Equilibrium Unemployment Dynamics in a Panel of OECD Countries*

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Abstract

We focus on the equilibrium unemployment rate as a parameter implied by a dynamic aggregate model of wage and price setting. The equilibrium unemployment rate depends on institutional labour market institutions through mark-up coefficients. Compared with existing studies, the resulting final equation for unemployment has a richer dynamic structure. The empirical investigation is conducted in a panel data framework and uses OECD data up to 2012. We propose to extend the standard estimation method with time dummies to control and capture the effects of common and national shocks by using impulse indicator saturation (WG-IIS), which has not been previously used on panel data. WG-IIS robustifies the estimators of the regression coefficients in the dynamic model, and it affects the estimated equilibrium unemployment rates. We find that wage co-ordination stands out as the most important institutional variable in our dataset, but there is also evidence pointing to the tax wedge and the degree of compensation in the unemployment insurance system as drivers of equilibrium unemployment.

I. Introduction

The concept of equilibrium unemployment in the OECD area has been the subject of both analytical and empirical research. One influential analytical approach, which also underlies our research, combines a model of monopolistic price setting among firms with collective bargaining over the nominal wage level; see Layard et al. (2005) and Layard and Nickell (2011). Intuitively, when the system is not in a stationary situation, nominal wage and price adjustments constitute a wage–price spiral that leads to increasing or falling inflation. According to Layard et al. (2005), the equilibrium for real wages requires that unemployment becomes equal to the non-accelerating inflation rate of unemployment (NAIRU). The equilibrium unemployment rate is not interpretable as a completely invariant parameter that is pinned down once and for all by autonomous market structures. Specifically, it can vary with changing labour market institutions.

We attempt to bridge the gap between the formal, but static, theoretical framework, which is common to several empirical studies, Belot and van Ours (2001), Bassanini and

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Duval (2006), Nickell et al. (2005) and Blanchard and Wolfers (2000) among others, and the dynamic specification of the estimated models. Our theoretical model leads to a cointegrated vector autoregression (VAR) with three endogenous variables; the rate of unemployment, the real exchange rate and the wage share. The VAR-based equilibrium rate of unemployment is also a NAIRU because it is determined jointly with a constant rate of inflation. Our main parameter of interest is the equilibrium rate of unemployment, rather than all the other parameters of the full system. We therefore base our empirical model on the main aspects of a final equation for the unemployment rate that is implied by the VAR.

The theoretical model gives some guidelines for the empirical specification that differ from those in the conventional empirical literature. First, the autoregressive order of the final equation is shown to be three, whereas the custom has been to impose first-order dynamics. The higher-order dynamics can be verified or refuted by econometric testing, which we do in the empirical parts of the paper. Second, the expected sign and magnitudes of the autoregressive coefficients stem from the theoretical model. Third, the theoretical model also has implications for the error term, which is affected by foreign price and technology shocks, which, when not controlled for, will be disturbances. We therefore allow for a heteroscedastic and autocorrelated error term in the empirical part of the analysis. Our dynamic theory model also predicts that changes in institutional features will affect unemployment rates gradually, not by sudden shifts, which is realistic and in line with earlier studies.

One of the contributions of this paper is to estimate the equilibrium level of unemployment for each of twenty OECD countries from a panel dataset, dependent on the current level of the labour market institutions. The estimation results, based on data for the period 1960 to 2012, show that different levels of wage-bargaining co-ordination, the generosity of the unemployment insurance system and tax wedges may have contributed towards different levels of equilibrium unemployment. We find that the equilibrium unemployment rates range from just above 3 percentage points to 13 percentage points.

The result is based on an estimation method that controls for all major shocks that have hit the individual unemployment rates in the sample. It includes the onset of the credit crisis and the transformation of the financial crisis into an international job crisis. To account for both national and common shocks, we have augmented the model using country-specific indicator variables of years that represent temporary location shifts in the mean of the individual unemployment rates (see Doornik (2009) and Johansen and Nielsen (2009)). This is a first application of so-called impulse indicator saturation to an econometric panel data model. We suggest this as a way of obtaining robust estimators of the regressions coefficients (and the standard errors) of the theoretical explanatory variables in the model and as a means to investigate the effect of shocks on the equilibrium unemployment rates.

We consider that our findings are in line with earlier panel data analysis in which equilibrium unemployment is determined by labour market institutions. Specifically, Nickell et al. (2005) find a strong role for institutional variables in explaining the increase in unemployment rates in Europe over the their sample period, 1960 to 1995. This corresponds to our analysis, where the average equilibrium unemployment increased by 4 percentage points during the 1970s and 1980s because of institutional changes. However, we do not find any evidence of a declining average equilibrium unemployment in the last part of our sample period. In our results, the average equilibrium unemployment rate has remained high throughout the 2000s and to date, despite the fact that many countries have adopted the recommendations of the OECD's job market study about reforms that make labour markets more flexible.

The literature is inconclusive about which particular institutional factors are most important for the equilibrium unemployment rate. This is also our experience, as different estimation methods points to somewhat disparate results. However, the weight of the evidence points in the direction of co-ordination (in the process of wage setting) and the generosity of the insurance system as the most important factors behind the explained part of equilibrium unemployment. On the other hand, the degree of employment protection is not a key determinant, in contrast to recent arguments in the debate about what causes high unemployment rates.

We should not be too disappointed to find that a broader spectrum of institutional variables is not found to be robust across estimators and significant, given this is predicted by the theory we use. In our model, institutions affect unemployment via their impact on the wage and price mark-up coefficients, and these effects on unemployment can cancel out. Bjørnstad and Kalstad (2010) observe that if the price mark-up coefficient is positively linked to the degree of co-ordination, the net effect of increased co-ordination on unemployment may be ambiguous. Following Bowdler and Nunziata (2007) and Bjørnstad and Kalstad (2010), we model price formation as a function of wage-determining factors and find that, indeed, the effect of co-ordination is positive on the price mark-up.

The paper is organized as follows. The dynamic model for equilibrium unemployment derived from the theoretical dynamic model of the wage and price spiral is presented in section II. The data for the evolution of labour market institutions are presented in section III. Econometric estimation issues and the method for deriving location shifts using panel data are briefly discussed in section IV. The results from the estimated dynamic unemployment equations are presented in section V, variables and dynamics and then extend this to controlling for location shifts. The estimated equilibrium unemployment rates are presented in section VI. We summarize our results in section VII, where we also discuss some extensions and provide suggestions for further work.

II. A framework for equilibrium unemployment

To estimate and test hypotheses about the equilibrium rate of unemployment (u^*), it is helpful to specify a medium-run dynamic macromodel for open economies.

Unemployment, real exchange rate and real wage

A simple dynamic relationship between u_t, the rate of unemployment in period t and re_t, the logarithm of the real exchange rate is given by:

$$u_t = c_u + \alpha\, u_{t-1} - \rho\, re_{t-1} + \epsilon_{u,t}, \;\; \rho \geqq 0, -1 < \alpha < 1, \tag{1}$$

re_t is defined in such a way that an increase in the real exchange rate leads to improved competitiveness (depreciation). This increases exports and, thereby, GDP increases and unemployment falls. Hence, $\rho \geqq 0$. $\epsilon_{u,t}$ contains all other variables that might affect u_t. The simplest interpretation of (1) is that it represents a stylized dynamic aggregate demand relationship, where the effects of other variables, e.g., the real interest rate, have been subsumed in the disturbances $\epsilon_{u,t}$, $t = 1, 2, ..., T$, which is therefore autocorrelated in general.

The importance of wage-setting institutions for long-term unemployment performance is a key point of the Layard–Nickell model. In (1), the link to the supply-side, wage and price setting, is represented by the real exchange rate. In appendix A (available online), we present a dynamic model for wage and price setting that contains the relationships that represent Layard and Nickell's wage and price setting curves as cointegrating relationships. The model includes nominal trends in foreign prices and the nominal exchange rate, as well as the real trend in labour productivity. These trends add realism to the model because unit-root type non-stationarities in wages and prices are hard to reject empirically. On the other hand, logical consistency requires cointegration because unemployment is stationary only when the real exchange rate is without stochastic trends. However, stochastic trends are present in the processes for the nominal price levels and for the exchange rate.

As shown in the Appendix A (available online), the wage–price model can be written as a dynamic system for re and the logarithm of the wage share (ws). When we combine this result with the equation for u_t we obtain the VAR:

$$\underbrace{\begin{pmatrix} re_t \\ ws_t \\ u_t \end{pmatrix}}_{\mathbf{y}_t} = \underbrace{\begin{pmatrix} l & -k & n \\ \lambda & \kappa & -\eta \\ -\rho & 0 & \alpha \end{pmatrix}}_{\mathbf{R}} \underbrace{\begin{pmatrix} re_{t-1} \\ ws_{t-1} \\ u_{t-1} \end{pmatrix}}_{\mathbf{y}_{t-1}} + \underbrace{\begin{pmatrix} e & 0 & -d \\ \xi & -1 & \delta \\ 0 & 0 & c_u \end{pmatrix}}_{\mathbf{P}} \underbrace{\begin{pmatrix} \Delta pi_t \\ \Delta a_t \\ 1 \end{pmatrix}}_{\mathbf{x}_t} + \underbrace{\begin{pmatrix} \epsilon_{re,t} \\ \epsilon_{ws,t} \\ \epsilon_{u,t} \end{pmatrix}}_{\epsilon_t} \tag{2}$$

The first two rows of the autoregressive matrix $\mathbf{R}$ contain reduced-form coefficients that are known expressions of the parameters of the model of the supply side (see the Appendix A (available online)). The third row contains the parameters of (1). For $\mathbf{y}_t$ to be stationary, none of the eigenvalues of $\mathbf{R}$ can have moduli on the complex unit circle. In the following, we assume a causal VAR, in which case the necessary and sufficient condition is that all the eigenvalues have moduli strictly less than one, Brockwell and Davies (1991, Ch. 3). This type of stationarity is secured by the assumptions about the parameters that we make in (1) and the Appendix A (available online).

The second term, $\mathbf{Px_t}$, in (2) shows that foreign price growth, Δpi_t, and exogenous productivity growth, Δa_t, play a role in the dynamic behaviour of $\mathbf{y}_t$. Foreign price growth (Δpi_t) affects the vector of the real variable $\mathbf{y}_t$, as dynamic price homogeneity is not imposed from the outset, unlike long-run price homogeneity. As shown in the Appendix (available online), dynamic price–wage homogeneity implies that the coefficients e and ξ in the $\mathbf{P}$ matrix are both restricted to zero.

The right column of $\mathbf{P}$ contains the three intercepts, of which d and δ are reduced forms, given by the expressions in the Appendix A (available online). The vector $(\epsilon_{re,t}, \epsilon_{prw}, \epsilon_{u,t})$ contains the VAR disturbances, which are reduced-form expressions of the structural disturbances.

Equilibrium rate of unemployment

The stable steady-state solutions of the endogenous variables correspond to their unconditional expectations. For the rate of unemployment in particular, we define the equilibrium rate by $u^* = E(u_t)$. It is given by:

$$u^* \equiv E(u_t) = \mathfrak{d}_{ss} - \mathfrak{e}_{ss}\, g_{pi} - \underset{\pm}{\mathfrak{b}_{ss}} g_a. \tag{3}$$

$\mathfrak{e}_{ss}$ is zero when the system is dynamically homogenous; otherwise, it is positive. The symbol $\pm$ below $\mathfrak{b}_{ss}$ indicates that the long-run impact of productivity growth can be zero. The "intercept" $\mathfrak{d}_{ss}$ is important in the following and it serves a point to write it in terms of the structural parameters (see the Appendix A (available online)):

$$\mathfrak{d}_{ss} = [\rho\,(m_w + m_q) + c_u\,\omega\,(1 - \phi)]/\Omega \tag{4}$$

with $\Omega > 0$ given the assumptions of the model.

It is a main thesis of the Layard–Nickell model that increased mark-ups lead to a higher NAIRU (see, e.g., Layard and Nickell (2011, Ch. 2)). This prediction is encompassed by our model, because m_w and m_q are the mark-up coefficients in the cointegrating wage and price curve relationships. However, in our model, the intercept in the aggregate demand equation, c_u, also affects the equilibrium rate, u^*, directly. As explained in the Appendix (available online), $\omega \geq 0$ is the wedge-coefficient in wage formation, and $\phi > 0$ is a parameter that measures the degree of openness of the economy (the share of domestic goods in consumption). g_{pi} and g_a in (3) are the drift terms of the foreign nominal trend and the productivity trend, respectively. As anticipated, the nominal growth rate drops out of the expression for u^* if there is dynamic price homogeneity (see the Appendix A (available online) for details).

m_w is a parameter that researchers think of as conditioned by the social order. It is not regarded as invariant to changes in wage-setting institutions and to other labour market reforms. Nunziata (2005) finds evidence of a monotonous relationship between co-ordination in wage setting and the real wage, Bowdler and Nunziata (2007). In our model, this entails that m_w is a declining function of co-ordination. The price mark-up, m_q, has received less attention and it is often regarded as a more autonomous parameter than the wage mark-up. An interesting exception is Bjørnstad and Kalstad (2010), who use panel data and find that the price mark-up is significantly higher in countries with a high degree of wage co-ordination compared with unco-ordinated countries. This may dampen the negative effect of co-ordination on unemployment that would otherwise be a consequence of a reduced wage mark-up. Bowdler and Nunziata (2007) investigate the effect of co-ordination on inflation, and show that co-ordination has an effect via an interaction term with the size of unionization.

In the empirical section, we introduce measurements of institutional variables for OECD countries. The mechanism that we mainly have in mind is that institutional variation can explain differences and the evolution in equilibrium unemployment through the two mark-up coefficients. This is the same hypothesis that Nickell et al. (2005) investigates. Our contributions are that the dynamics of the model are made explicit, we have reviewed and revised the operational measure of institutional attributes and the time series that we use are longer than those of the existing studies.

Finally, (4) reminds us that there is no contradiction between including institutional variables that, in theory, mainly explain variations in the mark-ups, and other type of variables that capture (or represent) changes in the intercept c_u. In this paper, we use indicator variables in the way that we explain in section IV. to represent such shifts. At this point, it might be noted, however, that even if the breaks in c_u are impermanent (not a step-function), we expect that they will have an effect on equilibrium unemployment dynamics, as (2) show.

The final equation model

Multi-equation econometric models of wage and price setting exist for single countries (see Bårdsen and Nymoen (2003) and Akram and Nymoen (2009) (Norway), and Bårdsen and Nymoen (2009) (USA), Schreiber (2012) and Bowdler and Jansen (2004)). Few of these studies have focused on the implied equilibrium unemployment rate, and there are no genuine panel data studies. We leave it to future work to take this approach to panel data, anticipating that progress can be made using the approach of Arellano (2003, Chap. 6), where a bivariate VAR for employment and wages is estimated for a (micro) panel. In this paper, we base the econometric study on the final equation for the unemployment rate implied by the structural VAR model. This equation determines the equilibrium rate as a parameter, and it avoids the difficulty of identifying the parameters of the wage and price equations.

To formalize our approach we make use of the implied third-order dynamics for u_t of the VAR model (2). This final equation model can be written as:

$$u_t = \beta_0 + \beta_1 u_{t-1} + \beta_2 u_{t-2} + \beta_3 u_{t-3} + \varepsilon_{u,t} \tag{5}$$

The autoregressive coefficients can be expressed in terms of the VAR parameters:

$$\begin{aligned} \beta_1 &= \alpha + \kappa + l \\ \beta_2 &= -\left[\alpha l(1-\kappa) + \kappa(\alpha + l) + n\rho + \lambda k\right] \\ \beta_3 &= \alpha\lambda k + \rho(n\kappa - \eta k) \end{aligned} \tag{6}$$

It follows from the assumptions in Appendix A (available online) that β_1 is positive, and that it may well be larger than one. Our theory suggests that the second autoregressive parameter is expected to be negative because all the coefficients inside the brackets are positive. We note that it may be reasonable that $\beta_1 > -\beta_2$ because the additional terms in β_2 are products of factors that are less than one. The third autoregressive coefficient, β_3, is likely to be markedly smaller in magnitude than the first two coefficients: $\alpha\lambda k$ is a small number and $\rho(n\kappa - \eta k)$ may be negative.

The characteristic roots associated with (5) are the same as the eigenvalues of $\mathbf{R}$ in the VAR. Hence, if the VAR is stationary, it follows that the process for u_t given by (7) is also stationary, and vice versa. Therefore, in our framework, the equilibrium rate of unemployment is determined uniquely. If we set the drift terms in a_t and pi_t to zero in order to simplify notation, we obtain:

$$u^* = \frac{\beta_0}{1 - \beta_1 - \beta_2 - \beta_3} \equiv [\rho\,(m_w + m_q) + c_u\,\omega\,(1 - \phi)]/\Omega \tag{7}$$

with:

$$1 - \beta_1 - \beta_2 - \beta_3 > 0 \tag{8}$$

because the low frequency characteristic root is inside the unit circle, a consequence of stationarity.[1] (6) shows how the autoregressive coefficients depend on the underlying parameters and (7) shows that β_0 is directly related to the institutionally determined parameters m_q and m_w that we discussed above. The final equation is comparable with the single-equation models used in the existing panel data studies of unemployment and institutions, which were cited in the Introduction.

One additional implication worth noting is the final equation disturbance $\epsilon_{u,t}$:

$$\begin{aligned}\varepsilon_{u,t} = {} & -(l-\kappa)\epsilon_{u,t-1} + (\lambda k + l\kappa)\varepsilon_{u,t-2} - \rho\epsilon_{re,t-1} + \rho\kappa\epsilon_{re,t-2} + k\rho\epsilon_{ws,t-2} \\ & - \rho e\Delta pi_{t-1} + \rho(\xi k + ek)\Delta pi_{t-2} + kp\Delta a_{t-2}\end{aligned} \tag{9}$$

It is made up of two-period moving-averages of the VAR disturbances, but also of the random shocks to import prices and productivity.

In section IV., we discuss the estimation methods we have used to estimate (5) on the panel dataset that we present in the next section.

III. Data

In this section, we present the evolution of the unemployment rates for 20 OECD countries (listed in Table 3) over the period 1960 to 2012. We also present the variables for the labour market institutions that are important for equilibrium unemployment. See the Data Appendix C (available online) for the detailed documentation.

The unemployment rate

The standardized unemployment rate from OECD Economic Outlook (2013) is used as a primary data source for the unemployment rate in the OECD countries.

There have been substantial changes in the unemployment rates of the OECD countries in the period 1960 to 2012. Figure 1 shows the unemployment rates in all countries, together with the average unemployment rate. Figure 1 illustrates that a rise in unemployment occurred in the early 1970s, together with an increase in the dispersion of unemployment rates between countries. The difference between the highest and the lowest unemployment rates across countries in 1995 is larger than it is in 1960. Then, from

[1] As seen from (7) and the equation in the Appendix (available online), $e = \xi = 0$ in the case of dynamic price–wage homogeneity. This confirms the result above about u^* being independent of $E(\Delta pi_t) = g_{pi}$ in that reference case. Note that $\mathfrak{d}_{ss}$, β_j, $j = 1, 2, 3.$ and Ω are invariant to the homogeneity restriction

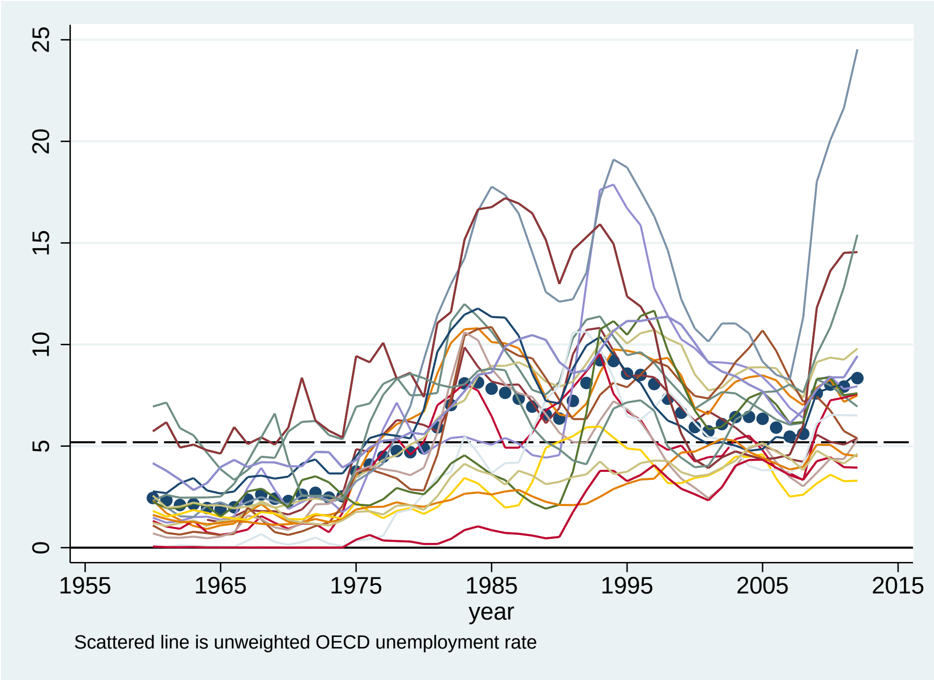

Fig. 1. Unemployment rates in the OECD countries. Per cent

1995 until the onset of the financial crises, both the average unemployment rate and the variation in unemployment rates across countries decreased. For instance, Norway and Switzerland have relatively low unemployment rates after the financial crisis and throughout the period compared with most other countries in the sample. Ireland and Spain, on the other hand, are examples of countries with high levels of unemployment after the financial crisis, but also in some years prior to the crisis. Germany experienced an upward trend in unemployment prior to the financial crisis, but the rate has subsequently fallen. None of the countries in the dataset has seen a declining global trend in unemployment over the entire period.

Note that even if one might question the accuracy of some of the data at the beginning of the period, given that it shows very low unemployment rates in, for instance, New Zealand and Switzerland, our main results are robust to the exclusion of these countries. The results are found in a previous version of this paper (see Sparrman (2011) Chapter 3 for details).

Institutional factors

The main hypothesis to be tested is whether the equilibrium rate of unemployment has been affected by changes in labour market institutions over the sample period. Institutional changes are measured by indices for employment protection (EPL), the benefit replacement ratio (BRR), benefit duration (BD), union density ($UDNET$), the tax wedge (TW) and the degree of co-ordination in wage setting (CO). As noted, these indicators are assumed to capture a gradual evolution in the wage mark-up coefficient (m_w) and, to a lesser degree, in the price mark-up (m_q) and the constant (c_u). As such, the institutional variables are essential in explaining the trend-like behaviour of the data for unemployment rates, within the framework of our model.

The Data Appendix C (available online) contains tables for the actual development for each variable for each country in the sample period. The last row of each table contains the unweighted average for each institutional variable.

The tax wedges are calculated from actual tax payments. The total tax wedge is equal to the sum of the employment tax rate (t1), the direct tax rate (t2) and the indirect tax rate (t3). Appendix, Table C6, illustrates a steady increase in the tax wedge in most OECD countries over the period 1960 to 2012. The largest increases over the period are found in Sweden, Spain and Portugal, and the highest tax wedges, larger than 50 per cent, are found in the Nordic countries and in Austria, Belgium, France, Italy, Spain and Germany. In Belgium, Canada, Finland, New Zealand, Sweden and the United States (US), there was a small decline in tax wedges toward the end of the sample period.

The time series for employment protection measures the strictness of the employment protection for employees. The overall indicator for employment protection is measured on a scale from 0 (low) to 5 (high). The average of the overall employment protection indicator is 2.1 for the whole time period, over all countries in the sample. Appendix, Table C2, shows an average decline in the strictness of employment protection since the beginning of the 1970s. However, there are large differences in the development of employment protection between countries. Countries with the highest levels of employment protection in the 1970s, including Belgium, France, Germany, Italy and Portugal, experienced a decline in the measure towards the end of the sample period. By contrast, the measure increased slightly from a very low level (0.57) in the period 1973 to 1979 in Australia and remained unchanged for Canada over the whole sample period.

The benefit replacement ratio is a measure of how much each unemployed worker receives in benefits from the government in their first period of being unemployed. There has been a steady increase in the average benefit ratio for the period 1960 to 1995 and a stable development since then (see Appendix, Table C4). There are large differences in unemployment benefits between the OECD countries, with the lowest benefits found in the United Kingdom (UK) and Australia, where unemployment benefits were in the range of 0.15 to 0.21 in the period 2009 to 2012. The highest benefits provided are in The Netherlands and Switzerland, with benefits for employees above 70 per cent in the first year of being unemployed. Some countries have reduced the benefits during the sample period; see, for instance, Austria, Canada, Denmark, New Zealand, Sweden and the UK. Italy has increased its benefits and is now in line with the average level of benefits.

Benefit duration is a measure of the unemployment benefits for recipients who have been unemployed more than one year, relative to benefits during the first year. For instance, Canada and Japan stop payments after one year and the index is then equal to zero (see Appendix Table C5). If benefits are the same for the first four years of unemployment, the value of the index is equal to one. Australia and New Zealand are the only countries where the benefits have increased over time and, hence, the index is larger than one for some time periods. In the period 2009 to 2012, there are no countries with increasing benefits over time. We observe that the average benefit duration increased over the sample period until 1995, and that the level has been steady since then. However, there are some variations between countries, i.e., Denmark, France, Ireland, Italy, Portugal, Spain and Switzerland have increased benefit duration in the latter sample period, whereas Germany, The Netherlands and Norway have decreased benefit duration.

We are interested in the effect of co-ordination on unemployment, i.e., through wage moderation. We use the index in Visser (2011), which measures whether co-ordination actually results in wage moderation at all times. The index is based on former work by Kenworthy (2001). The co-ordination index is shown in Appendix, Table C1. The average co-ordination rate shows a declining trend throughout the sample. As with the other indicators for the labour market, there is considerable variation between countries. The highest level of co-ordination in the period 2009 to 2012, equal to four, is found in Belgium, Germany, Italy, Japan and Norway. The lowest levels are found in Canada, the UK and the US.

Union density rates are constructed using the number of union members divided by the number of employed. Trade union density rates are based on surveys, wherever possible. Where such data were not available—in European Union countries, Norway and Switzerland—trade union membership and density were calculated using administrative data adjusted for inactive and self-employed members by Professor Jelle Visser, from the University of Amsterdam. Appendix Table C3 shows that union density has declined since the beginning of the 1970s in most countries. The exceptions are Norway and Belgium, where the measure has remained stable over the whole period, and Denmark and Finland, where the measure has increased.

We also investigate the interaction between the following pairs of institutional variables: benefit duration and the benefit replacement ratio; co-ordination in wage setting and union density; and co-ordination and the tax level. These interaction terms are measured as the deviations from country-specific means. For instance, the interaction between co-ordination and tax is equal to $(CO - \overline{CO})(TW - \overline{TW})$, where $\overline{CO}$ and $\overline{TW}$ are the country-specific means of that variable.

IV. Econometric model and estimation methods

The equilibrium rate, u^*, is an identified parameter of the final equation of the theoretical model, and the equilibrium dynamics of u_t follows from that equation. An adaptation of (5) to our case, with panel data and institutional variables affecting the equilibrium unemployment rate through the mark-up coefficients is:

$$u_{it} = \beta_{0i} + \beta_1 u_{it-1} + \beta_2 u_{it-2} + \beta_3 u_{it-3} + \beta_4 Z_{it-1} + \beta_5 Z_{it-2} + \varepsilon_{uit} \tag{10}$$

We have added the subscript i for country i, Z_{it} is a vector that contains the institutional variables for country i in year t, and β_4 and β_5 are corresponding (row) vectors of parameters. Z_{it} contains the six institutional factors, EPL, BRR, BD, $UDNET$, TW and CO.

The distributed lag in Z_{it} is motivated by theory, whereby an institutional reform in period t affects wage and price setting in $t + 1$ and this leads to an unemployment response in the final equation model in period $t + 2$. We also include the interaction terms just mentioned, which earlier studies (e.g., Belot and van Ours (2001); Nickell et al. (2005)) have shown to be of importance, but we do introduce explicit notation for those interaction terms in (10). To reduce collinearity, and to obtain a direct estimate on the "level effect" of an institutional variable that affects u^*, we estimate the model in terms of $\Delta Z_{i,t-1}$ and $Z_{i,t-2}$, which is a re-parameterization of (10) and does not affect the properties of the disturbances.

The number of cross-section units is 20, and the initial sample length is from 1960 to 2012; hence, $i = 1, 2, ... \ 20$ and $t = 1960, 1961, ..., 2012$. Because it is difficult to find consistent operational definitions of all the institutional variables, and because of dynamics, the sample used in the estimations is unbalanced and we lose a few annual observations. Typically, the longest time series is 50 observations, and the shortest is 46 observations.

We follow the study of OECD unemployment by Nickell et al. (2005) and use the within-group estimator (WG hereafter), also called the least squares dummy variable method, as the reference estimator. This reflects the fact that our main purpose is to estimate regression coefficients, and the equilibrium unemployment rate as a derived coefficient, free from unobserved heterogeneity bias. As the time dimension is relatively large, the sample realizations of the individual effects β_{0i} may be treated as parameters that are jointly determined with the common parameters in (10). In this setting, the WG bias will be small if the disturbances not are autocorrelated (see, e.g., Judson and Owen (1999)). A formalization of the key assumption of the fixed-effect model is as follows:

$$E(\varepsilon_{uit} \mid \beta_{0i}, u_i^{t-1}, Z_i^{t-1}) = 0 \tag{11}$$

where $u_i^{t-1} = (u_{i0}, u_{i1}, \ldots, u_{it-1})$ and $z_i^{t-1} = (Z_{i0}, Z_{i1}, \ldots, Z_{it-1})$. (11) implies that the disturbances are not autocorrelated.

While, generally, we would like assumption (11) to be accepted by the data (as it is important for the T-consistency of the estimators, and validates standard inference procedures at least as a guideline), our theory (9) implies that the disturbance ε_{ut} contains moving averages of the VAR disturbances as well as random shocks to import prices and productivity. Therefore, one may ask: does empirical acceptance of (11) contradict the relevance of the theory? However, of course, a failure to reject the null hypothesis of no autocorrelation with the use of standard misspecification tests does not prove that the theoretical disturbances are without autocorrelation. As illustrated in the numerical example in the Appendix A (available online), the misspecification tests do not reveal the autocorrelation in that artificial dataset. Equally important, the ordinary least squares (OLS) estimaton gives an accurate estimate of the u^* parameter in that one-off estimation. We would like to investigate the robustness of the WG estimator by using several different estimators. First, we use standard modifications of the WG estimator that allow for heteroscedasticity and autocorrelation, which are implied by our theory. Second, we present results from the difference generalized method of moments (GMM) estimator. This estimator addresses the underlying issue of finite sample bias of dynamic panel models (see Arellano (2003, Ch. 6.3) and (Baltagi, 2008, Ch. 8)).

In addition, we supplement the standard panel data estimators by the impulse indicator saturation (WG-IIS) estimator. This estimator has been shown to improve the size and the precision of the estimated coefficients of explanatory variables in econometric time series models; cf. Johansen and Nielsen (2009).

The indicator saturation divides the sample into two sets and saturates first one set, and then the other set with zero–one indicators for the observation. The indicators are tested for significance, taking the estimated indicator coefficients from the other half of the sample as given, and observations are deleted if the t-ratios are significant. In this way, the indicators are selected, and the number of indicators retained are often relatively few and interpretable (among others, see Hendry (1999) and Castle et al. (2014)). When the model is augmented by the selected individual-specific location-shift indicators and estimated by OLS, we obtain the IIS estimators of the original regression coefficients. When the error distribution is symmetric, and under the null hypothesis that there are no location shifts, the IIS estimator is centred around the same value as the OLS estimator; however, there is an efficiency loss because irrelevant explanatory variables are included in the model. Against that loss, there is the potential of gains both in the centring and in the efficiency under the alternative of breaks and/or outliers, which is empirically highly relevant for the actual unemployment rates, as we have seen.

Johansen and Nielsen (2009) extend the IIS estimator to stationary autoregressive distributed lags, but the literature still lacks a formalization of the panel data version of the IIS estimator. We propose to use the method to add robustness to the WG estimator. Concretely, the WG-IIS estimator is obtained by first applying within-group transformation of the dataset, and second, by impulse saturating the transformed dataset, as explained in the previous paragraph. In the algorithm for indicator saturation that we use, the original regressors (the autoregressive part and the institutional variables) are not selected over, see Doornik (2009), Doornik and Hendry (2009, Ch 14.8). As a further test of robustness, we also detect large outliers and represent them by indicator variables.

It might be noted that all our estimated models are augmented by time dummies in the same way as in Nickell et al. (2005) for example. In the present context, those dummies are interpretable as common location shifts in the unemployment rate, and location shifts that are identified from *Autometrics* represent heterogeneity in the location shifts. Of course, it is possible that no such dummies will be found to be significant and that the conventional times dummies are sufficient in terms of the robustness of the WG estimator of the regression coefficients.

V. Empirical results

We first give the results for the WG estimation of equation (10). Second, we present difference-GMM-based estimation results for a parsimonious version of equation (10). Finally, we show how the WG estimator changes when we account for shocks by using the WG-IIS estimator.

Results for the WG-based estimation

Table 1 contains the results for the WG estimator in four versions, first without any "whitening of the residuals", and then with robust estimation of the coefficient standard errors. There are no sign of autocorrelation in the error terms (see the test results in the lower part of Table 1, columns one and two). However, for completeness, we present feasible generalized least squares (GLS) estimators for the case of residual heteroscedasticity and autocorrelation, as indicated by the column headings in columns three and four in the same table.

The WG estimator shows that the two first lags of the unemployment rate have signs that are consistent with theory and are significant, as judged by the p-values of the coefficient. The third lag is insignificant and the estimated coefficient is negative in the "pure" WG estimation, but it is not robust in the two GLS estimations in columns three and four. The GLS estimators show that the third lag of unemployment is positive and it is significant at the 10 per cent level. These results are in line with the predictions from theory, in particular that the first autoregressive coefficient can be larger than one, and that the second is expected to be negative, but much smaller in magnitude.

When we look at the results for institutional variables there appears at first to be very little significance, in particular for the lagged levels variables that are important factors behind secular changes in the estimated equilibrium and long-run unemployment. The only level variables that stand out are the tax wedge (TW) and the interaction term between the benefit replacement rate (BRR) and benefit duration (BD). There is some evidence of a broader impact of institutional factors in the estimation results that control for heteroscedasticity and autocorrelation. If we apply a 10 per cent significance level, both co-ordination (CO) and the benefit replacement ratio (BRR) are now significant, in addition to the tax wedge (TW) and the interaction term between (BRR) and (BD). The signs of the estimated regression coefficients are reasonable: increased tax levels and an increased level and length of benefits increase equilibrium unemployment. The coefficient of co-ordination, which we showed was unsigned from theory (because of offsetting effects on the mark-ups in wage and price setting), has a negative regression coefficient in the estimation that corrects for autocorrelation. It is interesting to see if this empirical determination of the sign is robust when we compare the results with other estimation methods.

Difference-GMM-based estimation

As noted in section IV., the WG estimator of equation (10) would lead to biased and inconsistent estimates in the cross-section dimension even if the error term ε_{it} is not serially correlated.

An important alternative to the WG estimator relies on transforming model (10) to first differences to eliminate the individual effects. This leads to difference-GMM-based

TAB. 1. Within-group estimation results

	Dependent variable: Unemployment rate (u_{it}). Percent											
	WG			WG, robust[a]			WG heterosc.[b]			WG autocorr. [c]		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
u_{it-1}	1.37	0.03	0.00	1.37	0.12	0.00	1.43	0.03	0.00	1.39	0.03	0.00
u_{it-2}	-0.47	0.05	0.00	-0.47	0.20	0.02	-0.60	0.05	0.00	-0.57	0.05	0.00
u_{it-3}	-0.03	0.03	0.39	-0.03	0.10	0.78	0.05	0.03	0.10	0.06	0.03	0.07
Δ EPL$_{it-1}$	-0.26	0.25	0.31	-0.26	0.36	0.47	-0.36	0.21	0.09	-0.21	0.24	0.37
EPL$_{it-2}$	-0.06	0.07	0.37	-0.06	0.08	0.44	-0.01	0.06	0.85	-0.08	0.07	0.25
Δ BRR$_{it-1}$	-0.05	0.87	0.96	-0.05	0.53	0.93	-0.44	0.74	0.55	0.33	0.82	0.69
BRR$_{it-2}$	0.27	0.25	0.28	0.27	0.28	0.34	0.35	0.20	0.09	0.48	0.24	0.05
Δ BD$_{it-1}$	-0.27	0.52	0.61	-0.27	0.27	0.32	0.01	0.42	0.98	-0.43	0.49	0.39
BD$_{it-2}$	-0.15	0.17	0.35	-0.15	0.18	0.39	-0.10	0.13	0.45	-0.30	0.16	0.07
Δ Interaction - BRR and BD$_{it-1}$	3.39	1.88	0.07	3.39	2.04	0.10	2.56	1.49	0.09	3.54	1.77	0.05
Interaction - BRR and BD$_{it-2}$	1.90	0.60	0.00	1.90	0.65	0.00	1.41	0.50	0.00	2.12	0.59	0.00
Δ Interaction - CO and UDNET$_{it-1}$	0.12	0.21	0.56	0.12	0.25	0.62	-0.16	0.18	0.36	-0.16	0.20	0.43
Interaction - CO and UDNET$_{it-2}$	0.05	0.17	0.77	0.05	0.20	0.80	0.01	0.14	0.94	-0.08	0.16	0.62
Δ Interaction - CO and TW$_{it-1}$	0.43	0.32	0.17	0.43	0.37	0.25	0.69	0.27	0.01	0.42	0.30	0.15
Interaction - CO and TW$_{it-2}$	0.35	0.26	0.18	0.35	0.36	0.33	0.23	0.24	0.33	0.21	0.26	0.41
Δ UDNET$_{it-1}$	2.98	2.08	0.15	2.98	3.12	0.34	1.30	1.79	0.47	2.61	1.98	0.19
UDNET$_{it-2}$	0.32	0.32	0.32	0.32	0.31	0.30	0.13	0.30	0.66	-0.01	0.33	0.97
Δ CO$_{it-1}$	0.02	0.03	0.61	0.02	0.03	0.52	0.02	0.03	0.56	-0.02	0.03	0.49
CO$_{it-2}$	-0.03	0.03	0.39	-0.03	0.05	0.59	0.01	0.03	0.77	-0.05	0.03	0.05
Δ TW$_{it-1}$	1.98	1.64	0.23	1.98	2.31	0.39	2.52	1.42	0.08	1.58	1.50	0.29
TW$_{it-2}$	2.04	0.64	0.00	2.04	0.78	0.01	1.08	0.53	0.04	1.86	0.65	0.00
Tot. obs and the number of countries	994	20		994	20		994	20		994	20	
Standard errors of residuals	0.6			0.6			0.7			0.6		
χ^2 of all explanatory variables.[d]	34.95	(0.01)		452.86	(0.00)		32.35	(0.02)		35.10	(0.01)	
χ^2 of institutional variables (level).[d]	34.95	(0.01)		452.86	(0.00)		32.35	(0.02)		35.10	(0.01)	
χ^2 of institutional variables (interaction).[d]	14.35	(0.03)		14.95	(0.02)		16.02	(0.01)		16.53	(0.01)	
1st order autocorrelation[d]	1.15	(0.25)		0.56	(0.58)		.	(.)		.	(.)	
2nd order autocorrelation[d]	0.27	(0.79)		0.23	(0.82)		.	(.)		.	(.)	

Generalised least squares. Each equation contains country and time dummies.

a) Generalised least squares adjusting the standard errors to be robust to intragroup correlation.

b) Generalised least squares withing group allowing for heteroscedastic error terms.

c) Generalised least squares within group allowing for autocorrelated country specific error terms.

d) Numbers in parenthesis are p-values for the relevant null.

Variables:

The benefit replacement ratio (BRR), union density (UDNET) and employment tax wedge (TW) are proportions (range 0-1),
benefit duration (BD) has a range (0-1,1) employmetn protection (EPL) range (0-5) and coordination (CO) ranges (1-5).
All variables in the interaction terms are expressed as deviations from the sample mean.

estimation because, in our case, for example, Δu_{it-1} becomes correlated with $\Delta\varepsilon_t$ and needs to be instrumented (see Arellano (2003, Ch 6.3)).

The first available instrument correlated with Δu_{it-1} and uncorrelated with the error term (assuming no autocorrelation) is unemployment in period $t-2$. The instrument can also work as an instrument for later periods. The number of instruments is therefore quadratic in t. Differences can also be used as instruments. The first available instrument in differences is Δu_{it-2}, but this variable is highly correlated with the second endogenous variable Δu_{it-2}. The first instrument is Δu_{it-4}.

The use of difference-GMM on equation (10) is not straightforward when the panel consists of many time periods, as the number of instruments then becomes very large. This problem is known as instrument proliferation (see, for instance, Bowsher (2002) and Roodman (2009)). This problem affects both the one- and two-step estimators. Disregarding the loss of possible instruments resulting from multicollinearity problems, the instruments become quadratic in T and, in our case, the number of available instruments is 47^2. The literature notes three problems with many available instruments, and this concurs with our experience: First, we find that applying the difference-GMM method overfits the endogenous variable, and the GMM result approaches the OLS results. Second, the number of sample moments used to estimate the optimal weighting matrix for the identifying moments between the instruments and the errors, $var[z'\epsilon]$, is equal to 47^4, which also causes a bias in the Sargan and Hansen tests, where the Sargan test is always rejected whereas the Hansen test has implausibly high p-values (see Roodman (2009) and Bowsher (2002)).

To work around these issues, we have followed the recommendations in Roodman (2009). Specifically, we have reduced the number of moment conditions by restricting the number of instruments in two ways. First, the number of lags available as GMM instruments is reduced to 4. Second, the number of moments to be estimated is reduced by collapsing the instrument matrix, which reduces the number of variances per country.

Despite taking these measures to reduce available instruments, the GMM estimation results in too large a difference between the WG estimator and the GMM estimator, i.e., the difference cannot be interpreted as only being a correction of the finite i bias in the fixed-effect model[2]. For instance, the estimated coefficient of the lagged endogenous variable changed by a factor of three when the collapse option was used in the estimation. On the other hand, the estimated values of the institutional variables were close to the WG estimator. If the number of lags is reduced, the lagged endogenous variable is close to the WG estimator, whereas there are large changes in the numerical values of the estimates of the exogenous variables; for instance, the interaction between benefit replacement and benefit duration changed sign (see Table 2). Therefore, the reduction in available instruments is probably insufficient to obtain good estimators in this complex model. Furthermore, in this case, the Hansen test indicated rejection, with an implausibly high p-value, equal to one. Note also that the theory gives no clear guidelines as to how the number of lags available as instruments can be reduced before the "collapse" function is applied to the model. We have attempted using a different number of lags, but all choices reveal the erratic behaviour of the estimates and large variations (not reported because of space considerations).

Our results might suffer from a weak instrument problem caused by the transformation to first differences when the unemployment rate is close to a random walk (see Mátyás and Sevestre (2008, Ch. 8)). Then, past unemployment levels convey little information about future changes in the same variable, and untransformed lags may be weak instruments for transformed variables. If past changes are better predictors of current levels than past levels are of current changes, then new instruments are more relevant.

In addition, several of the WG estimators of the coefficients in equation (10) are

[2]All results from the difference-GMM method are available upon request

Tab. 2. Within-group and difference-GMM estimation results

	Dependent variable: Unemployment rate (u_{it}). Percent								
	WG			GMM, one step			GMM, two step		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
u_{it-1}	1.40	0.03	0.00	1.24	0.03	0.00	1.22	0.10	0.00
u_{it-2}	-0.59	0.05	0.00	-0.47	0.05	0.00	-0.46	0.12	0.00
u_{it-3}	0.06	0.03	0.05	-0.00	0.03	0.96	0.01	0.11	0.91
Δ EPL_{it-1}	-0.17	0.23	0.47	-0.25	0.32	0.43	-1.55	1.57	0.32
BRR_{it-2}	0.58	0.23	0.01	3.25	0.59	0.00	7.67	16.15	0.63
BD_{it-2}	-0.23	0.15	0.14	-0.62	0.35	0.08	-1.54	3.95	0.70
Δ Interaction - BRR and BD_{it-1}	2.88	1.55	0.06	-1.47	2.23	0.51	-3.57	13.64	0.79
Interaction - BRR and BD_{it-2}	1.95	0.56	0.00	-0.55	1.46	0.70	9.07	25.40	0.72
Δ Interaction - CO and TW_{it-1}	0.23	0.25	0.35	0.55	0.31	0.08	1.65	1.73	0.34
CO_{it-2}	-0.04	0.02	0.04	-0.04	0.04	0.29	-0.02	0.08	0.82
Δ TW_{it-1}	1.65	1.49	0.27	-2.95	2.01	0.14	-11.82	16.61	0.48
TW_{it-2}	1.83	0.61	0.00	1.04	1.05	0.32	-0.03	6.66	1.00
Tot. obs and the number of countries	994	20		974	20		974	20	
Standard deviation of residuals	0.65			0.73			0.77		
χ^2 of all explanatory variables.[a]	29.33	(0.00)		76.40	(0.00)		8.00	(0.53)	
χ^2 of institutional variables (level).[a]	25.90	(0.00)		68.42	(0.00)		0.25	(1.00)	
χ^2 of institutional variables (interaction).[a]	15.31	(0.00)		3.55	(0.31)		0.93	(0.82)	
1st order autocorrelation[a]	1.07	(0.29)		-24.87	(0.00)		-2.18	(0.03)	
2nd order autocorrelation[a]	-0.04	(0.97)		-0.11	(0.91)		0.09	(0.93)	
Sargan test[a]				934.51	(0.00)		934.51	(0.00)	
Hansen test[a]							16.38	(1.00)	

a) Numbers in parenthesis are p-values for the relevant null.

Variables:

The benefit replacement ratio (BRR), union density (UDNET) and employment tax wedge (TW) are proportions (range 0-1), benefit duration (BD) has a range (0-1,1) employmetn protection (EPL) range (0-5) and coordination (CO) ranges (1-5).

All variables in the interaction terms are expressed as deviations from the sample mean.

insignificant. The implication of applying the difference-GMM method on a model with insignificant variables is not clear. In practice, a straightforward implementation of the difference-GMM means that we instrument the lagged unemployment rate with many weak instruments. It is not clear how this method will affect the weighted matrix or how the resulting estimates should be interpreted. In order to be able to compare the WG and difference-GMM-based estimators, we have therefore chosen to report the WG estimator, and the one- and two-step GMM estimators for a simplified model, where only significant variables are entered as explanatory variables that do not use the collapse function.

Table 2 contains the results from models where we have omitted the institutional variables that regularly have p-values > 0.10 in the WG-based estimations. Results for the one- and two-step Arellano–Bond estimation are reported with lagged levels of unemployment as GMM instruments. The main impression is that there are small differences between the results for the two estimation methods, and that all variables have the same sign (compare the results under "WG", "GMM, one step" and "GMM, two step" in Table 2). Taken at face value, this shows that the "bias problem" of the WG estimator does not constitute a major issue for the parsimonious model. This is as expected for a sample like ours, where the time series are quite long and there are no roots "on" the unit circle.

Results for the WG-IIS estimator

We suggest using an estimator with location-shift dummies as a robust WG estimator. The approach is analogous to the analysis of the IIS estimator for dynamic regression models in time series, where Hendry et al. (2008) and Johansen and Nielsen (2009) have shown that exclusion of shocks might bias the estimators.

Structural breaks are identified empirically by two automatized procedures, dubbed IIS and large outliers, see Doornik (2009). The first procedure is the IIS method. This selection algorithm first adds dummies for each year to the model and then selects

TAB. 3. Years with location shifts. Results from impulse saturation and large outliers, no selection of lags of unemployment and the institutional variables in Table 1

	Impulse Indicator Saturation (2.5 %)	Large Outlier (5 %)
Australia	1975 1976 1977 1982 1983 1984 1990 1991 2009	1984 1991
Austria	2009	
Belgium	1975 1981 1993 2002 2009 2011 2012	
Canada	1970 1975 1977 1982 1991 2009 2010	1982 2009
Denmark	1975 1976 1979 1981 1985 1986 1994 1997 2005 2009	1994 2009
Finland	1967 1969 1979 1991 1992 1993 1997 2009	1979 1991 1992 1993 1997 2009
France	1984 1996 2000 2009	2009
Germany	1967 1968 1975 1981 1982 1983 1993 1996	
	2002 2003 2005 2009	
Ireland	1966 1967 1971 1972 1975 1976 1977 1978 1979 1980 1981	1967 1972 1976 1978 1981 1991 1998 2009
	1982 1983 1986 1990 1991 1993 1995 1997 1998 2008 2009 2011	
Italy	1986 2012	
Japan	1970 1984 1985 1990 1991 1992 1997 2009	
Netherlands	1981 1982 1983 1993 2012	
New Zealand	1983 1984 1988 1989 1991 1995 2009	1991
Norway	1976 1989 2006	
Portugal	1970 1975 1976 1987 1993 1998 2009 2011 2012	1980 2009 2012
Spain	1978 1980 1981 1982 1983 1984 1988 1992 1993 1999	1984 1993 2008 2009 2012
	2002 2008 2009 2011 2012	
Sweden	1971 1991 1992 1993 1996 1998 2009	1993
Switzerland	1976 1981 1991 2009	
UK	1980 1981 1988 1991 2009	1981
USA	1970 1974 1975 1976 1980 1982 1984 1985 1991 2008 2009	1976 1982 2009

automatically to produce a final model with a smaller set of significant structural breaks. The properties of the algorithm of automatized variable selection are documented (e.g., Doornik (2009) and Castle et al. (2012)). The second procedure is the large outliers selection algorithm, which adds dummies for the years with significant outliers.

These two selection algorithms are not pre-programmed for panel data models, but we have worked around this by using the within-transformation on all the data in equation (10). Impulse saturation and large-outlier selection are then applied to the transformed dataset. The set of institutional variables, and the three lags of the unemployment rate are not selected over. We used a significance level of 5 per cent when large outliers was used, and of 2.5 for impulse saturation, Doornik (2009). As shown in Table 3, very few outliers were found to be significant, while impulse indicator saturation retained a quite large number of indicators for some countries. Ireland and Spain are examples of countries with high levels of unemployment in some years and high volatility over time. In the results, and IIS in particular, these two countries stand out, with a large number of retained indicator variables.

When we look at the structural breaks across the countries, we find that most of the estimated breaks are interpretable with reference to known events and shocks in economic history. For example, the years 1980–1983 are represented by one or more indicators in the unemployment rates of 12 countries (ten European countries plus Australia and the US). These years followed the stagflation in the 1970s, the two oil-price shocks, widespread closures in traditional manufacturing in many OECD countries and a marked tightening of monetary policy in the US.[3] The oil-price shocks are represented by separate dummies in the results for several countries, the US in particular. Another concentration of breaks occur between 1989–1992. In the case of the three Scandinavian countries, Finland, Norway and Sweden, the causes of these rises in unemployment involved a housing price crash and severe banking crises (Norway and Sweden), and the collapse of trade with the Soviet Union (Finland). The current job crisis, with its origins in the credit crisis that started in 2008, is represented by indicator variables for all countries for 2009, except Norway.

Although the majority of the indicators in Table 3 represent increases in the rate

[3]The Federal Reserve increased the interest rate to 20 per cent early in the 1980s.

of unemployment, there are also examples of intermittent reductions. Some of these represent the effects of the well-known housing and credit market booms (for example, those that occurred in the UK in 1988 and in Norway in 2006). As mentioned, there are also effects of "bubbles" that burst at a later stage, for example, in the UK in 1991 and the US in 2008.

The estimation results in the column labelled WG-IIS in Table 4 are comparable with the WG in Table 1 above. The estimated residual standard error of WG-IIS (0.3) is a good deal lower than the WG residuals (0.6) (see the lower part of Table 4 and Table 1). This is also illustrated by comparing the residuals of the two models in Appendix, Figures B3 and B4 (available online). Lower standard residual errors make interval forecasts more precise. For the institutional variables, one change is that the interaction terms between BRR and BD are no longer significant when WG-IIS is used, showing that the WG estimate for this interaction term is not robust. The second column of Table 4 uses robust estimation of the coefficient standard errors. The interaction term between CO and TW is included in the group of significant institutional variables with this approach. The third column in Table 1 is a variant, where we treat the coefficients of the location-shift dummies as known, and combine them into one single variable before estimation by WG. There are no large changes in the results obtained with this estimator.

As assumed, there are no substantial changes when we assume that the error term follows an autoregressive process, nor in the estimated coefficients for the explanatory variables (see Table 4, columns one and four). There are, however, some more significant explanatory variables when the error term is allowed to follow an autoregressive process.

TAB. 4. Within-group impulse indicator saturation estimates, WG-IIS

	Dependent variable: Unemployment rate (u_{it}). Percent											
	WG-IIS			WG-IIS, robust[a]			WG-IIS stacked[b]			WG-IIS autocorr. [c]		
	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value	Coef.	Std	p-value
u_{it-1}	1.38	0.02	0.00	1.38	0.04	0.00	1.39	0.03	0.00	1.37	0.02	0.00
u_{it-2}	-0.54	0.04	0.00	-0.54	0.07	0.00	-0.55	0.05	0.00	-0.53	0.03	0.00
u_{it-3}	0.08	0.02	0.00	0.08	0.03	0.02	0.09	0.02	0.00	0.08	0.02	0.00
Δ EPL$_{it-1}$	-0.06	0.16	0.70	-0.06	0.20	0.76	-0.07	0.17	0.69	-0.08	0.13	0.55
EPL$_{it-2}$	-0.04	0.04	0.42	-0.04	0.06	0.56	-0.02	0.04	0.64	-0.03	0.04	0.39
Δ BRR$_{it-1}$	-1.18	0.54	0.03	-1.18	0.46	0.01	-1.19	0.39	0.00	-1.25	0.47	0.01
BRR$_{it-2}$	0.18	0.16	0.27	0.18	0.24	0.46	0.13	0.17	0.45	0.20	0.14	0.15
Δ BD$_{it-1}$	-0.12	0.32	0.72	-0.12	0.35	0.74	-0.14	0.28	0.62	-0.17	0.27	0.54
BD$_{it-2}$	-0.07	0.11	0.53	-0.07	0.12	0.59	0.01	0.10	0.92	-0.04	0.09	0.66
Δ Interaction - BRR and BD$_{it-1}$	-0.57	1.16	0.62	-0.57	0.99	0.56	-1.09	0.84	0.19	-0.45	1.00	0.65
Interaction - BRR and BD$_{it-2}$	0.51	0.38	0.19	0.51	0.42	0.23	0.43	0.34	0.20	0.50	0.34	0.14
Δ Interaction - CO and UDNET$_{it-1}$	0.45	0.15	0.00	0.45	0.17	0.01	0.39	0.11	0.00	0.43	0.13	0.00
Interaction - CO and UDNET$_{it-2}$	0.19	0.11	0.07	0.19	0.09	0.04	0.13	0.06	0.03	0.17	0.09	0.06
Δ Interaction - CO and TW$_{it-1}$	0.01	0.22	0.96	0.01	0.23	0.96	0.18	0.16	0.26	-0.04	0.19	0.82
Interaction - CO and TW$_{it-2}$	-0.29	0.17	0.10	-0.29	0.11	0.01	-0.24	0.08	0.00	-0.30	0.15	0.04
Δ UDNET$_{it-1}$	1.19	1.30	0.36	1.19	2.06	0.56	1.41	1.64	0.39	1.39	1.12	0.22
UDNET$_{it-2}$	0.14	0.21	0.51	0.14	0.18	0.45	0.17	0.12	0.15	0.08	0.18	0.65
Δ CO$_{it-1}$	0.00	0.02	0.97	0.00	0.03	0.98	0.01	0.02	0.68	-0.00	0.02	0.92
CO$_{it-2}$	-0.06	0.02	0.00	-0.06	0.03	0.06	-0.04	0.02	0.02	-0.06	0.02	0.00
Δ TW$_{it-1}$	1.88	1.05	0.07	1.88	1.58	0.23	1.93	1.36	0.16	1.85	0.89	0.04
TW$_{it-2}$	0.41	0.39	0.30	0.41	0.58	0.49	0.32	0.45	0.48	0.28	0.35	0.42
Tot. obs and the number of countries	994	20		994	20		994	20		994	20	
Standard errors of residuals	0.3			0.3			0.3			0.3		
χ^2 of all explanatory variables.[d]	38.54	(0.00)		32003.78	(0.00)		2453.08	(0.00)		53.07	(0.00)	
χ^2 of institutional variables (level).[d]	38.54	(0.00)		32003.78	(0.00)		2453.08	(0.00)		53.07	(0.00)	
χ^2 of institutional variables (interaction).[d]	15.87	(0.01)		17.35	(0.01)		55.87	(0.00)		20.13	(0.00)	
χ^2 of impulse saturation shocks.[d]	2069.84	(0.00)		5774.07	(0.00)		2954.79	(0.00)		2752.94	(0.00)	
1st order autocorrelation[d]	1.09	(0.28)		0.89	(0.28)		1.06	(0.29)		.	(.)	
2nd order autocorrelation[d]	-0.54	(0.57)		-0.57	(0.57)		-0.88	(0.38)		.	(.)	

Generalised least squares. Each equation contains country and time dummies.

a) Generalised least squares adjusting the standard errors to be robust to intragroup correlation.

b) Generalised least squares with the IIS estimator as one variable.

c) Generalised least squares within group allowing for autocorrelated country specific error terms.

d) Numbers in parenthesis are p-values for the relevant null.

Variables:

The benefit replacement ratio (BRR), union density (UDNET) and employment tax wedge (TW) are proportions (range 0-1), benefit duration (BD) has a range (0-1,1) employmetn protection (EPL) range (0-5) and coordination (CO) ranges (1-5).

All variables in the interaction terms are expressed as deviations from the sample mean.

The average fit of unemployment with the WG-IIS estimator from column one in

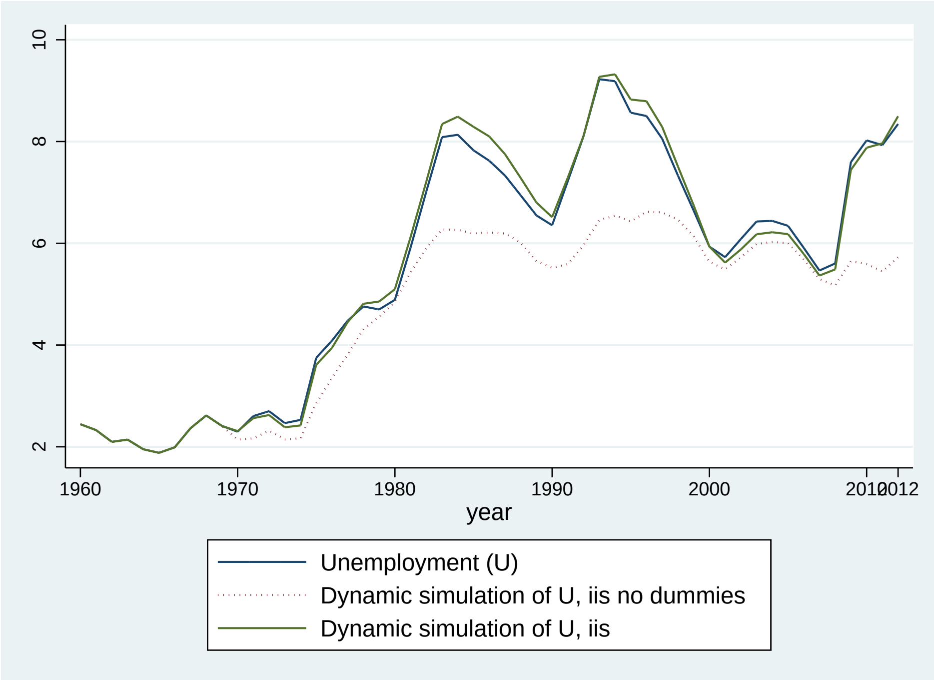

Fig. 2. Simulated unemployment rates in the OECD countries. WG-IIS estimates are used for the coefficients. Per cent

Table 4 is illustrated in Figure 2, together with a simulation without the impulse indicator saturation, but with the time dummies. As we start the simulation in 1970, the dotted line shows $\hat{u^*}$ over the sample. The figure shows that the simulated average unemployment increased by nearly 4 percentage points from the 1970s to the mid-1980s and has remained around 6 per cent throughout the sample period. There is no sign of a decline in the simulated equilibrium rate towards the end of the sample period, as one would expect from the fact that several countries then lowered the level of several of the included institutional variables; cf. in particular, the development of co-ordination, the benefit replacement ratio and tax wedges in Appendix Tables C1, C4 and C6 (available online). A simulation for each country in the sample is presented in Appendix, Figures B1 and B2 (available online).

VI. Model-based empirical equilibrium unemployment

Table 5 shows the WG-based estimated coefficients of the institutional determinants of the equilibrium rate u^*. The standard WG estimation results are to the left of the table, and the two WG-IIS versions are presented in columns two and three. One interesting finding is that the use of the robust estimator changes the results for CO. The coefficient of CO is insignificant with the WG estimation, but it is larger (in absolute value) and significant in the two WG-IIS estimations. There is also a notable sign change in the results for the interaction term between CO and TW, which switches from positive (WG) to negative (WG-IIS). Other coefficient estimates differ quite a lot between the estimation methods, although the signs are preserved. The estimated coefficients of TW, and the interaction term between BRR and BD are the main other examples of this. In sum, it appears that the main effect of changing from WG to WG-IIS is the centring of the estimates rather than the estimated standard errors.

The economic interpretation would have carried more weight if the precision of the estimates had been better. The hypothesis that receives most support from our results is

Tab. 5. Estimated long-run coefficients of the institutional variables

	WG (Tab. 1)		WG-IIS (Tab. 2)		WG-IIS stacked (Tab. 2)	
	Coeff.	St.err.	Coeff.	St.err	Coeff.	St.err.
EPL	-0.50	0.568	-0.413	0.50	-0.25	0.46
BRR	2.12	2.05	2.00	1.83	1.59	1.64
BD	-1.21	1.36	-0.76	1.21	0.11	1.10
UDNET	2.52	2.63	1.58	2.35	2.06	2.11
CO	-0.20	0.23	-0.64	0.21	-0.54	0.18
TW	16.09	5.26	4.65	4.46	3.93	4.18
BRR,BD interaction	14.93	4.80	5.81	4.24	5.34	3.90
CO,TW interaction	2.78	2.18	-3.28	2.06	-2.98	1.79
CO,UDNET interaction	0.38	1.35	2.21	1.18	1.65	1.06

that a higher degree of co-ordination in wage formation, over a period of time, lowers the equilibrium rate of unemployment. This is supported by the significance of the WG-IIS estimated coefficient of the variable, as well as by the other robust estimation methods in Table 5.

Another significant result suggesting the need for reform—although the supportive evidence is more in terms of a tendency when we average across estimation methods—is that lower compensation levels can reduce unemployment in the longer term; see both the coefficients of BRR and the interaction terms with benefit duration (BD). Third, the coefficient of the average tax wedge is consistently positive, although it is scaled down with WG-IIS estimation and with a t-value a little above or below one.

In Figure 3, we show the results (again for the average OECD country) of three simulations, where all the explanatory variables are the same, but where the coefficients estimates are from WG, WG-IIS and WG-IIS stacked. The simulations start in 2013 and end in 2014. The institutional variables are prolonged into the simulation period with their end-of-estimation sample values, to focus on the impact of the coefficient estimates on the simulated equilibrium rates. The graphs show that the two versions of the WG-IIS estimate provide some more grounds for optimism about the level of the equilibrium rate conditioned by the "no-change-in-institutions" situation. As we might expect, the 2012 level of the unemployment rate is as much as 2 percentage points higher than the estimated equilibrium rate.

Figure 4 shows the simulations results for the individual countries in four panels. The categorization is conducted mainly with reference to the institutional characteristics. In panel a), the UK stands out as the only country with above-average simulated equilibrium unemployment. The US, despite experiencing increased unemployment during the credit crises, has the lowest estimated u^*, together with New Zealand. Panel b shows that, for the Nordic countries, Denmark and Finland have estimated equilibrium levels above the average, whereas Norway and Sweden have lower estimated u^*s. Panel c shows the so-called PIIGS (Portugal, Ireland, Italy, Greece and Spain) countries with high current and estimated long-run unemployment rates. It is interesting to note that, although Spain performed worst at the end of the sample period, Spain's equilibrium rate is about the same as Italy's and Portugal's. Finally, panel d shows that France and Belgium have the largest estimated long-run unemployment rates in continental Europe. Germany, the Netherlands, Belgium and, in particular, Switzerland and Austria are all estimated to "land" well below the average u^* in our sample. Appendix B (available online) contains graphs for each country.

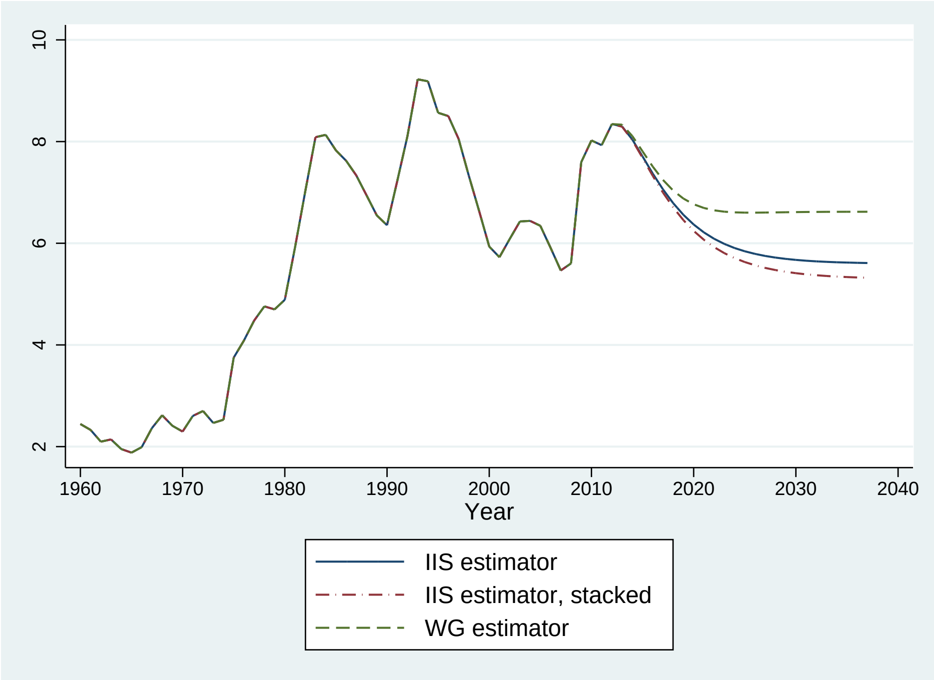

Fig. 3. Average simulated equilibrium unemployment rates for the sample of OECD countries. WG and WG-IIS estimators. Per cent

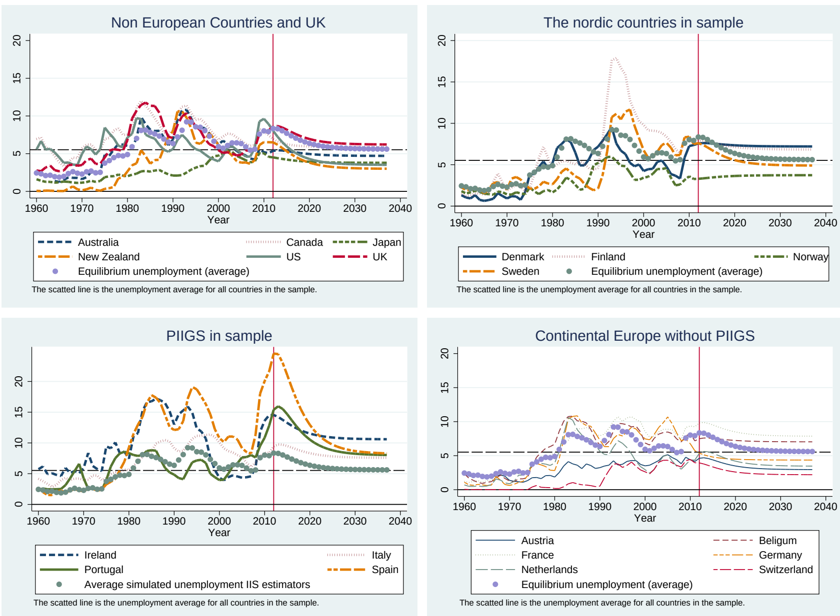

Fig. 4. Country specific simulated equilibrium unemployment rates, based on WG-IIS estimation. Actual unemployment rates shown for comparison

VII. Summary and discussion

The equilibrium rate of unemployment is an important parameter in economic models that are used for forecasting and as an aid for policy analysis and decision making. Realistic estimates of equilibrium rates are sought by those responsible for monetary and fiscal policy planning, and by analysts of the macroeconomic performance of different national economies. There is a range of methodological approaches to the estimation of equilibrium rates of unemployment, which is a positive because no single line of research is likely to be able to give a complete picture or to be without weaknesses.

In this paper, we have contributed to the literature that uses panel data information about institutional features in several OECD countries to estimate dynamic models for the rate of unemployment. Compared with earlier studies that use the Layard–Nickel model as a reference, we have provided a more detailed theoretical determination of the equilibrium rate, where the steady state is the solution to the dynamic model that determines the real-wage share and the real-exchange rate, jointly with unemployment. Estimation of this system, either as a structural model or as a VAR, is a promising option for future work, but in this work we have chosen to follow the earlier panel data literature and estimate single-equation models for the rate of unemployment. Unlike the earlier studies, the dynamics and inclusion of the institutional variables in our final equation follow precisely from the theoretical VAR model.

In terms of an estimation method, we propose that there is a role for the IIS method—which has been developed in time-series econometrics as a robust estimator—in the estimation of panel dynamic data models with many time-series observations. In our case, the impulse indicator saturation estimator is a robust version of the standard WG estimator. It is clearly related to the WG estimator with time dummies because WG-IIS includes only significant time dummies for individual countries after selection. We propose that this gives a relevant method for modelling, e.g., heterogeneous responses to common and national shocks.

The empirical results confirm that the role for higher-order dynamics may have been underestimated in earlier studies, which use first-order dynamics (before "autocorrelation correction"). This allows for complex dynamics while, at the same time, our estimation results are supportive of stable roots, so that all our models have a well-defined equilibrium rate.

The results that we report confirm earlier findings that labour market institutions are important for the estimated OECD equilibrium unemployment rate, but our results suggest that both the size and degree of precession may have been overestimated with the sample used by the earlier studies. In addition, changes in operational definitions may have played a role. Our own empirical evidence supports that the conclusion that, in particular, improved co-ordination in wage setting is a significant dimension of labour market reform that, over time, may reduce the level of unemployment. The tendency in the empirical evidence also supports the view that the generosity of the unemployment insurance system has been a driver of equilibrium unemployment over the sample. The tax wedge belongs to the same category of variables, whereas there is little evidence in our results that employment protection has been an important determinant of unemployment.

It is not surprising that our model shows that the unemployment in all countries (except Norway) was above the estimated equilibrium rate, and the explanation is found in the impact on the real economy of the credit crisis. Therefore, a simple prediction of our model is that actual rates of unemployment are likely to fall towards the estimated equilibrium levels, but only in the absence of new negative shocks.

A more interesting prediction is that it is indeed possible to reduce the long-run level of unemployment through institutional reform. It is true that not all the coefficients of the institutional variables are significantly different from zero, but the weight of the evidence (across estimation methods) validates such a conclusion.

Hence, advocates of the thesis that there is no viable alternative to labour market reforms for today's unemployment-stricken countries, find support for their view in our study. However, there is another interpretation of our results that is also relevant. It is that while reforms of labour market institutions are important, it remains essential to have policy instruments "ready" to be able to repair the damages of collapses in other markets or countries, before they get transformed into job and incomes crises.

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Additional information

Additional supporting information may be found in the online version of this article:

Appendix A **Additional notes and results for the theoretical framework of section II.**

Appendix B **Additional results from the estimation of equation (10)**

Appendix C **Data appendix**

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