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Julian Messina, Chiara Strozzi and Jarkko Turunen

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RECent: c/o Dipartimento di Economia Politica, Viale Berengario 51, I-41100 Modena, ITALY
Phone +39 059 2056856, Fax +39 059 2056947, E-mail: RECent@unimore.it

Real Wages over the Business Cycle:

OECD Evidence from the Time and Frequency Domains[1]

Julian Messina (University of Girona, FEDEA and IZA)[2]

Chiara Strozzi (University of Modena and Reggio Emilia and IZA)

Jarkko Turunen (European Central Bank)

Abstract:

We study differences in the adjustment of aggregate real wages in the manufacturing sector over the business cycle across OECD countries, combining results from different data and dynamic methods. Summary measures of cyclicality show genuine cross-country heterogeneity even after controlling for the impact of data and methods. We find that more open economies and countries with stronger unions tend to have less pro-cyclical (or more counter-cyclical) wages. We also find a positive correlation between the cyclicality of real wages and employment, suggesting that policy complementarities may influence the adjustment of both quantities and prices in the labour market.

Keywords: real wages, business cycle, dynamic correlation, labour market institutions

JEL codes: E32, J30, C10

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[2] Corresponding author: Julian Messina, Address: University of Girona, FCEE, Campus de Montilivi, 17071 Girona, Spain. Phone: (+34)972-418779. Fax: (+34)972-418032. Email: julian.messina@udg.edu

1. Introduction

Empirical evidence about the direction and the extent of the response of aggregate real wages to business cycle fluctuations is inconclusive. In particular, the available results differ according to the data and methods that have been used. For example, in their survey of the literature, Abraham and Haltiwanger (1995) find that typical discrepancies across measures relate to differences in the data used, such as the wage measure, deflator, business cycle indicator, data frequency, sample period and sectors covered; and to methods, such as the precise measure of co-movement and the extent to which dynamics of real wages and output are taken into account. As a result, little is known about true cross-country variation in the adjustment of real wages over the business cycle and its potential determinants.

We provide evidence of differences in the adjustment of aggregate real wages in the manufacturing sectors over the business cycle across a large sample of OECD countries covering a time period of more than 40 years, starting from the 1960s. Our paper contributes to the empirical knowledge about aggregate real wage cyclicality in several ways. First, we use cross-country data to evaluate qualitative conclusions emerging from survey evidence on real wage cyclicality. In particular, we analyze the importance of differences in data and methods in determining cross-country differences in measured real wage cyclicality. We evaluate three dimensions that have been found to be important in previous literature: the deflators used to construct real wages, the measure of the business cycle and the methodology used to measure cyclicality.

Second, we provide a first systematic cross-country evidence of real wage cyclicality using empirical approaches that properly take into account the dynamic nature of the aggregate time-series under consideration. Most studies in this literature have measured co-movement between real wages and the cycle using a static approach.[3] However, a number of authors beginning with Neftci (1978) have stressed that accounting for the dynamic properties of the data series, such as persistence over time, may matter for correctly understanding real wage cyclicality.[4] The dynamic properties of the data can indeed vary substantially across data series and countries, and as shown in Den Haan (2000) evidence on cyclicality based on simple static measures can be misleading. Den Haan argues that the measured cyclicality of prices depends on whether co-movement is measured over the short or the long run. A priori, this dimension is potentially even more relevant for measuring real wage cyclicality, as nominal wage contracts tend to be fixed for an extended time period. When nominal wages are rigid in the short-term, measured cyclicality of real wages in the short run is likely to be dominated by changes in the deflator as adjustment through the wage-setting process becomes evident only with a lag. We use two dynamic approaches: the time domain approach proposed by Den Haan (2000) and the frequency domain approach proposed by Croux *et al.* (2001). In addition to properly taking into account the dynamics of the data series these methods also allow us to evaluate different business cycle horizons as an additional dimension that may result in variation across countries. Few studies have used these methods to study real wage cyclicality so far. The exception is the short note by Den Haan and Sumner (2002), where real wages appear pro-cyclical in the G7 countries and more so at longer horizons. In a related paper Lamo et al. (2007) focus on the

[3] The chosen cyclicality measure has been either the unconditional correlation coefficient between the cyclical component of real wages and an indicator of the cycle or the coefficient of OLS regressions of a (de-trended) real wage series on a (de-trended) business cycle series. In both cases, only the contemporaneous values of real wages and the cycle have been taken into account.

[4] Within this branch of literature, most contributions have adopted distributed lag or VAR models to focus on the dynamic response of real wages to business cycle indicators, or have used larger structural VAR models with identifying restrictions to study the reaction of real wages to different types of shocks. A review of the available empirical literature on aggregate real wage cyclicality can be found in Messina et al. (2006).

cyclicality of consumption, compensation and employment in the public sector. Camacho *et al.* (2006) apply the methods used here to measure business cycle co-movement in European countries.

Finally, we show that a measure of real wage cyclicality that is clean of systematic differences in data and methods differs across countries in a meaningful way. Cross-country results of real wage cyclicality in European countries based on static correlation analysis of detrended series can also be found in Christodoulakis *et al.* (1995). We extend these cross-country analyses by using dynamic methods, measuring the impact of different data and methods, and documenting country heterogeneity after controlling for these differences.

Our study does not address two potentially important factors that could lead to cross-country differences in the adjustment of real wages over the business cycle: the composition of shocks and changes in the composition of the labour force. First, the adjustment of real wages over the business cycle is likely to depend on the nature of the shock, with supply shocks leading to predominantly pro-cyclical and demand shocks leading to counter-cyclical responses. We cannot exclude that the measures of real wage cyclicality across countries that we derive are affected by country specific shocks. However, as average real wage cyclicality is here measured over an extended time period, the impact of country specific shocks is likely to be less important. Instead, owing to institutional diversity, differences in real wage cyclicality across countries are more likely to reflect differences in the labour market response to common shocks. Our robustness analysis using two different subsamples supports this view. Second, following Solon *et al.* (1994) a number of studies based on micro data have found that changes in the composition of the labour force over the business cycle are important and may lead to lower (less pro-cyclical) aggregate estimates of real wage cyclicality. In addition, the extent of real wage adjustment may vary across other disaggregated dimensions, such as regions, firm type (e.g. firm size and ownership). The homogeneity of the manufacturing sector across countries along these dimensions is likely to mitigate the potential impact of composition effects on cross country comparisons of real wage cyclicality. Further, we argue that measuring real wage cyclicality at the macroeconomic level remains important for understanding the aggregate business cycle facts. In the absence of micro data that are both sufficiently comparable across countries and cover long time periods, cross-country comparisons of real wage cyclicality are only possible using macroeconomic data.

Our findings suggest that data and methods indeed matter for observed real wage cyclicality, thus confirming previous survey evidence. Among the several dimensions that we test, differences in the type of deflators used result in largest and most robust differences across measures. In particular, real wage cyclicality is significantly more negative (more countercyclical) when the wage measure is deflated using producer prices, as opposed to the other deflators. While the use of dynamic methods is likely to provide more accurate measures of real wage cyclicality than static measures, we find that whether co-movement is measured in the short or the long run is not an important determinant of differences in real wage cyclicality across countries. Furthermore, even after controlling for differences in data and methods, country differences in real wage cyclicality remain important. Summary measures and cluster analysis point clearly to grouping of the countries into three groups: countries with mainly pro-cyclical real wages, countries with mainly counter-cyclical real wages and countries with either a-cyclical real wages or with very different patterns of cyclicality across deflators. These results indicate a more complex grouping of countries than a basic categorisation of countries to Continental European, Anglo-Saxon and Nordic labour market types would suggest. In particular, more open economies tend to show counter-cyclical wages. Moreover, our evidence points to a positive correlation between the cyclicality of real wages and the cyclicality of employment, suggesting that policy complementarities may influence the adjustment of both quantities and prices in the labour market. An exploration of possible structural determinants points to few robust associations between real wage cyclicality and measures of labour and product market indicators.

The rest of the paper is organized as follows. In section 2 we describe our dataset. In section 3 we introduce the methodology we use in our empirical investigation. In section 4 we present our evidence on real wage cyclicality and systematically evaluate the importance of different data and method dimensions on the results. In section 5 we discuss the summary measures of wage cyclicality obtained and relate them to measures of wage rigidity and employment cyclicality. Finally, we conclude in section 6.

2. Data

Our sample consists of 18 OECD countries and includes eleven continental European economies (Austria, Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Spain and Sweden), six Anglo-Saxon countries (Australia, Canada, Ireland, New Zealand, UK and the US) and Japan. [5]

We focus on data for the manufacturing sector. This focus has both advantages and disadvantages. A key advantage is that given limitations of data on the services sector, our focus on manufacturing allows us to benefit from long, high frequency data series that are of good quality. An important disadvantage is that manufacturing is a small and declining part of the overall economy in most OECD countries, limiting the scope of our results. However, manufacturing production tends to vary more over the business cycle than services, and as a result the measures of real wage cyclicality we derive may represent upper bounds for the whole economy. For evidence using compensation per employee for the whole economy, see Messina et al. (2006).

The variables we consider are: nominal wages/earnings in manufacturing, Consumer Price Index (CPI) deflator, Gross Domestic Product (GDP) deflator, Producer Price Index (PPI) deflator, manufacturing employment and industrial production. The data frequency is quarterly. Data are available, depending on country and indicator, at most from the 1960s and at least from the early

[5] Data for Germany before unification has been constructed using growth rates for West Germany.

1980s to 2004 (see Table A1 in Appendix for details). All series have been seasonally adjusted following the X12 procedure. The source of the data is the OECD Main Economic Indicators database. Summary statistics point to substantial persistence over time and large variation in the dynamic properties of the data across countries. Table 1 shows the standard deviation and the first-order autocorrelation of band-pass filtered series.[6] All series are highly persistent, as shown by the average (across countries) first-order autocorrelation, which is consistently above 0.9. Volatility of nominal wages, as measured by the average standard deviation is comparable to that of the price series, except for the more volatile PPI series. As a result, real wages deflated by PPI deflator appear almost twice as volatile as real wages derived using the GDP deflator or the CPI deflator. At the same time, the maximum standard deviation of nominal and real wage measures is nearly five times larger than the minimum, suggesting that there are substantial differences across countries. Focussing on the business cycle indicators, industrial production is on average more volatile than employment. As expected, these indicators of the business cycle tend to be more volatile than nominal or real wages (with exception of the PPI deflated real wage measure).

3. Empirical methodology

We construct measures of real wage cyclicality that vary in four different dimensions: the deflator, the measure of the business cycle, the method, and the horizon at which co-movement is measured.

The three different deflators we consider are the CPI deflator, the GDP deflator and the PPI deflator. These deflators are distinct in terms of the concept that they measure. Roughly speaking, the PPI refers to the prices at the factory gate. As a result, the PPI is strongly influenced by the evolution of input prices such as raw materials, as well as the cost of capital and labour. The GDP deflator at market prices measures the price of domestic value-added (including indirect taxes) and is based on

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[6] We use the Baxter-King band-pass filter, with a frequency band of 1.5 years to 8 years per cycle.

national accounts data. By definition it refers to prices of a broader set of goods than just manufacturing products, including the prices of domestically produced private and public services. Finally the CPI is based on a representative basket of final consumption goods. This includes imported consumption goods. Beyond the precise definition of the price index in terms of goods covered, an important difference across the deflators thus refers to the extent that they reflect mark-ups (prices over input costs) at different stages of production. In particular, compared to the PPI, the CPI is likely to be most influenced by the cyclical evolution of mark-ups at later stages of the production and distribution chains. Results in Table 1 suggest that the cyclical component of consumer price inflation is substantially less volatile (as measured by the standard deviation) and more persistent (as measured by first order autocorrelation) than the PPI. Furthermore, on average the standard deviation of the GDP deflator and the CPI is of the same magnitude as the standard deviation of nominal hourly wages (not shown).

The two indicators of the business cycle we consider are industrial production and employment, as is standard in the literature on real wage cyclicality.

The two methods we adopt are the time domain approach proposed by Den Haan (2000) and the frequency domain approach proposed by Croux *et al.* (2001).[7] The time domain measure proposed by Den Haan (2000) is based on the degree of co-movement between VAR forecast errors at different horizons. This approach takes into account the dynamic nature of the macroeconomic data series under consideration through the inclusion of lagged variables in the VAR. Furthermore, the methodology can accommodate both stationary and non-stationary variables and thus does not require additional filtering. Den Haan considers a two variable VAR model in standard form:

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[7] To derive the time domain estimates we rely on the Matlab codes available at Wouter Den Haan's homepage (http://weber.ucsd.edu/~wdenhaan/soft.html), while to derive the frequency domain estimates we use a version of the Matlab codes available at Mario Forni's homepage (http://www.economia.unimore.it/forni_mario/matlab.htm), modified to generate bootstrapped confidence bands.

$$\begin{bmatrix} x_t \\ y_t \end{bmatrix} = A_0 + \sum_{i=1}^{m} A_i \begin{bmatrix} x_{t-i} \\ y_{t-i} \end{bmatrix} + \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} \quad (1)$$

where x_t and y_t are two random variables; A_0 is a 2 × 1 vector of constant terms or a matrix of deterministic coefficients; A_i are 2 × 2 matrices of coefficients; v_{1t} and v_{2t} are two error terms which are assumed to be serially uncorrelated but that can be correlated with each other, and *m* is the total number of lags included in the model. After estimating the model it is straightforward to calculate the *k*-period ahead forecasts errors for the two variables (see Den Haan, 2000 for a detailed description). In our empirical analysis, the Den Haan methodology is applied by estimating a number of bivariate VAR models where *x* and *y* refer to real wages and a measure of the business cycle. The VAR is estimated in terms of first log differences with the lag length and the deterministic components chosen by the Schwarz Information Criterion allowing for a maximum number of 8 lags. Bootstrapped standard errors based on 1000 replications were used to construct 90 percent confidence bands and standard errors.

The measure proposed by Croux *et al.* (2001) is a correlation defined within the frequency domain. Consider two zero-mean real stochastic processes, *x* and *y* and let $S_x(\lambda)$ and $S_y(\lambda), -\pi \leq \lambda \leq -\pi$, be the spectral-density functions of *x* and *y* and $C_{xy}(\lambda)$ the co-spectrum. Then the dynamic correlation between *x* and *y* at frequency λ is given by:

$$\rho_{xy}(\lambda) = \frac{C_{xy}(\lambda)}{\sqrt{S_x(\lambda) S_y(\lambda)}} \quad (2)$$

This measure is a real number that takes values between -1 and 1 and allows computing the correlation between two series for each band of frequencies. Consistent with the time domain approach, dynamic correlations were calculated using data in first log differences. Following Croux *et al.* (2001), we applied a standard block bootstrap technique to construct 90 percent confidence bands and standard errors. The number of replications was set at 1000 and the length of the blocks was chosen to equal 12 quarters.

It should be noted that both the time domain and the frequency domain measures have a specific relationship with static unconditional correlation. In the case of the time domain, the correlation coefficient of the forecast errors of a bivariate VAR will converge to the static unconditional correlation coefficient of the two series as the forecast horizon goes to infinity. In the frequency domain case, the dynamic correlation between two processes over a frequency band is identical to the static correlation of the same processes after suitable pre-filtering (e.g. using the band-pass filter). For a given frequency interval $[0, \pi]$, the static unconditional correlation is then the simple mean of the dynamic correlation over that interval.

The measures of co-movement stemming from the time and frequency domain methods that we adopt improve upon static correlations derived from cyclical components of economic time series that have been filtered using standard methods. These methods include, within the time domain the Hodrick-Prescott (HP) filter, and within the frequency domain, the Baxter-King (BK) bandpass filter. Both filters extract a business cycle component of the economic time series by removing low frequency trends (and for the BK band-pass filter also high frequency noise). Both standard methods present some problems. The HP filter is subject to an arbitrary choice of the smoothing parameter λ. The BK filter is only an approximation of the ideal band-pass filter since, like all moving-average smoothers, a number of observations at the beginning and the end of the sample are lost in its computation. More generally, alternative filters extract different types of information from the original series and, as a result, conclusions drawn from detrended data tend to vary widely across detrending methods (Canova, 1998).

The dynamic methods we adopt overcome these disadvantages of the standard methods. In addition, within the time domain, the Den Haan method accommodates both stationary and non-stationary processes and does not require identifying assumptions that are usually needed for VAR decompositions. In addition, the use of dynamic methods allows us to add a further dimension to our measures of real wage cyclicality: the horizon at which co-movement is measured. This has been shown to be a crucial element when evaluating co-movement of macroeconomic variables in different contexts (see Den Haan, 2000, Camacho et al., 2006 and Nath, 2004). The horizon could indeed matter for real wage cyclicality for a number of reasons. For example, as argued in the introduction, it is possible that the longer-term nature of nominal wage contracts allows wages to adjust to changes in business cycle conditions only with a significant lag.

We consider three business cycle horizons. In the time domain method, we look at 1.5, 4 and 8 year ahead forecast errors, while in the frequency domain method we look at the frequencies between 0 and $\pi/3$, 0 and $\pi/8$ and 0 and $\pi/16$ (corresponding to 1.5, 4 and 8 years periodicities in the time domain). We define these horizons as short, medium and long-run business cycle horizons. These horizons are in line with the findings of both the NBER and CEPR business cycle dating committees about the length of business cycles in the United States and the euro area respectively.[8]

As illustrative examples, the measures of co-movement and their 90% confidence bands are shown in Figures 1 and 2 for two countries in our sample: Spain and the United States. In both cases the measures are based on real wages deflated by the GDP deflator and IP as the business cycle indicator. The figures show that both dynamic methodologies are consistent and suggest that real wages are countercyclical in Spain and pro-cyclical in the United States. Interestingly, beyond the very first quarters, the horizon at which co-movement is measured does not appear to matter for the cyclicality of real wages in these two countries. In the time domain case, the short run correlation for Spain (United States) is -0.318 (0.521), the medium run correlation is -0.236 (0.515) and the long run correlation is -0.235 (0.503). In the frequency domain case, the short run correlation for Spain (United States) is -0.250 (0.472), the medium run correlation is -0.278 (0.491) and the long run correlation is -0.285 (0.495). We have checked that the limit of the time domain measure corresponds to the average of the frequency domain measure, confirming what theory predicts.

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[8] Details of these findings are available on the NBER and CEPR websites.

In sum, the variation across data and methods allows us to calculate altogether 36 different measures of real wage cyclicality for each of the 18 countries in our sample, resulting in a total of 648 observations. The 36 different measures of correlation between real wages and the business cycle are derived calculating real wage cyclicality along the four dimensions we consider: deflator (3), measure of the business cycle (2), method (2), and business cycle horizon (3).

4. Real wage cyclicality across data, methods and countries

In order to understand whether differences in data and methods are relevant in measuring real wage cyclicality, we first study kernel distributions of our real wage cyclicality measures along the four relevant dimensions. This first step is useful in determining simple patterns in our data without having to resort to parametric assumptions and/or specific assumptions about the type of relationship between different ways of measuring real wage cyclicality (e.g. linearity).

Before analysing our correlation measures we transform them using the Fisher transformation. This transformation removes the constraint that the correlation coefficient has to lie between -1 and 1, thus bringing the distributions of correlation coefficients closer to the normal distribution and allowing for statistical inference. Specifically, the transformation that we apply is given by:

$$z = \frac{1}{2}\log(\frac{1+r}{1-r}) \tag{3}$$

where r is the correlation coefficient. The transformed variable z is used in the analysis.

The distributions of our Fisher transformed correlation measures are shown in Figure 3. We use Gaussian kernels to smooth the histograms, choosing the optimal bandwidth that minimises the asymptotic mean integrated squared error.

These results show that the deflator used to construct the measures of real wage matters. There seems to be a clear ranking between the three deflators used: wages deflated by the PPI are most counter-cyclical and CPI deflated wages are most pro-cyclical. Correlations derived from GDP deflated wages lie between these two and are centred around zero. These differences in the distributions are confirmed by a Kolmogorov-Smirnov (KS) test of equality of distributions. The KS test of equality of PPI and GDP deflated wages yields a value of 0.254 (with a p-value less than 0.01), while the same test contrasting GDP and CPI deflated wages also rejects the null of equality of distributions (value: 0.4074; p-value less than 0.01). This finding confirms and extends previous evidence summarised in Abraham and Haltiwanger (1995). The differences in the cyclicality of real producer (PPI deflated) and consumer (CPI deflated) wages thus points to a possible role for imperfections in the goods market in determining differences in real wage cyclicality across countries.[9]

In contrast to the differences across deflators, other dimensions in the data and methods have a much smaller impact on the distribution of real wage cyclicality measures. The forecast horizon or frequency window does not seem to vary across the different dimensions we study. The KS test fail to reject the null of equality of short and medium run distributions (p-value: 0.717) and medium vs. long run distributions (p-value: 0.998). Similarly, the choice of the cycle measure (employment vs. IP) has little effects on the measurement of cyclicality (p-value of KS test: 0.352). The distribution of correlations generated using the frequency domain method is somewhat more peaked around zero

[9] These results may also relate to the impact of different compositional effects across sectors on price inflation. First, it seems reasonable to expect a closer relationship between producer price inflation and business cycle measures that are specific to the manufacturing sector, than inflation in the economy as a whole. In particular, overall CPI is likely to reflect also fluctuations at the business cycle frequency that are specific to the services sector. Second, lower cyclicality of consumer prices is also in line with empirical evidence that consumer prices are adjusted somewhat less frequently than producer prices, possibly in part because of the higher labour intensity in the services sector (see Alvarez et al. 2006).

than the distribution of measures from the time-domain methodology. This difference in distributions is confirmed to be statistically significant by the KS test (p-value: 0.003).

Regression analysis allows us to go a step further and test for the impact of differences across data and methods by taking into account measurement uncertainty in the construction of real wage cyclicality measures. Moreover, it also allows us to determine whether differences across countries are statistically significant once differences in data and methods are controlled for. We perform our regression analysis by pooling our Fisher transformed estimates of wage cyclicality and estimate different specifications following the general model:

$$c = \beta X + w \tag{4}$$

where X is a (T x K) matrix of exogenous variables including indicator variables for data characteristics, methods and countries, c is a (T x 1) vector of estimates of real wage cyclicality pooling countries along the four dimensions of real wage cyclicality of our data, and w is a (T x 1) vector of stochastic elements that meets all the usual Gauss-Markov assumptions. A problem arises when estimating equation (4), since c is not observed. Instead we observe an unbiased estimate $\hat{c}$, such that $\hat{c} = c + u$ and u is a vector of sampling errors from the first stage dynamic estimation techniques. The problem is such that even if $\mathrm{E}(ww') = \sigma^2 I$ the error covariance for this second stage regression will not be homoskedastic, and hence OLS estimates will be inefficient. Hanushek (1974) shows that, under certain regularity conditions, the variance of the first stage estimates provides sufficient information to construct Feasible General Least Squares estimates of the second stage parameters that are asymptotically efficient. Accordingly, the standard errors of the estimates we report below are adjusted to take into account the measurement uncertainty associated with the dependent variable.

Table 2 shows the regression results following expression 4. It should be noted that, after running the regressions with the z-Fisher transformed coefficients, we undo the transformation to present in the Tables standard correlation coefficients and summary measures of wage cyclicality. Specifically, we recover the correlation coefficients by applying the z to r transformation:

$$r = \left(e^{z} - e^{-z}\right) \Big/ \left(e^{z} + e^{-z}\right) \tag{5}$$

Since this transformation is not linear we approximate the standard errors of the coefficients using the delta method. The regression analysis confirms that among the four dimensions of our measures of cyclicality, the type of deflators used result in by far the largest statistically significant differences. In particular, real wage cyclicality is significantly more negative (more countercyclical) when measured using the PPI deflator, as opposed to the other deflators. In addition, in three out of four regressions the results show that the time domain method produces somewhat more countercyclical results than the frequency domain method. When country dummies are not included, other dimensions are also weakly statistically significant (see columns 1 and 3). However, these results are not robust to the inclusion of country dummies in the regression (see columns 2 and 4). The country dummies (not shown in the Table) are always jointly significant, even after controlling for differences across data and methods. We also ran regressions with interaction terms between the data and method dummies and found that they were in all cases not statistically significant (see columns 3 and 4). Therefore we drop them in the subsequent analysis.

The country dummies from column (2) in Table 2 are shown in the first column of Table 3. Most country dummies are large compared to the direct effects of data and methods shown above and highly statistically significant. In order to complete the test for the influence of data and methods on differences across countries in measured real wage cyclicality, we sequentially add interaction effects between the set of country dummies and the four different data and methods dimensions. These regressions are shown in columns (2) to (5) of Table 3. The results show that differences in the business cycle horizon can be ignored as a determinant of measured real wage cyclicality (see column 2). Most of the real wage adjustment appears to take place within the first year and a half of the initial business cycle impulse. This fact, together with the relatively mild responses of wages to the business cycle even at longer-term business cycle frequencies suggests substantial stickiness in the wage setting process across the board. Regarding the measure of the business cycle, we find that the interactions between the IP dummy and country effects are jointly statistically significant (see column 4). This contrasts with the non-significant average effect of the business cycle dummy obtained in Table 2, suggesting that in some countries measured real wage cyclicality is higher when industrial production rather than employment is used as the measure of the business cycle, while the opposite occurs in other countries. Both effects seem to compensate each other when we average across countries. Confirming previous results, the interactions with the dummies for the different deflators are also highly significant (see column 5).

We conclude that as regards the impact of data and methods, the methodology used, the deflator and the measure of the business cycle do matter for understanding measured real wage cyclicality. However, the only sizeable effect that is consistent across countries relates to the different deflators. In particular, when the dummy variables for the business cycle measure and the type of methodology are separately interacted with country dummies (columns 3 and 4), the country effects are very similar to our benchmark estimates when such interactions are excluded (column 1) and we fail to find a systematic pattern across countries along these different dimensions.[10] Instead, when the dummy variables for PPI and GDP deflators are interacted with country dummies (column 5) the country effects differ significantly from the benchmark case. In particular, and with the exception of New Zealand, summary country measures of CPI wages (column 5) are more pro-

[10] Note that the dummy variable for the business cycle measure takes the value of 1 when industrial production is used to measure the business cycle. Therefore, the country main effects in column 4 can be interpreted as summary measures of real wage cyclicality using employment as the measure of the business cycle. Similarly the dummy for method takes the value 1 when the time domain based method is used, and therefore the country main effects in column 3 refer to real wage cyclicality using the frequency domain based method.

cyclical than the country effects obtained when averaging across the three deflators (column 1), although differences for some particular countries are not statistically significant at standard levels of significance. Finally, note that even after data and method differences are controlled for, country differences remain important and statistically significant.

We turn next to investigate these country differences. We do this by constructing summary measures of real wage cyclicality for each country controlling for the impact of data and methods. We argue that looking at differences across different definitions of the real wage is economically meaningful. Since the results of our statistical analysis also suggest that differences across deflators are important, we separate the analysis of country effects between measures GDP, CPI and PPI deflated wages. In order to construct our summary indicators we run regressions following equation (4), controlling for the characteristics of the cyclicality measures for the three different sub-samples depending on the deflator. The summary country measures correspond to the coefficient of the country dummies in each of these regressions, where the constant term is excluded.

Figure 4 presents our three summary measures of real wage cyclicality using the different deflators. The difference between the cyclicality of consumer real wages (pro-cyclical in most countries) and producer real wages (counter-cyclical in most countries) is noticeable. Furthermore, the cyclicality of real wages measured using the GDP deflators is somewhere in between that of the consumer and producer wages, but closer to the former. The rank correlation between the cyclicality of consumer real wages and real wages deflated by the GDP deflator is substantially larger (0.88), than the equivalent correlation between cyclicality of consumer and producer real wages (0.38). As regards the country ranking, Germany and Japan are the only two countries that exhibit consistently pro-cyclical real wages over the business cycle irrespective of the deflators used. Real wages are also mainly pro-cyclical in the United Kingdom and the Unites States. At the other extreme, Ireland, Spain, Canada and New Zealand exhibit consistently counter-cyclical behaviour of real wages irrespective of the deflators used.[11] For some countries, such as France and Sweden, real wage cyclicality changes sign from pro-cyclical to counter-cyclical when moving from consumer to producer wages. This sign switch suggests that price mark-ups at the wholesale and retail levels have been more countercyclical in these countries.

Differences in real wage cyclicality across countries could reflect different prevalence of supply versus demand shocks over time. In particular, when wages are not fully flexible over the business cycle, pro-cyclicality of real wages could be the outcome of price declines (increases) as a result of dominant positive (negative) aggregate supply shocks. Counter-cyclicality of real wages could in turn reflect price changes in response to dominant demand shocks. Indeed, there is some empirical evidence that the type of shocks matters for the cyclicality of real wages. For example, Fleischman (1999) finds that real wages in the US tend to be counter-cyclical in response to labour supply and aggregate demand shocks, and pro-cyclical in responses to productivity and oil shocks. However, as average real wage cyclicality is here measured over an extended time period, the impact of different types of shocks is likely to be less important. Instead, we speculate that differences in real wage cyclicality across countries have more to do with diversity in institutional frameworks governing labour markets in the countries in our sample. In particular, differences in real wage cyclicality across countries are more likely to reflect differences in the labour market response to common shocks owing to differences in institutions.

To evaluate this dimension and to establish robustness of our results to differences in the composition of shocks over time, we have redone all of the analyses in the paper using a shorter sample, restricting the time period for all countries to begin from 1980. While in some cases the

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[11] An analysis of the cyclicality of nominal wages (not presented here for the sake of brevity) confirms that these results are driven by countercyclical nominal wages in these countries.

results for the 1980s onwards sample point to somewhat less pro-cyclical real wages, overall, the results are not substantially different from those for the whole sample.

5. Interpretation of the results

Cluster analysis of all the measures of real wage cyclicality is useful in further uncovering the underlying cross-country patterns in the different types of real wage cyclicality observed above. The algorithm underlying cluster analysis searches sequentially for the most similar pairs of countries in terms of real wage cyclicality. We focus here on hierarchical clustering, following the criteria proposed by Ward to construct clusters of countries.[12] Cluster analysis is performed on our sample of 648 measures of real wage cyclicality. The results are presented in terms of a tree diagram, where the height between clusters represents the dissimilarity between them.

The results indicate that the countries can be divided into three main groups (see Figure 5).[13] These groups of countries can be characterised as countries with mainly pro-cyclical real wages (Germany, Japan, the UK and the US), countries with mainly countercyclical real wages (Ireland, Spain, Canada and New Zealand) and the rest of the world "RoW" with very different patterns of cyclicality. These results indicate a more complex grouping of countries than suggested by a standard categorisation of countries to Continental European (sometimes further divided into Core and Mediterranean types), Anglo-Saxon and Nordic labour market types would suggest (see Sapir, 2006 and OECD, 2008). In addition, Germany, a continental European country with relatively strong unions and rigid labour market structures, is grouped together with Anglo-Saxon countries such as the UK and the US that have weaker unions and are usually considered more flexible. The consistent

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[12] Other criteria such as agglomerative complete, weighted and average linkage produced similar results.

[13] There is no definitive test on the optimal number of clusters. However, the Calinski and Harabasz pseudo-F index and the Duda and Hart Je(2)/Je(1) test (not shown in the paper) support the existence of three clusters, as visual inspection of the tree diagram seem to suggest.

pro-cyclicality of real wages in Germany, the UK and US is, however, in line with microeconomic evidence presented in Peng and Siebert (2007). The Anglo-Saxon group itself is strongly divided in terms of the extent of measured real wage cyclicality to countries with pro or counter-cyclical real wages. The RoW could be further divided into at least two groups, where one cluster is formed by countries with largely a-cyclical real wages (including most Nordic countries with relatively strong labour unions, centralized bargaining and compressed wage structures) and those that exhibit pro-cyclical consumer wages but counter-cyclical producer wages (including other large Continental European countries, such as France, Italy and the Netherlands with strong unionization but less centralized bargaining and stronger employment protection legislation).

We have also evaluated the relationship between the different summary measures of wage cyclicality and key structural features and labour market institutions. Since we are going to be studying the role of labour market institutions and policies that might influence jointly the cyclicality of employment and wages, in the remainder of the paper we use a new set of summary measures of wage cyclicality where we exclude employment as an indicator of the cycle. This helps us in isolating the impact of each institution on the determination of wages.

An important indicator that helps to explain the country clustering that we previously highlighted is trade orientation. The graphs on the left hand side of Figure 6 show the association between producer and consumer wages and the average import penetration in the goods sector of each country during the sample period.[14] The countries grouped under the heading "mainly procyclical wages" (marked with a diamond) tend to show very low values of import penetration in the goods sector. The opposite is true for countries with strongly counter-cyclical wages. The

[14] See the Appendix for a definition of the variables used in the text. For the sake of clarity of exposition we exclude GDP deflated wages from the figure, since the follow pretty closely consumer wages and results for both variables are very similar. Very similar results, available upon request, are obtained with other indicators of trade openness such as the import penetration of goods and services, or total external trade as a share of GDP.

negative association between trade openness measured by import penetration and wage cyclicality is very similar in the cases of producer wages (correlation coefficient: -0.49; p-value: 0.04) and consumer wages (corr.: -0.46; p-value: 0.06). The graphs on the right show the correlation between wage cyclicality and a different measure of trade openness based on policies towards free trade. The picture is broadly consistent with that observed for trade openness, although the correlations are somewhat weaker (-0.36 and -0.35 for producer and consumer wages, respectively). This negative association between openness and wage cyclicality is supportive of the theoretical predictions in Aizenman (1985), who shows that wage indexation and hence real wage rigidity is more likely to arise in more open economies.

Is the aggregate cyclicality of real wages related to measures of real wage rigidity? It is not easy to find systematic cross-country measures of wage rigidity to compare with our estimates. The International Wage Flexibility Project has recently measured real wage rigidity at the individual rather than aggregate level in several OECD countries. We therefore compare our results with those in Dickens et al. (2008) and find evidence of a negative correlation between their measure of downward real wage rigidity and our summary measures of real wage cyclicality. The correlation is strongest for the cyclicality of GDP deflated wages at -0.50 (with a p-value of 0.06 for 14 countries), suggesting that countries with more rigidities in real wages at the individual level are less like to have pro-cyclical and more likely to have counter-cyclical real wages at the aggregate level. While this evidence is only tentative, we would not expect a perfect correlation between the two set of measures as they relate to different samples (e.g. in terms of sectors and time periods covered) and, more importantly, to fundamentally different (although related) concepts.

The counter-cyclicality of wages observed in many European countries in our sample can be rationalized by wage bargaining models (e.g. McDonald and Solow, 1981), where the insensitivity of real wages to shocks naturally leads to counter-cyclical movements of real wages. Indirectly, unionisation may also lead to more segmentation in the labour market, thus increasing the share of workers whose wages are insulated from changes in macroeconomic conditions (see Bertola, Blau and Kahn, 2007). We have explored the role of unionization in Figure 7. We use two different indicators of union bargaining power: the extent of union coverage and an index of legal strikes. Depending on the set of norms and regulations governing the labour market, there are large differences within countries between the coverage of wages negotiated by unions and the extent of unionisation (as measured by the number of union members in the total labour force).[15] For this reason, the indicator of union coverage is a better indicator than union density for evaluating the impact of unions in wage negotiations across countries. The index of legal strikes, obtained from Botero et al. (2004), measures the legal framework ruling strike rights and ranges from 0 to 1, increasing with the ease of organising legal strikes. We expect higher union bargaining power in those countries with generous strike legislation. Our measures of consumer and producer wages are negatively correlated with both measures of union bargaining power. The correlation is always significant at the 5% level in the case of producer wages (corr: -0.49 with union coverage and corr: -0.46 with the index of legal strikes) but only significant in the case of consumer wages when the indicator of unionization is legal strikes (corr: -0.60, p-value less than 0.01). Hence, our results suggest that wages tend to be less pro-cyclical (more counter-cyclical) in countries with stronger union bargaining power.

The results presented above are indicative of a negative association between unionization or trade openness and the cyclicality of wages. We have also explored the relationship between our measures of wage cyclicality with other institutional variables typically considered in cross-country studies of the determinants of employment or unemployment, such as the generosity and duration of unemployment benefits, the stringency of employment protection legislation, the extent of

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[15] Well-known examples are those of France and Spain where collective agreements cover more than three quarters of the labour force but less than 15 per cent of workers belong to a union.

coordination in wage setting and the degree of product market regulations.[16] In all cases we failed to find a statistically significant correlation. We speculate that this result may reflect important non-linearities and complementarities. For example, while centralisation and coordination in bargaining is expected to reduce the adjustment of real wages to business cycle conditions Calmfors and Driffill (1988) argue that this relationship may be non-linear and consequently impossible to detect within our handful of observations. Moreover, it is possible that labour market institutions are binding only when they operate in bundle with other policies. Again, this complementarity of institutions, which has been shown to be important for other macroeconomic outcomes (see e.g. Belot and Van Ours, 2004), cannot be easily detected with a limited number of country observations.

Our final exercise aims at shedding some light on the importance of complementarity by focussing on the relationship between cyclicality of real wages and cyclicality of employment. On the one hand, firm profit maximization suggests that if employment and hours worked can not be fully adjusted to shocks such as fluctuations in product demand (e.g. because of substantial adjustment costs), then wages per hour worked may need to adjust more within the firm. This implies that also at the aggregate level lower cyclicality of employment should increase the need for wages to adjust to macroeconomic conditions, thus leading to a negative association between the cyclicality of real wages and employment within countries. On the other hand, labour market institutions might complement each other and thus attenuate fluctuations in both wages and employment. For instance, the ability of employment protection policies to insure workers from labour market income risk (Bertola, 2004) would be seriously limited if wages were adjusted at will by firms. Bertola and Rogerson (1997) have argued this is the reason we observe strong unions and binding minimum wages limiting wage adjustments in countries with strong employment protection. Hence, bundles of

[16] See Nickell et al. (2005) for a discussion of these indicators and their impact on unemployment in OECD countries.

policies might limit both employment and wage adjustment, leading to a positive association between employment and wage cyclicality across countries.

To explore this dimension, we proceed to measure the cyclicality of employment following the same methodology used for wages. Hence, we derive measures of employment cyclicality in the time and frequency domain and at different business cycle horizons, with industrial production as the measure of the business cycle. We also construct a summary measure of the cyclicality of employment in each country using the regression approach that was used to construct summary measures of real wage cyclicality. As before, we drop from the analysis employment as an indicator of the cycle to construct measures of wage cyclicality in order to avoid a spurious correlation with our measures of employment cyclicality.

Using these summary measures we find that the cyclicality of employment and real wages are positively associated (see Figure 8).[17] More (less) pro-cyclical consumer and producer wages appear to go hand in hand with more (less) pro-cyclical employment. Relating these results to the country clustering found in the previous section, countries with mainly pro-cyclical real wages (Germany, Japan, the UK and the US) also exhibit clearly pro-cyclical employment. Other country clusters show more diversity. In particular, countries with mainly counter-cyclical real wages (Ireland, Spain, Canada and New Zealand) exhibit different degrees of employment cyclicality. The positive association between the cyclicality of real wages and employment appears robust. Regressing the cyclicality of consumer wages on the cyclicality of employment, we find that the coefficient on the cyclicality of employment is positive and statistically significant after controlling for data and method dummies (see Column 2 in Table 4). The coefficient of employment cyclicality remains positive and statistically significant also after adding cluster dummies (see Column 3 in Table 4) suggesting that the positive association between employment and wage cyclicality holds within

[17] Again, the results based on a shorter sample from the 1980s onwards are not substantially different.

clusters. We also obtain mainly positive estimates in the case of producer wages (except for column 6 when we control for clusters of countries), although the coefficients are not significant at standard levels. Overall, our findings provide some support to the hypothesis of complementarity of institutions simultaneously limiting employment and wage fluctuations.[18]

6. Summary

Little is known about true cross-country variation in the adjustment of real wages over the business cycle and its potential determinants. This is partly due to the heterogeneity in data characteristics and methods used in the various investigations. In this paper, we have contributed to the existing empirical literature by providing consistent comparative evidence of aggregate real wage cyclicality in the manufacturing sector for a large sample of OECD countries. We find that data and methods indeed matter for observed real wage cyclicality. However, among the different dimensions that we test, only differences in the type of deflators used result in robust differences that are consistent across countries. In particular, real wage cyclicality is significantly more negative (more countercyclical) when measured using the PPI deflator, as opposed to the other deflators. This evidence is in line with the evidence summarised in Abraham and Haltiwanger (1995) and points to an important role of the mark-up between producer and consumer prices in determining real wage cyclicality. Other dimensions, including the horizon at which cyclicality is measured are either not statistically significant or do not show similar effects across countries.

We find that country differences in real wage cyclicality remain important even after controlling for differences in data and methods. The evidence presented in this paper thus suggests

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[18] Note that country dummies are not included in this regression. Including country dummies would limit the variation that is available to explain the association between the cyclicality of real wages and employment to differences in data and methods. However, standard errors are robust to heteroskedasticity and allow for correlation of the summary measures of wage and employment cyclicality within countries.

that variation in existing estimates of real wage cyclicality is not only due to differences across data and methods used. Our summary measure of real wage cyclicality that is clean of differences in data and methods confirms that the cyclicality of real wages varies substantially across countries. We argue that this heterogeneity may reflect fundamental differences in the functioning of labour markets, rather than a different composition of demand and supply shocks. Summary measures and cluster analysis point to a possible grouping of the countries to three groups: countries with mainly pro-cyclical real wages, countries with mainly counter-cyclical real wages and countries with either a-cyclical real wages or with very different patterns of cyclicality depending on the definition of the wage variable. These results also indicate a more complex grouping of countries than a basic categorisation of countries to Continental European, Anglo-Saxon and Nordic labour market types would suggest. Two structural features, the extent of union bargaining power and the degree of trade openness help understanding these clusters. Our evidence suggests that more open economies and countries with stronger unions tend to have less pro-cyclical (or more counter-cyclical) wages. Finally, our evidence also points towards a positive association between real wage cyclicality and the cyclicality of employment, in line with the view that policy complementarities could play an important role in determining labour market outcomes.

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Table 1. Dynamic properties of data

	Hourly Wages				Cycle indicators		Deflators		
	Nominal	GDP defl	CPI defl	PPI defl	IP	E	GDP defl.	CPI	PPI
Standard deviation									
Average	0.014	0.012	0.011	0.021	0.025	0.017	0.015	0.014	0.024
Minimum	0.005	0.005	0.005	0.009	0.019	0.008	0.008	0.008	0.016
Maximum	0.024	0.028	0.017	0.051	0.036	0.029	0.028	0.020	0.048
Autocorrelation (AR1)									
Average	0.939	0.901	0.916	0.912	0.909	0.934	0.943	0.958	0.930
Minimum	0.856	0.805	0.882	0.837	0.875	0.906	0.898	0.941	0.903
Maximum	0.959	0.950	0.945	0.948	0.937	0.957	0.967	0.972	0.957

Note: All variables are cycle components extracted from the original log series using a Baxter-King band-pass filter (1.5-8 year frequency). The summary statistics reflect averages, minima and maxima across countries in our sample. IP refers to industrial production and E to employment in manufacturing.

Table 2. Determinants of real wage cyclicality: effects of data and methods

	(1)	*(2)*	*(3)*	*(4)*
Medium Run	-0.047	0.011	-0.089	0.005
	(2.25)*	(0.69)	(2.17)*	(0.13)
Long Run	-0.046	0.013	-0.091	0.003
	(2.19)*	(0.77)	(2.21)*	(0.09)
Time Domain	-0.072	-0.035	-0.117	-0.045
	(4.00)**	(2.58)**	(2.97)**	(1.36)
Industrial Production	-0.069	-0.032	-0.081	-0.006
	(3.87)**	(2.34)*	(2.09)*	(0.17)
GDP Deflator	0.144	0.203	0.069	0.166
	(7.03)**	(12.78)**	(1.67)	(4.62)**
CPI Deflator	0.203	0.258	0.140	0.231
	(10.06)**	(16.59)**	(3.47)**	(6.60)**
Time Domain * Medium Run			0.050	0.013
			(1.10)	(0.40)
Time Domain * Long Run			0.048	0.012
			(1.06)	(0.36)
Industrial Prod. * GDP Defl.			0.068	0.026
			(1.51)	(0.77)
Industrial Prod. * CPI Defl.			0.036	-0.004
			(0.80)	(0.13)
Industrial Prod. * Time Domain			-0.020	-0.046
			(0.53)	(1.68)
Industrial Prod. * Medium Run			0.022	-0.016
			(0.48)	(0.48)
Industrial Prod. * Long Run			0.019	-0.019
			(0.43)	(0.56)
GDP Defl. * Medium Run			0.064	0.007
			(1.18)	(0.18)
GDP Defl. * Long Run			0.069	0.013
			(1.28)	(0.31)
CPI Defl. * Medium Run			0.075	0.017
			(1.38)	(0.43)
CPI Defl. * Long Run			0.084	0.026
			(1.56)	(0.64)
GDP Defl. * Time Domain			0.078	0.040
			(1.72)	(1.21)
CPI Defl. * Time Domain			0.068	0.035
			(1.51)	(1.03)
Country Effects	No	Yes**	No	Yes**
Wald Test			1.81*	0.50
Adjusted R^2	0.14	0.58	0.16	0.58
Observations	648	648	648	648

Note: The dependent variable is real wage cyclicality (mean: -0.0056; s.d.: 0.267). FGLS regressions accounting for heteroskedasticity due to uncertainty in first-stage estimates (see the text for details). Absolute value of t-statistics in parentheses. * and ** denote significance at the 5% and 1% level respectively.

Table 3: Determinants of real wage cyclicality: country effects

	(1)	(2)	(3)	(4)	(5)
Australia	-0.165	-0.123	-0.158	-0.192	-0.029
	(4.96)**	(2.31)*	(3.81)**	(5.07)**	(0.63)
Austria	-0.110	-0.107	-0.137	-0.197	0.061
	(3.39)**	(2.05)*	(3.41)**	(5.16)**	(0.84)
Belgium	-0.182	-0.160	-0.216	-0.180	-0.071
	(5.51)**	(3.03)**	(5.24)**	(4.55)**	(0.97)
Canada	-0.403	-0.398	-0.380	-0.423	-0.296
	(15.13)**	(9.39)**	(11.29)**	(13.99)**	(4.54)**
Denmark	-0.207	-0.177	-0.182	-0.281	-0.207
	(6.49)**	(3.45)**	(4.44)**	(7.65)**	(2.95)**
Finland	-0.097	-0.105	-0.132	-0.006	0.085
	(3.02)**	(2.06)*	(3.28)**	(0.17)	(1.20)
France	-0.009	0.014	-0.020	0.115	0.231
	(0.28)	(0.25)	(0.46)	(2.89)**	(3.29)**
Germany	0.157	0.139	0.115	0.184	0.253
	(4.98)**	(2.76)**	(2.85)**	(5.13)**	(3.76)**
Ireland	-0.352	-0.345	-0.313	-0.251	-0.287
	(12.33)**	(7.53)**	(8.43)**	(7.14)**	(4.20)**
Italy	-0.182	-0.179	-0.129	-0.122	-0.041
	(5.73)**	(3.52)**	(3.20)**	(3.19)**	(0.57)
Japan	0.139	0.083	0.094	0.019	0.300
	(4.21)**	(1.58)	(2.19)*	(0.48)	(4.54)**
Netherlands	-0.143	-0.165	-0.148	-0.082	0.019
	(4.30)**	(3.16)**	(3.52)**	(2.03)*	(0.26)
New Zealand	-0.308	-0.318	-0.211	-0.319	-0.435
	(9.96)**	(6.46)**	(5.18)**	(8.94)**	(7.16)**
Norway	-0.116	-0.128	-0.164	-0.007	-0.119
	(3.63)**	(2.51)*	(4.19)**	(0.18)	(1.67)
Spain	-0.367	-0.370	-0.340	-0.361	-0.325
	(12.75)**	(8.10)**	(9.13)**	(10.49)**	(4.99)**
Sweden	-0.132	-0.140	-0.128	-0.185	0.067
	(4.11)**	(2.76)**	(3.21)**	(5.08)**	(0.93)
United Kingdom	0.056	0.071	0.034	-0.071	0.149
	(1.68)	(1.36)	(0.83)	(1.73)	(2.09)*
United States	0.081	0.072	0.067	-0.014	0.123
	(2.49)*	(1.39)	(1.64)	(0.36)	(1.73)
Data and methods fixed effects	Yes**	Yes**	Yes**	Yes**	Yes**
Interactions with country effects:					
• Business cycle frequencies	No	Yes	No	No	No
• Method	No	No	Yes**	No	No
• Cycle Measure	No	No	No	Yes**	No
• Deflators	No	No	No	No	Yes**
Observations	648	648	648	648	648
Adjusted R^2	0.58	0.56	0.60	0.67	0.74

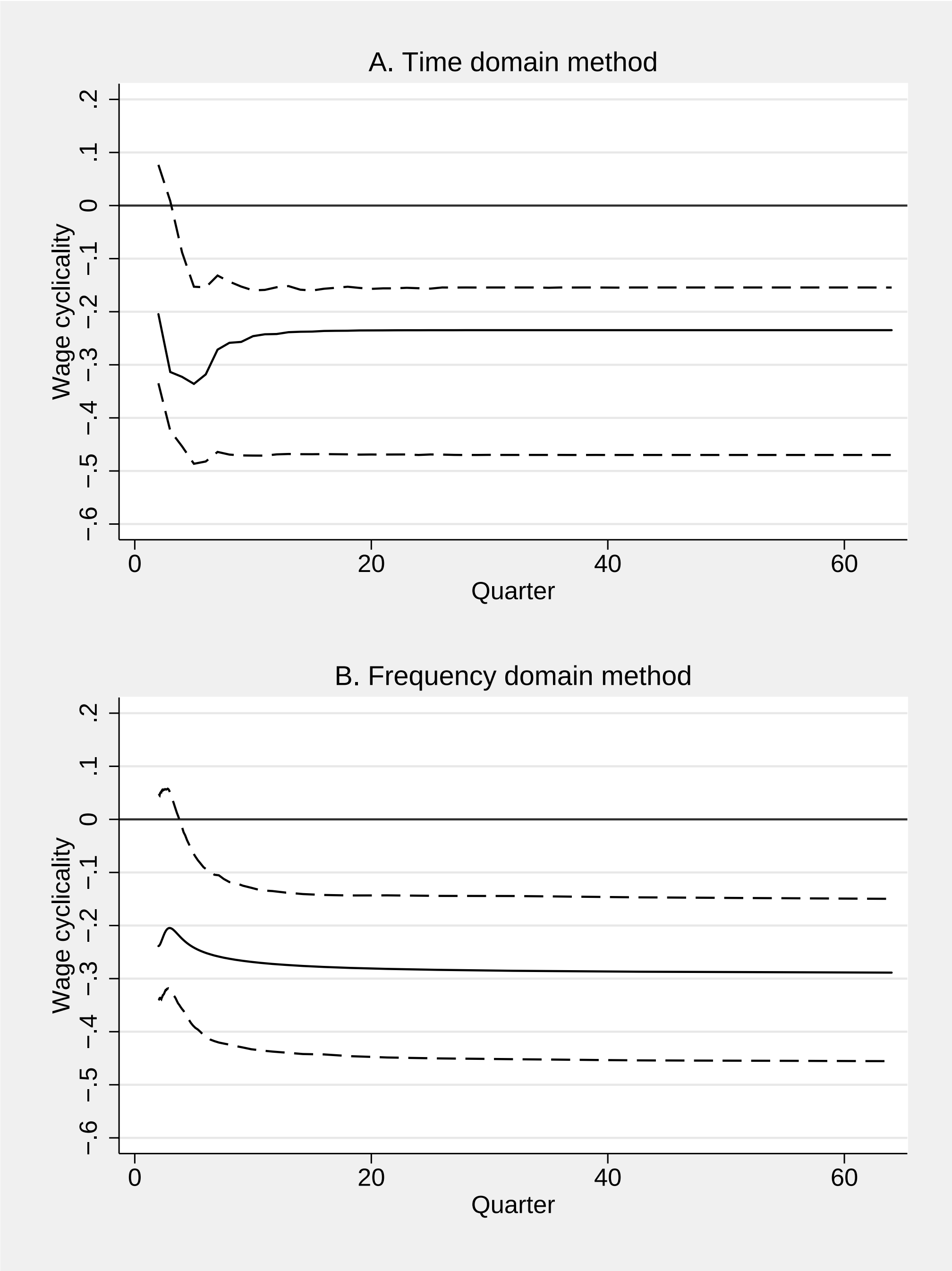
Note: The dependent variable is real wage cyclicality (mean: -0.0056; s.d.: 0.267). FGLS regressions accounting for heteroskedasticity due to uncertainty in first-stage estimates (see the text for details). Absolute value of t-statistics in parentheses. * and ** denote significance at the 5% and 1% level respectively. Column 1 presents the country effects corresponding to the specification presented in Column 2 of Table 2. The baseline category (shown in the table) for the interaction with country effects in columns 2-5 is: short run (column 2), frequency domain (column 3), employment (column 4) and CPI deflator (column 5).

Table 4. Wage and employment cyclicality

	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
	Consumer Wages			Producer Wages		
Employment Cycle	0.437	0.443	0.336	0.183	0.193	-0.072
	(2.74)*	(2.66)*	(3.78)**	(1.57)	(1.58)	(0.64)
Medium Run		-0.009	-0.005		-0.011	-0.002
		(0.54)	(0.32)		(0.97)	(0.17)
Long Run		-0.010	-0.005		-0.019	-0.007
		(0.52)	(0.29)		(1.31)	(0.46)
Time Domain		-0.065	-0.062		-0.079	-0.074
		(1.21)	(1.19)		(1.94)	(1.95)
Cluster 2			0.172			0.332
			(2.57)*			(5.08)**
Cluster 3			-0.371			-0.101
			(11.42)**			(1.18)
Observations	108	108	108	108	108	108

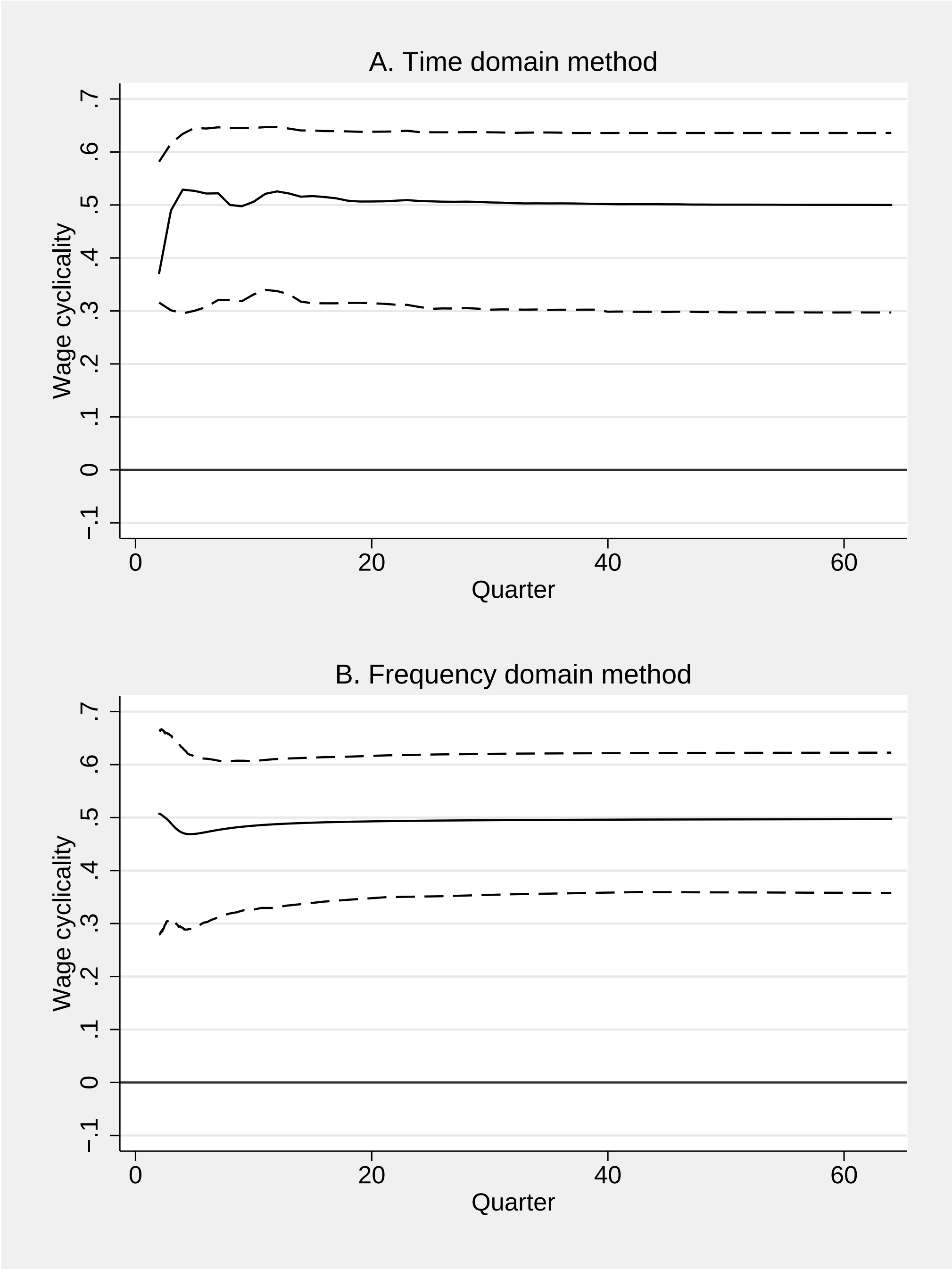
Note: Standard errors are robust to any form of heteroskedasticity, and are allowed to be clustered within countries. Cluster 2: takes value 1 for countries with mostly pro-cyclical wages as suggested in the text. Cluster 3: takes value 1 for countries with mostly counter-cyclical wages as suggested in the text.

Figure 1: Selected measures of co-movement for Spain
- Real wages (GDP deflated) and industrial production as cyclical indicator -



Note: Dashed lines represent 90% confidence bands. Bootstrapped standard errors based on 1000 replications were used to construct confidence bands and standard errors.

Figure 2: Selected measures of co-movement for the United States
- Real wages (GDP deflated) and industrial production as cyclical indicator –



Note: Dashed lines represent 90% confidence bands. Bootstrapped standard errors based on 1000 replications were used to construct confidence bands and standard errors.

Figure 3. Distribution of wage cyclicality estimates across different dimensions

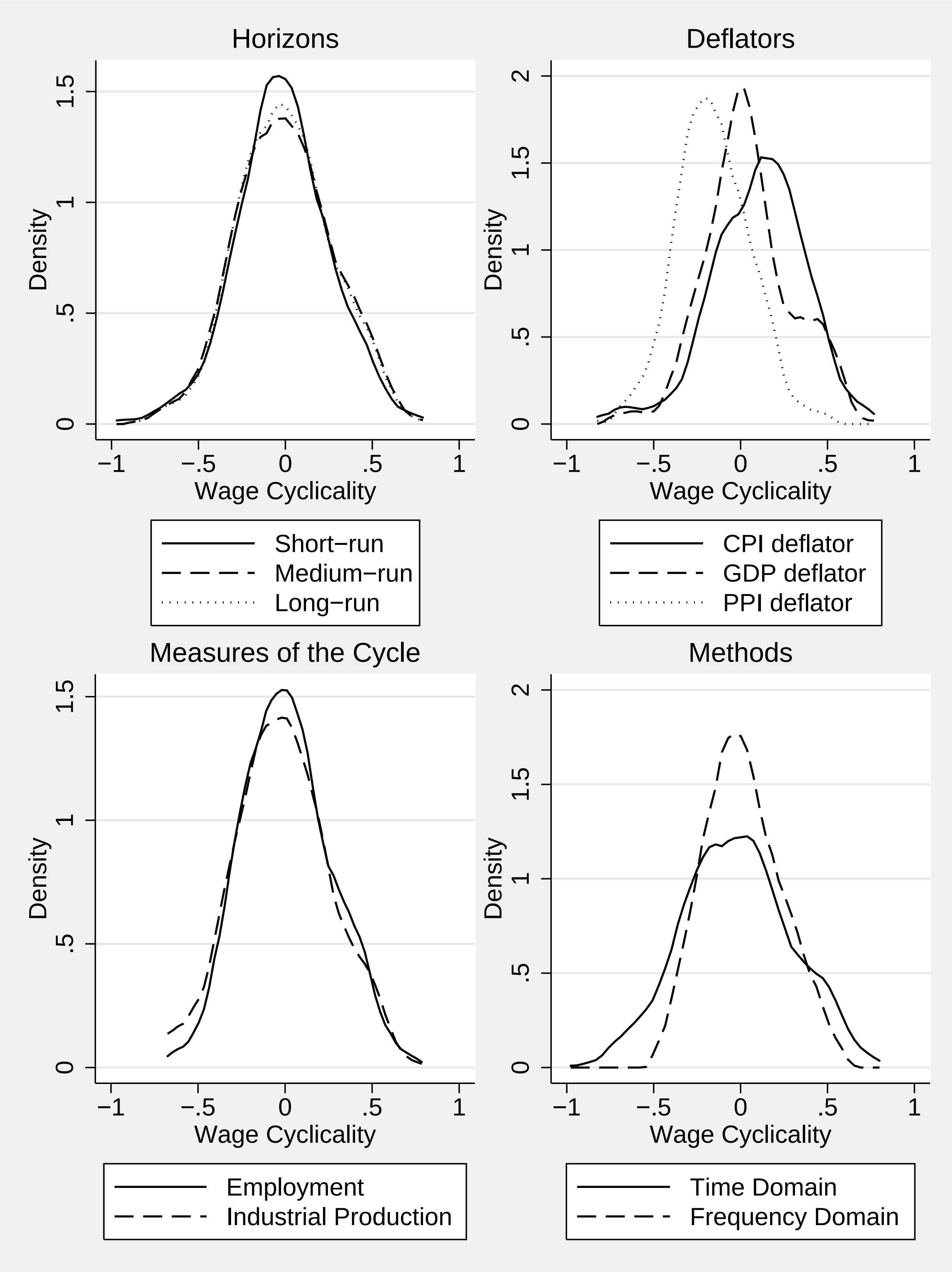


Figure 4. Summary measures of real wage cyclicality

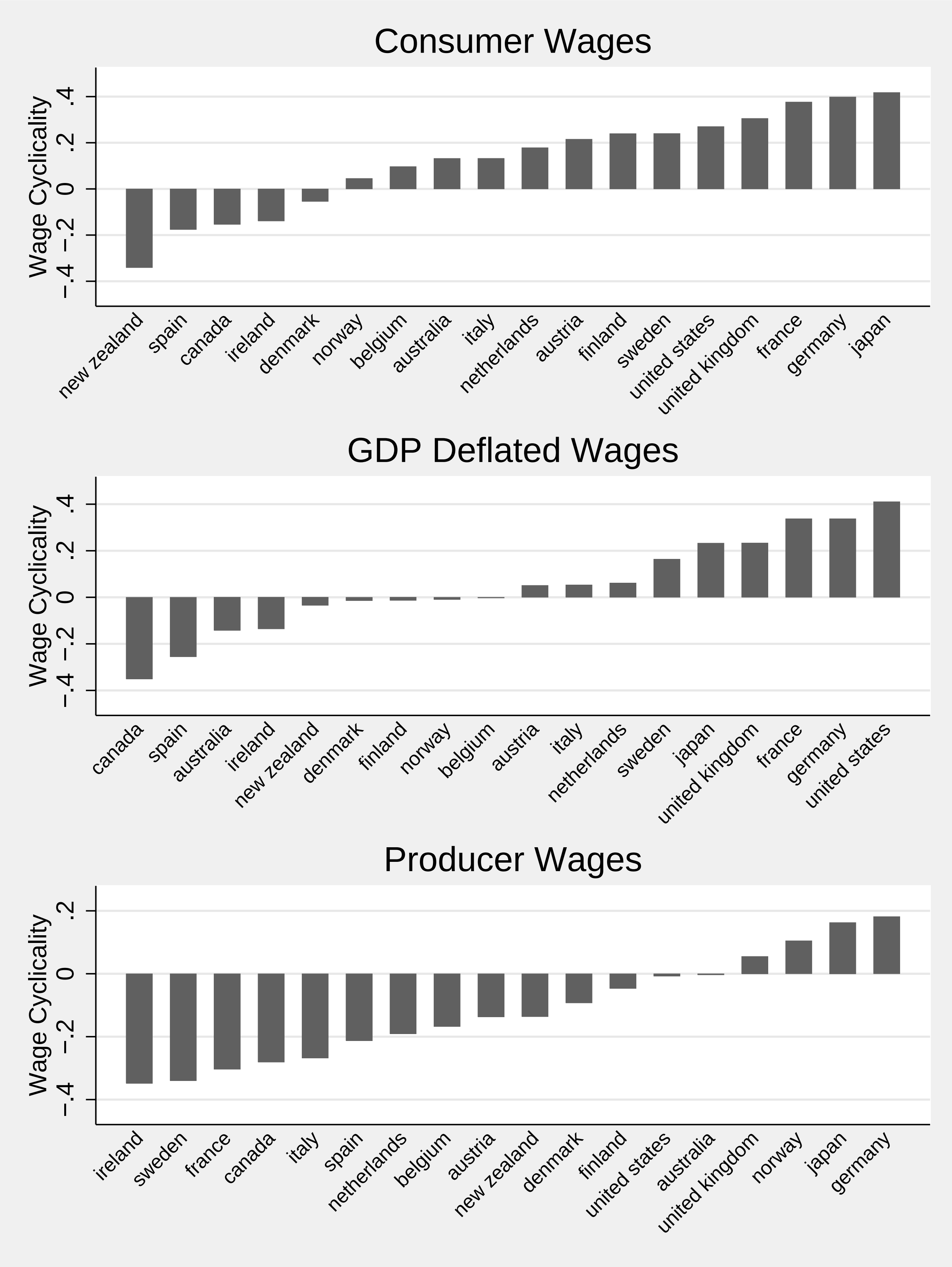


Figure 5. Cluster analysis

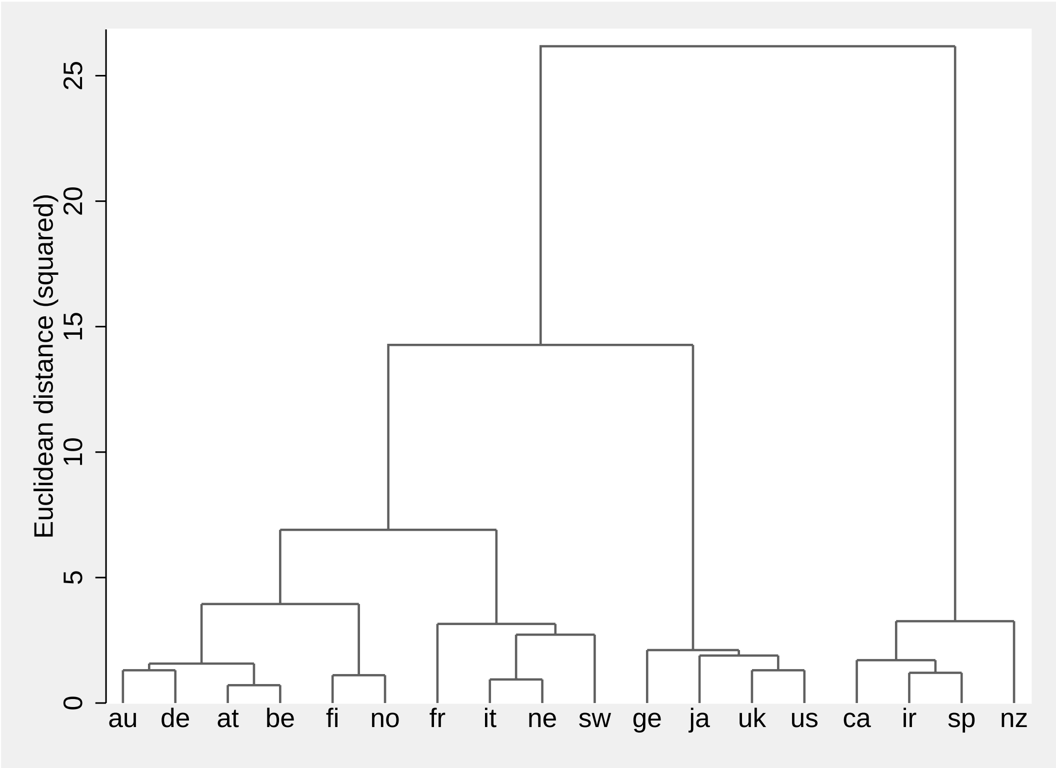
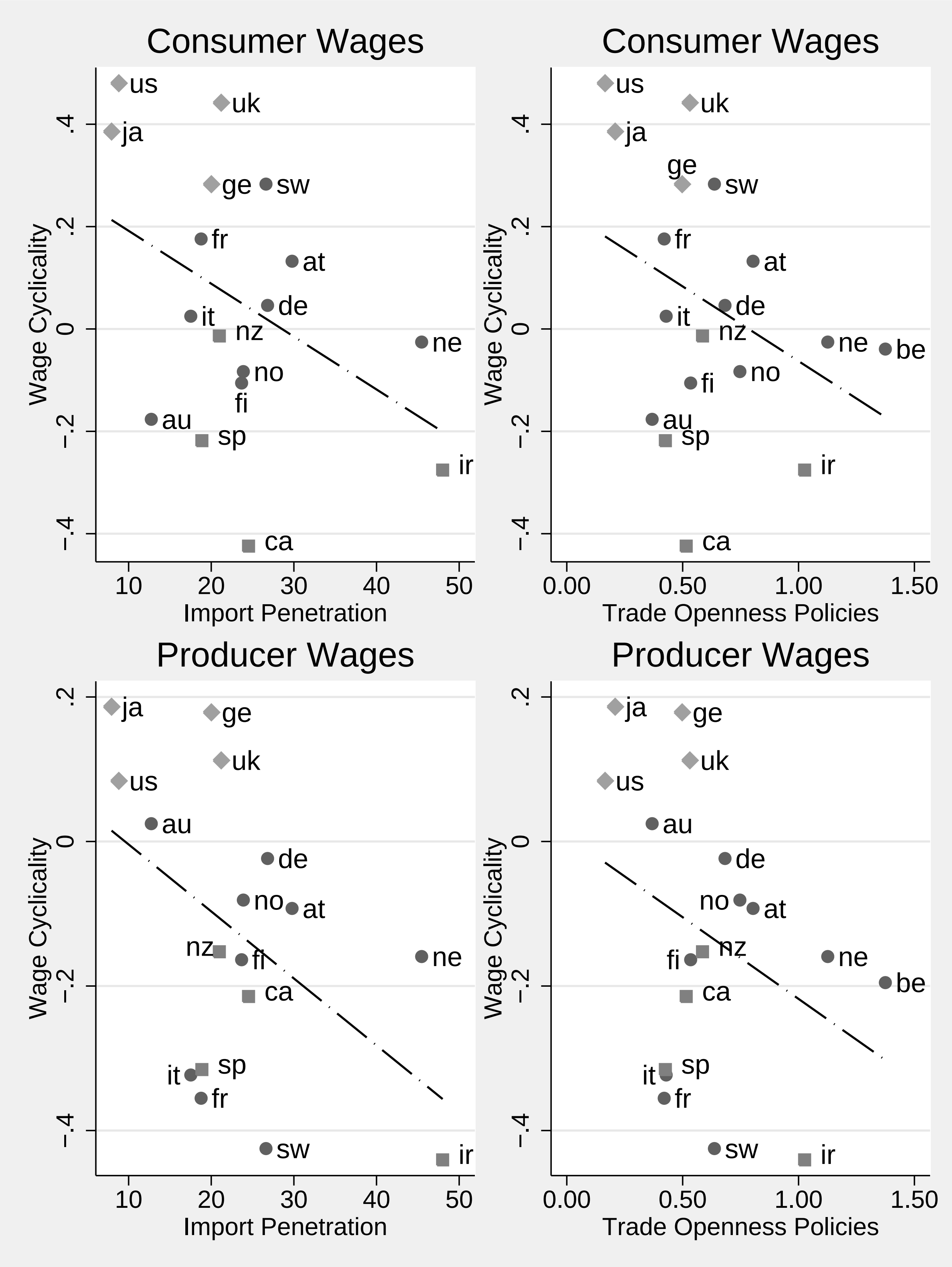
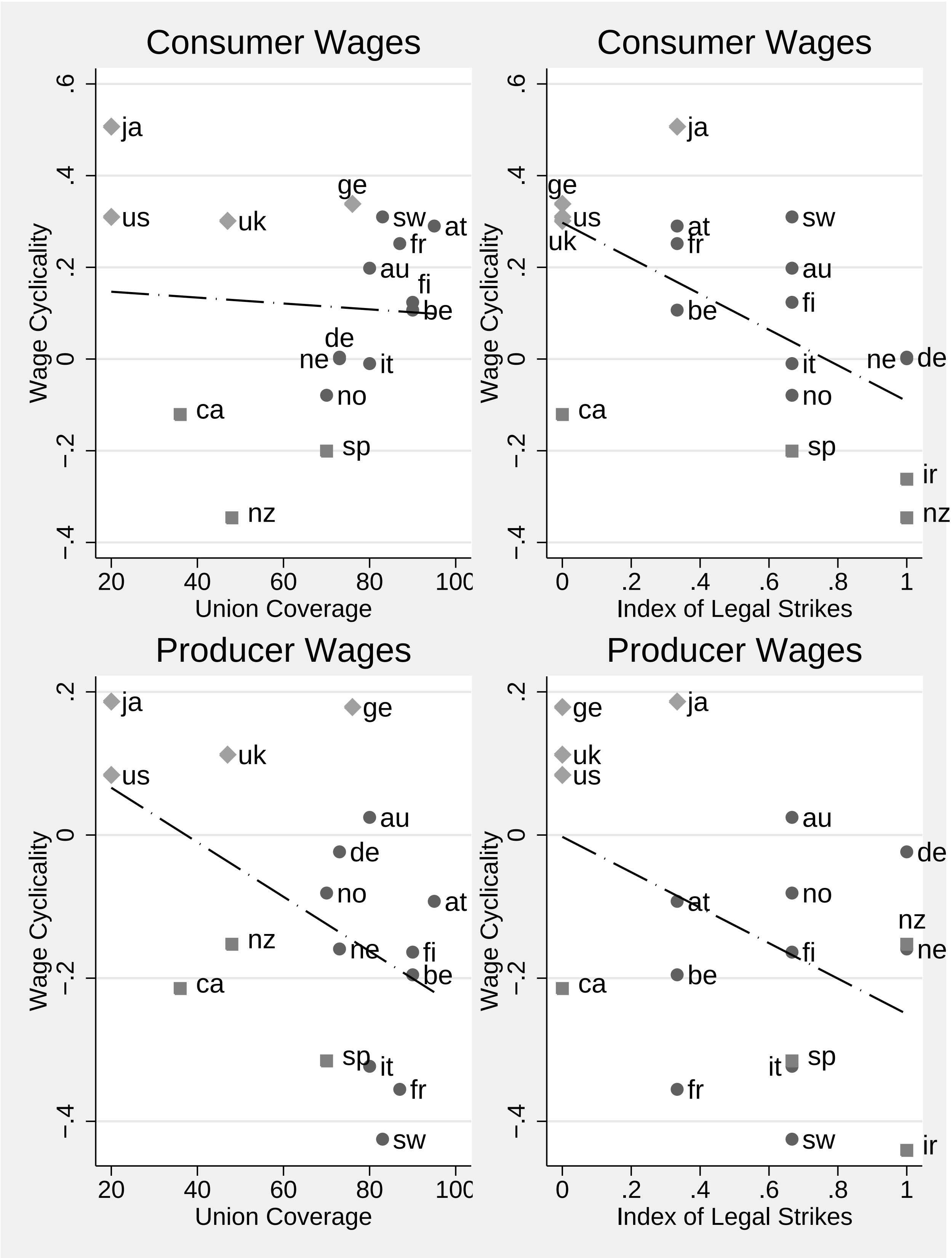


Figure 6. Wage cyclicality and trade openness



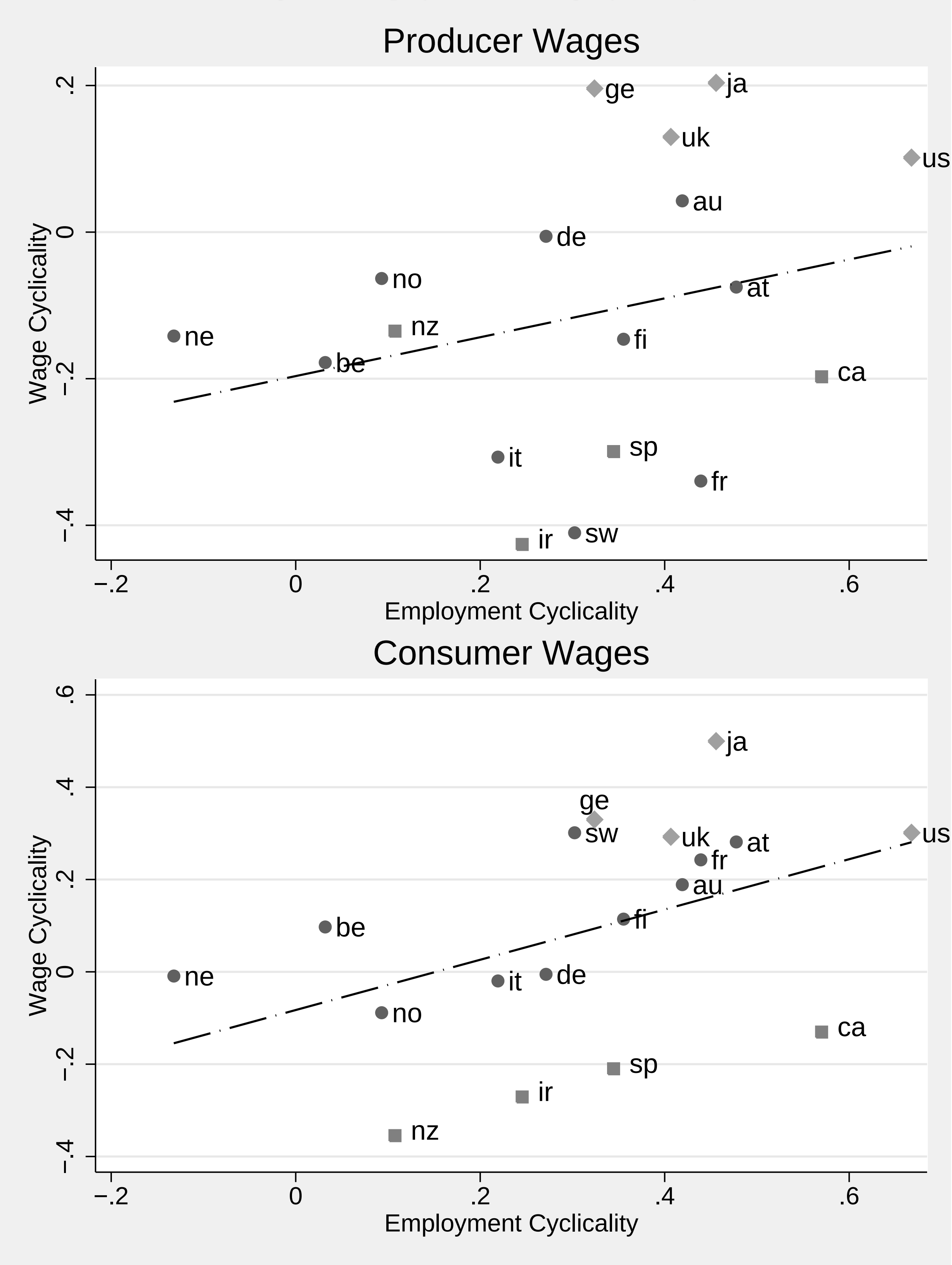
Note: Each symbol corresponds to one of the country clusters identified in Figure 5. The ♦ represents countries with mainly pro-cyclical wages, ■ stands for countries with mainly counter-cyclical wages and ● for the "RoW" cluster.

Figure 7. Wage cyclicality and labour relations



Note: Each symbol corresponds to one of the country clusters identified in Figure 5. The ♦ represents countries with mainly pro-cyclical wages, ■ stands for countries with mainly counter-cyclical wages and ● for the "RoW" cluster.

Figure 8. Employment and wage cyclicality



Note: Each symbol corresponds to one of the country clusters identified in Figure 5. The ♦ represents countries with mainly pro-cyclical wages, ■ stands for countries with mainly counter-cyclical wages and ● for the "RoW" cluster.

Appendix 1: Sample coverage

	Real wages (GDP deflated)		Real wages (CPI deflated)		Real wages (PPI deflated)		Employment		Industrial production	
	Start	End	Start	End	Start	End	Start	End	Start	End
Australia	1983q4	2004q1	1983q4	2004q1	1983q4	2004q1	1976q3	2004q1	1960q1	2005q4
Austria	1960q1	2003q4	1960q1	2003q4	1960q1	2003q4	1988q1	2004q1	1960q1	2005q4
Belgium	1980q1	2004q1	1980q1	2004q1	1980q1	2004q1	1981q1	2003q2	1960q1	2005q4
Canada	1961q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1961q1	2005q4
Denmark	1977q1	2003q4	1971q1	2003q4	1974q1	2003q4	1977q1	2003q4	1977q1	2005q4
Finland	1960q1	2004q1	1960q1	2004q1	1975q1	2004q1	1960q1	2004q1	1960q1	2005q4
France	1963q1	2003q4	1960q1	2003q4	1960q1	2003q4	1970q1	2004q1	1963q1	2005q4
Germany	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2005q4
Ireland	1960q1	2003q4	1960q1	2003q4	1968q1	2003q4	1960q1	2003q4	1960q1	2005q4
Italy	1960q1	2004q1	1960q1	2004q1	1981q1	2004q1	1970q1	2004q1	1960q1	2005q4
Japan	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2005q4
Netherlands	1970q1	2004q1	1970q1	2004q1	1971q1	2004q1	1987q1	2003q4	1960q1	2005q4
New Zealand	1989q1	2004q1	1989q1	2004q1	1989q1	2004q1	1989q1	2004q1	1961q2	2005q4
Norway	1960q1	2004q1	1960q1	2004q1	1977q1	2004q1	1972q1	2004q1	1960q1	2005q4
Spain	1981q1	2003q4	1981q1	2003q4	1981q1	2003q4	1976q3	2004q1	1970q1	2005q4
Sweden	1971q1	2004q1	1971q1	2004q1	1982q1	2004q1	1963q1	2004q1	1960q1	2005q4
United Kingdom	1963q1	2004q1	1963q1	2004q1	1963q1	2004q1	1980q1	2003q4	1960q1	2005q4
United States	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2004q1	1960q1	2005q4

Note: The table reports for each country the available coverage for the series on real wages and indicators of the cycle. Instead of showing the coverage separately for nominal wages, GDP deflator, CPI deflator and PPI deflator, we show directly the coverage of the correspondent real wage series. The reason is that the real wage series are the object of the analysis. The source is the OECD Main Economic Indicators database.

Appendix 2: Description of institutional variables

- Import penetration is defined as the share of imported goods on total internal demand. Source: OECD ANA database.

- Trade openness policies are measured as the fraction of years from 1960 to 1998 that the country does not interfere with foreign trade, as compiled by Sachs and Warner (1995) measured on a (0, 1) scale. A country is considered open if it satisfies all of the following criteria: (1) nontariff barriers cover less than 40 percent of trade; (2) average tariff rates are less than 40 percent; (3) the black market premium was less than 20 percent during the 1970s and 1980s; (4) the economy is not socialist; and (5) the government does not control major exports through marketing boards.

- Union coverage data for all countries with the exception of New Zealand, Spain and Ireland come from Golden and Wallerstein (2002). Coverage for New Zealand and Spain is reported in OECD (1997) and for Ireland in Holden and Wulfsberg (2008).

- Legal strikes is the average of three indicator variables: (1) Wildcat strikes are legal, (2) Political strikes are legal, (3) Sympathy / solidarity / secondary strikes are legal. Source: Botero et al. (2004).

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