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Equilibrium Unemployment Dynamics in a Panel of OECD Countries

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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 to existing studies, the resulting final equation for unemployment has 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 is not previously used on panel data. WG-IIS robustifies the estimator of the regression coefficients in the dynamic model, and it affects the estimated equilibrium unemployment rates. We find that wage coordination stands out as the most important institutional variable in our data set, 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.

Keywords: *OECD area unemployment, dynamics, structural breaks, equilibrium unemployment, wage setting, NAIRU , labour market institutions, automatic variable selection*

JEL classification: *C22,C23, C26,C51, E02, E11 E24*

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1 Introduction

The concept of equilibrium unemployment in the OECD area has been subject to 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), equilibrium of real wages require that unemployment becomes equal to the Non-Accelerating Inflation Rate of Unemployment (NAIRU). The equilibrium unemployment rate is not interpretable as a complete 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 for several empirical studies, i.e., Nickell et al. (2005a), Belot and van Ours (2001) and Bassanini and Duval (2006) see Nickell et al. (2005b) and Blanchard and Wolfers (2000), and the dynamic specification of the estimated model. Our theoretical model leads to a cointegrated 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 since it is determined jointly with a constant rate of inflation. Our main parameter of interest is the equilibrium rate of unemployment and not 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 which is implied by the VAR.

The theoretical model gives some guidelines for the empirical specification which are different from the conventional empirical literature. First, the autoregressive order of the final equation is shown to be three, while 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. Secondly, 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 in the 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 affect unemployment rates gradually and 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 on a panel data set, 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 coordination, 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 a estimation method that control 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 to become an international job crisis. To account for both national and common shocks we have augmented the model by 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 to investigate the effect of shocks in the equilibrium unemployment rates.

We consider that our findings are in line with the earlier panel data analysis where equilibrium unemployment is determined by labour market institutions. Specifically, Nickell et al. (2005b) 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 due to institutional changes. However, we do not find any evidence of a declining average equilibrium unemployment in the last part of our sample period, which could have been a consequence of that many countries have accommodated the recommendations from the OECDs job marked study and reduced the level of the labour market institutions. In our results the average equilibrium unemployment rate has remained high so far in the 2000s.

The literature is inconclusive about which particular institutional factors that are most important for the equilibrium unemployment rate. This is also our experience, since different estimation methods points to somewhat disparate results. However, the weight of the evidence point in the direction of coordination (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. The degree of employment protection on the other hand is not a key determinant, and contradicts arguments in the resent debate of what causes high unemployment rates.

Maybe we should not be too disappointed to find that not a broader specter of institutional variable are found to be robust across estimators and significant, this is as 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 can have canceling effects on unemployment. Bjørnstad and Kalstad (2010) observe that if the price mark-up coefficient is positively linked to the degree if coordination, the net effect of increased coordination on unemployment may be ambiguous. Following Bowdler and Nunziata (2007) and Bjørnstad and Kalstad (2010) we model price formation as function of wage determining factors and find that the effect of coordination is indeed 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 2. The data for the evolution of labour market institutions are presented in section 3. Econometric estimation issues and the method for deriving location shifts on panel data are briefly discussed in section 4. The results from the estimated dynamic unemployment equations, which include first only institutional variables and dynamics, and then also by controlling for location shifts, are presented in section 5 and the estimated equilibrium unemployment rates in section 6. We summarize in section 7 where we also discuss some extensions and give suggestions for further work.

2 A framework for equilibrium unemployment

In order to be able to estimate and test hypothesis about the equilibrium rate of unemployment (u^*), it is helpful to specify a medium-run dynamic macro model for open economies.

2.1 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

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

re_t is defined in such a way that an increase in the real exchange rate leads to improved competitiveness. This increases export, and thereby GDP increases and unemployment falls, hence $\rho \geqq 0$. $\epsilon_{u,t}$ contains all other variables which 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 main point in the Layard-Nickell model. In (1) the link to the supply-side and wage and price setting is represented by the real exchange rate re_t. In appendix A.1 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 as cointegrating relationships. The model includes nominal trends in foreign prices and the nominal exchange rate as well as a real trend in labor productivity. These trends add realism to the model, since unit-roots test typically do not reject when applied to wages and prices. On the other hand, logical consistency requires cointegration, since unemployment is stationary only when the real exchange rate is without stochastic trends. Stochastic trends are however, present in the processes for the nominal price levels and for the exchange rate.

As shown in the appendix, the wage-price model can be written as a dynamic system for the logarithm of the wage share ws and the real exchange rate (re). When we combine this result with the equation for the rate of unemployment (u_t) we obtain the VAR:

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

The two first 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 appendix A.1. The third row contains the parameters of (1). For $\mathbf{y}_t$ to be stationary, the eigenvalues of $\mathbf{R}$ must have moduli outside the 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 appendix A.1.

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

The left column of $\mathbf{P}$ contains the three intercepts, of which d and δ are reduced form, given by the expressions in the appendix. The vector $(\epsilon_{re,t}, \epsilon_{prw}, \epsilon_{u,t})$ contains the VAR disturbances which are reduced form expression of the structural disturbances. The expressions are in the appendix.

2.2 Equilibrium rate of unemployment

The stable steady-state solutions of the endogenous variables correspond to their unconditional expectations.[1] For the rate of unemployment in particular, we define the equilibrium rate by $u^* = E(u_t)$. It is given by

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

[1]When the moduli of the characteristic roots of $\mathbf{R}$ are inside the unit circle, the VAR is covariance stationary, and the deterministic version of the system is also globally asymptotically stable.

$\mathfrak{c}_{ss}$ 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 appendix A.1):

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

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

It is a main thesis in the Layard-Nickell model that increased mark-ups leads to 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 co-integrating wage- and price curve relationships. However, in our model, also the intercept in the aggregate demand equation c_u affects the equilibrium rate u^* directly. As explained in the appendix, $\omega \geq 0$ 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 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 coordination in wage setting and the real wage, Bowdler and Nunziata (2007). In our model this entails that m_w is a declining function of coordination. The price mark-up m_q have 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 multi-country data and find that the price mark-up is significantly higher in countries with a high degree of wage coordination compared to uncoordinated countries. This may dampen the negative effect of coordination on unemployment, that would otherwise be a consequence of reduced wage mark-up. Bowdler and Nunziata (2007) investigated the effect of coordination on inflation, and show that coordination 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 evolution in equilibrium unemployment through the two mark-up coefficients. This is the same hypothesis that Nickell et al. (2005b) investigated. Our contribution is that the dynamics of the model are made explicit, we have reviewed and revised the operational measure of institutional attributes, and time series are longer than 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 a way that we explain in section 4 to represent such shifts. At this point it might be noted however, that even if the breaks in c_u are non-permanent (not a step-function) we expect that they will have an effect on equilibrium unemployment dynamics, as (2) show.

2.3 The final equation model

Multi-equation econometric models of wage- and price setting exists for single countries, see Bårdsen and Nymoen (2003) and Akram and Nymoen (2009) (Norway), Bårdsen and Nymoen (2009b) (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 for future work to take this approach to panel data, anticipating that progress can be made by 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.1 that β_1 is positive, and that it may well be larger than one. The second autoregressive parameter is expected to be negative, since all the coefficients inside the brackets are positive from theory. We note that it may be reasonable that $\beta_1 > -\beta_2$, since 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. In our framework, the equilibrium rate of unemployment is therefore uniquely determined. 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

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

since the low frequency characteristic root is inside the unit-circle, as a consequence of stationarity.[2] (6) shows how the autoregressive coefficients depend on the underlying parameters, (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 to the single equation models used in the existing panel data studies of unemployment and institutions cited in the introduction.

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

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

hence 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 4 we discuss the estimation methods we have used to estimate (5) on the panel data set that we present in the next section.

[2] As seen from (7) and the equation in the appendix, $e = \xi = 0$, in the case of dynamic price-wage homogeneity. This confirms the results 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

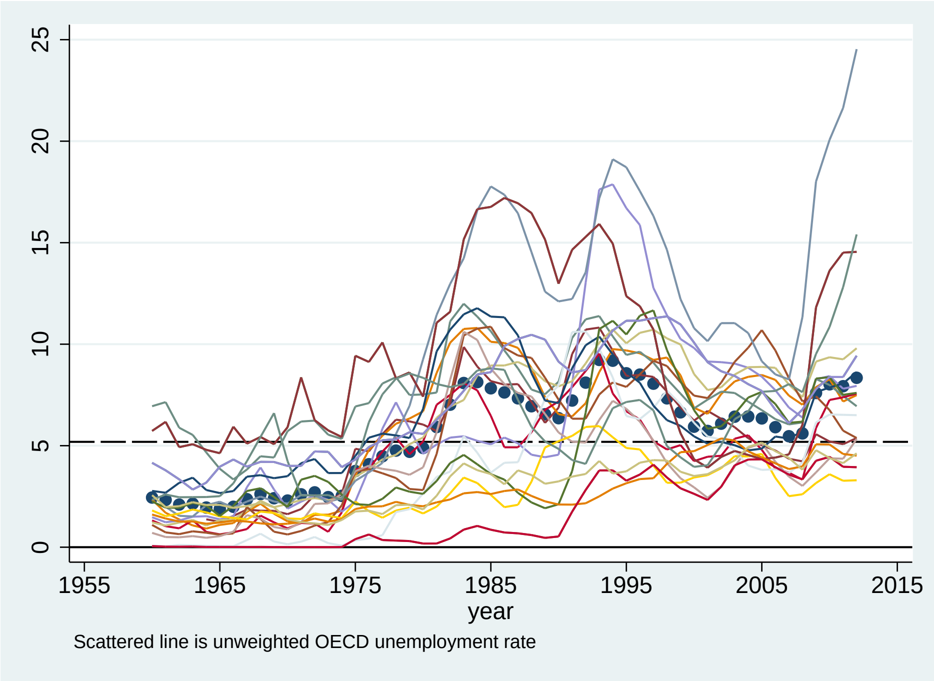

Figure 1: Unemployment rate in the OECD countries. Percent

3 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 which are important for equilibrium unemployment.

3.1 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, see the data appendix for details.

There have been a 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 the rise in unemployment in the early 1970s occurred together with an increase in the dispersion. The difference between the highest and the lowest unemployment rate in 1995 is larger than in 1960. From 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 a relatively low unemployment rate after the financial crisis and throughout the period compared to most other countries in the sample. Ireland and Spain, on the other hand, are examples of countries with high levels of unemployment post the financial crisis, but also in some years prior to the crisis. Germany had an upward sloping trend prior to the financial crisis, but the unemployment rates has subsequently come down. Non of the countries in the data set has seen a declining global trend in unemployment.

Note that even if one might question the accuracy of the data in the beginning of the period, showing very low unemployment rates in for instance New Zealand and Switzerland, are our main results 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.

3.2 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), benefit replacement ratio (BRR), benefit duration (BD), union density ($UDNET$), tax level (TW) and the degree of coordination of wage setting (CO). As noted, these indicators are assumed to capture gradual evolution in the wage mark-up coefficient m_w, and to a lesser degree also m_q and 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 contains a detailed description of all variables and their sources. The appendix also contains tables for the actual development for each variable for each country in the sample period. The last row of each table contains the unweighed average for each institutional variable.

The tax wedge 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 tax wedge in most OECD countries over the period 1960 to 2012. The largest increases are found in Sweden, Spain and Portugal, and the highest tax wedges, larger than a 50 percent, are found in the Nordic countries and in Austria, Belgium, France, Italy, Spain and Germany. In Belgium, Canada, Finland, New Zealand, Sweden and United States, there was a small decline in the tax wedges toward the end of the sample period.

The time series for employment protection measure the strictness of the employment protection for the employee. 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 and over all countries in the sample. 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 between countries. Countries with the highest levels of employment protection in the 1970s, like Belgium, France, Germany, Italy and Portugal, have experienced a decline in the measure towards the end of the sample period, while it has been increasing slightly from a very low level (0.57) in the period 1973 to 1979 in Australia and 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 the first period when 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, cf., appendix Table C4. There are large differences in unemployment benefits between the OECD countries, the lowest benefits are found in United Kingdom and Australia, with unemployment benefits in the range of 0.15 to 0.21 in the period 2009 to 2012. The highest benefits are provided in Netherlands and Switzerland with benefits above 70 percent in the first period. Some countries have reversed the benefits during the latter sample period, see for instance Austria, Canada, Denmark, New Zealand, Sweden and the United Kingdom. Italy has increased the benefits and are now on 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. Canada and Japan stop the payments after one year and the index is then equal to zero, cf., 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 increase 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 has increased over the sample period until 1995 and the level have been steady since then. There are however some variations between countries, i.e. Denmark, France, Ireland, Italy, Portugal, Spain and Switzerland has increased duration in the latter period, while Germany, Netherlands and Norway has decreased duration.

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

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, trade union membership and density in European Union countries, Norway and Switzerland were calculated using administrative data adjusted for non-active and self-employed members by Prof. Jelle Visser, University of Amsterdam. Appendix Table C3 shows that union density has declined since the beginning of the 1970s in most countries. Norway and Belgium are two exceptions where the measure has been stable over the whole period, and Denmark and Finland where the measure has increased.

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

4 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 which 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 (employment protection), BRR (benefit replacement ratio), BD (benefit duration), $UDNET$ (union density), TW (tax wedge), and CO (degree of coordination of wage setting). The distributed lag in Z_{it} is motivated by theory: An institutional reform in period t affects wage and price setting in $t+1$ and this leads to unemployment response in the final equation model in period $t+2$. We also include interaction terms just mentioned, which earlier studies have shown to be of importance Belot and van Ours (2001) and Nickell et al. (2005a), but we do introduce explicit notation for those interaction terms in (10).

In order to reduce collinearity, and to get a direct estimate on the 'level effect' of an institutional variable that affect u^* we estimate the model in terms of $\Delta Z_{i,t-1}$ and $Z_{i,t-2}$, which is a re-parametrization of (10) and do 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 loose a few annual observations. Typically, the longest time series is still 50, and the shortest is 46 observations.

We follow the study of OECD unemployment by Nickell et al. (2005b) and use the Within-Group estimator (WG hereafter), also called the least squares dummy variable (LSDV) method, as the reference estimator. This reflects that our main purpose is to estimate regression coefficients, and the equilibrium unemployment rate as a derived coefficient, that are free from unobserved heterogeneity bias. Since the time dimension is relatively large, the sample realizations of the individual effects $\beta_0 i$ 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

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

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 we generally would like assumption (11) to be accepted by the data (since it is important for 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 on may ask: does empirical acceptance of (11) contradict the relevance of the theory? However, a failure to reject the null hypothesis of no autocorrelation with the use of standard misspecification does of course not prove that the theoretical disturbances are without autocorrelation. As illustrated in the numerical example in the appendix, the miss-specification tests do not reveal the autocorrelation in that artificial data set. Equally important, the OLS estimation 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 allows for heteroscedasticity and autocorrelation, which also are implied by our theory. Second, we present results from difference GMM-based estimation. 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 shown to improve the size and the precision of the explanatory variables in time series, cf. Johansen and Nielsen (2009). The indicator saturation divides the sample in 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 over, and the number of indicators retained are often relatively few and interpretable, see Hendry (1999) and Castle et al. (2013) among others.

When the model is augmented by the selected individual specific location-shift indicators and estimated by OLS, we get the IIS estimators of the original regression coefficients. When the error distribution is symmetric, and under the null that there is no location shifts, the IIS estimator is centered around the same value as the OLS estimator, but there is an efficiency loss, since irrelevant explanatory variables are include in the model. Against that there is the potential of gains both in the centering and in the efficiency under the alternative of breaks and/or outliers which is empirically highly relevant for actual unemployment rates as we have seen.

Johansen and Nielsen (2009) have extended the IIS estimator to stationary autoregressive distributed lags, but the formalization of the panel data version of the IIS estimator is still not in the literature. We propose to use the method as a robustification of the WG estimator. Concretely, the WG-IIS estimator is obtained by first applying within-group transformation of the data set, second impulse saturating the transformed data set and then selecting location-shift by the computer algorithm *Autometrics*, 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. They are "fixed" in the sense of the computer program *Autometrics* that we use, see Doornik (2009), Doornik and Hendry (2009, Ch 14.8).

It might be noted that all our estimated models are augmented by time-dummies, in the same way as in Nickell et al. (2005b) 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 finding that no-such dummies are found to be significant and that the conventional times dummies are sufficient in terms of robustifying WG estimator of the regression coefficients.

5 Empirical results

We first give the results for WG estimation of equation (10). In section 5.2 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 ISS estimator.

5.1 Results for 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, column one and two. However, and for completeness, we present feasible GLS estimators for the case of residual heteroscedasticity and autocorrelation as indicated by the column headings in column three and four in the same table.

The WG estimator shows that the two first lags of the unemployment rate have theory consistent signs and are significant 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 column three and four. The GLS estimators show that the third lag of unemployment is positive and it is significant at the 10 percent level. These results are in line with the predictions from theory, in particular that the first autoregressive coefficient can be 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 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 percent significance level, both coordination (CO) and benefit replacement ratio (BRR) are now significant in addition to tax wedge (TW) and the interaction term between (BRR) and (BD). The signs of the estimated regression coefficients are reasonable: Increased tax levels and increased level and length of benefits increase equilibrium unemployment. The coefficient of coordination, 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

Table 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.

corrects for autocorrelation. It is interesting to see if this empirical determination of the sign is robust when we compare results with other estimation methods.

5.2 Difference GMM-based estimation

As noted in section 4, 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 effect. This leads to difference GMM based estimation, since in our case for example Δu_{it-1} becomes correlated with $\Delta\varepsilon_t$ and need to be instrumented, see Arellano (2003, Ch 4)

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. Also differences can 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, the number of instruments then becomes very large. This problem is known as instrument proliferation see for instance Bowsher (2002) and Roodman (2009). The problem affects both the one-step and the two-step estimators. Disregarding the loss of possible instruments due to multicollinearity problems, the instruments grows quadratic in T, and in our case, the number of available instruments is 47^2. The literature notes three problems with many available instruments, this is also 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 always reject while the Hansen test has implausible high p-values, see Roodman (2009) and Bowsher (2002). This is also our experience.

In order 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 experience from GMM estimation show a much too large difference between the WG estimator and the GMM estimator, i.e. they cannot be interpreted as only being corrections of the finite i bias in the fixed effect model[3]. For instance, the estimated coefficient of the lagged endogenous variable changed by a factor of three when the collapse option where used in the estimation. On the other hand, the estimated values of the institutional variables where close to the WG estimator. If the number of lags are reduced, the lagged endogenous variable is close to the WG estimator, while 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, cf., Table 2. The reduction in available instruments is therefore probably not sufficient to obtain good estimators in this complex model. Also in this case did the Hansen test show rejection with an implausible high p-value equal to one. Note also that the theory gives no clear guidelines for the choice of how to reduce the number of lags available as instruments before the "collapse" function is applied to

[3]All results from the Difference GMM method are available upon request

the model. We have tried with a different number of lags, but all choices reveal erratic behavior of the estimates and large variations (not reported for space consideration).

Table 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.

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 for current levels than past levels are of current changes, new instruments are more relevant.

In addition, several of the WG estimators of the coefficients in equation (10) are 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 estimator for a simplified model where only significant variables enter as explanatory variables not using 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 all variables have the same sign (compare the results under "WG", "GMM, onestep" 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.

Table 3: Years with location shifts. Results from Autometrics treating three lags of unemployment and the full set of institutional variables in Table 1 as fixed.

	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

5.3 Results for the WG-IIS estimator

We suggest to use an estimator with location shift dummies as a robustified WG estimator. The approach is analogue 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 in *Autometrics*, see Doornik (2009). The first procedure, is the impulse indicator saturation (IIS) method. This selection algorithm first adds dummies for each year to the model and then selects 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 in 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.

Autometrics is 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 is then applied using *Autometrics*, which is part of PcGive in OxMetrics, see Doornik and Hendry (2009). The set of institutional variables, and the three lags of the unemployment rate are not selected over. We used a significance level of 5 percent when the method of large outliers was used and 2.5 for impulse saturation, see Doornik (2009). As shown in Table 3, very few outliers were found 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 1980-83 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 USA.[4] The oil price shocks are represented by separate dummies in the results for several countries, the USA 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-crises 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 of unemployment, there are also examples of intermittent reductions. Some of these represent the effects of the well-know housing and credit market booms (for example the UK in 1988 and 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 USA in 2008.

The estimation results in the column labeled IIS in Table 4 is comparable to the WG in Table 1 above. The estimated residual standard error of IIS (0.3) is a good deal lower than the WG residuals (0.6), cf., 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. Lower standard residual errors makes more precise interval forecasts. For the institutional variables, one change is that the interaction terms between BRR and BD is no longer significant when WG-IIS is used, showing that the WG estimate for this interaction term is non-robust.

The second column of Table 4 uses robust estimation of 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 terms follows a autoregressive process, nor in the estimated coefficients for the explanatory variables, cf., column one and four in Table 4. There are however, some more significant explanatory variables when the error term is allowed to follow a autoregressive process.

The average fit of unemployment with the WG-IIS estimator of column one in Table 4 is illustrated in Figure 2 together with a simulation without the impulse indicator saturation, but with the time dummies. Since 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 1970s to the mid 1980s and has remained on around 6 percentage points 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 have lowered the level of several of the included institutional variables, cf. in particular the development benefit replacement ratio, coordination and tax wedges in appendix Table C4, C1 and C6. A simulation for each country in the sample is presented in appendix Figures B1 and B2.

6 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, and the two WG-IIS versions are presented in column two and three. One interesting finding is that the use of the robust estimator changes the results for the coordination index CO. The coefficient of CO is insignificant with WG estimation, but it is larger (in absolute value) and significant in the two WG-IIS estimations. There is also a notable sign-change

[4] The FED increased the interest rate to 20 percent at early in the 1980.

Table 4: Impulse indicator saturation WG estimates

	Dependent variable: Unemployment rate (u_{it}). Percent											
	IIS			IIS, robust[a]			IISstacked[b]			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.

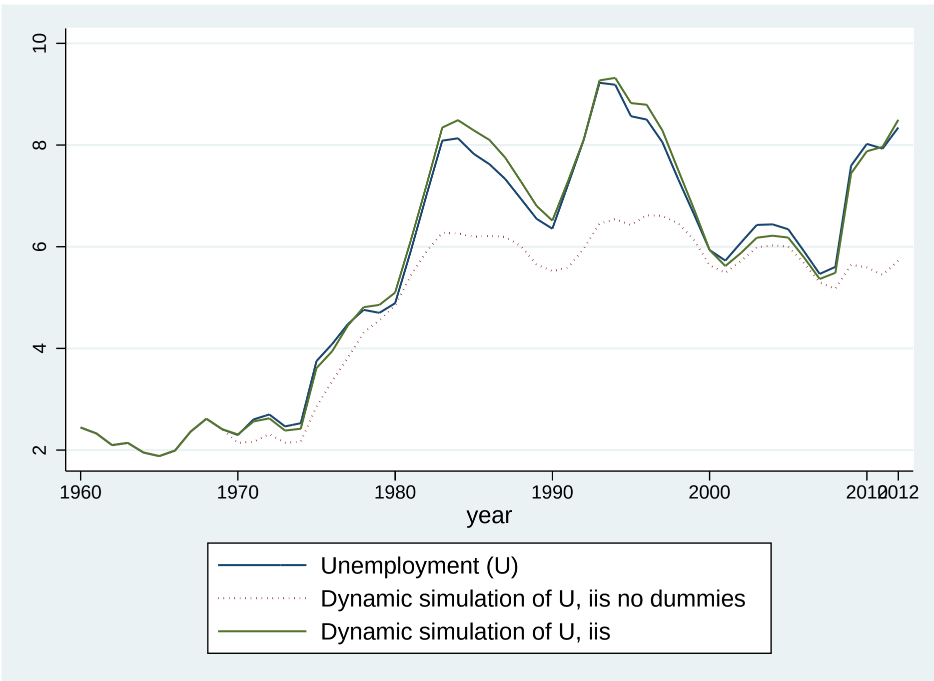

Figure 2: Simulated unemployment rate in the OECD countries. WG-IIS estimates are used for the coefficients. Percent

Table 5: Estimated long-run coefficients of the institutional variables.

	WG (Tab. 1)		WG-IIS (Tab. 2)		WG-IISs (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

in the results for interaction term between CO and TW, which switches (WG) from positive to negative (WG-IIS). Some other coefficient estimates also differ quite a lot although the signs are preserved. The estimated coefficients of the tax-level, TW, and the interaction term between benefits and duration (BRR and BD- interaction) are the main other examples of this. In sum, it appears that the main effect of changing from WG to WG-IIS is the centering of the estimates rather than in 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 by our results is that higher degree of coordination 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 line of reform, where the supportive evidence is more in terms of a tendency when we average across estimation methods, is that lower compensation levels (BRR) can reduce unemployment in the longer term, cf., both the coefficients of BRR and the interaction terms with benefit duration (BD). Third, the coefficient of the average tax-rate TW is consistently positive, although it is scaled down with WG-IIS estimation and with a t-value a little above or below one.

In 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 the "stacked" version of WG-IIS. 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 version of the WG-IIS estimate gives some more grounds for optimism about the level of the equilibrium rate conditioned by "no- change-in-institutions". As we might expect, the 2012 level of the unemployment rate is as much as 2 percentage points higher then the estimated equilibrium rate.

Figure 4 shows the simulations results for the individual countries in four panels. The categorization is mainly done with reference to institutional characteristics. In panel a), UK stands out as the only country with above average simulated equilibrium unemployment. USA, despite the rise 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 level above the average, while Norway and Sweden have lower estimated u^*s. Panel c shows the so called PIIG countries with high current and estimated long-run unemployment rates. It is interesting to not that although Spain performed worst at the end of the sample, Spain's equilibrium rate is about the same as Italy's and Portu-

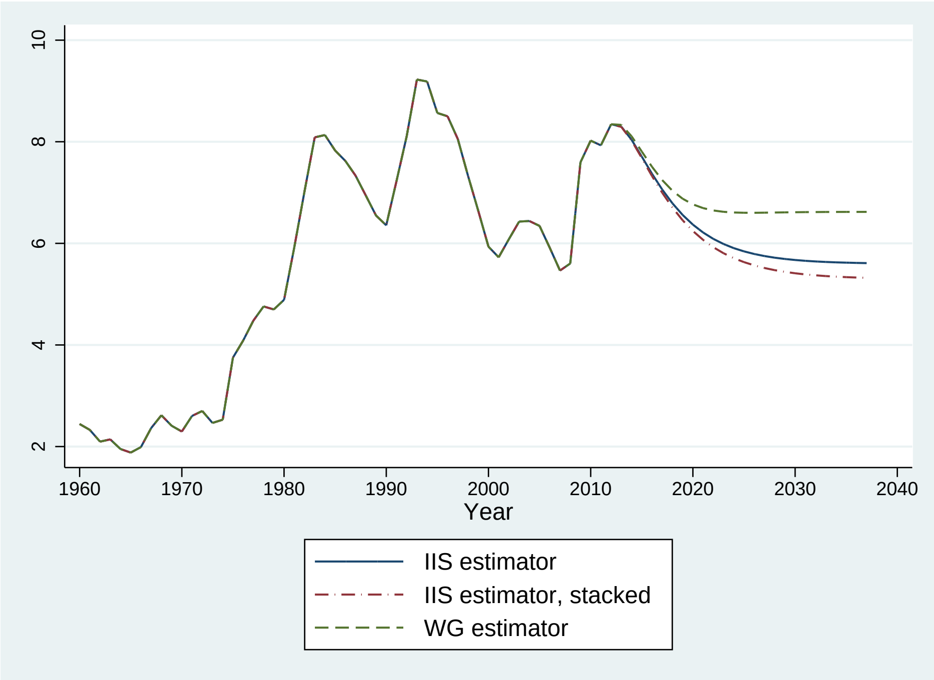

Figure 3: Simulated unemployment rate out of sample in the OECD countries. WG-IIS estimator. Percent

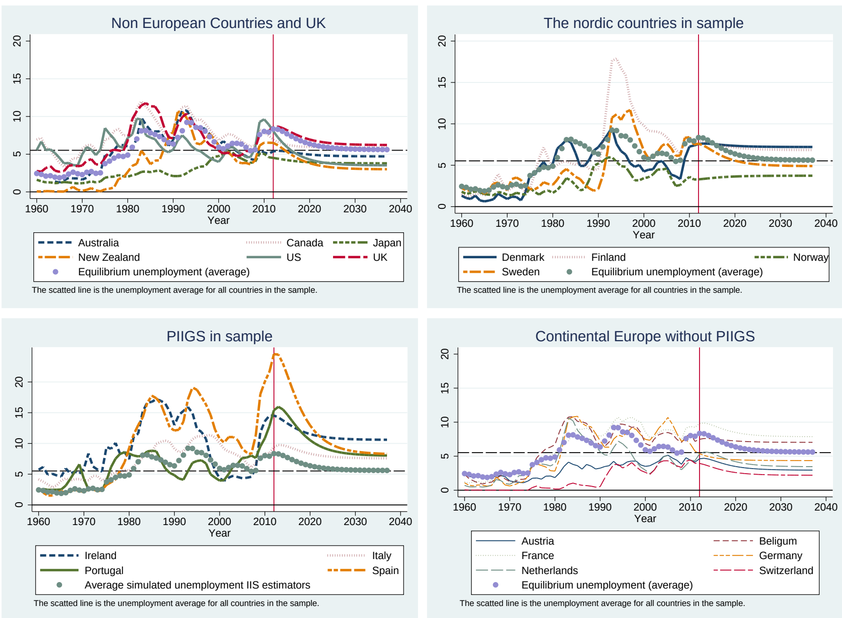

Figure 4: The actual unemployment rate, the WG-IIS model and the WG-IIS model with only country and time specific effects for each country in the panel

gal'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 contains graphs for each country.

7 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 decisions. Realistic estimates of equilibrium rates are sought after 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 good, since 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 dynamics models for the rate of unemployment. Compared to earlier studies that use the Layard-Nickel model as a reference, we have a more detailed theoretical determination of the equilibrium rate where the steady-state is the solution to the dynamic model which 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 offering we have chosen to follow the earlier panel data literature and estimate single equation models for the rate of unemployment. Unlike earlier studies, the dynamics and inclusion of the institutional variables in our final equation follows precisely from the theoretical VAR model.

In terms of estimation method, we propose that there is a role for the method of impulse indicator saturation, which has been developed in time series econometrics as a robust estimator, also in the estimation of panel dynamic data model with many time series observations. In our case the impulse indicator saturation estimator is a robust version of the standard WG (LSDC) estimator. It is clearly related to the WG estimator with time dummies, since WG-IIS includes significant time dummies for individual countries after selection by the computer program *Autometrics*. We propose that this gives a relevant method for modelling e.g., heterogenous 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 at the same time as 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. Changes in operational definitions may also have played a role. Our own empirical evidence support that, in particular, improved coordination in wage setting is a significant dimension of labour market reform that may over time reduce the level of unemployment. The tendency in the empirical evidence is also supporting 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, while there is little evidence in our results that employment protection has been an important determinant.

It is not surprising that our model shows that the unemployment in all countries (expect Norway) was above the estimated equilibrium rate, and the "explanation" is found in the impact on the real economy of the credit crisis. A simple prediction of or model is therefore 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 by 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 which 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.

Appendix A Additional notes and results for the theoretical framework of section 2

A.1 A dynamic model for wage-and-price setting

(1) is linked to wage and price setting, and the institutional changes in that part of the economy, through the real exchange rate, re_t. Since this variable is measured in logarithms we write it as $re_t = pi_t - q_t$ where pi_t denotes an index (in log scale) of import prices and q_t is an index of "home" producer prices.

q_t will be an endogenous variable in the domestic wage-and price, while pi_t is represented as random-walk with drift:

$$pi_t = g_{pi} + pi_{t-1} + \varepsilon_{pit} \tag{A1}$$

This equation represent a common nominal trend in our model. We also include a common real trend, for the log of average labour productivity a_t:

$$a_t = g_a + a_{t-1} + \varepsilon_{at} \tag{A2}$$

ε_{at}, and ε_{pit} are assumed to be innovations with zero expectations.

We now model wage and price setting as conditional on pi_t and a_t, equations (A1) and (A2) therefore imply that q_t, the (log of the) price level on domestic products, and w_t, the (log of) wage compensation per hour will be non-stationary, integrated of order one, $I(1)$, in a common notation.

Let p_t denote the logarithm of the consumer price index, and p_t is defined by:

$$p_t = \phi q_t + (1 - \phi) pi_t \tag{A3}$$

where the parameter ϕ measures the share of imports in total consumption.

We next define two theoretical (latent) real wage variables: The optimal real wages from the point of view of the firms, rw_t^f, and the bargained real wage, rw_t^b. They are given by the following two equations:

$$rw_t^f \equiv w_t^{ef} - q_t^f = -m_q \qquad\qquad + \; a_t + \vartheta\, u_t \tag{A4}$$

$$rw_t^b \equiv w_t^b - q_t^{eb} = \quad m_w + \omega \left(p_t^{eb} - q_t^{eb}\right) + \iota\, a_t - \varpi\, u_t. \tag{A5}$$

Equations (A4) and (A5) can seen as open-economy versions of the relationship for price- and wage-setting in Layard et al. (2005, p 13).[5]

In the price-setting equation (A4), q_t^f denotes the price level set by the firm on basis of expected nominal marginal labour costs $w_t^{ef} - a_t$. w_t^{ef} denotes the *expected* hourly wage cost. In the wage setting equation (A5), w_t^b denotes the nominal wage outcome and q_t^{eb} and p_t^{eb} are the price expectations that affect that bargaining outcome. A main implication of the bargaining model is that elasticity ι with respect to productivity is close to unity, as in Nymoen and Rødseth (2003). The standard assumption about the sign of the coefficient for unemployment ϖ is that it is non-negative, hence $-\varpi < 0$. The coefficient ω is called the wedge-coefficient since $\left(p_t^{eb} - q_t^{eb}\right)$ is the wedge between expected consumer and producer real wages (when we abstract from tax rates). The wedge coefficient is assumed to be non-negative, $\omega \geq 0$.

We proceed by making the assumption that rw_t^f and rw_t^b are *co-integrated* with the actual real wage. Similarly, on the price side, it is reasonable that equation (A4) captures the secular trend in the actual price q_t, but not the period-to-period changes in the price level.

[5]See also Sørensen and Whitta-Jacobsen (2010, Ch 12 and 17), Blanchard (2009, Ch 6).

We assume that the wage and price expectations errors $w_t^{ef} - w_t$, $q_t^{eb} - q_t$ and $p_t^{eb} - p_t$ are $I(0)$. The expectation variables in equations (A4) and (A5) is then replaced by w_t, q_t and p_t, without changing the order of integration. Cointegration therefore carry over to observable variables and we have the following the equilibrium correction model with reference to the Granger-Engle (1987) representation theorem:s

$$\Delta q_t = c_q + \psi_{qw}\,\Delta w_t + \psi_{qpi}\,\Delta pi_t - \varsigma\, u_{t-1} + \theta_q\, ecm_{t-1}^f + \varepsilon_{qt}, \tag{A6}$$

$$\Delta w_t = c_w + \psi_{wq}\,\Delta q_t + \psi_{wp}\,\Delta p_t - \varphi\, u_{t-1} - \theta_w\, ecm_{t-1}^b + \varepsilon_{wt}, \tag{A7}$$

where ε_{qt}, and ε_{wt} are innovations and all parameters are assumed to be non-negative. The error correction terms, ecm_{t-1}^f and ecm_{t-1}^b, are consistent with equations (A4) and (A5), with $w_t^{ef} = w_t$, $q_t^{eb} = q_t$ and $p_t^{eb} = p_t$ imposed. They are defined by

$$ecm_t^f = w_t - q_t - \ a_t \qquad\qquad - \vartheta\, u_t + m_q \tag{A8}$$

$$ecm_t^b = w_t - q_t - \iota\, a_t - \omega\,(p_t - q_t) + \varpi\, u_t - m_w, \tag{A9}$$

see Bårdsen et al. (2005) and Bårdsen and Nymoen (2009a) for similar derivations.

The dynamic model of the wage-price spiral is:

$$\begin{aligned}\Delta q_t &= (c_q + \theta_q\, m_q) + \psi_{qw}\,\Delta w_t + \psi_{qpi}\,\Delta pi_t - \mu_q\, u_{t-1}\\ &\quad + \theta_q\,(w_{t-1} - q_{t-1} - a_{t-1}) + \varepsilon_{q,t},\end{aligned} \tag{A10}$$

$$\begin{aligned}\Delta w_t &= (c_w + \theta_w\, m_w) + \psi_{wq}\,\Delta q_t + \psi_{wp}\,\Delta p_t - \mu_w\, u_{t-1}\\ &\quad - \theta_w\,(w_{t-1} - q_{t-1} - \iota\, a_{t-1}) + \theta_w\,\omega\,(p_{t-1} - q_{t-1}) + \varepsilon_{w,t},\end{aligned} \tag{A11}$$

$$\Delta p_t = \phi\,\Delta q_t + (1-\phi)\,\Delta pi_t, \tag{A12}$$

where equation (A12) is the result of taking the difference on both sides of the definition in equation (A3).[6] Note that in equations (A10) and (A11), notations $\mu_q = \theta_q\,\vartheta + \varsigma$ and $\mu_w = \theta_w\,\varpi + \varphi$ are used for the coefficients on u_{t-1}.

(A10)-((A12) can be re-formulated as a (open) VAR model for the real exchange rate re_t and the log of the wage share $ws_t = w_t - q_t - a_t$. This conditional VAR is found in the two two first rows of (2) in the main text. For re_t the coefficients are:

$$\begin{aligned}
l &= 1 - \theta_w\,\omega\,\psi_{qw}\,(1-\phi)/\chi,\\
k &= (\theta_q - \theta_w\psi_{qw})/\chi,\\
e &= 1 - (\psi_{qpi} + \psi_{qw}\,\psi_{wp}\,(1-\phi))/\chi, \quad = 0 \text{ if dynamic homogeneity}\\
n &= (\mu_q + \mu_w\,\psi_{qw})/\chi,\\
d &= (m_q\,\theta_q + c_q + (m_w\,\theta_w + c_w)\,\psi_{qw})/\chi,
\end{aligned}$$

where the denominator is: $\chi = 1 - \psi_{qw}(\phi\psi_{wp} + \psi_{wq}) > 0$. For ws_t the coefficients are:

$$\begin{aligned}
\lambda &= \theta_w\,\omega\,(1-\psi_{qw})(1-\phi)/\chi,\\
\kappa &= 1 - (\theta_w\,(1-\psi_{qw}) + \theta_q\,(1-\psi_{wq} - \phi\,\psi_{wp}))/\chi,\\
\xi &= (\psi_{wp}\,(1-\psi_{qw})(1-\phi) - \psi_{qpi}\,(1-\psi_{wq} - \phi\,\psi_{wp}))/\chi, \qquad = 0 \text{ if dynamic homogeneity}\\
\eta &= (\mu_w\,(1-\psi_{qw}) - \mu_q\,(1-\psi_{wq} - \phi\,\psi_{wp}))/\chi,\\
\delta &= ((m_w\,\theta_w + c_w)(1-\psi_{qw}) - (m_q\,\theta_q + c_q)(1-\psi_{wq} - \phi\,\psi_{wp}))/\chi.
\end{aligned}$$

[6] For coefficients ψ_{wq}, ψ_{qw} and ψ_{wp}, ψ_{qpi}, the non-negative signs are standard in economic models. Negative values of θ_w and θ_q imply an explosive evolution in wages and prices (hyperinflation), which is different from the low to moderately high inflation scenario that we have in mind for this paper.

By inspection, it is clear that all coefficient are non-negative for reasonable values of the structural coefficients. The exception is δ which can be both positive and negative. The first two VAR disturbances are

$$\epsilon_{re,t} = (\varepsilon_{q,t} + \psi_{qw}\,\varepsilon_{w,t})/\chi \quad \text{and} \quad \epsilon_{ws,t} = (\varepsilon_{q,t}\,(1-\psi_{wq}-\phi\,\psi_{wp}) - \varepsilon_{w,t}\,(1-\psi_{qw}))/\chi,$$

while the third is identical to ε_{ut} in the unemployment equation.

The steady-state solution for u_t, the equilibrium rate of unemployment is given by (3) with coefficients

$$\mathfrak{e}_{ss} = \rho\,(\theta_q(1-\psi_{wq}-\psi_{wp}) + \theta_w(1-\psi_{qw}-\psi_{qpi}))/(\theta_q\,\theta_w\,\Omega), \tag{A13}$$
$$\mathfrak{b}_{ss} = \rho\,(\theta_q - \theta_w\,\psi_{qw})/(\theta_q\,\theta_w\,\Omega), \tag{A14}$$
$$\mathfrak{d}_{ss} = [\rho\,(m_w + m_q + c_w/\theta_w + c_q/\theta_q) + c_u\,\omega\,(1-\phi)]/\Omega. \tag{A15}$$

with $\Omega = \omega\,(1-\phi)\,(1-\alpha) + \rho\,(\varpi + \vartheta)$.

For simplicity, the expression (4) for $\mathfrak{d}_{ss}$ in the main text is for the case where the drift terms c_w and c_q have been set to zero.

A.2 A numerical example

To illustrate the properties of the structural model, and of a simple one-off estimation of the equilibrium rate, we have used the VAR representation (2) in the main text, together with (A1)-(A3) to generate a data set (T=200) for re_t, ws_t, u_t, pi_t, a_t and p_t using parameter values that give stationarity, and with a single location shift in period 150. The structural disturbances are Gaussian and independent. We then estimate the structural form (not the VAR) on a data set that ends in period 160, and simulate the estimated structural form dynamically over a period that starts in period 160 and ends in period 200. The dynamic simulation is stochastic (1000 replications). The average of the solution paths represent the estimated expectations of the endogenous variables. Since we have estimated the true model, the solution converges to the imputed steady-state values of the endogenous variables.

The figure contains four panels with blue graphs of the actuals (i.e., the computer generated data) for re_t, ws_t, Δp_t (i.e., inflation) and u_t. The dashed green line is the model shows the average of the simulated solutions. The stable equilibrium nature of the solutions are evident. The red dotted lines are upper and lower 95 % prediction intervals around the solution. There are no structural breaks after period 150, so when two actuals for inflation are significantly outside the prediction interval, they are the result of tail observations ("black swans"), and are not the result of location shifts.

The line representing the solution for u_t is stable at 1.28 % showing that this the true u^* for this structure. If we use the computer generated data for u_t and estimate a third order autoregressive model over the sample period 1-160 by OLS, we obtain:

$$\begin{aligned} u_t \;=\; & \underset{(0.08)}{0.7728}\; u_{t-1} - \underset{(0.0994)}{0.2247}\; u_{t-2} + \underset{(0.0797)}{0.1379}\; u_{t-3} \\ & + \underset{(0.0922)}{0.4094} \end{aligned} \tag{A16}$$

with standard errors below the estimates. The estimated equilibrium rate from this simple final equation model is:

$$\hat{u}^* = \underset{(0.36)}{1.304} \tag{A17}$$

which we see is a good estimate of the true equilibrium rate. Nevertheless, the final equation diagnostics shows that there are significant first order ARCH effects, and departures

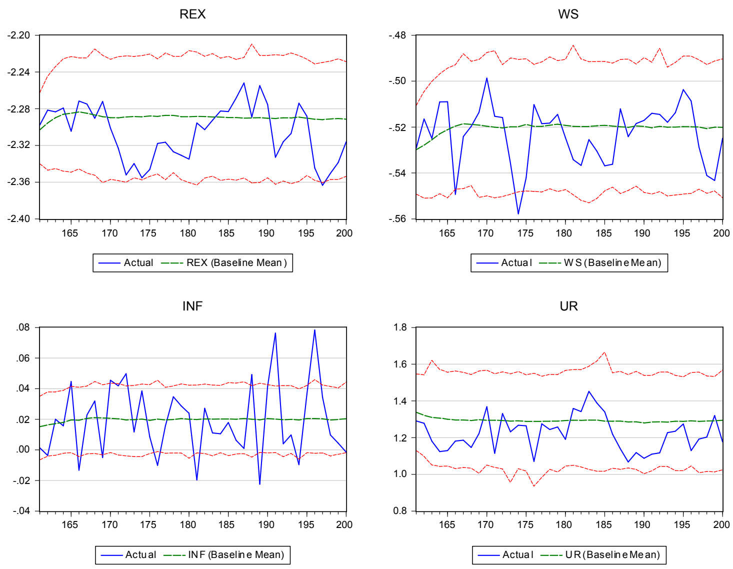

Figure A1: Data from calibrated version of theoretical model, and stochastic simulation based on estimated true structural form. Means and $\pm$ 2 standard errors shown for four of the model's endogenous variables: Panel a) The real exchange rate. Panel b) The wage-share c) Panel c) Inflation. Panel d) The unemployment rate.

form normality (Jarque-Bera test). Since we know that there is a location shift in the data these significant misspecification test are not surprising.

When we use Autometrics to select location break indictor variables, retaining the intercept and the three lags, and setting the 'target size' to 0.001 Autometrics includes two dummies for period 151 and period 152. This finding is correct since the effect of the structural impulse in period 150 is carried over to the following period(s) by the dynamics. The result is

$$\begin{aligned} u_t \quad = \quad & \underset{(0.094)}{0.7195}\ d151_t + \underset{(0.106)}{0.5123}\ d152_t + \underset{(0.0709)}{0.4946} \\ & + \underset{(0.0686)}{0.5338}\ u_{t-1} + \underset{(0.0768)}{0.01005}\ u_{t-2} + \underset{(0.0607)}{0.06647}\ u_{t-3} \end{aligned}$$

For this model the are no significant misspecification tests, even though we know that there must be a moving-average in the true disturbance of the final equation model, i.e., like (9) since it is the calibrated theory model that has generated the data.

The ISS estimated equilibrium rate is

$$\hat{u}^*_{IIS} = \underset{(0.30)}{1.269} \tag{A18}$$

showing that in this case, the main effect of the robustification is to lower the standard error of the estimated equilibrium rate.

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

This appendix contains additional information to section 5 and 6. Figures B1 and B2 shows the fit using WG-IIS estimators derived in section 5. Figures B3 and B4 contain the residuals from a simulation using WG and WG-IIS estimators. Finally, the equilibrium unemployment rates, given the level of institutions in 2012 and the WG-IIS estimators of Table 4, are presented in figures B5 and B6.

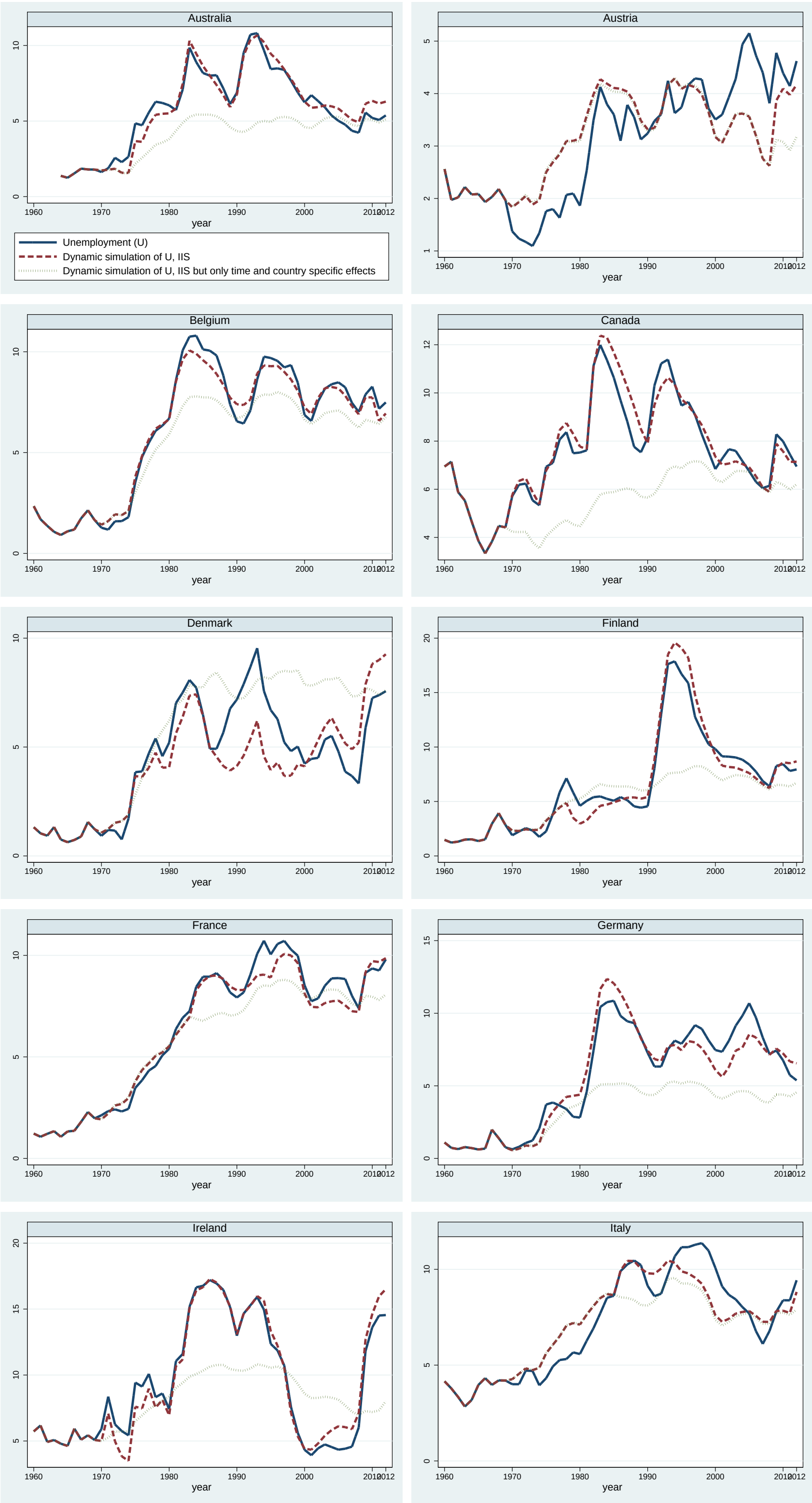

Figure B1: The actual unemployment rate, the WG-IIS estimated model and the WG-IIS estimated model with only country and time specific effects for each country in the panel

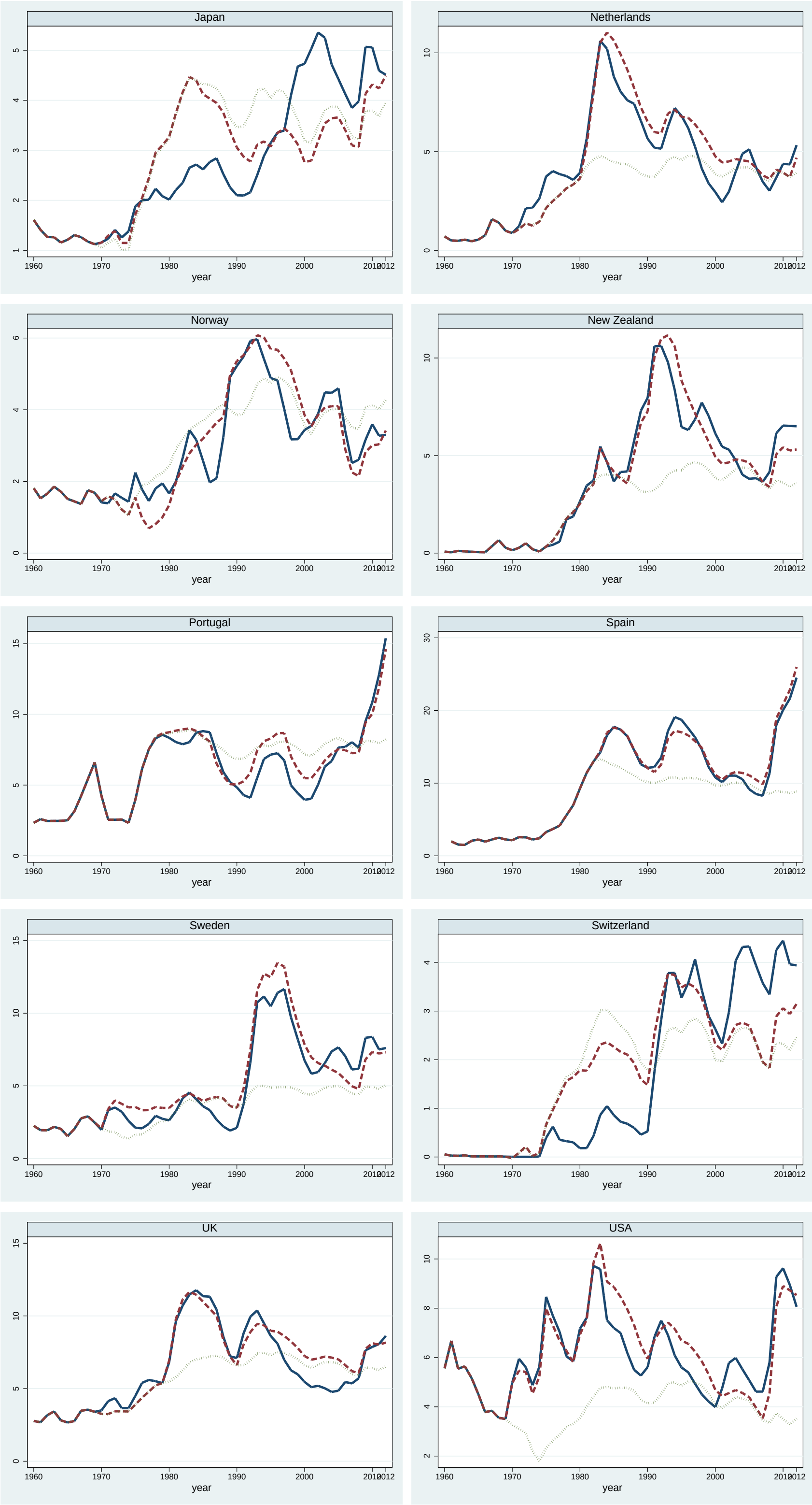

Figure B2: The actual unemployment rate, the WG-IIS estimated model and the WG-IIS estimated model with only country and time specific effects for each country in the panel, cont.

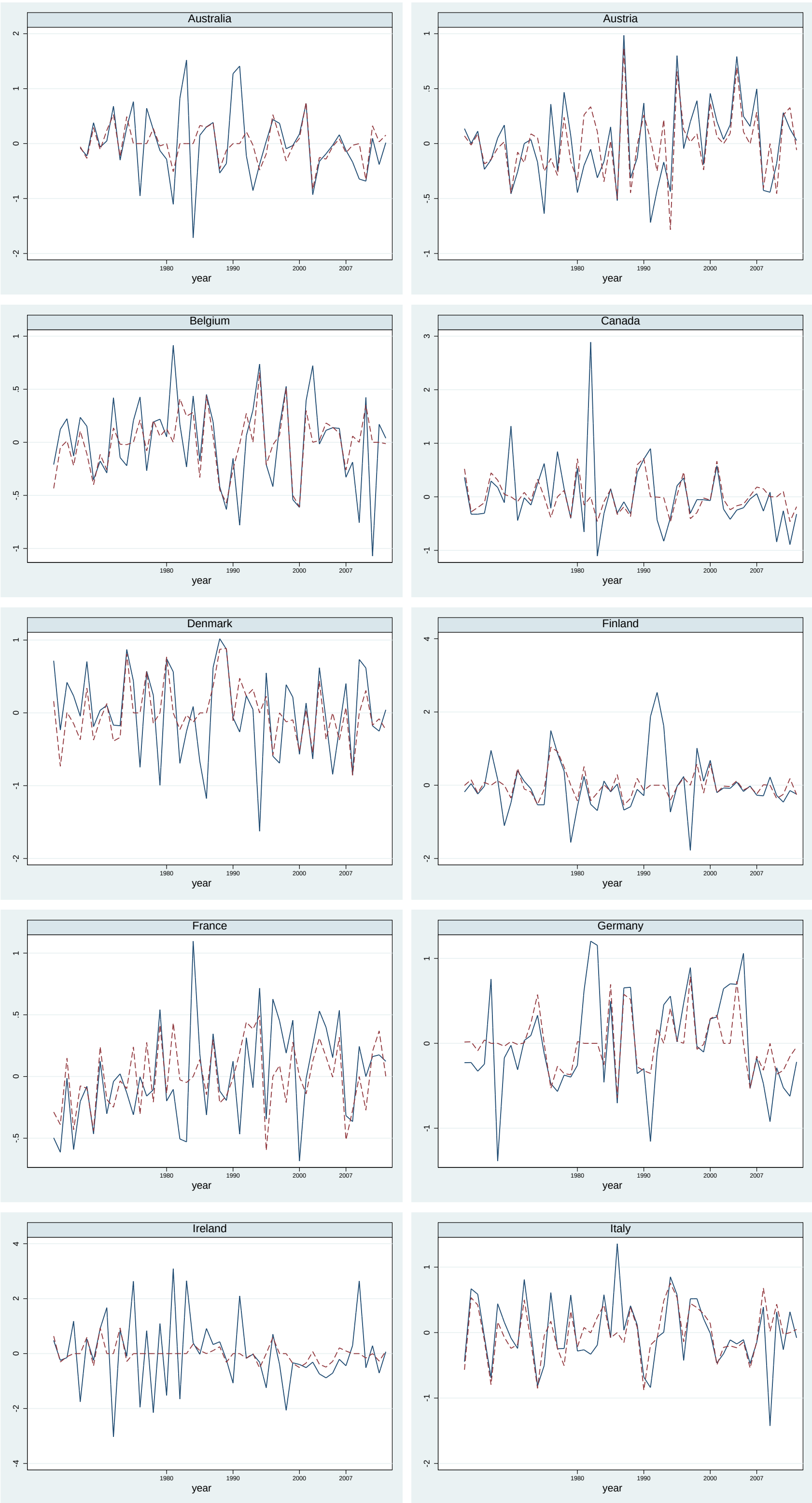

Figure B3: Residuals from the WG-IIS estimated model and the WG estimated model

Figure B4: Residuals from the WG-IIS estimated model and the WG estimated model, cont.

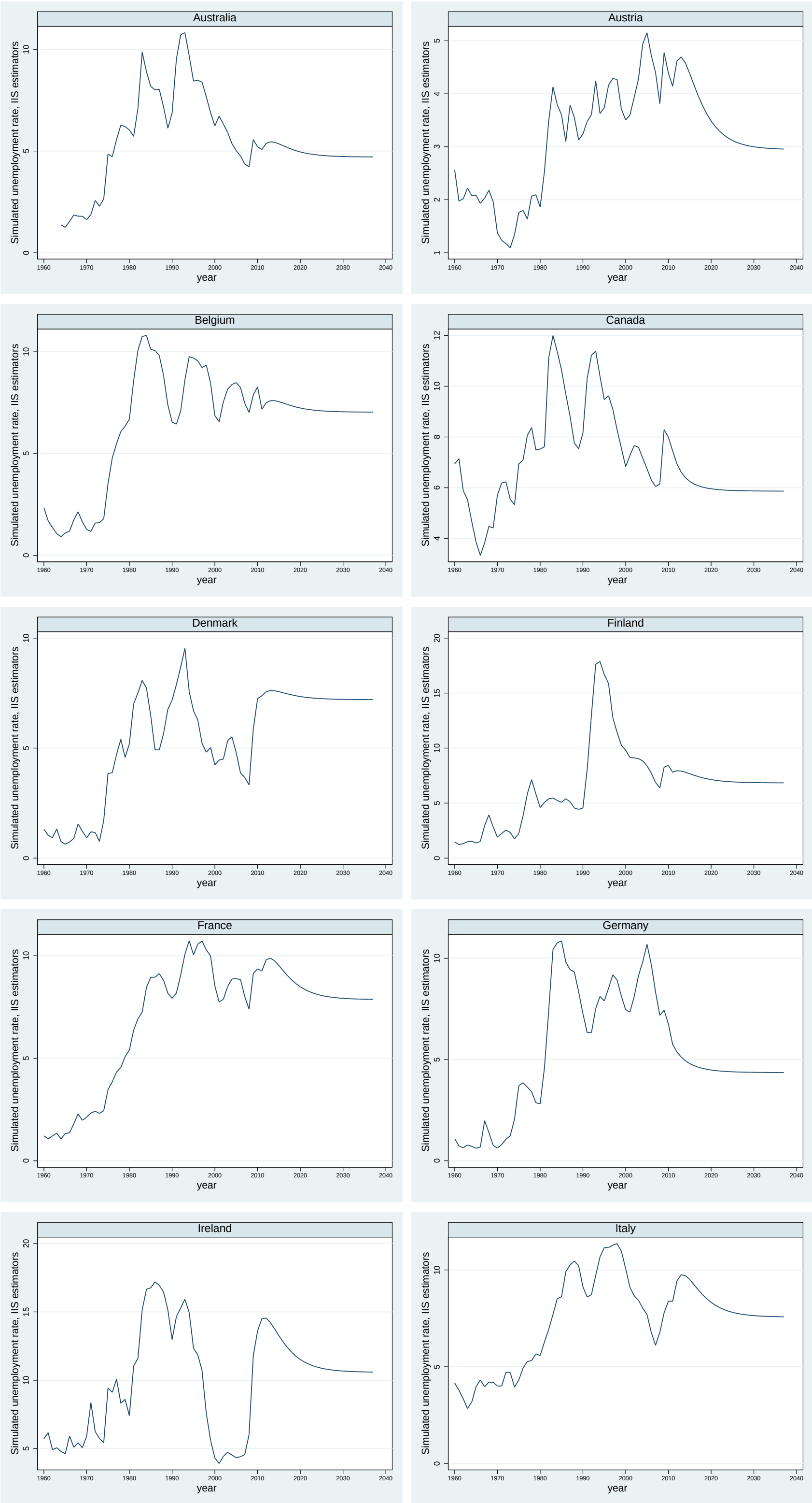

Figure B5: Simulations of the WG-IIS estimated model

Figure B6: Simulations of the WG-IIS estimated model, cont.

Appendix C Data Appendix

The data appendix describes where the data used in this thesis is found and how the variables are build up.

C.1 Unemployment Rate (UNR):

The standardized unemployment rate (UNR) from OECD (2013b). The data are prolonged backward for some countries, using the growth rate of numbers found in older versions of OECDs Economic Outlook: Germany before 1992 using numbers from OECD (2011) for Former Federal Republic of Germany, Ireland before 1990 using OECD (2012b) and Spain before 1967 using OECD (2005).

C.2 Coordination of Wage bargaining:

The coordination of wage bargaining is based on Kenworthy's 5-point classification of wage-setting coordination scores. The index ranges on a scale from 1 to 5. It was found in: ICTWSS Data on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts in 34 Counties Between 1960 and 2007, Version 3 Visser (2011b).

The actual development in coordination is found in Table C1.

C.3 Employment Protection Legislation Indicator (EPL):

The time series for employment protection measures the strictness of the employment protection for the employer. The OECD indicators of EPL are synthetic indicators of the strictness of regulation on dismissals of individuals or groups, and the procedures involved in hiring workers on fixed-term or temporary work agency contracts. They are compiled of 21 items, and divided into three areas: Individual dismissal of workers with regular contracts, additional cost for collective dismissals, regulations of temporary contracts, and an overall measure of EPL which is an average of the indicators for regulations for temporary and regular contracts. The overall measure for employment protection is measured on a scale from 0 to 5. Strictness is increasing in scale.

The data is from OECD (2010). The time series for employment protection is prolonged backwards for all the countries in the sample before 1985 except for New Zealand which was prolonged before 1990, by the growth rate of the measure of employment protection, *ep*, in the Nickell (2006) database.

The actual development in employment protection is found in Table C2.

C.4 Trade Union Density:

Trade union density corresponds to the ratio of wage and salary earners that are trade union members, divided by the total number of wage and salary earners (OECD Labor Force Statistics). Density is calculated using survey data, wherever possible, and administrative data adjusted for non-active and self-employed members otherwise.

Union density rates are constructed using the number of union memberships divided by the number of employed. The main data source is OECD (2012c), where they have mainly calculated the trade union density index based on surveys. When data were unavailable, they have used administrative data adjusted for non-active and self-employed members. The union density is prolonged by last known observation for New Zealand, Portugal and Spain before respectively 1970, 1977 and 1980. The interaction terms between union density and coordination are prolonged by the last known observation for these countries.

The actual development in union density is found in Table C3

Table C1: Average coordination in the OECD countries

Country	**Yr6064**	**Yr6572**	**Yr7379**	**Yr8087**	**Yr8895**	**Yr9601**	**Yr0208**	**Yr0912**
Australia	3.00	3.00	3.00	3.38	3.00	2.00	1.71	2.00
Austria	5.00	5.00	5.00	4.38	4.00	4.00	4.00	4.00
Belgium	4.00	4.00	3.71	4.63	4.25	4.50	4.29	4.00
Canada	1.00	1.00	2.71	1.00	1.00	1.00	1.00	1.00
Denmark	5.00	5.00	5.00	4.00	3.00	3.50	3.29	3.00
Finland	4.60	4.75	4.14	3.75	3.38	3.83	3.71	3.00
France	2.40	2.00	2.43	2.38	2.00	2.00	2.00	2.00
Germany	4.00	4.00	4.00	4.00	4.00	4.00	4.00	4.00
Ireland	1.00	1.75	3.14	1.75	4.25	5.00	5.00	2.00
Italy	2.00	2.00	2.57	2.75	3.00	4.00	4.00	4.00
Japan	4.00	4.00	4.00	4.00	4.00	3.50	3.00	3.00
Netherlands	4.60	3.88	3.86	4.38	4.00	4.00	4.00	4.00
New Zealand	4.00	4.25	5.00	4.38	1.75	1.00	1.57	2.00
Norway	5.00	4.50	4.43	3.50	4.50	4.33	4.00	4.00
Portugal	5.00	5.00	5.00	2.38	3.25	3.17	2.57	3.00
Spain	5.00	5.00	5.00	3.75	3.00	3.00	4.00	3.75
Sweden	5.00	5.00	5.00	4.25	3.63	3.00	3.00	3.00
Switzerland	4.00	4.00	4.00	4.00	4.00	3.00	3.00	3.00
UK	1.00	3.50	4.00	1.00	1.00	1.00	1.00	1.00
USA	1.00	1.50	1.57	1.00	1.00	1.00	1.00	1.00
Total	3.53	3.66	3.88	3.23	3.10	3.04	3.01	2.84

Table C2: Average employment protection in the OECD countries

Country	**Yr6064**	**Yr6572**	**Yr7379**	**Yr8087**	**Yr8895**	**Yr9601**	**Yr0208**	**Yr0912**
Australia	0.56	0.56	0.57	0.80	0.94	1.19	1.18	1.15
Austria	2.21	2.21	2.21	2.21	2.21	2.21	1.97	1.93
Belgium	3.27	3.27	3.38	3.25	3.15	2.33	2.18	2.18
Canada	0.75	0.75	0.75	0.75	0.75	0.75	0.75	0.75
Denmark	2.10	2.10	2.10	2.25	2.29	1.50	1.50	1.50
Finland	1.48	1.80	2.33	2.33	2.22	2.08	2.01	1.96
France	3.89	4.11	4.11	3.35	2.91	2.99	3.05	3.04
Germany	3.49	3.55	3.62	3.39	3.16	2.46	2.11	2.12
Ireland	0.65	0.65	0.78	0.92	0.93	0.93	1.08	1.11
Italy	3.81	3.88	3.91	3.74	3.57	2.79	1.86	1.89
Japan	1.71	1.71	1.80	1.88	1.84	1.52	1.43	1.43
Netherlands	2.37	2.37	2.70	2.78	2.73	2.42	2.08	1.95
New Zealand	0.86	0.86	0.86	0.86	0.86	1.06	1.46	1.40
Norway	3.95	3.95	3.33	2.92	2.87	2.65	2.62	2.69
Portugal	4.19	4.19	4.19	4.19	3.95	3.67	3.48	2.88
Spain	3.82	3.82	3.82	3.82	3.62	2.96	2.99	2.98
Sweden	3.87	3.81	3.54	3.48	3.11	2.28	2.19	1.87
Switzerland	0.76	0.79	1.05	1.14	1.14	1.14	1.14	1.14
UK	0.51	0.52	0.57	0.60	0.60	0.63	0.75	0.75
USA	0.00	0.00	0.00	0.10	0.21	0.21	0.21	0.21
Total	2.26	2.25	2.28	2.24	2.15	1.89	1.80	1.75

Table C3: Average union density in the OECD countries

country	Yr6064	Yr6572	Yr7379	Yr8087	Yr8895	Yr9601	Yr0208	Yr0912
Australia	0.49	0.45	0.49	0.46	0.38	0.27	0.21	0.18
Austria	0.67	0.63	0.59	0.53	0.45	0.38	0.32	0.28
Belgium	0.40	0.42	0.52	0.52	0.54	0.53	0.53	0.52
Canada	0.28	0.30	0.35	0.37	0.36	0.31	0.30	0.29
Denmark	0.58	0.59	0.71	0.79	0.76	0.76	0.71	0.69
Finland	0.35	0.47	0.66	0.69	0.77	0.77	0.72	0.70
France	0.20	0.21	0.21	0.15	0.10	0.08	0.08	0.08
Germany	0.34	0.32	0.35	0.35	0.32	0.26	0.21	0.19
Ireland	0.45	0.48	0.53	0.52	0.48	0.40	0.34	0.35
Italy	0.25	0.32	0.48	0.45	0.39	0.36	0.34	0.35
Japan	0.34	0.35	0.33	0.30	0.25	0.22	0.19	0.19
Netherlands	0.39	0.37	0.37	0.30	0.25	0.24	0.20	0.19
New Zealand	0.57	0.56	0.63	0.60	0.40	0.23	0.21	0.21
Norway	0.60	0.57	0.54	0.58	0.58	0.55	0.54	0.55
Portugal	0.61	0.61	0.61	0.47	0.28	0.23	0.21	0.20
Spain	0.08	0.08	0.08	0.10	0.15	0.16	0.15	0.16
Sweden	0.70	0.68	0.75	0.80	0.82	0.81	0.74	0.68
Switzerland	0.34	0.31	0.32	0.26	0.23	0.21	0.19	0.18
UK	0.39	0.41	0.46	0.47	0.38	0.31	0.29	0.26
USA	0.29	0.27	0.24	0.19	0.15	0.13	0.12	0.11
Total	0.42	0.42	0.46	0.44	0.40	0.36	0.33	0.32

C.5 Benefit Replacement Rates:

The benefit replacement rate is a measure of how much each unemployed worker receives in benefit from the government. The OECD gives information about the unemployment benefits for year 1, the average of year two and three, and the average of year four and five for unemployed person in different family situations and with different initial income level. The three different family types are: Single, with a dependent spouse and with a working spouse.

The different income levels are: 67 percent and 100 percent of average earnings. The measures used are:
Brr67a1: First year benefit replacement rate for workers with, 67 percent of average earnings and the average over family types.
Brr67a2: Benefit replacement for the second and third year, with 67 percent of average earnings and the average over family types.
Brr67a4: Benefit replacement for the fourth and fifth year, with 67 percent of average earnings and the average over family types.
Brr100a1: First year benefit replacement rate for workers with, 100 percent of average earnings and the average over family types.
Brr100a2: Benefit replacement for the second and third year, with 100 percent of average earnings and the average over family types.
Brr100a4: Benefit replacement for the fourth and fifth year, with 100 percent of average earnings and the average over family types.

The benefit replacement rate is calculated by taking the average of brr67a1 and brr100a1:

$BRR = \frac{brr61a1+brr100a1}{2}$

The actual development in the sample period is presented in Table C4

The data source is provided from OECD by e-mail, OECD (2012a).

Table C4: Average benefit replacement ratio in the OECD countries

Country	Yr6064	Yr6572	Yr7379	Yr8087	Yr8895	Yr9601	Yr0208	Yr0912
Australia	0.18	0.15	0.22	0.23	0.26	0.25	0.22	0.21
Austria	0.16	0.16	0.28	0.34	0.36	0.40	0.39	0.40
Belgium	0.39	0.37	0.55	0.51	0.48	0.46	0.46	0.51
Canada	0.40	0.40	0.60	0.57	0.58	0.47	0.39	0.44
Denmark	0.36	0.53	0.79	0.78	0.74	0.67	0.66	0.61
Finland	0.19	0.25	0.37	0.43	0.58	0.53	0.53	0.52
France	0.47	0.51	0.46	0.59	0.58	0.59	0.61	0.60
Germany	0.43	0.42	0.39	0.39	0.38	0.37	0.40	0.40
Ireland	0.21	0.24	0.39	0.51	0.41	0.35	0.42	0.54
Italy	0.10	0.07	0.04	0.02	0.09	0.43	0.56	0.52
Japan	0.36	0.38	0.33	0.28	0.30	0.35	0.40	0.40
Netherlands	0.42	0.65	0.65	0.68	0.70	0.70	0.71	0.71
New Zealand	0.39	0.30	0.27	0.30	0.29	0.28	0.26	0.25
Norway	0.12	0.12	0.24	0.53	0.62	0.63	0.64	0.64
Portugal	0.00	0.00	0.14	0.39	0.65	0.66	0.67	0.65
Spain	0.17	0.50	0.58	0.74	0.69	0.65	0.63	0.62
Sweden	0.24	0.31	0.70	0.85	0.87	0.77	0.74	0.68
Switzerland	0.17	0.12	0.27	0.50	0.67	0.72	0.75	0.75
UK	0.27	0.35	0.34	0.27	0.22	0.20	0.17	0.15
USA	0.22	0.23	0.28	0.31	0.25	0.29	0.31	0.51
Total	0.26	0.30	0.40	0.46	0.49	0.49	0.50	0.51

C.6 Benefit Duration:

The benefit duration is a measure of how long benefits last when you are unemployed and how the amount given changes over the duration. Benefit duration is given by the equation:

$$BDj_{it} = \alpha \frac{brrja2_{it}}{brrja1_{it}} + (1-\alpha)\frac{brrja4_{it}}{brrja1_{it}} \tag{C1}$$

where $\alpha = 0.6$, $j = \{67, 100\}$, $i = 1, 2...20$ and $t = 1960, 1961...2007$. $brrja1_{it}$ is the benefit replacement rate in year 1, $brrja2_{it}$ is the benefit replacement rate in year 2 and 3, and finally, $brrja4_{it}$ is the benefit replacement rate in years 4 and 5. $\alpha = 0.6$ gives more weight to the second and third year as compared to the fourth and fifth year. The index is calculated for both employment situations, i.e. 67 percent and 100 percent of average earnings. The average of $bd67_{it}$ and $bd100_{it}$ is used as an indicator of benefit duration, i.e. BD_{it}. If benefit duration stops after one year, then brr67a2 = brr67a4 = 0, and BD67 = 0. If benefit provision is constant over the years, then brr67a1 = brr67a2 = brr67a4, and BD67 = 1. However, some countries increase payments over time and the value of benefit duration is above one.

The actual development in benefit duration is found in Table C5.

The data source is provided from OECD by e-mail, OECD (2012a).

C.7 Tax Wedge:

Tax wedge is equal to the sum of the employment tax rate, the direct tax rate and the indirect tax rate. The rates described here 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), as given in Equation (C.7).

Table C5: Average benefit duration in the OECD countries

Country	Yr6064	Yr6572	Yr7379	Yr8087	Yr8895	Yr9601	Yr0208	Yr0912
Australia	1.01	1.01	1.03	1.02	1.03	1.01	1.00	1.00
Austria	0.00	0.00	0.58	0.75	0.74	0.70	0.71	0.71
Belgium	1.00	0.99	0.78	0.79	0.78	0.79	0.85	0.78
Canada	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Denmark	0.46	0.50	0.60	0.62	0.69	0.95	0.78	0.80
Finland	0.00	0.04	0.65	0.60	0.52	0.55	0.58	0.56
France	0.30	0.24	0.20	0.35	0.49	0.50	0.51	0.53
Germany	0.57	0.57	0.61	0.62	0.61	0.60	0.48	0.41
Ireland	0.66	0.77	0.46	0.38	0.55	0.76	0.74	0.74
Italy	0.00	0.00	0.00	0.00	0.00	0.26	0.45	0.45
Japan	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Netherlands	0.03	0.35	0.49	0.65	0.70	0.66	0.41	0.26
New Zealand	1.02	1.01	1.02	1.03	1.04	1.01	1.00	1.00
Norway	0.00	0.02	0.43	0.49	0.50	0.63	0.83	0.35
Portugal	0.00	0.00	0.00	0.07	0.35	0.46	0.49	0.58
Spain	0.00	0.00	0.00	0.19	0.28	0.37	0.38	0.38
Sweden	0.00	0.00	0.04	0.05	0.05	0.14	0.40	0.53
Switzerland	0.00	0.00	0.00	0.00	0.07	0.29	0.20	0.18
UK	0.89	0.63	0.54	0.69	0.71	0.80	0.83	0.80
USA	0.08	0.16	0.19	0.16	0.19	0.21	0.21	0.28
Total	0.30	0.31	0.38	0.42	0.46	0.53	0.54	0.52

$TW = t1 + t2 + t3$

t1: is equal to employers actual wage cost calculated by the sum of wages received by employees and taxes paid by the employer to the government.

$t1 = \frac{EC}{IE - EC}$

EC-Employers Social Security contributions
Social contributions are the actual or imputed payments to social insurance schemes to make provision for social insurance benefits. They may be made by employers on behalf of their employees or by employees, self-employed or non-employed persons on their own behalf. The contributions may be compulsory or voluntary and the schemes may be funded or unfunded.

IE- Compensation of employees
Compensation of employees is made up of two components: Wages and salaries payable in cash or in kind: These include the values of any social contributions, income taxes, etc., payable by the employee even if they are actually withheld by the employer and paid on behalf of the employee.

t2 Direct Tax Rate

$t2 = \frac{IT + WC}{HRC}$

WC- Employees social security contributions
Social contributions are the actual or imputed payments to social insurance schemes to make provision for social insurance benefits. They may be made by employers on behalf of their employees or by employees, self-employed or non-employed persons on their own behalf. The contributions may be compulsory or voluntary and the schemes may be funded

or unfunded.

IT- Income tax
Current taxes on income, wealth, etc.
HCR- Current receipts of households
Current receipts of households consist of all income to a household, whether monetary or in kind received by the household or by individual members of the household. It includes income from employment, investments, current transfers, etc.

t3 Indirect Tax Rate

$t3 = \frac{TX-SB}{CC}$

TX-Indirect taxes
Taxes on consumption goods.

SB- The value of subsidies
Value of subsidies paid by government.
CC- Final consumtion
Final consumption expenditure for entire economy.

All variables were found in National Accounts, OECD (2013a). EC (NFD12R), IE (NFD1R), WC (NFD61P-NFD12R), IT (NFD5P), HCR (NFB5GR) and SB (NFD3P) were found in Table 14.A (Non-Financial accounts by sector) in household sector for all except SB which was found in general government sector. TX(D2) was found in Table 10 , general government sector. CC(P3) was found in Table 1.

This series is extended backwards with the growth rate of the series for tax wedge used in Nymoen and Sparrman (2012) before 1995 for: Austria, Belgium, Denmark, France, Germany, Netherlands, Norway, Portugal and Sweden. It is extended backwards before the 1990 for UK and Italy, before 1975 for Finland, before 2002 for Ireland, Before 2000 for Spain and before 1998 for US. Australia, Canada, New Zealand and Switzerland are replaced for the entire time series.

The development in the actual tax rates are found in Table C6.

Table C6: Average tax rate in the OECD countries

Country	**Yr6064**	**Yr6572**	**Yr7379**	**Yr8087**	**Yr8895**	**Yr9601**	**Yr0208**	**Yr0912**
Australia	0.23	0.26	0.31	0.35	0.34	0.36	0.37	0.36
Austria	0.43	0.48	0.50	0.52	0.52	0.54	0.54	0.53
Belgium	0.46	0.49	0.50	0.52	0.57	0.59	0.58	0.55
Canada	0.30	0.37	0.38	0.40	0.46	0.47	0.44	0.43
Denmark	0.42	0.52	0.52	0.57	0.58	0.61	0.62	0.58
Finland	0.41	0.45	0.50	0.52	0.63	0.66	0.63	0.59
France	0.67	0.60	0.56	0.60	0.63	0.65	0.64	0.63
Germany	0.43	0.44	0.46	0.45	0.46	0.48	0.49	0.50
Ireland	0.33	0.38	0.40	0.43	0.45	0.48	0.52	0.48
Italy	0.39	0.37	0.36	0.43	0.53	0.64	0.63	0.63
Japan	0.23	0.24	0.26	0.31	0.33	0.33	0.35	0.35
Netherlands	0.44	0.48	0.54	0.56	0.52	0.52	0.54	0.55
New Zealand	0.32	0.32	0.29	0.32	0.40	0.37	0.38	0.38
Norway	0.50	0.54	0.56	0.57	0.55	0.58	0.56	0.56
Portugal	0.17	0.21	0.22	0.28	0.34	0.38	0.40	0.39
Spain	0.21	0.26	0.32	0.41	0.46	0.47	0.51	0.47
Sweden	0.36	0.48	0.57	0.65	0.70	0.72	0.71	0.68
Switzerland	0.16	0.17	0.21	0.21	0.21	0.24	0.24	0.25
UK	0.39	0.44	0.42	0.46	0.43	0.43	0.44	0.44
USA	0.22	0.23	0.24	0.24	0.24	0.25	0.22	0.20
Total	0.35	0.39	0.41	0.44	0.47	0.49	0.49	0.48

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