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The Wage Curve: the Lessons of an Estimation Over a Panel of Countries

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ABSTRACT

This paper gives a systematic comparison of private wage behavior in industrialized countries. A wage curve with nominal rigidities is estimated on a panel of 16 countries according to a specific method based on GMM and factor analysis. First, we show that the employment rate is a better indicator of labor market tensions than the unemployment rate. Second, the main difference across countries is the reaction of the wage rate to changes in productivity and in the employment rate. Third, we find evidence of some nominal rigidity in wage behavior and of a positive but small effect of the wedge.

Keywords: International comparison, labor market, wage flexibility.

JEL classification: C33, J30.

RÉSUMÉ

Les marchés du travail des pays des pays industrialisés ont des caractéristiques très différentes en termes de réglementation, de rôle et de poids des partenaires sociaux, d'indemnisation du chômage, etc. Dans quelle mesure ces disparités s'accompagnent-elles, au niveau macro-économique, de comportements profondément différents d'un pays à l'autre ? Cette question est importante dans le cadre de l'UEM. Une crainte est en effet que des différences de comportement sur le marchés du travail ne donnent des effets asymétriques à un choc initialement commun. Par ailleurs, la mesure de la capacité des marchés du travail européens à s'ajuster face à un choc est importante. Beaucoup d'économistes estiment en effet que l'intégration monétaire européenne pourrait connaître des difficultés sérieuses en présence d'une trop grande rigidité du marché du travail dans certains ou tous les pays de l'UEM, par exemple à la suite d'un choc asymétrique puissant.

Ainsi, la comparaison de la formation des salaires dans les différents pays européens, aux Etats-Unis et au Japon peut être riche d'enseignements pour le fonctionnement de l'UEM. Nous proposons ici l'estimation d'une équation des salaire sur un panel de 16 pays industrialisés, comprenant tous les pays participant à l'Euro. L'intérêt de recourir à un panel de pays est double. D'abord, dans la mesure où ces pays présentent certaines similarités structurelles, l'estimation du comportement de chaque nation bénéficie de l'information qu'apportent ses 15 partenaires, et nous pouvons obtenir ainsi des résultats plus solides et plus précis. Ensuite, cela nous permet d'identifier des différences structurelles robustes dans la formation des salaires entre pays.

Dans une première section, nous présentons une modélisation théorique pouvant donner lieu à une estimation économétrique simple de la formation des salaires. Celle-ci repose sur une *wage curve* où le coût du travail est une fonction de sa productivité, des prix, du coin salarial et du taux de chômage. Nous complétons cette équation en introduisant un élément de rigidité nominale : certains contrats de salaire fixent celui-ci pour plus d'une année, et sont donc fonctions du prix courant, mais aussi du prix futur anticipé.

La seconde section présente la méthodologie économétrique. L'équation de salaire est estimée simultanément sur les 16 pays, en supposant que certains paramètres ont une valeur commune dans plusieurs nations. La méthode d'estimation la plus naturelle est celle des moindres carrés généralisés. Mais, la présence de variables endogènes et anticipées nous conduit à recourir à des variables instrumentales et aux moments généralisés (GMM). Une nouvelle difficulté apparaît alors : la matrice de covariance des chocs heurtant les différents pays est de grande taille et estimée sur relativement peu de périodes. Pour limiter l'imprécision qui résulterait d'une estimation directe, nous introduisons un certain degré de structure dans les covariances entre chocs en supposant qu'elles peuvent être modélisées par un petit nombre de facteurs communs. Un autre aspect économétrique important est la stratégie de tests emboîtés, allant du général au particulier, qui permet de déterminer si les paramètres prennent ou non des valeurs différentes dans les pays.

La troisième section présente les résultats. Tous d'abord, le taux d'emploi est un bien meilleur indicateur des tensions sur le marché du travail que le taux de chômage. Ensuite les principales différences entre pays tiennent à la réponse des salaires à l'évolution de la productivité et aux tensions sur le marché du travail. Toutefois, si l'élasticité du coût du travail à l'emploi varie assez nettement entre les pays, dans les pays européens elle reste globalement de l'ordre des résultats de Blanchflower et Oswald (1995) pour lesquels l'élasticité du salaire au taux de chômage est d'environ 0,1. En revanche, le taux d'emploi

n'a pas d'impact significatif sur l'évolution du salaire aux Etats-Unis, où les ajustements s'effectuent davantage par des flux de mains d'œuvre entre Etats que par une modification des salaires. Les autres paramètres de l'équation ont des valeurs communes dans tous les pays de l'échantillon. On remarque alors que la durée des contrats de salaire est relativement longue, ce qui implique une certaine rigidité nominale, et les anticipations de prix sont assez statiques. Le coin salarial a un effet positif mais faible sur le coût du travail : une hausse des cotisations sociales est donc principalement supportée par une baisse de salaire touché par le salarié, conformément aux résultats de Cotis et Loufir (1990).

SUMMARY

This paper offers an estimation of private wage behavior on a panel of 16 industrialized countries, including all the EMU countries. Using a panel estimation helps us to get more robust and precise empirical findings: as these countries share some common structural features, each country estimation benefits from information brought by its 15 partners. Second, panel estimation allows us to identify deep structural differences between countries. This kind of analysis is particularly important as industrialized countries' labor markets display great heterogeneity concerning wage bargaining processes, degrees of job protection, and provision of replacement incomes, etc.

The first section proposes a simple formalization of wages setting, based on a wage curve in which the labor cost depends on labor productivity, prices, the wedge between real labor cost for firms and the purchasing power of nominal wages for wage earners, and the unemployment rate. We also introduce nominal rigidities in this equation: some wage contracts are longer than one year and depend not only on current prices but also on anticipated ones.

Section 2 is dedicated to the econometric method. The wage equation is estimated simultaneously for the 16 countries, assuming that some parameters have the same value in several countries. The presence of anticipated variables requires the use of instrumental variable or GMM methods, instead of traditional generalized least squares. This raises a new problem: the covariance matrix of the shocks hitting the countries has a large dimension and is estimated on a rather short time period. To improve the precision of the estimation, we assume that the shocks hitting the countries at the same time can be represented by a limited number of common factors. Another important aspect of our approach is the nested test strategy, from the less constrained model to the most one, designed to evaluate if each parameter is country specific or not.

The third section gives the results. We show that wage contacts are fairly long (which implies some nominal rigidity) and that price expectations are quite static. The wedge has a positive but small effect on wages: an increase in social contribution mainly results in smaller earnings for the workers. The employment rate is a better indicator of labor market tensions than the unemployment rate. The elasticity of the wage cost to the employment rate clearly differs across countries. However our results are not inconsistent with Blanchflower and Oswald (1995) findings of an elasticity of the wages to the unemployment rate of around 0.1. Lastly, in some countries, the wage behavior is clearly at odds with our specification, because of very specific labor market institutions (Spain) or because the country has experienced a strong economic shock (Finland).

The Wage Curve: The Lessons of an Estimation over a Panel of Countries

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INTRODUCTION

This paper offers an estimation of private wage behavior on a panel of 16 industrialized countries, including all the EMU countries. We are first interested in the estimation of the ability of wage adjustments to constitute a mechanism for macroeconomic stabilization. This question is important for the EMU countries that have lost national autonomy over monetary policy. The theory of optimum currency areas suggests that wage flexibility could be a way to cope with the asymmetric shocks that these countries are likely to face (for example, because of their different specialization). Using a panel estimation helps us to get more robust and precise empirical findings: as these countries share some common structural features, each country estimation benefits from information brought by its 15 partners. Second, panel estimation allows us to identify deep structural differences between countries. This kind of analysis is particularly important as industrialized countries' labor markets display great heterogeneity concerning wage bargaining processes, degrees of job protection, and provision of replacement incomes, etc. (See OECD (1994), Cadiou and Guichard (1999)). Therefore, labor markets are likely to lead an initially symmetric shock to have asymmetric consequences; this is also an important source of concern for the EMU.

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Section 2 is dedicated to the econometric method. The wage equation is estimated simultaneously for the 16 countries, assuming that some parameters have the same value in several countries. The presence of anticipated variables requires the use of instrumental variable or GMM methods; instead of traditional generalized least squares. This raises a new problem: the covariance matrix of the shocks hitting the countries has a large

[a] CEPII and IMF.
[#] CEPREMAP and TEAM.
The authors thank a lot Agnès Bénassy, Philippe Jolivaldt and Sevestre for their remarks and suggestions. The paper was presented at seminars given at the University of Paris X Nanterre, at Humbold Universität and at a workshop organised by ENEPRI. The discussions which followed our presentation were extremely useful for us. The remaining imperfections of this paper are of their responsibility.

dimension and is estimated on a rather short time period. To improve the precision of the estimation, we assume that the shocks hitting the countries at the same time can be represented by a limited number of common factors. Another important aspect of our approach is the nested test strategy, from the less constrained model to the most one, designed to evaluate if each parameter is country specific or not.

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I - THE THEORETICAL BACKGROUND

Each year, firms and workers are assumed to agree on nominal wage contracts. Some of them expire during the current year (their length is shorter than one year); others run out the following year (their length is shorter than two years[1]). Each contract is identified by two indices: *t* the year it is concluded and *i* which is equal to 1 if the contract expires during the current year, 2 if it expires during the following year. The theoretical models of wage setting (bargaining models, search models, efficient wage models) justify the following equation[2]:

$$Log(WC_{t,i} / PE_{t,i}) = w_0 + w_1 Log(Y_t / E_t) + w_2 Log(WDG_t) + w_3 Log(UN_t) + \varepsilon_{t,i} \quad (1)$$

WC is the private wage cost set by the contract. PE is the average expected price level during the contract. Y is the private output, E represents private employment, UN is the unemployment rate[3] (in percent) and WDG is the wedge. The error term e includes many other variables influencing the bargaining process, that are very likely to play an important role. They include the bargaining power of unions, the market power of firms, social laws that modify the power of insiders (for instance the layoff rules), leisure utility, etc. Some economists believe that the level and the duration of employment benefits are important determinants of wages through their effects on reservation wages. Other economists think that the degree of centralization of wages bargaining matters a lot. We were not able to find reliable indicators of these complex multi dimensional variables for the 17 countries which we investigated. These unobservable variables probably exhibit great persistence.

[1] However these contracts might also be quite short, for instance a contract concluded at the end of year t and expiring at the beginning of year t+1.

[2] See, for instance, Layard, Nickel and Jackman (1991), Blanchflower and Oswald (1994), Pissarides (1998).

[3] As shown later, the unemployment rate is a poor indicator of labour market desiquilibrium. We replace it with the employment rate.

We consider two expected prices. The first is associated with contacts running out during the current year and is the current price of private output.

$$Log(PE_{t,1}) = Log(P_t)$$

The second is associated with contracts that will expire during the following year; it is a weighted average of the current price and the price expected for the following year.

$$Log(PE_{t,2}) = qLog(P_t) + (1-q)Log({}_tP_{t+1}^a) \qquad (2)$$

Let call p the proportion of wage contracts expiring during the year they have been concluded and W_t the average wage cost during this year. We then have the identity:

$$Log(W_t) = [(1-p)WC_{t-1,2} + pLog(WC_{t,1}) + (1-p)Log(WC_{t,2})]/(2-p) \qquad (3)$$

As WC and PE are not observable, we should eliminate them from the equation. Using (1) to (3), we get:

$$\begin{aligned} &Log(W_t) = \\ &\{ (1-p)[qLog(P_{t-1}) + (1-q)Log({}_{t-1}P_t^a) + w_0 + w_1Log(Y_{t-1}/E_{t-1}) + w_2Log(WDG_{t-1}) + w_3Log(UN_{t-1})) + \varepsilon_{t-1,2}] \\ &+ p[Log(P_t) + w_0 + w_1Log(Y_t/E_t) + w_2Log(WDG_t) + w_3Log(UN_t) + \varepsilon_{t,1}] \\ &+ (1-p)[qLog(P_t) + (1-q)Log({}_tP_{t+1}^a) + w_0 + w_1Log(Y_t/E_t) + w_2Log(WDG_t) \\ &+ w_3Log(UN_t) + \varepsilon_{t,2}] \}/(2-p) \end{aligned} \qquad (5)$$

With $(1-p)/(2-p) = r$ and $p/(2-p) = 1-2r$; with $\eta_{t-1,t}$ the forecast error made at t-1 for price level in t ($\eta_{t-1,t} = Log(P_t) - Log({}_{t-1}P_t^a)$) that, under the assumption of rational expectation, follows a martingale difference. Then (5) becomes:

$$\begin{aligned} &Log(W_t) = w_0 + w_1[rLog(Y_{t-1}/E_{t-1}) + (1-r)Log(Y_t/E_t)] \\ &+ w_2[rLog(WDG_{t-1}) + (1-r)Log(WDG_t)] + w_3[rLog(UN_t) + (1-r)\log(UN_{t-1})] \\ &+ rqLog(P_{t-1}) + (1-r)Log(P_t) + r(1-q)Log(P_{t+1}) \\ &+ r(1-q)(\eta_{t-1,t} + \eta_{t,t+1}) + r\varepsilon_{t-1,1} + (1-2r)\varepsilon_{t,1} + r\varepsilon_{t,2} \end{aligned} \qquad (6)$$

Hence, the nominal wage cost in t depends on a weighted average (with the same weights) of productivity, wedge and employment rate in *t-1* and *t*. It also depends on prices of years *t-1* and *t*, but the relation is more complicated[4].

We mentioned previously that it is likely that the error terms $\varepsilon_{t,1}$ and $\varepsilon_{t,2}$ are highly persistent. So, we assume that they can be represented by an AR (1) process of parameter ρ. We tried to estimate this parameter and always got a value very close to 1. Therefore, we decided to assume that the error term follows an integrated process of order 1. That means that the unobservable variables are I (1), which is quite a reasonable assumption. This implies that the observable variables of equation (6) are not cointegrated**.** However,

[4] The value for p and q are not set *a priori* but are estimated.

the values of the elasticities in this equation are of great interest for the economist. We then estimate the first difference of (6), which also allows us to eliminate a constant that is very likely to vary across countries. Indeed, the national unemployment rate may also depend on some unobserved specific features such as a low bargaining power of long-term unemployed persons[5].

II - THE ECONOMETRIC METHOD

Our aim is to estimate the first difference of (6) on a panel of countries. The specification of this wage equation is the same for all countries, but the values of the parameters may differ. Panel estimations bring more information on behaviors than time series estimations[6]. However, this kind of estimation raises some problems.

The data

The sample of 16 industrialized countries includes Germany, Austria, Belgium, Canada, Spain, the United States, Finland, France, Greece, Italy, Ireland, Japan, the Netherlands, Portugal, the United Kingdom, Sweden[7]. The data are annual, taken from OECD Economic Outlook and expressed in logarithms.

The estimation period goes from 1982 to 1997 (the same for each country). The data are available on a longer period, but we decided to start in 1982 to prevent bias due to structural breaks in the equations: European countries implemented numerous reforms in the late 70's and at the beginning of the 80's. We consider that the cross-country dimension compensates for this relatively short period. Of course, economic shocks and labor market reforms also occurred during the 80's and 90's, but it seemed unwise to shorten the period more. Dummies were added to deal with the German reunification, as the data are available for the whole period neither for the Western part nor for the whole Germany. This is equivalent to excluding two years (1991 and 1992) out of the German equation.

WDG, the wedge between real labor cost for firms and the purchasing power of nominal wages for the wage earners is defined as (pc/(p*(1-sscr)*(1-tlr))*(1+vatr90)). pc is the CPI, sscr the social security contribution rate (for both employer and employees), tlr the tax rate on labour income, vatr90 the VAT rate for the base year. Morever, as the unemployment rate is a debated indicator of the tensions on the labor market we also used the employment rate (employment/population).

[5] In the new equation, the error term is an ARMA of probably low order. As said in the second section we could not retain this in the estimation, but the tests suggest this is not an important problem.

[6] When studying one country with quarterly data the number of observations is four times larger, but, even if it brings information on short term dynamics, it does not help much for the estimation of long term parameters, that are of greater economic interest (in our case the long run elasticity of wages to unemployment or wedge). Moreover, the quarterly data are often seasonally adjusted; that deteriorate the informative contend of the series and is a source of bias in the estimations ant the related tests. (see Hendry (1995, p. 559-565)).

[7] Two countries of the EU are missing (Denmark and Luxembourg) because the data were not available. The estimation method, defined in section 2, has been implemented on TSP 4.4. The program, with detailed comments, is available upon request.

On the whole, our estimations concern 16 countries and 15 years.

The problems

Our approach poses some problems that are not exactly the same as for traditional panel techniques and also that differ from time series econometrics[8]. We have a longer time period than traditional panel estimations, a smaller sample of individuals, a special interest for the comparison of parameters across individuals, and few constraints on the covariance matrix of error terms. The non-linearity of the equation and the presence of expected variables are additional difficulties.

Our problem was to estimate on a panel of I countries, indexed by i, and on a period of T years, indexed by t, the following system of I equations:

$$f_i(y_{it}, y_{i,t-1}, x^a_{1i,t+1}, x_{2it}, x_{3it}; \alpha_i) = \varepsilon_{it}; i = 1,.., I; t = 2,.., T-1 \tag{7}$$

I and T are of the same order of magnitude, and not very high. The y_{it} are the endogenous variables, the x_{jit} are the exogenous variables, f_i is a function representing the behavior associated to country i, the α_i are the parameters of this function. Some of these parameters are country specific; the others are common to some countries or to all of them. ε_{it} is the error term of null expected value[9].

Our answers

We assume that the error terms of a common year are correlated, and call Ω, of typical element ω_{ij}, their covariance matrix. This assumption is consistent with an interpretation of error terms as correlated random shocks affecting the different domestic economies. As explained below we also put some structure in these correlations to increase the number of degree of freedom of our estimation. However, the structure usually proposed by the error component models seems too restrictive and too poor for our needs.

On the other hand, we assume the error terms to be time independent. This assumption is questionable, and a more reasonable one would be that the error terms follow a stationary process (possibly after differentiating the equation). In that case the estimation would be quite easy if this process was auto regressive (see Kmenta (1986, 1997)). However, we find it better to take this eventuality into account by adding lagged endogenous and exogenous variables. This lowers the number of degrees of freedom, but avoids the introduction of *ad*

[8] See for instance the very good manual of Baltagi (1995).

[9] Our problem is different enough from the one of Pesaran and Smith (1995). We directly estimate the various values that a given parameter can take in the various countries without supplementary assumptions. For Pesaran and Smith, the differences between the values are random, and they estimate the expected value and the variance of each coefficient over the set of all countries. They could also compute in their analytic frame the optimal forecast of the values that the coefficients take in the various countries. Thus, if we use the terminology of the econometrics of panels, our approach is similar to that of the models with fixed-effects, and the approach of Pesaran and Smith is similar to the one of component errors models. This last approach gives more precise estimations, if the stronger assumptions it requires are valid. A very interesting extension of the approach followed by Pesaran and Smith, with an application, is given by Li, Maddala and Trost (1996).

hoc unjustified constraints in the dynamics of the error terms[10]. However, this is not a solution if the error terms have a moving average component, as it is the case for our theoretical model[11]. We were not able to deal with this issue, but the tests presented below indicated that this inability does not have important drawbacks.

Let us begin by assuming that all the explanatory variables are predetermined, i.e. that the e_{it} are independent of the contemporaneous and past values of the explanatory variables, and that the functions f_i have the property that system (7) can be rewritten:

$$y_{it} = g_i(y_{i,t-1}, x^a_{1i,t+1}, x_{2it}, x_{3it}; \alpha_i) + \varepsilon_{it}; i = 1,.., I; t = 2,.., T-1 \tag{8}$$

In this case system (8) can be easily estimated by generalized nonlinear least squares. However, we prefer to make more general assumptions. Thus, we assume that variable x_{2it} is predetermined, but that this property is not shared by variable x_{3it}. The endogeneity of x_{3it} prevents the nonlinear least square estimators of the parameters of being consistent[12]. Moreover, variable $x^a_{1i,t+1}$ represents the forecast at time t of variable x_{1i} for time $t+1$. As this variable is not observed, we follow a suggestion by Wickens (1981), and substitute it by its observed value at time $t+1$: $x_{1i,t+1}$. Thus, we introduce a supplementary error in the equation, which bears on the foreseen value of an explanatory variable for a future time[13]. We have then to estimate a model with errors on variables, and in this case least squares estimators are non-consistent.

To overcome these difficulties we allocate to each national equation a set of instrumental variables. For instance, if the anticipated variable is predetermined, natural instruments for the equation of country i are $y_{i,t-2}$, x_{1it}, x_{2it} and $x_{3i,t-1}$. In the rest of the paper we will assume that there are n instruments per country. Then, we will use the two step GMM method which is presented in Appendix 1.

[10] Our personal experience with tests of common root (COMFAC) is that they conclude in almost all cases to the rejection of the null hypothesis.

[11] The approximation of mean average by the introduction of more lagged variables in the equation would only be valid if we could introduce a very high number of lags, what would make the estimation impossible or imprecise. The GMM, which we present in Appendix 1, usually considers this case by substituting in the definition of Φ the empirical covariance of vector V_t by a weighted average of its empirical autocovariances for the nearest lags and leads. Th difficulty with this approach is we must estimate the autocovariances of the errors of the system of equations (7), that is several matrices of dimension ( ) on a sample of size $T-2$. As T is of the same order as I, these estimations can be only very imprecise, and the GMM method would then appear as little reliable. In Appendix 2, we develop a practical method based on factor analysis to overcome this difficulty for the covariance of errors terms. But we have not found any mean to extend this method to the autocovariances.

[12] For instance in the model of first section, productivity increases cannot be considered as exogenous, : Germany was a counter example where higher wages led to productivity increases.

[13] And then which follows a difference of martingale, independent of the explanatory and explained variables at and before time.

Another difficulty is that the estimation of the covariance matrix of the system of equations, Ω, is very imprecise: $\hat{e}_{it}$ is observed for $t = 2,..,T-1$, which makes $T-2$ observations. Yet I is of the order of $T-2$. Thus this matrix is almost singular, or even singular if the number of observed years is smaller than the number of countries. We use factor analysis to put some structure in this matrix; that is some interdependence between the shocks hitting the countries in a way that would appear natural to economists. Doz (1998, page 85-161) gives an introduction of factor analysis which is at the same time simple and rigorous, and we base on it here (Appendix 2).

We could complete this factor analysis of the covariance matrix, by an estimation of the factors, which would raise a problem of identification. We could also look for an economic interpretation of these factors. However, this is out of the subject of this paper.

The principles of the tests

There are two kinds of important tests that are easy to implement:

a) *The over identifying restriction test of Hansen:*

If the model and the instruments are valid, the objective function of the second step of the GMM $(\iota'\hat{V}\Phi^{-1}\hat{V}'\iota)/(T-2)$, where the hat identifies estimated values, follows a χ^2 with s degrees of freedom[14].

b) *Tests of restrictions on parameters, especially of equality of parameters between countries*:

A ^ identifies the results of the estimation of the non-constrained model. A ˜ identifies the results of the estimation of the model constrained by r equalities between parameters. Then, the statistics of the likelihood ratio $LR = (\iota'\tilde{V}\tilde{\Phi}^{-1}\tilde{V}'\iota - \iota'\hat{V}\tilde{\Phi}^{-1}\hat{V}'\iota)/(T-2)$ follows a c^2 with r degrees of freedom[15]. This expression uses twice the estimation of Φ got at the last step of the GMM estimation of the constrained model. Then, the estimation of the non-constrained model is very simple and is made in one step[16].

To test for the equality or the difference of the values of each parameter between countries, we choose to progress from general to specific. In a first series of null hypotheses, we assume that all the coefficients of the model differ between countries, except one. Then we

[14] s is the difference between the numbers of instruments and parameters (see Appendix 1).

[15] Another solution would have been to use a Lagrange test: $LM = \iota'V\Phi^{-1}\Delta(\Delta'\Phi^{-1}\Delta)^{-1}\Delta'\Phi^{-1}V'\iota/(T-2)$, where Δ is the matrix of the partial derivatives of $V'\iota/(T-2)$ relatively to the parameters of the unconstrained system of equations. Under the null hypothesis, LM follows a χ^2 with r degrees of freedom, where r is the number of constraints tested on parameters. However, a test of the likelihood ratio is a little easier to implement.

[16] In the case where the non-constrained model is just identified, the statistics of this test is numerically equal to the statistics of the over identification test of Hansen (for a common value of Φ of course).

continue with two coefficients common to all countries, etc. until the tests induces us to stop[17].

We give the example of our strategy in the case of an equation including three parameters to estimate: a_i, b_i and c_i. We test hypotheses relative to the equality of the values taken by one of these parameters over the total of all countries (which can be denoted, for instance: $c_i = c_j$), against the alternative that this parameters takes values which differ among countries ($c_i \neq c_j$). To retain the null hypothesis, we require that the likelihood ratio test again the alternative hypothesis, and the Hansen tests under the null hypothesis, have both p-values larger than 5%. Our strategy of nested tests is given in the following diagrams. At each step, a null hypothesis is accepted if it is not rejected against any of the associated alternative hypothesis. If, for a given alternative hypothesis, one of the associated null hypotheses is not rejected, this alternative hypothesis is rejected. This second criteria can be criticized. Indeed, an alternative hypothesis H1 can fail to reject the null hypothesis H0, but H0 is rejected against another alternative hypothesis H1'. By rejecting H1, we implicitly assume that this hypothesis includes wrong features, which do not appear in H1'. At the end of the series of nested tests, it may be possible to retain several configurations of equalities and differences of parameters between countries. Then, we could try to choose between them by using non-nested tests of the J kind (see Davidson and MacKinnon (1993), chapter 11), or by economic arguments.

First step

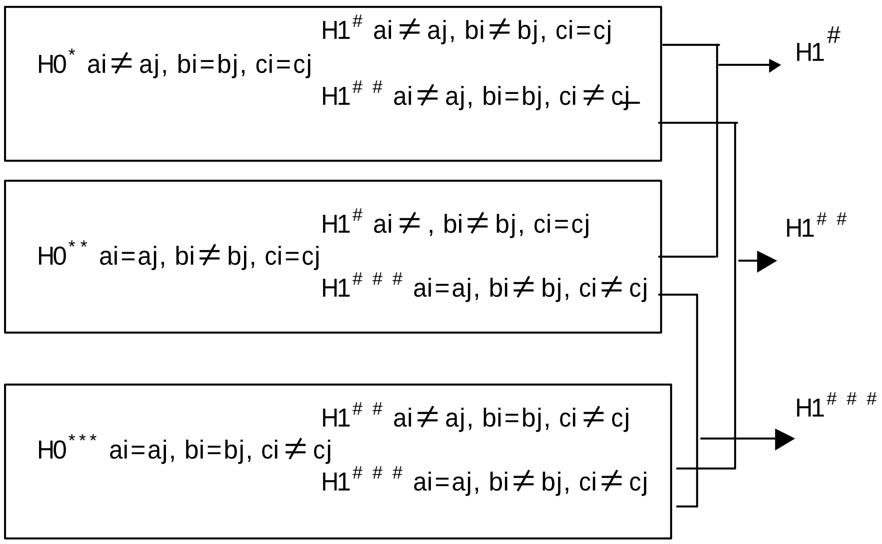
The alternative hypothesis H1 is that the three parameters differ between countries. The three null hypotheses $H0^{\#}$, $H0^{\#\#}$ and $H0^{\#\#\#}$ are that one of these parameters takes the same value across countries. If the tests of these null hypotheses reject the three null hypotheses, we retain the alternative assumption. Otherwise, we go to the second step.

$$
\left.
\begin{array}{ll}
H0^{\#}\ ai \neq aj,\ bi \neq bj,\ ci = cj & H1\ ai \neq aj,\ bi \neq bj,\ ci \neq cj \xrightarrow{H1} \\
H0^{\#\#}\ ai \neq aj,\ bi = bj,\ ci \neq cj & H1\ ai \neq aj,\ bi \neq bj,\ ci \neq cj \xrightarrow{H1} \\
H0^{\#\#\#}\ ai = aj,\ bi \neq bj,\ ci \neq cj & H1\ ai \neq aj,\ bi \neq bj,\ ci \neq cj \xrightarrow{H1}
\end{array}
\right\}
\begin{array}{l}
H1 \\ ai \neq aj, \\ bi \neq bj, \\ ci \neq cj
\end{array}
$$

[17] Turner and Seghezza (1999) answer in a slightly different way to this problem. First, the equations are estimated independently for each country, with systematic break tests. Then dummy variables, related to these breaks are introduced in the equations. Finally, these equations are simultaneously estimated on the set of all countries, and equality tests of the various parameters between countries are systematically made. To be more precise let us consider a given parameter. The authors start by assuming that it takes different values for each country and test the equality of this value for the two countries the nearest, with a Wald test. Then, if they accept this equality they reestimate the model under this constraint, and they test the equality to this common value of both countries to the nearest value got for the other countries, etc.

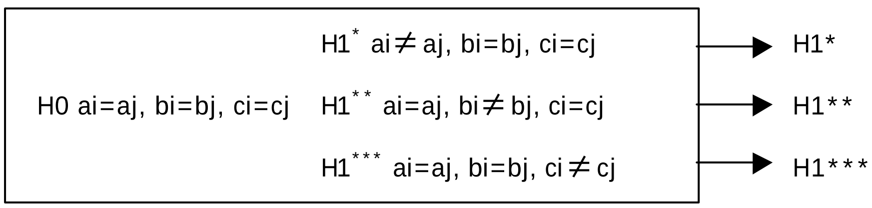
Second step

In the following diagram we assume that each of the three null hypotheses of step 1 has not been rejected, and now represents as many alternative hypotheses denoted by H1#, H1## et H1### (otherwise, some of these alternative hypotheses should not appear in the diagram). Each alternative hypothesis assumes that two parameters take different values across countries and is associated with two null hypotheses where only one of these parameters changes across countries. The total number of possible null hypotheses is equal to three: H0*, H0** and H0***. Each of them is associated with two alternative hypotheses.



Third step

The diagram corresponds to the case where the three null hypotheses of the second step have not been rejected. Then, they become as many alternative assumptions denoted H1*, H1** and H1***. Now, we have only one null hypothesis, H0 where the three parameters take common values across countries. If H0 is rejected against one or several alternative hypotheses, we retain this (these) last hypothesis. Otherwise, we retain the null hypothesis H0.



This series of nested tests is used to establish if the values of the various parameters change or not with countries. To do that in a rigorous way we must keep the same instruments for all the tests. The instruments we choose are the constant term, and all the explanatory and explained variables with a lag of three periods. Indeed, the explanatory variables, and so indirectly the explained variables, appear in the equations with a lag of two periods. Then, we have retained six instruments per country, so a total of: 1+5*16=81 instruments. Taking a higher number of instruments would increase the number of degrees of freedom, but would raise the risk of small sample biases. A supplementary problem is that it would result in bad statistics for the over identification test of Hansen.

III - RESULTS

The implementation of the test strategy was constrained by difficulties we met in the convergence of the GMM algorithm when all the parameters, and all but one, were country-specific. This was probably due to the low numbers of degrees of freedom in these cases. Therefore, even if the algorithms had converged, the results would have been too fragile to be considered. We started the process of nested tests by assuming that 2 parameters (out of 5) are country specific under the null hypothesis, and three under the alternative hypothesis. As, in the end, the best model is a model with a small number of parameters varying across countries, this drawback should not be really embarrassing. After having run the series of nested tests, we reached the conclusion that the best model retains identical parameters across all countries[18].

The estimation results are presented in table 1. They are disappointing. First, q is greater than 1 (even if we cannot reject its equality to 1). Moreover, the unemployment rate is not significant. Therefore, we reestimated the model, using the employment rate instead of the unemployment rate. It is very common to consider that the former (computed over people of age between 15 and 64) is a better indicator of the tensions on the labor market. In particular, it takes into account discouraged unemployed people who retire from the labor market, and hence from unemployment statistics, but would work if a position were offered to them. The share of these people in the working age population is likely to fluctuate with the economic cycle and to vary across countries. This indicator also takes into account other categories of people who leave unemployment without finding a job (because of training programs for instance), or young people who keep studying because they cannot find a job, or people who leave their job without becoming unemployed (through a pre retirement scheme).

[18] To limit the bias sue to our inability to estimate some model, we tested the null hypothesis of this model was tested against the alternative hypothesis of one country-specific parameter in only one country. The null hypothesis was never rejected.

Table 1: The results with the unemployment rate

	Coeff	Student
w1	0.60	33
r	0.28	3.0
w2	0.32	8.1
w3	-0.004	-0.9
q	1.1	4.5

p-value of Hansen test 30%

In that case, the best model was a model where both the productivity and employment rate parameters are country specific (Table 2). This model was not rejected against any model with one more country-specific parameter, with p-values ranging from 5 à 20%[19]. The models where only one parameter is country specific were rejected against the retained model, with p-values of the order of 1%.

On the one hand, the common parameters have the expected signs and reasonable values. r is equal to 0.4, meaning that only 25% of the wages contracts expire the year they have been concluded. The expected price only plays a small role in the price used as reference for the longest contracts (30% against 70% for current prices). The elasticity of wage cost to the wedge (w_2) is of 0.16, implying that an increase in social security contributions (of employers or employees) results in a small increase in wages costs and is borne by employees (whose net earnings fall). This result confirms the conclusion by Cotis and Loufir (1990) in the case of France. We have not separated in the wedge the respective contributions of direct taxes, indirect taxes and the terms of trade. This is theoretically legitimate when we consider a single representative worker, but would be less convincing if we separated workers, for instance by skills or by the size or the sector of the firms they are working in. Moreover, our equation only considers the average rate of the wedge. The marginal rate should have a significant and different effect, but we were unable to compute this rate.

On the other hand, productivity has a positive and significant impact on wages in all the countries under review except Greece and Spain (where the impact is negative, but not significant), and Portugal (where it is positive but not significant). Among the countries of the Euro area productivity shocks have weak effects on wages costs in France, in Italy and in the Netherlands, stronger effects in Germany, Finland and especially in Austria. The US and Japan do not significantly differ from the average of the Euro area. The effects of productivity shocks are especially strong in the UK.

The employment rate has no significant impact on wages in Belgium, Canada, Sweden, Ireland and the United States. It has a significant positive effect on wages in Italy, the United Kingdom, the Netherlands, Austria, Greece and Germany, and not very significant

[19] Then the null hypothesis of this model was tested against the alternative hypothesis of one more country-specific parameter in only one country. The null hypothesis was rejected in only one case: the coefficient q in the Netherlands, with a P-value of 0.6%. However because of the high value of the parameter we neglect this result.

in France. The traditional result of a high wage flexibility in Italy is confirmed[20]. However, our findings on the UK differ from most studies that conclude that wage flexibility in this country is quite low. In the same way Germany appears here as one of the most flexible countries in the Euro area when it is generally considered as in an intermediary position. The very high flexibility we get for Greece is likely to hide a specification problem in the equation.

Usually, wage flexibility is estimated regarding the unemployment rate and not to the employment rate. Everything being equal, this elasticity is $-w_{3i}\bar{u}_i$, with $\bar{u}_i$ the average unemployment rate for country i. in the case we use the employment rate, we can apply the same formula under the strong assumption that the participation rate is stable. We give these elasticities in the third column of Table 2. Their values are consistent with the conclusion of Blanchflower and Oswald (1995) on individual data sets (that is –0.1 in most countries)[21].

The employment rate has no significant effect on the US. In this country the flow of labor between states are important and very sensitive to local conditions on the labor market (Blanchard and Katz (1992)). Thus, the average unemployment rate in this country is an indicator of little significance, with little effects on average wages (Thomas (1994))[22]. It is interesting to notice that the two more liberal labor markets (the US and the UK) present opposite levels of flexibility. On the other hand the labor markets of Germany and Italy, which are usually assumed to be less liberal, present a high degree of flexibility, similar to the one in the UK.

In Finland, and to a larger extend in Spain, this effect is negative and significant. In both cases, we can find some piece of explanation in the specificities of the labor market working or the economic shocks that have hit the countries. Finland experienced a very strong economic cycle at the end of the 80's and the beginning of the 90's. After a period of sustained growth, it was hurt by a very deep financial and economic crisis from 1990 to 1993. During this period Finland experienced a strong fall in employment (a 10 points decrease in the employment rate); the unemployment rate jumped from 4 to 20% in 1994. Honkapohja and Koskela (1999) give a clear analysis of this recession. They estimate a wage equation (on annual data) very close to ours, but taking into account additional variables. Unfortunately, they do not study the contribution of each variable to the rigidity of the wage cost and to the deep deterioration of employment. The comments added to their article show that there is no consensus on this issue. As this shock was a very specific event, our results should not be interpreted as meaning that the Finish labor market would react very differently from the others in face of a common shock in Europe.

Franks (1994) got the same findings as us for Spain. He estimated Spanish wage equation from 1976 and 1992 and found a significant negative impact of productivity and a positive and significant impact of unemployment. He interprets this result as a consequence of the

[20] Layard, Nickell and Jackman (1991), Tyrvaïnen (1995), Mc Morrow (1996), Sinclair and Horsewood (1997), Roeger and in't Veld (1997), Cadiou, Guichard and Maurel (1999))

[21] The ranking of countries is quite different from Blanchflower and Oswald (the diversity of estimation periods in their study is however an important obstacle to national comparisons).

[22] Migration between states is much easier in the US than in Europe. But in this area the adjustment of the labor market can take place at the level of the participation rate, especially female participation. This can explain why unemployment plays a much less significant rôle in our estimations than the rate of employment.

great rigidity of Spain's labor market and of its deep duality. The 80's were characterized by deregulation of the labor market designed to increase its flexibility, mainly through the introduction of fixed term contracts. These new contracts met a big success with employers and they are now representing 30% of employment. However, by reducing the risk for the long-term employees to lose their job, they have increased their bargaining power[23].And deconnected this power from the values taken by the rates of unemployment or employment. The important power of these insiders is an essential explanation of the fact that the degradation of employment did not result in a fall of real wages. On the other hand the strong decrease in unemployment of these last years has not induced an increase in wages. The Spanish case is hence more problematic than the Finish ones: its particular institutions prevent the labor market from constituting a source of adjustment to macroeconomic shocks; moreover they may lead to very different reactions of the Spanish economy in face of a common shock in Europe.

Estimations have also been made by constraining the coefficient having wrong signs to be equal to zero. Unconstrained and constrained results are given in Table 2. Both results are similar.

Table 2: The results with the employment rate

	Non constrained estimation			Non significant and negative parameter constrained to 0	
	Coefficient	student	$-w_{3i}\overline{u}_i$	coefficient	student
Common parameters					
Q	0.69	5.66		0.53	5.59
R	0.43	5.08		0.49	7.35
W2	0.16	2.46		0.24	5.20
W1					
Austria	0.75	9.09		0.61	8.71
Belgium	0.46	2.36		0.52	2.89
Canada	0.89	9.72		0.81	11.47
Finland	0.66	17.70		0.67	19.91
France	0.36	3.51		0.37	4.17
Germany	0.57	6.25		0.49	5.86
Greece	-1.22	-1.54		0	Const.
Ireland	0.61	5.63		0.67	6.85
Italy	0.27	2.83		0.38	5.15
Japan	0.51	4.16		0.39	3.34
Netherlands	0.29	1.98		0.18	1.19
Portugal	0.19	0.86		0.31	1.49
Spain	-0.08	-0.35		0	Const
Sweden	0.54	2.20		0.53	2.31
UK	0.91	4.97		1.03	5.79

[23] See Bentolila and Dolado (1994).

United States	0.62	2.37		0.75	2.7
		W3			
Austria	2.16	6.42	-0.11	1.86	6.07
Belgium	0.41	1.00	-0.05	0.51	1.34
Canada	-0.57	-0.96	0.06	0	Const
Finland	-0.12	-1.85	0.01	0	Const
France	0.48	1.50	-0.05	0.60	2.12
Germany	0.80	3.53	-0.06	0.77	3.82
Greece	4.00	2.20	-0.33	5.87	3.84
Ireland	0.03	0.19	0	0	Const
Italy	0.70	3.28	-0.07	1.00	5.35
Japan	0.46	1.30	-0.01	0.52	1.52
Netherlands	0.64	4.05	-0.05	0.42	2.53
Portugal	-0.42	-1.30	0.03	0	Const
Spain	-0.84	-2.94	0.16	0	Const
Sweden	0.09	0.30	0	0.09	0.32
UK	0.98	4.19	-0.09	0.94	4.23
United States	0.06	0.21	0	0	Const

p- value of Hansen test: 8%

IV - CONCLUSION

The comparison of the results of the estimation of a wage setting equation between countries shows some similarities : average length of wages contracts, nominal rigidities, effects of the wedge. The main differences are in the impact of tensions on the labor market and of productivity. These differences are limited for the core countries of the EU (if we except Germany). However, there does not seem to exist simple relationship between these differences and the institutional features of the labor market. There are many reasons to think that the heterogeneity of the labor market is as strong inside countries that between countries. Moreover differences in the labor markets between countries bear on many dimensions and can hardly be summed up by a few macro economic indicators *a fortiori* by some general features. This may explain why it is difficult to find fully reliable differences in wages equations between countries, and to explain these differences in economic terms. In general, these differences look limited enough not to induce very different macro economic adjustments between countries. Anyway, national macro economic policies had yet to deal with very different situations on the national labor markets. A common monetary policy and the harmonization of fiscal policy should not make this task much more difficult.

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APPENDIX 1

Estimation of the system of equations (7) by GMM

Let W_i be the matrix of observations for the instruments related to i, of size ($T-2, n$). Let W_{it} be its typical line. Then, we define by $V_t = (e_{1t}W_{1t} .....e_{It}W_{It})$ the line vector of size In, and by V the matrix with typical line V_t and dimension ($T-2, In$). The moment's condition is:

$$EV_t = 0 \qquad \text{(A1.1)}$$

We approximate the theoretical moments by the empirical moments and we get:

$$V'\iota = 0 \qquad \text{(A1.2)}$$

where $\boldsymbol{i}$ is a column vector of 1 with dimension $T-2$. Condition (A1.2) cannot be exactly checked in most cases where the total number of instruments is larger than the number of parameters to estimate. The difference between the number of instruments and the number of parameters, denoted by s, is called the degree of over identification of the estimation. Thus, we try to minimize the distance between $V'\iota$ and 0, by using a distance matrix A, of dimension: (In, In), which is symmetric and positive definite. Thus, we minimize relatively to parameters the expression:

$$\iota' VAV' i \qquad \text{(A1.3)}$$

The efficient choice of the matrix A is: $A = \Phi^{-1}$, with $\Phi = [1/(T-2)] \underset{T\to\infty}{pl} (V'V)$, where pl represents the probability limit. We have assumed that the error terms of a same date have the covariance matrix Ω, which is independent of time (time homoscedasticity), and with typical element w_{ij}. Then, Φ has for typical element: $\omega_{ij} W_i' W_j / (T-2)$. To compute A, we must invert this matrix, of dimension (In, In)[24].

In practice we proceed through two steps. In the first step, we assume the matrix Ω to be proportional to the identity matrix. Thus, A is the block diagonal matrix, with typical

[24] In the computation of Φ we could expect to get estimators of the parameters and of their covariance matrix robust to heteroscedasticity by substituting in the second step, for the computation of Φ, $\hat{\omega}_{ij} W_i' W_j / (T-2)$ by $\sum_{t=2}^{T-2} W_{it}' \hat{\varepsilon}_{it} \hat{\varepsilon}_{jt} W_{jt} / (T-2)$. The problem is that this new estimator is the sum of Φ $T-2$ matrices of dimensions (In, In), but of rank 1. Indeed the matrix indexed t is the product of the column vector V_t' by its transpose. Consequently, the rank of the estimator of is at most equal to $T-2$. In most applications it will be less than In, and matrix Φ will be singular so non-invertible. Thus it seems impossible to build estimators robust to heteroscedasticity for our problem.

block: $(W_i^{'}W_i)^{-1}$. We minimize criteria (A1.3), and thus we get a first value for the parameters and the residuals. Then, we compute an estimator of the w_{ij}, denoted by $\hat{w}_{ij}$, as the empirical covariance between residuals $\hat{e}_{it}$ related to country i and residuals $\hat{e}_{jt}$ related to country j. In the second step, we estimate Φ by its typical block: $\hat{\omega}_{ij}W_i^{'}W_j\,/(T-2)$. We estimate A by Φ^{-1}, and we make the above minimization. The second step may be iterated several times.

The covariance matrix of estimated parameters (time $(T-2)^{1/2}$) is asymptotically equal to $(\Delta'\Phi^{-1}\Delta)^{-1}$, where Δ is the matrix of the partial derivatives of $V'\iota/(T-2)$ relatively to the parameters.

APPENDIX 2

Estimation of the covariance matrix

e_t represents here the vector of error terms for the set of all nations (of dimension I) and for $t = 2,..,T-1$. We denote in the same way the random vector, its realization and its estimation. We make the following assumptions where the number of factors is equal to f :

$$e_t = \Lambda F_t + u_t \tag{A2.1}$$

F_t represents column vectors of dimension f and u_t column vectors of dimension I, that are random.

Λ is a matrix of dimension (I, f) and is certain.

$EF_t = Eu_t = 0$, $E(u_t u_t^{'}) = D = diag(d_1,.., d_t)$ [25], $E(F_t u_\tau^{'}) = 0$, $\forall t, \tau$.

$E(F_t F_\tau^{'}) = E(u_t u_\tau^{'}) = 0, \forall \tau, t \neq \tau$, $E(F_t F_t^{'}) = U$ [26].

Then, we deduce:

$$\Omega = \Lambda\Lambda' + D \tag{A2.2}$$

Instead of having to estimate the $I(I+1)/2$ parameters of Ω, we just have to estimate the $(f+1)I$ parameters of Λ and D (actually the improvement is meaningful only when the number of factors is much smaller than half the number of countries). We can show that the maximum likelihood estimators of Λ and D, denoted by $\hat{\Lambda}$ and $\hat{D}$, are given by conditions:

$A = \sum_{t=2}^{T-1} (\varepsilon_t - m)(\varepsilon_t - m)'/(T-2)$, where m is the arithmetic mean vector of the e_t over the estimation period.

[25] $diag$ means a diagonal matrix with the following diagonal elements.

[26] U represents an identity matrix, which in this case is of dimension (f, f).

$1+\gamma_1,..,1+\gamma_I$, are the real positive eigenvalues of $A\hat{D}^{-1}$, which are assumed to be different and ranked by decreasing values (actually, the f first γ_i must be positive for the computation to be possible),

Γ is the diagonal matrix of dimension (f, f) with diagonal elements: $g_1, . g_f$.

the f columns of $\hat{\Lambda}$ are the f first eigen vectors of $A\hat{D}^{-1}$ (related to the f largest eigenvalues) which are normed to check for the identification condition: $\Gamma = \hat{\Lambda}\hat{D}^{-1}\hat{\Lambda}$).

The estimation procedure is iterative. First, we give an initial value to $\hat{D}$: D_0. Then we compute the eigenvalues and the eigen vectors of AD_0^{-1},and consequently Λ_0. Then, we compute D_1 which is the diagonal matrix, the diagonal elements of which are the same as for $A-\Lambda_0\Lambda_0^{'}$, and we start again. This procedure appears to converge easily in applications, although to our knowledge there do not exist mathematical results proving this property. More sophisticated estimation methods exist and are given by Doz[27].

The choice of the initial value D_0 is a supplementary problem. We denote by R_i^2 the square of the multiple correlation coefficient between the *ith* component of e_t and the $I-1$ other components, and by a_{ij} the typical element of matrix A. Then, we choose: $d_{i0} = a_{ii}(1-R_i^2)$.

Another difficulty is the choice of the number of factors f . A simple method is to compute a matrix of the same dimension as A, the non diagonal terms of which represent the correlations between the components of vector e_t, and the diagonal terms of which are the R_i^2. Then we make a principal component analysis of this matrix, and we keep as many factors as there exists non-negligible positive eigenvalues.

This *a priori* test is sufficient at the beginning of a succession of iterations of GMM, when the fact that matrix Φ may be a little wrong bears no serious consequences. However an *a posteriori* test of the validity of the choice of the number of factors, more rigorous, must be made at the last step of GMM. This test, of the likelihood ratio kind, uses as null hypothesis that the number of factors is equal to f . The alternative hypothesis is that there does not exist any constraint on the covariance matrix Ω. The statistics of the test is:

[27] The empirical covariance A and its estimated approximation $\hat{\Omega}$ have the same diagonal. This results from the fact that the factor representation does not change variances, but simplifies the structure of the covariances by assuming that it results from a small number of common factors.

$$\xi = -(T-2)\sum_{j=p+1}^{I}\ln(1+\gamma_j) \qquad \text{(A2.3)}$$

This statistics asymptotically verifies a χ^2 with a number of degrees of freedom equal to $[(I-p)^2-(I+p)]/2$. Bartlett suggests substituting, in the expression of $\boldsymbol{x}$, the number of observations $T-2$ by $T-2-(2I+5)-2p/3$, when the number of observations is low, which is the situation we face here.

APPENDIX 3

The autocorrelation of error terms

The econometric methodology presented in section 2, and its implementation in section 3, neglect the inter temporal dependence of error terms. This dependence has two unfavorable consequences. First, our choice of instruments can become unjustified, that is the moment condition (A1.1) of Appendix 1 may become non verified. Secondly, our choice of the weights matrix $A = \Phi^{-1}$, is no more efficient, and the covariance matrix of the estimated parameters, as the Hansen statistics, are no more computed adequately. For reasons we gave in footnote 24, it seems impossible to find a solution for the second problem. By taking the explained and explanatory variables with a lag as high as three years as instruments we limited the consequences of the first problem. We still have to evaluate if three years makes a long enough lag. To do that we reestimated the last equations with the constant term and the explanatory and explained variables lagged four years as instruments. Then, we computed the Hansen statistics and found its *p-value* to be high (equal to 14,7%). Then, we added to the lists of instruments the explained variable with a lag of three years. The *p-value* of the Hansen test was still high (19%). Then, we computed the difference between both Hansen statistics. If the new instrument is valid this difference follows a $\boldsymbol{c}^2$ with 16 degrees of freedom (the number of supplementary variables time the number of countries). We got a *p-value* of 51% and we concluded that using the explained variable with a lag of three years as an instrument was justified.

We made the same test for each explanatory variable, and we got a high *p-value* (more than 20%), in each case but for the price P for which the *p-value* was low (0,04%). An interpretation is that nominal inertia is longer than what is assumed by our specification. Thus, equation (6) would omit a price variable with a lag of two years, which would invalidate our choice of the price with a lag of three years as an instrument (we must remember that the estimated equation is the first difference of equation (6)). However, it seems difficult to find a theoretical or empirical justification of the presence of the price with a lag of two years in the equation determining average wages.

We did not do this test for all the variables simultaneously. Indeed, in this case the number of instruments becomes equal to 11 per country, which represents a total of: 1+10*16=161. With such a high number of instruments, the Hansen test produces very low *p-values*.

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