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The Unemployment Benefit System and Wage Flexibility in EMU: Time-varying Evidence in five Countries

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Abstract

A large number of European countries still cope with historically high unemployment rates. One line of research that has been followed to explain European unemployment and the differences among European countries is the impact labour market institutions. One important channel through which labour market institutions may affect unemployment is the responsiveness of wages to unemployment, commonly referred to as (real) wage flexibility. It has been shown that cross-country differences in labour market institutions can account for differences in wage flexibility, but there is not any consistent econometric work that explores the relationship between *changes* in labour market institutions and wage flexibility over time *within* countries. This is the issue addressed in this paper. Wage flexibility is defined as the coefficient on unemployment in a 'bargaining-augmented' wage equation, explaining (real) wage growth. We investigate the role of unemployment benefits in determining the degree of (real) wage flexibility. To this end we estimated a wage equation in a time-varying parameter framework for five core EMU countries. In Italy the unemployment benefit system is very limited. For the four other countries, the results show that, except for Belgium, wage flexibility is not related in a significant way to the generosity of the unemployment benefit system. This insight runs counter to the conclusions offered by cross-section studies. We therefore tentatively conclude that we should not be too optimistic about the effect of reform of the unemployment benefit system on wage flexibility and that such reform – in order to be really effective– should be radical.

Introduction

A large number of European countries still cope with historically high unemployment rates. Several different lines of research have been followed to explain European unemployment and the differences among European countries. For example, OECD (1994) focuses on labour market institutions. Because labour market institutions generally favour insiders, the interests of outsiders will hardly be taken into consideration at the bargaining table and cyclical unemployment tends to become structural. This argument is in line with the observation of rising equilibrium unemployment accompanying the rise in unemployment in Europe. On the other hand, Blanchard (1999) and Blanchard & Wolfers (2000) stress the interaction between shocks and institutions. Shocks explain the rise in unemployment, while institutions account for the differences among European countries. Blanchard also acknowledges the possibility that changes in labour market institutions over time have influenced unemployment. However, in a 'traditional' framework, such as Layard *et al.* (1991), with unemployment as the dependent variable regressed upon shocks and time-varying institutions, little evidence is found for the effect of institutions on unemployment.

One important channel through which labour market institutions may affect unemployment is the responsiveness of wages to unemployment, commonly referred to as (real) wage flexibility. Layard *et al.* (1991), Scarpetta (1996), Nickell (1997) and, more recently, Pentecost and Sessions (2002) have shown that cross-country differences in labour market institutions can account for differences in wage flexibility, but there is not any consistent econometric work that explores the relationship between *changes* in labour market institutions and wage flexibility over time *within* countries. This is the issue addressed in this paper. Wage flexibility is defined as the coefficient on unemployment in a 'bargaining-augmented' wage equation, explaining (real) wage growth. A time-varying parameter framework is then introduced to allow this coefficient to vary with various labour market institutions.

Since the establishment of EMU member countries have been deprived of an independent monetary and exchange rate policy, so wage flexibility in European labour markets has even become more important as the need for alternative adjustment mechanisms to deal with (asymmetric) shocks has increased. In this framework, an appropriate redesigning of labour market institutions to enhance wage flexibility, would not only reduce European unemployment, but also contribute to a smooth functioning of EMU. Our results, based on the role of changes in the generosity of unemployment benefits, show that reform of the unemployment benefit schemes should not be expected to modify the degree of wage flexibility in a significant way. This conclusion stands in contrast to the insights about the role of unemployment benefit schemes in the malfunctioning of the European labour markets offered by cross-section studies. However, these studies generate a flexibility measure based on estimation over a sample period under the assumption that flexibility is constant during the period. In a separate second stage, cross-country differences in flexibility are related to labour market institutions. Policy implications of the second stage then advise countries to adjust labour market institutions in order to increase flexibility. Clearly this is somewhat at odds whith the (implicit) first stage assumption of constancy, unless labour market institutions did not change during the sample period.

The remainder of this paper is organised as follows. In Section 2 we estimate a basic wage equation. In Section 3 we re-estimate this equation in a time-varying parameter framework, allowing the generosity of the unemployment benefits to influence the parameter of the unemployment rate in the wage equation. Section 4 concludes with our main findings.

The wage equation

Theoretical models (such as bargaining models) suggest a negative relationship between the level of the wages and the unemployment rate, given the reservation wage and the level of productivity. This is the so-called *wage curve* relationship. Empirical findings, however, suggest a *Phillips curve* relationship between wages and unemployment, i.e. a negative relationship between the rate of change of wages and the unemployment rate. Blanchard and Katz (1999) reconcile these theoretical and empirical specifications of the wages-unemployment relationship by interpreting the reservation wage as depending on productivity and lagged wages. This results in the following specification:

$$\Delta w_t = c_w + \Delta pc_t^e + \varphi(w_{t-1} - pc_{t-1} - z_{t-1}) + \beta u_t + \delta \Delta z_t + \varepsilon_t \qquad (1)$$

where w and pc are the logarithms of the wage and (consumption) price level, ^{e} denotes expectations, u is the unemployment rate and z is the logarithmic labour productivity. Δx stands for a growth rate. Wage growth is determined by inflation expectations, the level of the unemployment rate (the Phillips curve effect), the change in productivity and an ‘error correction’ term, $(w_{t-1} - pc_{t-1} - z_{t-1})$, implying an adjustment of real wages to (trend) labour productivity over time. In fact real wages adjust to marginal productivity, but assuming a Cobb-Douglas production function, marginal productivity ($\partial Y_t / \partial L_t$) equals average productivity ($Y_t / L_t = z_t$). Inflation expectations are assumed to be a convex combination of current and lagged inflation (adaptive expectations):

$$\Delta pc_t^e = \alpha \Delta pc_t + (1-\alpha)\Delta pc_{t-1} = \alpha \Delta\Delta pc_t + \Delta pc_{t-1} \qquad (2)$$

The closer α to one, the larger the influence of current inflation or institutionalized indexation (α=1 is (contemporaneous) full indexation) and consequently a small effect of lagged inflation. Substituting (2) in (1) and adding the difference between consumer and output price inflation (to test for a terms of trade effect) and changes in the unemployment rate (to test for possible hysteresis effects), we have the following estimable specification:

$$\begin{aligned}\Delta(w_t - pc_{t-1}) = c_w + \alpha\,\Delta\Delta pc_t + \beta\,u_t + \gamma\,\Delta u_t + \varphi(w_{t-1} - pc_{t-1} - z_{t-1}) + \delta\,\Delta z_t \\ + \theta(\Delta p_t - \Delta pc_t) + \varepsilon_t\end{aligned} \tag{3}$$

Note that in this specific setting with adaptive expectations, the impact of unemployment on nominal (Δw_t) and real ($\Delta(w_t - pc_{t-1})$) wage growth is interchangeably, as can be seen from (3) and (1). A specification along these lines is also estimated in OECD (1997) and Lauer (1999). A theoretical justification for this “bargaining augmented Phillips curve” can already be found in Knoester and Van der Windt (1987). Wage growth in the private sector (Δw) is shown to be the outcome of negotiations between unions and employers, more specifically a weighted average of wage growth claims of unions and wage growth offers of employers. Unions’ claims are assumed to reflect compensation for changes in consumer prices (Δpc), labour productivity growth in the private sector (Δz). Employers’ offers are derived from marginal productivity conditions for profit maximising firms. The wage offers are shown to include compensation for changes in *output-prices* (Δp) and changes in labour productivity (Δz). Finally the Phillips curve effect is introduced by the assumption that the respective bargaining power of unions and employers depend on the labour market situation, reflected by the unemployment rate (u). Additionally, the burden of direct taxes and social security contributions for employees and employers’ social security contributions will also be taken into account at the bargaining table.[1]

[1] We econometrically tested for the effect of these tax variables, but they proved to be insignificant, perhaps because the variables are not disaggregated enough or because of a too limited sample size (only starting in 1970) (see Plasmans *et al.* (1999)) We opted not to go into further detailed tax measures because –given the limited sample size- we prefer a parsimonious specification.

Equation (1) is estimated by OLS.[2] A "general-to-specific" approach[3] is used to identify the variables significantly influencing wage growth (for later use in a time-varying framework). The sample consists of annual observations for the period 1960-1995 and the variables are taken from the OECD Statistical Compendium, Economic Outlook (1998). *w, pc, p* and *z* are expressed in logarithms (prices and wages are denoted in home currency), the unemployment rate is expressed in levels (as a decimal). Wages and productivity refer to the private sector. For the construction of the error correction term trend labour productivity (based on a Hodrick-Prescott filter) is used rather than actually measured productivity.

Table 1: Wage equations - dependent variable: $\Delta(w_t - pc_{t-1})$ (standard errors between brackets)[4]

	Belgium[a]	**France**	**Germany**[a]	**Italy**	**the Netherlands**
Sample	1962-1995	1963-1995	1962-1995	1962-1995	1962-1995
Const.	0.2387	0.3354	0.3573	3.2791	0.2205
	(0.0741)	(0.0851)	(0.1228)	(0.5801)	(0.0472)
Dummy	-0.0315	-	-0.0336	-	-
	(0.0084)		(0.0097)		
$\Delta\Delta pc_t$	0.5752	0.8966	1.2092	1.0508	0.9157
	(0.1017)	(0.1176)	(0.1647)	(0.1067)	(0.1021)
u_t	-0.6382	-0.7110	-0.5117	-1.3655	-0.6483
	(0.0762)	(0.0954)	(0.1144)	(0.1554)	(0.1000)
Δp_t - Δpc_t	-	0.3921	0.9741	1.2589	-
		(0.2251)	(0.1537)	(0.3461)	
Δz_t	0.2257	-	0.7324	-	0.5272
	(0.1420)		(0.0607)		(0.1474)
w_{t-1}-pc_{t-1}-z_{t-1}	-0.0837	-0.1364	-0.1710	-0.3614	-0.0893
	(0.0375)	(0.0409)	(0.0612)	(0.0654)	(0.0247)
Adj. R^2	0.88	0.92	0.94	0.89	0.88
DW	1.85	1.75	1.78	2.16	1.76

[a] dummy: 1962-1969=1

[2] Using instruments to account for possible simultaneity bias and inconsistency does not alter the results significantly, which is also confirmed by OECD (1997).
[3] See Hendry (1995); the technique used also accounts for possible multicollinearity among the explanatory variables.
[4] Estimating a Seemingly Unrelated Regression model yields similar results. Pooling the data adds little to our analysis, Wald tests reject the hypothesis of equal coefficients across countries. (see Plasmans *et al.* (1999)).

The results are shown in table 1. The adjusted R^2 is high for all wage growth equations and the Durbin-Watson statistic is above the inconclusive region for all equations, except for Germany, where it is close to the upperbound (1.08-1.89). For Belgium and Germany a dummy variable was included in the regression for the 1960-69 period. The negative sign of the coefficient on this variable is consistent with the Newell-Symons "wage explosion" dummy for the seventies (see Layard *et al.* (1991)).

Changes in the unemployment rate turned out to be insignificant (γ=0), pointing to the absence of hysteresis effects. This is in line with Lauer (1999) and OECD (1997), who using different sample periods, also find no or only marginal evidence for hysteresis. Hence, $|\beta|$ can be interpreted as wage flexibility (nominal and real wage flexibility on the β-parameter can be used interchangeably). The unemployment rate enters significantly in all five wage equations. In Italy the point estimate of the $|\beta|$-coefficient (1.37) is significantly higher than in the other countries. Since the Italian unemployment benefit system is extremely strict (see OECD (1994)) this might be a crude indication that generosity matters for the level of real wage flexibility. The estimated measure of flexibility in the other countries is about 0.6, which implies that an increase of one percentage point in unemployment increases the downward pressure on wages exerted by unemployment of 0.6%. The error correction term always enters significantly with the expected negative sign. If real wages were higher than indicated by productivity, this has a mitigating effect on wage growth in the next period.

Except for Belgium, the coefficient on inflation growth is close to one. This implies that inflation is entirely passed on into wage growth in the same period (year). In Belgium lagged inflation still has an impact on wages. Changes in productivity have a significant positive influence on wages in Belgium, Germany and The Netherlands. Further, the terms of trade variable, $(\Delta p_t - \Delta pc_t)$, turns out to be significant for wage formation in France, Germany and Italy, but not for Belgium and The Netherlands. The degree to which firms can influence the prices of their output is probably smaller in small open economies, hence the insignificance of the difference between output-prices and consumer prices.

The wage equation in a time-varying parameter framework

It is now standard to link wage flexibility to labour market institutions. In the theoretical and empirical literature much attention has been devoted to the role of unemployment benefits, active labour market policies, level of wage bargaining, employment protection legislation and types of wage contracts (Scarpetta (1996), Vinals and Jimeno (1996), Nickell (1997)). Given our definition of (real) wage flexibility as the responsiveness of (real) wages to unemployment, the (absolute value of the) parameter β ($\gamma=0$ for all countries) in equation (1) can be interpreted as a measure of (real) wage flexibility. Equation (2) reflects the consensus view about the relationship between (real) wage flexibility and labour market institutions:

$$|\beta| = f(\overset{-}{GUS}, \overset{+}{ALMP}, \overset{\pm}{CWB}, \overset{-}{EPL}, \overset{-}{ADWC}, \overset{+}{SWC}) \qquad (4)$$

with: GUS: generosity of the unemployment benefit system
ALMP: active labour market policies (expenditure as a percentage of GDP)
CWB: degree of centralisation of wage bargaining
EPL: degree of employment protection legislation
ADWC: average duration of wage contracts
SWC: synchronisation of wage contracts

The signs in equation (4) reflect the expected theoretical effect of the various labour market institutions on real wage flexibility: negative for the generosity of unemployment benefits, employment protection and the average duration of wage contracts; positive for active labour market policies and the degree of synchronisation of wage contracts. The degree of centralisation of wage bargaining is assumed to have an ambiguous effect on wage bargaining due to the well-known hump-shaped relationship. A thorough discussion of the role of these labour market institutions is beyond the scope of this paper and can be found in Calmfors and Driffil (1988), Heylen (1993), Johanssen (1999), Nickell (1997), Scarpetta (1996) and Vinals and Jimeno (1996).

Since wage flexibility is influenced by labour market institutions we cannot expect this parameter to be constant over time if the underlying institutions change over time. This hypothesis can be modelled in the following manner[5]:

$$\Delta(w_t - pc_{t-1}) = c_w + \beta_t u_t + C_X X_t + \varepsilon_t \qquad (5)$$

$$\text{with} \quad \beta_t = c_\beta + C_V V_t + \upsilon_t$$

$$\text{and} \quad X_t := [\Delta\Delta pc_t, (w_{t-1} - pc_{t-1} - z_{t-1}), \Delta z_t, (\Delta p_t - \Delta pc_t)]$$

$$V_t := [\mathrm{GUS}_t, \mathrm{ALMP}_t, \mathrm{CWB}_t, \mathrm{EPL}_t, \mathrm{ADWC}_t, \mathrm{SWC}_t]$$

The above model can be characterised as a systematically varying-parameter framework. Time series for most of the variables in the *V*-vector not being available, we had to restrict ourselves to one institutional variable, viz. the generosity of the unemployment benefit system. The restriction to one institutional variable may limit the scope of the analysis, since labour market outcomes are often related to a complex combination of labour market institutions. The reform of one institution or characteristic can result in an off-setting reaction by wage setters and policy makers with respect to the other institutions. If, for example, unemployment benefits are made less generous, wage setters may bargain for longer notice periods and higher severance payments or conclude longer-term employment contracts.

With this caveat in mind the model to be estimated then becomes:

$$\Delta(w_t - pc_{t-1}) = c_w + \beta_t u_t + C_X X_t + \varepsilon_t \qquad \text{with } \varepsilon_t \overset{i.i.d.}{\sim} (0,\sigma_\varepsilon^2) \tag{6a}$$

where β_t is stochastic and is assumed to vary according to:

$$\beta_t = c_\beta + c_{GUS}\, GUS_t + \upsilon_t \qquad \text{with } \upsilon_t \overset{i.i.d.}{\sim} (0,\sigma_\upsilon^2) \text{ and } \upsilon_t \text{ independent of } \varepsilon_t\ \forall t \tag{6b}$$

This model analyses the effect of (changes in) the generosity of the unemployment benefit system on real wage flexibility. Since β_t is expected to be negative and because a more generous system is assumed to decrease wage flexibility, we expect c_{GUS} to be positive and $c_\beta<0$ with $|c_\beta|>c_{GUS}$.

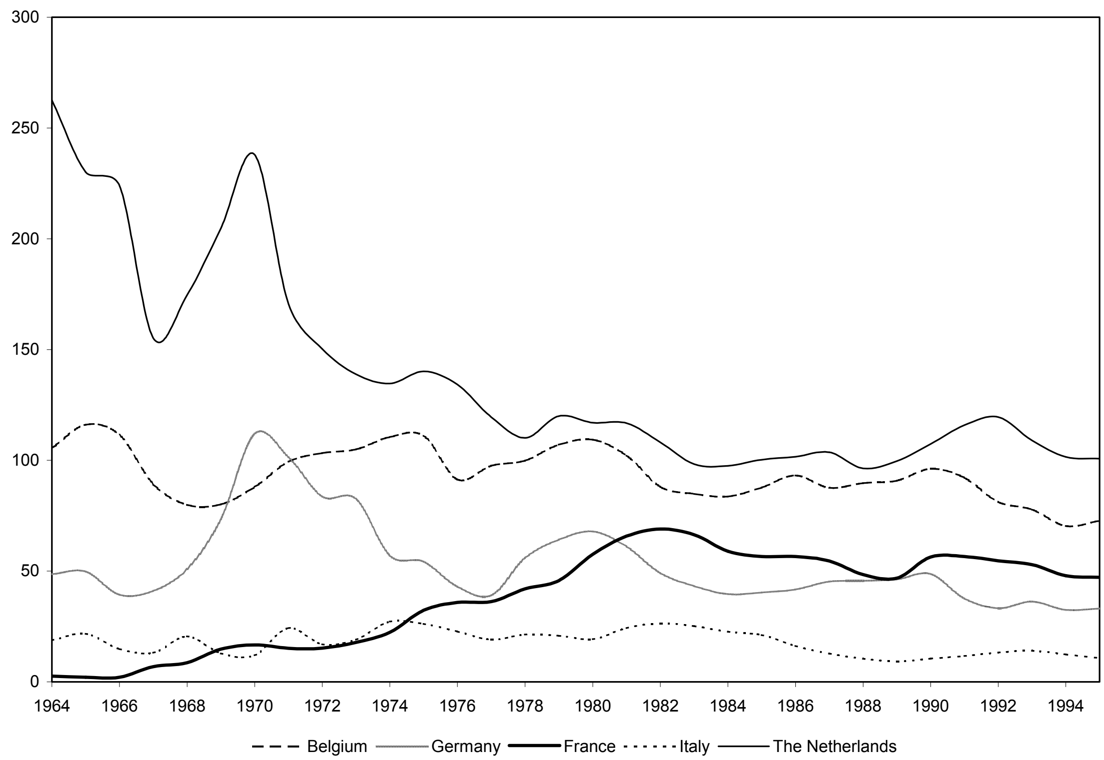
As a measure of the generosity of the unemployment benefit system we use expenditures on unemployment benefits as a percentage of GDP, divided by the number of unemployed expressed as a percentage of total population (where GDP and unemployment benefits are measured both in value or both in volume). This is equivalent to expenditures per unemployed as a percentage of per capita GDP. This implies that when unemployment benefits are not adjusted to an increasing welfare level, measured by per capita GDP, generosity declines. By using per capita GDP, a potential bias towards small countries is ruled out. Data for total expenditures on unemployment benefits were taken from Eurostat and calculated back with growth rates of comparable data obtained from the International Labour Organization (ILO). Additional data for the transformations were taken from OECD (1998).

From figure 1 it can be inferred that the Netherlands have the most generous system of the five countries studied. The Belgian system is also fairly generous and since the mid seventies almost

[5] Since the tax variables and changes in unemployment proved to be insignificant for all countries we do not include them in X_t.

as generous as the Dutch. At the opposite side we observe the almost non-existent Italian benefit system. Since benefits in Italy are so small and since there are no large changes in the system, it does not make sense to perform a time-varying estimation for Italy.[6] The benefit systems of France and Germany are somewhere in between these two extremes.

Figure 1: Expenditures on unemployment benefits per unemployed, as a percentage of GDP per capita



Source: own calculations with data from OECD (1998) and Eurostat, "Social protection expenditures and receipts" (calculated back with growth rates of comparable data obtained from ILO, "The cost of social security")

The variables of the OLS regressions in table 1 are used for the maximum likelihood estimation of model (6a-b) in a time varying parameter setting. In the estimation we use (lagged) logarithms and (lagged) changes in logarithms as transformations of the generosity measure, GUS in (4b). In the appendix the joint likelihood function for a time-varying model with one explanatory variable

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[6] We did test for the presence of a time-varying relationship between the coefficient of unemployment and the generosity of unemployment benefits, but the results were as expected, i.e. there is no relationship.

in parameter equation (*cf.* (4b)) is derived, where error terms are assumed to be normally and independently distributed. It is shown that the maximisation of the likelihood function boils down to the maximisation of a function F with respect to the variance-ratio, i.e. the ratio of the variance of the error term of the wage flexibility equation and the variance of the error term of the wage equation. This maximisation can be solved by a simple random search procedure. A grid search procedure is used to find the variances-ratio that (indirectly) maximises the likelihood function. This enables us to find the most accurate estimates. To facilitate the comparison with the OLS estimations we will also report an average value of β_t (i.e. β_{av}).

In addition we test four hypotheses, viz.

$H_0(1)$: $\lambda = 0$ (where λ is the ratio of the variances of υ and ε, cf. appendix);

$H_0(2)$: $\lambda = 0$, $c_{GUS} = 0$;

$H_0(3)$: $\lambda = 0$, $c_{GUS} = 0$, $c_\beta = 0$, and

$H_0(4)$: $\lambda = 0$, $c_{GUS} = 0$, $c_\beta = 0$, $C_X = 0$.

If all four hypotheses are rejected we have a significant time varying parameter model. Rejection of (null) hypothesis 3 implies that the unemployment rate is a significant determinant of wage growth in the time-varying model. Rejection of (null) hypothesis 4 (and acceptance of the first three hypotheses) implies a significant impact of the variables in the *X*-vector.

Since the selection of variables for the wage equation in the varying parameter model are based on the estimations of the previous section, it is not surprising that hypotheses 3 and 4 are rejected for all four (five) countries. The main interest lies in hypotheses 1 and 2. If hypothesis 1 is not rejected our equation explaining the unemployment coefficient is only approximately stochastic. If at the same time hypothesis 2 is rejected GUS has nevertheless a significant non-stochastic influence on the unemployment parameter. This would imply an (approximately) exact relationship between β and the generosity of the unemployment benefit system. If neither is rejected, we have the OLS-case with a constant coefficient on the unemployment rate.

The estimation results for equation (6b) are presented in table 2. Results for c_w and C_X in (6a) are fairly close to the results in table 1 and therefore not shown here.[7] Hypotheses 1 and 2 are only rejected in the case of Belgium. This implies that in France, Germany, (Italy) and The Netherlands changes in the generosity of the unemployment benefits do not have a significant impact on wage flexibility over the period 1960-1995. In Belgium, however, the generosity of unemployment benefits is a significant determinant of wage flexibility (over the period 1960-1995). For three of the four transformations of the generosity measure, we find the expected positive impact. The Δlog-transformation is wrongly signed but β_{av} is still negative. As expected, the estimated values of the parameters are approximately the same as their OLS-counterpart (-0.63) in table 1. Figure 2 shows the estimated time-varying wage flexibility based on $log(GUS_{t-1})$ and $\Delta log(GUS_{t-1})$. The series based on the differences jumps up and down a lot, whereas the series based on levels (logarithms) shows a much smoother pattern (as could be expected). As the standard deviations of both series do not differ very much (0.11 and 0.14, respectively), this finding suggests a stronger reaction to changes in generosity than to levels of generosity (also reflected in a higher value of the likelihood function).

[7] See Plasmans *et al.* (1999) for the full results.

Table 2: Estimation results for the time-varying parameter model

	GUS Trans-formation[e]	RWF (β) c_β	c_{GUS} [c]	β_{av} [d]	Max. F [b]
Belgium	*log*	-0.8809	0.0605**	-0.6059*	157.36
	log_{-1}	-1.2278	0.1375**	-0.6027*	157.41
	Δlog	-0.6274	-0.1312**	-0.6271*	157.36
	Δlog_{-1}	-0.5931	0.6112**	-0.5948**	158.90
	RR_{-1}	-1.4939	2.0845**	-0.6276**	159.14
France	*log*	-0.9599	0.0511	-0.7948	165.47
	log_{-1}	-0.9454	0.0471	-0.7976	165.42
	Δlog	-0.7091	0.0218	-0.7071	165.33
	Δlog_{-1}	-0.7122	0.2146	-0.6968	166.61
Germany	*log*	-1.5947	0.2852	-0.4750	165.72
	log_{-1}	-0.6942	0.0499	-0.4986	165.11
	Δlog	-0.5424	0.2147	-0.5419	165.40
	Δlog_{-1}	-0.5031	-0.1411	-0.5037	165.27
the Netherlands	*log*	4.0279	-1.0412	-1.1128	147.02
	log_{-1}	-0.1480	-0.1154	-0.7194	146.46
	Δlog	-0.6946	-0.6599	-0.6859	146.75
	Δlog_{-1}	-0.6375	0.8384	-0.6487	147.01

[a] columns under the heading C_β and C_{GUS} refer to the estimated coefficient as represented in (4b), β_{av} is the average of the time-varying value of β_t

[b] Maximum value of the loglikelihood function *F*, see appendix for a description of *F*.

[c] *, ** denote rejection of $H_0(2)$ at the 5% and the 1% level, respectively

[d] *, ** denote rejection of $H_0(1)$ at the 5% and the 1% level, respectively

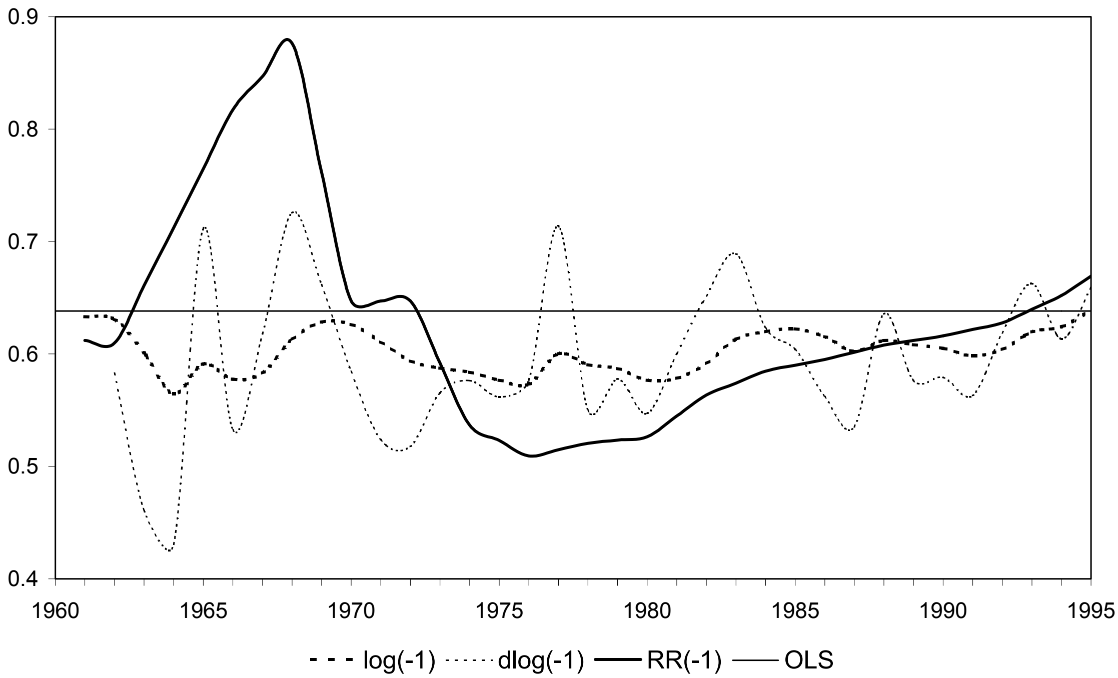
[e] *RR* stands for the replacement rate measure of the OECD (*cf. infra*)

A way of testing the robustness of these results consists in using the OECD replacement ratio series[8]. We generated a series with yearly observations by linearly interpolating (a quadratic or cubic interpolation did not alter the results) the two-yearly replacement ratio series and re-estimated model (4a-b) with the (lagged) replacement ratio as the generosity measure. The values for the test statistics for $H_0(1)$ and $H_0(2)$ are in general higher for all countries, but the conclusions are very similar: only in the case of Belgium does the replacement ratio significantly

[8] These series were taken from the OECD database on 'Unemployment Benefit Entitlements and Replacement Rates'. The 'replacement *ratio*' is calculated as an average replacement rate (benefits over income) over three periods of an unemployment spell (1st year, 2nd & 3rd year and 4th & 5th year), three family types (single, with dependent spouse and

influence real wage flexibility (see table 2, only Belgium reported). The resulting time-varying wage flexibility series for Belgium (RR) is also shown in figure 2.

Figure 2: Estimated time-varying real wage flexibility, $|\beta_t|$,based on log(GUS), Δlog(GUS$_{-1}$) and the lagged replacement ratio as compared to the time independent OLS-estimate in Belgium



Given the results above, we can conclude that, except for Belgium, there is no evidence of a relationship between wage flexibility and the generosity of the unemployment benefit system in the countries investigated. There are four possible explanations for these findings: 1) there is no relationship in reality; 2) a relationship exists but the changes in generosity during the period 1960-1995 have been too small to significantly affect flexibility (*cf.* Italy); 3) the data set is too limited to draw any firm conclusions and 4) there are interactions with other labour market institutions that are not present in the estimated model (e.g. the comprehensive disability insurance for employees (WAO) in the Netherlands). The first explanation is at odds with a large

with spouse in work) and two different levels of previous earnings (at 100% and at 66.7% of average earnings). For number of empirical studies using cross-section data. This leaves us with the three remaining explanations. Taking into account that explanations 3 and 4 prevent us from reaching firm conclusions, we tentatively conclude that for a number of countries changes in the benefit system will only affect real wage rigidity if the reform is substantial.

Conclusions

In this paper we have investigated the role of unemployment benefits in determining the degree of (real) wage flexibility. To this end we estimated a wage equation in a time-varying parameter framework for five core EMU countries. In Italy the unemployment benefit system is very limited. For the four other countries, the results show that, except for Belgium, wage flexibility is not related in a significant way to the generosity of the unemployment benefit system. This insight runs counter to the conclusions offered by cross-section studies. We therefore tentatively conclude that we should not be too optimistic about the effect of reform of the unemployment benefit system on wage flexibility and that such reform –in order to be really effective– should be radical.

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Appendix: Estimation of time-varying parameter models

Consider the following model:

$$y_t = \alpha_0 + \beta_t x_t + \gamma \overline{w}_t + \varepsilon_t \tag{A1}$$

with: $$\beta_t = \alpha + \delta z_t + \upsilon_t \tag{A2}$$

Assume that $\begin{pmatrix} \varepsilon_t \\ \upsilon_t \end{pmatrix} \overset{i.i.d.}{\sim} N(0, \begin{pmatrix} \sigma_\varepsilon^2 & 0 \\ 0 & \sigma_\upsilon^2 \end{pmatrix})$ and that x_t and $\overline{w}_t$ are uncorrelated with ε_t, so that

by defining $\lambda := \frac{\sigma_\upsilon^2}{\sigma_\varepsilon^2}$, we find from substitution of (A2) in (A1):

$$\mathrm{E}[y_t] = \mathrm{E}[\alpha_0 + \alpha x_t + \delta x_t z_t + \upsilon_t x_t + \gamma \overline{w}_t + \varepsilon_t]$$

$$= \alpha_0 + \alpha x_t + \delta x_t z_t + \gamma \overline{w}_t$$

$$\mathrm{var}(y_t) = \mathrm{E}[y_t - \mathrm{E}\{y_t\}]^2 = \mathrm{E}[\upsilon_t x_t + \varepsilon_t]^2 = x_t^2 \sigma_\upsilon^2 + \sigma_\varepsilon^2$$

$$= (1 + \lambda x_t^2)\sigma_\varepsilon^2$$

The joint loglikelihood-function for n observations is then given by:

$$\log L(\alpha_0, \alpha, \delta, \gamma, \sigma_\varepsilon, \lambda) = const - \frac{n}{2}\log(\sigma_\varepsilon^2) - \frac{1}{2}\sum_{t=1}^{n}\log(1 + \lambda x_t^2)$$

$$- \frac{1}{2\sigma_\varepsilon^2}\sum_{t=1}^{n}\frac{(y_t - \alpha_0 - \alpha x_t - \delta x_t z_t - \gamma \overline{w}_t)^2}{1 + \lambda x_t^2} \tag{A3}$$

With $r_t := \sqrt{1 + \lambda x_t^2}$ this reduces to:

$$\log L(\alpha_0, \alpha, \delta, \gamma, \sigma_\varepsilon, \lambda) = const - \frac{n}{2}\log(\sigma_\varepsilon^2) - \frac{1}{2}\sum_{t=1}^{n}\log(r_t^2)$$

$$- \frac{1}{2\sigma_\varepsilon^2}\sum_{t=1}^{n}\frac{(y_t - \alpha_0 - \alpha x_t - \delta x_t z_t - \gamma \overline{w}_t)^2}{r_t^2} \qquad \text{(A4)}$$

Now determine the optimal estimator for the constant term α_0. The first order condition gives:

$$\left.\frac{\partial \log L}{\partial \alpha_0}\right|_{\wedge} = 0 \quad \rightarrow \quad \hat{\alpha}_0 = \frac{\widetilde{y} - \hat{\alpha}\widetilde{x} - \hat{\gamma}\widetilde{\overline{w}} - \hat{\delta}\widetilde{z}}{R} \qquad \text{(A5)}$$

where: $R := \sum_{t=1}^{n}\frac{1}{r_t^2}$ (A6)

$$\widetilde{y} := \sum_{t=1}^{n}\frac{y_t}{r_t^2} \qquad \text{(A7)}$$

$$\widetilde{x} := \sum_{t=1}^{n}\frac{x_t}{r_t^2} \qquad \text{(A8)}$$

$$\widetilde{\overline{w}} := \sum_{t=1}^{n}\frac{\overline{w}_t}{r_t^2} \qquad \text{(A9)}$$

$$\widetilde{z} := \sum_{t=1}^{n}\frac{x_t z_t}{r_t^2} \qquad \text{(A10)}$$

Define the sum of squared residuals (SSQ) as follows:

$$SSQ := \sum_{t=1}^{n}\frac{\left[(Ry_t - \widetilde{y}) - \alpha(Rx_t - \widetilde{x}) - \gamma(R\overline{w}_t - \widetilde{\overline{w}}) - \delta(Rx_t z_t - \widetilde{z})\right]^2}{R^2 r_t^2} \qquad \text{(A11)}$$

The loglikelihood-function then can be written as:

$$\log L(\hat{\alpha}_0, \alpha, \delta, \gamma, \sigma_\varepsilon, \lambda) = const - \frac{n}{2}\log(\sigma_\varepsilon^2) - \frac{1}{2}\sum_{t=1}^{n}\log(r_t^2) - \frac{1}{2\sigma_\varepsilon^2}SSQ \qquad \text{(A12)}$$

From which we deduce: $\hat{\sigma}_\varepsilon^2 = \dfrac{SSQ}{n}$ (A13)

The loglikelihood-function (A12) strongly resembles the loglikelihood-function of the following linear relationship:

$$\frac{Ry_t - \tilde{y}}{Rr_t} = \alpha \frac{Rx_t - \tilde{x}}{Rr_t} + \gamma \frac{R\overline{w}_t - \tilde{\overline{w}}}{Rr_t} + \delta \frac{Rx_t z_t - \tilde{z}}{Rr_t} + \varepsilon_t \qquad \text{(A14)}$$

For a given value of λ we find from OLS (A14) the estimators for α, γ and δ which can be substituted back in the loglikelihood-function. This results in:

$$\log L\left(\hat{\alpha}_0, \hat{\alpha}, \hat{\delta}, \hat{\gamma}, \hat{\sigma}_\varepsilon, \lambda\right) = const - \frac{n}{2}\log\left(\hat{\sigma}_\varepsilon^2\right) - \frac{1}{2}\sum_{t=1}^{n}\log\left(r_t^2\right) - \frac{n}{2} \qquad \text{(A15)}$$

which is of the form: $\log L\left(\hat{\alpha}_0, \hat{\alpha}, \hat{\delta}, \hat{\gamma}, \hat{\sigma}_\varepsilon, \lambda\right) = \tilde{c} + F(\lambda)$ (A16)

where: $F(\lambda) = -\dfrac{n}{2}\log\left(\hat{\sigma}_\varepsilon^2\right) - \dfrac{1}{2}\displaystyle\sum_{t=1}^{n}\log\left(r_t^2\right)$ (A17)

Maximisation of the likelihood-function (A3) is thus equal to the maximisation of (A17) with respect to λ, which can be solved by e.g. a random search procedure.