Investigaciones Regionales

# Wage dynamics in Spain: evidence from individual data (1994-2001)

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**ABSTRACT:** In this paper, we test the hypothesis of a wage curve against a Phillips curve for Spain, within a dynamic framework that allows for both of these, and for more general alternatives. To this end, we use data from the European Community Household Panel, providing micro-information for the period 1994-2001. The results indicate that, contrary to the situation in other European countries, the wage adjustment occurs in just one period, with the elasticity of wages to unemployment being close to the «empirical law of economics» of –0.1.

**JEL Classification**: J30, J60, J64, R23.

**Keywords**: Phillips curve, Wage Curve, dynamic panel data, GMM.

## Dinámica salarial en España: Evidencia a partir de datos individuales (1994-2001)

**RESUMEN:** En este artículo, contrastamos para España la hipótesis de una curva de salarios frente a la curva de Phillips en un marco dinámico que permite éstas y otras alternativas más generales. Para ello utilizamos datos del Panel de Hogares de la Unión Europea, el cual proporciona información individual para el periodo 1994-2001. Los resultados indican que, al contrario de lo observado en otros países europeos, el ajuste de los salarios tiene lugar en un solo periodo, siendo la elasticidad de largo plazo de los salarios a variaciones en el empleo próxima a la «ley empírica de la economía» de –0,1.

**Clasificación JEL**: J30, J60, J64, R23.

**Palabras clave**: Curva de Phillips, curva de salarios, panel de datos dinámico, MGM.

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## 1. Introduction

The dynamics of wage adjustment is a controversial issue. The negative relationship of the Phillips curve, between the growth rate of wages and the unemployment rate, became the cornerstone of the Keynesian synthesis and macro-econometric modelling, since it seemed to capture well the price and wage adjustment mechanisms. Economic authorities, by following a «fine tuning» policy, were then able to select an appropriate combination of inflation and unemployment rates. The Phillips curve has been widely used to model the supply side of the economy, such that, when confronted with the demand side, their intersection determines equilibrium values of product and price. Under this Phillips curve view, wages and prices tend to adjust to excess demand in such a way that, sooner or later, the economy moves towards the equilibrium locus. Thus, supply and productivity shocks have no long-run effects on real variables.

The Phillips curve, however, has been disputed, from the micro-economic perspective in the US by Blanchflower and Oswald (1994), who claim that «the Phillips curve is probably a mis-specified aggregate wage curve», and consequently, «the idea of a Phillips curve may be inherently wrong». In Europe, and in other OECD countries, the hypothesis of a Phillips curve had already been empirically challenged since the mid-1980s (Grubb, 1986), in favour of a dynamic wage relationship. This evidence has theoretical foundations, since non-competitive theories of the labour market predict a negative relationship at the micro stage between the level of wages and the unemployment rate (see Shapiro and Stiglitz, 1984). This relationship, called «the wage curve», represents an equilibrium locus of wages and unemployment, resulting from the optimising behaviour of the agents involved in the bargaining process (see Blanchflower and Oswald, 1994; Card, 1995). The wage curve is depicted as 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 it is confronted with the demand labour curve, determines an equilibrium wage above that of the labour market, and an «equilibrium unemployment rate» (Woodford, 1994). The growing acceptance of the wage curve in modelling labour markets has led to its being reinterpreted 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, 2011). Under the wage curve model, both aggregate supply and productivity shocks will have permanent effects on unemployment and output.

Thus, the discussion about whether unemployment is related to growth or to the level of wages is not meaningless, but has powerful consequences for our understanding of the labour market and of the economy as a whole. First, in determining the dynamic effects, if any, of demand and supply wage variables on the natural rate of unemployment, and second, in providing an empirical guide for policy modelers to appraise the effects of shocks on price inflation and on the inflation-unemployment trade-off. In 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 (Montuenga-Gómez and Ramos-Parreño, 2005), labelling this alternative view as dynamic wage curve.

It is also of great interest in retrieving evidence about the degree of wage flexibility in a country, especially in those countries belonging to a monetary union. It is well known that both exchange rate and monetary policies are no longer independent in the European Monetary Union (EMU), so that membership of the EMU imposes further requirements on factor market flexibility. The European Commission (2003, p. 155) calls attention to this by stating that «the formation of EMU is often taken to put further demands on the flexibility of wages to compensate for the lack of (national) instruments to deal with economic disturbances. If wages are too rigid, the necessary adjustment will come slowly and with considerable economic and social costs».

The aim of this paper is to shed some light on this debate for Spain by using individual data to make comparisons with other EU countries. This paper is then similar in nature to those of Bell *et al.* (2002), Iara and Traistaru (2004), Blanchflower and Oswald (2005) and Baltagi *et al.* (2009), extending previous evidence to the case of Spain (where no prior studies exist). Results show that, contrary to other countries, the Spanish wage curve seems to be static, with little or no autoregression in pay. Section 2 surveys the literature and describes the main concerns of our research. Section 3 presents the empirical specifications addressed in the applied analysis and describes the data base. Section 4 presents the empirical results, and Section 5 lays out our conclusions.

## 2. The wage curve and the Phillips curve

The idea of a wage curve in micro-economic terms can be opposed to the existence of a Phillips curve in aggregate terms (Blanchflower and Oswald, 1994). First, the 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 normally obtained from disaggregated data of longitudinal household or individual surveys, whereas the Phillips curve is usually estimated with macro-time series of unemployment and wage inflation. A further difference lies in the economic meaning of each concept. The Phillips curve is a set of disequilibrium points that represent the adjustment process in a competitive model of the labour market. In contrast, 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. Recently, some effort has been devoted to reconcile both approaches in a unified framework, see for example Blanchard and Katz (1999) or Whelan (2000). Similarly, Campbell (2008), by developing an efficiency wage model, shows that the wage equation looks like a wage curve when regional economies are modelled, and looks like a Phillips curve at the national level.

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

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

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 $\varepsilon$ is the remainder error term. Time-period effects control for all those variables that vary over time but that are common to all regions (*i. e.* business cycle variables), whereas variables that are time-invariant but specific to each region, such as endowments, amenities, facilities, etc., are contemplated by including regional fixed effects.

This double logarithmic expression has been justified as providing the best results (see Blanchflower and Oswald, 1994) and has been widely applied. The coefficient $\beta$ is, therefore, the elasticity of wages with respect to unemployment, with a negative estimated value, thereby demonstrating the existence of a wage curve. The inclusion of regional fixed effects allows us to capture any permanent component of the relationship between wages and unemployment, so that the unemployment coefficient $\beta$ is only reflecting the temporary component of that relationship. Expressions like (1) have been estimated for many countries, showing, as a general result, that wage elasticity to unemployment lies in the range (–0.20, –0.05) for most cases. This general result has given rise to calls to recognise «an empirical law of economics» [1].

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

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

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

This implies testing the extreme alternatives of a competitive modelling of the labour market, the Phillips curve, $\rho = 1$, in which supply shocks do not have permanent effects on unemployment and income compared to a non-competitive framework, wage curve $\rho = 0$, by which wages adjust almost instantaneously to changes in un-

[1] Extensive surveys of this literature can be found in Blanchflower and Oswald (2005), Nijkamp and Poot (2005) and Montuenga-Gómez and Ramos-Parreño (2005). Some of the most recent country-specific contributions are Livanos (2010) for Greece, Ammermuller et al. (2010) for Germany and Italy, and Deller (2011) for the US.

employment, so that supply shocks have persistent effects on output and unemployment. However, there is a wide range of possible values, $0 < \rho < 1$, as a reflection of an intermediate situation, indicating that supply shocks have permanent impact on wages, but some time is required to exert such impact; more rapidly when $\rho \to 0$, more slowly when $\rho \to 1$. This is the dynamic 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 $\rho$ is close to zero. This result suggests that wages adjust rapidly to the unemployment rate, which constitutes the starting point in claiming the death of the Phillips curve. These authors 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 their use of micro data is considered to be most appropriate in unveiling the truth. This conclusion presented a significant challenge to the predominant evidence shown by the aggregate studies for the case of the US, always favourable to the Phillips curve, and spurred empirical analysis in order to study the phenomenon of wage persistence in greater depth.

Blanchard and Katz (1997, 1999) disputed those results, arguing that the type of data used and the measurement of the dependent variable were inadequate, since the samples from the March CPS are too small to properly measure the yearly wage variations in each state, and the use of annual earnings may be contaminated by the effect of worked hours. These two factors may bias the estimate of the autoregressive parameter $\rho$ downwards. In order to control for this, the authors employed data from the merged Outgoing Rotation Group (ORG) in the CPS, which presents a larger sample size (almost twice as large as the simple CPS) and reduces the measurement error in the computation of the hourly wages. They then apply a two-step procedure to estimate the parameter $\rho$, obtaining estimates above 0.90, close to one. Later research for the US, including Bell (1996), Whelan (2000) and Blanchflower and Oswald (2005), conclude that this parameter is estimated to be strictly positive, but significantly lower than 1. That is to say, a dynamic wage curve exists.

Using regional data, a dynamic wage curve has also been found in several non-US countries —for example, Germany (Pannenberg and Schwarze, 2000; Ammermuller *et al.,* 2010) and Norway (Dyrstad and Johansen, 2000)—. With individual data, Bell *et al.* (2002) for the UK, Iara and Traistaru (2004) for Bulgaria and Poland, and Baltagi *et al.* (2009, 2012) for Germany, have also found that the wage curve is dynamic. A different result is obtained in studies of the Nordic countries. For example, Albaek *et al.* (2000) analyse a group of Nordic countries and find that the estimate of the autoregressive parameter is close to one, favourable to the Phillips curve, but $\beta$ is, however, non-significant, which leads to the rejection of both the wage and the Phillips curve specifications. They argue that centralised-type negotiations, such as are found in these countries, may generate this kind of result [2].

[2] By contrast, Barth *et al.* (2002) for Norway and the UK obtain evidence in favour of a static wage curve, even though they recognise that their sample period may be too short to avoid the dynamic bias of the autorregressive parameter.

Overall, at the microeconomic level, the US has more auto-regression in wages than many other nations, but there is much evidence across nations of a dynamic wage curve [3]. It seems, then, that the relationship between wages and unemployment is more appropriately determined by a dynamic specification, in which unemployment has an influence that lingers over time and wages. We test this hypothesis for the Spanish case on the basis of individual data from the second half of the 1990s. The following section presents the empirical specifications and describes the database.

## 3. Empirical specification and data

The relationship between wages (or wage inflation) and unemployment, for Spain, has been studied in the aggregate —Dolado and Jimeno (1997), Galí and López-Salido (2001), Bentolila *et al.* (2008)— and on a regional basis —Jimeno and Bentolila (1998), Bande and Karanassou (2009)—. The most significant findings are the elevated hysteresis in the rate of unemployment, the low wage elasticity to unemployment, the permanent, and widening, unemployment differences across regions, and both low inter-regional mobility and flows into and out of the participation status. This evidence is corroborated with individual data by García and Montuenga (2003), who estimate a static wage curve for Spain. It is our interest now to provide some evidence for the dynamic wage adjustment to unemployment shocks by allowing for a more general framework. To do this, we employ the specification derived in Bell *et al.* (2002), based directly on individual data, to make use of the panel properties of our database. Thus, the estimated equation is

$$ln(w_{irt}) = a_i + \rho\, ln(w_{irt-1}) + f_r + r_t + b\, X_{irt} + \beta\, ln(u_{rt}) + \gamma\, t_r + \varepsilon_{irt} \qquad (3)$$

where the inclusion of regional trends, $t_r$, takes into account regional differences in the evolution of wages, as suggested by Bell *et al.* (2002). Nevertheless, this inclusion is tested empirically in our approach. The estimated value of $\rho$ captures the dynamic behaviour of the relationship between unemployment and wages; $\beta$ expresses the short-run elasticity and $\beta/(1-\rho)$ the long-run elasticity. Analysis of wage dynamics of this sort is used in Bell *et al.* (2002), Iara and Traistaru (2004), and Baltagi *et al.* (2009, 2012).

Our 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 panel that offers micro-information on wages and individual characteristics for more than one year in Spain. From an international perspective, this is a short period of analysis (see Bell *et al.,* 2002; Baltagi *et al.,* 2009) making the measurement of the dynamic character of the wage curve more difficult [4]. This is a potential caveat

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[3] At the macro level, Madsen (2009), however, finds support for the Phillips curve for 18 OECD countries.

[4] However, it is similar to those in Barth *et al.* (2002) and Iara and Traistaru (2004), for example.

of our study, which cannot be solved with the existing data for Spain [5]. 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 [6].

The regional dimension of the data base is reduced, since it is provided only at the NUTS 1 level, resulting in only 7 regions being considered [7]. This also reduces the total number of degrees of freedom to 56 (7 regions times 8 years) [8]. In order to enlarge this number, and then 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 provides up to 448 different unemployment rates (7 regions by 8 years by 4 age groups by 2 genders). The unemployment rate will be considered exogenous, since prior studies (García and Montuenga, 2003) have demonstrated its predetermined nature for Spain [9].

A final comment is worthwhile before describing the estimation procedure. Some recent studies have been concerned with the spatial influence of neighbouring regional unemployment rates on individual wages (see Iara and Traistaru, 2004; Longhi *et al.,* 2006; Deller, 2011; and Baltagi *et al.,* 2012). In Spain, most wage bargaining takes place at the sectoral provincial level (NUTS 3), which is clearly more disaggregated than what is available in our data. Moreover, given the size 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 it difficult to determine the existence of interdependence of unemployment rates across units. All this has convinced us not to consider the problem of spatial autocorrelation, while retaining confidence in our estimates.

As for the estimator used, the inconsistency of the Least Square Dummy Variable (LSDV) estimator arises from the short-period available data. Since the lagged dependent variable appears as an additional explanatory regressor, it leads to an asymptotic correlation between the dependent variable and the error term, generating a negative bias in the estimated value of the autorregressive coefficient of order 1/T,

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[5] There is no other data set providing information on wages at the individual level for a longer span of years.

[6] It is not possible to distinguish between normal working wages and overtime wages, which may bias the estimation (Black and FitzRoy, 2000). Detailed information for individual wages in Spain is not available in our data set (and in no other dataset in panel data form). Additionally, it must be noted that, as usual, the survey-provided wages tend to be lower than the actual earnings. Both facts are inescapable limitations of our data.

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

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

[9] 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 (Baltagi *et al.,* 2009).

where T is the number of sample periods (see Nickell, 1981) [10]. In this case, the GMM estimator is the best choice for controlling this bias (Arellano and Bond, 1991) [11]. Although some other estimators have also been suggested (Kiviet, 1995), they perform no better than GMM for T < 10. As a consequence, one first alternative is to apply the Arellano-Bond GMM procedure, also called «difference GMM», which involves first taking differences in Equation (3) in order to remove the fixed effects, and then running the estimate using all lags of the variables in levels, as instruments. 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. This can be checked by examining the $m_1$ and $m_2$ tests for serial correlation in the first-differenced residuals, following Arellano and Bond (1991).

However, efficiency can be dramatically improved, provided that first differences of instrumenting variables are uncorrelated with the fixed effects, so that the number of instruments is augmented (Arellano and Bover, 1995; Blundell and Bond, 1998). Operationally, a system of two equations is built —the original equation as well as the transformed equation— that is known as «system GMM». In the system GMM estimator, the differenced equations, using level instruments, are combined with equations in levels using differences as instruments. Blundell and Bond (1998) show that first differences of the series may be uncorrelated with the industry-specific effects under stationarity. This allows the use of lagged differences as instruments for the levels equation. One further advantage of this estimator is that its efficiency is not affected in the cases in which the dependent variable is close to a random walk. In such cases, difference GMM performs poorly, since past levels convey little information about future changes, so that untransformed lags are weak instruments for transformed variables. Errors in the two-step estimation of system GMM are corrected according to Windmeijer (2005), so that they are superior to robust one-step. Finally, we employ Sargan-Hansen tests of over-identifying restrictions for the GMM estimates.

## 4. Results of the estimation

In table 1, we report the estimates of the relevant coefficients of equation (3) and their corresponding robust standard errors. Estimates presented here are com-

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[10] The value of the bias corresponds to the case in which the lagged endogenous is the only regressor. In any case, 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.

[11] Arellano and Bond (1991) show that their test is preferred to the OLS, Within Groups, and Anderson-Hsiao difference and levels estimators using Monte Carlo simulations. The difference GMM exhibits the least bias and variance in estimating the parameter of interest.

puted using the program *xtabond2* in Stata[12]. We begin by showing the results corresponding to the inconsistent estimators OLS and LDSV. As indicated by Roodman (2009), these two estimators produce extreme values, between which the true parameter should lie. In the following two columns, the estimates according to the difference GMM proposed in Arellano and Bond (1991) are presented: the one-step GMM with robust standard errors, and the two-step GMM, exploiting 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 (3) is MA (1) with unit root. Two-step GMM uses the estimated residuals of one-step GMM to construct a weighting matrix that yields a two-step GMM estimator, which, in turn, is robust to general cross-section and time-series heteroskedasticity. Both GMM estimators hinge on the assumption that there is no second-order serial correlation for the disturbances of the first differenced equations, which is checked by the respective tests $m_1$, $m_2$, presented below the estimates. The next columns present the estimates obtained from the «system GMM» suggested by Arellano and Bover (1995) and Blundell and Bond (1998), which are shown to be more efficient, considering alternative specifications. The preferred specification is chosen according to the Sargan-Hansen test of over-identifying restrictions, which is also helpful in assessing the validity of the instruments.

We have explored diverse specifications using the flexibility provided by the *xtabond2* program. The unemployment rate has been introduced in levels and also one-period lagged, resulting in the latter being non-significant; we have also tested whether it is endogenous, or not, with the values produced by the Sargan-Hansen tests rejecting the null. Similarly, the length of lags of the variables used as instruments has been chosen based on these tests, obtaining that the null of validity of instruments is accepted when only the dependent variable is considered as endogenous, and instruments are restricted to take the second to the fifth lagged values. All specifications appear to capture the relevant dynamics, since no second order residual correlation is evident. Regional, time, and individual fixed effects are found to be jointly significant and are retained in the regressions. By contrast, regional trends are found not to be individually significant, and they are excluded from final estimations. Whereas, in most of the specifications, Sargan-Hansen tests reject the null that the over-identifying restrictions are valid, in the two final columns this hypothesis is not rejected. Hence, we restrict our comments on these estimated values, though point estimates do not vary significantly across different specifications.

The unemployment coefficient is about –0.07, within the range of the typical finding in the literature (see Nijkamp and Poot, 2005), approaching the «empirical law of economics» of –0.1. The coefficient of the lagged dependent variable is found to be statistically non-significant (it is, however, in the «difference system» estimation, but with a very low value, below 0.1). This indicates, for the period

[12] For more on this matter, see Roodman (2009).

1994-2001, almost no autoregression in pay, with wages thereby adjusting very rapidly to changes in unemployment. From an economic point of view, this result can be interpreted as the Spanish labour market being relatively more sensitive to supply shocks than in other countries. In this sense, the labour market in Spain has been characterised by both an enormous rate of temporary contracts, over one third in this period, and an increasing relevance of immigration (Bentolila *et al.,* 2008), providing an important tool in rapid wage adjustment to shocks. This result is not at all normal in other countries. Whereas values, more or less closer to unity, have been obtained for the US, in Europe most countries are between the 0.3 coefficient in Germany and 0.5 in the UK (see Baltagi *et al.,* 2009; Bell *et al.,* 2002; and Montuenga-Gómez and Ramos-Parreño, 2005). Only in Norway (Barth *et al.,* 2002) and Romania (Iara and Traistaru, 2004) have values similar to ours been found [13].

Since these two features, high temporary rates and the increasing participation of immigrants, are peculiar to Spain, we have focused on their potential influence. To check for this, we have run several regressions similar to that in the last two columns, distinguishing, first, between workers with permanent contracts against those with fixed-term contracts, and second, between native and immigrant workers. In the case of permanent workers, the coefficient of the lagged dependent variable is 0.44, within the range of global estimates for other EU countries, and the unemployment coefficient is non-significant, probably reflecting the low responsiveness of permanent worker payments to changes in unemployment during that period. By contrast, in the case of fixed-term employees, the autoregressive parameter is non-significant, with the unemployment coefficient being statistically significant with a value of –0.111, thereby reflecting higher wage flexibility with respect to the case of permanent workers. More noticeably, the unemployment coefficient of immigrants becomes –0.204, being non-significant in the case of native workers [14]. These both results may support the view of a more responsive labour market in Spain. Although wage flexibility is reduced, the elasticity of wages to unemployment is, in absolute values, above –0.1 for both temporary and immigrant workers, providing the Spanish labour market with a rapid adjustment of wages; that is to say, the reaction of wages to unemployment is low, but it is rapid. However, it must be noted that the period of analysis may be too short to capture the dynamic nature of wage adjustments as, for example, in the case of Norway and the UK (Barth *et al.,* 2002). The unavailability of data samples covering a longer period prevents us from achieving undisputed conclusions.

[13] *A special case is Italy, for which Ammermuller et al. (2010) find a negative autoregressive coefficient for the period 1991-2004. The authors argue this result is probably due to the continuous decline observed in Italian real wages since 1992.*

[14] Overall results are not presented to preserve space, but they are available from the authors.

**Table 1.** Estimation results of the dynamic wage equation

| | *OLS* | *LDSV* | *Difference GMM* | | *System GMM* | | | |
|---|---|---|---|---|---|---|---|---|
| | | | *GMM1* | *GMM2* | *GMM1* | *GMM2* | *GMM1* | *GMM2* |
| $Ln(w_{irt-1})$ | 0.322*** (0.004) | –0.064*** (0.063) | 0.088*** (0.015) | 0.098*** (0.015) | 0.068 (0.092) | 0.072 (0.095) | 0.073 (0.089) | 0.081 (0.094) |
| $Ln(u_{rt})$ | –0.012 (0.016) | –0.087*** (0.023) | –0.069*** (0.033) | –0.072*** (0.033) | –0.078*** (0.025) | –0.081*** (0.026) | –0.065*** (0.023) | –0.067*** (0.024) |
| *Secondary education* | 0.093*** (0.009) | 0.012 (0.015) | –0.002 (0.020) | –0.002 (0.020) | 0.117*** (0.018) | 0.119*** (0.013) | 0.126*** (0.012) | 0.120*** (0.013) |
| *Higher education* | 0.191*** (0.011) | 0.030* (0.022) | 0.006 (0.026) | 0.005 (0.026) | 0.267*** (0.031) | 0.274*** (0.017) | 0.275*** (0.016) | 0.273*** (0.017) |
| *Experience* | 0.125*** (0.001) | 0.036*** (0.006) | 0.015*** (0.006) | 0.011** (0.006) | 0.024*** (0.004) | 0.023*** (0.002) | 0.024*** (0.002) | 0.025*** (0.013) |
| *Experience squared* | –0.016*** (0.002) | –0.050*** (0.007) | –0.067*** (0.015) | –0.065*** (0.015) | –0.034*** (0.007) | –0.034*** (0.004) | –0.033*** (0.003) | –0.035*** (0.004) |
| *Married* | 0.052*** (0.008) | 0.070*** (0.022) | –0.022 (0.034) | –0.019 (0.034) | 0.087*** (0.019) | 0.088*** (0.012) | 0.094*** (0.012) | 0.090*** (0.012) |
| Long–run elasticity | –0.017 | –0.082 | –0.092 | –0.080 | –0.084 | –0.087 | –0.070 | –0.073 |
| Regional fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Time fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Individual fixed effects | No | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| F tests (p-values) | 188.75 (0.000) | 51.15 (0.000) | 30.85 (0.000) | 33.58 (0.000) | 223.23 (0.000) | 225.18 (0.000) | 227.51 (0.000) | 231.25 (0.000) |
| $m_1$ (p-values) | | | –17.19 (0.000) | –17.14 (0.000) | –17.30 (0.000) | –17.00 (0.000) | –5.65 (0.000) | –17.30 (0.000) |
| $m_2$ (p-values) | | | 0.40 (0.689) | 0.56 (0.575) | 0.57 (0.571) | 0.47 (0.641) | 0.08 (0.933) | 0.15 (0.882) |
| Sargan test (p-values) | | | 125.16 (0.000) | 62.93 (0.000) | 120.31 (0.000) | 69.73 (0.000) | 21.29 (0.265) | 21.29 (0.265) |
| Hansen test (p-values) | | | 61.34 (0.000) | 45.62 (0.000) | 73.96 (0.000) | 42.57 (0.000) | 12.03 (0.846) | 12.03 (0.846) |
| Number of instruments | | | 55 | 55 | 63 | 43 | 58 | 58 |

Control variables include: marital status, 3 educational levels; 8 occupational categories; experience and experience squared; 2 industry categories, gender, 5 firm size categories (detailed description of all these variables in the Appendix. Clustered by unemployment rates (see text) standard errors in parentheses. In GMM2, the Windmeijer (2005) correction is included. GMM1 is one-step GMM. GMM2 is two–steps GMM. The number of observations is NT=27,954.
***, **, * indicates significance at 1%, 5%, 10% level, respectively.
$m_1$ tests for AR (1) in first differences
$m_2$ tests for AR (2) in first differences
Sargan-Hansen tests. The hypothesis null is that over-identifying restrictions are valid.

## 5. Conclusions

The aim of this article has been to study the relationship between individual wages and local unemployment rates in Spain, considering a dynamic specification. The existing literature for Europe has shown that labour markets are better modelled by a wage curve representation in which wages are linked to the level of unemployment. Usually, the effect of unemployment on wages is persistent, so that some time is required to exert an inverse influence on wages; i.e. a dynamic wage curve is present.

We have estimated this dynamic wage curve for Spain, using individual data, coming from the eight waves of the ECHP for the period 1994 to 2001, using a specification that is common in empirical studies. We take into account both the reduced time dimension available and the subsequent bias arising from the estimation of a dynamic panel data model with fixed effects. Thus, we have used GMM estimators to test the degree of sluggishness in the response of wages to changes in unemployment rates.

Estimated results seem to reveal that, contrary to most earlier empirical research for other countries, a static wage curve models well the case of Spain, since the autoregressive parameter is non-significantly different from 0, and with the elasticity wages to unemployment of –0.07, close to the –0.1 «empirical law of economics» posited by Blanchflower and Oswald (1994, 2005). This can be interpreted as wages being low-degree sensitive to changes in unemployment, but wages adjusting very rapidly to such changes, at least during the period under consideration. Accordingly, supply shocks will impact wage bargaining and price/wage inflation, so that they will have permanent effects on the unemployment rate. This may be due to the high flexibility provided by temporary and immigrant workers to the Spanish labour market during the period. However, this result must be treated with caution, given that the period analysed, first, may be too short to capture the dynamic behaviour of wages and, second, coincides with an expansive phase of the Spanish economy, during which unemployment fell sharply, employment increased strongly, and real wage growth was controlled. The availability of longer data bases with individual information would allow us to obtain more robust conclusions.

Looking back in time, the advent of the Great Recession in 2008 has had a strong impact on the Spanish economy, with an explosion of unemployment rates, whereas real wages have grown only moderately (Bentolila *et al.,* 2012). During these last years, there has been a quantitative adjustment in the labour market, such that non-permanent workers and immigrant groups have borne most of the charge in such adjustment. Thus, the temporary rate has decreased from 35% in 2006 to 25% in 2011, representing almost two million fewer workers with temporary contracts in the period. Similarly, non-Spanish employees have been reduced by 500,000 between 2007 and 2012. When new information becomes available, it would be of great interest to repeat the present study, comparing results obtained under both scenarios: in booms and in recessions. An exercise as such will be helpful in understanding the adjustment mechanisms in the Spanish labour market.

## Appendix

The sample from the European Community Household Panel (ECHP) is made up of 17,908 individuals who were surveyed personally. The final size of the sample is reduced to 5,779 employees, forming an overall sample of 27,954 observations. Some individuals have been discarded: those who are not workers (including the self-employed), workers in agriculture and fishing, civil servants, and members of the military. The survey provides information on earnings, as well as job and personal characteristics. Specifically, the variables we have 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 are then deflated by the corresponding weighted regional CPI, which is own-elaborated at the NUTS 1 level from the NUTS 2 information 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.
— *Fixed-term contract*: Employed with a temporary contract = 1. Employed with a permanent contract = 0.
— *Immigrant*: Non-Spanish worker = 1. Spanish worker = 0.
— *Education level of the employee*: This includes 3 categories: primary or 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: manager, professional technician, supporting professional technician, administrative, simple services, qualified craftsman and technician, assembler, and non-qualified worker.
— *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 (NUTS I and NUTS II) disaggregation

| *Spain* | *NUTS I* | *NUTS II* |
|---|---|---|
| Region 1 | North West | Galicia, Asturias, Cantabria |
| Region 2 | North East | Basque Country, Navarre, La Rioja, Aragón |
| Region 3 | Community of Madrid | Community of Madrid |
| Region 4 | Center | Castilla-León, Castilla-La Mancha, Extremadura |
| Region 5 | East | Catalonia, Comunidad Valenciana, Balearic Islands |
| Region 6 | South | Andalusia, Murcia, Ceuta and Melilla |
| Region 7 | Canary Islands | Canary Islands |

## References

Albaek, K.; Asplund, R.; Blomskog, S.; Barth, E.; Gudmundsson, B.; Karlsson, V., and Madsen, E. (2000): «Dimensions of the wage-unemployment relationship in the Nordic countries: Wage flexibility without wage curves», *Research in Labor Economics*, 19, 345-381.

Ammermüller, A.; Lucifora, C.; Origo, F., and T Zwick (2010): «Wage flexibility in regional labour markets: Evidence from Italy and Germany», *Regional Studies,* 44 (4), 400-421.

Arellano, M., and Bond, S. (1991): «Some tests of specification for panel data: Monte-Carlo evidence and an application to employment equations», *Review of Economic Studies*, 58, 277-297.

Arellano, M., and Bover, O. (1995): «Another look at the instrumental variables estimation of error components models», *Journal of Econometrics*, 68, 29-51.

Baltagi, B.; Blien, U., and Wolf, K. (2009): «New evidence on the dynamic wage curve for Western Germany: 1980-2004», *Labour Economics*, 16, 47-51.

— (2012): «A dynamic spatial panel data approach to the German wage curve», *Economic Modelling*, 29(1), 12-21.

Bande, R., and Karanassou, M. (2009): «Labour market flexibility and regional unemployment rate dynamics: Spain 1980-1995» *Papers in Regional Science*, 88(1), 181-207.

Barth, E.; Bratsberg, B.; Naylor, R., and Raaum, O. (2002): *Explaining variations in wage curves: theory and evidence,* memorandum 03/2002, University of Oslo.

Bell, B. (1996): *Wage curve or Phillips curve?,* Nuffield College, University of Oxford. Mimeo.

Bell, B.; Nickell, S., and Quintini, G. (2002): «Wage equations, wage curves an all that», *Labour Economics*, 9, 341-360.

Bentolila, S.; Cahuc, P.; Dolado, J. J., and Le Barbanchon, T. (2010): «Unemployment and Temporary Jobs in the Crisis: Comparing France and Spain», *The Economic Journal 122,* F155-F187.

Bentolila, S.; Dolado, J. J., and Jimeno, J. F. (2008) «Does Immigration Affect the Phillips Curve? Some Evidence for Spain», *European Economic Review,* 52, 1398-1423.

Black, A., and Fitzroy, F. (2000): «Earnings curves and wage curves», *Scottish Journal of Political Economy*, 47(5), 471-486.

Blanchard, O. (2011): *Macroeconomics*, Updated 5th edition, Prentice Hall, Englewood Cliffs.

Blanchard, O., and Katz, L. (1997): «What do we know and we do not know about the natural rate of unemployment», *Journal of Economic Perspectives*, 11(1), 51-73.

— (1999): «Wage dynamics: Reconciling theory and evidence», *American Economic Review, AEA Papers and Proceedings*, 89(2), 69-74.

Blanchflower, D., and Oswald, A. (1994): *The Wage Curve*, Cambridge, MIT Press, MA.

— (2005): *The wage curve reloaded,* NBER WP 11338.

Blundell, R., and Bond, S. (1998): «Initial conditions and moment restrictions in dynamic panel data models», *Journal of Econometrics*, 87, 11-143.

Campbell, C. M. (2008): «An efficiency wage approach to reconciling the wage curve and the Phillips curve», *Labour Economics,* 15, 1388-1415.

Card, D. (1995): «The wage curve: A review», *Journal of Economic Literature*, 33, 785-799.

Deller, S. (2011): «Spatial Heterogeneity in the Wage Curve», *Economics Letters,* vol. 113 (3), 231-233.

Dolado, J. J., and Jimeno, J. F. (1997): «The causes of Spanish Unemployment: A Structural VAR Approach», *European Economic Review*, 41(7), 1281-1307.

Dyrstad, J., and Johansen, K. (2000): «Regional wage responses to unemployment and profitability: empirical evidence from Norwegian manufacturing industries», *Oxford Bulletin of Economics and Statistics*, 62(1), 101-117.

European Commission (2003): «Wage Flexibility and Wage Interdependence in EMU: Some Lessons from the Early Years», in *The EU Economy: 2003 Review*, pp. 153-200. DG ECFIN/391/03-EN, Brussels, November.

Galí, J., and López-Salido, D. (2001): «Una nueva curva Phillips para España», *Moneda y Crédito* (212, 265-310). Published in English as «A new Phillips curve for Spain» in *Empirical Studies of Structural Changes and Inflation. Bank for International Settlements Papers* (vol. 3, 174-203), Basel, Switzerland.

García, I., and Montuenga, V. (2003): «The Spanish wage curve: 1994-1996», *Regional Studies,* 37, 929-945.

Grubb, D. (1986): «Topics in the OECD Phillips curve», *Economic Journal*, 96, 55-79.

Iara, A., and Traistaru, I. (2004) «How flexible are wages in accession countries?», *Labour Economics*, 11(4), 431-450.

Jimeno, J. F., and Bentolila, S. (1998): «Regional Unemployment Persistence: Spain, 1976-1994», *Labour Economics*, 5(1), 25-42.

Kennedy, S., and Borland, J. (2000): «A wage curve for Australia?», *Oxford Economic Papers*, 52, 774-803.

Kiviet, J. (1995): «On bias, inconsistency and efficiency of various estimator in dynamic panel data models», *Journal of Econometrics*, 68, 53-78.

Livanos, I. (2010). «The wage-local unemployment relationship in a highly regulated labour market: Greece», *Regional Studies*, 44(4), 389-400.

Longhi, S.; Nijkamp, P., and Poot, J. (2006): «Spatial heterogeneity and the wage curve revisited», *Journal of Regional Science*, 46, 707-731.

Madsen, J. (2009): «The dynamics of income shares and the wage curve-Phillips curve controversy», *Scottish Journal of Political Economy*, 56(1), 45-72.

Montuenga-Gómez, V., and Ramos-Parreño, J. M. (2005): «Reconciling the wage curve and the Phillips curve», *Journal of Economic Surveys*, 19, 735-765.

Nickell, S. (1981): «Biases in dynamic models with fixed effects», *Econometrica*, 49, 1399-1416.

Nijkamp, P., and Poot, J. (2005). «The last word on the wage curve?», *Journal of Economic Surveys*, 19, 421-450.

Pannenberg, M., and Schwarze, J. (2000): «Wage dynamics and unemployment in Germany: Evidence from regional panel data», *Labour*, 14(4), 645-656.

Roodman, D. (2009): «How to do xtabond2: An introduction to difference and system GMM in Stata», *Stata Journal*, 9(1), 86-136.

Shapiro, C., and Stiglitz, J. (1984): «Equilibrium unemployment as a discipline device», *American Economic Review*, 74, 433-444.

Whelan, K. (2000): «Real wage dynamics and the Phillips curve», Federal Reserve Board Finance and Economic, Discussion Series Paper No. 2000-2.

Windmeijer, F. (2005): «A finite sample correction for the variance of linear efficient two-step GMM estimators», *Journal of Econometrics*, 126, 25-51.

Woodford, M. (1994): «Structural slumps», *Journal of Economic Literature*, 32 (4), 1784-1815.

## *Comment on «Wage Dynamics in Spain: Evidence From Individual Data (1994-2001)»,* by Inmaculada García-Mainar and Víctor M. Montuenga-Gómez

José Luis Roig *

The analysis of wage flexibility is a highly relevant issue both from the academic and policy perspectives. It is even more crucial for countries such as Spain that belong to a single currency area. How do wages adjust to unemployment variations in Spain? This is the question that «Wage dynamics in Spain: evidence from individual data (1994-2001)» addresses. To answer the question, a *dynamic wage curve* is estimated with data from the European Community Household Panel for the entire period covered by this panel: 1994-2001.

The idea behind the dynamic wage curve is that the Phillips curve that establishes an aggregated dynamic relation between unemployment levels and changes in wages in a competitive labour market framework and the wage curve that assumes a micro relationship between regional unemployment rates and wage levels in a non-competitive labour market framework only represent extreme cases. In the Phillips curve, full wage persistence is assumed, whereas the wage curve assumes exactly the opposite. In some countries, neither of these approaches may be able to properly model the adjustment of wages to unemployment. Instead, a mixed approach would be more effective. The dynamic wage curve extends the wage curve by introducing a lagged wage term as an explanatory variable. A non-significant coefficient of this variable would mean that the wage curve explains the relationship between unemployment and wages, whereas a coefficient that is close to one would mean that wage adjustment is explained by the Phillips curve. Intermediate coefficient values would support a partial persistence adjustment. Empirical evidence from the United Kingdom, Germany and other countries supports the hypothesis of partial persistence.

The article by García and Montuenga contributes to this literature by analysing the Spanish case. The contribution is interesting because it increases the available evidence on a non-settled topic and because the institutional context of the Spanish labour market is different from that of other European countries due to regulations that, according to many authors, generate rigidities to adjustment in the labour market and favour the creation of a sort of dual market in which some groups of workers bear the burden of differentially worse contractual conditions (temporary contracts) that allow flexibility at the margins of the market.

The econometric modelling is well developed and uses a GMM system to address endogeneity issues in a dynamic panel. The results show that the wage curve rules the wage dynamics in Spain. The authors obtain a coefficient of –0.07 for the

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unemployment rate. This is not far from the coefficient of –0.1 that constitutes the «empirical law of economics» of Blanchflower and Oswald. Most importantly, the lagged wage term is not significant in the preferred estimations and is very low in the other presented estimations. In the words of the authors, «this result can be interpreted as that the Spanish labour market is relatively more sensitive to supply shocks than in other countries». The given explanation for this sensitivity is the presence of groups with lower bargaining power in the labour market. The authors note two of these groups: workers in temporary contract jobs and immigrants. The results of segmented estimations for these groups of workers (temporary *vs.* permanent contract; native *vs.* immigrant) tend to corroborate this hypothesis. The degree of flexibility and non-persistence of temporary and immigrant workers is larger than that of their permanent and native counterparts.

In my view, the main weakness of the paper lies in the short time period available for the estimations. Eight years may not be enough to fully capture the dynamics of the market, and this limitation is rightly acknowledged by the authors. There is also a second question that should be taken into account. The period of analysis runs from 1994 to 2001, whereas immigration in Spain began to reach significant levels only after 2001 or later. Thus, the importance of the immigrant workers in the market during the period of analysis is likely to be very limited [15].

All in all, this is a nice piece of work on a disputed topic and is extremely relevant given the problems that the Spanish labour market is currently facing. The results will need to be corroborated when longer and more recent panel data become available.

[15] Nonetheless, it could be interesting to estimate the model with the panel data from the EU-SILC available for 2004 to 2010. The number of observations is now very similar to that of the ECHP used in this article, but the cyclical characteristics of this period are richer and immigrant workers account for a significant share of the market.