Nominal Wage Rigidity in Europe: estimates and institutional causes.

Orietta Dessy
Università di Milano

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Abstract

In this paper we construct and compare different measures of nominal wage rigidity for the EU countries using the 1994-2000 waves of the European Community Household Panel. The observed distributions of nominal wage changes show a relevant percentage of nominal wage cuts and freezes across countries. When measurement error is taken into account in an econometric model of wage changes appropriately estimated, it explains almost the totality of wage cuts observed. Therefore the extent of nominal wage rigidity is quite high in Europe. Institutional causes of wage rigidity are investigated, finding an "hump-shaped" relationship between nominal wage flexibility and both employment protection legislation and coordination. On the other hand, an "u-shaped" impact of union coverage on measures of downward wage rigidity is found .

JEL Codes: J30, J31, J51.

Keywords: Nominal wage rigidity, measurement error, intercountry comparison, institutions in labor markets .

1 Introduction

The target of low inflation for monetary policy has been always quite debated in the literature, given the recent position of the European Central Bank. The dispute is based on a crucial assumption on nominal wage determination. Typically employment depends on the level of real wages, representing the cost of labour of the firm: firms hit by a positive idiosyncratic demand shock may want to rise wages and increase employment, whereas firms hit by a negative demand shock may want to cut costs (reduce real wages) and reduce employment. Substantial real wage reductions can however only be achieved by slowing down nominal wage growth below the inflation level, or (if inflation is too low) by cutting nominal wages. If nominal wages were completely flexible there would be no real impact of inflation decreases on output and employment. According to this view (Ball and Mankiw, 1994, Gordon, 1996), any downward wage rigidity that may exist would be the result of an inflationary environment, and the society would adapt to a zero inflation policy without large and persistent effects on output and unemployment. On the contrary, it is argued (Tobin, 1972, Holden, 1994, Akerlof, Dickens and Perry, 1996, 2000) that when nominal wages are downwardly rigid and inflation is low, firms may have difficulties in cutting costs through wage adjustments and may turn to lay-offs instead, which would result in higher unemployment. In this context, it may be appropriate that the ECB relaxes its inflation target to increase wage flexibility and reduce unemployment.

There is a quite widespread literature on the effects of low-steady inflation on wage formation[1], based on the assumption of nominal rigidity in wages. This assumption has been usually tested using aggregated macro-data. The recent availability of individual panels of different nature (survey, administrative files, interviews) has given rise to a relevant number of papers aimed instead at measuring the extent of nominal wage rigidity at the micro-level. All the existing studies are based on the analysis of individual nominal wage change distributions. Evidence is available for the US and for a number of European countries. However, the different characteristics of the data used induce difficulties in inter-country comparisons. The recent availability of the European Community Household Panel (ECHP), collected by Eurostat since 1994, seems to overcome this problem, since in theory it presents the unique feature that the same questionnaire is asked in 15 countries of the EU. Moreover, the data cover most of the '90s, a period of relatively low and stable inflation in Europe. This makes the analysis of

[1] See Holden (2004) for a review.

wage rigidity with the ECHP particularly interesting for policy purposes, as the phenomenon of downward nominal wage rigidity can induce real effects when the level of inflation is low. Moreover, the inter-country dimension can help in finding possible explanation for the existence of nominal wage rigidity.

Some preliminary evidence on individual nominal wage change distributions, based on the first three waves of the ECHP, has been given in Dessy (2002). In that study, the institutional characteristics of labour markets were considered as possible explanations for the extent of nominal rigidity observed, measured as the frequency of wage cuts. Although alternative explanations for nominal wage rigidity have been explored in the literature, mainly based on fairness consideration and money illusion, the institutionalist view (Holden 1994, 1999, 2004; Groth and Johansson, 2001) seems to be the preferred interpretation for the European Central Bank (ECB). In fact, according to ECB (2003), 'structural labour market reforms are expected for reducing the role of downward nominal wage rigidity and sustain the low inflation target'. Another paper where the issue of institutional explanations for downward nominal wage rigidity is taken up from an empirical point of view is Holden and Wulfsberg (2004), but data are aggregated at industry level over the 1973-1999 period.

In this paper we extend the analysis in Dessy (2002) to 7 instead of 3 waves of the ECHP, covering the 1994-2000 period and fifteen EU countries: Germany, France, UK, Italy, Spain, Netherlands, Belgium, Luxembourg, Ireland, Denmark, Sweden, Finland, Austria, Portugal and Greece. The availability of more data for each country allows us to overcome three important limits of that preliminary analysis: 1) the construction of measures of wage rigidity that take into account measurement errors; 2) an extended analysis of the impact of institutional variables on wage rigidity, allowing for a certain variability of institutions; 3) the evaluation of the impact that nominal wage rigidity might have had on unemployment in Europe.

Although the ECHP has the great advantage of allowing inter-country comparisons, as most of the individual surveys available it is subject to two kind of measurement error: rounding behaviour of individuals and reporting errors. The best way for determining measurement errors characteristics is to carry out validation studies that compare data from different sources, normally survey and administrative data. This task is clearly very difficult for the ECHP, and to date no comprehensive study is available for correcting precisely wages from measurement errors in this data[2]. We therefore follow

[2] Hanish and Rendtel (2003) consider rounding in the German and Finnish ECHP wage

a structural approach, in which measurement error is modelled according to the classical assumptions. In particular, we use a simplified version of the Altonji and Devereux (2000) model, very similar to the one considered by Fehr and Goette (2002), for estimating nominal wage rigidity measures comparable across countries. We find that, although the observed percentage of nominal wage cuts and freezes is quite relevant and different across the European countries, when measurement error is taken into account it explains almost the totality of wage cuts observed. Therefore there is a quite high degree of nominal wage rigidity in Europe.

Following a two-step procedure, we then carry out a regression approach, treating the institution variables as cardinal (Bean 1990, Scarpetta 1996, Holden and Wulfsberg (2004)). This approach leads to the following strong results. First, there emerges a significant non-linear impact of the employment protection legislation variable (*epl*) on nominal wage flexibility. Such effect always comes under the form of a "hump-shaped" relationship between *epl* and hourly wage cut frequencies, however estimated. Second, we find a significant "u-shaped" impact for the coverage variable (*pcov*), with the decreasing portion of the curve predominating over the increasing. Third, we find an "hump-shaped" impact for coordination (*coord*), with the increasing portion of the curve predominating over the decreasing. These results are robust to: a) the choice of the wage cut frequency variable; b) the choice of the centralisation variable; c) the inclusion of time dummies and all macroeconomic controls; and finally d) the treatment of the possible endogeneity of the macro variables within an instrumental variable estimation framework. For the centralization variables, instead, we are unable to report robust and significant results.

To get a deeper understanding on the implications of the estimated non-linearities, we supplement our empirical analysis with some simple comparative statics from the regression estimates for sizeable changes in the institutional variables. We find that one standard deviation increase or decrease from the average value of *epl* brings about in either case a strong reduction in expected wage cut frequencies, by around 20 and 30 percentage points respectively. For *coord* instead, the increasing part of the relationship turns out to be predominant, so that one standard deviation increase leads to only small rises in expected cut frequencies, no higher than 2 points; on the other hand one standard deviation less of *coord* implies a reduction by around 12 points. For *cov* we observe a reduction of around 10 points in

data. Peracchi and Nicoletti (2003) analyse the distorsive impact of imputation and non-response on income data in the ECHP.

expected cut frequencies when it increases by one standard deviation and a stronger increase, by more than 30 points, when it decreases.

The economic interpretation of the foregoing results is that a higher degree of nominal wage flexibility is supported by a labour market regulated by not too strict employment protection rules, with a moderately small percentage of workers covered by collective agreement and a sufficiently high degree of consensus between the collective bargaining partners. The insignificance of any of the centralisation variables in the context of a general empirical model is not surprising, and even tends to confirm the widespread consensus about the relatively higher importance of coordination and coverage. This wisdom is clearly summarised in OECD (1997), where it is remarked that "even relatively centralised bargaining will have little impact if few workers are covered."

The structure of the paper is the following: in Section 2 we summarise the existing literature on nominal wage rigidity. Section 3 gives some information on the ECHP. Section 4 introduces wage distributions for all countries, whereas Section 5 considers wage change distributions, presenting the frequencies of nominal wage cuts and freezes observed in the data. Section 6 deals with the estimation of measures of wage rigidity. Section 7 carries out the regression analysis. Section 8 concludes.

2 Related literature

The interest in analyzing individual panel data in order to assess whether wages are rigid or flexible dates back to McLaughlin (1994)'s paper. From that date on, a number of similar analyses have been carried out both for the US and some European Countries. There is no agreement among researchers about the extent of downward wage rigidity even for the US, where most of the analyses have been carried out using the same data-set, the Panel Survey of Income Dynamics (PSID). Intercountry comparisons are difficult because surveys are collected according to different criteria, and therefore the information available both for defining the subsample and the variable of interest may differ across countries.

The existing evidence from individual surveys seems to support the idea that wages are rigid in nominal, and not in real terms. But it is not clear the extent of downward nominal wage rigidity at individual level. Using data from the PSID, McLaughlin (1994) finds that on average over 1976-86 there were 17% stayers experiencing wage cuts and 7% had zero nominal pay growth. He uses two measures for wages: *earnings* and *hourly earnings* concluding that wages in the US are quite flexible.

Instead of pooling together periods of high and low inflation, subsequent studies have focused on yearly wage changes, finding evidence of a spike at zero in the distribution of nominal wage changes. This spike is taken as evidence of nominal wage rigidity. But, although being quite asymmetric around zero, wage changes distributions are not completely downwardly rigid. Kahn (1997) distinguishes between 10.6% wage cuts for *wage* earners and 24.3% cuts for *salary* earners, finding also a strong evidence of 8% nominal wage rigidity during 1971-88. Also, according to Card and Hyslop (1997), despite many individuals in the PSID report wage cuts, there is clear evidence of nominal wage rigidity. In particular, they find that the spike at zero *hourly* wage changes spans from 7% in a 10% inflation environment to 15% when inflation fell to 5%, both for salary and wage earners. Therefore, the spike at zero nominal wage changes is highly sensitive to the rate of inflation.

Data at firm level (Altonji and Devereux (2000)'s personnel file of a large firm, Bewley (1998)'s interview study that involves 300 business people) show much higher levels of wage rigidity, measured as *hourly* wage, and basically no wage cuts. Using different methodologies, McLaughlin (1994), and Card and Hyslop (1997) argue that measurement errors can not explain all the percentages of wage cuts observed in individual survey's data, whereas by estimating an econometric model Altonji and Devereux (2000) explain all the wage cuts observed in the PSID with measurement error. Therefore, in the PSID, measurement errors apparently reduce the observed percentage of nominal wage rigidity and increase the percentage of wage cuts.

Similar analyses carried out in some European countries seem to give different results. Goux (1997) compares two different sources of data available for France: the 1976-92 Déclarations Annuelles de Donnèe Sociales (DADS), an administrative, potentially error-free data-set, and the 1990-96 French Labor Force Survey (LFS). Using *annual earnings* as a measure of wages, and therefore not controlling for the number of hours, she finds that the amount of wage cuts is similar in the two data-sets and affects approximately 25% of stayers employed full-time[3].

Smith (1999) and Nickell and Quintini (2001) examine the UK using different data sources. Smith (2000) analyses the British Household Panel Survey (BHPS). She uses *weekly earnings* as a measure for wages and stayers

[3]Interestingly, a good percentage of wage cuts for the stayers can be explained with one of the following: 1) better working conditions, 2) decrease in annual bonus, and 3) 4-digit change in occupation.

who do not change the number of hours worked as the sample of interest. She finds that the percentage of employees whose wage is constant from one year to the next is 9% during 1991-6 and the percentage of wage cuts is on average 23%. Using the unique feature of the BHPS, that allows to consider the subsample of people whose payslip has been checked by the interviewer, Smith (1999) focuses on the employees whose reported earnings are error-free, and who do not receive bonuses or overtime pay, finding that only 1% of them had zero pay growth and 18% received wage cuts. Therefore, contrary to what has been found for the US, measurement errors seem to be the main source of wage rigidity, increasing the percentage of nominal wage rigidity observed in survey data. As a consequence wages appear much more flexible in the UK than in the US. Using the UK New Earnings Survey (NES) from 1997 to 1999, Nickell and Quintini (2001) find on average lower percentages of both no change in wage and wage cuts than Smith (2000) for the same 1991-96 period, but they use a different measure of wage. In the NES data are provided by employers and come directly from payroll records, which ensures a high degree of accuracy. The measure of the nominal *hourly wage* rate used is the weekly pay of those whose pay is unaffected by absence, excluding overtime pay, divided by weekly hours excluding overtime hours. Again, only full-time employees not changing job are considered.

Fehr and Goette (1999) analyze *earnings per working hour* of stayers in the Swiss LFS during the period of very low inflation 1991-96. They find 12% of rigid wages and 25% of wage cuts. Allowing for measurement errors they estimate Altonji and Devereux (2000)'s econometric model obtaining results similar to those found for the US: measurement errors can explain most of the observed wag cuts, that actually turn in no wage changes.

3 Data

The ECHP is a recent large-scale longitudinal study set up and funded by the European Union. The great advantage of the ECHP is that information is given not only at household, but also at individual level. In the first wave (1994) a sample of about 60,500 nationally representative households - i.e. approximately 130,000 adults aged 16 years and over - were interviewed in the then 12 Member states. Austria (in 1995) and Finland (in 1996) have joined the panel since then. From 1997 onwards, similar data are available for Sweden. In fact, ECHP UDB variables were derived from the Swedish Living conditions Survey and are now included in the ECHP UDB. In wave 2, EU-13 samples totalled some 60,000 and 129,000 adults. For the fourth wave of the ECHP, i.e. in 1997, the original ECHP surveys were stopped in three countries, namely Germany, Luxembourg and the United Kingdom. In these countries, existing national panels were then used and comparable data were derived from the German and UK survey back from 1994 onwards and for the Luxembourg survey back from 1995 onwards. Consequently, two sets of data are available for the years 1994 to 1996 for Germany and the UK, and 1995-1996 for Luxembourg. Eurostat recommends the use of the original ECHP data for any analysis covering only the years 1994-1996 for countries with two different data-sets for the same year. However, for longitudinal analysis covering more years, the converted data-sets should be used. In this chapter we use all the sources available for each country, so that when there are two data-sets for the same country they can be compared.

Although the same questionnaire, centrally designed by Eurostat, is asked in all the countries belonging to the project, different interviewing methods are carried out on different countries. The recommended method is telephone or proxy interview, but in Greece, Netherlands, Portugal and the UK interviews are carried out, at least partly, using computer assisted personal interviewing (CAPI). This heterogeneity between countries can affect the quality of individual salary and earnings reported.

The sample of interest

To facilitate comparisons with previous studies on wage rigidity we concentrate on employees, excluding self-employed from our analysis. Employees are detected as people reporting wages. The sample we are interested in is composed of *stayers,* i.e. employees who do not change firm. Since the firm identifier is not available, one possible way for selecting stayers is to use the information about employees' monthly status, considering only individuals who have been continuously employed from one wave to the next. A further check for employees not changing sector and occupation is advisable, although there is still the possibility of keeping in the sample employees changing employer, but not occupation and sector, without experiencing any unemployment period. Unfortunately, for some countries (Denmark, Belgium, Luxembourg PSELL, Ireland and Sweden) the information on sector and occupation is missing for a number of waves. Since we decided to use all the information available to get as close as possible to the precise definition of stayers, the sample is not defined homogeneously across countries. A summary of how stayers have been defined in the various countries is given in Tab. 3.1.

Unfortunately we can not distinguish employees paid by the hour from those paid weekly. But we have quite detailed information about the type of employment contract. In particular, we know whether the employee is working part-time or full-time[4]. We consider only the sub-sample of *stayers full-time,* the majority of observations in our sample.

Although most of the previous analyses in this field of research focused on the private sector, we pull the public and private sectors together, as wages in the two sectors turn out to be highly correlated in all the European countries.

Measures of wages and hours

In the first 3-waves version of the ECHP, used in Dessy (2002) only net wages were available. The current 7-waves version we are working with gives instead both net and gross salary and earnings. As explained in the survey, the ideal measure to work with for measuring wage rigidity would be the base hourly wage. As in most of the individual surveys, in the ECHP base hourly wages are not reported. However, two measures of labour earnings are available: "current wage and salary earnings" (i.e. earnings received in the month of the interview) are given both gross and net of individual taxes; and "total wage and salary earnings" (referred to the year before the interview). We decided to take "current gross wage and salary earnings" as the most useful measure of wages for two reasons: 1) "current net wages" can be subject to individual, familiar, or institutional shocks; and 2) the number of months which "total wage and salary earnings" is referred to is not reported.

Since the number of "weekly hours worked in the main job" (always in the month in which the interview was taken) are known, it is also possible to calculate "hourly current earnings" dividing monthly wages by the number of hours. This is clearly only a proxy for the contracted base wage, since it is biased from overtime hours, overtime pay, monthly bonuses and premia. The bias given from this variable part of labour earnings is what we call "reporting error". Another way of getting closer to a measure of the increase in basic wages, adopted in Smith (2000), is to study pay growth when there are no hours changes[5]. For the purposes of a validation study, Smith's method is better than ours because it does not change the value directly reported by individuals, and allows to study their rounding behaviour. Dividing wages by the number of hours can hide rounding error. But, at the same time, focusing on individuals not changing the number of hours worked can induce

[4] From 1995 on, we also know the type of contract (permanent, fixed-term or short-term, casual with no contract, other working arrangements) and, for temporary contracts, the length of the contract.

[5] Clearly in this case both total and hourly earnings changes coincide.

strong sample selection biases, especially in countries where the number of hours is quite flexible. Also, if hours are reported with error, selecting individuals on the basis of this information does not help in eliminating this second source of measurement error. Moreover the fact that, although employees keep constant the number of hours worked from one period to the next, overtime pay or benefits can change over time, makes the observed measure of wage for this sample still biased by reporting errors. Since from our trials we realised that applying Smith's method we were losing many observations without changing qualitatively our results, we decided to divide monthly earnings by the number of hours, as in the majority of previous works on the subject, and then introduce formally a measurement error in a structural model framework.

Since all the above measures are referred to the month in which the interview was taken, we checked that comparing two different months of the year does not bias our results. Quite often, restricting the sample to people whose interview was taken in no more than two months difference in the two years period considered, reduces dramatically the number of observations[6]. In qualitative terms, however, our results do not seem to change significantly. In order not to loose too many observations, we therefore decided to keep in the sample people who reported their wages in different months for consecutive interviews.

4 Wage distributions in the Europanel

We start our analysis on wage rigidity by presenting wage, hours and hourly wage distributions. Wage dynamics are shown in the next section. We find useful to separate the two issues because this helps in explaining the impact of the two components of measurement error (rounding and reporting errors) on the unobserved base hourly wage. In particular, looking at the values directly reported by individuals gives us an idea of the extent of rounding in the ECHP survey. Comparing monthly and hourly wage change distributions together with hours changes is useful instead for understanding the impact of the number of hours on hourly wage changes.

In the ECHP we do not know whether net and/or gross wages have been directly reported by individuals, therefore in Fig. 3.1 we present, in the first column, the distributions of gross wages, in the second column net wages and, in the third column, the distribution of gross wages divided by the number of hours. We show only one year for each country (1995

[6] In Germany, for example, the month of the interview is not reported.

or 1996) because the shape of the distributions does not change relevantly over time. Comparing the three distributions we can see that, although their general shape changes across countries, none of them is pretty smooth: they all present spikes at rounded values. In all the European countries the percentage of rounded wages[7] is quite high, about 80% in almost all countries.

The characteristics of rounding error in the ECHP have been analysed for Germany and Finland by Hanish and Rendtel (2001, 2002). The reason why only these two countries are considered is that for them the authors have access to the original release of the panel, the so-called Production Data Base (PDB)[8], that is richer of information than the UDB, although the original variables can differ across countries. For Finland, survey data can also be compared with administrative records. Hanish and Rendtel (2002) find that rounding errors on personal gross wages are quite relevant: they are related to the level of wages (better to the number of digits), and to many individuals characteristics. This has an error effect on income quantiles and derived statistics like the Gini coefficient and poverty measures, but also on wage equation estimates, where measurement error is assumed to follow the classical assumptions. Rounding error has also some impact on wage mobility, i.e. growth rate of labour earnings: small wage changes are often rounded to zero, and exceptional changes are often under-reported.

Although the results in Hanish and Rendtel (2001, 2002) let us skeptical about making the usual normality assumptions for rounding errors in the ECHP Finnish and German panels, we are to date not sure that the same results are valid for all countries. But an overall validation study of the ECHP is out of the scope of this paper, therefore we will assume that rounding errors are normally distributed in all countries. We consider each country separately, and measurement error, when not modelled according to the classical assumption, will be taken into account in the meta-analysis framework carried out in a second step, with country-specific effects.

The reason why we present both gross and net wage distributions is that in the PDB often only one of the two has been reported. Therefore many values have been imputed when converting the original Production Data Base (PDB) in the user-friendly version User Data Base (UDB), accessible

[7] By rounded wages we mean wages ending with as many zeros as are the number of chiphers of the national currency minus two. This 'rule of the thumb' has been used in Hanish and Rendtel (2001, 2002).

[8] This research is part of CHINTEX, an EU-sponsored reserch project on the harmonisation of panel surveys.

to researchers[9]. Nicoletti and Peracchi (2004) deal explicitly with this issue, trying to evaluate the impact of imputation both on wage and wage change distributions. They use a variable contained in the household file, indicating whether the value has been imputed or not. Selecting only single person households they can evaluate the impact of imputation methods on wages and salary reported at individual level. They find that, although the imputation procedure distorts wage distributions, the percentage of imputed values is not very high. Therefore, there are not major problems for wage distributions. As far as wage change distributions are concerned, imputed values increase the percentage of extreme values. The consequence is that, whereas the mean of the distribution is highly biased, the median is a less distorted location measure of wage change distributions. We take into account of this problem by eliminating 1% of observations in both tails of wage change distributions[10]. As far as rounding errors are concerned, both gross and net wage distributions present spikes at rounded values, therefore the percentage of rounded wages is quite high for both measures.

As we can see from Fig. 3.1 dividing monthly wages by the number of hours does not cancel out completely the existence of many spikes in wage distributions. However, hourly wage distributions are overall more flexible and smoother than gross and net wages directly reported by individuals.

It is interesting to notice the particular shape of wage distributions in countries, such as France, Luxembourg and Portugal, where a minimum wage is fixed at the national level. There is a clear drop on the left of the minimum wage and a little spike where the distribution starts, indicating the quite high number of people getting the minimum wage. In countries such as Greece, Spain and the Netherlands the phenomenon is less pronounced, probably because the level of the minimum wage fixed is very low. Wage levels lower than the minimum wage are quite common in stayers full time wage distributions. Often they are interpreted as measurement errors, but sometime they can be explained with particular contractual arrangements. We therefore keep all the observations in our sample.

In Appendix 1 (Tab. A1.1 - A1.3) we give descriptive statistics of gross wage distributions, number of weekly hours and hourly wages for every year in each country. We can see that, on average, wages are increasing in all countries. This is not surprising, as we are working with nominal wages that usually follow the inflation rate. On the contrary, the distribution

[9] See Peracchi (2002) and Nicoletti and Peracchi (2002) for a detailed evaluation of the ECHP data.

[10] Cuts of the tails of wage change distributions are widespread in this literature, for eliminating outliers.

of the number of hours is quite stable over time for each country. The average number of weekly hours is 41 hours and the standard deviation is 6.5. However, there are countries where hours are more flexible (the UK), and countries (such as Luxembourg and Portugal) where the number of hours is more rigid than the average. As a consequence, hourly wage distributions are overall increasing over time.

Summarising, the most important feature of wage distributions in the ECHP is the pervasive phenomenon of rounding in reported wages. Dividing wages reported by the number of hours does not eliminate the high percentage of rounded wages. Although the empirical evidence available for Finland and Germany is contrary to assuming that rounding errors are normally distributed, we are not able to validate the data for all countries, and therefore will stick to the classical assumptions for rounding in the estimation of measures of wage rigidity.

5 Wage change distributions: observed measures of wage rigidity

In this section we focus on the characteristics of hourly gross wage distributions in each country. In fact, this is the measure of wages closer to the base wage contracted, although subject to both reporting and rounding errors. Our purpose is to construct a first data-set that collects respectively the percentage of wage cuts, no wage changes, and wage rises observed for each country.

According to the descriptive approach, we are interested in three features of wage change distributions: 1) a spike at zero nominal wage changes as evidence of nominal wage rigidity; 2) a spike at the rate of inflation for real wage rigidity; 2) symmetric drops around zero for menu-costs effects; 4) the percentage of wage cuts, and symmetry of the distribution around zero for downward wage rigidity.

Fig. 3.2 shows wage change distributions for all countries, all years. A bar has been drown at zero and at the rate of inflation for every year-change.

The histograms show that in all the European countries nominal wage changes have a prominent spike at zero. We also observe a sharp drop for little wage changes in stayers' distributions, with higher positive changes of wages more likely to occur. For most countries, there is clear evidence of downward nominal wage rigidity as the distributions are asymmetric. At the same time, wages are not completely downwardly rigid across the European countries: the percentage of wage cuts reported are quite high.

In most of the countries we observe a second, small spike in the nearby of the rate of inflation: real wage cuts are much more frequent than nominal wage cuts. From a first inspection of qualitative characteristics of wage change distributions we therefore conclude that: 1) there is evidence of nominal wage rigidity; 2) there can be a certain extent also of real rigidity; 3) there is no support for the menu-costs theory; 4) wages are not completely downwardly rigid.

In this paper we focus on nominal wage rigidity issues, and therefore we are particularly interested in exploring the exact percentage of wage cuts and the frequency of no wage changes observed. As we can see, there are interesting differences among countries from a quantitative point of view. In particular the extent of the spike at zero varies across countries, but it is important to notice that the spike is constructed around zero, and therefore it includes small positive and negative wage changes. We discuss inter-countries differences referring to Table 3.2, which gives wave by wave the precise figures for the percentage of cuts, freezes and rises in monthly wages, hours and hourly wages[11].

First of all we can notice that, dividing monthly wages by the number of hours, the percentage of rises does not change much in all countries, whereas spikes decrease and cuts rise. Therefore, considering the number of hours worked induces downward wage flexibility. It seems that people tend to increase the number of hours worked for keeping their total labour earnings constant, or not letting them fall down dramatically. Normally, when in administrative data the number of hours is not observed, as in Fehr and Goette (2003), Devicienti (2002), Knoppik and Beissinger (2001), they are the only component of measurement error and are modelled with the classical assumptions. But this might be incorrect if the impact of hours is asymmetric on wage change distributions.

Clearly, the impact of changes in hours on downward wage flexibility is stronger in countries where hours are more flexible. For example, in Germany, the UK, Belgium, Spain and Ireland, where only less than 50% of employees do not change the number of hours, the spike at zero hourly wage changes is less than half of the spike for monthly wage changes. Instead in Denmark, the Netherlands, Luxembourg, France, Italy, Greece, Portugal, Austria and Finland, where more than 50% of employees do not change the number of hours, the reduction of the spike at zero when dividing by the number of hours is less pronounced.

[11] Detailed descriptive statistics for the same distributions can be found in Appendix 1 (Tab. A1.4 - A1.6).

Comparisons with previous results

In general, we can say that wage change distributions from the ECHP bear the same features as the distributions constructed from similar panel data in the US and other European countries. On average, the percentages of rigid wages and wage cuts in Europe are not far away from those observed in the US for similar rates of inflation, but there are enormous differences across countries.

The numbers that we find found for the UK are different from Smith (2000)'s, although the panel used is the same. For the years 1994-95 and 1995-96, before controlling for the payslips and therefore correcting rounding errors, Smith (2000) finds respectively 9.4% and 7.8% wages unchanged and 22.5% and 23.4% wage cuts. But she focuses on individuals not changing the number of hours, probably a small sample in the UK, where the number of hours worked is quite flexible. Over the '90s we observe instead 32% cuts and 2% freezes. Nickell and Quintini (2003) find far less cuts (20% on average) in the error-free New Earnings Data and on average 2% no wage changes over the '90s, but the measure they observed is the base hourly-wage not distorted by overtime pay, bonuses and premia. This can explain the higher proportion of cuts in the BHPS than in the NES.

Comparisons with Goux (1997) for France can be carried out only for monthly wages. She considers gross earnings for full-time workers in the French Labour Force Survey (LFS) finding respectively, in 1994-95 and 1995-96, 11.5% and 12% full-time workers whose earnings did not change and 27% and 28% wage cuts. Our first wage change computed for France gives an unreliable percentage of 80% of cuts, probably due to data problems for which to date we do not have any clear explanations. However, in general in the French ECHP we observe higher percentages of wage cuts and lower percentages of rigid wages than what found in the French LFS.

A suspiciously high increase of wage freezes is observed in Greece between 1999 and 2000 (almost the double of the previous year-change). Proportions of wage cuts observed higher than 44% will be considered outliers and eliminated from the inter-country analysis in the next section.

In Belgium, Borgijs (2001), finds about 20% cuts and 12% freezes in the '90s. Therefore, with the respect to what found in the ECHP, about 10% less cuts and 5%-6% more no wage changes. But, although dividing by the number of hours, he considers net and not gross wages. As a consequence the two results are not directly comparable.

Conclusions on the observed measures of nominal wage rigidity

A spike at zero nominal wage changes and a relevant frequency of nominal wage cuts seem to be common characteristic of the distributions of nominal wage changes constructed from survey-data in all the ECHP countries. We can therefore conclude that there is evidence in Europe of nominal wage rigidity, although wages are not completely downward rigid. A first ranking of the EU countries can be based on: 1) the extent of the spike; 2) the percentage of hourly cuts observed. Countries with the highest percentages of zero wage changes are Austria and Italy, followed by Finland, Denmark, Belgium, Portugal, and the Netherlands. Germany, Luxembourg and Greece have a slightly smaller percentage of wage rigidity. The countries with the most flexible wages turn out to be France, the UK, Ireland, and Spain. There may be over time small changes of the above ranking.

Looking at hourly wage cuts, we can rank Spain, Germany, the UK, France, and Belgium among countries with the highest percentage of cuts, followed by Austria, Italy, and then Ireland, Finland, Luxembourg and the Netherlands. Among countries where wage cuts are more rare, we can mention, in decreasing order: Greece, Denmark, and Portugal[12]. Therefore, the unit of measure used is crucial for inter-country comparisons.

There are two major limits in using the observed percentages of wage cuts and freezes for cross-country comparisons. First, for comparative purposes we need to assume that measurement errors have the same characteristics across countries. But, if rounding depends on the number of digits in wage levels, this might not be the case.

The second limit is referred to the assumption on the counterfactual for identifying wage rigidity. For direct comparisons across countries, according the descriptive approach, we are implicitly assuming that the counterfactual distribution is the same across countries. This is not necessarily true, because due to country-specific characteristics, the hypothetical distribution supposed to be observed in a perfectly flexible regime would be not only smooth, but also with different shapes across countries.

We therefore try to estimate the percentages of wage cuts and wage freezes using a structural approach, in which 1) the counterfactual is estimated country by country using observable individual characteristics; and 2) measurement error is taken into account. Unfortunately, the only way we can model measurement error in this context is by introducing the classical assumptions.

[12] Notice that for these rankings we have considered national surveys for Germany, the UK and Luxembourg. The descriptive results are quite different however if we consider the first three waves of the ECHP panel for the above countries.

6 Estimating nominal wage rigidity in the EU countries: a structural approach

There might be reasons - e.g. efficient nominal wage contracts, nominal fairness standards and nominal loss aversion - that render nominal wage cuts costly for the firms. Therefore firms will not implement all desired wage cuts and, as a consequence, there will be a difference between the desired or notional wage cut and actually implemented wage cuts. Our estimates of the extent of nominal wage rigidity in Europe are based on an easier specification of the Altonji and Devereux (2000) (AD) model proposed by Fehr and Goette (2003) in its initial simplified version in which only the threshold α is estimated, and the links with the theoretical model behind AD are relaxed.

A plausible justification for abadoning the original AD specification of the econometric model is that in the EU countries an application of the MM model is difficult to interpret. In fact, although valid for contracts that can be renegotiated only by mutual consent, the MM model does not take into account the fact that wages are determined at different levels in the European countries. Since the role of unions is ignored, we might question about the applicability of hold-up models in the EU countries.

According to the initial specification of the AD structural model, actual (or observed) wage changes follow notional wage dynamics only when the change is positive. Wage cuts are implemented only if they are larger than a threshold-level α. If wage cuts are below this threshold, they are not implemented and workers are affected by nominal wage rigidity. The general structure of the model that we estimate is the following:

$$\Delta y_{it} = \begin{cases} x_{it}\beta + e_{it} & if \quad 0 \leq x_{it}\beta + e_{it} \\ 0 & if \quad -\alpha \leq x_{it}\beta + e_{it} \leq 0 \\ x_{it}\beta + e_{it} & if \quad x_{it}\beta + e_{it} \leq -\alpha \end{cases}$$

where Δy_{it} is the observed log nominal wage change of individual i in period t, $x_{it}\beta + e_{it}$ is the notional wage change that would be implemented in absence of downward nominal wage rigidity, x_{it} are a set of variables that are likely to affect wage growth, e_{it} represents the usual error term. As we can see, when wage cuts are implemented they follow exactly the outside option of employees. This is different from what implied by the MM model, and specified in the AD model, according to which wage cuts, when implemented, follow the outside option of the firm $(\lambda + x_{it}\beta + e_{it})$. In a sense, our specification of the model gives more power to workers, which is probably the case in Europe.

Introducing measurement error m_{it}, which can be interpreted as rounding and reporting error in the ECHP, the model becomes:

$$\Delta y_{it} = \begin{cases} x_{it}\beta + e_{it} + m_{it} & if \quad 0 \leq x_{it}\beta + e_{it} \\ m_{it} & if \quad -\alpha \leq x_{it}\beta + e_{it} \leq 0 \\ x_{it}\beta + e_{it} + m_{it} & if \quad x_{it}\beta + e_{it} \leq -\alpha \end{cases}$$

Since both e_{it} and m_{it} are i.i.d as Normal with mean zero, the parameters we estimate are: α, σ_e, σ_m and β.

In our empirical estimates below it is important that x_{it} contains variables that capture business cycle variation in wages, and individual characteristics correlated with wage growth. Variables normally used in the literature are: labour market experience, age, tenure, and observable skills. The inclusion of these variables is suggested by many papers (e.g. Topel, 1991), and in previous estimations of the wage change model above they resulted very significant. Unfortunately in the ECHP we found very difficult to find variables useful for explaining wage dynamics. It is not possible to calculate tenure for all employees because the information is not precise for long-term stayers. As a consequence, experience can not be included in the x_{it} vector and we use worker's age as a proxy. All other observable skills and firm characteristics (education, occupation, sector, firm size, etc.), when available, unfortunately resulted never significant in the ECHP data, and it was more efficient to eliminate them from the model. We keep only the sex dummy. Business cycle factors are captured by time dummies.

Therefore the model that we estimate in all countries includes only a few variables in x_{it}: age, sex, and time dummies.

Results

From a technical point of view our model is a switching regime model with unobserved threshold, that is estimated with maximum likelihood methods. The specification of the likelihood estimated can be found in Appendix 2. The model converges nicely in all countries to a global maximum (different initial values have been tried), and the shape of the likelihood function is increasing and concave as expected.

The basic results are displayed in Table 3.3. First of all, as we can see the extent of measurement error is quite high in our survey data. Our estimate of the standard deviation σ_m ranges between 4 and 10 percent. This is anyway lower than standard errors obtained from validation studies for the

US, that are never below 10 percent. In Switzerland Fehr and Goette (2003) find a standard deviation of measurement error between 6 and 7 percent. AD could not estimate the significance of σ_m.

Thresholds in the European countries are: 0.1 for Austria; 0.14 for Portugal, Spain and Denmark; about 0.17 for Germany, Italy, Greece and Finland; and about 0.20 for Netherlands, Belgium, Luxembourg, UK, Ireland, and France.

However, comparing thresholds directly across countries is not correct because the underlying counterfactual distribution can be different across countries. For obtaining measures of nominal wage rigidity directly comparable we need to calculate country by country the percentage of sweep-ups and freezes implied by the model, and therefore corrected for measurement error. Given estimates of the model parameters we calculate, year by year for each country, the probability that $x_{it}\beta + e_{it} < -\alpha$ conditional on x_{it}. We then take the average of the probabilities over the sample members. Given estimates of β and α from the model, we calculate the probability that a worker with a given x takes a nominal wage cut and, hence, the proportion of workers that take wage cuts. Similarly, we use the model to estimate the proportion that have a nominal wage freeze in each year.

Tab. 3.4 compares the observed and estimated proportion of wage cuts and freezes in the ECHP. In all countries it is clear that most of the observed wage cuts are turned into wage freezes. Therefore measurement errors explain a very high proportion of the observed wage cuts. As a consequence, the estimated extent of nominal wage rigidity is very high across the European countries. According to the estimated proportion of cuts, we can rank in an increasing order of flexibility: France, Luxembourg, Netherlands, Belgium, Denmark, UK, and Finland among countries with quite rigid wages (below 10% of estimated cuts); Germany, Ireland, Portugal, and Italy are in between (below 15% estimated cuts); Spain and Austria present quite an high percentage of estimated cuts (between 16% and 22%). In Greece wages have become more an more flexible over the time period (from 6% to 18%), whereas in the other countries the percentage of cuts is quite stable over time.

If we consider the probability estimated of having a wage freeze, the most rigid countries are Belgium, France, Netherlands, and Germany (more than 40%), followed by Luxembourg, Denmark, UK, Italy, Finland and Ireland (between 39% and 30%), and then by Greece, Spain, Portugal and Austria (less than 30%).

Conclusions on estimates of wage rigidity

Estimates of a simplified version of the AD model in the European countries show quite high degrees of downward nominal wage rigidity. However, there is high variability across the European countries. With respect to the observed frequencies of wage cuts and freezes, the estimated ones exhibit lower percentages of cuts and higher freezes. Therefore in the observed data the extent of downward wage rigidity is underestimated and measurement errors explain almost all wage cuts observed.

7 The impact of institutions on wage rigidity

7.1 Theoretical framework and related literature

In the literature, two alternative explanations of the existence of downward nominal wage rigidity have been proposed. The most common explanation, advocated by Blinder and Choi (1990) and Akerlof at al. (1996), is that employers avoid nominal wage cuts because both they and the employees think that a wage cut is unfair. The other explanation, proposed by MacLeod and Malcomson (1993) in an individual bargaining framework and Holden (1994) in a collective agreement framework, is that nominal wages are given in contracts that can only be changed by mutual consent. As argued by Holden (1994), the two explanations are likely to be complementary.

Based on a theoretical framework allowing for bargaining over collective agreements as well as individual bargaining, Holden (2004) argues that workers who have their wage set via unions or collective agreements have stronger protection against a wage cut, thus the extent of downward nominal wage rigidity is likely to depend on the coverage of collective agreements and union density. For non-union workers, the strictness of employment protection legislation (*epl*) is key to their possibility of avoiding a nominal wage cut.

Groth and Johannsson (2001), consider a model with heterogeneus agents, wage setting by monopoly unions and monetary policy conducted by a central bank. They show that the duration of nominal wages is u-shaped in the level of centralisation, with intermediate bargaining systems yielding more flexible nominal wages than both decentralised and centralised systems.

Although there is now a fairly large and growing number of studies estimating the extent of wage rigidity in many countries, different methods and data make it in general difficult to compare the degree of downward nominal wage rigidity across countries. However, similar data and measures from a number of countries is needed in order to explore the institutional causes of wage rigidity using country-specific characteristics. The analysis carried out in the chapter 3 is useful for this purpose since it adopts the same method for estimating the extent of downward nominal wage rigidity in a number of countries for which data of similar nature are available. Accordingly, we find evidence of downward wage rigidity in all the EU countries.

Many economists think of nominal rigidities as related or caused by labour market institutions. As documented by OECD(1999), labour market institutions differ considerably among European countries, and it is therefore interesting to investigate the existence of DNWR for individual countries.

Summary of theoretical predictions to be tested

According to Holden (2004): EPL and union density have a significant negative effect on the incidence of nominal wage cuts and so has inflation, in a non-linear way. High unemployment reduces the incidence of wage cuts.

According to Groth and Johannsson (2001): it exists an hump-shaped relationship between wage cuts and level of centralisation of wage bargaining.

7.2 Data on institutions

Our empirical analysis is focussed on the following institutions characterising a national labour market: the body of employment protection laws; the degree of centralisation of collective bargaining; the proportion of employees covered by collective agreement; and finally the degree of consensus/coordination among bargaing parteners.

Labour economists, in an attempt to produce precise evaluations of the role of national labour market institutions in influencing macroeconomic performances, have constructed measures to provide a numerical description and corresponding rankings of countries for each of the above institutions.

Centralization describes the locus of the formal structure of wage bargaining. Typically three levels of bargaining are considered:1) centralized or national bargaining, which may cover the whole economy; 2) intermediate bargaining, where unions and employers' associations negotiations cover particular industries or crafts; and 3) decentralized or firm-level bargaining between unions and management. There are three alternative measures for centralization of wage bargaining. The CENTR variable taken from OECD (1997) is an OECD Secretariat estimate updating table 5.1 of OECD (1994), CENTRCD taken by Boeri, Brugiavini and Calmfors (2001) and CENTRLN by Nickell and Layard (1999). Since each yields a different ranking of European countries, we try them separately in our regressions.

The variable labelled COORD indicates the degree of coordination/consensus between the collective bargaining partners. COORD is an OECD Secretariat estimates constructed from combined information taken from Visser's (1990) classification of trade union coordination, the Calmfors and Driffil (1988) index and information gathered by the OECD on employers' associations.

As we can see from Tab. 4.6, for the percentage of employees covered by collective agreement the two sources considered (OECD (1997), and Cesifo Forum (2001)), give very similar measures, summarized in the variable labelled PCOV, which is for use in our regressions.

The strictness of employment protection legislation, captured by the variable EPL, is taken from OECD (1999). We do not consider union density, namely the percentage of employees belonging to the union in each country, as an esplicative variable, to keep an adequate level of model parsimony because there is widespread agreement that coverage matters much more than union density in determining wages.

In an attempt to identify a pure impact of institution measures, distinct from country specific time variant economic policies and macroeconomic effects, we have included into the model specification some important macroeconomic variables possibly capturing such effects. First, we consider the national unemployment rate as calculated by the OECD *Standardised unemployment rate*, URATEST. This is a variable containing data on the national unemployment rate adjusted to ensure comparability over time and across countries. We also consider the national inflation rate, as calculated by the percentage annual variation in the *Consumer price index* (OECD) and the OECD estimates of pecentage annual variation in *Labour productivity in the business sector*, that is total economy less the public sector.

Table .4.1 reports some descriptive statistics for the main variables used in the empirical analysis.

Table 4.1: Descriptive statitistics for the regression variables

Variable	Mean	Std. dev.	Min	Max
cmhgw (observed)	28.64	6.16	13.62	42.64
cmhgw (estimated)	10.79	6.77	1.30	38.45
centr	2.04	0.26	1.5	2.5
centrcd	2.97	1.46	1	6
centrln	9.06	3.50	5	17
coord	2.15	0.55	1	3
epl	2.34	1.04	0.5	3.7
pcov	0.79	0.16	0.40	0.99
uratest	8.33	3.64	2.30	18.80

Various sources indicated in text

As a preliminary analysis of the impact of institutions we work out rank correlations between either dependent variable and the institution measures. We employ Spearman's correlation, which is actually the Pearson's correlation between the ranks generated by the variables of interest. Results are reported below.

For observed wage cut frequencies we have the following coefficients with their significance level (as indicated by the probability value of the corresponding t statistics)

- CENTR: Spearman's rho = -0.0324, p-value for t = 0.8073;
- CENTRCD: Spearman's rho = -0.0172, p-value for t= 0.8916;
- CENTRLN: Spearman's rho = -0.1428, p-value for t= 0.2565;
- COORD: Spearman's rho = -0.0317, p-value for t= 0.8117;
- EPL: Spearman's rho = -0.2028, p-value for t= 0.0898;
- PCOV: Spearman's rho = 0.2310, p-value for t= 0.0642.

For the AD wage cut frequency we obtain

- CENTR: Spearman's rho = 0.3070, p-value for t = 0.0103;
- CENTRCD: Spearman's rho = 0.2871, p-value for t= 0.0119;
- CENTRLN: Spearman's rho = -0.0167, p-value for t= 0.8863;

- COORD: Spearman's rho = 0.4642, p-value for t= 0.0001;
- EPL: Spearman's rho =0.4419, p-value for t= 0.0000;
- PCOV: Spearman's rho = 0.2270, p-value for t= 0.0486.

Results are quite different btween the two dependent variables. While for observed frequencies there is a weakly significant negative correlation with EPL and positive with PCOV, for AD frequencies all correlation terms except CENTRLN are significant, in addition the relationship with EPL switches sign. However, at this simple level of analysis it is impossible to shed light about the distinctive impact of institutions. The multivariate regression analysis below is more promising in this respect.

7.3 Regression results

The OECD (1997) suggests two approaches to the empirical analysis of institutions, one based on non-linear specifications, treating institution measures as cardinal, and the other based upon dummy variables comprising the effect of subsets of countries with common measured institutional characteristics. The former is simpler, but has the drawback of maintaining cardinality for institution variables, which is clearly restrictive. The latter avoids this problem treating institution measures as purely ordinal, but still brings about the problem that country grouping is necessarily arbitrary and, moreover, it is not able to shed light on the non linear effects. We therefore prefer the former approach, which has the merit to detect hump-shaped or u-shaped effects very easily.

We consider a linear projection of the frequencies of gross hourly wage cuts (CMHGW) on both linear and squared terms of the institutional measures: COORD; PCOV; CENTR, (CD), (LN); and COORD. This specification is general enough to capture simultaneous non-linear effects, such as u-shaped and hum-shape correlations with the dependent variable. For each centralisation measure we consider a different model, thus our baseline model is the following

$$E^{*}\left(y|x\right)=\alpha_{o}+\sum_{i\in I}\left(\beta_{i}x_{i}+\beta_{i,i}x_{i}^{2}\right) \qquad (1)$$

where y is CMGHW and I_C={C, COORD, EPL, PCOV} and C=CENTR, CENTRCD, CENTRLN.

In the empirical application model (1) is supplemented with time dummies and URATEST; CPI; and LABPROD in an attempt to capture macroeconomic shocks, policy effects and growth.

Only in the presence of zero correlation between the random part of y and the explanatory variables, will OLS provide best linear unbiased estimates for the β's. There are three sources of randomness to be concerned with when modelling wage cuts using observed or estimated frequencies. The first, common to both data-sets used, is caused by the occurence of idiosyncratic aggregate shocks which may affect the wage change distribution as a whole in a given region. It may be partly controlled the inclusion of time dummies. The second arises at the micro level and is given by individuals misreporting and rounding their earnings. The third type of randomness, referenced to as measurement error, stems from the lack of information about the structure of earnings in surveys, and often also in administrative data. In our case it is of particular concern since it is usually difficult to isolate the contracted hourly wages from total earnings. We have attempted to account to some extent both these last types of errors in the econometric implementation of the Altonji and Devereoux model. The inclusion of the macroeconomic variables may be of concern for their likely correlation with the first source of randomness, which will be dealt with by using an Instrumental variable (2SLSL) estimator instrumenting the macroeconomic variables by their lags up to the fifth. For all specifications the usual tests of instrument validity (Sargan test of overidentifying restrictions and F tests on the joint significance of instruments in the first stage regression) support our choice of instruments. For all specifications, estimation methods implemented and tests see Appendix 3, Tables ???? report results solely for the OLS regression with the macro variables.

We can single out the following set of results common to all specification tried. First, there emerges a significant non-linear impact of EPL on nominal wage flexibility. Such effect always comes under the form of a "hump-shaped" relationship between EPL and hourly wage cut frequencies, however estimated. Second, we find a significant "u-shaped" impact for PCOV with the decreasing portion of the curve predominating over the increasing. Third, we find an "hump-shaped" impact for COORD with the increasing portion of the curve predominating over the decreasing. These results are robust to a) the choice of the wage cut frequency variable; b) the choice of the centralisation variable; c) the inclusion of time dummies and all macroeconomic controls; and finally d) the treatment of the possible endogeneity of the macro variables within an instrumental variable estimation framework. For the centralization variables, instead, we are unable to report significant results using the observed frequencies. With the AD estimated frequencies the impact of centralisation is although significant, not robust to the different choices of the centralisation variable.

This results confirm, with some required qualifications, the theoretical prediction by Holden. In particular the predicted negative effect of EPL begins to bite from a point of intermadiate strictness. For PCOV, instead, the predicted negative impact holds since the beginning over a large portion of the sample.

Tab. 4.2 Observed frequencies, CENTR

Variable	Coefficient	(Std. Err.)
coord	87.877	(68.338)
coord2	-17.393	(13.940)
centr	-23.311	(145.536)
centr2	5.260	(32.579)
epl	124.127**	(40.305)
epl2	-25.351**	(7.867)
pcov	-818.649†	(457.986)
pcov2	429.517	(258.489)
uratest	3.422**	(0.767)
uratest2	-0.130**	(0.041)
cpi	0.665	(0.790)
labprod	1.103†	(0.596)
Intercept	165.788	(110.673)

N	57
R^2	0.772
$F_{(17,39)}$	33.788

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Tab. 4.3 Observed frequencies, CENTRCD

Variable	Coefficient	(Std. Err.)
coord	88.905	(67.269)
coord2	-17.651	(13.939)
centrcd	1.116	(6.866)
centrcd2	-0.185	(1.188)
epl	123.989**	(34.723)
epl2	-25.305**	(6.810)
pcov	-858.605*	(399.512)
pcov2	455.445†	(231.005)
uratest	3.421**	(0.782)
uratest2	-0.129**	(0.041)
cpi	0.668	(0.771)
labprod	1.111†	(0.595)
Intercept	152.820**	(56.183)

N	57
R^2	0.772
$F_{(17,39)}$	33.925

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Tab. 4.4 AD frequencies, CENTR

Variable	Coefficient	(Std. Err.)
coord	479.057**	(77.912)
coord2	-97.418**	(16.076)
centr	471.561**	(145.601)
centr2	-102.579**	(33.034)
epl	271.295**	(40.082)
epl2	-52.415**	(7.791)
pcov	-3478.288**	(494.262)
pcov2	1963.820**	(283.643)
uratest	-0.734	(0.569)
uratest2	0.001	(0.027)
cpi	-1.403	(0.952)
labprod	-0.666	(0.588)
Intercept	109.530	(110.992)
N	69	
R^2	0.637	
$F_{(18,50)}$	11.683	

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Tab. 4.5 AD frequencies, CENTRCD

Variable	**Coefficient**	**(Std. Err.)**
coord	313.375**	(94.450)
coord2	-62.768**	(19.573)
centrcd	-25.804**	(6.264)
centrcd2	4.709**	(1.040)
epl	196.625**	(41.293)
epl2	-38.244**	(8.051)
pcov	-1755.594**	(513.447)
pcov2	929.869**	(296.594)
uratest	-0.254	(0.601)
uratest2	-0.010	(0.029)
cpi	-1.220	(0.961)
labprod	-0.663	(0.589)
Intercept	245.539**	(69.977)
N	69	
R^2	0.628	
$F_{(18,50)}$	10.44	

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

The qualititative evidence from the regression models is clear-cut, suggesting a significant hump-shaped relationship between cut frequencies on one hand and employment protection legislation and coordination on the other; and a u-shaped relationship between cut frequencies and coverage. Nonetheless, direct inspection of coefficients does not help to draw as much precise quantitative conclusions, since for the institutional variables no clear unit of measure is available. Given the nature of the institutional measures we compute a discrete partial effect for each variable. More specifically, we work out the variation in the linear projections of cut frequencies caused by one standard deviation from the sample mean in the institutional measure of interest.

The population partial effect for one standard deviation increase is given by

$$\begin{aligned} PE_i^{+} &= E^{*}\left(y|\mu_i+\sigma_i,\cdot\right)-E^{*}\left(y|\mu_i,\cdot\right)=\left(\beta_i+\beta_{i,i}2\mu_i\right)\sigma_i+\beta_{i,i}\sigma_i^2, \\ i &= centr, epl, coord \text{ and } pcov, \end{aligned}$$

whereas for one standard deviation decrease we have

$$\begin{aligned} PE_i^{-} &= E^{*}\left(y|\mu_i-\sigma_i,\cdot\right)-E^{*}\left(y|\mu_i,\cdot\right)=-\left(\beta_i+\beta_{i,i}2\mu_i\right)\sigma_i+\beta_{i,i}\sigma_i^2, \\ i &= centr, epl, coord \text{ and } pcov. \end{aligned}$$

These are both estimated by their consistent sample analogs

$$\begin{aligned} \widehat{PE}_i^{+} &= \left(\widehat{\beta}_i+\widehat{\beta}_{i,i}2\widehat{\mu}_i\right)\widehat{\sigma}_i+\widehat{\beta}_{i,i}\widehat{\sigma}_i^2, \\ \widehat{PE}_i^{-} &= -\left(\widehat{\beta}_i+\widehat{\beta}_{i,i}2\widehat{\mu}_i\right)\widehat{\sigma}_i+\widehat{\beta}_{i,i}\widehat{\sigma}_i^2, \\ i &= centr, epl, coord \text{ and } pcov. \end{aligned}$$

The presence of the variance in the partial effects formula is due to the fact that we are considering a discrete variation equal to the standard deviation, which is not necessarily small.

We report results only from the models for the "observed" frequencies of wage cuts including all macro variables and time dummies and using the *centrcd* measure of centralisation. Results on estimates and t statistics for $\widehat{PE}_i^{+}$ and $\widehat{PE}_i^{-}$ are reported in Table 4.6 For the sake of simplicity, they are computed by supposing the institution variables as fixed across repeated samplings, so that $\widehat{\mu}_i$ and $\widehat{\sigma}_i$ are held fixed too. Although this does not seem implausible given the particular nature of the variables considered, it is nonetheless one potential source of randomness that is neglected, and which may lead to underestimating the relevant standard errors. For ease of interpretation we also report the estimated extreme points of the curve $\widehat{x}_i^* = -\widehat{\beta}_i/2\widehat{\beta}_{i,i}$ and locate each of them in comparison with $\widehat{\mu}_i$. This is useful to understanding whether the local averaged partial effect $\beta_i + \beta_{i,i}2\mu_i$ is positive or negative, and "bell" or "u" shape of the estimated curve is actually relevant given the observed cross-national heterogeneity in institution measures. For example, in the presence of a "bell" shaped curve, if the maximum lays to right of the mean point by more than one standard deviation, then an increasing monotonic curve would actually prevail over a larger region of the sample. The opposite would happen with a u-shaped curve.

Below we summarise results without making reference to the type of regression model, IV or OLS, since they are very similar. One standard deviation increase or decrease from the average value of *epl* brings about in either case a strong reduction in expected wage cut frequencies, by around 20 and 30 percentage points respectively. For *coord* instead, the increasing part of the relationship turns out predominant, so that one standard deviation increase leads to only small rises in expected cut frequencies, no higher than 2 points; on the other hand one standard deviation less of *coord* implies a reduction by around 12 points. For *cov* we observe a reduction of around 10 points in expected cut frequencies when it increases by one standard deviation and a stronger increase, by more than 30 points, when it decreases. This evidence is robust to the several other specifications, estimation methods and variables tried.

Table 4.6 Partial effects, CENTRCD, all macro vars. and time dummies

Variable	$\widehat{PE}_i^+$ $\widehat{PE}_i^-$	t	$\widehat{x}_i^*$	$\widehat{\mu}_i$	$\widehat{\sigma}_i$
coord	1.76	1.69	2.52	2.15	0.55
	-12.27	-1.51			
centrcd	-0.36	-0.12	3.02	2.97	1.46
	-0.42	-0.17			
epl	-21.75	-4.81	2.45	2.34	1.04
	-32.78	-3.20			
pcov	-10.92	-2.32	0.94	0.79	0.16
	33.10	2.77			
uratest	2.89	2.62	13.22	8.33	3.64
	-6.32	-5.06			

Table 4.7 Partial effects, CENTRCD, IV regression with all macro vars. end. and time dummies

Variable	$\widehat{PE}_i^+$ $\widehat{PE}_i^-$	z	$\widehat{x}_i^*$	$\widehat{\mu}_i$	$\widehat{\sigma}_i$
coord	1.72	1.99	2.51	2.15	0.55
	-13.73	-1.90			
centrcd	-0.74	-0.28	2.99	2.97	1.46
	-0.78	-0.35			
epl	-22.05	-5.31	2.46	2.34	1.04
	-34.16	-3.64			
pcov	-10.71	-2.79	0.93	0.79	0.16
	34.97	3.22			
uratest	2.87	2.99	13.27	8.33	3.64
	-6.21	-5.63			

8 Conclusions

In this paper we have analysed wage dynamics at the individual level using the 1994-2001 data from the ECHP survey, with particular emphasis on constructing wage rigidity measures for inter-country comparisons. First of all, a simple descriptive analysis of wage change distributions detected the existence of nominal wage rigidity in Europe, through the presence of spikes at zero nominal wage changes and asymmetry of the distributions around zero in all the countries. However, wages were found to be not completely downwardly rigid, since the percentage of observed cuts was relevant in Europe. No particular evidence was found for menu costs, whereas some evidence of real wage rigidity was detected in some countries.

However, the existence of measurement error in the two forms of rounding and reporting errors was documented in the data, therefore a proper estimation procedure, based on a simplified version of the AD model, allowed us: 1) to take into account measurement errors; and 2) to construct measures of wage rigidity comparable across countries. Our first result was that in all the European countries measurement error modelled according the classical assumptions explains a relevant proportion of the observed wage cuts, that are nominal wage freezes instead. Therefore the estimated extent of nominal wage rigidity is higher than the observed one in all the EU countries. This result is in line with previous findings from estimations of similar models in other countries.

At the same time, the use of the ECHP data allows us to construct two measures of nominal wage rigidity (the percentage of cuts and freezes) comparable across countries. If these frequencies are the observed ones, inter-country comparisons can be carried out only under very restrictive assumptions on the counterfactual distributions and measurement errors. If we introduce simplifying assumptions on measurement errors, we can estimate for each year and each country the probability of cuts and freezes, conditional on some individual variables observed, that can be directly compared across countries. We find that the percentage of observed cuts is between 13% and 38%, whereas observed freezes are between 1% and 24%. On the contrary estimated cuts vary between 4% and 22%, and the estimated freezes are between 20% and 44%.

We then investigated the importance of institutional characteristics of the labour market in explaining the extent of nominal wage rigidity. Our regression results are the following. First, there emerges a significant "hump-shaped" impact of the employment protection legislation variable (*epl*) on nominal wage flexibility. Second, we find a significant "u-shaped" impact for the coverage variable (*pcov*), with the decreasing portion of the curve predominating over the increasing. Third, we find an "hump-shaped" impact for coordination (*coord*), with the increasing portion of the curve predominating over the decreasing. These results are robust to a) the choice of the wage cut frequency variable; b) the choice of the centralisation variable; c) the inclusion of time dummies and all macroeconomic controls; and finally d) the treatment of the possible endogeneity of the macro variables within an instrumental variable estimation framework. For the centralization variables, instead, we are unable to report robust and significant results.

This results partly confirm the theoretical predictions by Holden (2004).

Although the analysis carried out so far in this paper was focused on measuring nominal wage rigidity in the EU countries, further investigation is needed to explore the consequences of nominal wage rigidity in Europe.

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Tab. 3.1 Information used for defining stayers full-time by country in the ECHP

Country	Waves available	Monthly status=employed	No change in sector	No change in occupation
Germany GSOEP	1-7	*	not available	*
Germany	1-3	*	*	*
Denmark	1-7	*	many missing in wave6	many missing in wave 4
Netherlands	1-7	missing	*	*
Belgium	1-7	*	many missing in wave 6 and 7	many missing in wave 6 and 7
Luxemburg PSELL[1]	2-7	*	many missing in waves 1-5	many missing in waves 1-5
Luxembourg	1-3	*	*	*
France	1-7	*	*	*
UK BHPS	1-7	*	*	*
UK	1-3	*	*	*
Ireland	1-7	*	many missing in waves 1-7	many missing in waves 1-7
Italy	1-7	*	*	*
Greece	1-7	*	*	*
Spain	1-7	*	*	*
Portugal	1-7	*	*	*
Austria	2-7	*	*	*
Finland	3-7	*	*	*
Sweden[1]	4-7	missing	many missing	missing

1. Only Net wage available; Sweden excluded.

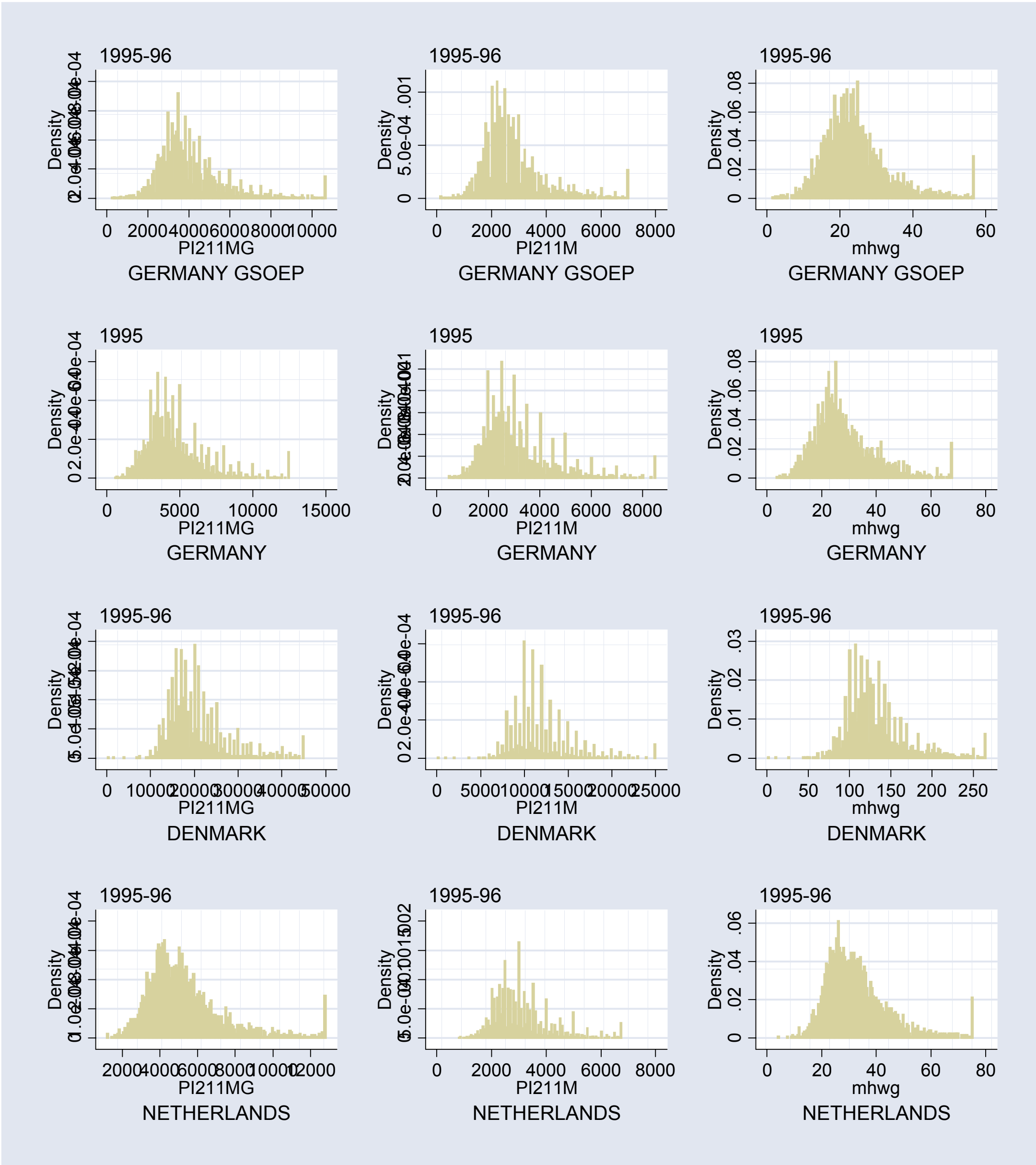

Fig 3.1 Gross, Net and Gross Hourly Wage distributions in the ECHP for Stayers full-time.

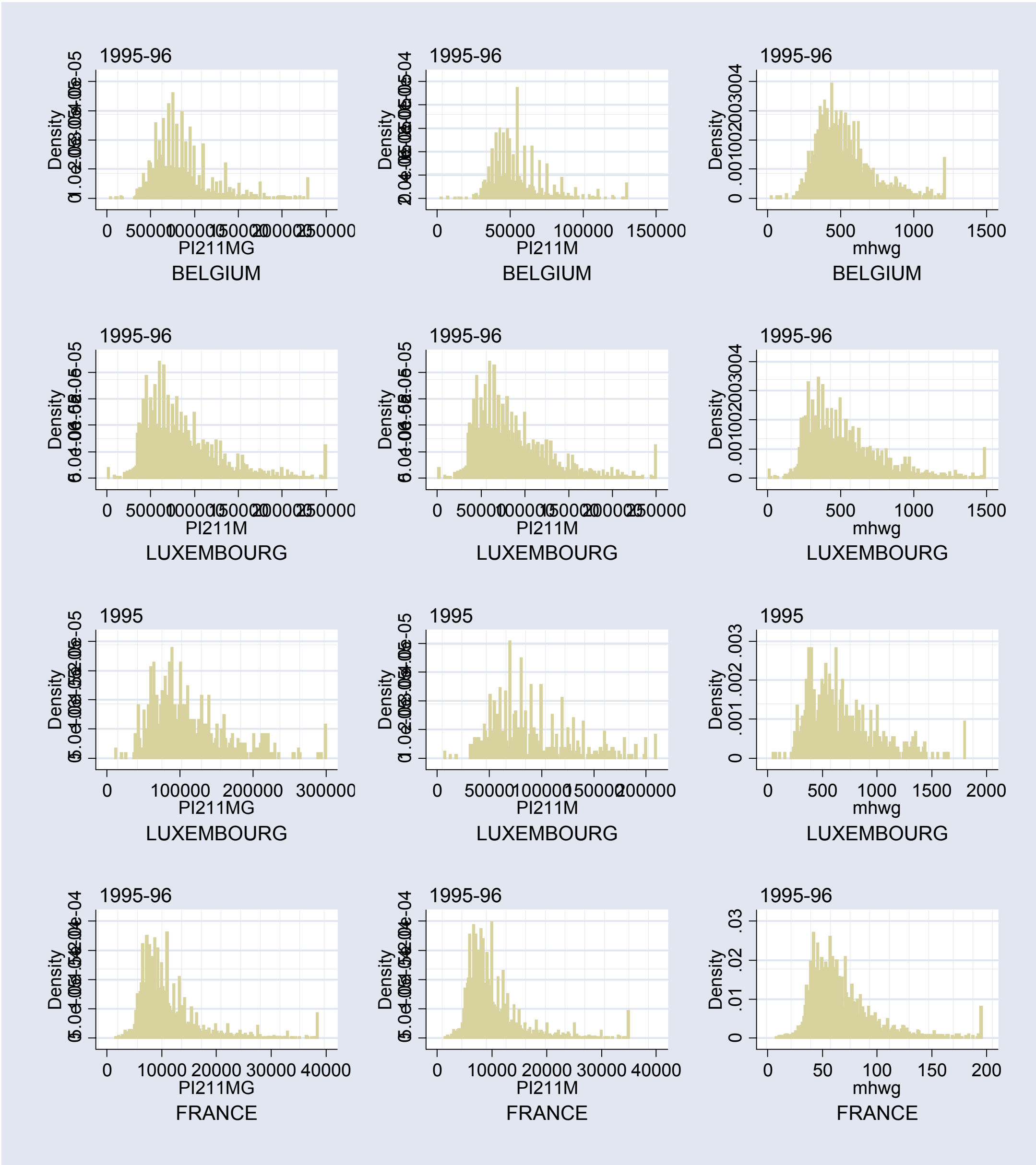

Fig. 3.1 continued

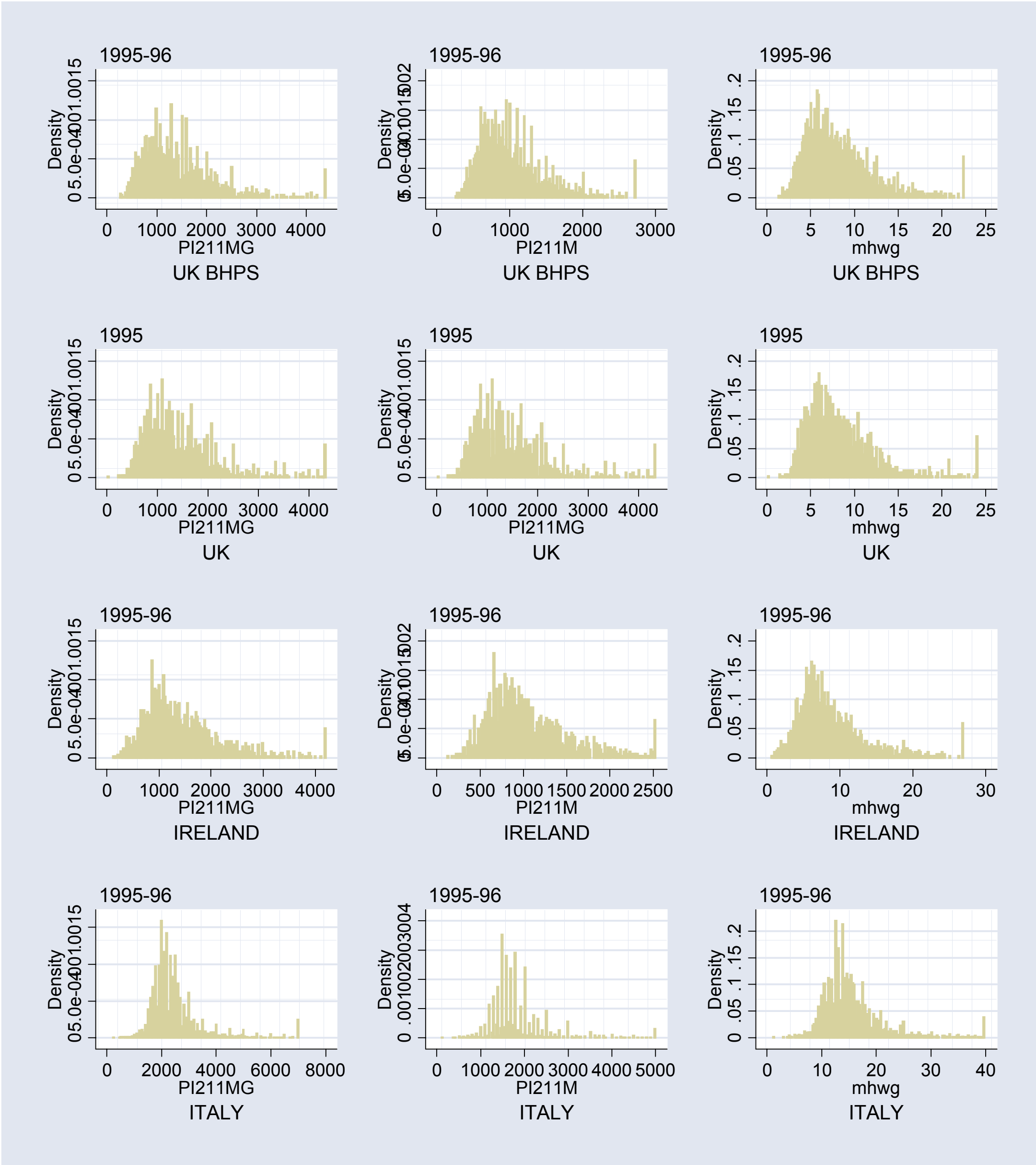

Fig. 3.1 continued

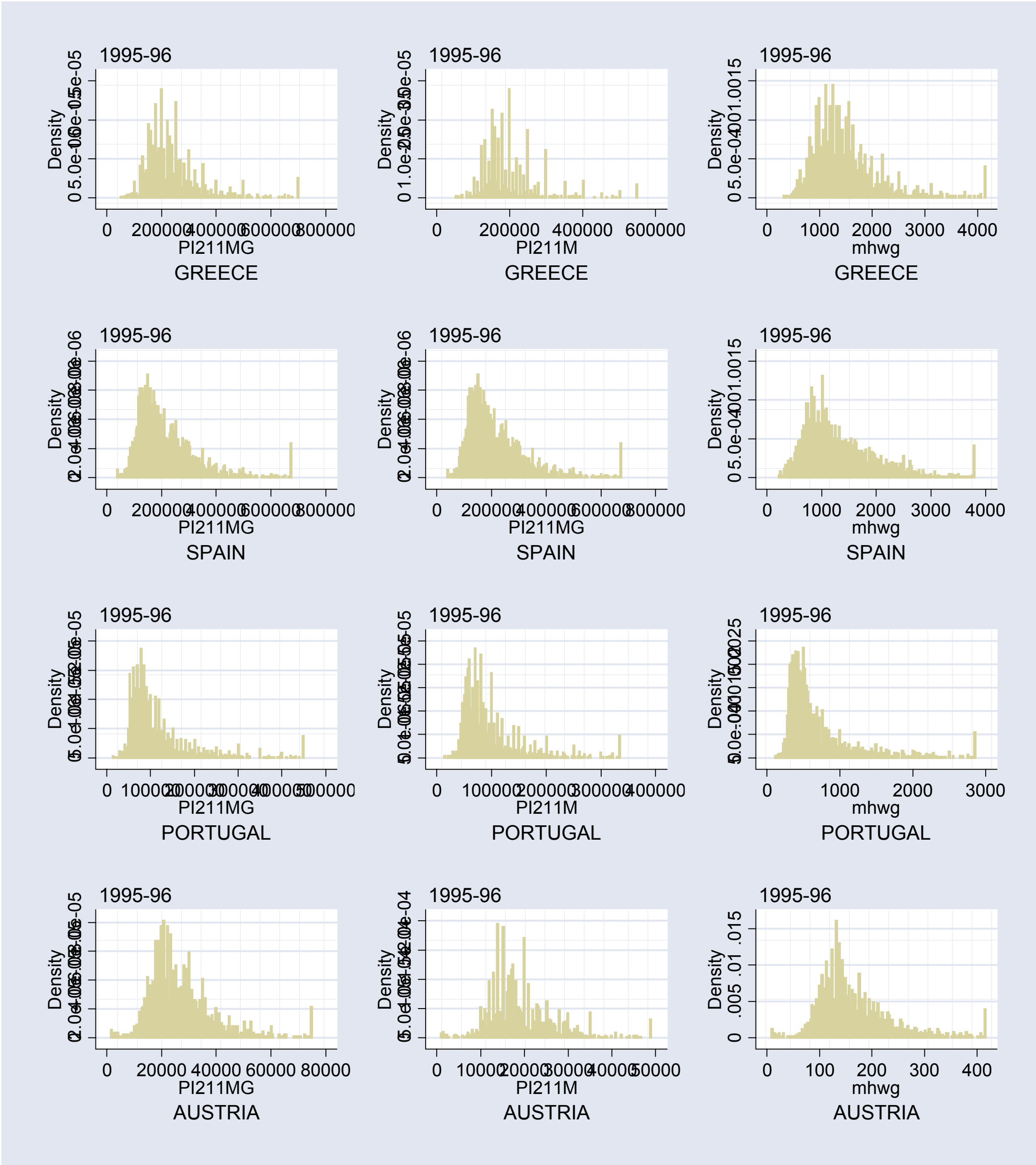

Fig. 3.1 continued.

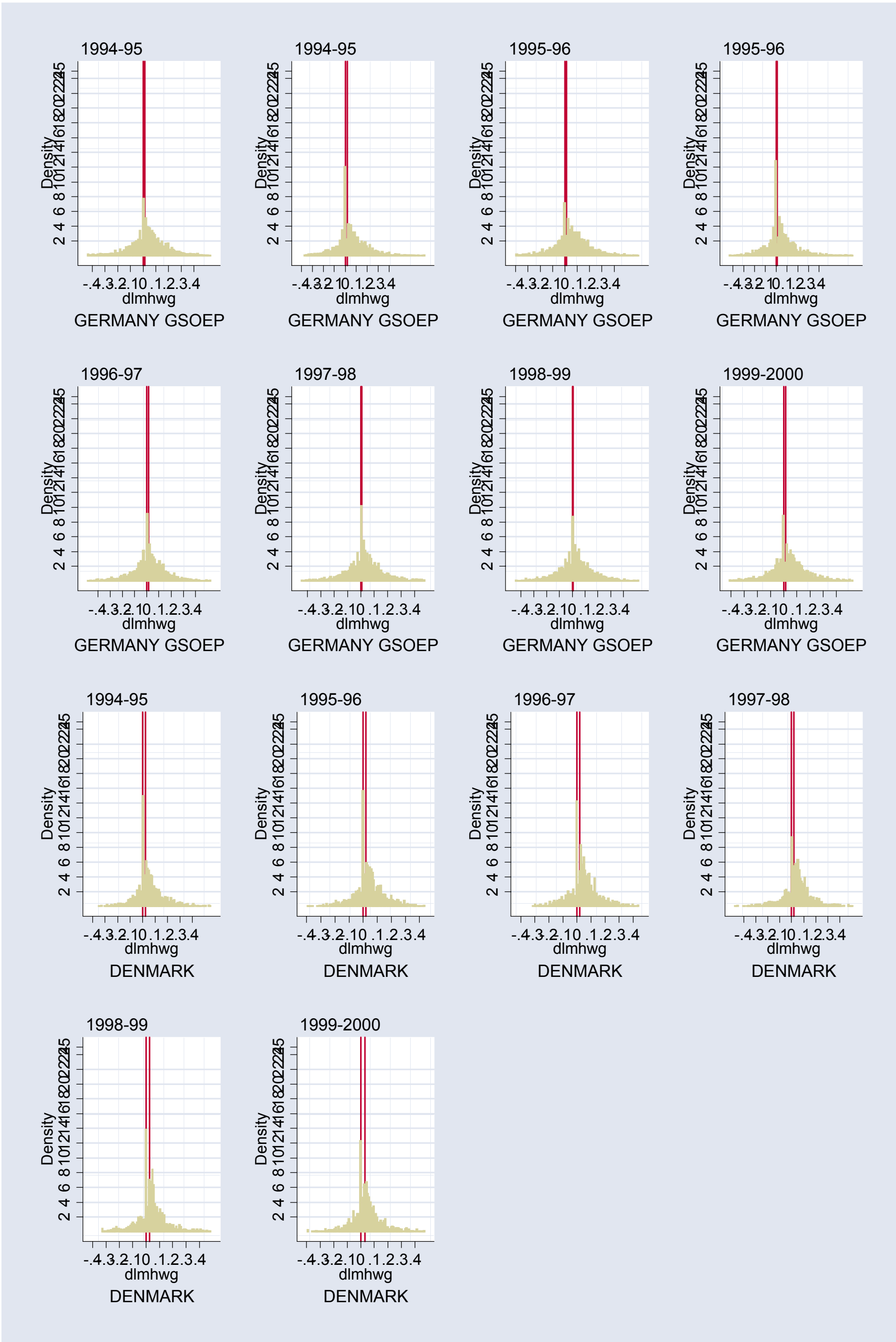

Fig. 3.2: Gross Hourly Wage Changes Distributions in the ECHP.

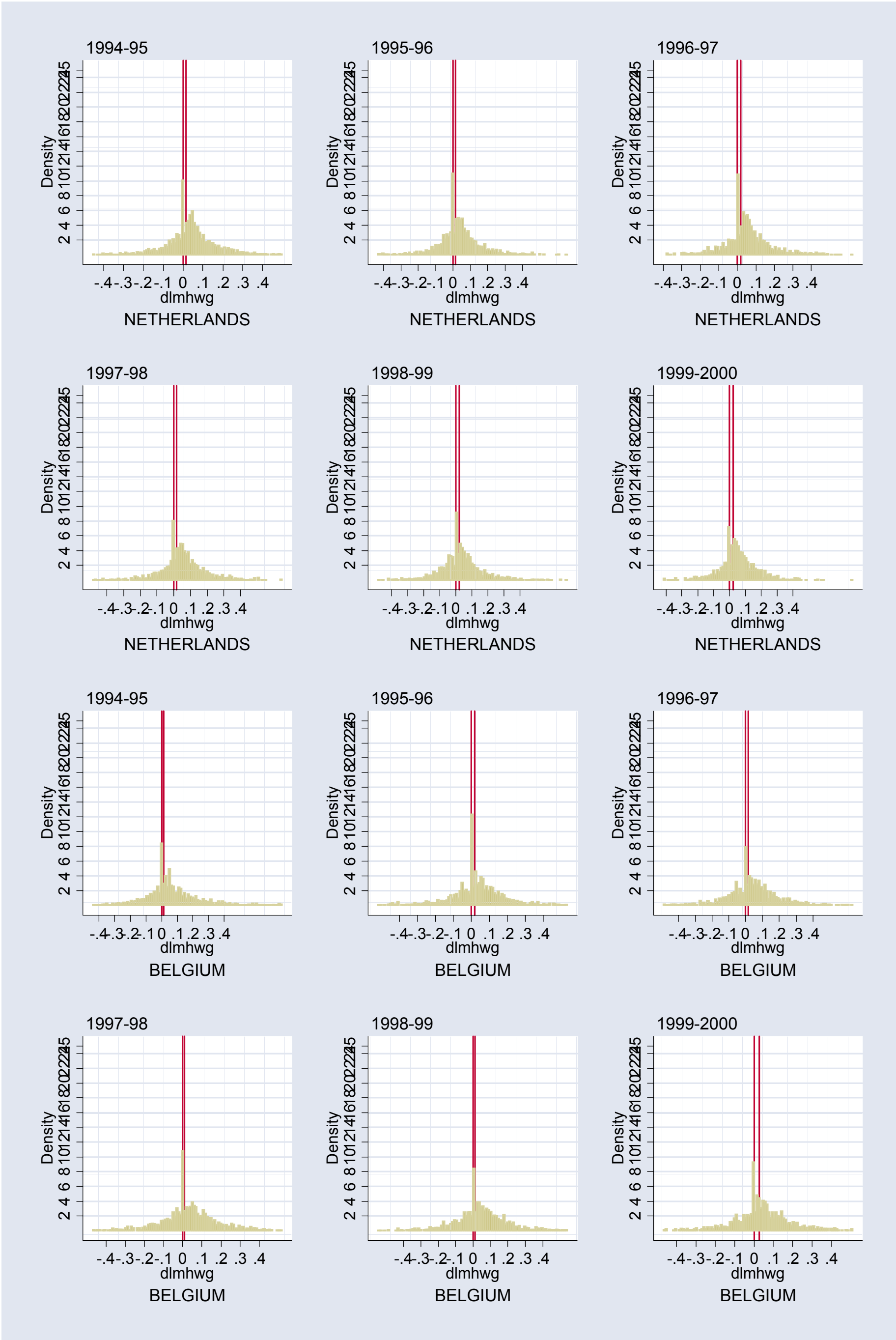

Fig. 3.2 continued.

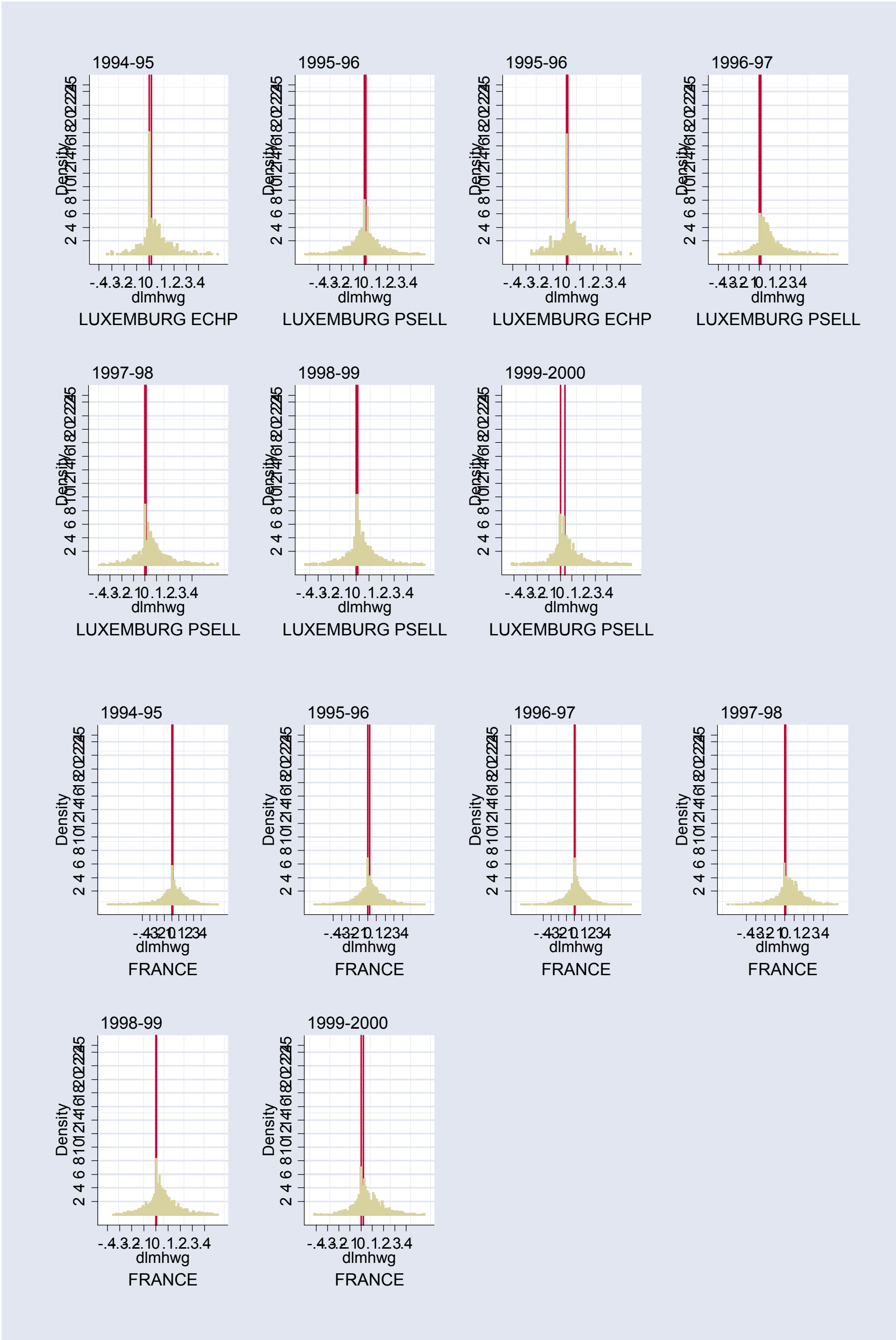

Fig. 3.2 continued.

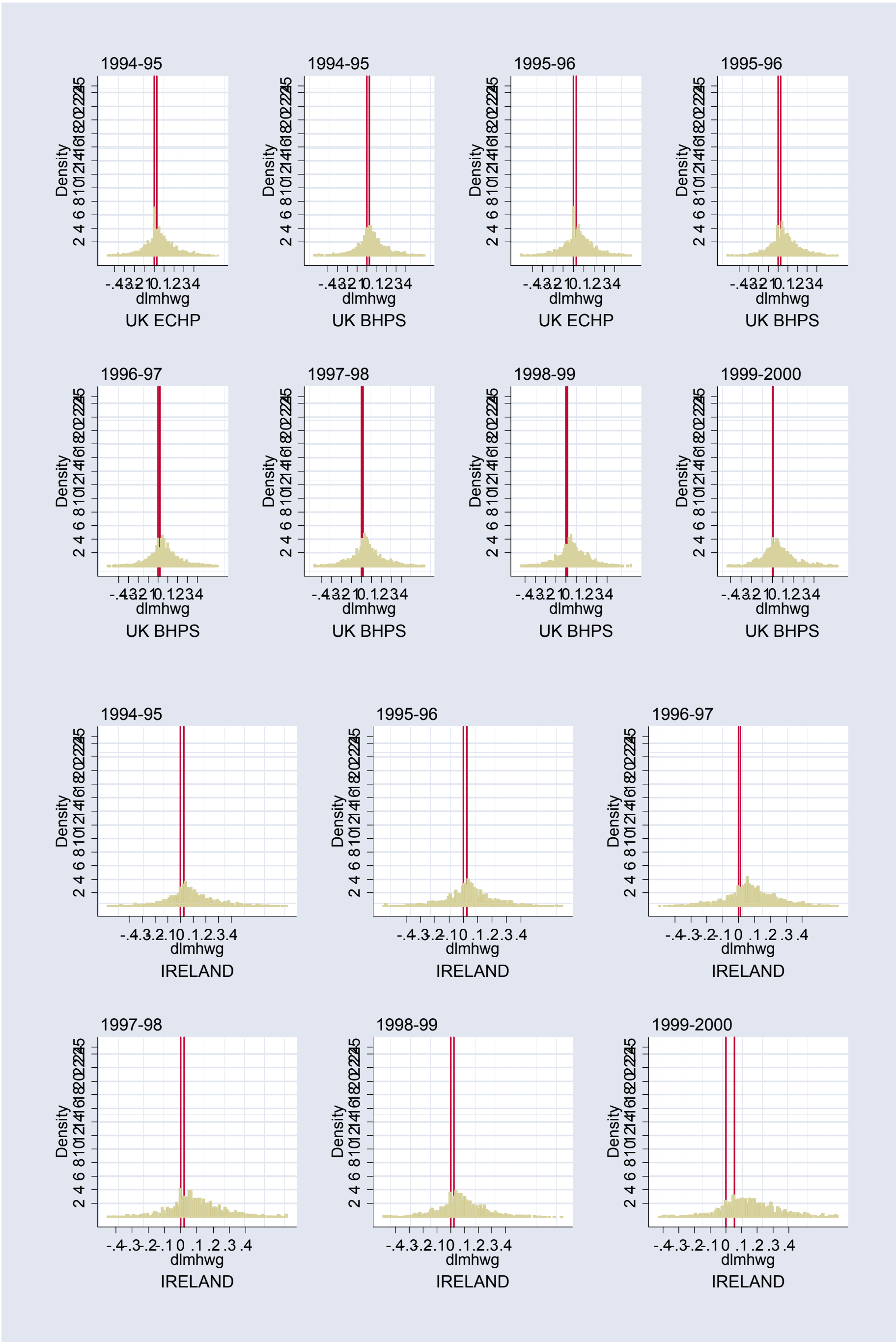

Fig. 3.2 continued.

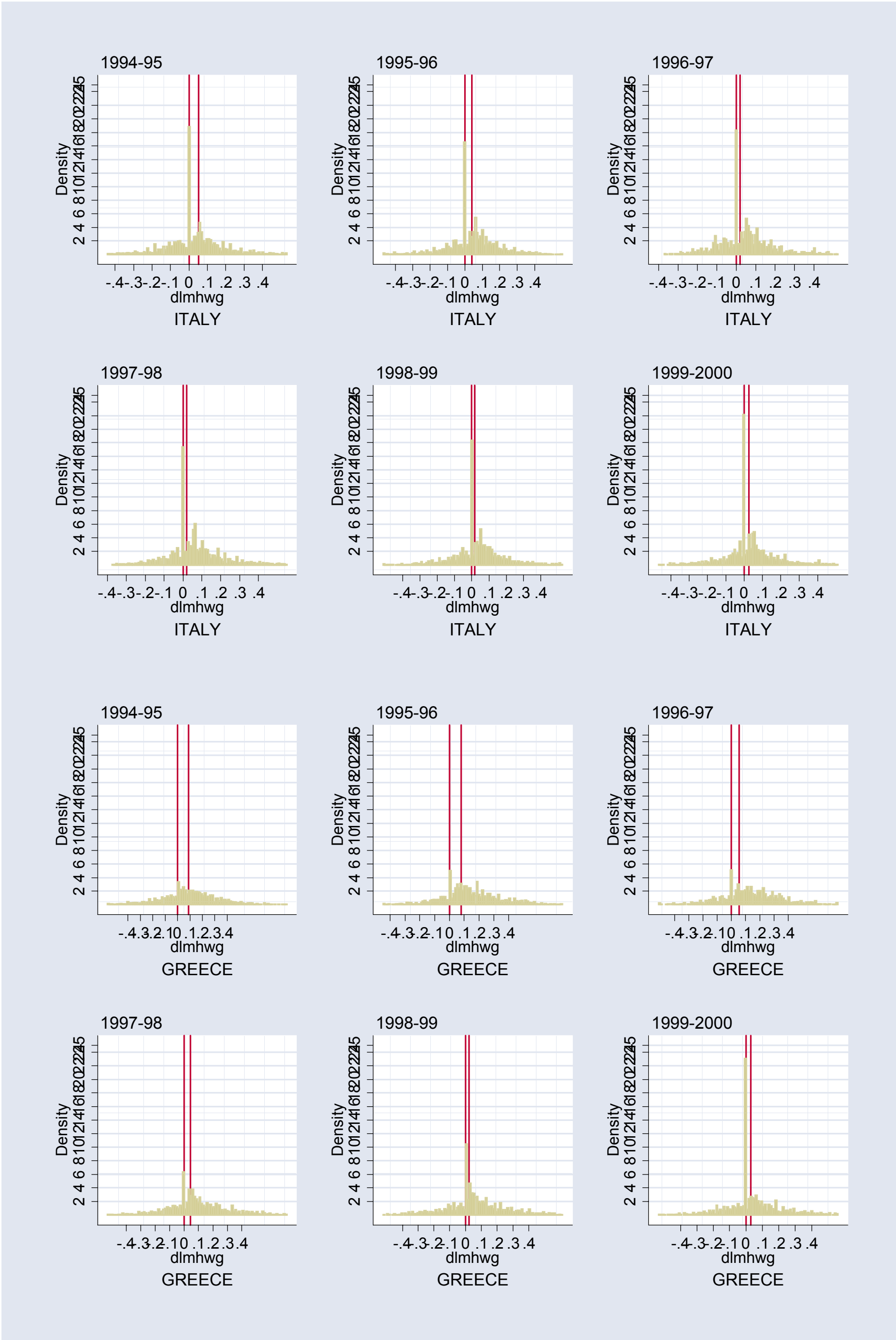

Fig. 3.2 continued.

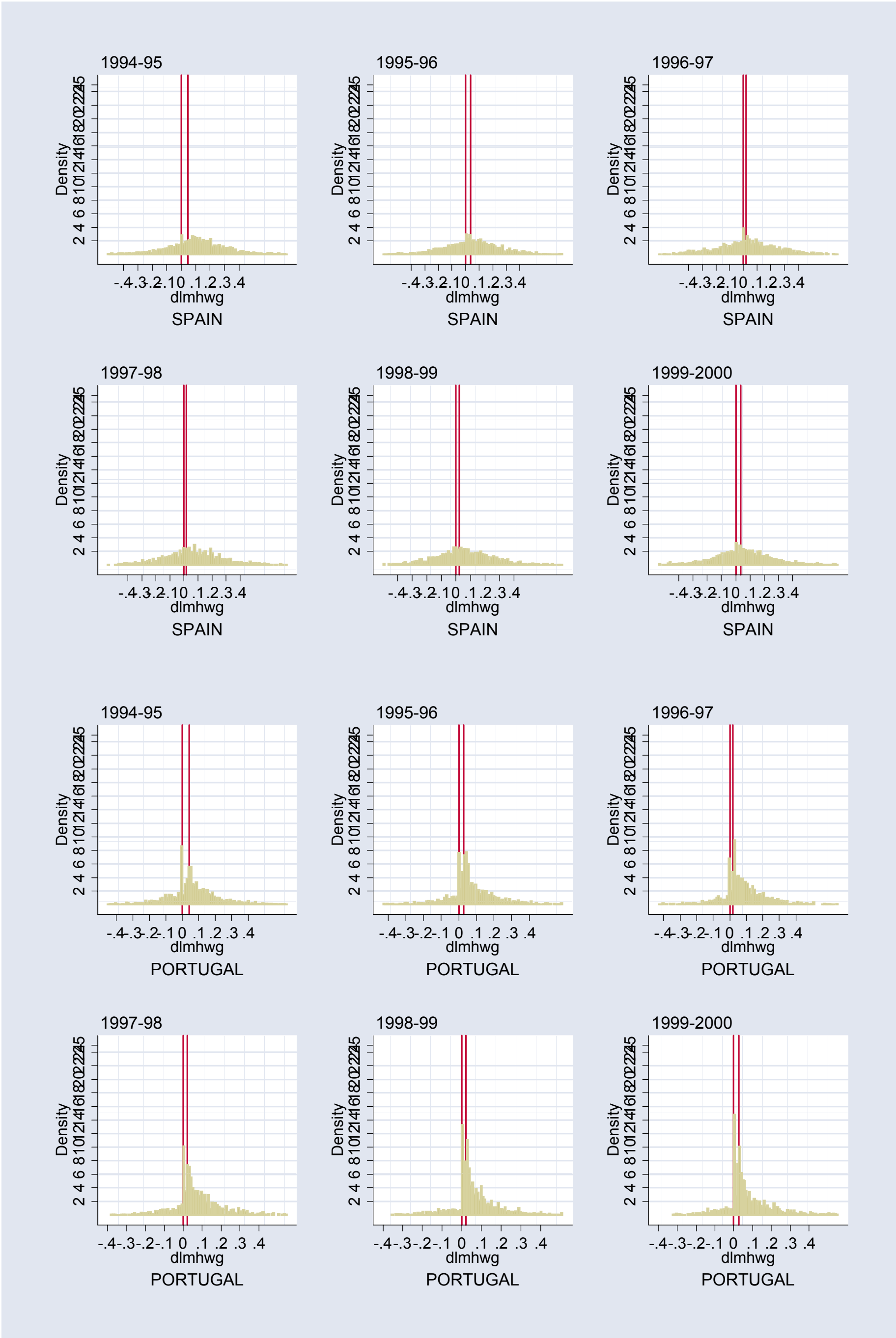

Fig. 3.2 continued.

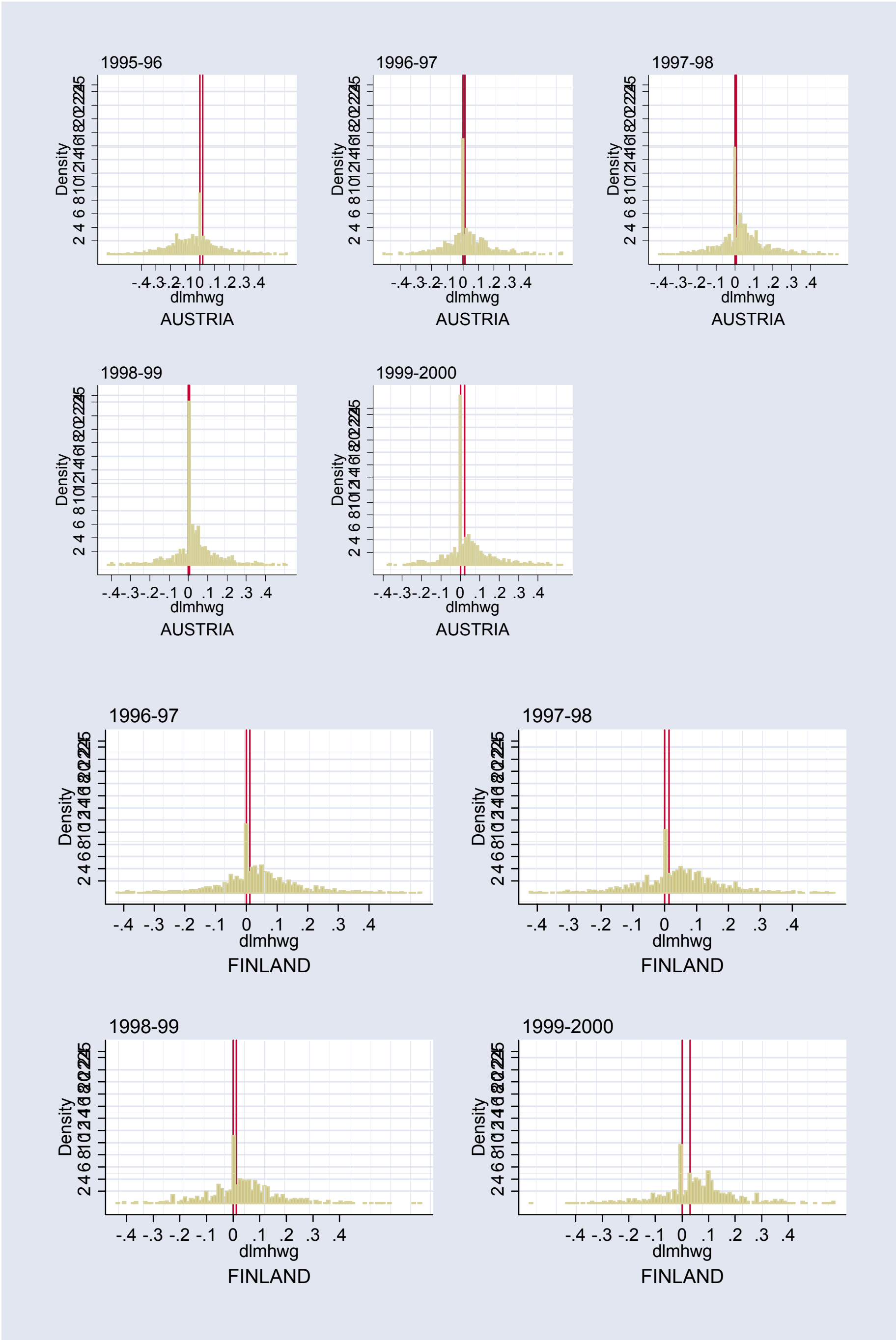

Fig. 3.2 continued.

Tab. 3.2: Percentage of wage cuts, freezes and rises observed in the ECHP

		Monthly Wages			Hours			Hourly Wages		
Country	wave	dmw<0	dmw=0	dmw>0	dh<0	dh=0	dh>0	dhw<0	dhw=0	dhw>0
GERMANY gsoep	2	22.23	13.47	64.29	28.48	37.45	34.07	32.39	5.41	62.2
	3	21.34	13.53	65.14	36.76	34.12	29.12	28.04	5.38	66.57
	4	27.27	16.34	56.39	27.66	37.89	34.45	38.33	6.86	54.81
	5	28.19	17.98	53.83	33.38	37.39	29.23	36.53	7.18	56.29
	6	27.43	17.41	55.16	31.42	36.95	31.63	36.92	6.74	56.34
	7	23.19	16.29	60.52	31.31	36.56	32.12	33.11	6.46	60.43
average		24.7879	15.7364	59.0596	31.3586	36.7051	31.7012	34.0356	6.29825	59.3023
DENMARK	2	21.28	19.02	59.7	27.82	60.95	11.23	22.4	11.5	66.1
	3	17.86	18.04	64.09	16.2	68.31	15.49	23.46	12.29	64.25
	4	16.01	14.54	69.45	17.61	67.01	15.38	21.75	9.63	68.62
	5	17.52	11.04	71.44	14.69	67.62	17.69	24.74	7.84	67.42
	6	18.8	13.63	67.58	15.33	69.42	15.26	23.48	9.88	66.64
	7	18.77	13.83	67.41	15.48	68.53	15.99	25.3	9.75	64.95
average		18.3049	14.7701	66.4998	17.3911	66.9124	15.0344	23.4896	10.046	66.3139
NETHERLANDS	2	21.3	13.59	65.11	21.17	57.39	21.44	28.7	8.13	63.18
	3	23.62	16.11	60.26	22.58	58.37	19.05	29.39	9.81	60.8
	4	19.43	13.96	66.61	25.85	56	18.15	23.91	8.56	67.53
	5	20.19	12.94	66.87	26	54.17	19.83	25.48	7.14	67.37
	6	29.62	14.51	55.86	23.62	56.42	19.95	34.28	8.07	57.65
	7	19.26	11.93	68.81	21.29	57.5	21.21	24.17	6.97	68.87
average		21.9745	13.7797	63.7578	23.3376	56.6255	19.9052	27.4281	8.06009	64.1031
BELGIUM	2	21.97	18.9	59.13	28.87	44.06	27.07	29.74	8.35	61.91
	3	23.19	24.69	52.12	30.11	42.52	27.37	32.06	10.94	57
	4	22.8	15.86	61.34	28.77	42.65	28.58	32.96	7.25	59.79
	5	24.21	18.8	56.99	27.57	43.8	28.63	34.37	9.04	56.6
	6	22.7	17.53	59.78	30.82	39.68	29.5	32.54	7.33	60.13
	7	22.74	17.09	60.17	29.98	41.85	28.17	31.5	7.39	61.12
average		22.9252	18.6206	58.1721	29.3337	42.4017	28.2082	32.1637	8.28883	59.3917
LUXEMBURG psell	3	43.51	6.22	50.26	5.92	87.86	6.21	44.5	5.06	50.44
	4	15.03	6.7	78.27	6.16	86.6	7.24	18.18	5.68	76.15
	5	25.47	6.88	67.66	10.12	86.56	3.32	24.42	6.18	69.41
	6	30.39	5.99	63.62	4.24	91.08	4.67	31.44	5.37	63.19
	7	28.74	4.78	66.48	4.71	91.79	3.5	28.55	4.4	67.05
average		27.0744	6.06543	64.6081	5.936	88.7501	4.75851	28.1681	5.30375	64.6521
FRANCE	2	80.51	2.07	17.42	24.67	54.73	20.6	76.07	0.93	23
	3	33.62	8.86	57.52	21.81	56.54	21.65	37.6	5.56	56.84
	4	12.62	1.95	85.44	21.68	54.83	23.49	17.89	1.14	80.97
	5	23.8	6.83	69.37	24.46	55.45	20.09	27.79	3.95	68.26
	6	27.15	7.76	65.09	25.4	56.83	17.77	30.23	4.64	65.14
	7	27.66	8.17	64.17	34.62	49.13	16.25	27.27	4.35	68.38
average		29.1262	4.99257	54.0038	25.1141	54.5219	19.8284	32.4716	2.78833	56.3994
UK bhps	2	25.94	6.61	67.45	31.57	34.57	33.87	33.7	2.95	63.35
	3	26.3	5.67	68.03	32.74	33.56	33.7	33.41	2.07	64.52
	4	24.18	4.5	71.32	33.18	34.79	32.03	32.91	1.81	65.28
	5	25.58	5.68	68.74	34.85	32.01	33.14	31.46	2.05	66.49
	6	26.41	4.97	68.62	34.81	31.59	33.6	32.86	1.68	65.46
	7	26.73	5.59	67.68	36.23	31.36	32.42	33.1	1.77	65.13
average		25.8429	5.4641	68.6282	33.8611	32.9507	33.1196	32.8989	2.01721	65.0313

Tab. 3.2 continued.

		Monthly Wages			Hours			Hourly Wages		
Country	wave	dmw<0	dmw=0	dmw>0	dh<0	dh=0	dh>0	dhw<0	dhw=0	dhw>0
IRELAND	2	25.51	4.71	69.78	32.04	41.69	26.27	32.19	1.59	66.22
	3	28.44	4.25	67.31	26.91	46.55	26.55	34.13	2.02	63.86
	4	23.32	3.37	73.32	26.65	48.05	25.3	29.11	1.89	69
	5	19.12	3.73	77.15	24.83	51.81	23.36	25.57	2.49	71.95
	6	24.07	6.46	69.46	25.58	49.52	24.9	31.64	3.3	65.06
	7	24.1	1.88	74.02	32.65	46.5	20.85	24.79	0.85	74.36
average		23.9272	3.8094	71.7644	27.9476	47.2478	24.457	29.3668	1.86734	68.306
ITALY	2	27.34	22.15	50.5	18.87	61.07	20.06	32.85	14.78	52.37
	3	25.11	17.29	57.6	19.47	60.56	19.97	29.54	12.24	58.21
	4	23.55	17.46	58.99	17.8	61.54	20.66	30.25	11.97	57.78
	5	23.61	20.83	55.56	25.04	52.92	22.04	28.39	12.21	59.4
	6	25.04	24.61	50.35	24.2	54.32	21.48	30.72	15.5	53.78
	7	24.6	25.85	49.56	20.13	55.33	24.54	32.54	16.16	51.31
average		24.8437	21.1144	53.6287	20.748	57.5155	21.4039	30.6746	13.7038	55.3874
GREECE	2	25.35	7	67.65	29.69	45.31	25	28.98	3.92	67.1
	3	13.63	9.98	76.4	27.01	53.77	19.22	17.54	5.72	76.74
	4	13.45	9.1	77.45	22.04	53.59	24.37	21.13	5.97	72.9
	5	20	10.53	69.47	20.4	57.17	22.43	26.9	7.13	65.96
	6	22.43	15.86	61.71	24.79	54.03	21.18	27.95	10.35	61.71
	7	21.22	31.32	47.46	16.36	63.15	20.49	28.2	24.01	47.79
average		18.8062	12.2172	65.8381	22.953	54.2416	22.0195	24.7117	7.86767	64.6508
SPAIN	2	23.48	3.13	73.4	26.22	47.6	26.18	27.41	1.61	70.99
	3	36.3	2.81	60.89	27.82	47.59	24.59	37.94	1.08	60.98
	4	42.72	2.78	54.5	27.94	47.88	24.18	42.64	1.39	55.97
	5	33.13	2.88	63.99	25.98	48.79	25.23	36.77	0.91	62.32
	6	36.8	1.92	61.28	27.66	50.44	21.9	38.24	1.12	60.64
	7	33.71	3.22	63.07	24.68	50.08	25.24	38.62	1.69	59.69
average		33.8203	2.75301	62.6107	26.6896	48.7163	24.5154	36.6107	1.26834	61.6034
PORTUGAL	2	16.85	14.4	68.75	22.8	56.91	20.29	24.16	8.24	67.6
	3	14.29	9.3	76.41	22.16	62.12	15.73	18.48	6.35	75.17
	4	12.76	11.55	75.69	36.85	53.04	10.11	15.58	6.44	77.98
	5	12.69	12.22	75.09	25.69	62.52	11.79	16.84	8.48	74.68
	6	8.44	13.31	78.26	13.28	75.74	10.98	13.62	10.65	75.73
	7	9.61	14.85	75.55	15.64	75.35	9	13.94	11.75	74.31
average		12.1153	12.4566	74.8979	21.5207	63.715	12.4696	16.7697	8.42484	74.1735
AUSTRIA	3	50.6	16.79	32.61	22.08	48.26	29.65	55.09	9.28	35.63
	4	27.75	25.55	46.7	23.08	56.78	20.15	34.35	16.99	48.66
	5	21.61	21.36	57.03	19.07	60.08	20.85	27.54	13.64	58.81
	6	18.15	27.05	54.79	18.96	64.05	16.99	24.32	20.8	54.88
	7	14.32	30.72	54.95	16.34	65.26	18.4	22.34	23.15	54.5
average		23.954	23.7863	48.249	19.7576	58.5512	20.801	30.9316	15.96	49.7565
FINLAND	4	22.44	17.18	60.37	21.88	54.32	23.8	31.03	9.76	59.21
	5	21.25	15.66	63.08	22.91	54.79	22.3	29.01	8.85	62.14
	6	22.55	20.77	56.68	20.33	60.53	19.14	29.87	12.04	58.1
	7	18.12	18.63	63.25	24.79	55.56	19.66	24.1	10.6	65.3
average		21.0098	17.9625	60.7856	22.4193	56.2469	21.1398	28.3724	10.2466	61.1243

Tab. 3.2 continued

		Monthly Wages			Hours			Hourly Wages		
Country	wave	dmw<0	dmw=0	dmw>0	dh<0	dh=0	dh>0	dhw<0	dhw=0	dhw>0
GERMANY echp	2	19.12	29.37	51.5	35.25	41.62	23.13	26.35	12.08	61.58
	3	22.01	27.67	50.32	30.79	46.67	22.55	29.65	13.03	57.32
average		20.5142	28.5073	50.9066	32.9446	44.0727	22.8382	27.9513	12.546	59.4118
LUXEMBURG echp	2	22.12	20.22	57.66	15.72	76.57	7.7	24.33	14.69	60.98
	3	27.18	16.97	55.85	9	82.65	8.35	29.49	12.36	58.15
average		24.5198	18.5239	56.7478	11.8945	79.5519	8.01842	26.786	13.4747	59.5482
UK echp	2	22.13	12.4	65.47	30.18	38.01	31.81	30.33	5.67	64
	3	23.3	11.25	65.46	29.87	40.14	29.99	30.72	5.31	63.97
average		22.7075	11.811	65.465	30.0246	39.0605	30.8866	30.5244	5.48705	63.985

Tab. 3.3: Estimates of the econometric model

GERMANY GSOEP							GERMANY ECHP						
log pseudo-likelihood=	10361.972									log pseudo-likelihood=			2632.336
				Numbe	r of obs =	17885.000					Numbe	r of obs =	5072.000
				Wald	chi2(7) =	188.240					Wald	chi2(3) =	19.670
log pseudo-likelihood=	361.972			Prob	> chi2 =	0.000	log pseudo-likelihood=	632.336			Prob	> chi2 =	0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]	dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta							beta						
age	-0.001	0.000	-6.720	0.000	-0.001	-0.001	age	-0.001	0.000	-2.670	0.008	-0.001	0.000
female	0.000	0.003	-0.080	0.939	-0.005	0.005	female	-0.002	0.005	-0.400	0.686	-0.012	0.008
time3	0.023	0.004	5.670	0.000	0.015	0.031	time3	-0.016	0.005	-3.460	0.001	-0.025	-0.007
time4	-0.021	0.004	-4.780	0.000	-0.030	-0.012							
time5	-0.016	0.004	-3.800	0.000	-0.025	-0.008							
time6	-0.015	0.004	-3.780	0.000	-0.023	-0.007							
time7	0.001	0.005	0.230	0.817	-0.008	0.010							
_cons	0.043	0.006	7.150	0.000	0.032	0.055	_cons	0.071	0.011	6.550	0.000	0.050	0.092
se							se						
_cons	0.137	0.001	117.590	0.000	0.135	0.140	_cons	0.153	0.002	66.480	0.000	0.148	0.157
sm							sm						
_cons	0.073	0.002	36.980	0.000	0.069	0.077	_cons	0.049	0.007	6.950	0.000	0.035	0.063
alpha							alpha						
_cons	0.177	0.007	26.670	0.000	0.164	0.190	_cons	0.128	0.017	7.480	0.000	0.094	0.161

NETHERLANDS

log pseudo-likelihood= 7515.987

Number of obs = 10751.000
Wald chi2(7) = 144.970
log pseudo-likelihood= 515.987
Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.002	0.000	-8.990	0.000	-0.002	-0.001
female	0.007	0.004	1.780	0.074	-0.001	0.014
time3	0.002	0.005	0.340	0.736	-0.008	0.011
time4	0.024	0.005	4.280	0.000	0.013	0.034
time5	0.023	0.006	4.040	0.000	0.012	0.034
time6	0.000	0.006	0.040	0.969	-0.011	0.011
time7	0.025	0.006	4.350	0.000	0.014	0.036
_cons	0.064	0.009	7.390	0.000	0.047	0.082
se						
_cons	0.135	0.002	69.970	0.000	0.131	0.138
sm						
_cons	0.066	0.002	28.390	0.000	0.061	0.070
alpha						
_cons	0.202	0.011	17.630	0.000	0.179	0.224

BELGIUM

log pseudo-likelihood= 4410.398

Number of obs = 9340.000
Wald chi2(7) = 62.330
log pseudo-likelihood= 410.398
Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-3.620	0.000	-0.001	0.000
female	-0.007	0.004	-1.550	0.120	-0.015	0.002
time3	-0.049	0.008	-6.330	0.000	-0.064	-0.034
time4	-0.036	0.008	-4.710	0.000	-0.052	-0.021
time5	-0.046	0.008	-5.950	0.000	-0.061	-0.031
time6	-0.034	0.008	-4.390	0.000	-0.050	-0.019
time7	-0.040	0.008	-5.150	0.000	-0.055	-0.025
_cons	0.071	0.011	6.250	0.000	0.049	0.094
se						
_cons	0.160	0.002	78.210	0.000	0.156	0.164
sm						
_cons	0.084	0.003	29.860	0.000	0.078	0.089
alpha						
_cons	0.230	0.011	20.080	0.000	0.208	0.253

Tab. 3.3 continued

LUXEMB0URG PSELL

log pseudo-likelihood= 4992.065

Numbe r of obs = 8279.000
Wald chi2(6) = 260.000
log pseudo-likelihood= 992.065 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.001	0.000	-7.280	0.000	-0.002	-0.001
female	-0.008	0.004	-1.710	0.088	-0.016	0.001
time3	0.089	0.006	13.990	0.000	0.077	0.102
time4	0.060	0.006	9.850	0.000	0.048	0.072
time5	0.031	0.006	5.280	0.000	0.020	0.043
time6	0.050	0.006	7.790	0.000	0.037	0.062
_cons	0.023	0.009	2.570	0.010	0.005	0.040
se						
_cons	0.148	0.002	66.320	0.000	0.144	0.153
sm						
_cons	0.068	0.003	22.070	0.000	0.062	0.073
alpha						
_cons	0.208	0.014	15.010	0.000	0.181	0.236

LUXEMBOURG ECHP

log pseudo-likelihood= 953.566

Numbe r of obs = 1240.000
Wald chi2(3) = 9.350
log pseudo-likelihood= 53.566 Prob > chi2 = 0.025

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-1.830	0.067	-0.001	0.000
female	-0.010	0.008	-1.310	0.189	-0.025	0.005
time3	-0.014	0.007	-2.040	0.041	-0.028	-0.001
_cons	0.072	0.016	4.570	0.000	0.041	0.102
se						
_cons	0.120	0.003	35.730	0.000	0.113	0.126
sm						
_cons	0.024	0.007	3.420	0.001	0.010	0.037
alpha						
_cons	0.069	0.015	4.690	0.000	0.040	0.098

Tab. 3.3 continued

UK BHPS

log pseudo-likelihood= 3850.800

Number of obs = 9981.000
Wald chi2(7) = 76.630
log pseudo-likelihood= 3850.801 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.002	0.000	-8.290	0.000	-0.002	-0.001
female	0.000	0.004	-0.080	0.933	-0.009	0.008
time3	0.009	0.007	1.310	0.191	-0.004	0.022
time4	0.012	0.007	1.680	0.092	-0.002	0.027
time5	0.014	0.007	1.910	0.056	0.000	0.028
time6	0.023	0.007	3.040	0.002	0.008	0.037
time7	0.018	0.007	2.370	0.018	0.003	0.032
_cons	0.077	0.010	7.800	0.000	0.058	0.096
se						
_cons	0.165	0.002	79.880	0.000	0.161	0.169
sm						
_cons	0.096	0.004	24.440	0.000	0.088	0.104
alpha						
_cons	0.229	0.015	15.070	0.000	0.199	0.259

UK ECHP

log pseudo-likelihood= 782.805

Number of obs = 3760.000
Wald chi2(3) = 31.610
log pseudo-likelihood= 782.805 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.002	0.000	-5.170	0.000	-0.002	-0.001
female	0.009	0.006	1.340	0.180	-0.004	0.021
time3	-0.009	0.006	-1.420	0.155	-0.021	0.003
_cons	0.083	0.013	6.180	0.000	0.056	0.109
se						
_cons	0.155	0.003	53.800	0.000	0.149	0.160
sm						
_cons	0.083	0.004	19.990	0.000	0.075	0.091
alpha						
_cons	0.206	0.015	13.490	0.000	0.176	0.235

Tab. 3.3 continued

IRELAND

log pseudo-likelihood= 795.975

Number of obs = 8389.000
Wald chi2(7) = 180.090
log pseudo-likelihood= 795.975 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.002	0.000	-9.440	0.000	-0.003	-0.002
female	0.001	0.006	0.090	0.927	-0.010	0.011
time3	-0.033	0.009	-3.740	0.000	-0.050	-0.016
time4	-0.009	0.009	-1.080	0.281	-0.026	0.008
time5	0.004	0.009	0.460	0.645	-0.013	0.021
time6	-0.013	0.009	-1.470	0.143	-0.031	0.004
time7	0.058	0.010	5.890	0.000	0.039	0.077
_cons	0.135	0.012	11.090	0.000	0.111	0.158
se						
_cons	0.203	0.002	87.590	0.000	0.199	0.208
sm						
_cons	0.093	0.005	18.750	0.000	0.083	0.103
alpha						
_cons	0.233	0.016	14.640	0.000	0.202	0.264

DENMARK

log pseud likelihood = 7777.589

Number of obs = 9349.000
Wald chi2(7) = 67.820
log pseudo-likelihood= 777.589 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-6.840	0.000	-0.001	-0.001
female	-0.009	0.003	-2.980	0.003	-0.014	-0.003
time3	-0.015	0.005	-3.170	0.002	-0.024	-0.006
time4	-0.008	0.005	-1.710	0.088	-0.017	0.001
time5	-0.013	0.005	-2.570	0.010	-0.023	-0.003
time6	-0.015	0.005	-3.050	0.002	-0.025	-0.005
time7	-0.013	0.005	-2.680	0.007	-0.023	-0.004
_cons	0.079	0.008	10.320	0.000	0.064	0.094
se						
_cons	0.117	0.001	80.650	0.000	0.114	0.120
sm						
_cons	0.051	0.002	29.200	0.000	0.047	0.054
alpha						
_cons	0.149	0.007	20.630	0.000	0.135	0.163

Tab. 3.3 continued

ITALY

log pseudo-likelihood= 7430.741

				Numbe	r of obs	= 16573.000
				Wald	chi2(7)	= 83.980
log pseudo-likelihood=		430.741		Prob	> chi2	= 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.001	0.000	-4.100	0.000	-0.001	0.000
female	0.000	0.003	-0.070	0.948	-0.006	0.006
time3	0.014	0.004	3.170	0.002	0.005	0.023
time4	0.024	0.006	4.180	0.000	0.013	0.036
time5	0.028	0.005	6.070	0.000	0.019	0.038
time6	0.004	0.005	0.900	0.366	-0.005	0.013
time7	-0.003	0.005	-0.660	0.508	-0.012	0.006
_cons	0.030	0.007	4.130	0.000	0.016	0.044
se						
_cons	0.159	0.001	130.010	0.000	0.156	0.161
sm						
_cons	0.070	0.004	16.720	0.000	0.061	0.078
alpha						
_cons	0.168	0.011	14.720	0.000	0.145	0.190

FRANCE

log pseudo-likelihood= 636.726

				Numbe	r of obs	= #######
				Wald	chi2(7)	= 1283.730
log pseudo-likelihood=		636.726		Prob	> chi2	= 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-6.910	0.000	-0.002	-0.001
female	0.002	0.003	0.580	0.564	-0.005	0.008
time3	0.148	0.006	26.580	0.000	0.137	0.158
time4	0.260	0.008	31.620	0.000	0.244	0.277
time5	0.181	0.006	29.620	0.000	0.169	0.193
time6	0.176	0.006	28.410	0.000	0.164	0.188
time7	0.195	0.007	29.400	0.000	0.182	0.208
_cons	-0.118	0.011	-11.110	0.000	-0.139	-0.097
se						
_cons	0.144	0.002	61.240	0.000	0.139	0.149
sm						
_cons	0.107	0.004	28.540	0.000	0.099	0.114
alpha						
_cons	0.234	0.020	11.630	0.000	0.195	0.274

Tab. 3.3 continued

GREECE

		1687.747				
				Numbe	r of obs =	8103.000
				Wald	chi2(7) =	315.460
log pseudo-likelihood=		687.747		Prob	> chi2 =	0.000
		Robust				
dlmhwg	Coef.	Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	0.000	0.000	-1.570	0.116	-0.001	0.000
female	0.007	0.005	1.370	0.170	-0.003	0.017
time3	0.048	0.008	5.700	0.000	0.031	0.064
time4	0.035	0.010	3.570	0.000	0.016	0.054
time5	-0.016	0.009	-1.810	0.071	-0.034	0.001
time6	-0.046	0.009	-5.400	0.000	-0.063	-0.030
time7	-0.075	0.008	-9.040	0.000	-0.091	-0.059
_cons	0.095	0.013	7.230	0.000	0.069	0.120
se						
_cons	0.201	0.002	97.040	0.000	0.197	0.205
sm						
_cons	0.075	0.009	7.990	0.000	0.057	0.094
alpha						
_cons	0.172	0.023	7.390	0.000	0.127	0.218

Tab. 3.3 continued

PORTUGAL

log pseudo-likelihood= 2.831

Number of obs = 12694.000
Wald chi2(7) = 50.710
log pseudo-likelihood= 2.831 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.001	0.000	-5.090	0.000	-0.001	0.000
female	-0.006	0.003	-2.190	0.029	-0.012	-0.001
time3	0.004	0.005	0.800	0.424	-0.006	0.014
time4	0.013	0.005	2.400	0.016	0.002	0.024
time5	0.001	0.005	0.140	0.890	-0.010	0.011
time6	-0.010	0.005	-2.020	0.043	-0.019	0.000
time7	-0.001	0.005	-0.240	0.810	-0.011	0.008
_cons	0.076	0.007	11.280	0.000	0.062	0.089
se						
_cons	0.146	0.002	81.210	0.000	0.143	0.150
sm						
_cons	0.047	0.005	10.010	0.000	0.037	0.056
alpha						
_cons	0.145	0.017	8.760	0.000	0.112	0.177

SPAIN

log pseudo-likelihood= 1398.558

Number of obs = 9840.000
Wald chi2(7) = 145.040
log pseudo-likelihood= 398.558 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-3.050	0.002	-0.001	0.000
female	-0.009	0.005	-1.890	0.059	-0.018	0.000
time3	-0.055	0.007	-8.360	0.000	-0.068	-0.042
time4	-0.080	0.008	-9.860	0.000	-0.096	-0.064
time5	-0.048	0.008	-6.310	0.000	-0.062	-0.033
time6	-0.049	0.008	-6.480	0.000	-0.064	-0.034
time7	-0.052	0.008	-6.840	0.000	-0.067	-0.037
_cons	0.122	0.011	10.840	0.000	0.100	0.144
se						
_cons	0.192	0.003	54.980	0.000	0.185	0.199
sm						
_cons	0.106	0.007	14.120	0.000	0.091	0.120
alpha						
_cons	0.145	0.013	10.990	0.000	0.119	0.171

Tab. 3.3 continued

AUSTRIA

log pseudo-likelihood= 398.835

Numbe r of obs = 6033.000
Wald chi2(6) = 246.660
log pseudo-likelihood= 398.835 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.001	0.000	-3.570	0.000	-0.001	0.000
female	-0.005	0.004	-1.280	0.202	-0.013	0.003
time3	0.066	0.007	9.240	0.000	0.052	0.080
time4	0.080	0.006	13.210	0.000	0.068	0.091
time5	0.070	0.006	12.130	0.000	0.059	0.082
time6	0.080	0.006	13.660	0.000	0.069	0.092
_cons	-0.028	0.009	-3.090	0.002	-0.046	-0.010
se						
_cons	0.147	0.002	74.220	0.000	0.143	0.151
sm						
_cons	0.035	0.003	10.470	0.000	0.028	0.041
alpha						
_cons	0.100	0.008	12.320	0.000	0.084	0.116

FINLAND

log pseudo-likelihood= 718.960

Numbe r of obs = 4596.000
Wald chi2(5) = 29.650
log pseudo-likelihood= 718.960 Prob > chi2 = 0.000

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-4.650	0.000	-0.002	-0.001
female	-0.007	0.005	-1.320	0.186	-0.017	0.003
time3	0.010	0.006	1.710	0.086	-0.001	0.022
time4	0.004	0.008	0.480	0.634	-0.012	0.020
time5	0.018	0.009	2.030	0.042	0.001	0.034
_cons	0.066	0.014	4.700	0.000	0.039	0.094
se						
_cons	0.142	0.003	56.070	0.000	0.137	0.147
sm						
_cons	0.070	0.004	16.970	0.000	0.062	0.078
alpha						
_cons	0.186	0.016	11.930	0.000	0.156	0.217

APPENDIX 1

In this appendix we enclose descriptive statistics of wage, hours, and hourly wage distributions (Tab. A1.1-A1.3). Descriptive statistics of wage, ours, and hourly wage change distributions follow (Tab. A.1.4-A1.6) insert tab1-33pagine

APPENDIX 2

In this appendix we illustrate the likelihood function that has been estimated for the econometric model presented in chap.3.

The basic model is

$$y_{it} = \begin{cases} x_{it}\beta + e_{it} + m_{it} & if \quad 0 \leq x_{it}\beta + e_{it} \\ m_{it} & if \quad -\alpha \leq x_{it}\beta + e_{it} \leq 0 \\ x_{it}\beta + e_{it} + m_{it} & if \quad x_{it}\beta + e_{it} \leq -\alpha \end{cases}$$

Assume that $e_{it} \sim N\left(0, \sigma_e^2\right)$, $m_{it} \sim N\left(0, \sigma_m^2\right)$ are independently distributed. Thus,$(e_{it} + m_{it}) \sim N\left(0, \sigma_e^2 + \sigma_m^2\right)$ and $E\left(em\right) = 0$.

For any observation (i, t) there are three possible mutually exclusive regimes, so that the likelihood function is given by

$$L_{it}\left(\vartheta\right) = L1_{it}\left(\vartheta\right) + L2_{it}\left(\vartheta\right) + L3_{it}\left(\vartheta\right),$$

where $\vartheta = \left(\beta, \alpha, \sigma_e, \sigma_m\right)$.

The term of the likelihood for regime 1, $L1_{it}\left(\vartheta\right)$, is derived from

$$y_{it} = x_{it}\beta + e_{it} + m_{it} \quad if \quad 0 \leq x_{it}\beta + e_{it}$$

to give:

$$\begin{aligned} L1_{it}\left(\vartheta\right) &= \Phi\left[\frac{x_{it}\beta + \left(\frac{\sigma_e^2}{\sigma_e^2+\sigma_m^2}\right)\left(y_{it} - x_{it}\beta\right)}{\sqrt{\frac{\sigma_e^2\sigma_m^2}{\sigma_e^2+\sigma_m^2}}}\right] \\ &\quad \frac{1}{\sqrt{\sigma_e^2 + \sigma_m^2}}\varphi\left(\frac{y_{it} - x_{it}\beta}{\sqrt{\sigma_e^2 + \sigma_m^2}}\right). \end{aligned}$$

So

$$L1_{it}(\vartheta) = \Phi\left[\frac{x_{it}\beta + \left(\frac{\sigma_e^2}{\sigma_e^2+\sigma_m^2}\right)(y_{it} - x_{it}\beta)}{\sqrt{\frac{\sigma_e^2\sigma_m^2}{\sigma_e^2+\sigma_m^2}}}\right] \frac{1}{\sqrt{2\pi}\sqrt{\sigma_e^2+\sigma_m^2}}\exp\left[-\frac{(y_{it}-x_{it}\beta)^2}{2\left(\sigma_e^2+\sigma_m^2\right)}\right] \tag{10}$$

The piece of the likelihood function for Regime 2 is obtained from

$$y = m_{it} \text{ if } -\alpha - x_{it}\beta \leq e_{it} \leq -x_{it}\beta.$$

So,

$$L2_{it}(\vartheta) = \left[\Phi\left(\frac{-x_{it}\beta}{\sigma_e}\right) - \Phi\left(\frac{-\alpha - x_{it}\beta}{\sigma_e}\right)\right]\frac{1}{\sigma_m}\varphi\left(\frac{y_{it}}{\sigma_m}\right)$$

Finally, the likelihood term for Regime 3 stems from:

$$y_{it} - x_{it}\beta = e_{it} + m_{it} \text{ if } e_{it} \leq -\alpha - x_{it}\beta$$

and is given by

$$L3_{it}(\vartheta) = \Phi\left[\frac{-\alpha - x_{it}\beta - \left(\frac{\sigma_e^2}{\sigma_e^2+\sigma_m^2}\right)(y_{it} - x_{it}\beta)}{\sqrt{\frac{\sigma_e^2\sigma_m^2}{\sigma_e^2+\sigma_m^2}}}\right] \frac{1}{\sqrt{\sigma_e^2+\sigma_m^2}}\varphi\left(\frac{y_{it}-x_{it}\beta}{\sqrt{\sigma_e^2+\sigma_m^2}}\right)$$

Thus, the log-likelihood can be formulated as follows

$$\ln L = \sum_{i=1}^{N}\sum_{t=1}^{T}\ln\left[L1_{it}(\vartheta) + L2_{it}(\vartheta) + L3_{it}(\vartheta)\right]$$