

The role of Institutions in Explaining Wage Determination in the Euro Area: a Panel Cointegration approach

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Abstract

Over the last 15 years, the evolution of labor costs has been very diverse across EMU countries. Since wages have important second-round effects on prices and competitiveness, and EMU countries do not have the tool of the nominal exchange rate to correct for such imbalances, understanding the determinants of the wage is a matter of increasing concern and debate. We estimate the equilibrium wage equation for the Euro Area over the period 1995-2010 using panel cointegration techniques that allow for cross-section dependence and structural breaks. The results show that the equilibrium wage has a positive and proportional relation with productivity and negative relation with unemployment, as expected. Moreover, real exchange rate appreciation triggers a drop in the real wage. We also include institutional variables in our analysis, showing that a more flexible labor market is consistent with long-run wage moderation. Allowing for a regime break, we find that in the second half of the sample the equilibrium wage equation can be interpreted as stationary unit labor costs. We argue that increased international competition made wage bargaining less influenced by domestic factors, so that wage developments became relatively more linked to productivity and the real exchange rate.

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1. Introduction

The wage equation is a crucial element of any comprehensive model of the macroeconomy. A large strand of literature, since Blanchflower and Oswald (1994), has estimated the relationship between wages and unemployment *à la* Phillips (1958), and more generally the equilibrium equation for the real wage.¹

In the present paper, we use empirical techniques recently introduced by the literature to provide an estimation of the long-run equilibrium wage equation for the Euro Area, also taking into account the role of labor market institutions.

As shown in Figure 1, over the last 15 years the evolution of the real wage across EMU countries has been very diverse. These differences, in turn, were reflected in aggregate unit labor costs. On the one hand, wage setting has important second-round effects on prices and, therefore, potentially on competitiveness: when a country records persistently high inflation - for example due to increasing unit labor costs - with respect to the other member states of the Monetary Union, it will experience real exchange rate appreciation and a progressive loss of competitiveness. On the other hand, EMU countries cannot use the tool of the nominal exchange rate to correct divergent price dynamics. For this reason, the only way to correct such inflation differential in absence of nominal exchange rate depreciation is via internal devaluation.

Understanding how wage determination works in the euro area is therefore a matter of primary concern, also given the importance that the European Commission has been giving to labor market reforms in several Member States.. At the same time, since EMU countries are also characterized by quite different labor market institutions, it is reasonable to think that the institutional framework, together with macroeconomic developments, has an important impact on wages.

Against this background, our paper studies the long-run determinants of real wages for EMU-11 countries, with a focus on labor market institutions. We extend a classic wage equation (Blanchard, 2000) to take into account also the role of labor market regulation and wage bargaining. Thus, one important contribution of this paper is that it is the first work estimating equilibrium wage equations using cointegration that also accounts, in the long-run estimation, for institutional factors. Including institutional variables in the cointegration relation implies assuming that changes in the institutional set-up, for example a reform that increases the flexibility of the labor market, have permanent effects on the level of the real wage. Second, we also estimate the wage equation separately for the “Core” and “Periphery” countries, given the importance that wage setting may have on competition and the creation of current account imbalances across the EMU.

¹ To name but a few, Nunziata (2005), Broersma et al. (2005), Marcellino and Mizon (2000), Alesina and Perotti (1997), Baltagi et al. (2000).

However, the main innovative feature of our analysis comes from the statistical design. We estimate the determinants of real wages using panel cointegration techniques that take into account the issue of cross-section dependence in the data and the presence of breaks in the series as well as in the cointegration relations. The issue of cross-section dependence is crucial when using macro data, especially for countries that are so much interlinked as EMU or OECD countries. Thus, we begin by testing for cross-section dependence using Pesaran's (2004) CD test and for the presence of a unit root in the data using the panel unit root tests CADF proposed by Pesaran (2007) and PANIC (Bai and Ng, 2004), which allow for such dependence. Moreover, we test for cointegration between the variables using panel cointegration tests which allow for an (unknown) break in the equilibrium relations, in particular the ones proposed by Banerjee and Carrión-i-Silvestre (2013). This is extremely important, considering the institutional changes national labor market in the Euro Area have been going through during the last 20 years, not to mention the introduction of the euro. We also efficiently estimate the long-run relationship (i.e. the wage equation) using the CUP-BC and CUP-FM estimators proposed in Bai et al. (2009).

The paper is organised as follows: section 2 discusses the theoretical background; section 3 presents the data used in the empirical analysis; section 4 introduces the panel unit root and cointegration tests performed, and their results; section 5 presents the results of the panel estimation of the long-run relationship. Section 6 reports the result of the estimation by groups. Section 7 concludes.

2. Theoretical background

In a world with perfect competition, wages adjust to clear the labor market. The observed real wage is thus the result of the bargaining process between employees' unions and employers.

On the labor supply side, employees' unions tend to push for wage increases above productivity; however, their bargaining power depends on the unemployment rate since wage demands by unions tend to be more moderate when unemployment is high. Thus, we can write

$$r_comp_t = f(prod_t, unemp_t); f'_{prod} > 0, f'_{unemp} < 0$$

where r_comp_t is the (log) real wage, measured as real compensation per hour worked; $prod_t$ is (log) labor productivity and $unemp_t$ is the unemployment rate.

On the labor demand side, employers tend to constrain the real wage, maximizing their mark-up on unit labor cost, where the latter is defined as $r_comp_t - prod_t$. The mark-up that employers will be able to extract from the real wage will in turn be related to the real exchange rate. The real exchange rate can affect labor costs in

different ways. First, a depreciated exchange rate increases the demand for domestic goods, thus raising labor demand and the real wage (Campa and Goldberg 2001). Second, depreciation increases the price of imported intermediate goods and thus production costs; to the extent that those goods are complement to labor, it will foster a reduction in labor demand and in the real wage (Robertson, 2003). Third, depreciation of the real exchange rate implies that imported goods are more expensive, which makes the consumer price index increase and real wage decline. If we define the real effective exchange rate as units of (trade-weighted) foreign goods per unit of domestic goods, the first effect would imply a negative relationship between $w_t - p_t$ and the real exchange rate, while the second and third would imply a positive relationship. On the demand side, we can therefore write

$$r_{comp}_t = g(prod_t, reer_t)$$

Where $g'_{prod} > 0$; $reer_t$ is the real exchange rate and therefore $g'_{reer} \geq 0$ depending on which of the channels described above prevails.

Finally, the observed equilibrium wage is the result of additional wage pressure factors, which can be more generally classified as wage setting institutions. In particular, a more flexible labor market should be associated with wage restraint: by increasing the bargaining power of insiders, higher employment protection may put upward pressure on bargained wages. However, previous literature that estimated the effect of employment protection has been focusing more on its impact on employment than wages. Moreover, other institutional factors such as the degree of coordination in wage bargaining and the involvement of unions in government decisions also affect the equilibrium wage².

Therefore, combining the labor supply and demand side and in analogy with existing work (Nickell, 1998; Bell et al. 2002; Nunziata 2005), the equilibrium wage equation can be written as a reduced-form specification suitable for estimation, incorporating both demand- and supply-side factors as well as labor market institutions:

$$r_{comp}_t = \beta_0 + \beta_1 prod_t + \beta_2 unemp_t + \beta_3 reer_t + \beta_4' \mathbf{Z}_t + \varepsilon_t \quad (1)$$

Where we expect, a priori, $\beta_1 > 0$, $\beta_2 < 0$ and $\beta_3 \geq 0$; $\mathbf{Z}_t$ is a vector of variables defining wage setting institutions.

In particular, the vector of variables defining wage-setting institutions we consider in the present paper is:

$$\beta_4' \mathbf{Z}_t = \gamma_1 WCOORD_t + \gamma_2 EPL_t + \gamma_3 UD_t$$

² See Calmfors and Driffill (1988), Nunziata (2005), Boeri et al. (2001), among others.

Where $WCOORD_t$ is the degree of wage coordination (a higher $WCOORD_t$ implies stronger coordination), EPL_t represents the degree of employment protection and UD_t is union density (i.e. the ratio of total union members to salaried employees).³

According to the Calmfors and Driffill (1988) hypothesis, the relationship between centralization in wage bargaining (i.e. $WCOORD_t$) and the aggregate wage is not linear. In particular, with dominant firm-level bargaining, wage claims have a direct effect on the firm's competition, and this will have a moderating effect on the unions' claims. However, at the same time, with economy-level bargaining and thus maximum co-ordination in wage bargaining, unions will internalize the cost of excessive wage claims and this will have a moderating effect on the wage. For this reason, the relationship between $WCOORD_t$ and the real wage should have an inverted-U shape. Our specification will not allow us to test for the Calmfors and Driffill (1988) hypothesis; however, the expected sign of γ_1 depends on such hypothesis. In fact, since $WCOORD$ in our data may take values from 1 (dominant firm-level bargaining) to 5 (dominant economy-wide bargaining)⁴ we can graphically represent the relationship between $WCOORD$ and the real wage suggested by the Calmfors-Driffill hypothesis as in Figure 2. Therefore, depending on whether the dominant level of wage bargaining within our sample is firm level (i.e. $WCOORD < 3$) or national level ($WCOORD > 3$), γ_1 will be positive or negative. Within our dataset, in all countries except France (also Ireland and Portugal, but for a very short period) $WCOORD$ has always taken values between 3 and 5. Therefore, we are in the right-hand side of Calmfors and Driffill's inverted-U curve. Based on this discussion, we would therefore expect $\gamma_1 \leq 0$. As far as EPL_t is concerned, instead, from the previous discussion we should expect stricter employment protection to result in higher wages, ceteris paribus and thus $\gamma_2 > 0$. Finally, since higher Union Density increases unions' bargaining power, it is often assumed that $\gamma_3 > 0$. However, this does not necessarily translate into higher wages. In fact, as discussed by Checchi and Nunziata (2011), national union leaders may internalize the effect of excessive wage pressure, and especially so when centralization is high, as it is the case in EMU-11 countries. In this sense, higher Union Density may trigger wage moderation and $\gamma_3 < 0$. We will come back to this point in Section 6.

3. The Data

³ We use Union Density in place of bargaining coverage, being aware that it should be the latter, rather than Unions' size, which affects wage determination in the long run, since, with higher coverage, wage increases are extended to a bigger pool of workers. However, UD is also a proxy of Unions' strength, since we would expect Unions' bargaining power to be stronger if they represent a higher share of workers. Moreover, available data on bargaining coverage are of poorer quality.

⁴ See Appendix 1 for details.

We use quarterly data from 1995Q1 until 2010Q4 on the group of countries generally referred to as EMU-11: Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain⁵. Macroeconomic data comes from Eurostat and the variables are all seasonally and working day adjusted. The wage is defined as employees' compensation per worker, and is therefore calculated as the ratio of total compensation over the number of persons employed:

$$comp_p = \frac{comp_emp}{emp_p}$$

Productivity is calculated as output per worker, and therefore is the difference between the log of GDP and the log of persons employed.

Institutional variables come from J. Visser's ICTWSS database⁶, except for EPL that is provided by OECD⁷. The ICTWSS database provides annual data on 34 countries (all out countries are included in the database) about trade unionism, wage setting, state intervention and social pacts from 1960 to 2010. Since both ICTWSS and OECD institutional data are annual, we had to transform them to quarterly: to that end, we used quadratic interpolation for *UD*; for the other institutional variables, whenever a change in the annual series was present we looked for a corresponding reform or regulatory intervention within the year and thus constructed the quarterly series: dates of labor market reforms and collective agreements were taken from European Commission, Directorate General for Economic and Financial Affairs and Economic Policy Committee LABREF database.

Table 1 reports summary statistics of the variables involved. A detailed description of the variables and data sources is provided in Appendix 1.

4. Panel Unit Root and Cointegration Tests

We first applied Pesaran's (2004) cross-section dependence test (CD Test henceforth) to the variables. This test is formulated having as null hypothesis cross-section independence, so that its rejection would mean that dependence is found among the individual countries in the group and should be accounted for in the remaining panel tests. This is the case for the four variables in Table 2: the null hypothesis of independence is clearly rejected⁸.

⁵ The other current member states of the EMU were excluded due to lack of data, or because the available time series were too short.

⁶ <http://www.uva-aiaa.net/208>

⁷ EPL is not available for Luxembourg, which was therefore excluded from the institutional analysis.

⁸ The test was performed in Stata using the *xtcd* code provided by Markus Eberhardt in his webpage.

Once we have found the presence of dependence in the variables, we have studied their order of integration using two different tests that account for dependence. Both are representative of the “second generation” panel unit root tests⁹.

First, we apply Pesaran’s (2007) CADF test. In this case, unlike previous tests that demeaned the series to correct for the existence of dependence among the members of a panel, he augments the standard DF or ADF regressions with the cross-section averages of the lagged levels as well as the first differences of the individual series. Based on this procedure, he suggests developing modified versions of the t-bar test of Im et al. (2003) IPS test, the inverse chi-squared test (the P test) of Maddala and Wu (1999), and the inverse normal Z test suggested by Choi (2001). Pesaran (2007) defines the test statistics for the model with constant, with trend and with constant and trend. The main advantage of Pesaran’s CADF test is that is simple and intuitive. Moreover, it is also valid for panels where N and T are of the same order of magnitude. This is frequently the case of panel unit roots applied to macro-variables.

Second, we also apply Bai and Ng (2004), a suitable approach when cross-correlation is pervasive, as in this case. Furthermore, this approach controls for cross-section dependence given by cross-cointegration relationships, potentially possible among our group of countries and variables — see Banerjee et al. (2004). This is clearly the case for wages, but also for the real effective exchange rate.

Bai and Ng (2004) make use residual factor models to take account of dependence. From a rather general set-up, they allow for the possibility of unit roots and cointegration in the common factors. However, they still assume that $N/T \rightarrow 0$, as N and $T \rightarrow 1$. They apply the principal component procedure to the first-differenced version of the model. Then, they estimate the factor loadings as well as the first differences of the common factors.

They decompose the $Y_{i,t}$ time series as follows:

$$Y_{i,t} = D_{i,t} + F_t' \pi_i + e_{i,t},$$

with $t = 1, \dots, T$, $i = 1, \dots, N$, where $D_{i,t}$ denotes the deterministic part of the model — either a constant or a linear time trend — F_t is a $(r \times 1)$ -vector that accounts for the common factors that are present in the panel, and $e_{i,t}$ is the idiosyncratic disturbance term, which is assumed to be cross-section independent. As stated above, unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model. For the estimated idiosyncratic component, they propose an ADF test for individual unit

⁹ See Breitung and Pesaran (2007) and Choi (2006) for a review of second-generation panel unit root tests.

roots and a Fisher-type test for the pooled unit root hypothesis (P_ε), which has a standard normal distribution. The estimation of the number of common factors is obtained using the panel BIC information criterion as suggested by Bai and Ng (2002), with a maximum of six common factors.

Bai and Ng (2004) propose several tests to select the number of independent stochastic trends, k_1 in the estimated common factors, $\hat{F}_t$. If a single common factor is estimated, they recommend an ADF test whereas if several common factors are obtained, they propose an iterative procedure to select k_1 : two modified Q statistics (MQ_c and MQ_f), that use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution. The null hypothesis of $k_1 = m$ is tested against the alternative $k_1 < m$ for m starting from $\hat{k}$. The procedure ends if at any step $k_1 = m$ cannot be rejected. The results from the application of the Bai and Ng (2004) statistics are summarized in Table 3.

The results of applying Pesaran (2007) tests are presented in the second and the third columns of Table 3¹⁰. The headings $CADF(4)^c$ and $CADF(4)^t$ correspond, respectively, to the model with constant and with trend, where the number of lags chosen is $p=4$. The null hypothesis of unit root cannot be rejected for any of the variables analysed in the case of the model with constant. When the specification includes a trend, the unit root null is rejected for the real effective exchange rate. The right-hand side of the table is devoted to the results of Bai and Ng (2004) panel unit root test. The number of chosen common factors is the maximum (six) for the two statistics and the first three variables. The exception this time is unemployment, where two factors are found in MQ_c and three in MQ_f . Concerning the unit root tests, we find again rejection in the idiosyncratic ADF test for the real effective exchange rate, whereas, according to the MQ tests all the common components are non-stationary.

Given our a priori theoretical expectations as exposed in section 2, we are looking for an equilibrium relationship between wages, productivity, the unemployment rate and the real (effective) exchange rate, which can then be augmented to include institutional variables as in equation (1).

Thus, the next step in our empirical strategy is to test for cointegration applying the Banerjee and Carrión-i-Silvestre (2013) test¹¹. Using factor models to account for cross-section dependence, they propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector. It is worth noticing that inference concerning the presence of cointegration can be affected by misspecification if the existence of

¹⁰ The test was performed using Piotr Lewandowski's pescadf.ado Stata code.

¹¹ See Appendix 2 for a description of the specified models and the test.

breaks is ignored. They propose a test formulated for six different specifications of the deterministic components including a constant, a trend and structural breaks. They then recover the idiosyncratic disturbance terms ($\tilde{\epsilon}_{it}$) through accumulation of the estimated residuals and test for the null of no cointegration against the alternative of cointegration with breaks using the ADF statistic. In our case we concentrate in models with common homogenous structural breaks.

As shown in Table 4, using the Banerjee and Carrion (2013) to test for non-cointegration, we apply the statistic based on the accumulated idiosyncratic components, Z_j^* . We present the test for all possible specifications; in all cases the null hypothesis of non-cointegration is rejected. Moreover, according to the information criteria¹², the appropriate model could be either Model 3 (constant and trend restricted in the cointegration relation, and a break in both) or Model 6 (the break affects the level, the trend and the cointegrating vector, i.e. a “regime break”). The break is found to be in 2004Q2 in Model 3 and 2004Q4 using Model 6¹³. The Banerjee and Carrion (2013) test thus suggests that the preferred models include a trend in the cointegration relation. While including a linear term in the equilibrium relationship may be criticized since it is difficult to interpret economically, on the other hand it allows us to define the cointegration as one between the real wage and trend-adjusted productivity (plus other variables), where the trend proxies technological progress¹⁴.

5. The wage equation of the EMU

Following the results of the Banerjee and Carrion-i-Silvestre (2013) test, we estimated equation (1) using the CUP-FM and CUP-BC estimators (Bai et al., 2009) for the model with a break in the deterministic components (Model 3) and the model with a regime break (Model 6). The results are reported in Tables 4 and 5 for equation (1) for the specification without and with institutional variables, respectively.

First of all, let us consider Table 5, where the results of the cointegration analysis using macroeconomic variables *prod*, *unemp* and *reer* are reported. All coefficients have the expected sign. Higher labor productivity is associated with higher real wages, while unemployment tends to make the real wage fall. However,

¹² Using AIC we would have chosen Model 6, whereas according to BIC the best specification would be Model 3. Therefore, we present the results for both models. Moreover, the comparison between the two models is also interesting, as Model 8 allows for a structural change in the cointegrating relationship. The information criteria results are available from the authors upon request.

¹³ We will discuss on the interpretation of this date in the following section.

¹⁴ Moreover, including a linear term is a standard procedure in cointegration analysis when the variables included seem to exhibit a deterministic trend, other than a stochastic trend, and the slope of such trend is different across variables. In this case, excluding the trend from the cointegration relation may leave us with residual non-stationarity. See also Juselius (2006).

looking at Table 5, panels a. and b. our results also suggest that such negative relation is limited to the period before the break. Moreover, note that the coefficient for the real exchange rate is negative and significant, meaning that a real depreciation pushes the real wage up. Following our discussion in Section 2, this means that the “increased demand of domestic goods” effect following a real depreciation outweighs the purchasing power and labor demand effects. Note, however, that the long-run coefficient for *unemp* is never significant if we take Model 3 as a reference. It is significant, and negative, when we allow for a break in the cointegrating vector. In other terms, in the second half of our sample (after 2004Q4), unemployment is no longer a determinant of the long-run wage¹⁵. We will come back to this point below.

As far as the institutional variables are concerned, the results are reported in Table 6. We estimated the model including, alternatively, *WCOORD* and *UD*, because the two variables appear to be collinear. More precisely, the impact of unions’ bargaining power on wages is likely to be compensated by coordination.¹⁶ Stricter employment protection is associated with higher real wage in the long run, the coefficient of *EPL* being positive and significant. Since it can be cumbersome to interpret what a coefficient of 0.017-0.021 means for an institutional variable such as *EPL*, table 7 reports the estimated long-run effect on wages of some relevant reforms of employment protection adopted in Europe over the last 20 years. For example, we estimate that the “Hartz reforms” in Germany in 2003 brought about a contraction in the real wage in the long run of about 0.7%, while the Treu Law in Italy was associated with a drop in the wage between 1 and 1.2% in the long run. Higher coordination in wage bargaining (higher *WCOORD*) is associated with a lower long-run real wage, given that the coefficient is negative. This confirms again our prior expectation of being “located” in the right-hand side of Calmfors and Driffill’s inverted-U curve. In this sense, higher centralization of wage bargaining would bring about a drop in real wages.

Union Density (*UD*) deserves a specific comment. The variable presents a long-run negative coefficient. This goes against what the *a priori* expectation suggesting that the relationship should be positive: the higher Union Density, the higher the bargaining power of unions and therefore the aggregate wage. Indeed, it confirms the opposite view that Union leaders may internalize the cost of excessive wage increases, as discussed in section 2 and suggested by Checchi and Nunziata (2011). In this sense, above a certain threshold of union density, an increase in it may trigger wage moderation, a result which is consistent with Checchi and Lucifora’s (2002) “good” view of unions, i.e. of unions being welfare-enhancing.

¹⁵ Indeed, we can see from Table 5 that unemployment is the only variable for which the coefficient representing the break in the CV is significant.

¹⁶ See Nunziata (2005), Nickell and Layard (1999) and Boeri, Brugiavini and Calmfors (2001).

An additional interesting result of our analysis comes from the estimation of Model 6, that allows for a break in the cointegration vector. We have already seen that, when accounting for this break, *unemp* becomes significant in the first half of the sample in Table 5. Moreover, Table 6 panel b. shows that, in absolute terms, the coefficient of all variables except *reer* became smaller after the break. The only variable, for which this change is proportionally small, and only significant at 10%, is productivity. The long-run coefficient for *reer* is, instead, three times bigger in the second half of the sample.

It is honestly difficult to interpret the 2004 break simply as far as the labor market is concerned. Indeed, the year was characterized by the Eastern enlargement of the EU, which was then completed in 2007 with the accession of Romania and Bulgaria. This supposedly brought about two shocks, on the goods production side and on the labor supply side. On one hand, firms in the Euro Area-11 (and, more generally, in the European Union-15) found themselves with increased product market competition due to the removal of all barriers *vis à vis* Central and Eastern Europe (CEE), although most of the process of trade barriers removal had already been completed in the previous years. On the other hand, free labor mobility from the CEE countries affected labor supply.

At the same time, the degree of openness of EMU-11 economies has largely increased over the sample period, with the exception of Ireland¹⁷. With firms facing increased international competition, it is realistic to think that wage bargaining was less influenced by domestic factors like unions' bargaining power and labor market institutions, and therefore wage developments became relatively more strictly linked to productivity developments, in an attempt to keep unit labor costs stable, and the real exchange rate.

6. Core and Periphery of the EMU

Estimating a single long-run equation for wage determination in the EMU-11, as we did in Section 5, has the advantage of using a higher number of observations in the cross-section dimension, but it entails the implicit assumption that the coefficients are the same across countries or, in other words, that by accounting for differences in (some) labor market institutions, we are controlling for all possible heterogeneity.

However, this need not be the case, and in order to dig deeper into the long-run determinants of the real wage in the EMU, we then adopted an alternative approach: we estimated the long-run regression, using the approach outlined in

¹⁷ For example, trade openness of Austria went from 62% in 1995-2004 to 84% in 2005-2010; in Germany, it went from 46% to 69.5%, in the Netherlands from 92% to 135% and in Italy from 36% to 45%.

section 5, for two different groups of countries, which we will call the Core and the Periphery. The Core is composed of Germany, Austria, Belgium, Luxembourg and the Netherlands, while the Periphery is composed of Spain, Italy, France and Ireland. By looking at Table 8, we can also say, from a purely descriptive perspective, that the Core's labor market is more "functional" while the Periphery's is more "dysfunctional": in fact, the former shows, on average, higher compensation per employee and labor productivity, lower unemployment rate and is more decentralized and flexible, while union density is higher.

According to the Banerjee and Carrion (2013) test, cointegration is present in both subsamples, with a break in the deterministic part in the Core (i.e., Model 3) and a regime break in the Periphery (i.e. Model 6)¹⁸.

The results of the estimation are reported in Table 9 (panels a.-c.). First of all, as far as the Core is concerned, an analysis which included all 6 core countries provided puzzling results, since the sign of the productivity coefficient is always negative and significant (Table 9, panel a.). On the other hand, the sign of *reer* is positive, i.e. the "price-level channel" and the "production costs channel" in Section 2 prevailed. The reason for the disappointing result concerning the coefficient of productivity in Table 9 a. may be related to the evolution of wages and productivity in Germany over our sample period. Especially since the early 2000s, Germany experienced a significant decrease in unit labor costs, which most likely contributed to its success in export markets. Such decrease was due, at the same time, to a fall in the real wage and an increase in productivity. Indeed, within our sample, the correlation between the two variables in Germany is equal to -0.76.

Therefore, in Table 9 panel b. we report the results of the CUP estimation of the Core excluding Germany. While other coefficients are generally unchanged, the long-run coefficient for productivity is no longer univocally negative and actually is positive and significant in the base model using the CUP-BC estimator.

In the case of the periphery, the model includes a structural break. The first thing to notice is that the long-run coefficient of productivity is significantly larger than one, suggesting that wages adjust more than proportionally to productivity in the long run. This was strengthened after the break.

There is a negative long-run relation between wages and the *reer*, but it became weaker after the break and possibly changed sign, a result that confirms that of Camarero et al. (2013). Moreover, the long-run coefficient of *EPL* is larger than in the Core, i.e. reducing employment protection in Periphery countries likely entails a larger reduction of the long-run real wage than it does in the Core. Nevertheless, this is not the case after the break, since the *EPL* coefficient became smaller. This may be due to a nonlinear relationship between *EPL* and *r_comp* which in our model is proxied by the break, so that increased flexibility of the labor market in Periphery countries induced by several reforms (In Italy in 1997Q1 and 2003Q1;

¹⁸ Results of the Banerjee and Carrion (2013) test in the two subsamples are available from the authors upon request.

in Spain in 2002) reduced the scope for a wage-moderating effect of further reforms of employment protection legislation.

Finally, while in the Core unemployment has a wage-moderating effect (i.e. there is a negative long-run relation between the real wage and the unemployment rate), this is not the case in the Periphery, although after the break there are signs of such negative relation having appeared or strengthened.

7. Conclusions

In his paper, we have estimated an equilibrium wage equation for the Euro Area over the period 1995-2010 using panel cointegration techniques which take into account the issue of cross-section dependence in the data and the presence of breaks in the series and in the cointegration relations. Moreover, we have also included institutional variables in the long-run equation, in order to show how a different design of the labor market may be consistent with wage moderation.

Our results confirm that the long-run real wage is positively related with productivity and negatively with unemployment; moreover, other things equal, when the real exchange rate is overvalued the real wage tends to be lower to compensate for the appreciation. A more flexible labor market is consistent with wage moderation, and our estimates suggest a significant wage-moderation effect of recent labor market reforms that have been taken since the 1990s.

The cointegration techniques adopted in the present work allowed us to include a break in the long-run relationships. The break was found to be in 2004 (second or fourth quarter, depending on the chosen model). While, on one hand, it does not seem possible to identify a specific event that triggered this break, on the other hand, the way that the wage equation was affected is very clear: in the second half of the sample, the long-run coefficient of all variables except productivity and the real exchange rate became smaller or even not significant. We argued in section 6 that increased product competition coming from international trade, as well as higher international labor mobility, both presumably resulting from the 2004 EU eastern enlargement, may have a role in explaining this break, together with the rapid increase in the degree of trade openness of each of the Euro Area-11 countries, except Ireland, over the sample period. Indeed, the changes in the cointegration vector implied by the 2004 regime break and shown in Table 5, panel b, suggest that firms became more concerned with the impact that real wage increases have on competitiveness.

In order to take into account possible heterogeneity of wage setting relations across EMU countries, we repeated the same approach for two groups, which we called the Core (Austria, Belgium, Finland, Germany, Luxembourg and the

Netherlands) and the Periphery (France, Ireland, Italy, Portugal and Spain), where, on average, the former has a more “functional” labor market (higher wages, lower unemployment, more flexibility and decentralization). Results suggest that, in the periphery, the long-run relationship between wages and productivity has been more than proportional; moreover, reforms heading towards labor market liberalization, other things equal, may have a higher impact on long-run wages there. Finally, our results have an additional, more general implication which is related to the Lucas (1976) critique: it is crucial to account for a break in the estimation of the long run real wage equation; failure to do so may reduce the power of cointegration tests, as discussed in the literature (see also Banerjee and Carrion-i-Silvestre 2006), and it may also lead to a misinterpretation of the estimated coefficients, as our discussion has demonstrated.

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Appendix 1. Series definition

<i>prices</i>	CPI index, seasonally adjusted using TRAMO/SEATS. Source: Eurostat.
<i>unemp</i>	unemployment rate. Source: Eurostat.
<i>comp_emp</i>	total compensation of employees. Source: Eurostat.
<i>emp_p</i>	thousands of persons employed. Source: Eurostat.
<i>comp_p</i>	Compensation per person employed. Calculated as: $\ln(comp) - \ln(emp_p)$
<i>gdp</i>	real GDP. Source: Eurostat.
<i>reer</i>	ULC-deflated real effective exchange rate. Source: IFS.
<i>WCOORD</i>	Coordination of wage bargaining. 5: economy-wide bargaining; 4: mixed industry- and economy-wide; ... 1= company-level bargaining. Source: ICTWSS Database.
<i>EPL</i>	Employment Protection Legislation. 0 = minimum employment protection; 5 = maximum employment protection. Source: OECD.
<i>r_comp</i>	real compensation per hour worked: $\ln(comp_emp) - \ln(emp_h) - \ln(prices)$
<i>prod</i>	labor productivity per hour worked: $\ln(gdp) - \ln(emp_h)$
<i>UD</i>	union density (in %). Source: ICTWSS Database.

Appendix 2 – The Banerjee and Carrion (2006) Cointegration Test with breaks and dependence

Banerjee and Carrion-i-Silvestre (2010) propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector, also accounting for the presence of cross-section dependence using factor models. They define a $(m \times 1)$ vector of non-stationary stochastic process, $Y_{it} = (y_{it}, x_{it})$ whose elements are individually $I(1)$ with the following Data Generating Process:

$$y_{it} = D_{it} + x'_{it}\delta_{it} + u_{it} \quad (A.1)$$

The general functional form for the deterministic term D_{it} is given by:

$$D_{it} = \mu_i + \beta_i t + \sum_{j=1}^{m_i} \theta_{ij} DU_{ijt} + \sum_{j=1}^{m_i} \gamma_{ij} DT_{ijt} \quad (A.2)$$

Where $DU_{ijt} = 1$ and $DT_{ijt} = (t - T_{it}^b)$ for $t > T_{it}^b$ and 0 otherwise, $T_{it}^b = \lambda_{ij}^b T$ denotes the timing of the j -th break, $j = 1, \dots, m_i$, for the i -th unit, $i = 1, \dots, N$, $\lambda_{ij}^b T \in \Lambda$, being Λ a closed subset of $(0,1)$. The cointegrating vector is a function of time so that

$$\delta_{it} = \begin{cases} \delta_{i1} T_{i0}^c < t \leq T_{i1}^c \\ \delta_{i2} T_{i1}^c < t \leq T_{i2}^c \\ \dots \\ \delta_{ij} T_{i,j-1}^c < t \leq T_{ij}^c \\ \dots \\ \delta_{in_i+1} T_{in_i}^c < t \leq T_{in_i+1}^c \end{cases} \quad (A.3)$$

with $T_{i0}^c = 0$ and $T_{in_i+1}^c = T$, where $T_{ij}^c = \lambda_{ij}^c T$ denoting the j -th time of the break, $j = 1, \dots, n_i$, for the i -th unit, $i = 1, \dots, N$, for the i -th unit, $i = 1, \dots, N$, $\lambda_{ij}^c \in \Lambda$.

Banerjee and Carrion-i-Silvestre (2006) propose six different model specifications:

Model 1. No linear trend - $\theta_{ij} = \beta_i = \gamma_{ij} = \mathbf{0} \ \forall i, j$ in (A.2) – and stable cointegrating vector - $\delta_{ij} = \delta_i \ \forall j$ in [A.3].

Model 2. Stable trend - $\theta_{ij} = \mathbf{0}; \beta_i \neq \mathbf{0} \ \forall i$ and $\gamma_{ij} = \mathbf{0} \ \forall i, j$ in (A.2) – and stable cointegrating vector - $\delta_{ij} = \delta_i \ \forall j$ in [A.3].

Model 3. Changes in level and trend - $\theta_{ij} \neq \mathbf{0}; \beta_i \neq \gamma_{ij} \neq \mathbf{0} \ \forall i, j$ in (A.2) – and stable cointegrating vector - $\delta_{ij} = \delta_i \ \forall j$ in [A.3].

Model 4. No linear trend - $\beta_i = \gamma_{ij} = \mathbf{0} \quad \forall i, j$ in (A.2) but presence of multiple structural breaks that affect both the level and the cointegrating vector of the model.

Model 5. Stable trend $\beta_i \neq \mathbf{0} \quad \forall i$ and $\gamma_{ij} = \mathbf{0} \quad \forall i, j$ in (A.2) with multiple structural breaks that affect both the level and the cointegrating vector of the model.

Model 6. Changes in the level, trend and in the cointegrating vector. No constraints are imposed on the parameters of equations (A.2) and (A.3).

The common factors are estimated following the method proposed by Bai and Ng (2004). They start by computing the first difference of the model; then, they take the orthogonal projections and estimate the common factors and the factor loadings using principal components.

In any of these specifications, Banerjee and Carrion-i-Silvestre (2013) recover the idiosyncratic disturbance terms ($\tilde{\epsilon}_{it}$) through accumulation of the estimated residuals and propose testing for the null of no cointegration against the alternative of cointegration with break using the ADF statistic.

The null hypothesis of a unit root can be tested using the pseudo t -ratio $t_{\tilde{\epsilon}}^j = (\lambda_i), j = c, \tau, \gamma$. The models that do not include a time trend (Models 1 and 4) are denoted by c . Those that include a linear time trend with stable trend (Models 2, and 5) are denoted by τ and, finally, γ refers to the models with a time trend with changing trend (Models 3 and 6). When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms.

TABLE 1

Summary Statistics

	<i>N. Obs.</i>	<i>mean</i>	<i>standard dev.</i>
<i>comp_h</i>	704	2.920	0.821
<i>unemp</i>	704	7.617	3.381
<i>prod</i>	704	1.481	0.116
<i>reer</i>	704	97.936	5.241
<i>UD</i>	672	33.860	18.031
<i>EPL</i>	576	2.420	0.718
<i>WCOORD</i>	704	3.554	0.957

TABLE 2

Pesaran's (2004) Cross-Section Dependence CD Test

	CD Test Statistic	P-value
<i>r_comp_{it}</i>	32.27	0.000
<i>reer_{it}</i>	34.19	0.000
<i>prod_{it}</i>	7.78	0.000
<i>unemp_{it}</i>	18.61	0.000

Note: Null hypothesis states that series are cross-section independent. $CD \sim N(0,1)$ under H_0 .

TABLE 3

Panel unit Root Tests

	CADF(4)^c	CADF(4)^T	r_c	MQ_c	r_f	MQ_f	Idiosync. ADF
<i>r_comp_{it}</i>	0.497 (0.690)	3.163 (0.999)	6	-46.062	6	-44.375	0.867 (0.807)
<i>reer_{it}</i>	-1.573 (0.058)	-2.390 (0.008)	6	-53.960	6	-54.128	0.353 (0.638)
<i>prod_{it}</i>	0.065 (0.526)	3.692 (1.000)	6	-41.309	6	-40.407	1.008 (0.843)
<i>unemp_{it}</i>	0.275 (0.608)	1.853 (0.968)	6	-46.841	6	-48.955	-2.164 (0.015)

P-values in parenthesis. Critical Values for the MQ statistic are tabulated by Bai and Ng (2004), Table I. r_c is the number of common factors in MQ_c; r_f is the number of common factors in MQ_f. The last column represents the unit root test on the idiosyncratic component as in Bai and Ng (2004).

TABLE 4

Banerjee and Carrion (2013) panel cointegration tests

$f(prod_{ijt}, reer_{ijt}, unemp_{ijt})$			
Model	Z_j^*	AIC	BIC
1	-6.85	-8.71	-8.55
2	-5.82	-8.78	-8.58
3	-4.72	-8.96	-8.67
4	-6.19	-8.53	-8.21
5	-5.41	-8.68	-8.31
6	-4.25	-8.99	-8.58

Note: Critical values of the Z_j^* are -2.824, -2.113 and -1.759 at 1%, 5% and 10% significance levels, respectively, for the model with constant, whereas -2.924, -2.240 and -1.835 are their equivalents in the model with trend.

TABLE 5

Estimation of the long run parameters – Base Long-Run Model

a. Model 3 – Break in constant and trend

	CUP - FM	CUP - BC
$prod_{it}$	0.660*** (3.074)	0.620*** (2.863)
$reer_{it}$	-0.093*** (4.864)	-0.091*** (4.715)
$unemp_{it}$	-0.065 (1.554)	0.056 (1.345)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, *: at 10% .

b. Model 6 – Regime Break

	CUP - FM	CUP - BC
$prod_{it}$	0.578** (2.779)	0.572** (2.713)
$reer_{it}$	-0.067*** (3.141)	-0.063*** (2.953)
$unemp_{it}$	-0.198*** (2.986)	-0.216*** (3.245)
$prod_{it} \times D0402s$	-0.023 (1.142)	-0.021 (1.058)
$reer_{it} \times D0402s$	0.042 (0.924)	0.037 (0.820)
$unemp_{it} \times D0402s$	0.415*** (5.275)	0.412*** (5.206)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, *: at 10% .

TABLE 6

Estimation of the long run parameters –Long-Run Model with institutional variables

a. Model 3 – Break in constant and trend

	Model with WCOORD		Model with UD	
	CUP - FM	CUP - BC	CUP - FM	CUP - BC
$prod_{it}$	1.093*** (4.183)	1.063*** (4.070)	1.296*** (3.667)	1.217*** (3.441)
$reer_{it}$	-0.087*** (5.698)	-0.088*** (5.750)	-0.098*** (4.719)	0.097*** (4.696)
$unemp_{it}$	-0.124** (2.726)	-0.114** (2.507)	0.001 (0.160)	0.000 (0.089)
EPL_{it}	0.017*** (5.837)	0.017*** (5.746)	0.021*** (5.146)	0.021*** (5.082)
$WCOORD_{it}$	-0.006*** (4.092)	-0.006*** (4.008)		
UD_{it}			-0.248*** (3.282)	-0.239*** (3.165)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, * : at 10% .

b. Model 6 – Regime Break

	Model with WCOORD		Model with UD	
	CUP - FM	CUP - BC	CUP - FM	CUP - BC
$prod_{it}$	1.331*** (4.796)	1.325*** (4.775)	1.265*** (3.420)	1.319*** (3.566)
$reer_{it}$	-0.056*** (3.452)	-0.055*** (3.410)	-0.061** (2.887)	-0.059** (2.789)
$unemp_{it}$	-0.251*** (4.899)	-0.253*** (4.942)	-0.135* (1.857)	-0.141* (1.931)
EPL_{it}	0.010*** (3.308)	0.011*** (3.399)	0.014*** (3.245)	0.014*** (3.432)
$WCOORD_{it}$	-0.003 (1.552)	-0.003* (1.739)		
UD_{it}			-0.307** (4.244)	-0.300*** (4.321)
$prod_{it} \times D0404s$	-0.036 (1.377)	-0.033 (1.236)	-0.114** (2.348)	-0.092* (1.930)
$reer_{it} \times D0404s$	-0.124*** (2.914)	-0.118** (2.775)	-0.137** (2.383)	-0.139** (2.425)
$unemp_{it} \times D0404s$	0.397*** (5.239)	0.395*** (5.211)	0.329*** (2.971)	0.334*** (3.029)
$EPL_{it} \times D0404s$	-0.036*** (7.880)	-0.033*** (7.202)	-0.012** (2.420)	-0.013** (2.687)
$WCOORD_{it} \times D0404s$	-0.014*** (5.651)	-0.013*** (5.101)		
$UD_{it} \times D0404s$			0.054*** (3.063)	0.049** (2.761)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, * : at 10% .

TABLE 7

The estimated effect of reforms of EPL on real wages

Country	Reform	Estimated long-run effect on real wages
Austria	"Abfertigung Neu", 2002	-0.48% / -0.59%
Belgium	Fixed-Term contracts and Temporary work agencies reform, 1997	-1.70% / -2.10%
Germany	"Hartz reforms", 2003	-0.60% / -0.74%
Ireland	Fixed term contracts reform (maximum number of renewals), 2003	+0.31% / +0.38%
Italy	"Legge Treu", 1997	-0.97% / -1.20%
Netherlands	"Wet Flexibiliteit en Zekerheid", 1998	-1.20% / -1.50%
Portugal	Reform of fair dismissals and collective dismissals	-0.99% / -1.21%
Spain	Ley 10/1994	-1.40% / -1.70%

TABLE 8

Summary Statistics: Core vs. Periphery

	<i>Core</i>		<i>Periphery</i>	
	N. Obs.	mean	N. Obs.	mean
<i>comp_p</i>	384	3.22	320	2.56
<i>unemp</i>	384	6.47	320	9.00
<i>prod</i>	384	1.52	320	1.43
<i>reer</i>	384	98.99	320	96.67
<i>UD</i>	368	41.50	304	24.61
<i>EPL</i>	288	2.23	288	2.61
<i>WCOORD</i>	384	3.67	320	3.41

TABLE 9

a. Estimation of the long run parameters – Core (with Germany)

	No inst. Variables		Model with WCOORD		Model with UD	
	CUP-FM	CUP-BC	CUP-FM	CUP-BC	CUP-FM	CUP-BC
$prod_{it}$	-0.891*** (-4.510)	-0.724*** (-3.680)	-0.680*** (-3.624)	-0.571*** (-3.021)	-0.782*** (-4.141)	-0.675*** (-3.530)
$reer_{it}$	0.123*** (5.056)	0.113*** (4.660)	0.108*** (4.532)	0.105*** (4.344)	0.109*** (4.614)	0.102*** (4.268)
$unemp_{it}$	-0.122* (-1.831)	-0.110 (-1.628)	-0.014 (-0.211)	-0.244*** (-0.359)	0.088 (1.224)	0.078 (1.072)
EPL_{it}			-0.009*** (-5.162)	0.004*** (2.410)	0.010*** (5.589)	0.004** (2.520)
$WCOORD_{it}$			-0.006** (-3.483)	-0.007* (-3.717)		
UD_{it}					-0.248*** (-4.528)	-0.239*** (-4.340)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, *: at 10% .

b. Estimation of the long run parameters – Core (without Germany)

	No inst. Variables		Model with WCOORD		Model with UD	
	CUP-FM	CUP-BC	CUP-FM	CUP-BC	CUP-FM	CUP-BC
$prod_{it}$	-0.533*** (-2.992)	0.772*** (4.125)	-0.639*** (-4.100)	0.259 (1.591)	-1.542*** (-11.630)	-0.146 (-0.793)
$reer_{it}$	0.134*** (4.847)	0.096*** (3.534)	0.153*** (5.802)	0.117*** (4.481)	0.206*** (7.500)	0.115*** (4.172)
$unemp_{it}$	-0.358*** (-5.052)	-0.338*** (-4.404)	-0.315*** (-4.599)	-0.161** (-2.151)	-1.176*** (-20.207)	-0.276*** (-3.420)
EPL_{it}			0.019*** (10.286)	0.001 (0.787)	0.024*** (12.086)	0.003 (1.414)
$WCOORD_{it}$			-0.010*** (5.905)	-0.014*** (-7.506)		
UD_{it}					0.293*** (5.633)	-0.059 (-1.150)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, *: at 10% .

c. Estimation of the long run parameters – Periphery

	No inst. Variables		Model with WCOORD		Model with UD	
	CUP-FM	CUP-BC	CUP-FM	CUP-BC	CUP-FM	CUP-BC
$prod_{it}$	1.889*** (4.930)	1.058*** (3.415)	3.378*** (16.902)	2.167*** (10.074)	2.474*** (15.013)	1.968*** (12.182)
$reer_{it}$	-0.084*** (-4.144)	-0.068*** (-3.296)	-0.435** (-2.184)	-0.060*** (-2.945)	-0.131*** (-9.622)	-0.156*** (-11.560)
$unemp_{it}$	0.046 (1.031)	-0.075 (1.535)	-0.013 (-0.274)	-0.161*** (-3.277)	0.400*** (10.384)	0.460*** (11.717)
EPL_{it}			0.038*** (13.731)	0.037*** (11.792)	0.055*** (32.646)	0.055*** (32.777)
$WCOORD_{it}$			-0.007*** (-4.990)	-0.001 (1.007)		
UD_{it}					-1.059*** (-33.682)	-1.151*** (-37.190)
$prod_{it} \times D0404s$	0.387*** (20.934)	0.280*** (13.878)	0.247*** (7.271)	0.278*** (8.226)	0.037 (1.616)	0.047** (2.087)
$reer_{it} \times D0404s$	0.135*** (3.545)	0.048 (1.233)	0.220*** (6.180)	0.110*** (3.089)	0.121*** (4.004)	0.077*** (2.580)
$unemp_{it} \times D0404s$	-0.154*** (-2.819)	-0.036 (-0.620)	-0.255*** (-4.759)	-0.198*** (-3.680)	0.056 (1.407)	0.034 (0.860)
$EPL_{it} \times D0404s$			-0.020*** (-3.844)	-0.034*** (-12.797)	0.003 (1.471)	0.007*** (3.566)
$WCOORD_{it} \times D0404s$			-0.005*** (-2.870)	-0.020*** (-11.402)		
$UD_{it} \times D0404s$					0.126*** (10.490)	0.135*** (11.262)

Absolute t-values in parenthesis. All models are estimated with 2 factors, as suggested by PCA. The bandwidth was chosen using Silverman's rule of thumb. *** : significant at 1%, ** : at 5%, * : at 10% .

FIGURE 1

Real compensation per employee in EMU-12 countries, 1995-2011

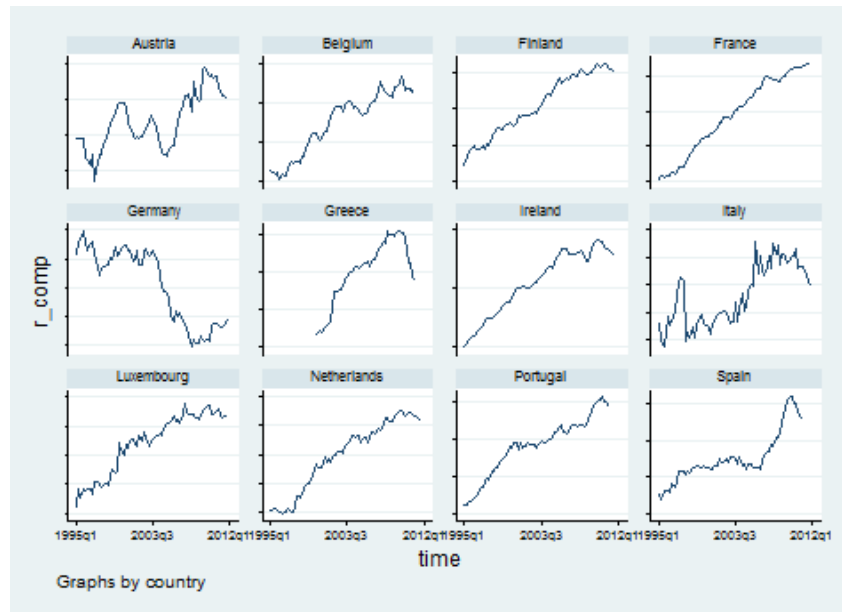


FIGURE 2

The Relationship between bargaining coordination and the real wage

