The wage dynamics in Spain: evidence from individual data (1994-2001)

INMACULADA GARCÍA-MAINAR*

and

VÍCTOR M. MONTUENGA-GÓMEZ**

***** Departamento de Análisis Económico, Facultad de Ciencias Económicas y Empresariales, Gran Vía, 2, 50005 Zaragoza, SPAIN. E-mail: igarcia@posta.unizar.es

** Departamento de Economía y Empresa, Universidad de la Rioja, Edificio Quintiliano, C/ La Cigüeña, 60, 26004 Logroño (La Rioja), SPAIN. E-mail: victor-manuel.montuenga@dee.unirioja.es

Abstract

In this paper, we test the hypothesis of a wage curve against a Phillips curve for Spain within a framework which allows for these both and more general alternatives. To this end, we use data from the European Community Household Panel, which provides microinformation for the period 1994-2001. The results indicate that a partial wage adjustment is at work, as in other European countries, and that the elasticity of wages to unemployment is close to the "empirical law of economics" of –0.1.

Keywords: Phillips curve; Wage Curve; Panel data; ECM

JEL classification: J31, J60, J64

Corresponding author: Inmaculada García-Mainar. Departamento de Análisis Económico, Facultad de Ciencias Económicas y Empresariales, Gran Vía, 2. 50005 Zaragoza (SPAIN). Tel +(34) 976 762635. Fax: +(34) 976 761996. E-mail: igarcia@posta.unizar.es

Introduction

The dynamics of the wage adjustment is a controversial issue nowadays. The negative relationship found by Phillips (1958) between the growth rate of wages and the unemployment rate became the cornerstone of the Keynesian synthesis and the macroeconometric modelling, since it captured adequately the price and wage adjustment mechanism. In this way, economic authorities, by following a "fine tuning" policy, were able to choose an appropriate combination of inflation and unemployment rates. The original formulation, and updated versions of the Phillips curve (see Friedman's, 1968, and Phelps', 1968, accelerationist and Lucas', 1973, price surprise hypotheses), have been widely used to model the supply side of the economy, such that, when confronted with the demand side, their intersection determines the product and price equilibrium values. Under this view, wages and prices tend to adjust when demand excess in such a way that, sooner or later, economy move towards the equilibrium locus. Therefore, supply and productivity shocks are assumed to have not long run effects on real variables.

However, the validity of the Phillips curve, in aggregate terms, has been recently disputed in the US given the particularly good results in inflation and unemployment during the last years of the past century (see Coen et al., 1999). Similarly, using microeconomic data, Blanchflower and Oswald (1990 and 1994) proved that in the US "the Phillips curve is probably a mis-specified aggregate wage curve", and consequently, "the idea of a Phillips curve may be inherently wrong". Nevertheless, recent studies has shown that, when using further amendments of the Phillips curve (by considering trends in the productivity growth and/or in the natural rate of unemployment), it fits well the US reality (Gordon, 1998, Ball and Moffitt, 2002). Simultaneously, the Phillips curve has also enjoyed a resurgence in the theory of monetary policy, see Clarida et al. (1999) for a summary.

By contrast, the hypothesis of a Phillips curve representing the relation between wages and unemployment in European and in other OECD countries has already been challenged since the mid-eighties (Grubb, 1986 and Layard et al., 1991), on favour of a dynamic wage relationship. This empirical evidence is not theoreticless founded. The modern non-competitive theories on the labour market predict a negative relationship at the micro stage between the levels of the wages and the unemployment rate (see Shapiro and Stiglitz, 1984, Mortensen and Pissarides, 1994), and the microeconomic evidence provided by Blanchflower and Oswald (1994), and subsequent studies, corroborate this finding. This relationship, called "the wage curve", represents an equilibrium locus of the pairs wage and unemployment resulting from the optimising behaviour of the agents involved in the bargaining process (see

Blanchflower and Oswald, 1995 and Blanchard and Katz, 1997). The wage curve represents an upward sloping quasi-labour supply curve (or surrogate labour supply, or wage setting curve, depending on the author), that lies to the left of, and is flatter than, the classical labour supply curve, in such a way that, when is confronted with the demand labour curve, determines an equilibrium wage, above the one that clears the labour market, and an "equilibrium unemployment rate" (see Lindbeck, 1993 and Woodford, 1994). The increasing relevance of the wage curve in the modelling of labour market has led to re-interpret this as a wages/unemployment space where a downward-sloping wage curve intersects with a horizontal or upward-sloping price curve, to derive a new aggregate supply curve (see Blanchard and Katz, 1997 and Suedekum, 2005).[1] Under this view, aggregate supply and productivity shocks will have permanent effects on unemployment and output.

On consequence, the discussion about whether unemployment is related to the growth or to the levels of wages is not meaningless, but it has powerful consequences in our understanding of the labour market and of the economy as a whole. Firstly, in determining the dynamic effects, if some, of demand and supply wage variables on the natural rate of unemployment. Some authors, e.g Blanchard and Katz (1999) consider that it is essential in the determination of the NAIRU, whereas others, e.g. Whelan (2000) and Bell et al. (2002) sustain that it only matters for its evolution over time and not for the value of the NAIRU itself (for a general overview on this matter, see the recent discussion in Bardsen and Nymoen, 2003). Secondly, it may help to ascertain the exact nature of the reservation wage, and the dependence (if exists) of current wages on lagged wages (Blanchard and Katz, 1997 and Ball and Moffitt, 2002). Thirdly, it provides and empirical guide for policy modelers to appraise the effects of shocks on the price inflation and on the inflation-unemployment tradeoff. If unemployment is related to wage changes (Phillips curve), supply shocks will only temporarily affect price inflation, whilst if unemployment is linked to wage levels, then such shocks will continue to impact wage bargaining and price inflation in later periods (Fares, 2002 and Madsen, 2002).

In the recent years, and given the European experience, some evidence seems to support the idea that both hypotheses, the static wage curve and the Phillips curve, are extreme cases, and that an inter-medium view is probably more appropriate. Labelling this alternative view as a dynamic wage curve, the aim of this paper is to cast some light on this debate for the

[1] Alternative specifications, derived from Harris-Todaro model (1970), use an upward-sloping zero migration equation to represent labour markets (see Blanchflower and Oswald, 1994, Bell et al., 2002, and Morrison et al., 2005).

Spanish case by using individual data. Besides, some interesting aspects of the wage dynamics in Spain will be investigated. Section 2 surveys the literature and describe the main concerns of our research. Section 3 presents the empirical specifications that is addressed in the applied analysis and describes the data base. Section 4 presents the empirical results and section 5, finally, concludes and anticipates some future extensions.

2. The wage curve and the Phillips curve

The idea of a wage curve in microeconomic terms can be opposed to the existence of a Phillips curve in aggregate terms. First, he wage curve is a negative relationship between the wage *level* and the unemployment rate, whereas the Phillips curve captures the negative relationship between the *growth* of wages (wage inflation) and the unemployment rate. Second, the wage curve is habitually obtained from disaggregated data of longitudinal household or individual surveys, whereas the Phillips curve is usually estimated with macroeconomic unemployment and wage inflation data. A further difference lies in the economic meaning of each concept. The wage curve represents a *locus* of equilibrium points, the wage/unemployment rate pairs that arise from the optimising behaviour of economic agents in non-competitive models of the labour market. By contrast, the Phillips curve is a set of disequilibrium points that represent the adjustment process in a competitive model of the labour market.

Habitually, a wage curve using individual data is estimated by adding into a Mincer-type wage equation, the log of the unemployment rate

$$ln(w_{irt}) = a + f_r + d_t + b\,X_{irt} + \beta\,ln(u_{rt}) + \varepsilon_{irt} \qquad (1)$$

where i represents individuals, r regions and t time periods and where w is the real wage, X a set of individual and labour characteristics (such as gender, education, occupation...), u the unemployment rate, f_r a set of regional fixed effects, d_t a set of time fixed effects and ε is the remainder error term. Time-period effects control for all those variables that vary over time but that are common to all the regions (i.e., business cycle variables), whereas variables that are time-invariant but particular to each region, such as endowments, amenities, facilities, etc., are contemplated by including regional fixed effects.

This double logarithmic expressions has been justified as to provide the best results (see Blanchflower and Oswald, 1994) and has been profusely applied since then. The coefficient β

is, therefore, the unemployment elasticity, with a negative estimated value, thereby demonstrating the existence of a wage curve. The inclusion of regional fixed effects allows to capture any permanent component of the relationship between wages and unemployment, so that the unemployment coefficient β is only reflecting the temporary component of that relationship. Expressions like (1) have been estimated for multitude of countries showing, as a general result, that wage elasticity to unemployment lies on the range (-0.20, -0.05) for most of the cases.[2]

To study the behaviour of wage dynamics, Blanchflower and Oswald (1994) add, as additional regressor, the lagged dependent variable, and test whether its associated coefficient is close to zero or to one. Therefore, an equation like (2) is estimated.

$$ln(w_{irt}) = a + \rho\, ln(w_{irt-1}) + f_r + d_t + b\, X_{irt} + \beta\, ln(u_{rt}) + \varepsilon_{irt}. \qquad (2)$$

With the estimate of the parameter ρ, the hypothesis of a Phillips curve can be tested straightforward. If its value is not significantly different from one, the null hypothesis could not be rejected, whereas if its value is close to zero, we would accept the alternative hypothesis of a wage curve.

In their tables 4.27 and 6.20 Blanchflower and Oswald (1994) show that, with data from the March CPS (Current Population Survey) for the US and from the GESS for the UK, the estimate of ρ is close to zero. This result suggests that wages adjust rapidly to the unemployment rate, which constitutes the starting point to claim the death of the Phillips curve. They argue that "the apparent autoregression in macro pay levels may be the result of aggregate error or measurement error or specification error or all three" (p. 284), and then the use of micro data is considered as most appropriate in unveiling the truth. This conclusion supposed a big challenge against the predominant evidence showed by the aggregate studies for the case of the US, always favourable to the Phillips curve (see King and Watson, 1994 and Roberts, 1995, 1997a for instance),[3] and spurred the empirical analysis in order to study deeply the phenomenon of wage persistence.

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[2] Among many others, Wagner (1994), Blanchflower and Oswald (1994), Bratsberg and Turunen (1996), Turunen (1998), Janssens and Konings (1998), Baltagi and Blien (1998), Kennedy and Borland (2000) and Montuenga et al., (2003). Exceptions to these general rule are the Nordic countries (Albaek et al., 2000) where parameter β is found to be non-significant, and some Eastern European and developing countries (see Baltagi et al., 2000, and Blanchflower, 2001), where the estimated values of coefficient β are considerably higher (in absolute terms). Extensive surveys on this literature can be seen in Poot et al., (2004) and Montuenga and Ramos (2005).

[3] The macro evidence for the US in the late 1990s led the authors concerned with the natural unemployment rate to criticise the existence of a Phillips curve (see Coen *et al*., 1999, for a summary). However, some others argued that the Phillips curve still existed but that it had temporarily shifted inwards by fortuitous supply shocks and labor market developments (Gordon, 1998 and Katz and Krueger, 1999). More recently, the studies by Staiger *et*

The type of data used by Blanchflower and Oswald (1994) and the measurement of the dependent variable in their estimation are criticised by Blanchard and Katz (1997). First, the samples from the March CPS are too small to adequately measure the yearly wage variations in each state. Second, the use of annual earnings may be contaminated by the effect of worked hours. These two factors may bias the estimate of the autorregressive parameter ρ downwards. In order to control for this, Blanchard and Katz (1997) employ data from the merged Outgoing Rotation Group (ORG) in the CPS, which presents the advantages of larger sample size (almost twice as big as the simple CPS) and of reducing the measurement error in the computation of the hourly wages. They apply a two-step procedure to estimate the parameter ρ. In the first one, individual wages are regressed on worker's characteristics and region-by-year fixed effects. In the second one, the obtained dummy coefficients are used as a measure of average regional wages and are regressed on regional and fixed time effects as well as the unemployment rate and the lagged wages.[4] In this way, the common group (Moulton, 1990) and the composition biases are minimised (Solon et al., 1994). The test works in a similar way as in Blanchflower and Oswald (1994). If the estimate of the autorregresive parameter is close to one, the null hypothesis of a Phillips curve will not be rejected, whereas if it is close to zero the hypothesis of a wage curve will not be rejected. In their Table 2, Blanchard and Katz (1997), obtain estimates for the parameter ρ above 0,90, close to one, what is considered as evidence on favour of a US Phillips. Even more, when using the same data set that Blanchflower and Oswald (1994), they obtain an estimated value for ρ of 0.26, similar to the found by these authors. Blanchard and Katz (1997) hence hypothesise that the null wage persistence found by Blanchflower and Oswald (1994) is only due to statistical problems, and that when wages are measured appropriately, the Phillips curve representation still holds in the US.[5] On consequence, this indicates that a Phillips curve is achieved either with microeconomic data or with aggregate data.

However, some problems arise in the estimation of equation (2), especially when using time-short micro data. In particular, the lagged dependent variable appears as an additional explanatory regressor, which leads to an asymptotic correlation between the dependent variable and the error term. This generates a negative bias in the estimated value of the autorregressive coefficient of order 1/T, where T is the number of sample periods (see

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al., (2002) and Ball and Moffitt (2002) support the validity of the Phillips curve for the US in this period when the univariate trends of unemployment rate and productivity growth are incorporated.

[4] This approach has been also advocated by Nickell and Bell (1996) and Card and Hyslop (1997).

Nickell, 1981).[6] Additionally, both the presence of regional fixed effects and the plausible autocorrelation in the residuals make necessary a more adequate procedure. One possibility is proposed in Card (1995) who suggest an aggregate specification, that can be interpreted as a differenced version of (1). This is

$$\Delta ln(w_{rt}) = g_t + a_1\, ln(u_{rt}) + a_2\, ln(u_{rt-1}) + b_1\, y_{rt} + b_2\, y_{rt-1} + e_{rt}, \tag{3}$$

where the regional fixed effects have disappeared because of time-differencing, y_{rt} stands for the productivity-related characteristics affecting wages, g_t for the re-defined time fixed effects and e_{irt} denotes the new residual. Empirically, if $a_2 = -a_1$ a wage curve exists, whereas if $a_2 = 0$, a Phillips curve is observed. In the case of the US, Card and Hyslop (1997) use regional corrected wages from the merged ORG in the CPS to obtain that a_2 is non-significant and then non-rejecting the Phillips curve interpretation.[7] However, this procedure has some shortcomings, since although it overcomes the technical problems mentioned above, it lacks to consider the possibility of a dynamic wage curve, limiting itself to only test the wage curve against the Phillips curve hypotheses. This leads to considere some other alternatives.

Thus, Bell (1996), with the same data set and the same procedure that Blanchard and Katz (1997), uses the Generalised Methods of the Moments (GMM) to control for the dynamic bias in the wage equation (Arellano and Bond, 1991). He obtains an estimated value for ρ of around 0.83, that is to say, close to, but significantly below, one.[8] At the same time, this author shows how wages have evolved differently across US states by the different behaviour of prices, labour productivity and some other reasons. If these wage differences across states are not explicitly controlled for, the autoregressive coefficient will be upwards biased, since neither the fixed state effects nor the time effects will capture them. Including these state trends and using the second lag of wages as an instrument for the first lag, the author estimates a value for ρ around 0.56. This result leads to consider that there exists a high autocorrelation in wages, but the autoregressive coefficient is significantly different from one.

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[5] A similar conclusion is put forward in Partridge and Rickman (1997) using different demographic statistics for the US.

[6] The value of the bias corresponds to the case in which the lagged endogenous is the only regressor. Besides, if there exist other predetermined regressors, such as the individual characteristics or the fixed effects, the bias will be even greater. However, when the sample size is large, the bias becomes negligible.

[7] By contrast, Devereux (2001) with US data drawn from the PSID, presents evidence on favour of the wage curve hypothesis since he finds that $a_2 = -a_1$ is not rejected and that a_2 is significant. Buettner (1999) uses German data for rejecting the Phillips curve, even though it cannot be also rejected that unemployment coefficients are different, producing on consequence, an ambiguous result of the test. Black and FitzRoy (2000), using equation (3) finds support for the wage curve in the UK. However, when extending the equation to include the change in unemployment, a dynamic wage curve is a best representation of the reality.

[8] A similar conclusion is reached by Barth et al., (2002) using the same approach.

Therefore, it is not a Phillips curve. It has to be better thought of as a relationship between wage levels and unemployment rate, in which there exists a considerable sluggishness and the adjustment to a new equilibrium is relatively slow. In other words, there exists a wage curve with a partial adjustment towards the equilibrium, a fact that makes it similar to the Phillips curve.[9]

On concluding, a Phillips curve is observed in the US with aggregate data, but this hypothesis cannot be entirely supported with micro data. Some authors have devoted effort in trying to reconcile both empirical results. Roberts (1997b) and Whelan (2000) have proposed two alternative derivations to obtain, from a micro wage curve, an aggregate Phillips curve (new-Keynesian in the case of Roberts, 1997b, accelerationist, in the case of Whelan, 2000). Both of these studies put to one side the important economic meaning of the autoregressive coefficient. This is not the case of the work by Blanchard and Katz (1999), which offers an in-depth analysis of the theoretical arguments that allow for a dynamic specification of the relationship between wages and unemployment. This will be used in our empirical Section 3. In particular, they derive an expression like the following.

$$(w\text{-}p)_{rt}= \delta + \xi\, \phi\, (w\text{-}p)_{rt-1} + (1 - \xi\phi)y_{rt} + \beta\, u_{rt} + \upsilon_{rt}, \qquad (4)$$

that can also be rewritten as

$$\Delta(w\text{-}p)_{rt}= \delta + (1\text{-}\xi\, \phi)\, (w\text{-}p\text{-}y)_{\,rt-1} + (1 - \xi\phi)\, \Delta y_{rt} + \beta\, u_{rt} + \upsilon_{rt}. \qquad (5)$$

where $(w\text{-}p)_{rt}$ is the regional real wage, u_{rt} the rate of regional unemployment, y_{rt} the regional labour productivity, and υ_{rt} the disturbance term. ϕ, belonging to (0, 1), indicates the influence of the reservation rate, b_{rt}, in the real wage in a simple wage setting equation like $(w\text{-}p^{e})_{rt} = \phi b_{rt} +(1 - \phi)y_{rt} + \beta\, u_{rt}$, and ξ, also belonging to (0, 1), captures the influence of the lagged real wage in a simple equation for the reservation wage like $b_{rt} = \alpha + \xi\, (w\text{-}p)_{rt-1} + (1 - \xi)\, y_{rt}$ (for more details, see Blanchard and Katz, 1999).[10] If $\xi\phi=1$ then (5) is equivalent to the Phillips curve, whereas if it is close to zero, it gives support for a wage curve.

Expression (5) is nothing more than a modified version of the Error Correction Mechanism (ECM) in Sargan (1994). The $(w\text{-}p\text{-}y)_{t-1}$ is the error correction term and the coefficient $(1\text{-}\xi\phi)$ indicates whether a deviation of the real wage from the equilibrium level determined by

[9] The same conclusion is obtained by Bell et al,. (2002) for the case of the UK. When regional trends are not included in the regression, the autoregressive parameter is close to one, whereas it is below 0.75 when they are introduced.

[10] Here, it is assumed that the reservation wage is homogenous of degree one in the real wage and productivity in the long-run. Otherwise, technological progress wille lead to a persistent trend in the unemployment rate.

labour productivity and unemployment rate causes variation in wages inflation. A negative value of this parameter represents that real wage will adjust, finally, to the level determined by the productivity and the unemployment rate, but this will take some time, allowing for a dynamic wage curve specification. Using aggregate data, Grubb (1986), Turner et al. (1996), OECD (1997) and Blanchard and Katz (1997) have estimated the parameter $(1-\xi\phi)$ to be around -0.25 in European countries, whereas the coefficient was not significantly different from zero in the US. This seems to confirm the validity of a US Phillips curve (see also Staiger et al., 2002), whereas the European case is characterised by a modified version of the Phillips curve with error correction but high autocorrelation.[11] The interpretation, according to (5), indicates that at least one (if not both) of the components in $\xi\phi$ is, in Europe, less than one. This is a credible result, according to Blanchard and Katz (1999), if we consider that trade unions play a more relevant role in Europe when negotiating real wages ($\phi<1$), and that the black economy is probably also more extended in Europe ($\xi<1$) (see also Barth *et al.*, 2002).[12]

Using regional or individual data, the existence of a dynamic wage curve has been also found in several non-US countries as, for example, Argentina (Galiani, 1999), the UK (Cameron and Mullbauer, 2000), Germany (Pannenberg and Schwarze, 2000), Norway (Dyrstad and Johansen 2000) and Canada (Fares, 2002). A different result is obtained when studying the Nordic countries. For example, Albaek *et al.*, (2000) analyse various Nordic countries and find that the estimate of the autorregressive parameter is close to one, favourable to the Phillips curve, but β is, however, non-significant, what leads to reject both the wage and the Phillips curve specifications. They argue that centralised-type of negotiation, as such existing in these countries, may generate these kind of results.[13]

In summary, the literature is not conclusive. Using macro data, a wage Phillips curve is supported in the US whereas a slightly modified error-correction models well the situation in

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[11] However, using a similar approach, Whelan (2000) concludes that a dynamic wage curve is also present at the US, since the error correction term is significantly negative (even though lower, in absolute values, than in European countries).

[12] According to this reasoning, the derivation of the Phillips curve for the US, stemming from a wage curve, is only possible when the labour productivity does not influence either the wage setting process, or the subjective valuation of the reservation wage. Recent contributions have questioned such a restrictive assumption (see Gordon, 1998 and Staiger *et al.*, 2002).

[13] Similarly, Bårdsen and Nymoen (2003) have also derived an error-correction model, which encompasses both the wage curve and the Phillips curve specifications, to test the NAIRU hypothesis for the Norwegian case. The results obtained, however, are non-conclusive in that the Phillips curve is rejected but the wage curve is not supported. By contrast, Barth et al., (2002) for Norway and the UK obtain evidence on favour of a static wage curve, even though they recognise that their sample period may be too short to avoid the dynamic bias on the autorregressive parameter.

Europe. By contrast, the micro evidence casts doubt on even the Phillips curve specification for the US. In general terms, it seems that the relationship between wages and unemployment is more appropriately determined by a dynamic specification, in which unemployment has an influence that lingers on over time and wages. We test this hypothesis for the Spanish case on the basis of individual data coming from the second half of the nineties. Next section presents the empirical specifications and describes the database.

3. Empirical specification and data

The relationship between wages and unemployment for Spain has been basically studied on the aggregate or regional basis, as in Dolado and Lamo, (1993) Dolado and Jimeno (1997) Jimeno and Bentolila (1998). The most relevant findings are the elevate hysteresis in the rate of unemployment, the low wage elasticity to unemployment, the permanent, and even widening, unemployment differences across regions, joined to the low interregional mobility and flows into and out of the participation status. This evidence has been recently corroborated with individual data by García and Montuenga (2003) that estimates a static wage curve for Spain. It is our interest now to provide some evidence on the dynamic wage adjustment to unemployment shocks by allowing for a more general framework. In this way, we employ the specification derived in Blanchard and Katz (1999) so that, the dichotomy between wage curve or Phillips curve can be relaxed to obtain a more precise description of the functioning of the Spanish labour market.

Equation (4) can be adapted to an empirical specification like (2), which is estimated based directly on individual data to make use of its panel properties (see Bell et al., 2002). Thus, the estimated equation is

$$ln(w_{irt}) = a_i + \rho\, ln(w_{irt-1}) + f_r + r_t + b\, X_{irt} + \beta\, ln(u_{rt}) + \gamma\, t_r + \varepsilon_{irt}. \qquad (6)$$

where the inclusion of regional trends, t_r, to take into account regional differences in the evolution of wages, is empirically tested. The estimated value of ρ will yield an estimate of the parameter $\xi\phi$ in (5). β expresses the short-run elasticity, whereas $\beta/(1-\rho)$ the long-run elasticity. Analysis of the wage dynamics of this sort are used in Pannenberg and Schwarze (2000) and Iara and Traistaru (2004).

Data comes from the European Community Household Panel (ECHP), which collects information about wages and personal characteristics from a sample of 17,908 surveyed individuals. The study covers the period 1994-2001, with the ECHP being the only survey that offers micro-information for more than one year with respect to Spain. The employment statistics come from the official Spanish Labour Force Survey (*Encuesta de Población Activa*). A detailed description of the data set can be found in the Appendix. Hourly wages are expressed in real terms by deflating the nominal values by the corresponding regional CPI.[14]

The regional dimension of the data base is quite reduced since it is provided only at the NUTS 1 level, implying that only 7 regions are considered.[15] This makes that the total number of degrees of freedom is also quite reduced, 56 (7 regions times 8 years).[16] In order to enlarge this number, and then to be able to obtain more precise estimates of the wage adjustment, regional unemployment rates are also expressed by gender and by age group (see Kennedy and Borland, 2000, García and Montuenga, 2003). This facilitates that up to 448 different unemployment rates are available (7 regions by 8 years by 4 ages groups by 2 genders). Unemployment rate will be considered as exogenous, since previous studies (Montuenga et al., 2003 and García and Montuenga, 2003) have demonstrated its character of predetermined for Spain.[17] This seems plausible given the high degree of persistence in labour demand and the notoriously sluggish response of unemployment to shocks of any kind.

A further couple of comments are worth before describing the estimation procedure. First, some recent studies have concerned about the spatial influence of neighbouring regional unemployment rates in individual wages (see Buettner, 1999, Longhi et al., 2003, Iara and Traistaru, 2004). In Spain, most of the wage bargaining takes place at the sectoral provincial level (NUTS 3), which is clearly more disaggregated than the available in our data. Moreover, given the large extension of the NUTS 1 regions, the possibility of commuting is unlikely. Finally, the way that unemployment rates are defined (region by age by gender) makes very difficult to figure out the existence of interdependence of unemployment rates on wages. All this has convinced us to not consider the plausible problem of spatial autocorrelation and, still, be confident on our estimates.

Secondly, Blanchard and Katz (1999) consider that estimating equation (5) in aggregate terms is more adequate since, given inter-regional mobility of workers, the wage in one state will depend not only on the particular state lagged wage, but also on the aggregate lagged

[14] It is not possible to distinguish between normal working wages and overtime wages, which may bias the estimation (Black and FitzRoy, 2000, Hart, 2003). However, this is an inescapable limitation of our data since detailed information for individual wages in Spain is not available in our data set (and in no other dataset in panel data form).

[15] The NUTS 1 regions are obtained from simple grouping of the 17 Spanish Autonomous Communities (see Appendix 1).

[16] Note that in the regression equation, the unemployment rate is defined at a higher level, regional, that the dependent variable, individual. On consequence, the regional dimension is the restricting factor in the availability of the degrees of freedom.

[17] This is a generalised finding elsewhere (see Blanchflower and Oswald, 1994, Bell, 1996, Black and FitzRoy, 2000, and Bell et al., 2002), except in Germany (see Baltagi and Blien, 1998, Baltagi et al., 2000).

wage. If this effect is not explicitly considered, it will remain hidden in the fixed time effects, leading to a negative bias in the value of ρ. This source of bias, however, is expected to be unimportant in the European countries, because of the lack of inter-regional mobility. As an alternative, instead of including fixed time effects, explicitly aggregate variables may be used to properly model such influences, even though a richer time dimension is required. Galiani (1999), for Argentina, and Bell *et al.*, (2002), for the UK, find that aggregate variables do not really influence the dynamic wage adjustment. On consequence, we rely on the estimation of the time fixed effects in equation (6).

As mentioned before, the inconsistency of the Least Square Dummy Variable (LSDV) estimator is probably to arise due to short period data available. In this case, the GMM estimator is the best choice for controlling it (Arellano and Bond, 1991). Although some estimators have also been suggested afterwards (Kiviet, 1995, Baltagi, 1995), they does not perform better than GMM for T<10, see Judson and Owen (1999). On consequence, we apply the Arellano-Bond GMM procedure, which include the following steps. Equation (6) is first-differenced in order to remove the fixed effects, and then estimate it using instrumental variables. As instruments, all lags of the variables in levels are used. Since these are correlated with differenced variables, but uncorrelated with difference error terms (unless the error terms in levels display serial correlation), they provide a set of valid instruments. While first order autocorrelation in the first-differenced residuals complies with the estimator's consistency requirements, it is necessary that the differenced error terms are free of second order autocorrelation. Arellano and Bond (1991) propose two GMM estimators (one-step GMM and two-step GMM), which exploit all available lagged values of the dependent variables as instruments. One-step GMM simply takes account of the fact that the first differenced error term of equation (6) is MA(1) with unit root. Two-step GMM uses the estimated residuals of one-step GMM to construct a weighting matrix, which yields a two-step GMM estimator, which is robust to general cross-section and time-series heteroscedasticity. Both GMM estimators hinge upon the assumption that there is no second-order serial correlation for the disturbances of the first differenced equations. Therefore we employ a Sargan test of over-identifying restrictions for the GMM estimates. Wald tests are robust to general heteroscedasticity.

4 Results of the estimation

In Table 1, we report the estimates of the relevant coefficients of equation (6) and their corresponding standard errors. We use three alternative estimators: the LDSV, the one-step

GMM and the two-step GMM. On principle, the one-step GMM estimation is the more adequate, even though we carry out Sargan test for choosing the best option.

We report the one-step GMM estimator with robust standard errors. Since the standard errors from the two-step GMM are frequently found downward biased (AB, 1991), for inference on single variables' coefficients we rely on the one-step estimator. For the choice between specifications, however, we use the Sargan test of over-identifying restrictions after the corresponding two-step GMM estimator (no robust Sargan test using one-step residuals is available). Consistency of the estimator requires the absence of second-order autocorrelation in the differenced residuals, that is checked by the respective tests developed by Arellano and Bond (1991).

Table 1. Estimation results of the dynamic wage equation

	GMM1	GMM1 robust	GMM2
Rho	0.088	0.088	0.098
	(9.66)	(5.67)	(6.44)
Beta	-0.068	-0.068	-0.073
	(-2.01)	(-2.03)	(-2.19)
Regional fixed effects	X	X	X
Time fixed effects	X	X	X
Individual fixed effects	X	X	X
Regional trends	X	X	X
Long-run elasticity	-0.075	-0.075	-0.081
m1	-49.59	-17.19	-17.14
	(0.0000)	(0.0000)	(0.0000)
m2	0.73	0.40	0.56
	(0.4643)	(0.6865)	(0.5757)
Sargan test	126.67		62.93
	(0.0000)		(0.0000)

The coefficient of the lagged dependent variable is clearly significant with a value lower than 0.1. This means that the reaction of wages to changes in unemployment rates is very fast, indicating that the Spanish labour market is very close to a competitive standard model. This result is not new. Montuenga et al. (2006), with the same database, show that, among five EU countries, Spain exhibited a labour market very close to the functioning of a spot labour market. In this sense, it must be noted that the period analysed, 1994-2001, coincides with a recovering phase of the Spanish economy in which employment increased, the unemployment rate reduced and wage groth was moderate. The unemployment coefficient is close to –0.07 in the range of the typical finding in the literature (see Nijkamp and Poot, 2005). The long-run elasticiy is hence –0.08.x

5. Conclusions

The aim of this article has been to study the relationship between individual wages and regional unemployment rates in Spain considering a dynamic specification. The existing literature for Europe, which includes the Spanish case, has show that labour markets are better modelled by a wage curve representation by which the level of the wages are linked to the level of unemployment. The evidence also shows that the effect of unemployment in wages in persistent so that the total impact is not fully translated in just one year, but it needs some time to exert an inverse influence on wages. This is called the dynamic wage curve. Our target is to estimate this for Spain with individual data coming from the eight waves of the ECHP 1994-2001 using an specification that is common in the empirical studies. This takes into account both the reduced time dimension available and the subsequent bias arising from the estimation of dynamic panel data models with fixed effects. Therefore, we have used a GMM estimator to test the degree of sluggishness in the responde of wages against changes in the unemployment rates.

Results seem to show that a statice wage curve models well the case of Spain since the autorregressive parameter is very close to 0. Consequently, the short and long-run unemployment elasticities are very similar and around –0.07, close to the –0.1 empirical lwa of economics claimed by Blanchflower and Oswald (1994, 2006), and that coincides with the mode found by Nijkamp and Poot (2005) in their meta-analytic study of the wage curve. This shows the relatively flexible labour market in Spain compared with other EU countries. However, this result must be taken with caution given that the period analysed coincides with an expansive phase of the Spanish economy during which unemployment reduced sharply, employment increased strongly and real wage growth waws contained.

It has been previously stated that the exact structure linking wages to unemployment is important to determine the NAIRU or, at least, its evolution over time. Some reasons make this exercise not interesting in our analysis. First, and fundamental, the time dimension of our study is only seven years, by which most of the implications on the NAIRU are seriously restricted. In this respect, aggregate studies using time-series information are more appropriate. Second, NAIRU rates are usually derivated from the dynamic wage curve together with a price setting formulation. The choice of this pricing rule has important implications. Thus, Whelan (2000) and Bell et al. (2002) use specifications that allow for obtaining an aggregate Phillips curve, for any microeconomic configuration of the wage curve. By contrast, the standard specification used by Blanchard and Katz (1999) predict difference behaviour whether a Phillips or wage-type curve fits the data. Third, it seems more reasonable in order to model the NAIRU behaviour to establish a general framework allowing for the simultanous consideration, not only of the price and wage setting equations but also the unemployment rate. This is, however, beyond of the scope of this paper and matter for future research.

APPENDIX

The sample from the European Community Household Panel (ECHP) is made-up of 17,908 individuals that were surveyed personally. The final size of the sample is reduced to 2,715 employees per year, forming an overall sample of 8,145. Some individuals have been discarded: those individuals who are not workers, as well as the self-employed, workers in agriculture or fishing, civil servants and members of the military. The survey provides information on earnings, as well as job and personal characteristics. In particular, the variables that have been used are:

- *Log real wage per hour.* Nominal wages are computed as the ratio between annual earnings and the number of hours worked in a week times the number of weeks worked in a year (50). They then are deflated by the corresponding weighted regional CPI, which is own-elaborated at the NUTS 1 level from the NUTS 2 information, that is provided by the Spanish National Statistic Institute.
- *Log unemployment rate*. The variable measures the unemployment rate by region by gender and by age group (the corresponding age groups being between 16 and 19, between 20 and 24, between 25 and 54 and over 55). The data are drawn from the Spanish Labour Force Survey.
- *Age*. This is used to proxy working experience. We also introduce it to the second power (divided by 100) to shape the decreasing returns on experience.
- *Gender*. Male=1 and female=0.
- *Marital status*. Married=1, otherwise=0.
- *Part-time work*: Working less than 30 hours per week=1. Working more than 30 hours=0.

- *Education level of the employee*: This includes 3 categories: primary and no formal education, secondary education, and university and technical education.
- *Occupation group*. This variable describes the type of specialisation of the employee, divided into 8 categories: managers, professional technician, supporting professional technician, administrative, simple services, qualified craftsmen and technician, assemblers, and non-qualified workers.
- *Seniority*. The number of years that a worker has been employed in his/her current position. This includes 3 categories: less than 2 years, between 2 and 10 years, and more than 10 years.
- *Type of activity*. In principle, this classifies into agricultural, industrial and service activities. However, once we eliminate agricultural workers, it becomes a dummy variable. Industry worker=1, Services worker=0.

The ECHP offers regional disaggregation for the seven NUTS I ("nomenclature of territorial units for statistics") areas of Spain (see Table A).

Table A. Regional (NUT I and NUT II) disaggregation in the five sample countries

SPAIN	NUT I	NUT II
Region 1	North West	Galicia, Asturias, Cantabria
Region 2	North East	País Vasco, Navarra, La Rioja, Aragón
Region 3	Community of Madrid	Comunidad de Madrid
Region 4	Center	Castilla-León, Castilla-La Mancha, Extremadura
Region 5	East	Cataluña, Comunidad Valenciana, Baleares
Region 6	South	Andalucía, Murcia, Ceuta y Melilla
Region 7	Canary Islands	Canarias

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