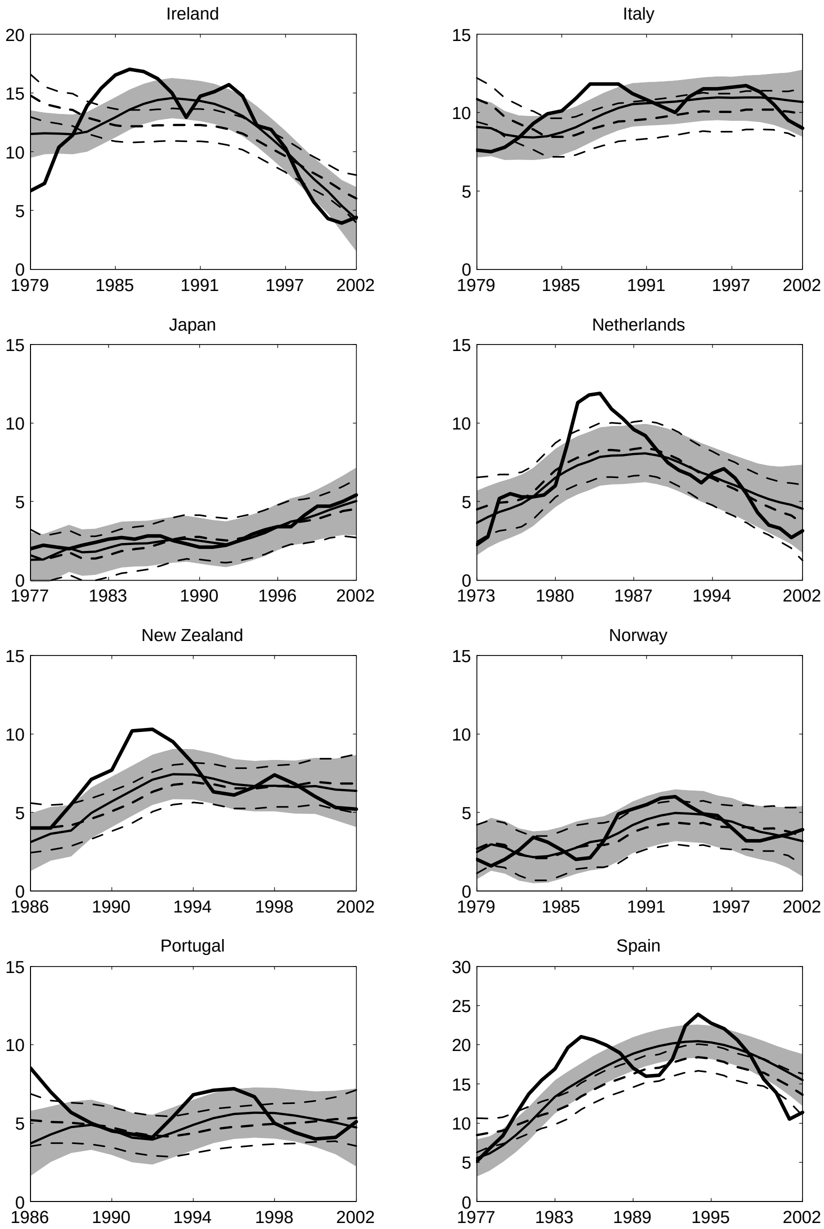


Figure 2: Unemployment and the NAIRU (cont.)

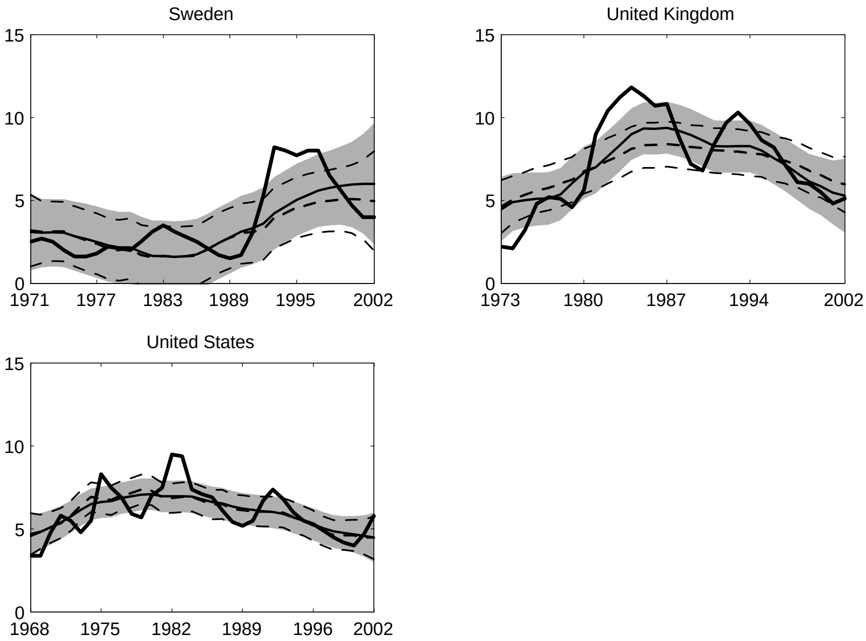
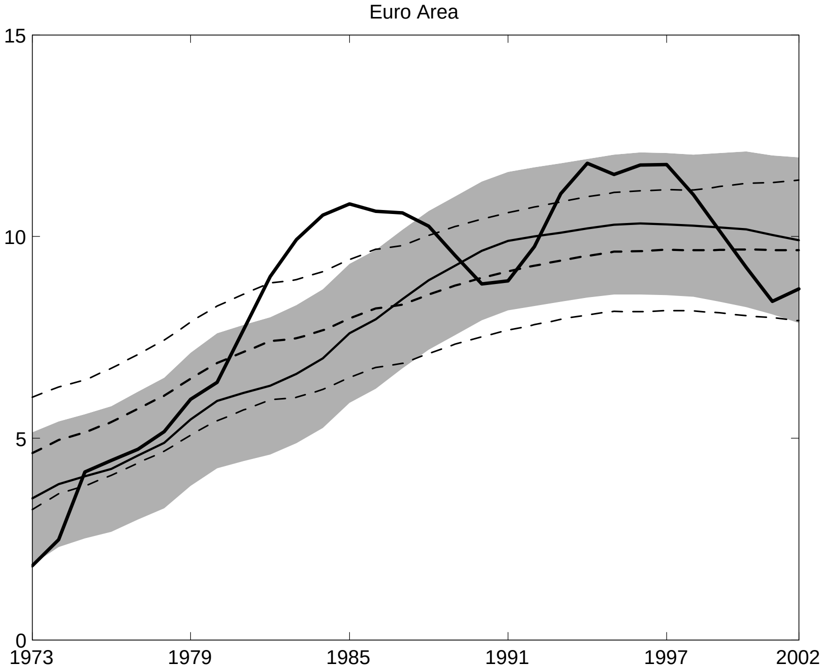


Figure 3. Unemployment and the NAIRU, Euro Area



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

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