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Wage rigidity at the micro-level in the European Union's countries: evidence and estimation issues

by

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ABSTRACT

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WAGE RIGIDITY AT THE MICRO-LEVEL IN THE EUROPEAN UNION'S COUNTRIES: EVIDENCE AND ESTIMATION ISSUES

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In this thesis we analyse the extent of wage rigidity in the European countries using data at the micro-level. After an extensive survey of the literature, we provide evidence of nominal wage rigidity in 14 EU countries using the individual surveys of the 1994-2000 European Community Household Panel. We follow both a descriptive and an econometric approach, taking into account reporting error issues in constructing appropriate measures of downward nominal wage rigidity. We find that the extent of nominal wage rigidity observed increases when reporting errors are modelled according to the classical assumptions. Therefore nominal wages are quite rigid in the EU countries, although measures are different across countries.

We therefore move to try to explore the causes of wage rigidity in Europe, focusing on the institutional characteristics of labour markets. We find that there is an hump-shaped relationship between employment protection legislation and nominal wage flexibility measures.

Then the French case is analysed in detail, comparing data of different nature (declarative and administrative). A validation study is carried out for the French Labour Force Survey (FLFS), showing that rounding behaviour of individuals does not follow the classical assumptions. This has an impact on the observed measures of wage rigidity: whereas the observed extent of wage rigidity in France is quite high in the FLFS, no evidence of wage rigidity is found in administrative data.

In the last chapter we question the appropriateness of measures of wage rigidity based only on individual data, and construct an appropriate matched employer-employee data set for France that allows to link individuals wage dynamics to measures of idiosyncratic, firm-level shocks. We therefore define wage rigidity as asymmetric adjustments of wages to firm-level shocks. Testing this definition on French data we conclude that, although the reporting-error free distribution of wage changes does not show evidence of wage rigidity in France, wage react asymmetrically to positive and negative shocks. Therefore, according to a more general definition, nominal wages are rigid in France.

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1 Introduction

In recent years a number of countries have adopted explicit inflation targets for monetary policy, reflecting a general agreement that monetary policy must ensure low inflation. The deliberate policy of low inflation has led to renewed interest among academics as well as policy makers for the contention of Tobin (1972) that if policy aims at low inflation, downward rigidity of nominal wages may lead to higher wage pressure, involving higher equilibrium unemployment (Akerloff et al, 1996, 2000, Holden, 1994, and Wyplosz, 2001). Other economists have been less concerned, questioning both the existence of downward nominal wage rigidity (DNWR), and the possible macroeconomic effects (Gordon, 1996 and Mankiw, 1996). The issue has also received considerable attention among policy makers, (ECB, 2003, OECD, 2002, and IMF, 2002).

To shed light on this issue, a fast growing body of empirical research has explored the existence of DNWR in many OECD countries. Due to the recent availability of individual panel data, particular attention has been devoted to the analysis of individual wage change distributions for employees staying in the same firm for constructing measures of DNWR. Assuming that, in absence of DNWR, wages would adjust freely to productivity changes of individuals, the distribution of wage changes should be smooth and symmetric. DNWR can be identified with the existence of a spike at zero wage change, a general asymmetry of the distribution around zero, and a very low percentage of wage cuts.

The quality of data available, the measure of wage reported, and the information available for selecting correctly the sample of stayers, can bias the shape of wage change distributions with relevant consequences for the construction of measures of DNWR and inter-country comparisons. Two major issues arise in using micro-data for determining the extent of DNWR. The first problem is linked to the measure of wage available. One measure that the researcher would like to use for studying wage rigidity would be the contracted hourly base-wage. Unfortunately very rarely this measure is available in panel data-sets, both of declarative and administrative source. Often measurement errors are modelled according to the classical assumption, but in this case appropriate validation studies should suggest the relevance of classical assumptions for measures of wage rigidity.

The second issue arising from the use of wage change distributions for the analysis of DNWR is the strategy used for estimating measures of wage rigidity. Since these measures arise from the comparison between the observed wage change distribution and a counterfactual in absence of rigidity, it is crucial to assess how the counterfactual is constructed. The counterfactual can either be based on statistical properties of the wage change distribution or be estimated from observed characteristics of firms and individuals.

This thesis analyses DNWR in the EU countries, emphasising both measurement and estimation aspects. After an introductory survey of the literature, summarising the existing results for the US and some European countries and the estimation methods previously used in the literature, two chapters are devoted to the analysis of 14 European countries in an intercountry perspective. Chapter 3 describes the data used, the recently available first 7 waves of the European Community Household Panel (ECHP). Although this data-set has the advantage of allowing intercountry-comparisons, serious problems of rounding when individuals report their wage can bias DNWR measures and intercountry comparisons. Different measures of DNWR are constructed country by country. In particular, the observed percentage of nominal wage cuts and freezes are compared with measures estimated from an econometric model that takes into account of measurement error using the classical assumptions. The result is that a quite high extent of DNWR is observed in the EU countries.

Chapter 4 is an attempt at exploring the causes of DNWR in Europe, considering the institutional characteristics of the European labour markets. Focusing on intercountry comparisons, a meta-analysis is carried out showing that there is a robust hump-shaped relationship between employment protection legislation and the measures of DNWR of Chapter 3.

The second part of the thesis focuses on the French case. It has been carried out at the French National Institute of Statistics (INSEE), during my two years stay. The unique opportunity of having access to individual data of different source and the possibility of matching them with firm data has allowed two kind of analyses. In Chapter 5 we present a validation study of the French Labour Force Survey, based on the direct comparison of wage declared by individuals and reported by firms for administrative purposes. In this way, the impact of rounding behaviour of individuals on measures of DNWR is analysed, concluding that the classical assumptions might be very distortionary in France for measuring the extent of DNWR: whereas using wages reported we would conclude that wages are rigid in France, on administrative data no evidence of nominal wage rigidity is found.

In Chapter 6 we argue that the concept of DNWR used so far in the literature might be very restrictive. It is based only on the observation of individual data, without linking individual wage dynamics to firm - level shocks. This is mainly due to the unavailability and the difficulty of constructing an appropriate data - set. We therefore match three individual data - sets of different nature with firm balance sheets, from which an appropriate measure of firm - level shocks is taken. We then introduce a more general definition of DNWR, according to which wages are rigid if they react asymmetrically to positive or negative firm - level shocks. This new definition of DNWR is tested on our French matched employer - employee data showing that, although the analysis of wage change distributions would not imply the existence of DNWR, wages do not react symmetrically to firm-level shocks, and therefore there is evidence of DNWR in France.

Chapter 7 concludes.

2 Survey of the literature

The target of low inflation for monetary policy has been recently quite debated in the literature, given the 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 raise 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, Akerloff, 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

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

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. In this chapter we survey the evidence available for the US and a number of European countries, focusing on the methodologies adopted by the authors.

Some of the empirical strategies aim at testing the existing theoretical micro-foundations of the nominal wage rigidity assumption. The chapter is therefore divided in two sections. Section 2.1 summarises the implications of models of nominal wage rigidity.

In section 2.2 we describe how the above implications have been tested in various countries. Section 2.3 concludes.

2.1 Models of nominal wage rigidity

Several different explanations for nominal wage rigidity have been proposed in the literature. According to Holden (2004) "we can distinguish three groups of models: 1) models of coordination failure: concern for relative wages; 2) fairness: wage cuts are viewed as unfair; 3) legal restrictions: wages are given in contracts that can only be changed by mutual consent."

Keynes (1936) introduced the *coordination failure* argument. The basic idea is that workers are concerned about relative wages, and thus oppose nominal wage cuts as this lead to lower relative wages. Workers are less opposed to the same reduction in real wages if it takes place via higher prices, as this does not affect real wages. Bhaskar (1990) provides additional micro-foundations for this idea, based on the assumption that workers' disutility of being paid less than others is greater than the utility gain of being paid more.

The *fairness* argument has characterised much of the first empirical work on the subject. Nominal wage cuts are not implemented by firms as both employers and employees think they are unfair. This idea involves money illusion, and therefore runs counter to the standard rationality arguments. However there is considerable evidence, mainly coming from personal interviews, documenting the existence of money illusion: Bewley (1999), Shafir, Diamond and Tversky (1997), Fehr and Tyran (2001).

Among many economists, mechanisms based on money illusion are met with considerable scepticism, based on the argument that rational agents care only about real variables, so that any effect of nominal variables due to money illusion will disappear over time. The *legal restrictions* argument, however, explains the nominal wage rigidity mechanism in models with rational agents, and is therefore particularly relevant in the literature. It is a fact of life that, in most industrialised economies, workers have their wage set in some type of contract, either an individual employment contract (in the US) or a collective agreement (in most of the European countries). Payment is typically specified in nominal terms although annual, partial indexation to the consumer price index is sometimes used, in particular in periods of high inflation. Such contracts, quite widespread in most Western European countries given the extensive coverage of collective agreements[2], are not adjusted continuously. There may be several reasons for the prevalence of rigid wage contracts. For individual employment contracts the following motiva-

[2]See Taylor (1999) and Boeri et al. (2001).

tions have been considered[3]: 1) to share risk; 2) to protect the parties' investments from hold-up; 3) to avoid renegotiation costs for small wage changes (menu-costs); and 4) to avoid opportunistic behaviour of employees (efficiency wages). All of them are plausible explanations of wage rigidity, but whereas models of risk sharing and efficiency wages predict real wage rigidity, hold-up can explain either real or nominal wage rigidity and menu-costs theories imply nominal wage rigidity. In particular hold-up models are the only ones, among the theories of individual employment contracts, that can explain downward nominal wage rigidity. Menu-costs predict nominal wage rigidity but not deal with the downward nominal wage rigidity concept. Moreover, whereas the first three groups of theories have clear predictions on the dynamics of wages, it is quite difficult to predict the implications of efficiency wage theories for wage changes. Therefore efficiency wage models can not be tested through the analysis of wage change distributions. On the contrary, some clear implications on nominal and real wage change dynamics can be investigated exploring some characteristics of individual wage change distributions. For this purpose, it is useful to clarify how wages are determined according to individual employment contracts theories of risk sharing, hold-up and menu-costs.

Malcomson (1999) shows that, allowing for renegotiation by mutual consent, as is common in Europe, the dynamics implied for wages by risk sharing and hold-up theories is the following. At the beginning of the employment relationship wage is determined at some level between the outside option of the employee (the minimum wage she is willing to accept) and the outside option of the firm (the maximum wage it is able to pay the employee). The

[3]These theories have been recently surveyed in Malcomson (1999), and we follow his approach in this very short summary.

employment relationship is efficient as long as the outside option of the employer is higher than the outside option of the employee, otherwise either the employee quits or she is fired. The outside options in these models act as constraints for the wage. It can be shown that, once a certain wage is contracted at the beginning of the employment relationship, it is not renegotiated (i.e. stays constant) until either the firm's or the employee's outside option becomes binding. If the employer's outside option is binding, the contracted wage will be renegotiated downward, whereas if the employee's outside option is binding the wage is renegotiated upward. In either case, the change in wage can be high or small, and not necessarily symmetric: what matters is that the wage change follows exactly the change in the outside options. For these reasons, we expect to observe wages generally constant for employees staying in the same firm (stayers) over time. Any changes in stayers' wages are explained as a change in their outside options, when they become binding. At the same time, we expect to observe flexible wages for movers. In fact an efficient separation takes place whenever the outside option of the employer becomes lower than the outside option of the employee. In particular, layoffs occur when the outside option of the firm is too low for the employee to accept it and quits are efficient for employees when their outside options become too high for the firm to be able to pay them. Therefore movers' wages are supposed to move either upward or downward. Movers' wages can never be constant according to these models as, if the firm or the employee were willing to accept the same wage, it would be efficient for them to keep their work relationship.

The difference between risk-sharing models and hold-up models rests in the motivations for writing a contract, that are crucial for the choice of the variable contracted. Contracts to allocate risk consider risk-averse employees whose purpose is to insure against any fluctuation in consumption. If financial markets cannot provide such insurance, employers may provide it instead, considering that they have better relevant information than financial markets. Since insurance to the employee will be given in the form of a constant *real* wage, these models predict *real* wage rigidity. Hold-up models instead emphasize the concern of the employer/employee in protecting their investment once they have incurred in it, after writing the contract. In presence of turn-over costs, this issue is relevant not only for specific but also for general investments by the firm or the employee. The constant wage predicted by these models can be either *real* or *nominal,* even though the presence of any cost in indexing the wage is considered to be a sufficient reason to contract a *nominal* wage. The analysis in Malcomson (1999) indicates that, in this case, if the contracted wage is sufficiently low that the firm's outside option constraint never binds, the wage is not renegotiated downwards, so hold-up theories can explain also downward nominal wage rigidity.

Menu costs theories predict that maximizing firms facing small menu costs (costs arising from renegotiating the nominal wage) will not find it profitable to renegotiate the wage for a small amount. Menu costs will be incurred in only when counterbalanced by sufficiently large nominal wage changes. Therefore, menu costs theories imply *nominal* wage rigidity but symmetric wage changes around zero for stayers. No explanation of completely downward wage rigidity is given by this theories.

The implications of theories of wage determination at individual level are summarised in Table 1. There are three important implications of theories of nominal wage rigidity that, according to Malcomson (1999) can be tested on wage change distributions: 1) no change in nominal wage (Nominal Wage

Rigidity, NWR), or no change in real wage (Real Wage Rigidity, RWR); 2) no nominal wage cuts (Downward Nominal Wage Rigidity, DNWR); and 3) asymmetry of small wage changes around zero (hold-up theories versus menu-costs).

Table 1 Summary of implications of theories of wage determination

THEORIES	STAYERS	MOVERS
Contracts to protect Investments (HOLD-UP) & renegotiation by mutual consent	1. NW Rigidity or RW Rigidity NW Rigidity if indexing is costly 2. N/R W changes can be high/small 3. N/R W changes can be asymmetric 4. DNW Rigidity	Flexible W
Contracts to Insure employees (Risk-Sharing) & renegotiation by mutual consent	1.-3. for Real Wages	Flexible RW
MENU COSTS	1. NW Rigidity 2. NW changes can be only high (drops near zero) 3. NW changes symmetric	Flexible NW

N=Nominal; R=Real; W=Wage

The problem in testing theories of individual employment contracts in the European countries is that they do not consider the role of trade unions in renegotiating the contract. In Europe, wages can be determined at different levels: national, sector, and firm-level. There are differences across the European countries about the degree of centralisation of wage bargaining and labour market institutions characteristics. Therefore, the individual wage observed is the result of all the levels of bargaining, and can be interpreted with the theories of individual employment contracts probably only for its very last stage of determination. Unfortunately, there are no theoretical models in which different levels of wage bargaining are taken into account and simultaneously considered. However, Holden (1994) shows that the same mechanism considered in MacLeod and Malcomson (1993) for hold-up models is valid for collective agreements. Therefore downward nominal wage rigidity is explained also in the collective agreements case. Other theories in which the parties involved in wage determination are not individuals but unions are: staggered contracts theories, that predict nominal wage rigidity, and insider-outsider models, that explain real wage rigidity.

It is important to notice that, while in a macro-context by wage flexibility we usually mean how wages react to unemployment, at the micro-level wage rigidity is defined as no change in wages. Testing how wages react to employment at the micro-level would require a model of simultaneous determination of labour earnings and number of people employed, to date not yet available. Therefore it is not easy to evaluate the impact of wage rigidities on employment at the micro-level.

As far as the causes of wage rigidity are concerned, from this section's discussion we can conclude that there are various explanations for why wages are downwardly rigid in the literature. Certainly using individual wage change distributions can help in testing the implications of the existing models but hardly they can allow to distinguish among the different causes of wage rigidity. This is particularly unfortunate both for scientific understanding and for the analysis of economic policy.

2.2 Micro-evidence of wage rigidity

The identification of wage rigidity, and the construction of relative measures of downward wage rigidity, requires the comparison between an observed distribution of wage changes (*actual* distribution) and an hypothetical distribution, supposed to hold in absence of wage rigidity (*counterfactual* or *notional* distribution). The counterfactual distribution is determined by employers and employees outside options behaviour, as well as by productivity. Each of the approaches proposed in the literature is based on different assumptions on the counterfactual distribution. In particular, we can distinguish three main methodologies that have been adopted in the literature: 1) what we call the simple descriptive analysis of wage change distributions; 2) the location approach; 3) the structural approach.

Descriptive analyses are implicitly based on the assumption of smoothness of the counterfactual distribution. The location approach assumes that in absence of rigidity wage change distributions are symmetric around the median. In the structural approach instead no particular assumption is made on the shape of the counterfactual distribution: it is estimated on the basis of the information available on employers and employees.

In this section we present the 3 methods above separately, discussing how they have been implemented in different countries and their results.

2.2.1 Descriptive analyses of wage change distributions

If wages were completely flexible they should accommodate any intertemporal change of employers' and employees' outside options. Employers' outside options depend on measures of demand shock, productivity, etc.. Workers' outside options depend instead on individual characteristics such as age, sex, experience, tenure, and so on. In the simplest initial framework we can suppose that, since the determinants of outside options are quite flexible over time, the *counterfactual* distribution, whichever its shape, should be smooth. As a consequence, a spike anywhere in the *observed* distribution of wage changes would be evidence of the existence of no change in wage at that point, signalling the existence of wage rigidity. In particular, if we consider nominal wage change distributions, a spike at zero could be interpreted as evidence of nominal wage rigidity, whereas a spike at the rate of inflation would indicate a certain extent of real wage rigidity. Moreover, in presence of nominal wages completely downwardly rigid, all negative notional wage changes would result in no wage changes and symmetric drops around zero would be evidence of menu-costs. The implications of theories of wage determination can thus be translated in characteristics of the *observed* distribution of nominal wage changes, that can be used for testing purposes. Table 2 shows the relationship between the implications of theories of wage determination and observable characteristics of wage change distributions. Since most of the empirical literature concentrates on stayers, we omit in the table the implications for movers. However, for movers, according to the theories of wage determination we would expect to observe quite flexible distribution, without spike at zero.

Table 2 Relationship between implications of theories of wage determination and stayers' nominal wage change distributions

THEORIES	IMPLICATIONS	NW CHANGE DISTRIBUTIONS CHARACTERISTICS
Contracts to protect Investments (HOLD-UP) & renegotiation by mutual consent (m.c.)	1. NW rigidity or RW rigidity NW rigidity if indexing is costly 2. N/R W changes can be high/small 3. N/R W changes can be asymmetric 4. DNW rigidity	1. Spike at zero (NW rigidity) or at the rate of inflation gp (RW rigidity) 2. No drops in a small interval around zero/gp 3. No symmetry around zero/gp 4. Rare wage cuts
Contracts to Insure employees (Risk-Sharing) & renegotiation by m.c.	1.-3. just for Real Wages	1-3 for the rate of inflation gp
MENU COSTS	1. NW rigidity 2. NW changes can be only high 3. NW changes symmetric	1. Spike at zero 2. Drops in a small interval around zero 3. Symmetric drops in a small interval around zero

N=Nominal; R=Real; W=Wage; gp=inflation rate; m.c.=mutual consent.

Empirical evidence

The first step for identifying the existence of wage rigidity is therefore a simple investigation of wage change distributions for employees staying in the same firm, same job, for at least two consecutive years. The ideal data-set for examining nominal wage rigidity would be a representative sample of firms personnel files including precise information on wages, individuals productivity and other individual characteristics. There is no study available in the literature with such a data-set. Bewley (1998), Altonji and Devereux (2000) and Fehr and Goette (2003) provide evidence from non-representative firm-level information based on interviews. In all the above studies the distribution of employees wage changes (measured in log wage differences) are completely downwardly rigid, exhibiting a prominent spike at zero and almost no wage cuts.

A more extensive analysis was introduced by McLaughlin (1994), on US Labour Force Survey data from the 1976-1986 Panel Study for Income Dynamics (PSID). The advantage in using individual surveys for studying wage rigidity is that the sample is representative of the population. The disadvantages consist in the difficulty in: 1) selecting correctly employees staying in the same firm, since normally firms identifiers are not available in individual surveys; and 2) getting a correct measure for the contracted base-wage, since individuals usually report total labour earnings. Usually, stayers are defined in this literature as employees not changing job, and sometimes checks of no change in sector or occupation are carried out. In the majority of the cases, a proxy for hourly wages is constructed dividing labour earnings reported by the number of hours. Sometimes instead, the sample is restricted to employees not changing the number of hours worked. Since both earnings and number of hours can be subject to reporting errors, this proxy is known to be possibly quite dirty. The impact of reporting errors can be studied either in a descriptive framework, with appropriate validation studies, or making assumptions on the distribution of reporting errors. This last method has been followed in the literature mainly in a structural approach framework, and will be discussed later on. The results of the descriptive analyses are given by country: we start from the US and then move to some European countries.

The first country for which individual wage change distributions have been analysed is the US. Pooling all the years together, McLaughlin (1994) observes 17% nominal wage cuts and 43% real wage cuts, therefore concluding that wages are quite flexible in the US. He does not find support for the menu-costs theory, since there are no drops of the distribution in a small interval around zero. Using the same data, but focusing on year by year changes, both Card and Hyslop (1997) and Kahn (1997) find instead quite relevant spikes at zero wage changes: from 6% to 10% in years of high inflation (the late '70s) to over 15% in the low inflation era (mid '80s). They therefore conclude that there is a certain degree of nominal wage rigidity in the US. At the same time, in the data many individuals, about 20% on average, experience wage cuts. Interestingly, Kahn (1997) considers two different measures of wages: hourly wages for wage earners and total wages for salary earners, finding that pay cuts are more frequent in this last case, and are not due to changes in hours. But Card and Hyslop (1997) do not confirm this result.

McLaughlin (1994) and Card and Hyslop (1997) take into account previous results from validation studies on earnings and hours reported in the PSID. Although earnings are better reported than hours, the measure they consider is hourly wages, calculated dividing total earnings by the number of hours. Simply correcting for measurement error does not alter qualitatively the above results: it just reduces slightly both the spike and the percentage of wage cuts.

The impact of measurement error from reporting earnings is taken into account for the first time in the nominal wage rigidity literature by Smith (2000). She analyses the first seven waves of the British Household Panel Survey, that cover the 1991-1996 period, taking advantage from a particular information available in this data-set: for a number of employees it is known whether wages have been reported by people or if they have been taken directly from the payslip shown to the interviewer. Considering monthly wages of employees who do not change the number of hours worked, Smith (2000) finds that, when wages are directly reported, wage change distributions in the UK exhibit spikes similar to those found in the US. Instead, when the sample is restricted to individuals showing their payslips, the spike at zero disappears. Therefore, reporting errors explain almost the totality of wage freezes in the UK. Also, comparing the percentage of wage cuts for the two samples, she finds that there is no evidence of perfect downward wage rigidity in reporting-error free data, although the percentage of cuts is lower than in reported data. In Smith (2002), a similar validation study is carried out specifically on the percentage of cuts, distinguishing between nominal and real cuts and different reasons why employees may accept wage cuts. Only 7% of nominal and 9% of real cuts are validated (instead of 28% nominal and 41% real cuts from reported data) which are not due to overtime, bonuses, and hours changes. Variables of satisfaction, human capital, and negative productivity shocks explain the probability of receiving a pay cut. Nickell and Quintini (2003), and Barwell and Schweitzer (2004) analyse the UK administrative data-base: the 1975-2000 New Earnings Survey (NES).

The quality of information about wages in the NES is very high for studying wage rigidity, as the hourly base wage is reported separately from the other components of earnings. Differently from what found in Smith (2000), spikes at zero are evident for almost every year wage change. The extent of the spike varies over time and is negatively related to the rate of inflation. Quite interestingly, Barwell and Schweitzer (2004) shed light on another spike in the nearby of the rate of inflation, therefore rising a certain interest in introducing measures of not only nominal but also real wage rigidity.

Goux (1997) compares the percentage of wage cuts in the 1990-1996 French Labour Force Survey (The Enquete Emploi) and the administrative data Déclarations Annuelles de Données Sociales (DADS) during the 1976-1992 time period. A negative relationship between the rate of inflation and the percentage of cuts is found. Having observed that for the overlapping years the percentage of cuts is roughly the same in the two data, although the measure for wages is not exactly the same, she uses the information available on job characteristics in the EE for explaining how employees can be induced to accept wage cuts. In particular, among full-time workers with pay cuts and without firm change: 34% have better working conditions with respect, for example, to night work; 22% face a decrease in their annual bonuses; 30% change 4-digit occupation; and more than 60% are in one, at least, of these three situations.

The Italian case is considered in Dessy (1998). In the Bank of Italy bi-annual Survey, during the time-period 1989 to 1995, extremely high spikes are observed at zero nominal wage changes both for stayers and for movers. Considering that the spikes at zero normally decrease when the time length of wage variation increases, this result shows much more rigidity in Italy than in most other countries. At the same time if, as in the UK case, the spike at zero is due to rounding, the very high spikes found both for stayers and movers can be considered as reporting error affecting the two categories of workers in exactly the same extent. Actually Devicienti (2002) finds no spike in stayers daily earnings from the administrative Istituto Nazionale della Previdenza Sociale (INPS) data.

The usual shape of wage change distributions from survey data are found by Borgjis (2001) in the Belgian 1993-1998 Panelstudie van Belgische Huishoudens (PSBH). Spikes at zero nominal gross wage change are on average 10% of observations whereas the percentage of cuts is 23%.

For Germany, there is evidence available only from administrative data: the 1975 to 1995 version of the IAB-Beshaftigtenstichprobe (IABS) reporting annual total compensation and no information on hours. Beissinger and Knoppik (2001) find that nominal wage change distributions exhibit a shape similar to those found for the UK by Barwell and Schweitzer (2004), i.e. it is a double-spiked distribution at zero and in the nearby of the rate of inflation. Wage cuts are not rare. Although not stressed in the papers, the existence of rounding phenomena in reporting data is documented in the IAB.

Fehr and Goette (2003) consider the Swiss case. Although they do not focus on the impact of the quality of data on wage change distributions, they compare the Swiss 1991-1998 Labour Force data with administrative files. The measure for wages is hourly wages for SLFD and total year compensation in the SIF sample, therefore the two distributions are not directly comparable. However, it is clear that administrative data are much less dispersed than survey data, i.e. the distribution of wage changes is more centred around zero than the SLFD. Both the distributions are asymmetric around zero, however in administrative data the asymmetry is much more pronounced. There is a striking discontinuity around zero and the pile-up of observations just above zero is very pronounced in SIF data.

The measurement of wage rigidity with micro-data has been recently the object of analysis of an international project, the International Wage Flexibility Project (IWFP) lead by W. Dickens (the Brookings Institution) and E. Groshen (The Federal Reserve Bank of New York). They have contacted most of the European centres where it is possible to have access to administrative data and have carried out parallel analyses of wage change distributions. The countries involved are: Germany, Italy, Sweden, US, Switzerland, Norway, Finland, Denmark, Belgium, France, Austria, UK, and Portugal. The purpose of the project is to derive a method to be applied in all countries for constructing simultaneous measures of nominal, real and institutional wage rigidity and therefore exploring different causes of wage rigidity. Many attempts have been carried out, based on both the location and the structural approach, but to date unfortunately none of them turned out to be completely satisfactory. However, the simple descriptive analysis has revealed that, on administrative data, the shape of the distributions is different across countries: not always the spike at zero is observed and the asymmetry can be more or less pronounced around zero. It seems quite clear that the position of wage change distributions follows the rate of inflation, therefore having a certain impact on the percentage of wage cuts observed. Strong measurement error problems arise however for intercountry comparisons purposes in the IWFP, since the unit of measure used for wages varies across countries. In particular, in many countries the number of hours is not observed.

Conclusions on descriptive analyses

Summarising, descriptive evidence on wage change distributions varies across countries and, for the same country, depends on the unit of measure defined, the length of time considered, and the quality of data used. However, the following stylised facts can be deduced from the analysis of nominal wage change distributions:

1. A spike at zero nominal wage changes is always present in individual survey data. The extent of the spike is lower in administrative files, and in some countries is even inexistent. The fact that it can be due to rounding error rises concern on relying on it as evidence of nominal wage rigidity, unless the quality of data is excellent.

2. In some EU countries it is argued that there is a spike also at the rate of inflation. This would be evidence of a certain stability of wages also in real terms.

3. At the same time, wages are not completely downwardly rigid. The percentage of wage cuts both from survey and administrative data is always significantly different from zero.

4. Although there are drops of the distribution of wage changes around zero, small wage rises are usually more frequent than small wage cuts. There seems therefore to be no evidence of menu-costs effects.

According to the above stylised facts we could accept, although not at the same extent, both theories of nominal and real wage rigidity. Complete nominal downward wage rigidity, however, is not observed in the data. Among the theories of nominal wage rigidity, menu-costs models are not supported by the empirical evidence of wage change distributions. However, we have to bear in mind that descriptive results are usually not referred to the base hourly wage contracted, therefore they might be biased by reporting and measurement errors.

2.2.2 Measures of location

Even though in the simple descriptive approach no distributional assumptions are made on the shape of the counterfactual distribution, it can be argued that, in absence of rigidities, not necessarily wage change distributions have to be smooth. In particular, the shape or degree of asymmetry of some distribution in themselves might not reveal much about the presence of nominal rigidity, since they may be characteristics of a particular counterfactual distribution. Similar considerations might hold for the spike at zero and the share of negative observations. Therefore in the literature it has been introduced an alternative method, still completely non-parametric, in which wage rigidity measures are calculated on the basis of the location of wage change distributions, and the relationship between different parts of the distribution. Basic contributions in this approach are: the histogram-location approach by Kahn(1997), and the symmetry approach by Card and Hyslop (1997). These approaches have been connected in a common analytical framework by Beissinger and Knoppik (2001), with some extensions.

Kahn (1997) models factual and counterfactual distributions as median-centred histograms. The whole distribution is divided in a number of bars of equal length, and therefore of constant distance, from the median. Assuming that variations in the shape of the counterfactual distributions can be caused only by shifts of the counterfactual distribution over the course of years[4], the bars constructed above can obviously shift over time. For example, the bar that is distant 3 percentage points left of the median might contain each year with a different probability the zero spike, or nominal wage cuts, or small wage cuts. Kahn therefore regresses each bar for each year on bar dummies

[4]This makes only sense if the shape of the distributions does not vary for other reasons.

for zero, for negative nominal change, and for 1 percent above and below zero. Hence she will be able to capture in some non-parametric form the importance of the various places in the distribution of interest (such as the spike at zero nominal wage change, the bars that surround the spike bar, the various bars strictly below zero nominal change). Her results are the following: 1) there is a large coefficient on zero nominal wage change that reflects the spike, and this is interpreted as evidence of nominal wage rigidity; 2) there are sizeable and negative coefficients on 1 percent dummies above and below zero nominal wage change which are consistent with menu-costs theories; 3) there is a large and negative coefficient on the negative dummy for wage earners (hourly pay) that reflects downward nominal stickiness; 4) however there is a positive coefficient on the negative dummy for salary earners, reflecting that pay changes are more likely if they entail a pay cut. Interestingly, she claims that this result is not due to changing usual hours.

Dealing with a complete model of the distributions and not only with indicators has the advantage of being able to assess the extent of rigidity and not only to observe its existence. The limit of Khan's analysis is that no assessment on the magnitude of measurement errors is incorporated in the analysis, and extending her approach on this direction is very difficult. The issue of incorporating measurement errors in a non-parametric framework is addressed by Card and Hyslop (1997).

Card and Hyslop (1997) follow another identification strategy than the histogram-location approach. They introduce in the literature the so-called symmetry-approach, based on the following assumptions: 1) in the absence of rigidities the distribution would be symmetrical around the median; 2) the upper-half of the distribution is unaffected by rigidities; 3) wage rigidities do not affect employment. The approach is therefore still non-parametric, but the counterfactual can change its shape over time, not only the position. Measurement errors are taken into account simply correcting the observed PSID values on the basis of the validation study results in Card (1996).The counterfactual can be constructed simply replicating on the left of the median the right part of the observed distribution. Comparing the actual and counterfactual distributions on the left of the median allows to construct some measures of wage rigidity both in terms of number of persons affected by wage rigidity and in terms of wage changes (i.e. those which, in the absence of rigidity, would have been different). They find that: a) the number of persons affected by such nominal wage rigidity amounts to 8 to 12 percent in the mid '80s; and b) the effects of such nominal wage rigidity on wage changes are such that wage changes have been 1 percent higher every year than they would have been in absence of rigidity during the same time period.

Knoppik and Beissinger (2001) relax the time-invariance assumption on the shape of the counterfactual distribution in the histogram-location approach à la Kahn. This extension is justified with the fact that the data they use are total labour earnings, and therefore can be subject to hours variability over the business cycle. Also, survey studies on firm wage policies report that some wage cuts do occur, but only under decidedly adverse market conditions, that vary over the business cycle. The results confirm the existence of strong downward wage rigidity in Germany.

Conclusions on measures of location

Considering the whole distribution of wage changes, and how different parts of this distribution are related, allows to construct more precise measures of wage rigidity than the simple estimation of the observed frequencies of wage cuts or freezes. However, some more restrictive assumptions have to be introduced on the counterfactual distribution. The measures of location presented in this section are all based on a non-parametric approach, and therefore do not rely on any specific assumption on the shape of the counterfactual.

The limit of the location-approach is that measurement errors can not be taken into account easily. Unless it is possible to correct the observed distribution on the basis of appropriate validation studies, or very good quality data are available, applying this methodology might not be completely satisfactory.

2.2.3 The structural approach

Taking into account formally of measurement error issues implies introducing assumptions on their distribution. This brings up a fully parametric approach for the specification of the counterfactual distribution. Usually measurement errors are modelled according to the classical assumptions, therefore the normality of wage change distributions is introduced. In the literature, this issue is taken into account in a so-called structural framework, introduced by Altonji and Devereux (2000). An appropriate econometric model is estimated, based on the MacLeod and Malcomson (MM) (1993) hold-up model, the only micro-economic foundation for nominal downward wage rigidity, in which the outside options, and therefore the counterfactual or notional distribution is estimated on the basis of individual observable characteristics of both firms and employees. The presence of measurement errors is formally taken into account, so that one can separate true wage changes from wage changes that merely reflect reporting errors or reductions in actual hours worked.

The Altonji and Devereux (AD) econometric model

First of all, AD distinguish between *notional* wage w_{it}^*, an optimal wage that the firm would like to implement if there were no downward rigidity in period t but possibly taking account of the fact that the wage chosen today will constrain later wage choices, and the *actual* wage w_{it}^0 that the firm actually implements at time t.

They model the change in the *actual* wage as a function of the change in the *notional* wage as follows:

$$w_{it}^0 - w_{it-1}^0 = \begin{cases} w_{it}^* - w_{it-1}^0 & if \quad 0 \leq w_{it}^* - w_{it-1}^0 \\ 0 & if \quad -\alpha < w_{it}^* - w_{it-1}^0 < 0 \\ \lambda + w_{it}^* - w_{it-1}^0 & if \quad w_{it}^* - w_{it-1}^0 \leq -\alpha \end{cases} \qquad (1)$$

Basically, the wage coincides with the notional wage, if the notional wage change implies a wage increase. If the notional wage change is a nominal wage cut of less than α, the model specifies that the actual wage change is zero. Nominal wage cuts occur when the notional wage change is sufficiently negative. α and λ are positive constants, to be estimated. The parameter λ is a positive constant that determines the response of wage changes to the notional wage change when a cut is appropriate. When $w_{it}^* - w_{it-1}^0 = -\alpha$, $w_{it}^0 - w_{it-1}^0 = \lambda - \alpha$.

Both w_{it}^* and w_{it}^0 are in logs. The log of the notional wage w_{it}^* is a function of a vector of explanatory variables x_{it}, a parameter vector β and a normally distributed error term ε_{it}:

$w_{it}^* = x_{it}\beta + \varepsilon_{it}$

Substituting in (1),we obtain:

$$w_{it}^0 - w_{it-1}^0 = \begin{cases} x_{it}\beta + \varepsilon_{it} - w_{it-1}^0 & if \quad 0 \leq x_{it}\beta + \varepsilon_{it} - w_{it-1}^0 \\ 0 & if \quad -\alpha \leq x_{it}\beta + \varepsilon_{it} - w_{it-1}^0 \leq 0 \\ \lambda + x_{it}\beta + \varepsilon_{it} - w_{it-1}^0 & if \quad x_{it}\beta + \varepsilon_{it} - w_{it-1}^0 \leq -\alpha \end{cases} \tag{2}$$

or, equivalently:

$$w_{it}^0 = \begin{cases} x_{it}\beta + \varepsilon_{it} & if \quad w_{it-1}^0 \leq x_{it}\beta + \varepsilon_{it} \\ w_{it-1}^0 & if \quad x_{it}\beta + \varepsilon_{it} \leq w_{it-1}^0 \leq \alpha + x_{it}\beta + \varepsilon_{it} \\ \lambda + x_{it}\beta + \varepsilon_{it} & if \quad \alpha + x_{it}\beta + \varepsilon_{it} \leq w_{it-1}^0 \end{cases} \tag{3}$$

It is important to keep in mind that the value of β is influenced by whether employers take into account the possibility of being constrained by downward rigidity in the future, when setting current wages. The value of β in a labour market characterized by nominal wage rigidity is likely to differ from the value of β when wages are perfectly flexible.

The wage model above contains as special cases both a model of perfect wage flexibility and a model of perfect downward nominal wage rigidity. The model of perfect wage flexibility is obtained with the following restrictions on the parameters: $\alpha = 0$ and $\lambda = 0$. For perfect downward wage rigidity model instead λ is arbitrary and α approaches ∞. Because both these models are nested in the general model, one can test whether the restrictions implied by either perfect rigidity or perfect flexibility are consistent with the data.

AD show that their general model also nests MM's model of wage contracts. They call $\underline{w}(t)$ the worker's outside option, and $\overline{w}(t)$ the firm's outside option. They consider the fixed wage contract with renegotiation discussed in section 2.1 above, that can be put in their framework as follows:

$$w_{it}^0 = \begin{cases} \underline{w}(t) & if \quad w_{it-1}^0 \le \underline{w}(t) \\ w_{it-1}^0 & if \quad \underline{w}(t) \le w_{it-1}^0 \le \overline{w}(t) \\ \overline{w}(t) & if \quad \overline{w}(t) \le w_{it-1}^0 \end{cases} \tag{4}$$

where obviously $w_{it-1}^0 = w^c$, the wage contracted. One can specify functional forms for $\underline{w}(t)$ and $\overline{w}(t)$, in terms of the x variables and regression error and so end up with a nominal rigidity model that is quite similar to the empirical model (3). In particular, choosing the following parametrisation of the MM model:

$$\begin{aligned} \underline{w}(t) &= x_{it}\beta + \varepsilon_{it} \\ \overline{w}(t) &= \alpha + x_{it}\beta + \varepsilon_{it} \quad where \quad \alpha > 0 \end{aligned} \tag{5}$$

AD get an expression very similar to (3):

$$w_{it}^0 = \begin{cases} x_{it}\beta + \varepsilon_{it} & if \quad w_{it-1}^0 \le x_{it}\beta + \varepsilon_{it} \\ w_{it-1}^0 & if \quad x_{it}\beta + \varepsilon_{it} \le w_{it-1}^0 \le \alpha + x_{it}\beta + \varepsilon_{it} \\ \alpha + x_{it}\beta + \varepsilon_{it} & if \quad \alpha + x_{it}\beta + \varepsilon_{it} \le w_{it-1}^0 \end{cases} \tag{6}$$

that is a special case of model (2), where $\lambda = \alpha$. Thus, the intuitive model can be seen to encompass the MM model as a special case. This is neither a perfect flexible model nor a perfect downward rigidity model: it is just one possible model of nominal wage rigidity.

Because wages are reported with error, we need to parametrise the *reported* wage w_{it} as a function of the true wage and the measurement error component u_{it}. AD use the following:

$$w_{it} = w_{it}^0 + u_{it}$$

Substituting this equality into model (2), we obtain the model expressed in terms of the *reported* wage rather than the true wage:

$$w_{it}-w_{it-1} = \begin{cases} x_{it}\beta + \varepsilon_{it} - w_{it-1} + u_{it} & if \quad 0 \leq x_{it}\beta + \varepsilon_{it} - w_{it-1} + u_{it-1} \\ u_{it} - u_{it-1} & if \quad -\alpha \leq x_{it}\beta + \varepsilon_{it} - w_{it-1} + u_{it-1} \leq 0 \\ \lambda + x_{it}\beta + \varepsilon_{it} - w_{it-1} + u_{it} & if \quad x_{it}\beta + \varepsilon_{it} - w_{it-1} + u_{it-1} \leq -\alpha \end{cases} \tag{7}$$

AD use alternative specifications for the distribution of u_{it}, which turn out to be crucial for their results because sometimes they change according to the hypotheses about the distribution of u_{it}.

The major problem posed for the estimation of this model is unobserved heterogeneity. Because unobserved ability is correlated with w_{it-1}, $Cov\left(w_{it-1}, \varepsilon_{it}\right) > 0$. Hence w_{it-1} cannot be treated as a predetermined variable in estimating the model. AD deal with this problem in two ways.

Method 1 (M1). They replace w_{it-1} with its conditional expectation given lagged values of x. More precisely, they approximate w_{it-1}^0 by modelling its expectation $\widehat{w}_{it-1}$ to be a linear function of x_{it-1}, x_{it-2}, and x_{it-3}. Substituting in model (7):

$$w_{it-1} = w_{it-1}^0 + u_{it-1} = \widehat{w}_{it-1} + \mu_{it-1} + u_{it-1}$$

After this substitution, the model that AD estimate is:

$$w_{it}-w_{it-1} = \begin{cases} x_{it}\beta + e_{it} - \widehat{w}_{it-1} + m_{it} & if \quad 0 \leq x_{it}\beta + e_{it} - \widehat{w}_{it-1} \\ m_{it} & if \quad -\alpha \leq x_{it}\beta + e_{it} - \widehat{w}_{it-1} \leq 0 \\ \lambda + x_{it}\beta + e_{it} - \widehat{w}_{it-1} + m_{it} & if \quad x_{it}\beta + e_{it} - \widehat{w}_{it-1} \leq -\alpha \end{cases} \tag{8}$$

where $m_{it} = u_{it} - u_{it-1}$ and $e_{it} = \epsilon_{it} - \mu_{it}$. It is assumed that m_{it} and e_{it} are normally distributed mean-zero random variables.

Model (8) is estimated by maximum likelihood.

Method 2 (M2). The model is estimated under the assumption that w_{it-1} is approximately equal to w^*_{it-1}. In this case the equation $w_{it-1} = X_{it-1}\beta + \varepsilon_{it-1} + u_{it-1}$ can be used to eliminate w_{it-1} from the model.

On a priori grounds, M1 is preferred to M2, even though the results are given for both the methods and sometimes they are completely different.

Results: First of all, measurement error seems to explain almost all the wage cuts observed. AD's results are in-line with validation studies carried out on the same data (PSID).

Likelihood ratio tests of the flexible wage model, reject the model of perfect flexibility versus the general model. Also, the perfect downward rigidity model is rejected against the general model. But the likelihood of the downward rigidity model is higher than for the perfectly flexible model, suggesting that it is a better description of reality.

AD also estimate the probability of wage cuts, wage rises and freezes but because of the variation of the estimates according to the methods used they do not draw strong conclusions about the size of the effect of downward rigidities on average wage changes of stayers.

Even though the analysis is conducted on the sample of stayers (and therefore can be affected by sample selection bias), the interest in measuring the extent of nominal wage rigidity rests on the impact of such rigidity on unemployment and therefore on quits, layoffs and promotions. AD deal with this problem simply estimating separate linear probability models for the probability of a layoff, a quit and a promotion. But again, the magnitude of their estimates is very sensitive to the choice between M1 and M2, therefore they are not able to say much about the above transitions.

Fehr and Goette (FG) version of the AD model

Fehr and Goette (FG) (2003) use a simplified version of the AD model, where $\lambda = 0$ and w_{it-1} is eliminated from the model (Method 2 of AD). The fact that $\lambda = 0$ does not allow to test the MM model, based on the assumption that $\lambda = \alpha$. Eliminating w_{it-1} is not as restrictive as it might seem at first instance, but certainly allows to estimate only a reduced form of the MM model. The advantage is that the FG specification is easier to estimate than the original AD version, and gives more robust results.

Whereas in their original model FG estimated a common α for all individuals, in the most recent version they have individual-time varying thresholds α_{it}. The model estimated is the following:

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

This model is similar to, but more general than, the Altonji and Devereux (2000) model because individual heterogeneity is taken into account whereas Altonji and Devereux (2000) impose the restriction that the threshold is the same for all workers. This restriction counterfactually implies that there are no wage change observations in the interval $[-\alpha_{it}, 0]$. By allowing for individual heterogeneity in wage cut thresholds some workers may have flexible wages while others have rigid wages. The main focus of the analysis is to estimate the mean μ_c and the variance σ_c of the distribution of thresholds. Workers with negative thresholds exhibit perfectly flexible wages. This model nests the AD model as a special case. In fact, if the variance of α_{it} goes to zero, the two models become identical. In addition to allowing for individual heterogeneity, they also allow for nonzero correlation between the error term e_{it} and the individual thresholds α_{it}, estimating the value of this correlation. This would capture the fact that in periods of firms financial distress individuals are more likely to accept wage cuts. Since changes in productivity are probably an important component of e_{it}, a positive correlation is expected between α_{it} and e_{it}. It is assumed that in every year a fraction p (that is estimated) of the individual data has no measurement error, but that the rest of the sample draws a normally distributed error. Measurement error is interpreted differently according to the data used: in the SLFS the fraction p of individuals states the correct income, whereas in the SIF sample the fraction p of individuals has no variation in hours. The outside option is estimated with variables x_{it} that capture business cycle variation in wages, and individual characteristics correlated with wage growth. Variables that capture business cycle variation are regional unemployment and year fixed effects. Variables that systematically affect wage growth across workers are labour market experience, age, tenure, and observable skills of workers. As an additional control, also the firm size is included.

Results: Analogously to the AD conclusions, FG find that downward wage rigidity is a persistent phenomenon in Switzerland, also in periods of low inflation. Their results are more robust than in the AD paper, and are valid in all the data sources considered. Moreover, as in AD measurement error explains almost all wage cuts observed.

Beissinger and Knoppik's model

Beissinger and Knoppik (BK) (2003) introduce the so called *proportional* *model* of downward wage rigidity. As in the AD model, there is an underlying notional or counterfactual wage change for individual i at time t, that can be explained as a set of variables arranged in a vector x_{it}:

$$\Delta y_{it}^{*} = x_{it}\beta + e_{it}, e_{it} \sim N\left(0, \sigma_e^2\right)$$

The actual wage change Δy_{it} is equal to the notional wage change, except in cases where the latter is negative and the person is affected by downward nominal wage rigidity. Whether this is the case is indicated by a random variable D_{it} which takes on the value one with probability ρ if there is downward nominal wage rigidity, and zero otherwise. Actual wage changes are therefore modelled as follows:

$$\Delta y_{it} = \begin{cases} x_{it}\beta + e_{it} & if \quad 0 \leq x_{it}\beta + e_{it} \\ 0 & if \quad x_{it}\beta + e_{it} < 0 \wedge D_{it} = 1 \quad where \quad \Pr\left(D_{it} = 1\right) = \rho \\ x_{it}\beta + e_{it} & if \quad x_{it}\beta + e_{it} < 0 \wedge D_{it} = 0 \end{cases}$$

This is therefore a model with proportional downward wage rigidity since the proportion ρ of notional wage cuts will be prevented by rigidity. The degree of rigidity in the model is captured by the parameter ρ, that has the advantage of estimating directly the proportion of wage cuts prevented by rigidity. In the AD and FG model instead, the sweep-ups were calculated after the estimation of the parameter α. This is a clear advantage of the BK specification of the model, which also: 1) takes account of the observation that cuts do happen, which would not be the case with the tobit model; 2) allows for small and moderate size cuts to happen, which is not the case in the AD and FG model, since in that case cuts occur below the threshold α.

At the same time, the BK approach presents the following limits: 1)

ρ is the same for all individuals, therefore differently from the FG model individual heterogeneity is not considered; 2) the fact that the parameter α cannot be estimated is a limit in terms of a rigorous structural approach, therefore the link with the theoretical MM model becomes weak and we can talk in this case only of a reduced form analysis; 3) nominal downward wage rigidity implies only the occurrence of freezes, and not of cuts, whereas in the AD and FG model even in a downward rigidity regime cuts could happen below the threshold α.

Since the BK model is estimated on German administrative data, in which yearly wages are given and no information is available on the number of hours, measurement error has to be taken into account. The model becomes:

$$\Delta y_{it} = \begin{cases} x_{it}\beta + e_{it} + m_{it} & if \quad 0 \leq x_{it}\beta + e_{it} \\ m_{it} & if \quad x_{it}\beta + e_{it} < 0 \wedge D_{it} = 1 \quad where \quad \Pr(D_{it} = 1) = \rho \\ x_{it}\beta + e_{it} + m_{it} & if \quad x_{it}\beta + e_{it} < 0 \wedge D_{it} = 0 \end{cases}$$

KB propose three alternative specifications for the error term m_{it}: 1) normal measurement error (NME); 2) mixed measurement error (MME); 3) contaminated mixed measurement error (CMME). In all estimated model variants KB obtain a high and highly significant value of ρ. For workers this varies between 0.46 and 0.72 and for salaried employees between 0.58 and 0.91. The results indicate therefore the presence of substantial downward nominal wage rigidity in Germany.

The same model has been estimated on Italian administrative data by Devicienti (2002), finding values of the parameter ρ between 0.51 and 0.68. Considering that the Italian wage change distributions do not exhibit any spike, this means that the majority of the observed cuts actually are measurement error. This conclusion is quite worrying in the Italian case, since the spike is generated, rather than explained, by the model.

Models of nominal, real and institutional wage rigidity

One of the attempts carried out by the IWFP has been to extend the AD model in order to take into account at the same time of a nominal and a real wage rigidity threshold. The likelihood in this case becomes quite complicated, details can be found in Dickens and Goette (2002). Although the approach has been abandoned because it was quite unsatisfactory for inter-country comparisons purposes, the German, the Italian and the English teams have continued to work at the model. Bauer, Bonin and Sunde (2003), as well as Devicienti, Maida and Sestito (2993), and Barwell and Schweitzer (2004) find a substantial extent of real wage rigidity with respect to nominal wage rigidity on administrative data.

Conclusions on the structural approach

The advantage of the structural approach is to estimate the counterfactual distribution, instead of just making assumptions on it. However, if measurement errors are taken into account, the normality assumption is introduced and we end up in a fully parametric approach. A general limit of this approach is that often, especially when working with administrative data, not many variables are available for predicting the counterfactual distribution behaviour. In general, the estimation results show that the extent of downward nominal wage rigidity is quite high in all the countries examined. Usually measurement errors explain the wage cuts observed. This implies that the spike observed at zero underestimates the percentage of rigid wages: all the wage cuts observed, when corrected for measurement error, become rigid nominal wages.

2.3 Conclusions

In this chapter we have surveyed the existing evidence on wage rigidity in a number of countries. Three methodologies have been considered: 1) the simple descriptive approach; 2) the non-parametric approaches; and 3) the structural approaches. The main problems are: a) how to treat measurement errors; and b) the assumptions on the counterfactual distribution.

The non-parametric approach is advised only on measurement-error free data, or when it is possible to correct precisely for measurement error with validation studies. Otherwise, it is better to model wage change dynamics in a structural framework, introducing the classical assumptions on measurement errors. However it is quite common, in this case, to explain almost the totality of wage cuts observed as measurement error.

Due to the contrasting results depending on the quality of data available, the methodology adopted, and the country considered, it is clear that further research is needed on the issue of measuring the extent of wage rigidity. Useful contributions can go in the following directions: collect new data, possibly of good quality, for extending the evidence available on wage change distributions especially in the European countries; explore more in detail causes and consequences of different types of wage rigidity; introduce new methods for defining and estimating wage rigidity.

3 Observed and estimated measures of nominal wage rigidity in the EU countries

The analysis of wage change distributions for determining the extent of nominal wage rigidity has been typically carried out for different countries separately. As discussed in the survey chapter, 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 wage rigidity with the ECHP data particularly interesting for policy purposes, as the phenomenon of downward nominal wage rigidity can induce real effects when the level of inflation is low.

The first purpose of this thesis is therefore to consider cross-country comparisons of wage rigidity measures for the EU countries following a systematic approach. The analysis of the first 7 waves (1994-2000) of the ECHP for measuring nominal wage rigidity covers two chapters and considers two different issues. In this chapter we describe the data, with particular attention to the information given about wages and hours worked, and the impact of measurement error on wage change distributions. Our purpose is to estimate the extent of nominal wage rigidity in the EU countries, carrying out inter-country comparisons. For this purpose we adopt the structural approach à la Altonji and Devereux (AD) (2000). In chapter 4 instead we try to explore the causes of downward wage rigidity in Europe using institutional, country-specific characteristics.

Considering that the principal purpose of this chapter is to introduce inter-country comparisons, the ECHP has the great advantage of covering fifteen EU countries: Germany, France, UK, Italy, Spain, Netherlands, Belgium, Luxembourg, Ireland, Denmark, Sweden, Finland, Austria, Portugal and Greece. However, the information on wages from individual surveys is not always of very good quality. As with most of the individual surveys available, the ECHP 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[5]. We therefore follow the structural approach, in which measurement error is modelled according to the classical assumptions. In particular, we use a simplified version of the AD model, very similar to the one considered by Fehr and Goette (2002), for calculating nominal wage rigidity measures comparable across countries.

The structure of the chapter is the following: in section 3.1 we give some information on the ECHP. Section 3.2 introduces wage distributions for all countries, whereas Section 3.3 considers wage change distributions, presenting the frequencies of nominal wage cuts and freezes observed in the data. Section 3.4 deals with the estimation of measures of wage rigidity. Section 3.5 concludes.

[5] Hanish and Rendtel (2003) consider rounding in the German and Finnish ECHP wage data. Peracchi and Nicoletti (2003) analyse the distortive impact of imputation and non-response on income data in the ECHP.

3.1 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 Table 3.

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[6]. 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 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) is known, it is also possible to calculate "hourly current earnings" dividing monthly wages by the number

[6] 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.

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[7]. 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 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

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

people whose interview was taken in no more than two months difference in the two years period considered, reduces dramatically the number of observations[8]. In qualitative terms, however, our results do not seem to change significantly. In order not to lose too many observations, we therefore decided to keep in the sample people who reported their wages in different months for consecutive interviews.

3.2 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 Figure 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 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:

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

they all present spikes at rounded values. In all the European countries the percentage of rounded wages[9] 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)[10], 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) make 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. Since an overall validation study of

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

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

the ECHP is out of the scope of this thesis, we will assume that rounding errors are normally distributed in all countries[11]. The further concern raised from the German and Finnish validation studies regards the correctness of inter-country comparisons: if rounding behaviour depends on the number of digits, and therefore on the currency of the country, it might be that also measures of wage rigidity are affected by the currency of the country. This would make direct inter-country comparisons not feasible. We therefore consider each country separately in this chapter. Measurement error, when not modelled according to the classical assumption, will be taken into account in a meta-analysis framework of the next chapter, 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 to researchers[12]. 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,

[11]We analyse this issue in detail for France in chapter 5.

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

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 this problem by eliminating 1% of observations in both tails of wage change distributions[13]. 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 Figure 1, dividing monthly wages by the number of hours does not cancel out completely the existence of many spikes in wage distributions[14]. 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 sometimes[15] they can be explained with particular contractual arrangements. We therefore keep all the observations in our sample.

In Appendix 1 (Tables A1.1 - A1.3) we give descriptive statistics of gross wage distributions, number of weekly hours and hourly wages for every year

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

[14]In Figure 1 censorig has been used in the upper tail.

[15]See the analysis carried out for France in Chapter 5.

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 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.

3.3 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 four 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.

Figure 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 thesis 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 4, which gives wave by wave the precise figures for the percentage of cuts, freezes and rises in monthly wages, hours and hourly wages[16].

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 while keeping their total labour earnings constant, or not letting them fall 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 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.

[16] Detailed descriptive statistics for the same distributions can be found in Appendix 1 (Tables 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 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 uses a wage variable which eliminates imputed and - calculated from - net values - and the gross wage is constructed from raw data. This makes a big difference, by eliminating some classical measurement error. 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 chapter.

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. Rankings 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[17].

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 we have seen in the previous section that, 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.

[17]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, but this issue was not pursued to avoid extending the discussion too far.

3.4 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 first attempt of estimating the extent of nominal wage rigidity in the EU countries has been an implementation of the original version of the Altonji and Devereux (AD) model presented in chapter2. Method 2 has been followed for treating past values of earnings. Although the likelihood estimated converges in all countries, we have found values for the parameter λ systematically bigger than 1, i.e. bigger than α since $0 < \alpha < 1$[18]. This result is difficult to interpret in terms of measures of wage rigidity as it would imply wage cuts higher than 100%. Introducing the restriction $\lambda < 1$ is not technically easy. Since, in order to nest the MacLeod and Malcomson (1993) hold-up model, the restriction $\alpha = \lambda$ was required, we can say that this was certainly not the case in our replication of the AD model. Therefore, our AD results could not be interpreted in terms of the MM model. For this reason we decided to move to an easier specification 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.

Another justification for abandoning 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

[18]Results are available under request.

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 we decided to 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 are very significant. Unfortunately in the ECHP we found it 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, also experience can not be included in the x_{it} vector. We use worker's age as a proxy for experience. 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 5. 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.

Table 6 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 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.

3.5 Conclusions

In this chapter we have analysed wage dynamics at the individual level using the 1994-2000 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[19]. 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.

[19]The alternative approach, based on the BK model, has not been implemented. The reason is that the BK model is estimated on administrative data, therefore results are not directly comparable with our data.

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%.

The limit of our procedure for carrying out inter-country comparisons rests on the specification of the econometric model. Unfortunately the observed variables on which we are conditioning our analysis are only sex, and age. The biggest problem with this approach is to find variables useful for explaining wage dynamics. In fact, the variables normally used for wage equations are not significant in wage change equations.

Also, the assumptions on measurement errors, on which inter-country comparisons are based, are quite strong. Probably more complex specifications for measurement errors are could be tried.

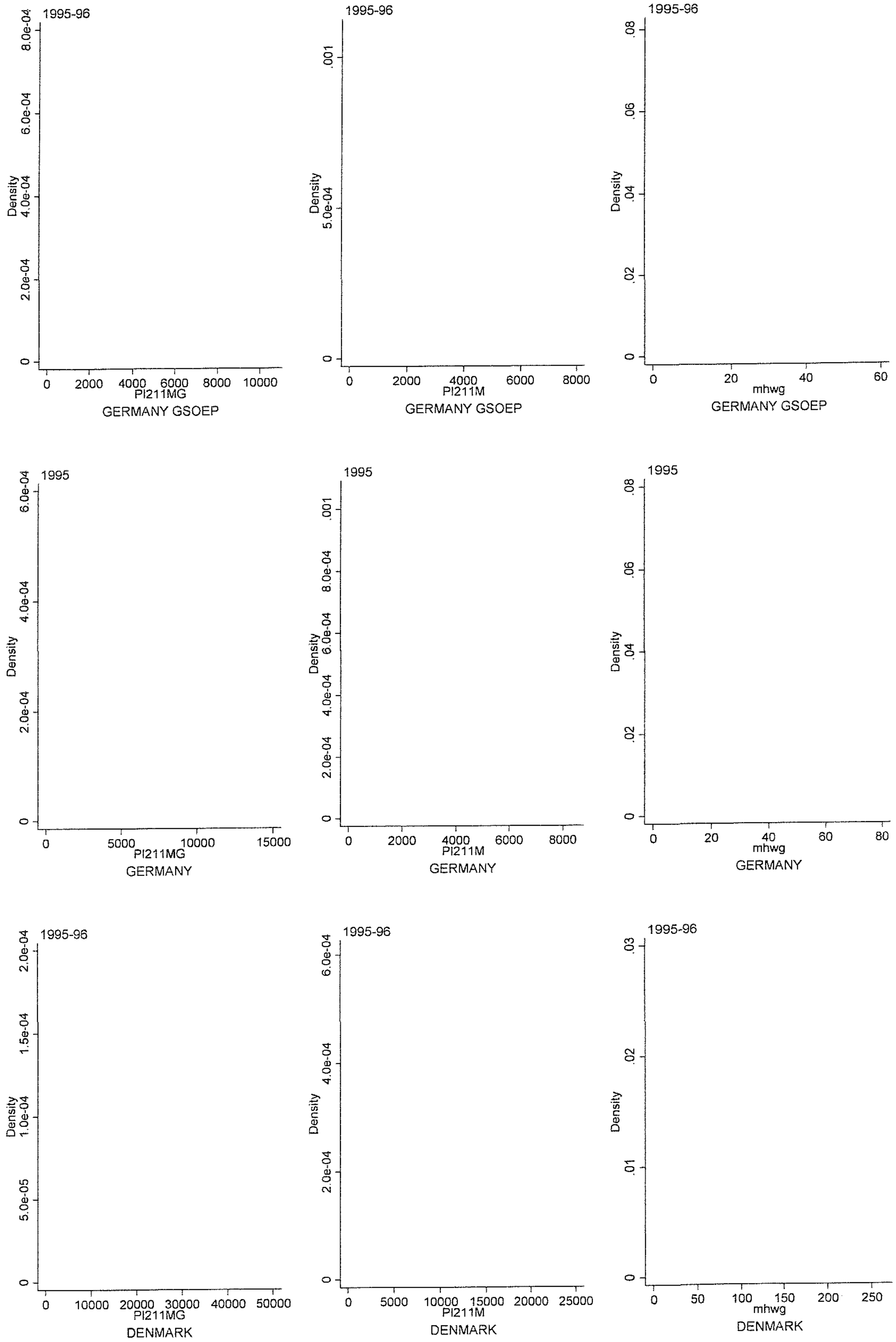
Although the analysis carried out in this chapter was focused on measuring nominal wage rigidity in the EU countries, further investigation is needed to explore the causes and the consequences of nominal wage rigidity in Europe. In the next chapter we consider the institutionalist explanation for nominal wage rigidity, often proposed given the pervasive role of unions and labour market institutions in wage determination in Europe.

Table 3 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

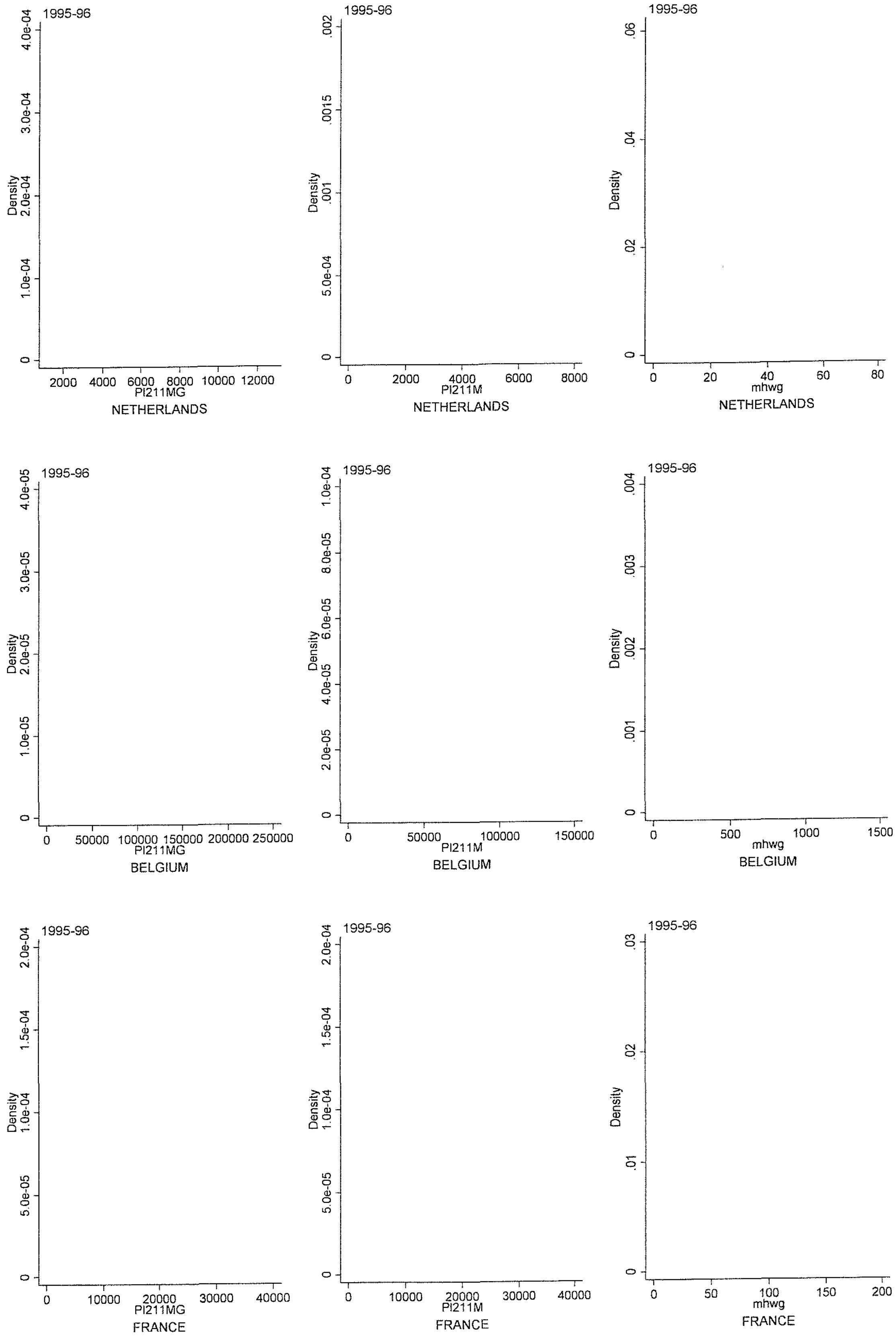
1. Only Net wage available; Sweden excluded.

Figure 1 : Gross, Net and Gross Hourly Wage Distributions in the ECHP for Stayers-full time



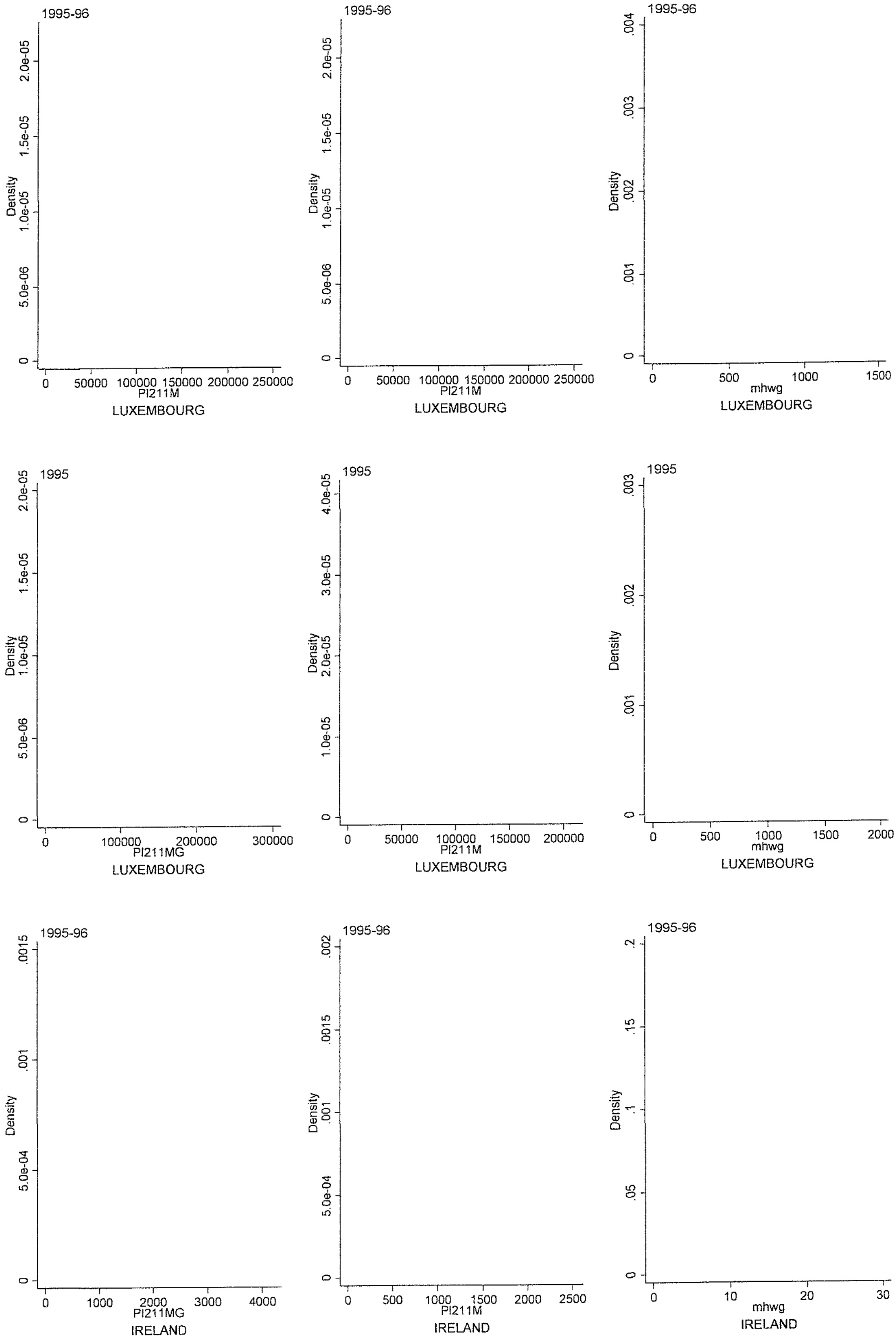
Source: ECHP.
Sample: Stayers full-time.

Figure 1 continued



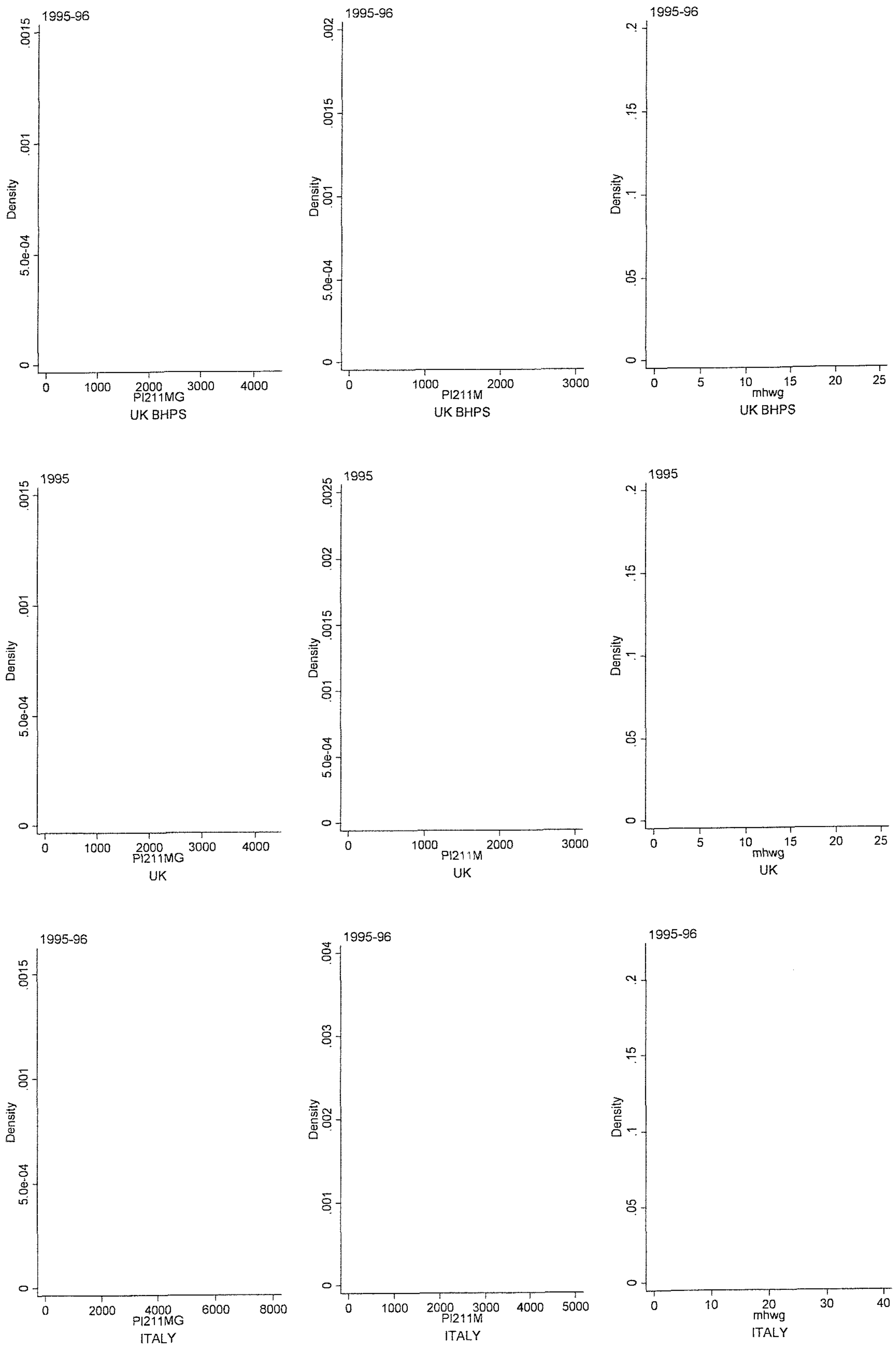
Source: ECHP.
Sample: Stayers full-time.

Figure 1 continued



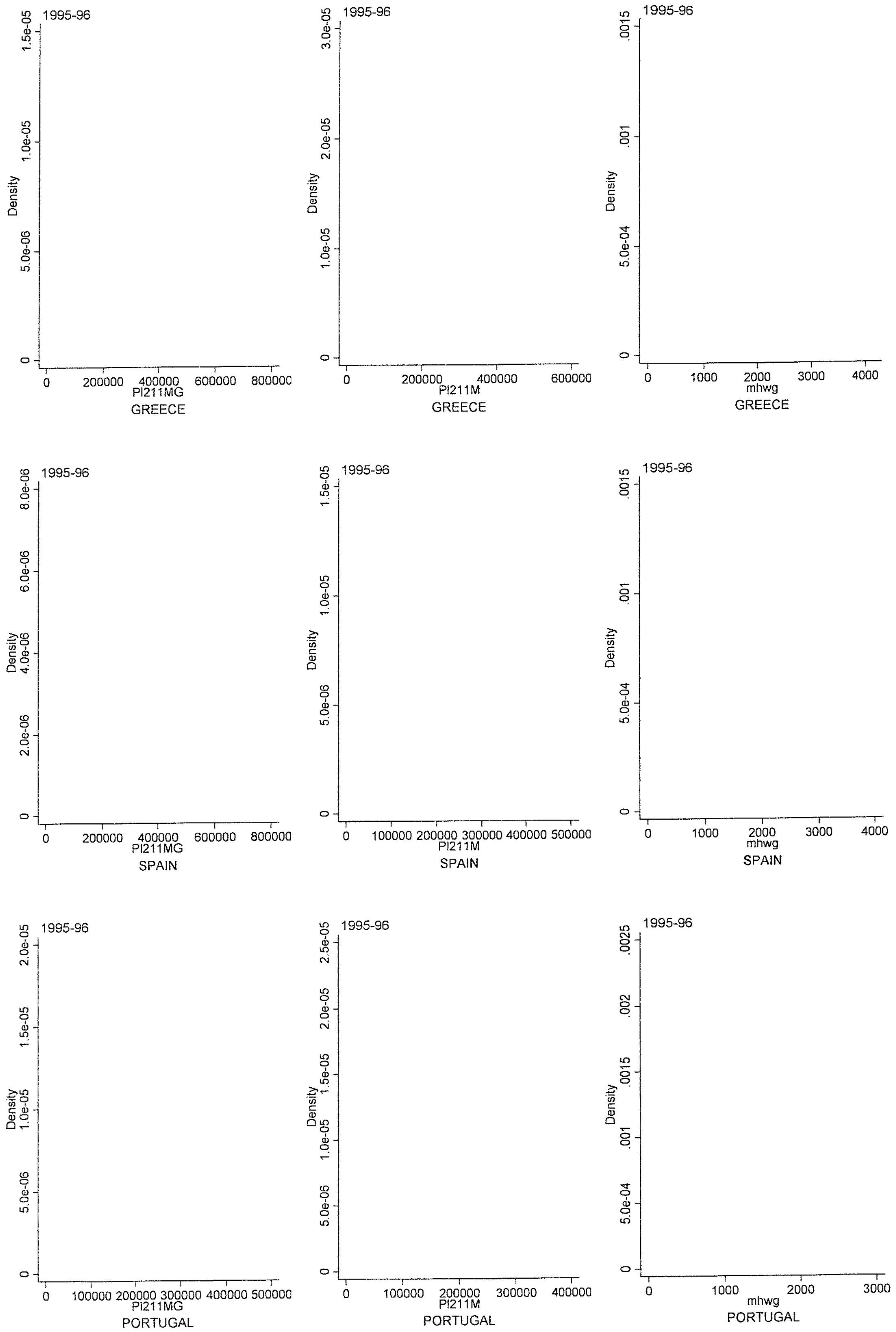
Source: ECHP.
Sample: Stayers full-time.

Figure 1 continued



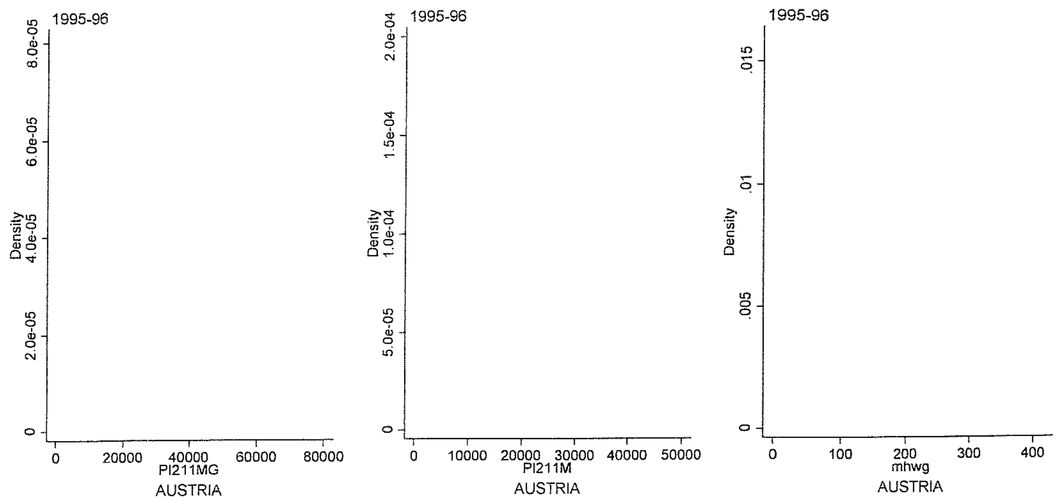
Source: ECHP.
Sample: Stayers full-time.

Figure 1 continued



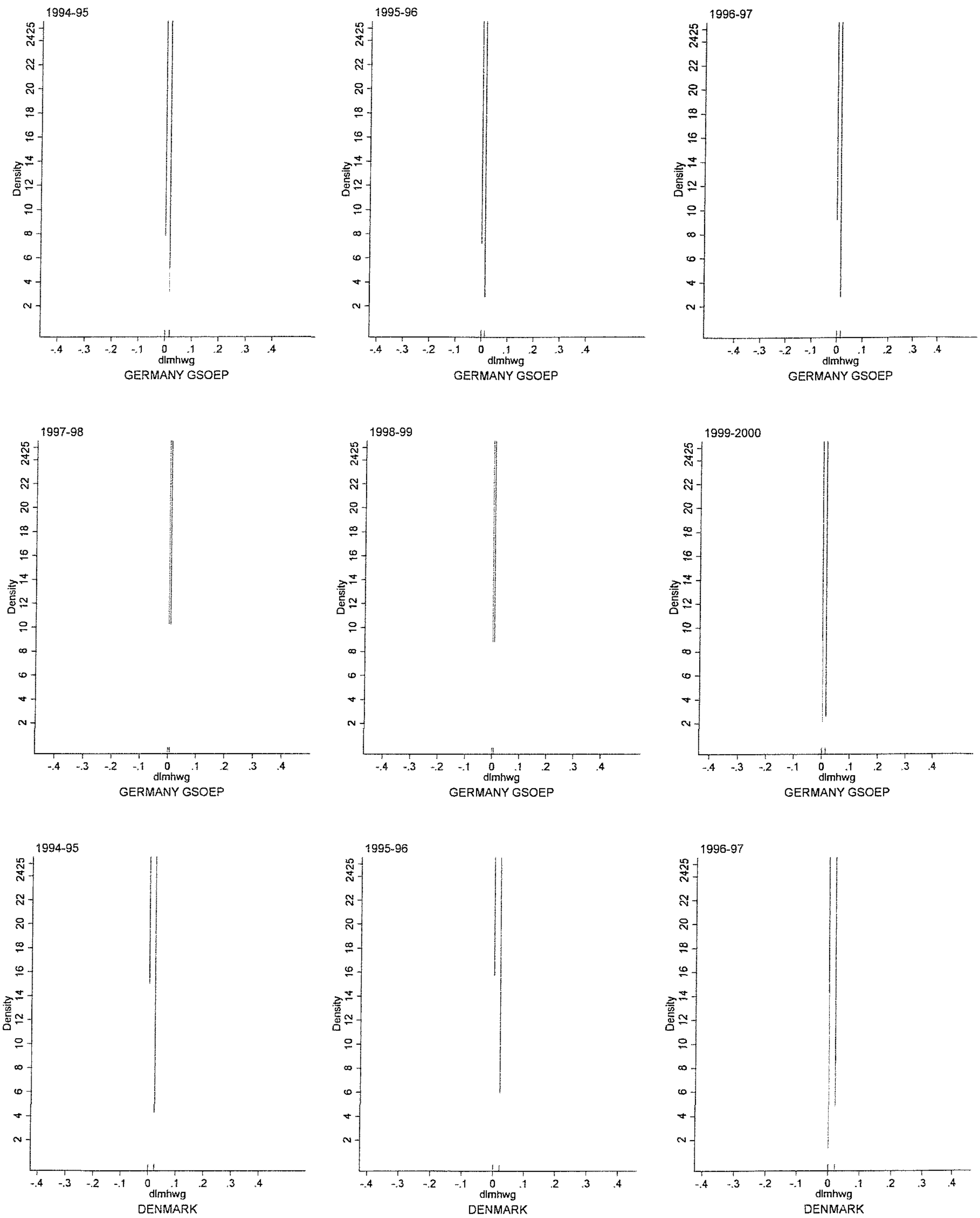
Source: ECHP.
Sample: Stayers full-time.

Figure 1 continued



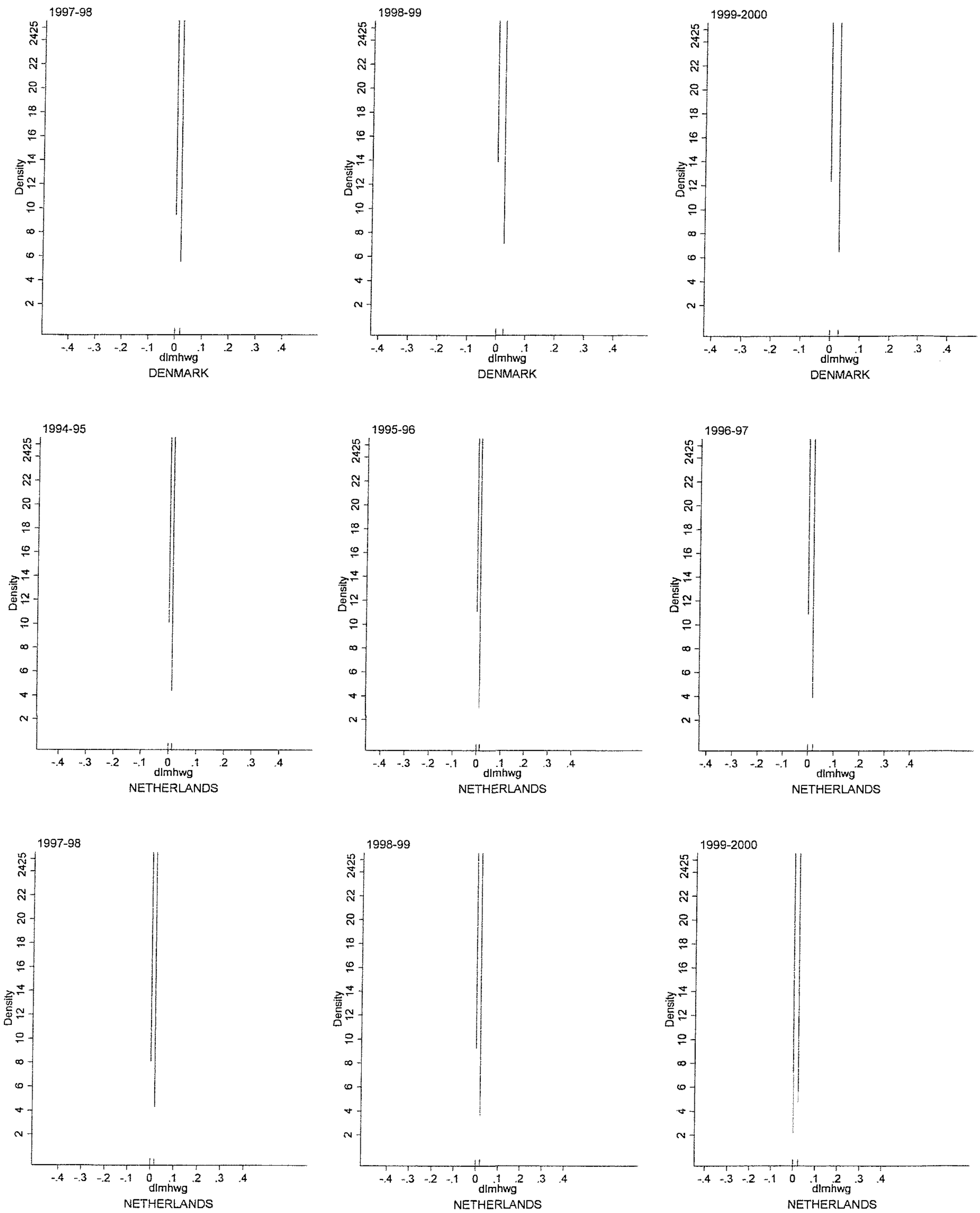
Source: ECHP.
Sample: Stayers full-time.

Figure 2: Gross Hourly Wage Change Distributions in the ECHP



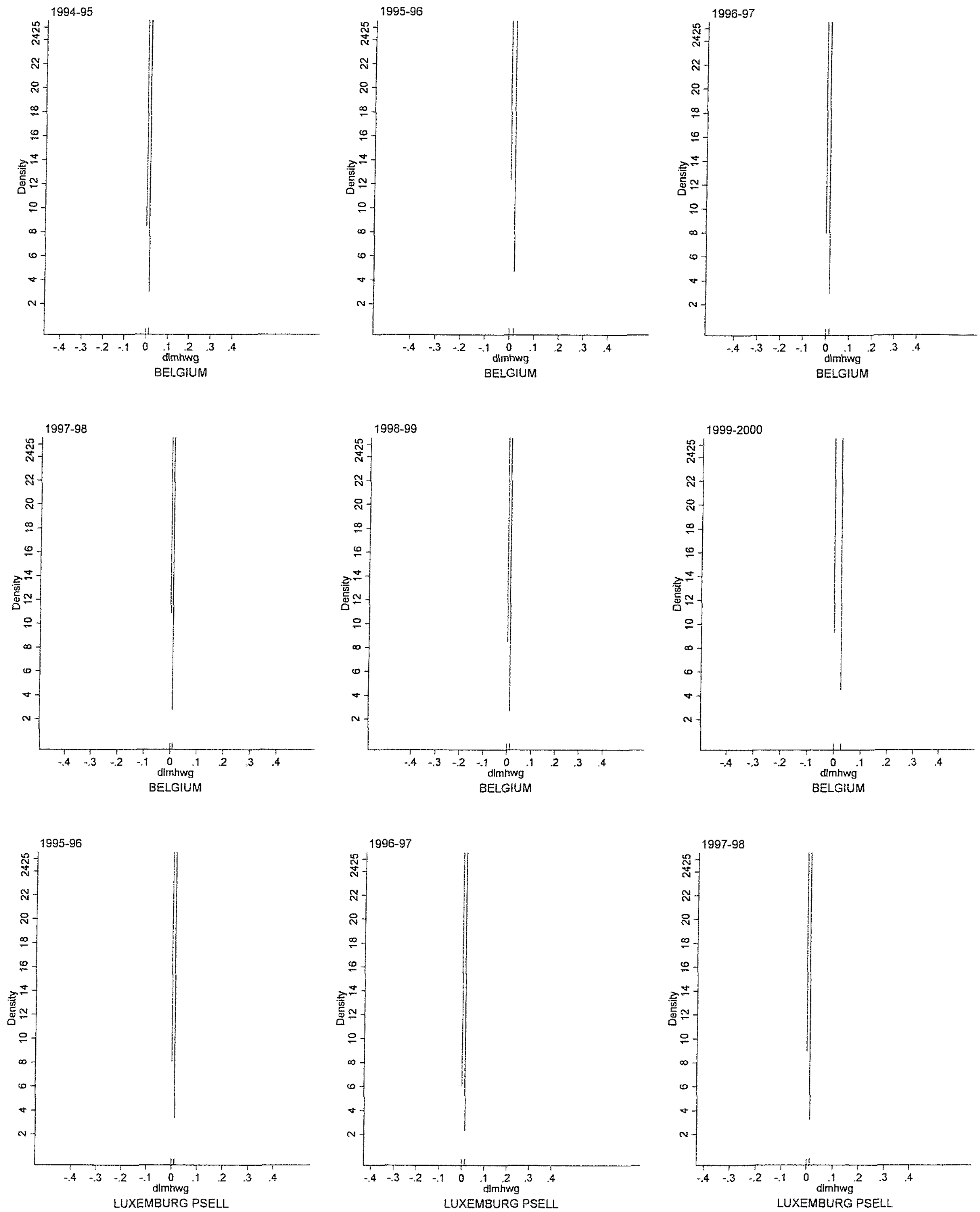
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



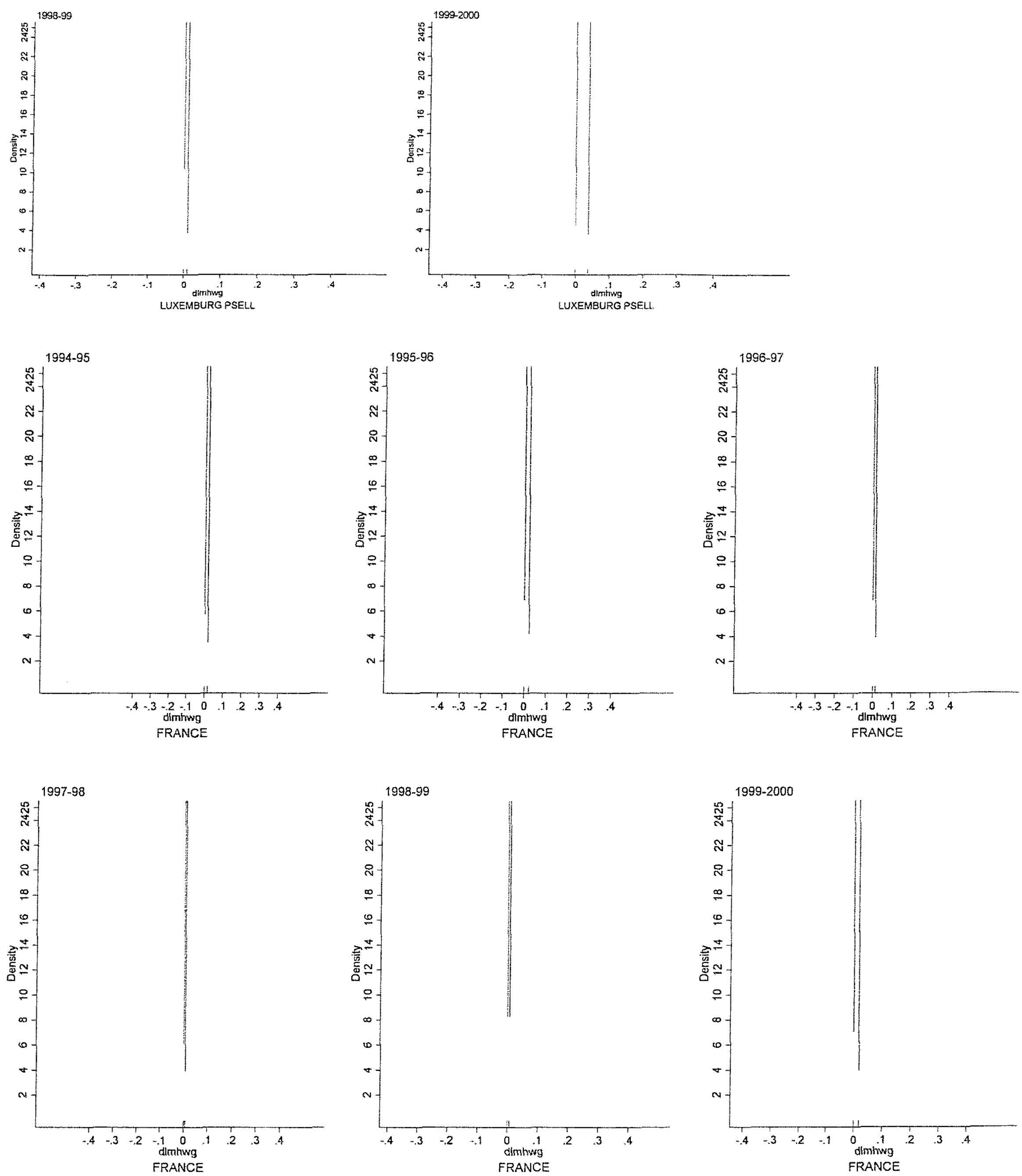
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



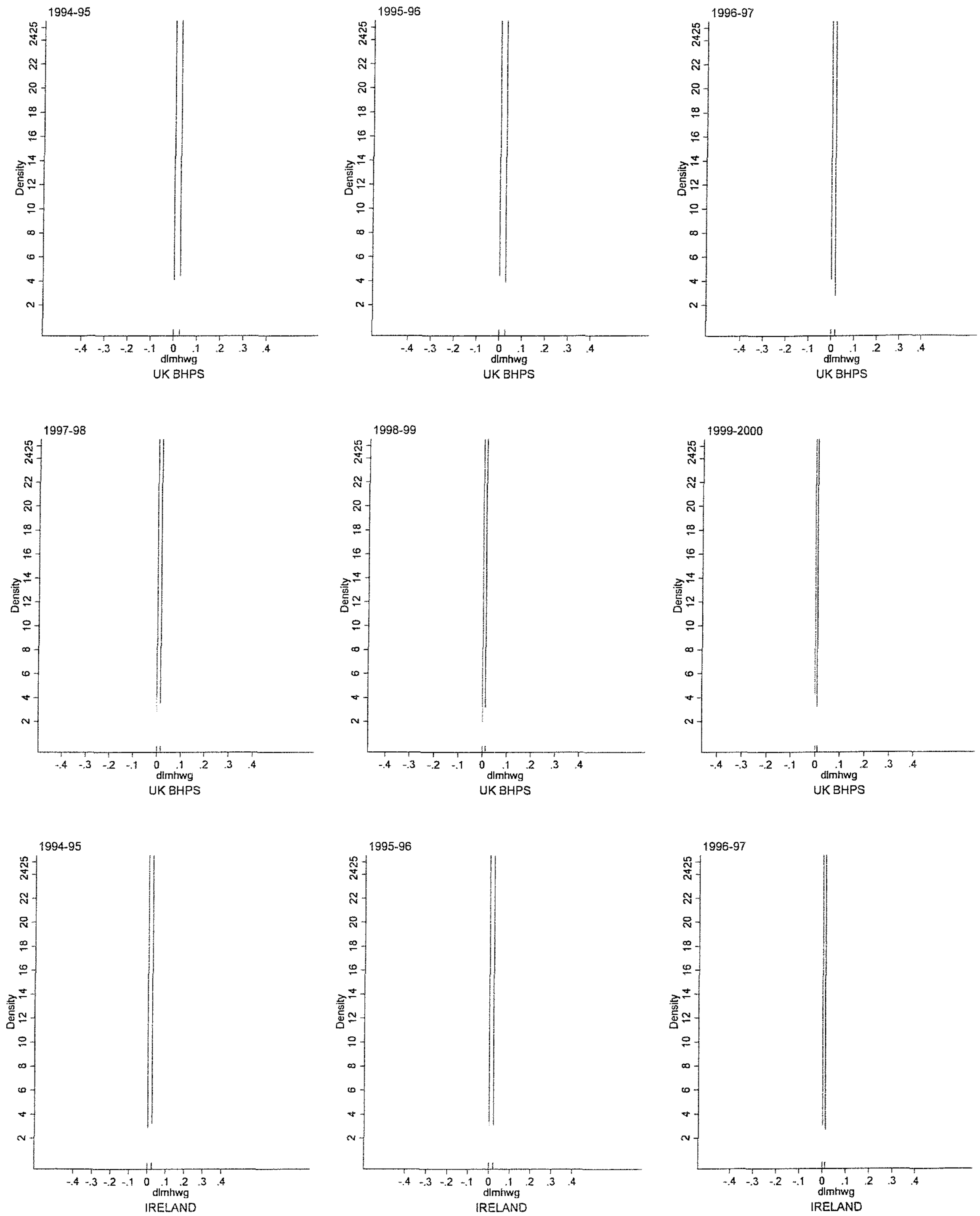
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



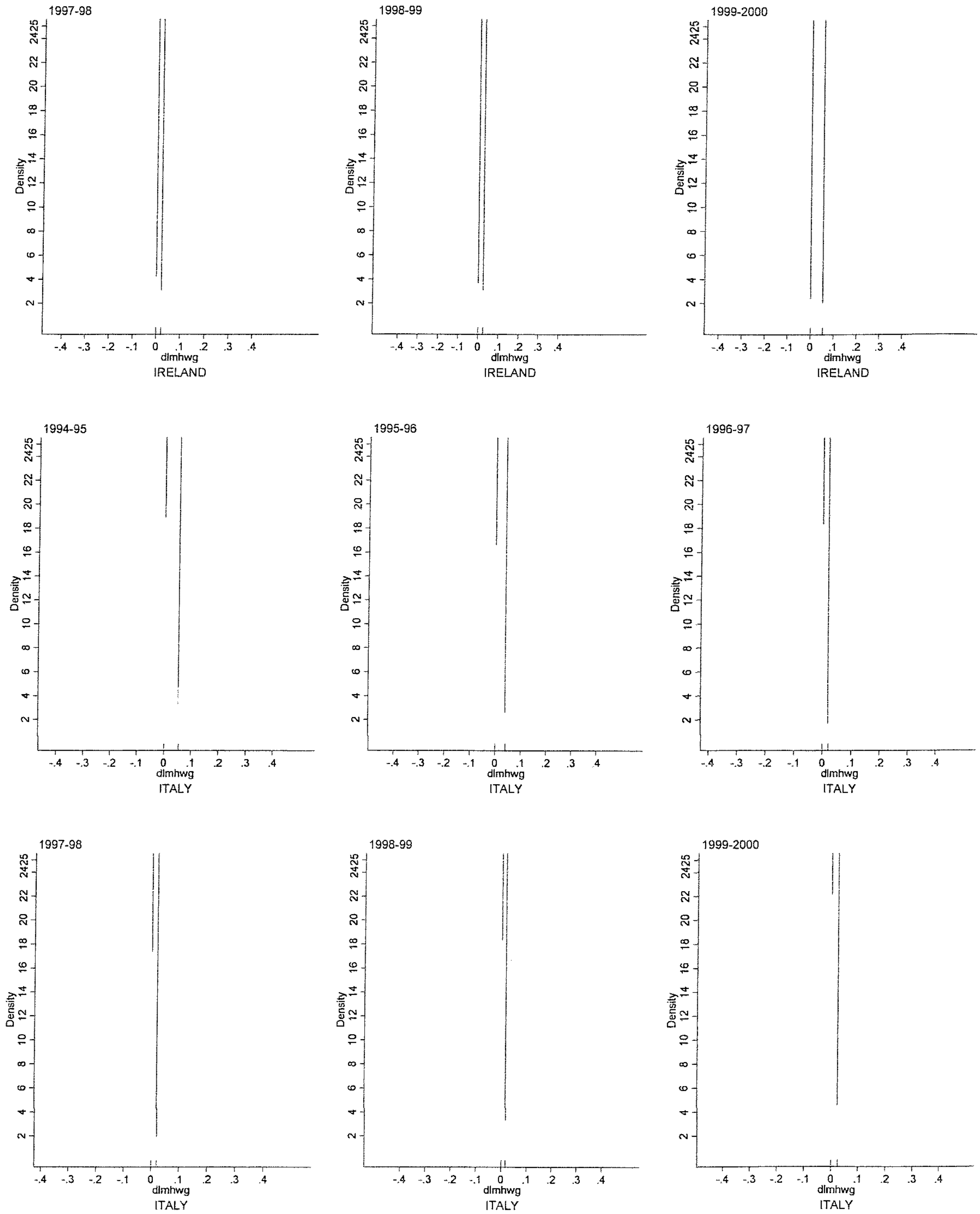
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



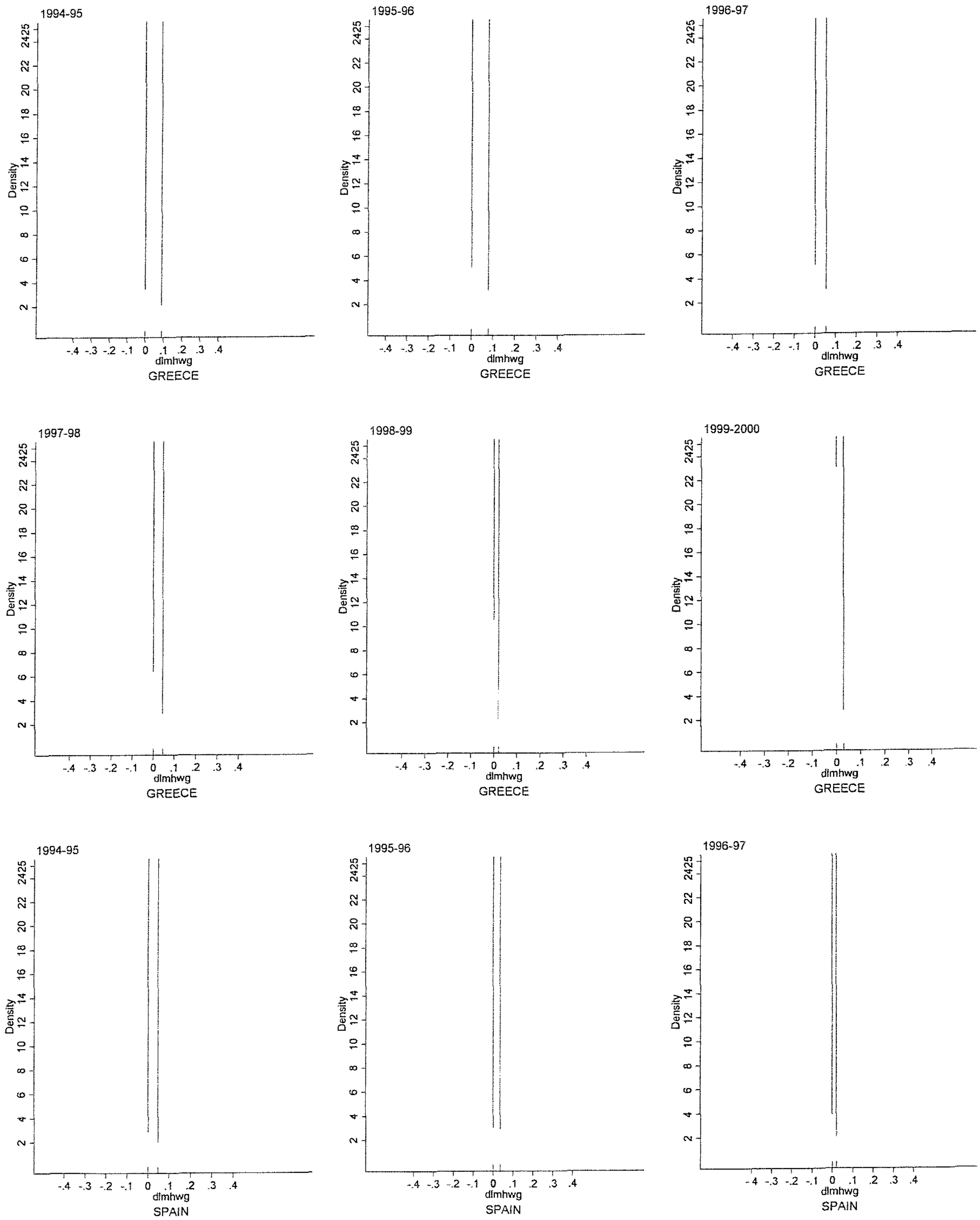
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



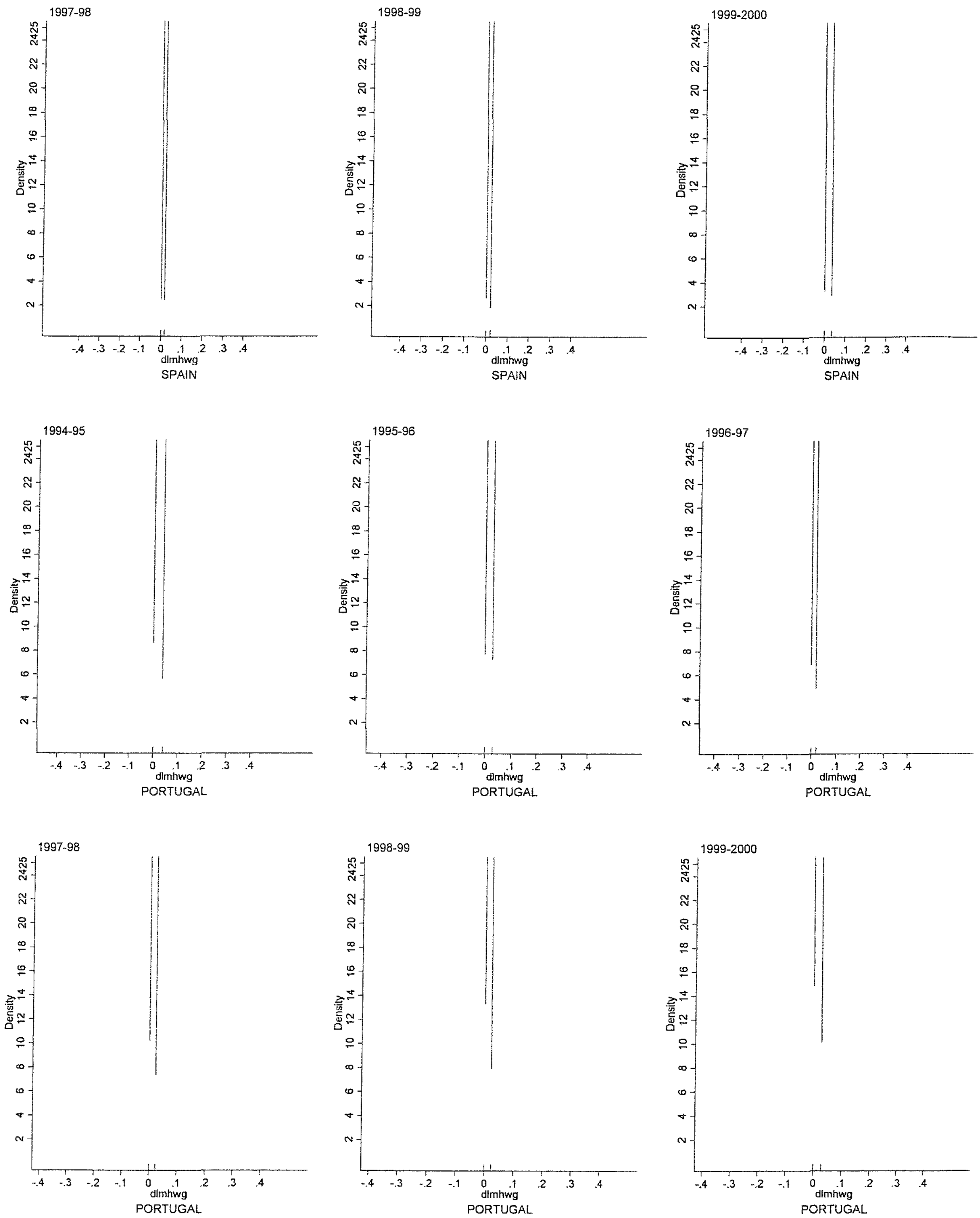
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



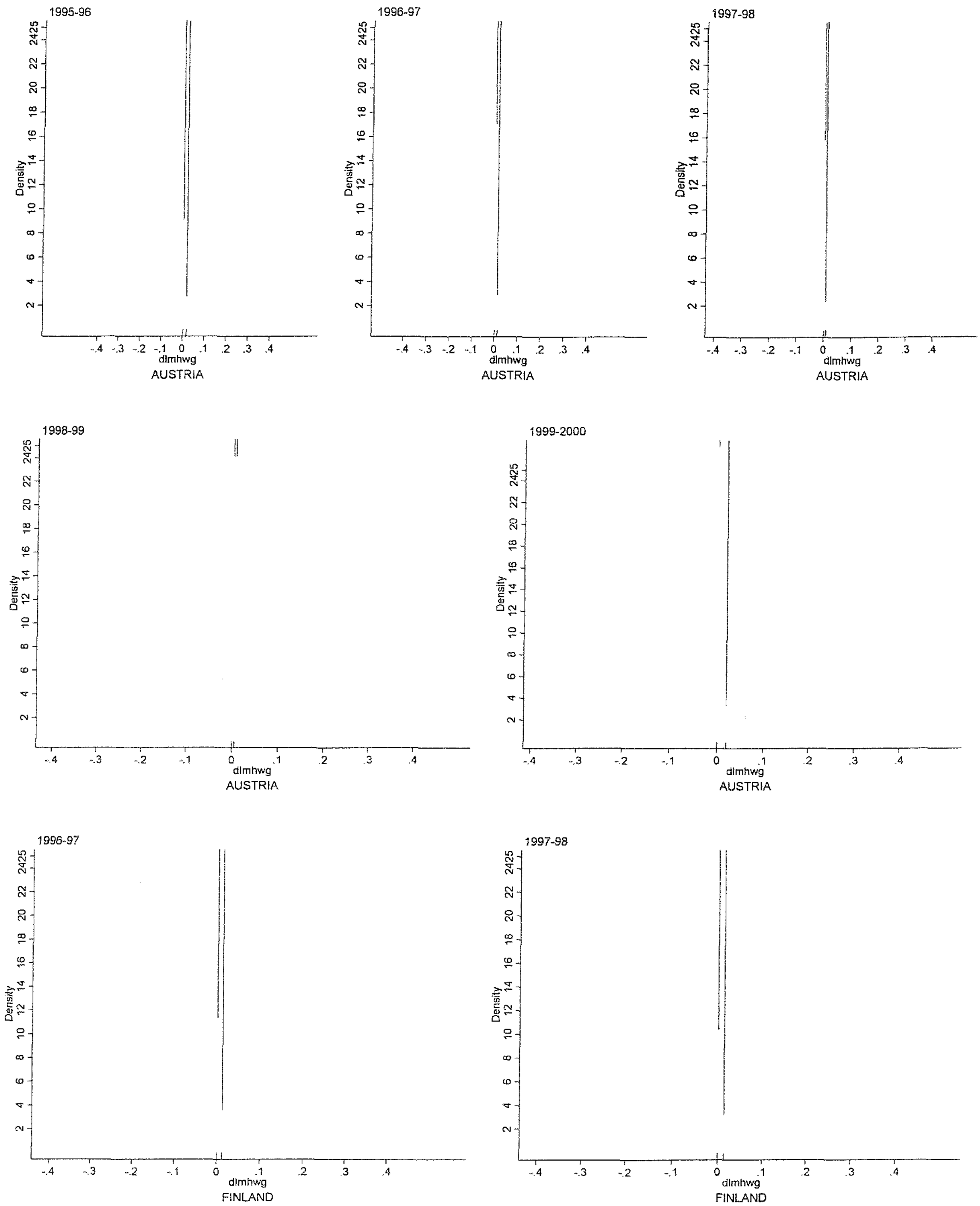
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



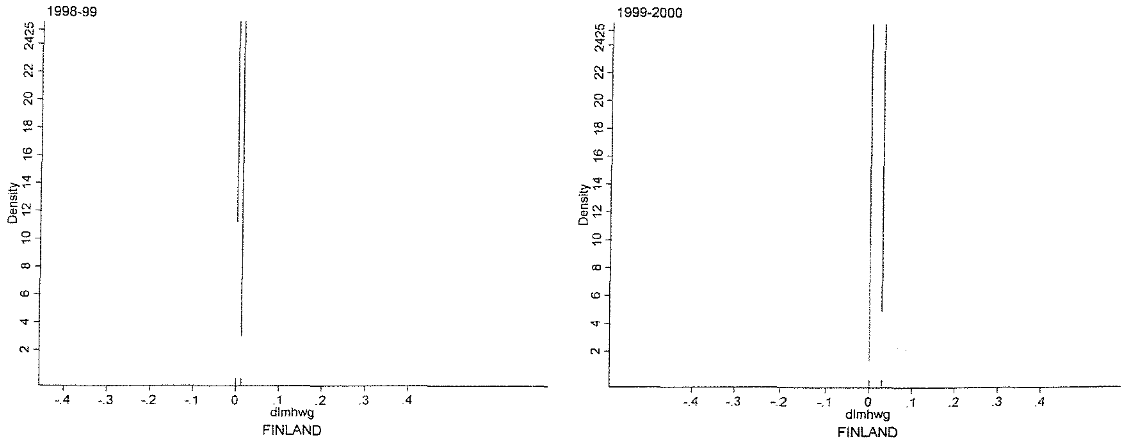
Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



Source: ECHP.
Sample: Stayers full-time.

Figure 2 continued



Source: ECHP.
Sample: Stayers full-time.

Table 4: 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

Source: ECHP
Sample: Stayers full-time

Table 4 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

Source: ECHP
Sample: Stayers full-time

Table 4 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

Source: ECHP
Sample: Stayers full-time

Table 5: Estimates of the econometric model

GERMANY GSOEP

log pseudo-likelihood= 10361.972

Number of obs = 17885.000
Wald chi2(7) = 188.240
Prob > chi2 = 0.000

log pseudo-likelihood= 361.972

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Con	Interval]
beta						
age	-0.001	0.000	-6.720	0.000	-0.001	-0.001
female	0.000	0.003	-0.080	0.939	-0.005	0.005
time3	0.023	0.004	5.670	0.000	0.015	0.031
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
se						
_cons	0.137	0.001	117.590	0.000	0.135	0.140
sm						
_cons	0.073	0.002	36.980	0.000	0.069	0.077
alpha						
_cons	0.177	0.007	26.670	0.000	0.164	0.190

GERMANY ECHP

log pseudo-likelihood= 2632.336

Number of obs = 5072.000
Wald chi2(3) = 19.670
Prob > chi2 = 0.000

log pseudo-likelihood= 632.336

dlmhwg	Coef.	Robust Std. Err.	z	P>z	[95% Conf.	Interval]
beta						
age	-0.001	0.000	-2.670	0.008	-0.001	0.000
female	-0.002	0.005	-0.400	0.686	-0.012	0.008
time3	-0.016	0.005	-3.460	0.001	-0.025	-0.007
_cons	0.071	0.011	6.550	0.000	0.050	0.092
se						
_cons	0.153	0.002	66.480	0.000	0.148	0.157
sm						
_cons	0.049	0.007	6.950	0.000	0.035	0.063
alpha						
_cons	0.128	0.017	7.480	0.000	0.094	0.161

Table 5 continued

NETHERLANDS

log pseudo-likelihood= 7515.987

				Numbe	r 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% Con	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

				Numbe	r 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

Table 5 continued

LUXEMBOURG 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

Table 5 continued

UK BHPS

log pseudo-likelihood= 3850.800

				Numbe	r 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

				Numbe	r 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

Table 5 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

Table 5 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

Table 5 continued

GREECE

1687.747

	Numbe	r of obs =	8103.000
	Wald	chi2(7) =	315.460
log pseudo-likelihood= 687.747	Prob	> chi2 =	0.000

dlmhwg	Coef.	Robust 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

Table 5 continued

PORTUGAL

log pseudo-likelihood= 2.831

				Numbe	r 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

				Numbe	r 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

Table 5 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

Table 6: Cuts and freezes observed and estimated in the ECHP

Country	year	cuts estimated	cuts observed	freezes estimated	freezes observed
GERMGSOEP	1995	8.865311	32.39	38.56437	5.41
GERMGSOEP	1996	6.46703	28.04	34.36423	5.38
GERMGSOEP	1997	11.6132	38.33	41.98494	6.86
GERMGSOEP	1998	10.96486	36.53	41.28718	7.18
GERMGSOEP	1999	10.88069	36.92	41.19454	6.74
GERMGSOEP	2000	8.839809	33.11	38.54275	6.46
DENMARK	1995	6.36667	22.4	33.60175	11.5
DENMARK	1996	8.12021	23.46	36.84304	12.29
DENMARK	1997	7.28543	21.75	35.39126	9.63
DENMARK	1998	7.92445	24.74	36.51619	7.84
DENMARK	1999	8.256849	23.48	37.05092	9.88
DENMARK	2000	8.07065	25.3	36.75442	9.75
NETHERLANDS	1995	7.13821	28.7	43.65794	8.13
NETHERLANDS	1996	6.98156	29.39	43.33933	9.81
NETHERLANDS	1997	5.1763	23.91	39.15111	8.56
NETHERLANDS	1998	5.30776	25.48	39.53344	7.14
NETHERLANDS	1999	7.35992	34.28	44.06739	8.07
NETHERLANDS	2000	5.27958	24.17	39.39874	6.97
BELGIUM	1995	5.02831	29.74	36.78928	8.35
BELGIUM	1996	9.07679	32.06	44.8882	10.94
BELGIUM	1997	7.894	32.96	43.03568	7.25
BELGIUM	1998	8.81879	34.37	44.51137	9.04
BELGIUM	1999	7.7512	32.54	42.78798	7.33
BELGIUM	2000	8.30419	31.5	43.71298	7.39
LUXPSELL	1997	3.78592	18.18	31.53377	5.68
LUXPSELL	1998	5.6539	24.42	37.00803	6.18
LUXPSELL	1999	8.2516	31.44	42.20193	5.37
LUXPSELL	2000	6.547	28.55	39.0582	4.4
FRANCE	1996	7.39111	37.6	49.5264	5.56
FRANCE	1997	1.30525	17.89	26.01824	1.14
FRANCE	1998	4.72251	27.79	43.25204	3.95
FRANCE	1999	5.1468	30.23	44.47836	4.64
FRANCE	2000	3.88299	27.27	40.42061	4.35
UKBHPS	1995	7.71139	33.7	40.43869	2.95
UKBHPS	1996	6.97268	33.41	39.07027	2.07
UKBHPS	1997	6.69382	32.91	38.51226	1.81
UKBHPS	1998	6.62455	31.46	38.38062	2.05
UKBHPS	1999	5.9903	32.86	36.97245	1.68
UKBHPS	2000	6.37916	33.1	37.83961	1.77
IRELAND	1995	8.9959	32.19	32.90269	1.59
IRELAND	1996	11.86574	34.13	36.34635	2.02
IRELAND	1997	9.77275	29.11	33.96003	1.89
IRELAND	1998	8.69644	25.57	32.49064	2.49
IRELAND	1999	10.12205	31.64	34.40035	3.3
IRELAND	2000	5.32062	24.79	26.29536	0.85
ITALY	1995	13.81144	32.85	34.90027	14.78
ITALY	1996	11.95472	29.54	33.24141	12.24
ITALY	1997	10.75507	30.25	31.98112	11.97
ITALY	1998	10.26781	28.39	31.42095	12.21
ITALY	1999	13.27807	30.72	34.45861	15.5
ITALY	2000	14.32683	32.54	35.30749	16.16

Source: ECHP
Sample: Stayers full-time

Table 6: continued

Country	year	cuts estimated	cuts observed	freezes estimated	freezes observed
GREECE	1995	10.43413	28.98	24.03564	3.92
GREECE	1996	6.75622	17.54	19.46063	5.72
GREECE	1997	7.63949	21.13	20.7187	5.97
GREECE	1998	11.9585	26.9	25.50687	7.13
GREECE	1999	15.24925	27.95	28.06855	10.35
GREECE	2000	18.77823	28.2	30.08117	24.01
SPAIN	1995	11.08309	27.41	20.93119	1.61
SPAIN	1996	17.50816	37.94	25.37229	1.08
SPAIN	1997	21.177	42.64	27.00953	1.39
SPAIN	1998	16.57476	36.77	24.86304	0.91
SPAIN	1999	16.76127	38.24	24.96724	1.12
SPAIN	2000	17.18608	38.62	25.19865	1.69
PORTUGAL	1995	9.50794	24.16	27.78492	8.24
PORTUGAL	1996	9.05513	18.48	27.21017	6.35
PORTUGAL	1997	8.086571	15.58	25.88449	6.44
PORTUGAL	1998	9.39015	16.84	27.63888	8.48
PORTUGAL	1999	10.67473	13.62	29.14487	10.65
PORTUGAL	2000	9.622331	13.94	27.92475	11.75
AUSTRIA	1997	22.80267	34.35	24.6468	16.99
AUSTRIA	1998	20.32306	27.54	23.76446	13.64
AUSTRIA	1999	22.12231	24.32	24.42402	20.8
AUSTRIA	2000	20.19582	22.34	23.71388	23.15
FINLAND	1997	9.24961	31.03	39.95563	9.76
FINLAND	1998	8.156	29.01	38.30311	8.85
FINLAND	1999	8.65764	29.87	39.06767	12.04
FINLAND	2000	7.29497	24.1	36.7665	10.6

Source: ECHP
Sample: Stayers full-time

4 The impact of institutions on nominal wage flexibility in Europe

In the previous chapter we have tested and measured nominal wage flexibility in the EU countries following two approaches: a statistical approach, based on the direct estimation of frequencies of hourly wage cuts from the observed wage changes; and a structural approach, which is an attempt to accommodate measurement error in the observed wage changes, estimating the frequencies of hourly wage cuts as predictions from estimates à la Altonji and Devereux.

It is well known that a number of institutional labour market characteristics, such as collective bargaining and employment protection legislation, are possible causes of nominal wage rigidity in Europe. 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'.

It is the purpose of the present chapter to address empirically the question of how labour market institutions affect hourly wage change distributions in the EU countries exploiting observable cross-country differences in labour market institutions. More specifically, using each of the two frequency estimates obtained in the previous chapter as a measure of nominal wage rigidity, we investigate their variation across countries in comparison with the available institution measures for: 1) centralization of wage bargaining;

2) coordination in wage bargaining; 3) employment protection legislation; and 4) coverage of collective agreements. These institutional variables are a product of the literature on macroeconomic performance (OECD (1997), OECD(1999), Nickell and Layard (1999); Boeri, Brugiavini, Calmfors et al. (2001) and Cesifo (2002)).and, although time invariant, they show a sufficiently high degree of cross-national heterogeneity to permit identifying an institution effect on nominal wage rigidity.

There are only a few papers taking up the issue of institutional explanations for downward nominal wage rigidity in Europe from an empirical point of view (Dessy (2002), Holden and Wulfsberg (2004)). The existing literature about the relationship between macroeconomic performance and collective bargaining is instead mainly focused on the impact of institutional variables on the rate of unemployment[20], failing to address the effects on individual wage changes.

In Dessy (2002) some preliminary results on the impact of institutions on nominal wage flexibility are provided for a past version of the ECHP. The hourly wage change distributions therein are based upon the sole measure of wage offered, as of that time, by the data set, that is "total net wage and salary earnings". This measure, however, presents a serious flaw in that it may obscure pure labour market variations with variations caused by the tax system. We extend the analysis in Dessy (2002) along two directions. First, we consider "gross current wage and salary earnings", made available in the last update of the ECHP, to obtain a more appropriate hourly wage change distribution. Second, we exploit the increased time dimension of the

[20]See Flanagan (1999) for a comprehensive survey about collective bargaining and macroeconomic performance; and Nickell and Layard (1999) for a survey about labor market institutions and unemployment.

new ECHP to obtain more precise estimates and control for the possible additional impact of time varying country specific macro variables, such as unemployment, inflation and labour productivity. Finally, in addition to the frequencies of wage cuts obtained from observed wage changes, we also use the expected wage cut frequencies and wage freezes predicted by the Altonji and Devereux model of nominal wage rigidity.

Holden and Wulfsberg (2004) apply basically the same methodology as Dessy (2002) on a sample of measures of downward nominal wage flexibility estimated on 14 countries. They use an unbalanced data-set of hourly nominal wages at industry level over the period 1973-1999.

In this chapter we follow 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 measured. 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 of 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 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 chapter is the following. In the next section we summarise the implications of the existing theories that explain downward wage flexibility with institutional issues. In section 4.2 we describe the institutional and macroeconomic variables used in the empirical analysis. Section 4.3 reports results for the statistical analysis. Section 4.4 presents regression results. Section 4.5 concludes.

4.1 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 Akerloff 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 heterogeneous 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 increases the incidence of wage cuts.

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

4.2 Data

Our empirical analysis is focused 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 bargaining partners.

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 institution.

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[21].

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 Driffill (1988) index and information gathered by the OECD on employers' associations.

As we can see from Table , for the percentage of employees covered

[21] Although recently new rankings have been proposed in Nickel, Nunziata, Ochel (2005), we have not used them in this analysis because measures of institutions do not vary enough over the period considered.

by collective agreements 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 explanatory 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 percentage annual variation in *Labour productivity in the business sector*, that is total economy less the public sector.

Table 7 reports some descriptive statistics for the main variables used in the empirical analysis.

Table 7: Descriptive statistics 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[22]. 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_{56} 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;

[22]Spearman's correlation is useful because: 1) it is non-parametric; 2) as a measure of linear association between ranks it is not necessarily a measure of linear association between the actual values.

- 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 between 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.

4.3 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, or at least the property that ratios of intervals are meaningful. The latter avoids this problem by treating institution measures as purely ordinal, but to avoid using too many dummies one may have to, for example, group countries, which may be arbitrary. We prefer the former approach, which will model hump or U-shaped or U-shaped relationship, although both approaches have their merits.

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), and (LN). This specification is general enough to capture simultaneous non-linear effects, such as hump and U-shaped correlations with the dependent variable. For each centralisation measure we consider a different linear projection E^*, thus our baseline model is the following

$$E^*\left(y|x\right) = \alpha_o + \sum_{i \in I_C} \left(\beta_i x_i + \beta_{i,i} x_i^2\right) \tag{9}$$

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

In the empirical application model (9) 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 occurrence 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 Devereux 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 (2SLS) 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[23], Tables 8-11 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

[23]For estimation methods implemented and tests see Appendix 3. Notice that, wheras for the observed measure of nominal wage rigidity we have 57 observations, the number of estimated measures are 69. In fact, estimates can be calculated also for the first year.

for PCOV. Third, we find an "hump-shaped" impact for COORD. 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 intermediate strictness. For PCOV, instead, the predicted negative impact holds since the beginning over a large portion of the sample.

Table 8 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%

Table 9 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%

Table 10 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%

Table 11 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 qualitative 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 12 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, in Table 13, 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 12 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 13 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			

4.4 Conclusions

In this chapter we have assessed the role of institutions in explaining nominal wage rigidity in Europe using empirical and estimated frequencies of wage cuts from the 1994-2000 ECHP survey.

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.

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

5 A validation study for measures of wage and wage changes in the French Labour Force Survey

In this chapter we introduce the analysis of the French case, that covers also the next, concluding chapter. The focus on this country comes from the unique opportunity of having access to the sources of data collected at INSEE, the French National Institute of Statistics[24]. The excellent quality of data available, and the possibility of matching between them, allowed us to contribute to two crucial issues of wage rigidity: 1) the relevance of measurement errors, for determining the extent of wage rigidity; and 2) the difficulty of linking wage changes to firm-level shocks, for measuring wage rigidity as the extent of wage adjustment at the micro-level. In this chapter we focus on the first issue, carrying out a validation study of the French Labour Force Survey (FLFS) with the intention of understanding the impact of individuals behaviour in reporting wages on measures of wage rigidity. The next chapter will deal with the second issue, introducing a new definition of wage rigidity.

As we have seen, the relationship between the actual and counterfactual distributions of wage changes is based on very strong assumptions. Particularly relevant are the assumptions on measurement error. Measurement error can arise for two reasons: 1) the unavailability of data on contracted

[24]Administrative and firm data used in this chapter and the next are considered sensitive data. Therefore access to them is allowed only under the strict supervision of INSEE administrators. I thank Pierre Biscourp and Nathalie Fourcade, both INSEE administrators, for supervising me while working on a representative sample and checking the final results on the whole dataset.

hourly wages; and 2) reporting and rounding errors in individual surveys. The first type of error is quite difficult to overcome since very rarely the base-wage is reported at individual level separately from bonuses, overtime pay and other benefits[25]. Contracted base-wages would be the best unit of measure to work with, since most of the theoretical models trying to explain wage rigidity are formulated in terms of this measure. Moreover, base-wages are not distorted by hours dynamics. Normally, when working with individual surveys, a proxy for the base-wage is constructed dividing total earnings by the number of hours worked, and a classical measurement error is attached to wage or wage change equations. If, as for most of the administrative data, hours are not available they are included in the error term.

The second type of error can be taken into account as long as appropriate validation studies of survey data are available, so that reported wages or estimates based on them can be appropriately corrected [26]. Otherwise, together with measurement errors of the first type, they are modelled following the classical assumptions[27]. As all the literature on validation studies shows quite clearly[28], however, reporting errors are far from following the classical assumptions: often they are correlated to the dependent or the explicative variables and this can seriously bias parameters' estimates. As Smith (1999) shows for the UK BHPS, reporting errors can bias wage change distributions and measures of wage rigidity because, contrary to what generally found when they are treated according to the classical assumptions,

[25]The only exception is the UK New Earnings Survey, analysed in Nickell and Quintini (2003) and Barwell and Schwartz (2004).

[26]e.g. in Card and Hyslop (1997).

[27]Akerloff et al. (1996), Altonji and Devereux (2000) and Fehr and Goette (2003).

[28]See Bound, Brown, Mathiowetz (2001) for a detailed survey.

they do not explain wage cuts but the spike at zero.

Since to date no detailed validation studies are available for France, we start our analysis of wage rigidity in this country with an effort to evaluate the bias induced by reporting errors on wage and wage change equations (and therefore test the relevance of classical assumptions on measurement errors) in the FLFS.

In particular, we collect all the available sources of information on wages in France: a 1:25 sample from the Déclarations Annuelles de Données Sociales (DADS), the Revenue Fiscales (RF) and the annual Enquete Emploi (EE), the FLFS. These three sources are based on completely different methods of collection of data: the first two are administrative data, collected respectively from statements of firms and individuals; the third one is the standard labour force survey, in which individuals report their earnings. The three panel data on employees are respectively matched with balance sheets of a panel of all firms that are subject to the so called Bénéfice Réel Normal (BRN) fiscal regime. The three resulting matched employer-employee data-sets allow us: 1) *to identify correctly stayers full-time using the firm* and establishment identifiers, for defining the sample of interest with respect to a representative number of firms; 2) to compare wage and wage change distributions for stayers full-time from different data-sources, for studying in detail the characteristics of rounding errors; 3) as we will see in the next chapter, to test wage rigidity with respect to firm-level measures of idiosyncratic shock taken from the BRN.

Analogously to Smith (1999) we explain the spike at zero nominal wage changes as rounding error. Studying in more detail the characteristics of rounding errors we realise that, with the FLFS, modelling reporting errors according to the classical assumptions can be extremely distortive. One solution would be to avoid using individual surveys for studies on wage rigidity and rely just on administrative data. The problem is that, even with administrative data, unfortunately it is not possible to isolate the base-wage from overtime pay, bonuses and premia. We are not able therefore to check whether also assuming that measurement error on hours follows the classical assumptions can be distortive for base-wage dynamics.

The chapter is organised as follows. In section 5.1 we explain how we constructed our matched data-sets. Section 5.2 and Section 5.3 present the results of our validation study on labour earnings in the EE for wage and wage changes distributions respectively. Section 5.4 concludes.

5.1 The construction of matched employer-employee data

The construction of an appropriate data-set for investigating wage rigidity in France is one important contribution of our work. Instead of using separately labour force survey and administrative data on wages, we match each of them with balance sheets of an extended sample of French firms. In this way we restrict the analysis to employees belonging to same firms in each data-set, and add information on the firm-side, therefore extending the sources of explanation for measurement errors. We start with a description of data on employees, then move to the proper matching with firm-level data. At the end, the sample of interest is defined.

5.1.1 Wages and hours: administrative and declarative data on employees

We have three sources of information available for labour earnings and hours:

- The Enquete Emploi (EE) is a source of declarative nature: it is the

French Labour Force Survey. Surveys are carried out for no more than three consecutive years, and every year employees report their net wages in March. This measure includes the base-wage plus monthly bonuses received by the individual for the same month. Annual bonuses and premia are also reported, together with two measures of hours: usual total weekly hours, and the number of hours worked the week before the interview. Data available cover the 1994-2000 period.

- The Déclarations Annuelles de Données Sociales (DADS), of administrative nature: the establishments employing must give to the fiscal administration, for social security purposes, some information on their employees. This information contains a measure of annual net labour earnings received by individuals during the employment period spent in the establishment considered. This measure includes bonuses and premia. The DADS give also the number of months and hours paid, so that it is possible to calculate a measure of net hourly earnings. This is different from the contracted base hourly wage, since it includes bonuses. We do not consider the full data set (that covers all the working population in France), but a sample of 1:25, the so called DADS panel, of all people born in October of even years. Since the number of hours is available from 1994, we consider the 1994-2000 waves.

- The enquete Revenue Fiscaux (RF), of administrative nature, consists in the statements made out by individuals for their income-tax return, collected by the Direction Général des Impots (Internal Revenue Service). Such statements are compulsory and supposed to be given by all working people. The subsample available is composed by all people being in the EE of March of the following year. For these people the two sources of data can therefore be matched. The only measure available for wages in the RF is net labour earnings, including bonuses and premia received during the year. Information on the number of hours worked can be retrieved from the matching with the EE. Data available cover only the period 1996-1999.

As we can see, the three sources above are referred to the same statistical unit (employees receiving wages), and allow us to calculate a measure of hourly wages. However, there are important differences among them. First of all, wage information is referred to different time-periods: the month of March of each year for the EE, and the whole year for the DADS panel and the RF. This difference can generate problems when comparing the three data, but only if wages are taken in levels: if wage dynamics are considered, differences between the three sources due to the reference period become negligible.

Moreover, the measure of wages is different in the EE with respect to the two administrative data sets. The net wage reported in the EE is the gross wage, net from social security contributions due by the employee (cotisations salarié), the specific income tax imposed on most types of income for financing the public social security scheme (Contribution Sociale Généralisée CSG) that can be in part subtracted from the income tax return, and another similar tax that is instead entirely subject to the income tax return (Contribution au Rembursement de la Dette Sociale CRDS). This is therefore in principle the sum indicated in the wage bill of employees. We define this measure of wage '*net-net*' in order to distinguish it from the information available in administrative data. In the DADS and RF the measure of wages available is what we define the *net* imposable wage, i.e. the wage net only from employees contributions and the deductible part of CSG. We reconstruct the *'net-net'* wage in the fiscal sources for comparative purposes. This is therefore the measure used in the paper, although whenever possible we do prefer using the declared original values.

The third important difference between our data-sets regards the method of collection of information. Our purpose is to evaluate the impact of the nature of data on individual wage changes distributions. As it has been often noticed in the literature, individual surveys can be biased by reporting errors, that can give rounded (for salaries) or normalised (for the number of hours) figures. Administrative data should not be subject a priori to the same kind of biases, since wages are reported by employers or individuals compulsorily for administrative purposes, and can eventually be checked through by the fiscal authority.

In this paper, as in most of the validation studies available in the literature, we assume that administrative data give the true value of wages[29]. This assumption allows to identify *reporting error* in the EE as the difference, at individual level, between the wage reported in the EE and the one declared in RF. It is important to notice that the same comparison is not possible between the DADS and the EE, since the sample is different and the two sources can be matched only through the firm identifier. The interest in considering the DADS rests in its bigger size compared to the EE and the matched RF-EE sample, and in giving an alternative measure of hours with respect to the EE.

[29]The only exception is Abowd and Stinson (2003), where also administrative data are not supposed ex ante to be correct.

5.1.2 Matching employers and employees data

Both the EE and the DADS report sex, age, occupation (two-digits), industry, type of contract (full-time, part-time), and tenure. They also collect the firm (SIREN) and the establishment identifier (SIRET), and the geographic code of the region where the firm is located.

The EE has the advantage of giving more detailed information than the DADS on employees and their job characteristics. For example, we know occupation (4 digits), duration of contract (permanent, temporary, interim, etc.), function (production, study, commerce, etc.), hours regime (regular, alternated, variable), title, and job characteristics (work over night, on Saturday or Sunday).

The availability of establishment and firm identifier in the DADS and EE (and consequently in the RF) allows us to match the two data separately with any source of information at the establishment or firm-level. Firm and establishment identifiers are a priori better recorded in the DADS (where the establishment itself gives the information) than in the EE, where the SIRET is codified from the address and the fiscal code reported by individuals. We consider two important variables for firms: those taken from their balance sheets, and describing the situation of the firm, and those revealing the geographic localisation of the employing establishment.

The first group of variables is taken from balance data of firms subject to the so called Bénéfice Réel Normal (BRN) fiscal regime. Are subject to BRN all the firms in the industrial and commercial sectors, whichever their size, declaring sales for at least 3 million francs. The threshold is 1,5 millions francs for services industries. We consider the years 1992-2000. The variables contained in each file describe the situation of the firm at the 31th of December of the corresponding year. The BRN data give information on: sales, number of employees, value-added, profits, financial variables etc..For the descriptive analysis of this paper, we have used only three variables: main activity of the firm, sales, and number of employees. We use the information on the number of employees for determining the size of the firm, and measure the idiosyncratic shock of the firm with the past and present percentage increase of sales. This variable is used as an explicative variable for measurement error.

Since, as it is well known, wage negotiations depend also on employees' bargaining power, in our wage and wage change equations we use as an indicator of employees' bargaining power the local unemployment rate, that is the unemployment rate in the local labour market to which employees and establishments belong. The 1999 census of the active population reports the total number of active and unemployed population at commune level. These data have been matched with DADS, where the code of the common where the establishment is placed is known. These codes can be appropriately aggregated at geographic level in 'travel - to - work area' (350 in France), representative of the basin of employment. The unemployment rate calculated for these 'markets' represents our local rate of unemployment.

5.1.3 The sample of interest

Our analysis is carried out on full-time employees of the private sector, staying in the same establishment in at least two consecutive years. This sample has the advantage of being typically used both for validation studies (since it selects the employees more likely to report correctly their wages) and in the empirical studies on nominal wage rigidity (due to the fact that the purpose is to study the measure in which wage changes are implemented at firm level). In this paper, since we merge all data on wages with firms balance sheets in the BRN, we restrict our samples also to the firms included in the BRN. As we can see from Table 14 employees characteristics after matching respectively the EE, RF and DADS with the BRN are homogeneous across data although, in the absence of a corresponding pre-matching table we cannot conclude in favour of no - selection problems. The utility of matching earnings data with the BRN consists in introducing a further, useful element of control in our validation study.

Although in the DADS we can select only employees working at least 80% of a full-time contract, this is not a major problem for our purposes since the number of hours and days paid during the year are known. Moreover, being labour earnings in the EE referred to the month of March, the comparison between the EE and the RF is feasible only restricting to employees staying in the same establishment continuously during the whole year. We can therefore compare the wage declared in the RF of year n to the wage reported in the EE either in March $n+1$ or in March n for employees staying in the same establishment between January n and March $n+1$[30]. We also exclude from the sample employees having a second job, whose earnings are included in the RF statement but not in the EE, where only the first job is considered. For comparison purposes we impose to the DADS sample the restriction of considering only employees staying in the same establishment continuously for two consecutive years. Every year, on average, our samples are 16,200 individuals in the EE, 6,500 in the RF, and 209,000 in the DADS.

The structure of the sample from the three sources of data, as far as both

[30]See Appendix 4 for details on the reference period for measures of wages.

employees and firms characteristics are concerned, are very similar. Figures are given in Table 14.

In the next two Sections we analyse the impact of measurement error first on wage distributions and then on wage change distribution.

5.2 Measurement error in wage distributions

In this Section we carry out a validation study for the EE survey, considering the impact of rounding on measures of wages. In the next Section, a similar analysis will be done on wage change distributions, and therefore on measures of wage rigidity. For the same sample of employees, full-time workers staying in the same firm, we construct a measure of earnings comparable across data sources and then match them. Assuming, as in most of the validation studies available, that administrative data are error-free, we determine the extent of rounding in the EE and study its characteristics. We find that measurement error is asymmetric and concentrated on the left of zero, therefore indicating a general tendency in the EE to under-report wages when rounding. Analysing wages in levels, we find that the extent of rounding is correlated to individuals and firms characteristics, as well as to the level of wages. We conclude therefore that, similarly to US data, in the French EE measurement error due to rounding does not follow the usual classical assumptions. As a consequence, modelling reporting errors according to the classical assumptions induces biases in the estimation of coefficients of wage equations.

In what follows we compare labour earnings distributions from the DADS panel, the EE, and RF. We consider three different measures: annual earnings, weekly hours, and resulting hourly wages. Descriptive statistics give an explorative analysis of our data. In particular, the comparison between different data allows us to study some characteristics of reporting errors in the EE.

5.2.1 Non-response, under -reporting and rounding in the EE

In the DADS and RF total annual labour earnings is the measure directly imputed . In the EE, instead, total earnings are calculated from wages declared in the month of March, to which the annual bonuses received are added. It is therefore possible to isolate annual bonuses in the EE, but not in DADS. We add annual bonuses to monthly wages in the EE in order to obtain a measure of earnings comparable in the three data sets.

Due to the characteristics of the RF sample with respect to the EE, a priori it is possible to construct a measure of annual wages in the EE, comparable with the DADS 1998 and RF 1998, using 1) the EE March 1998 file, 2) the March 1999 file, or 3) an average of the two dates. As we will see, individuals behaviour in answering the questionnaire is characterised by an attitude to under-report and round wages. The average of labour earnings in the EE between the values reported in March 1998 and March 1999 underestimates the presence of rounding with respect to the values in March 1998 and March 1999. For this reason we decided to avoid considering the third option. This choice can be also motivated with the fact that, in this paper, we are interested in the reporting behaviour of respondents to the questionnaire. We therefore prefer working with the variables directly declared. The under-reporting behaviour of individuals in the EE is evident comparing the 1998 EE to the DADS and RF for the same year. We can reasonably assume that wages do grow over time. Instead we can see that the value of March 1998 is on average lower than the value for the whole 1998. Moreover, wages declared in March 1999 in the EE are lower than the 1998 values in the DADS or RF so that we can reasonably deduce that there is evidence of underestimation of wages in the EE.

In Figure 3 we show the distributions of annual earnings in the three data-sets. We consider only one year, since the shape of the distributions does not change significantly over the short time-period considered. The DADS and RF distributions are very similar, therefore supporting our assumption that measures reported in administrative data are the correct one.

A bar is inserted in Figure 3 for compatible levels of the *net-net* statutory minimum wage (SMIC). We can see a little spike in proximity of the SMIC only in the EE, whereas wages smaller than the SMIC can be observed in all the data. This result is at first instance a bit strange, since we focus on full-time workers. The variable identifying full-time workers in the EE and RF can be reasonably supposed correct as long as qualitative variables are less subject to reporting errors than continuous variables. In these data annual wages lower than SMIC can be explained with the fact that in some firms the number of weekly hours collectively bargained can be less than 39, therefore employees earning the minimum hourly wage and working full-time have still a total wage lower than the monthly SMIC, calculated on the basis of 39 hours. Looking more in detail at workers' contracts characteristics, we can try to explain the part of the distribution below the minimum wage. In particular, one part of employees earning wages lower than the SMIC are apprentices, *stagiers* or assistants. Others are concierges or have jobs in which part of the remuneration consists in benefits. Finally, many people round the SMIC to a lower level. These three categories of workers make up about two-thirds of individuals in the lowest part of the distribution of earnings. It can be that the remaining third are people with handicaps, for which there is a cut in wage compensated by a state contribution to their employer[31]. We have decided to keep workers earning less than the minimum wage in the sample for comparative purposes with the DADS sample, where all employees working more than 80% of a full-time contract are aggregated. Therefore, in the DADS, people working between 80% and 100% of a full-time contract can constitute most of the wage distribution below the SMIC[32].

In the EE, wages are concentrated on rounded values: 15% of wages are reported at 5000 Francs and multiples, 27% at 1000 Francs and 50% at 100 Francs. This behaviour is not observed in the DADS panel and the RF (the above categories represent less than 1% of data). Moreover, it seems that wages are not rounded at random, but they are systematically truncated at the lowest rounded value. This is what we define 'under-estimation' of wage in the EE, and it increases with the level of wage (see Table 15).

We can notice a small tendency to under-estimate wages in the RF with respect to DADS, although smaller than in the EE. Besides tax evasion, a structural difference between the DADS sample and the RF can be a possible explanation. This problem does not emerge when comparing the EE with the RF since the two samples are identical by construction. Therefore, the comparison between the EE and the RF source is particularly useful for studying the characteristics of reporting errors[33]. In Figure 4 we show the

[31]Unfortunately we are not able to detect them.

[32]Averaging measurement errors on hours worked, hourly wages of these workers should be equal to the hourly SMIC; as we will show, instead, a significant proportion of employees earn an hourly wage lower than SMIC.

[33]The only difficulty in the EE-RF matching can be the definition of the period for which wages are calculated. This issue is discussed in Appendix 5.1.

distribution of measurement errors, assuming that RF gives the true value of annual earnings. Table 16 considers the average of measurement errors and its contribution to the variance of wages declared in the EE.

We can notice that measurement error in the EE is negative on average. These results are consistent with those obtained in Hagneré and Lefranc (2002), where the 1996-98 EE and the 1997-98 RF are compared, and with studies carried out on US data. The distribution of measurement error is clearly asymmetric, with a density more concentrated to the left (the mean is smaller than the mode). Considering a simple decomposition of the variance of wages in the EE, in Table 16 we can see that the variance of measurement error is 15% of the variance of wages reported. The correlation between measurement error and the declared figures is negative and high; therefore the variance of wages reported is lower than for true wages. We thus confirm that relative underestimation is increasing in the level of wages.

We then study non-response determinants in the EE, ceteris paribus. In particular, we consider the impact of individuals' or firms' characteristics on 1) the probability of non-response, 2) the probability of underestimate wages, and 3) the magnitude of reporting errors. Table 17 presents the results for the RF data at time n compared to the EE values in March $n+1$[34].

First of all, respondents can decide not to answer the question about their wage, or to give the values in brackets (we consider this eventuality as

[34]Here we are assuming that the difference between the measure of wages in the EE and in RF, linked to the time-period of reference (see Appendix 4 for a discussion), does not depend significantly from individuals or firms caracteristics so that the analysis of their impact is not biased.

a non-response). The absence of a clear relationship between non-responses and individuals or firms characteristics is crucial for the representativeness of the EE when working with wages. A logistic regression allows to measure the above relationship, analysing the probability of non-response as a function of individuals and firms characteristics.

A number of variables result to be significant. First of all, the probability of not reporting wages increases with the level of wages measured in the RF, and this can explain part of the under-estimation of wages observed in the EE. Ceteris paribus, this probability is higher for executives, intermediate professions and white-collars than for blue-collars, who are in particular under-represented in the EE. The probability of non-response increases non linearly with age: it is the highest for people between 30-40 years and the lowest for individuals aged more than 45. If the questionnaire is filled not personally by the respondent but by a third person, the probability of not answering the question about wages increases: it can actually be expected that employees themselves are the best informed, among people in their family, about their labour earnings. Also, the probability of non-response is the lowest for employees full-time with a permanent contract (Contract pour une Durée Indéterminée, CDI). Sex, tenure and firm size do not seem to be significant.

Summarising, the non-response behaviour in the EE does not seem to be independent from individuals and firms characteristics, and this is the first source for bias in this data-set.

Other sources of biases come from reporting errors. We have already shown a general tendency in the EE to under-report wages. Now, we analyse separately the probability of under-report wages through a logistic regression. A third regression will allow to study the magnitude of reporting error, with the variable analysed being the error in absolute value.

The probability of under-declaring wages, as well as the absolute value of reporting error, increases with the level of wages reported in the RF. Therefore, people earning high wages seem to be the most reluctant to declare their earnings and, even when they are willing to declare their wages, they have a clear tendency to under-estimate it.

The probability of under-reporting and the magnitude of the relative error are high for blue-collars, and decrease when we consider white-collars, intermediate professions and lastly executives. We can see in this case the power of carrying out an analysis ceteris paribus: descriptive statistics would show that under-reporting is higher for executives, whereas this is due only to the positive relationship between the level of wages and the probability of belonging to the category of executives.

Women under-declare their wages more often than men, but the absolute value of their error is smaller than for men. We observe the same characteristics for employees with CDI contracts, or whose tenure increases. The opposite is true for employees aged more than 45 and for people with irregular working times: if they under-report less frequently, the measure of their mistake in absolute value is bigger.

The probability of under-reporting and the measure of relative error increases with the size of the firm, whereas employees declaring their annual bonuses in the EE give quite accurate measures. It is possible to explain this result with mistakes on bonuses: some employees receiving annual bonuses do not report them in the EE, and this results in an higher probability of under-reporting for people not declaring annual payments lump-sum (wages in RF are defined including bonuses and premia). It can be also supposed that the quality of response to the two questions is correlated, since in general people declaring their annual premia and bonuses are automatically reporting their wages more precisely.

As expected, the quality of the information on wages decreases when it is given by a third person and the highest is the level of rounding the bigger the magnitude of reporting error[35]. The results on the relationship between the level of rounding and under-reporting are more difficult to interpret: one declaration multiple of 5000 or of 1000 is accompanied to a small probability of under-report wages whereas the probability is higher for people declaring a multiple of 500 or of 100 (the reference measure is a response non multiple of 100).

Structural models of nominal wage rigidity usually are based on very strong assumptions on measurement errors. Altonji and Devereux (2000) consider two specifications for wages in levels. According to the first case, measurement error is distributed as a Normal independent from the other variables of the model, and in particular from the level of wages. In the second case, it follows the above distribution with probability p, being zero with probability (1-p), so that it exists a probability different from zero of declaring the true wage. Similar assumptions on measurement errors are made in Fehr and Goette (2003), Beissinger and Knoppik (2003), Bauer, Bonin and Sunde (2003), and Devicienti (2002).

From our comparison of the three sources of data, we question the merit of these assumptions on such reporting errors. Their distribution is clearly complex and very different from the simplifying assumptions usually made

[35]We remind that the answers to the EE are annualised, i.e. wages reported in March every year are multiplied by 12 and anual bonuses and premia are added. This explains the frequency of roundings at high values, multiles of 5000 or 1000. For example, one person declaringa monthly wage multiple of 500 will have an annual wage multiple of 1000.

in the literature. Importantly, the distribution of measurement error is a function of the true level of wage and its distance from a rounded value, it depends on employees individual characteristics, the size of the firm where she works, and the survey characteristics. All these elements can induce biases in the estimated wage equations.

5.2.2 The difficult measure of hours

There is an important difference in the information on hours between the EE and DADS. Whereas in the EE the usual number of weekly hours *worked* is reported, in the DADS the number of hours *paid* during the year is given.

The measure of hours in the DADS is given for each employee in the data, since 1994. The number of hours *paid* represents the total number of hours for the period when the employee has been linked to the establishment by his contract, including holidays, illness and accidents at work. In particular, overtime hours are included in the number of hours paid. This does not exclude, of course, that employees can be paid lump-sum for their work overtime. In any case, this measure is close to the concept of number of hours paid.

In the EE, two measures for usual number of weekly hours are available: one is the '*usual*' number of hours *worked* without reference to an explicitly specified time period; the other is the '*effective*' number of hours *worked* the week before the survey. Employees who have an irregular working time are allowed not to answer the question . For the measure of *effective* number of hours, employees not working the week before, for example because of holidays, clearly can report zero number of hours. In principle, the *usual* number of hours -whenever available- is the preferred measure for calculating the average number of hours worked during the year, although it gives values clearly concentrated around the legal weekly number of hours (but the same holds true for the *effective* number of hours). In order to avoid selectivity bias we measure the number of hours using the *usual* number of hours whenever they are available, otherwise the *effective* number of hours is taken, as long as it is different from zero.

As we can see from Table 18, we usually observe a spike at 39 hours. However, this values concentrates an higher number of observations in the EE than in DADS. Therefore, the number of hours in the EE is apparently overestimated with respect to DADS data.

The difference between the EE and DADS can be explained with the fact that probably part of the number of overtime hours worked are not paid. In fact, the bias between the sources of data is due for the most part to executives, and to a certain extent to intermediate professions.

The RF hourly wage is calculated using the number of hours reported in the EE in March $n + 1$. In the three sources we observe always a non-negligible proportion of wages below the minimum wage (Figure 5), as for annual wages measures.

Hourly wages in the EE are lower than in the DADS, particularly for the last fractiles. This can be explained both with smaller annual wages in the EE than in DADS, and with an higher number of hours reported in the survey than in administrative data. The different measure used for hours explains the difference between RF and DADS. We can therefore claim that hourly wages are underestimated in the EE, but obviously it might be that the concept of hourly wage in the EE is closer to the effective remuneration of employees than the measure constructed in the DADS panel.

As we can see in Table 19, under-estimation of hourly wages in the EE is common to all professional groups, and in particular it is evident for executives, for whom the biases in annual wages and hours are the highest.

Summarising, under the assumption that DADS and RF give the true value of annual labour earnings, it seems that the EE is biased by important measurement errors, characterised by non-response, rounding and under-declaring behaviour of individuals.

In the next section we show that the above results have important consequences on the shape of wage change distributions. Therefore, measures and tests of wage rigidity can give very different results if based on data of administrative instead of declarative source.

5.3 Measurement error and wage change distributions: the impact on wage rigidity estimates

After considering the impact of rounding errors on wage distributions, we analyse wage change distributions. According to the classical approach, in fact, wage rigidity is identified with the presence of a spike at zero nominal wage changes. We show that the spike at zero in French data is evident only in the EE, and can be entirely explained as rounding error. Moreover, the error in wage changes turns out to be correlated to individuals', firms' characteristics and the level of earnings. Therefore measurement error does not follow the classical assumptions not only in wage equations but also in wage change equations, contrarily for example to the assumptions in Fehr and Goette (2004).

In what follows, first we consider annual labour earnings as a measure of wages comparable across data sources. Then we explain the impact of the number of hours on wage change distributions. Since a clear result from all the data considered is a strong evidence of wage cuts, in the last Section we take advantage from the information available in the EE, especially regarding job characteristics, for explaining more in detail wage cuts.

5.3.1 Annual wage changes

The starting point of the classical analysis of wage rigidity is to show the shape of wage changes distributions. First of all, we can compare different data. Figure 6 presents the distributions of the percentage change of annual wages between 1999 and 2000 in the DADS, RF and EE respectively. The different reference periods for the measures of earnings (the year for DADS and RF and the month of March in the EE) might be a problem when comparing wages in levels. However this should not matter when comparing wage changes since the unit of measurement is time-invariant within the same data-set.

The most shocking feature is the absence of a spike at zero in the DADS and RF, against a quite high percentage of zero wage variation in the EE.

According to our discussion in the previous paragraph, the spike at zero, often interpreted in the literature as evidence of nominal wage rigidity, is to be interpreted as reporting errors in this case. In particular, it can be interpreted as the propensity by individuals to round their wages in answering the questionnaire. Consider for example people who, at any date, round their wage at 1000 francs. If the true variation of labour earnings is sufficiently low, the rounded value will not be modified. Among the employees whose wage (including bonuses) does not change in the EE between 1999 and 2000, 58% give in 2000 a value multiple of 1000 francs, 20% a multiple of 500 francs, 21% a multiple of 100 francs (but not of 1000 or 500), and less than 1% another value.

In our graphs we have reported also the percentage change of the *net-net* SMIC (comparable to the measure of earnings considered for each data-set). This value coincides with a little spike in the EE, but the same behaviour is not visible in the DADS and RF.

Table 20 allows us to compare the sources of data more precisely. The percentages of employees receiving wage cuts are similar in the EE, DADS and RF, whichever year is considered (between 20% and 30%, depending on the year), even though the percentage of strong cuts is lower in the DADS than in the EE. These results do not support the existence of downward wage rigidity, according to the *classical* definition.

The proportion of wage rises instead is much higher in the DADS and RF than in the EE. The difference, about 12% every year, corresponds to the percentage of zero wage changes in the EE. It seems therefore that the spike at zero observed in the EE is the result of the rounding behaviour of individuals, apparently strongly asymmetric: employees seem to round only small changes upwards of their wages, whereas cuts are not rounded. This evidence is contrary to the assumption, often found in the literature on nominal wage rigidity, that rounding behaviour of individuals is symmetric around zero.

Obviously, this evidence is not sufficient to conclude that nominal wages are not rigid, but it simply induces to avoid interpreting the spike at zero from individual surveys as evidence of nominal wage rigidity. Therefore the choice of an appropriate source of data, not biased by reporting errors, is crucial when studying this issue. In fact, it can be a mistake to model rounding errors according to classical assumptions. It might be worse, to ignore them. In this regard, the DADS present a small spike at zero in

2000, and a visible asymmetry. Therefore we do not exclude the existence of nominal wage rigidity, but if this is the case, its extent will be certainly lower than what an analysis of the EE would suggest.

Unfortunately, our sources of data do not allow us to establish whether the high percentage of labour earnings cuts observed can be explained as bonuses and premia reductions or as cuts in the base-wage contracted. In fact, as explained in advance, the only variable available in the DADS and RF is total wages including bonuses and premia, and in the EE only annual bonuses are reported separately from monthly wages. It can be that downward wage adjustments are realised through bonuses and premia changes, since the basis-wage is written in the employment contract and cannot be renegotiated downward without mutual consent of the employer and the employee. Apart from contractual constraints, the fact that wage adjustments are realised through bonuses and premia or through basis-wage changes is not the same. If the basis-wage is actually downwardly rigid, we can think that after a long period of difficulties, those firms who have been using bonuses and premia as an instrument for reducing their wage costs will lose this instrument as long as the variable part of wages is decreasing. However, it is possible that during the 90s this phenomenon has been compensated by the further increase of the percentage of bonuses and premia in total wages, due to the purpose of firms to make their wage costs more flexible. If we ignore the possible negative effect of losing an instrument of adjustment for firms experiencing repeatedly negative shocks, we can assume that from the point of view of both the firm and the employee a wage cut induced by a reduction of bonuses and premia or by a cut in the base-wage are equivalent since both imply a reduction of labour earnings for the employee and of labour costs for the firm.

The fact that wage cuts are comparable across the three data sources used so far does not imply that wage cuts in the EE are correctly measured. The comparison between EE and RF allows us to establish whether wage cuts in the EE are correctly measured and can therefore be considered as true cuts. Our answer unfortunately is not positive, and raises concerns in using individual surveys for studying wages growth rates. Among the survey respondents who declare a wage cut in the EE, only a percentage between 25% and 35% have correspondingly wage cuts in the RF. The others declare wage raises (the percentage of no wage changes in the RF is zero)[36].

Both Figure 7, showing measurement error on wages growth rate in the EE in 1999, and Table 21, that gives a decomposition of the variance analogous to the one carried out for wages in levels, can help us in describing precisely measurement errors distribution. Measurement error is negative on average every year (between -1% and -2%) and negatively correlated to the true wage growth rate, therefore introducing an important bias in the distribution of wage changes in the EE. The variance of the true wage is only one third of the variance observed in the EE.

5.3.2 Hourly wage changes

From the point of view of employees, annual wages are probably the most useful variable for studying wage rigidity since it is a measure of their total labour earnings. However, from the point of view of the firm the link between total wages and the number of hours worked is direct. For this reason it is interesting to study the issue of wage rigidity using hourly wages. We

[36] The percentage is the same if we compare wages growth rates in the EE between March $n+1$ and March n to wages growth rate in the RF either between year $n+1$ and year n, or between year n and year $n-1$.

therefore devote this section to the analysis of hourly wage changes, although the presence of bias due to errors in reporting hours induces us to focus on annual wages as the preferred measure of wages.

Between 1995 and 1999 50% to 60% of employees experienced a yearly change of the number of hours paid in the DADS (and in the 30% of the cases it is a reduction of hours), whereas only 30% report hours changes in the EE (of which 15% are cuts). In 2000 the reduction of the legal number of weekly hours (RTT) is visible in the increase of the percentage of cuts declared: 54% in DADS and 38% in the EE.

The distributions of hourly wage growth rate in Figure 8 have similar characteristics to annual wages distributions. The comparison between Table 20 and Table 22 allows us to measure the impact of considering the number of hours worked.

Annual and hourly wage distributions differ only marginally in the DADS for the years before 1999. Moreover, if we assume that hours are correctly measured in the DADS, annual wage cuts are far from being explained as a reduction of the number of hours worked, with the hourly contracted wage constant. In 1999 and in particular in 2000, the RTT seem to increase hourly wages: the percentage of hourly wage rises in 2000 is 81% whereas the frequency of annual wage rises is 74%.

On the contrary, the percentage of zero wage changes in the EE decreases if we consider the number of hours, in any year. The phenomenon is accentuated in 2000 because of the RTT increasing effect on wages. Its impact on hourly wages however is lower in the EE, probably because it is referred to the month of March, and therefore less subject to a shock whose effects can start appearing only a few months after the date of the first of January 2000.

5.4 Wage cuts and jobs characteristics

Since an important fraction of workers experience wage cuts, in this Section we try to establish in which measure this change is correlated to a change in jobs characteristics such as: reduction of the number of hours paid, reducing working at night or during the week-ends, change of activity, bonuses.

The EE is rich of this kind of information, for wages we can use either the EE or the EE matched with the RF. Table 23 shows that a quarter of employees experiencing wage cuts have changed occupation between March 1998 and March 1999 (results are similar for the other years). But it is in general quite hard to establish an ordinal classification of professions, and therefore their relationship with wage changes.

On the contrary, job characteristics can change for the same kind of activity, and it is possible to talk about improving or worsening of job characteristics. These sort of changes can be linked to wage changes, as for example we all know that working in situations different from the normality (overnight, Sunday, Saturday, etc.) is usually compensated with pay higher than the basis-wage.

The EE collects some information on job characteristics. In particular, we know if the individuals work on Saturday, Sunday, overnight or during the evening. Each of those variables is reported with three modality: for example, the variable '*evening*' is given the value 1 if the employee usually works in the evening, 2 if she works in the evening only from time to time, and 3 if this never happens. The same criterion is applied for the other qualitative variables. We construct therefore an aggregate indicator called '*job conditions*', that reflects the characteristics of working times different from the usual ones, according to social norms, and then we sum up these three variables. We implicitly assume that the different variables have the same weight and that the fact, for example, of not working any more during the evenings can compensate working on Sundays. For describing the dynamics of job characteristics we consider changes of the variable that we have constructed in this way between two periods of time.

We find that every year about one third of the employees experience a change in their working conditions, both in the positive and negative directions. Table 23 compares the frequencies of wage cuts for employees whose working conditions improve (1 in the table) and the total sample of employees. As can be seen, cuts are more frequent among people in the first group. The same table referred to wages net from annual bonuses in the EE gives very similar results: the frequency of wage cuts is about 5 points higher for employees whose jobs conditions improve. This result seems to suggest that annual bonuses and premia do not adjust when jobs conditions improve.

We obtain the same result using the variable describing the type of working times of individuals: the same hours every day, alternate hours, working times changing from one day to the next.

It is not easy to isolate bonuses in our data: only in the EE annual premia and bonuses are reported separately from the hourly wage rate, however monthly wages do include monthly bonuses. From Table 23 we can see that in 1999 only 24% of wage cuts, including premia, reported in the EE correspond uniquely to annual premia cuts (3 in the table), being wages net from annual premia constant or rising.

At the same time, Table 23 shows that in the EE and RF in 1999 about 80% of annual wage cuts cannot be explained either with a reduction of the number of hours worked (2 in the table), or with a change of profession inside the same establishment (4 in the table).

If we cross the previous variables, we observe that for 66% of wage cuts in the EE and 56% in the RF employees have experienced an improvement of their working conditions, a reduction of their weekly working time, a reduction of their annual premia without change of their wage excluding annual premia, or have changed profession inside the same establishment. It is therefore hard to interpret the above results as explanations of wage cuts as long as that changes regard also those individuals whose wage is constant or rising, in the same proportion, for the RF (56%) and, in a lower extent, in the EE (53%).

The percentage of wage cuts among those employees who do not experience changes in the number of hours, jobs conditions and profession for the EE and RF, or of hours and CS for the DADS, is still high: about one quarter in the EE, between 12% and 28% in the DADS and RF according to the year considered. Results are shown in Table 24.

5.5 Conclusions

In this chapter we have introduced the analysis of the French case. Different sources of data have been collected, in order to assess the quality of the French LFS. All the data giving information on employees have been matched, through firm and establishment identifier, to firm data. The construction of a matched employer-employee data set is particularly useful for our validation study, since 1) it enables to define the sample of interest more precisely; and 2) it extends information on the firm for explaining rounding behaviour.

Comparing wage distributions for similar measures of wages, extensive phenomena of rounding are evidenced in the EE. The error distribution shows clearly that wages are systematically under-reported in the EE, with respect to the administrative sources DADS and RF. Moreover, reporting error is correlated to many individual and firm variables normally used in wage equations and is asymmetrically distributed. Therefore it is hard to assume classical measurement error in wage equations.

Measurement error does not follow the classical assumptions and is asymmetric also for wage change distributions, therefore using the classical assumptions can be distortive also for wage change equations . Although the methodology used in this chapter is based on previous work carried out for the US, our analysis does not follow entirely all the steps of similar validation studies: for the purposes of this thesis it has been focused on evaluating the impact of reporting errors on the shape of wage change distributions, and therefore on measures of wage rigidity. However, another aspect of rounding behaviour of individuals can be object of further research on the same data-set: the autocorrelation of measurement errors. The study of this issue can be very useful for trying to model appropriately rounding errors in the EE.

From the analysis of rounding errors in the EE, we deduce that the most appropriate data source to be used for the studying wage rigidity is certainly the DADS, where wages are rounding-error free. However, in the DADS we do not observe hourly wages contracted by individuals. Therefore there is still an error component whose characteristics we are not able to analyse. We therefore can investigate either annual earnings or a proxy of hourly wages dynamics. From the qualitative point of view, the results of the descriptive (or *classical*) analysis of wage rigidity are the following: 1) there is no evidence of either nominal or real wage rigidity in France; 2) there is no downward wage rigidity in France. The absence of spike at zero and the presence of wage cuts can not certainly be explained with rounding errors, but they can still be due to changes in hours, or in the flexible part of labour earnings (bonuses, overtime pay, benefits, etc.).

The point is that, although nominal wages turn out to be flexible, we are not able from the inspection of wage change distributions to measure if wages change as they *should*. In the next chapter we will therefore undertake an econometric approach for estimating wage rigidity in France.

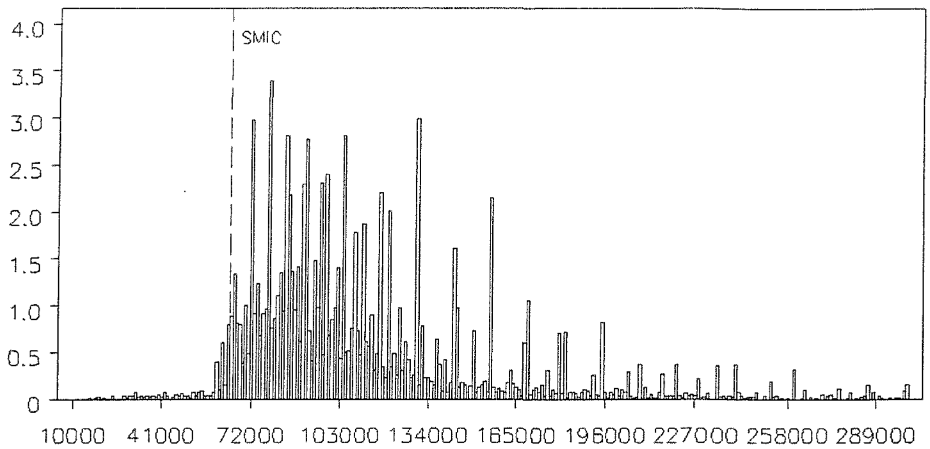
Table 14 : Structure of data after matching with the BRN

	Enquete Emploi	Revenus Fiscaux	DADS
men	71	72	70
women	29	28	30
Age< 30	17	14	17
30<age<40	33	32	34
40<age<45	17	17	17
age>45	34	37	33
industry	46	47	45
construction	9	8	9
commerce	18	18	19
other services	27	27	27
executives	12	13	14
interm-professions	23	24	24
white collars	18	17	18
blue collars	47	46	44
n.employees<20	18	18	17
20<n.employees<200	34	34	34
200<n.employees<1000	20	21	21
n.employees>1000	28	27	27
n.observations	97511	19523	1254720
years	1995-2000	1997-1999	1995-2000

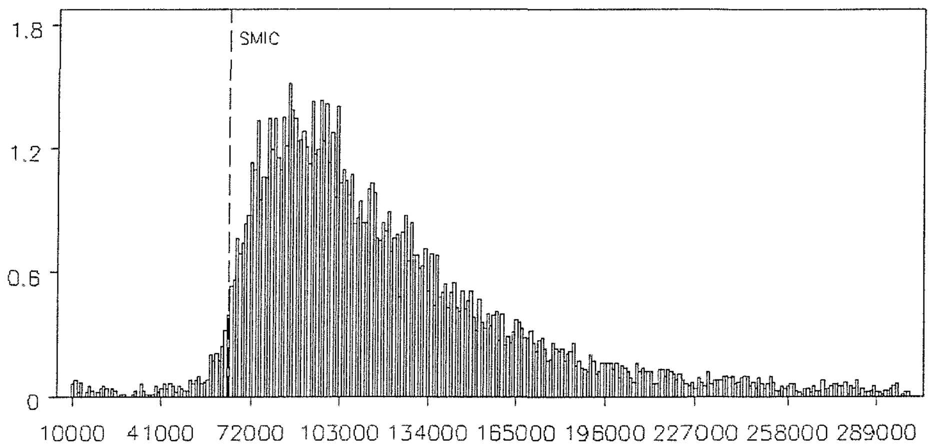
Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Figure 3: Distributions of annual wages

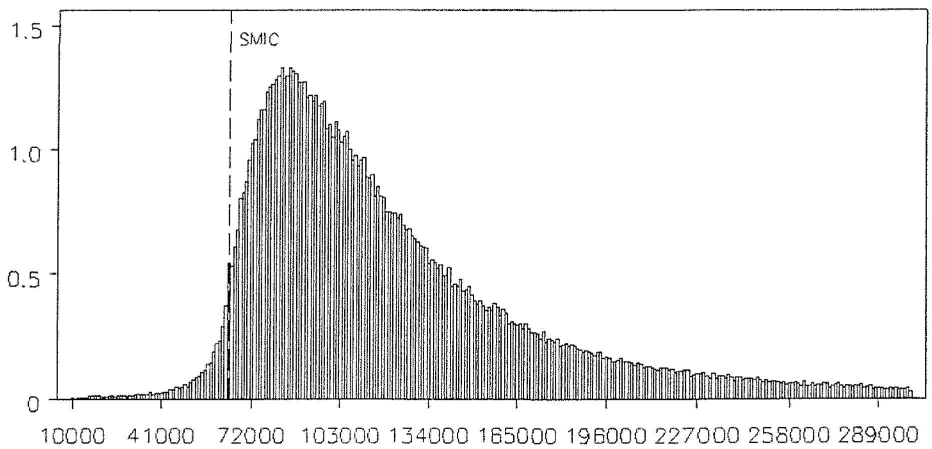
Enquête Emploi 1999



Revenus fiscaux 1998



DADS 1998



Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

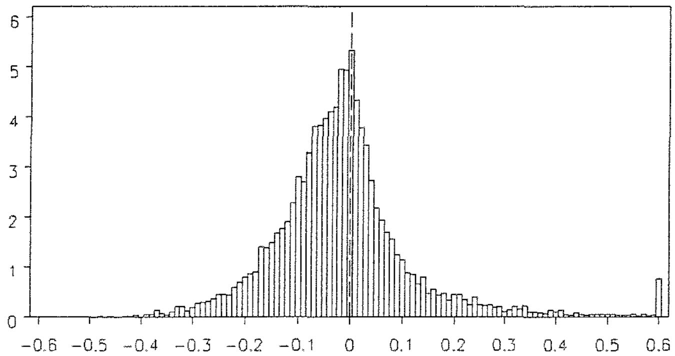
Table 15 Distribution of annual wages (in Francs)

	1st quartile	median	3rd quartile
Enquête Emploi 1999	84000	102000	130000
Enquête Emploi 1998	80600	98670	130000
Revenus fiscaux 1998	85756	106910	141669
DADS 1998	85645	108356	147210

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Figure 4 : Relative measurement error in the enquête Emploi

Enquête Emploi 1999 - Revenus fiscaux 1998



Enquête Emploi 1998 - Revenus fiscaux 1998

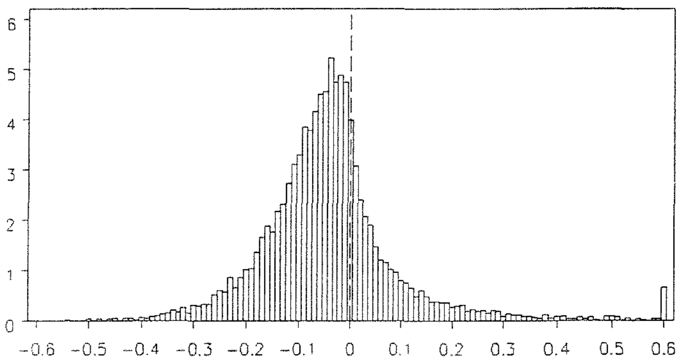


Table 16 : Mean and variance component of measurement error on wages in the Enquête Emploi

Revenus fiscaux	Enquête Emploi	Mean (EE-RF)/RF	Mean (EE-RF)	V(RF)/V(EE)	V(EE-RF)/V(EE)	Corr(EE-RF,RF)
1996	1997	-0.1%	-2113	1.11	0.12	-0.31
	1996	-2.7%	-4862	1.12	0.13	-0.33
1997	1998	-0.5%	-2350	1.07	0.12	-0.26
	1997	-3.8%	-6483	1.21	0.16	-0.42
1998	1999	-1.4%	-4006	1.23	0.17	-0.43
	1998	-4,0%	-6959	1.27	0.18	-0.47
1999	2000	-1,0%	-3568	1.22	0.18	-0.43
	1999	-4,3%	-7856	1.39	0.23	-0.55

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Table 17 : Measurement error in the Enquête Emploi and individual characteristics

	Probability of non-reponse on wages	Probability of under-report of wages	Relative error in absolute value
Logarithm of wage	0.640** (0.052)	1.871** (0.048)	0.032** (0.002)
Executives (ref. blue collars)	0.189** (0.061)	-1.599** (0.050)	-0.015** (0.002)
Intermediate professions (ref. blue collars)	0.197** (0.042)	-0.638** (0.031)	-0.008** (0.001)
White collars (ref. blue collars)	0.199** (0.049)	-0.141* (0.034)	-0.001 (0.002)
Women	0.019** (0.039)	0.275** (0.028)	-0.008** (0.001)
Age<30 (ref age 45)	-0.176** (0.058)	0.081* (0.040)	-0.008** (0.002)
30<age<40 (ref age 45)	-0.291** (0.042)	0.023 (0.030	-0.007** (0.001)
40<age<45 (ref age 45)	-0.162** (0.044)	0.032 (0.032)	-0.004** (0.001)
Irregular working hours	-0.041 (0.032)	-0.089** (0.023)	0.005** (0.001)
Answer given by a third person	0.332** (0.031)	0.173** (0.023)	0.011** (0.001)
Tenure	0.001 (0.002)	0.015** (0.001)	-0.0005** (0.0001)
Employees<20 (ref.1000<employees)	0.062 (0.047)	-0.293** (0.034)	-0.007** (0.002)
20<employees<200 (ref.1000<employees)	0.052 (0.043)	-0.137** (0.031)	-0.002 (0.001)
200<employees<1000 (ref.1000<employees)	-0.050 (0.048)	-0.077* (0.034)	-0.001 (0.002)
Contract different from CDI	0.334** (0.129)	-0.031 (0.092)	0.038** (0.004)
Declaration of annual bonuses in the Enquête Emploi		-0.491** (0.025)	-0.013** (0.001)
Declaration of annual wages multiple of 5000F in the Enquête Emploi		-0.207** (0.028)	0.023** (0.002)
Declaration of annual wages multiple of 1000F in the Enquête Emploi		-0.038** (0.019)	0.014** (0.002)
Declaration of annual wages multiple of 500F in the Enquête Emploi		0.160** (0.031	0.005** (0.002)
Declaration of annual wages multiple of 100F *in the Enquête Emploi*		0.186**) (0.020)	-0.002 (0.002)

Sources : enquête Emploi and Revenus fiscaux, from 1996 to 1999. RF data of year *n* are matched with the March, *n+1* enquête Emploi.

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – wages higher than the minimum wage (SMIC).

Table 18 : Weekly number of hours worked

		1stquartile	median	3rdquartile
Enquête Emploi	Executives	39	45	50
	Non- executives	39	39	39
DADS	Executives	38	39	39
	Non- executives	37	39	39

Sources : enquête Emploi 1999 and DADS 1998.

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

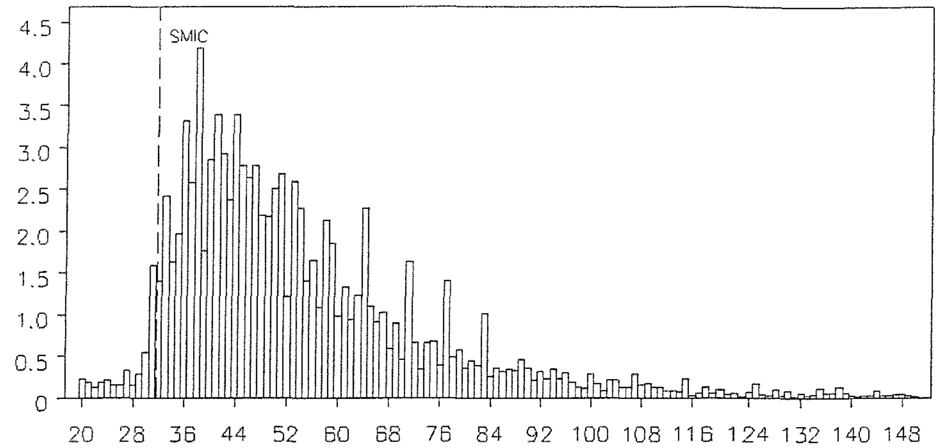
Table 19 : Distributions of hourly wages (in Francs)

		1stquartile	median	3rdquartile
Enquête Emploi 1999	Executives	71	88	110
	Non- executives	39	47	58
Enquête Emploi 1998	Executives	69	85	107
	Non- executives	38	46	57
Revenus Fiscaux 1998	Executives	72	90	115
	Non- executives	40	49	62
DADS 1998	Executives	89	113	151
	Non-executives	42	51	64

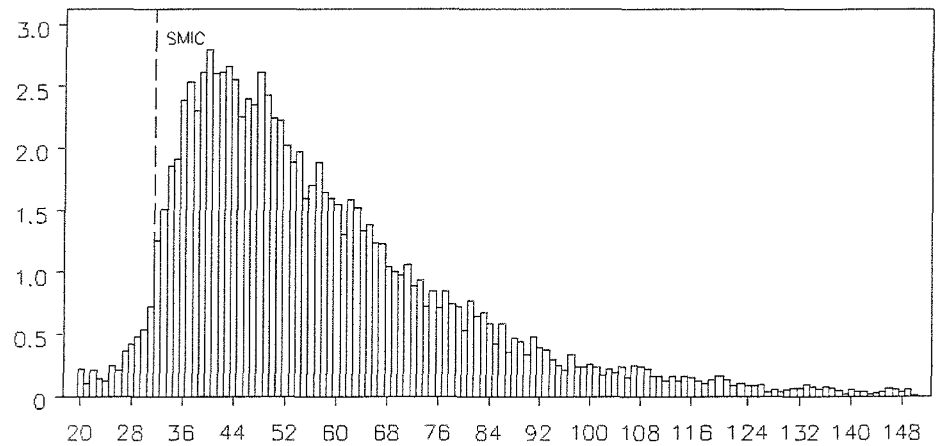
Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Figure 5 : Distributions of hourly wages in levels

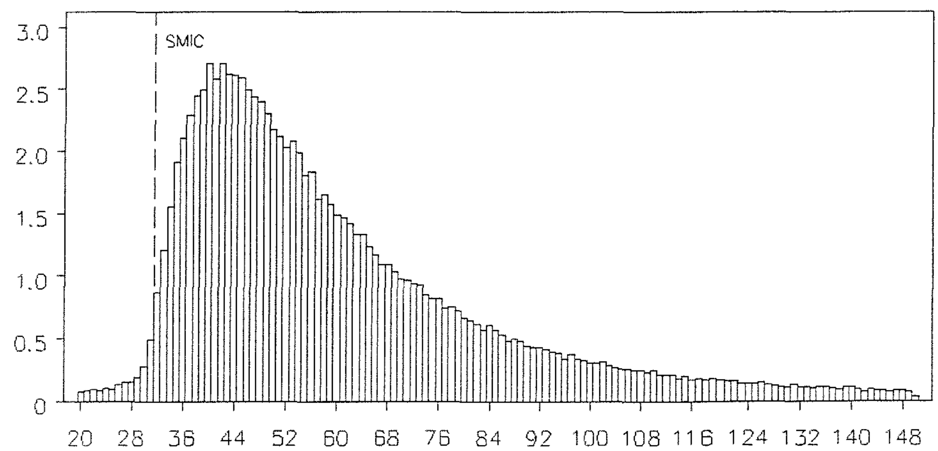
Enquête Emploi 1999



Revenus fiscaux 1998



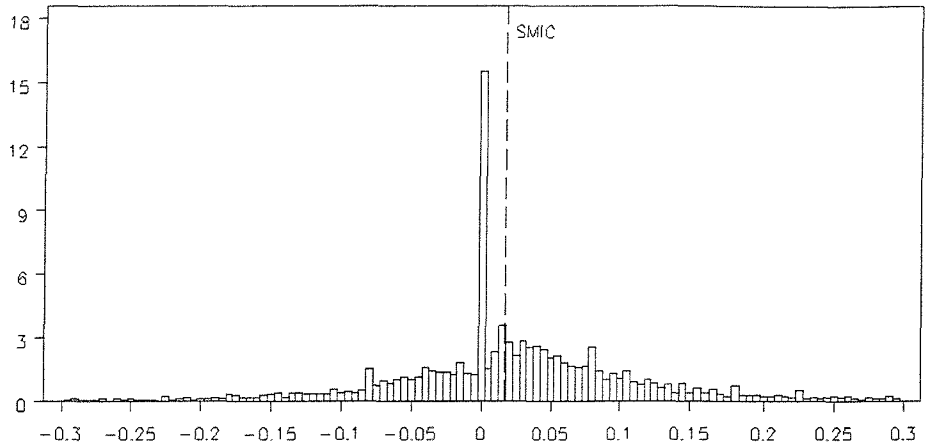
DADS 1998



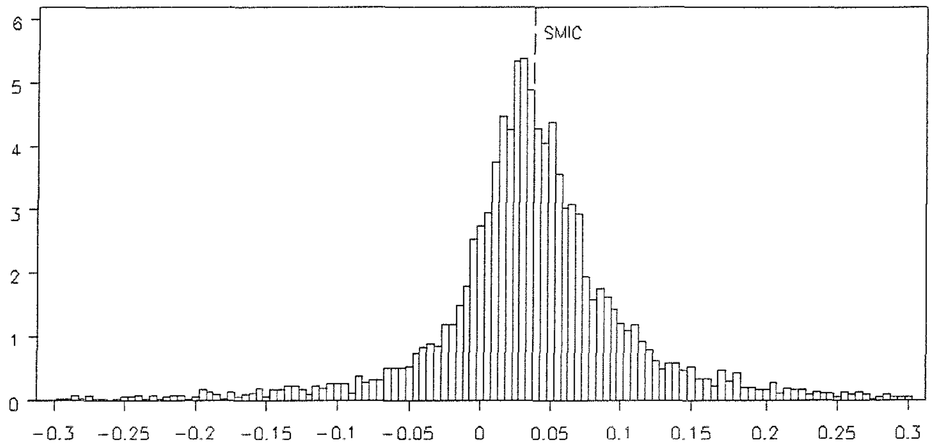
Sample : employees working full-time in the private sector, staying in the same establishment for two consecutive years – firms in the BRN sample.

Figure 6 : Distributions of one-year wage changes

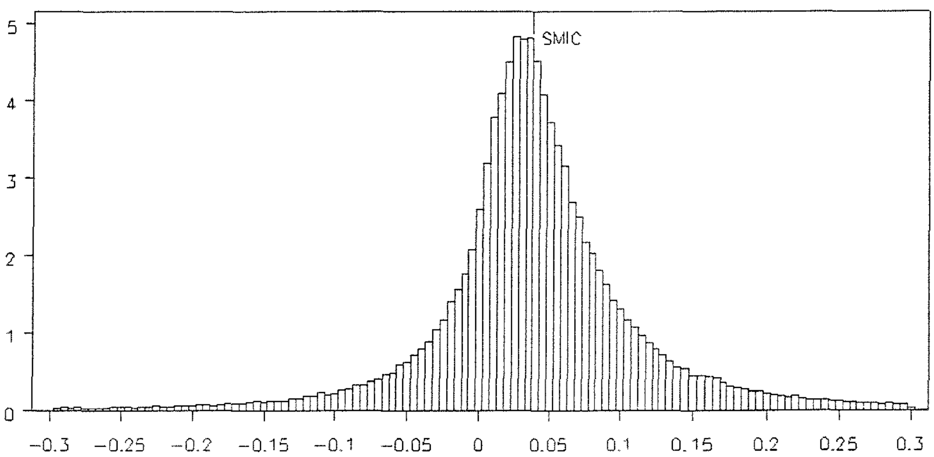
Enquête Emploi 1999



Revenus fiscaux 1998



DADS 1998



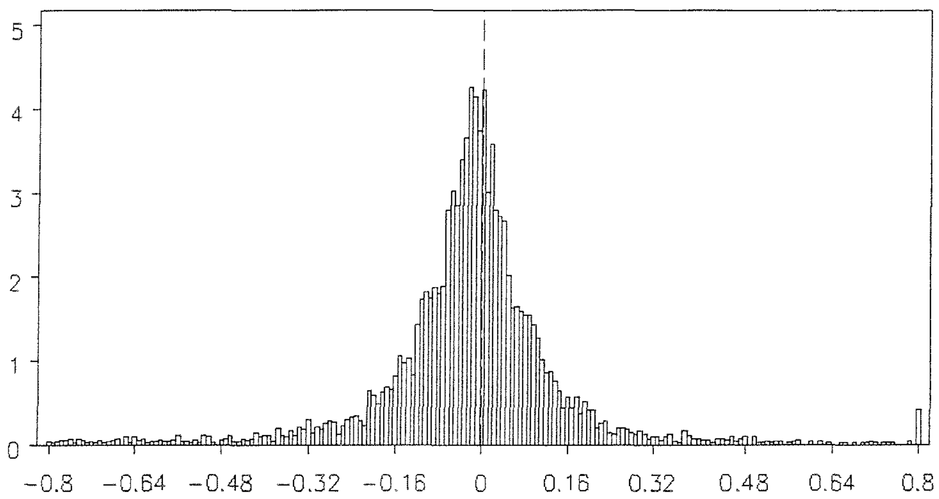
Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Table 20 : Percentages of zero, negative and strongly negative yearly wage changes.

	Enquête Emploi			Revenus fiscaux			DADS		
year	ΔLnW=0	ΔLnW<0	ΔLnW<5%	ΔLnW=0	ΔLnW<0	ΔLnW<5%	ΔLnW=0	ΔLnW<0	ΔLnW<5%
1995	12%	28%	16%				0%	16%	7%
1996	13%	29%	17%				0%	32%	12%
1997	12%	28%	16%	0%	20%	8%	0%	21%	9%
1998	12%	25%	15%	0%	20%	8%	0%	21%	9%
1999	15%	28%	16%	0%	28%	10%	0%	31%	11%
2000	14%	27%	15%				0%	26%	10%

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Figure 7 : Measurement error of the rate of growth of wages in the Enquête Emploi, Enquête Emploi 1999 - Revenus fiscaux 1998



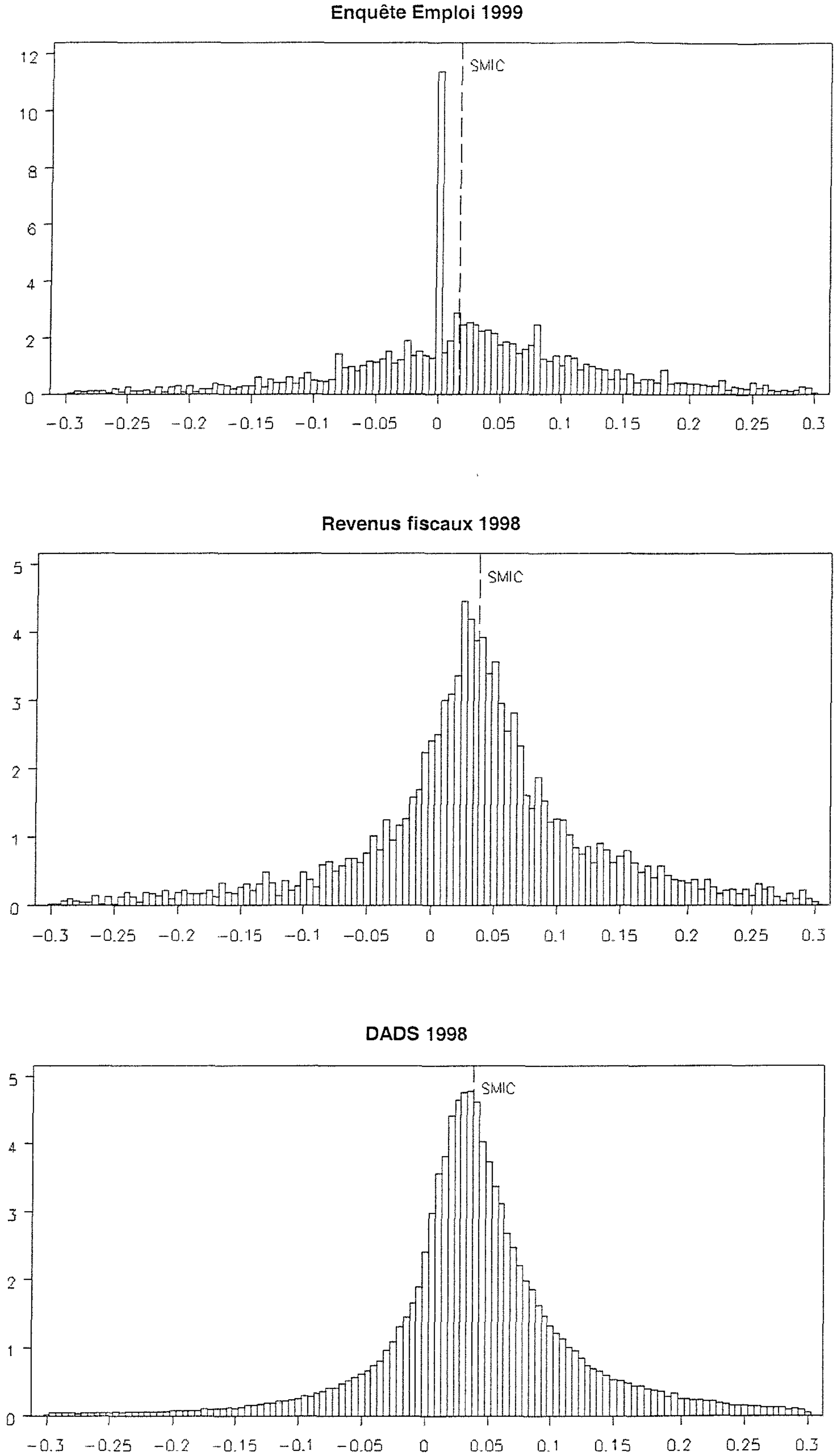
Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Table 21 : Mean and variance component of measurement error on wage growth in the Enquête Emploi

Revenus fiscaux	Enquête Emploi	Mean (EE-RF)	V(RF)/V(EE)	V(EE-RF)/V(EE)	Corr(EE-RF,RF)
1997	1998	-0.012	0.48	1.40	-0.54
1998	1999	-0.020	0.35	1.30	-0.48
1999	2000	-0.010	0.20	1.15	-0.36

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Figure 8: Distributions of hourly wage changes



Sample : employees working full-time in the private sector, staying in the same establishment for two consecutive years – firms in the BRN sample.

Table 22 : Percentages of zero, negative and strongly negative hourly wage changes

	Enquête Emploi			Revenus fiscaux			DADS		
year	ΔLnW=0	ΔLnW<0	ΔLnW<5%	ΔLnW=0	ΔLnW<0	ΔLnW<5%	ΔLnW=0	ΔLnW<0	ΔLnW<5%
1995	9%	32%	20%				0%	16%	8%
1996	9%	33%	21%				0%	32%	13%
1997	8%	32%	20%	0%	27%	15%	0%	21%	9%
1998	8%	30%	19%	0%	26%	14%	0%	21%	9%
1999	10%	32%	20%	0%	27%	14%	0%	27%	10%
2000	7%	26%	16%				0%	19%	8%

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

Table 23 : Wage cuts and changes in working conditions

	Enquête Emploi		Revenus fiscaux 1998	
	Individuals whose annual wages			
Proportion of employees whose	decrease	are constant or increase	decrease	are constant or increase
(1) job's conditions improve	21%	17%	20%	17%
(2) usual weekly number of hours decreases	22%	19%	20%	20%
(3) *annual* bonuses decrease without a decrease of wages(excluding bonuses)	24%	10%	15%	15%
(4) change of occupation according to standard classification	27%	25%	23%	24%
(1) or (2)	37%	32%	33%	33%
(1) or (2) or (3)	53%	38%	43%	42%
(1) or (2) or(3) or (4)	66%	53%	56%	56%

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

With the exception of the annual RF wage, all the information is taken from the Enquête Emploi. For the columns « Revenus fiscaux » we use the data matched with the Enquête Emploi.

Table 24 : Percentages of wage cuts among employees whose jobs conditions are unchanged

	Enquête Emploi	Revenus Fiscaux	DADS
1995	26%		12%
1996	27%		27%
1997	26%	19%	17%
1998	23%	19%	16%
1999	26%	28%	25%
2000	25%		20%

Sample : full-time employees of the private sector working in the same firm, same establishment for two consecutive years – firms in the BRN.

No changes of hours worked, jobs conditions and occupation for the EE and RF, of hours and CS for the DADS.

6 Estimating nominal wage rigidity in France with matched employer-employee data

The identification of wage rigidities based on the observed wage change distributions makes use in general of a restrictive definition of rigidity. According to this *classical* definition, at least part of wages downwardly rigid are interpreted as wage changes equal to zero. In this section we adopt a more general definition of rigidity: wage rigidities attenuate the transmission of productivity shocks to wages, and downward wage rigidity is defined as the transmission of the same shocks more completely upward than downward. Then we test the existence of downward nominal wage rigidity on French micro-data.

The kind of rigidity considered in this chapter is compatible with the absence of a spike at zero in wage change distributions. But its identification is more complex, since it needs the observation of an exogenous shock inducing an upward pressure to wages, a shock in the different direction inducing a downward pressure to wages, and the possibility of comparing a measure of the response of labour earnings to both these shocks. We now take advantage of the matching between sources of information on wage dynamics and employees characteristics (DADS, EE, RF) and the source of information on the firm (BRN) and use sales dynamics as a measure of the shock affecting wage changes.

We therefore try to test the existence of an asymmetry in earnings reactions to a positive or a negative shock in sales. For reaction of wages we mean the growth rate of wages. We also aim at testing the elasticity of wages to increases or decreases of the activity of firms. For completing the description of wage dynamics, and in particular of wage flexibility, we also analyse the impact on the probability of wage decreases of shocks on sales.

The above analysis is carried out conditionally on employees characteristics, and of their job in the firm. It allows us to describe the determinants of wage changes as a function of the observed characteristics of employers and employees. From this point of view the EE and RF are richer of information than the DADS. Nevertheless, the measure of wages in the EE, as we have seen, suffers of important reporting errors when studying wage dynamics. Concerning the RF, it presents the inconvenience that no more than one-period wage change is observed for each individual. It is therefore not possible to correct for the endogeneity bias linked to unobserved heterogeneity on wage growth. On the contrary, the DADS allow an estimation in the intra-individual dimension.

We start by introducing the theoretical motivations for a more general definition of nominal wage rigidity. Then, we analyse the DADS in section 6.2. In Section 6.3 we compare the results obtained with the RF and the EE. The RF allows for an estimation in the inter-individual dimension; the EE allows similarly an estimation in the intra-individual dimension, but starting from a dependent variable strongly biased from reporting errors. The comparison of these two last sources in the interindividual dimension allows therefore also to evidence the impact of reporting errors when using the EE for studying wage dynamics. Section 6.4 concludes.

6.1 The theoretical model

The theoretical background for our econometric model is based on the textbook Aggregate Demand (AD) - Aggregate Supply (AS) model[37]. We will

[37] See Blanchard (2003) as a general reference.

restrict to a very simple graphical analysis, giving the fundamental intuition behind our definition of wage rigidity.

We can start our discussion from the description of a typical long-run equilibrium using the AD-AS framework. In a long-run context, prices and wages are perfectly flexible by assumption. As a consequence, the AS-curve that describes the supply side of the economy is vertical at the level of output corresponding to the natural unemployment Y_n. Therefore the general level of prices of long-run equilibrium is completely determined by the demand side of the economy, represented by the downward sloping AD-curve. Figure 9 shows a typical long-run equilibrium position in the point E. As we can see in the same graph, any positive (or negative) shock of the demand results in an upward (or downward) shift of the AD curve that, in the long-run, has impact only on the level of prices therefore inducing inflation (or deflation).

The mechanism of adjustment in the short/medium - run, when prices and wages are by assumption not completely flexible, can be described as in Figure 10. In the short-run the AS-curve is upward sloping, showing the typical positive relationship between prices and output. Starting from the long-run equilibrium E, a positive shock of the demand shifts the AD-curve to the right. In the short-run the equilibrium moves from point E to E^+ because prices rise following the increase of demand. Also, since the production is increasing, employment is higher than the natural rate. This would increase employees' bargaining power, pushing wages upward in the medium run. But the following increase of prices, also driven by expectations of positive inflation, would slightly reduce the demand for goods. This would slowly shift the AS curve upward along the AD^+ curve, until the natural rate of unemployment is reached again. But, if wages were upward rigid, the expectations mechanism would not work, and E^+ would be also the long-run equilibrium, leaving employment higher than the natural rate.

Analogously, a negative shock would have symmetric effects, reducing employment and prices, and if wages were downwardly rigid the long-run equilibrium would be E^-, where the level of unemployment is higher than the natural rate in equilibrium. As a consequence, the existence of downward wage rigidity could be detrimental for employment. Therefore it is clearly crucial to measure its extent as a response of wages to shocks.

Often it is argued that, in this case, a zero inflation policy would be particularly dangerous. In fact, as we can see from Figure 10 (b), price-rigidity in the short-run would imply an horizontal long-run AS curve that, together with wage rigidity, would leave an equilibrium unemployment rate higher than in case (a) as a consequence of negative demand shock. It can therefore be argued that, in presence of wages downwardly rigid, it would be better to induce a certain degree of positive inflation through monetary policies. These would shift the AD to the right, back to the original position, and would be the only way to ripristinate the natural level of unemployment.

From the above discussion it is clear that a mere definition of downward wage rigidity as absence of wage cuts is not appropriate, even though the recent efforts of constructing measures based on individual data seem to go in a better direction with respect to the traditional use of aggregate data. Rather, we believe that measures of nominal wage rigidity based on the sensitivity of wages to firm-level shocks would be more meaningful in terms of a definition aimed at capturing the economic issues described above. In particular, if we look at a representative sample of firms in the economy at a certain time-period, some might be affected by positive and others by negative shocks. Therefore, a simple test of downward wage rigidity might consider the symmetry of wage reactions to positive or negative firm-level shocks. A symmetric reaction of wages to shocks in the two different directions would not be informative on the extent of wage rigidity. But if wages in the economy react less to negative shocks than to positive shocks we can deduce that we are really in a downward wage rigidity context, with potential negative consequences on long-run unemployment.

In the next section we translate this intuitive definition in an a statistical model, to be tested on the appropriately constructed matched employer-employee data of chapter 5.

6.2 DADS

We carry out two kinds of analyses of variance, explaining wage dynamics (annual or hourly) with employees, jobs and firms characteristics. This analysis allows to test the association between wage changes and one of the above characteristics. For describing wage dynamics, we consider separately first wage growth and then the probability of wage cuts. The explanatory variables are the same in both cases.

For each case, we estimate different coefficients for the link between wage growth and sales growth, depending on the latter increasing, decreasing or staying constant. If these coefficients are different, we conclude that there exists an asymmetry between a positive or a negative shock in sales on wage growth. The presence of downward rigidity is also captured by a smaller coefficient for sales decreases than for sales increases.

In both cases we correct for one part of the endogeneity of explanatory variables: the one associated to unobserved heterogeneity. The possible simultaneity with contemporary shocks is not considered. Therefore, at the present stage of the analysis, all our results have an economic interpretation in terms of rigidity only under the assumption of exogeneity of sales shocks at the firm level. This is an important limit, that could be eliminated in principle only by recourse to instrumental variables techniques. But unfortunately the DADS is not provided with such variables. We therefore select from the BRN the variable that to us looks like the *less endogenous*: sales.

Year by year regressions show that coefficients are stable across time. We therefore carry out the analysis of variance on the whole sample of years, controlling for the time-effects, but also for individual averages of time-dummies in order to take into account the fact that the panel is not balanced.

6.2.1 The statistical model

We restrict our sample to stayers $i = 1, ..., N$ in the same firm $j = 1, ..., J$ for at least two years. Then, consider the mapping $j(i) : i \to j$ assigning each workers i to the firm $j = j(i)$ employing her (since we are focusing on stayers $j(i)$ does not depend on time $t = 1, ..., T$). We try to explain her wage dynamics between two consecutive years, and the probability that such dynamics correspond to a wage cut, with a certain number of individuals characteristics, and of her jobs and firms characteristics. We therefore consider separately two different specifications of the statistical model: 1) in the continuous case, the dependent variable is wage growth; 2) in the discrete case, the dependent variable is the probability of receiving a wage cut. The explanatory variables can be constant or can vary over time; they can also be observable or unobservable.

In the continuous case, we estimate the following equation of general form:

$$Y_{it} = K\left(CA_{j(i)t}\right)+\beta^1 X^1_{it}+\beta^2 X^2_i+\gamma^1 Z^1_{j(i)t}+\gamma^2 Z^2_{j(i)}+\alpha_i+\phi_{j(i)}+\eta_t+\varepsilon_{it}+v_{j(i)t} \tag{10}$$

where:

$$\begin{aligned}
X^1 &= \text{vector of employees observable characteristics, that can vary over time} \\
Z^1 &= \text{vector of firms observable characteristics, that can vary over time.} \\
X^2 &= \text{vector of employees observable characteristics, that do not vary over time} \\
Z^2 &= \text{vector of firms observable characteristics, that do not vary over time} \\
\alpha_i &= \text{employees unobservable characteristics, that do not vary over time} \\
\phi_{j(i)} &= \text{firms unobservable characteristics, that do not vary over time} \\
\varepsilon_{it} &= \text{shock affecting the employees at date } t \\
v_{j(i)t} &= \text{shock affecting the firm at date } t
\end{aligned}$$

The dependent variable Y_{it} is the growth rate of wages: $\Delta \log W_{it}$. The term $K\left(CA_{j(i)t}\right)$, capturing our test of downward wage rigidity, is a function defined as follows:

$$K\left(CA_{j(i)t}\right) = \left(\alpha^+ 1_{\Delta \log CA_{j(i)t} \geq 0} + \alpha^- 1_{\Delta \log CA_{j(i)t} < 0}\right) * \Delta \log CA_{j(i)t}$$

where $CA_{j(i)t}$ is the change in sales, i.e. the demand/productivity shock at the firm-level. α^+ and α^- capture the impact of changes of sales on changes in wages, conditioning on being the shocks in sales respectively positive and negative. In this case, if wage dynamics reflect productivity changes, and these are linked to the activity of the firm, i.e. the demand-side of the economy, the expected sign for both coefficients should be positive (positive demand-shocks increase wages and negative demand-shocks decrease wages). In particular, if $\alpha^+ > \alpha^-$, wage changes are asymmetric and, according to our definition, we observe downward wage rigidity because wages react more to positive than to negative shocks.

In the discrete case the dependent variable is the probability for a worker to have a wage cut. We therefore consider a logit model specified as follows. Define:

$$\Psi_{it} = K\left(CA_{j(i)t}\right)+\beta^1 X_{it}^1+\beta^2 X_i^2+\gamma^1 Z_{j(i)t}^1+\gamma^2 Z_{j(i)}^2+\alpha_i+\phi_{j(i)}+\eta_t+\varepsilon_{it}+v_{j(i)t}$$

$$Y_{it} = \Pr(y_{it} = 1 \mid \Psi_{it}) = F\left(\Psi_{it}\right) \tag{11}$$

where $y_{it} = 1$ if wages decrease and $y_{it} = 0$ otherwise, and F is the cumulative logistic distribution:

$$F\left(z\right) = \frac{\exp\left(z\right)}{1+\exp\left(z\right)}.$$

In this case, the expected sign for α^+ and α^- is negative (positive shocks decrease the probability of wage cuts and negative shocks increase the probability of cuts), and $\mid \alpha^+ \mid > \mid \alpha^- \mid$ in case of downward wage rigidity.

The estimation of equation 10 by OLS (in the continuous case) or of the simple logistic regression 11 (in the discrete case) gives inconsistent estimates of the parameters of interest $(\beta^1, \beta^2, \gamma^1, \gamma^2)$ since correlation between characteristics of unobserved shocks and explanatory variables is very likely. This is also true in the second model if the shocks $\varepsilon_{it}, v_{j(i)t}$ are correlated over time.

It is possible to control for the existence of endogeneity biases caused by unobserved time invariant heterogeneity by exploiting the panel dimension of DADS and EE. This is unfortunately not possible with the RF, for which we never observe the same individual for more than two times. It is also possible to model the error term with a logit or probit model, as an autoregressive process of order one for taking into account the existence of autocorrelation. It is more complicated to control for the existence of simultaneity between explanatory variables and the idiosyncratic shocks $\varepsilon_{it}, v_{j(i)t}$. Both the DADS and the BRN are unfortunately not provided with valid instruments. The only possibility would be therefore the use of *'internal'* instruments: the lagged values of explanatory variables in the GMM method for example. But this approach, however, often gives not robust results and we do not pursue it.

In what follows therefore we consider the first two types of bias. In the continuous case as in the discrete case we model unobserved heterogeneity as a linear function of individual averages of explanatory variables:

$$\begin{aligned} \alpha_i &= \theta^1 \overline{X}_i^1 + \theta^2 X_i^2 + u_i \\ \phi_{j(i)} &= \sigma^1 \overline{Z}_{j(i)}^1 + \sigma^2 Z_{j(i)}^2 + v_{j(i)} \end{aligned} \tag{12}$$

In the continuous case, this approach corresponds to the Mundlak (1978) method. It gives an OLS estimator of β^1, γ^1 identical to the within estimator. Since the analysis is carried out only on employees staying in the same establishment for at least two consecutive years, the shift to the intra-individual dimension allows to eliminate individual and establishment effects not varying over time. This estimator therefore can deal with the unobserved heterogeneity bias, with the significance of parameters θ^1, σ^1 indicating the presence of the bias. Compared to the within estimator, the advantage of the Mundlak method is to allow the identification of parameters β^2, γ^2. In fact this is possible only if X, Z are not time invariant otherwise one can only identify $(\beta^2 + \theta^2)$ and $(\gamma^2 + \sigma^2)$. The impact of variables stable over time is therefore estimated without bias only in the absence of unobserved heterogeneity.

In the presence of autocorrelation on individual and firm shocks, the estimator is still consistent although it loses efficiency.

In the discrete case, the model used has been proposed by Chamberlain (1984). We estimate the following specifications:

- Logit model with random effects and equations 12.
- Logit model without random effects but with autocorrelated residuals.
- The conditional logistic regression (Chamberlain, 1980).

The above models give consistent estimates only under the assumption of independence of shocks, conditionally to individual effects. The third model allows us to control for the existence of a bias associated to an eventual temporal correlation of shocks without formulating an explicitly dynamic model, of difficult estimation. The third model takes advantage from the functional form of the logistic model for eliminating completely the individual effects, as would be done using a within estimator in the continuous case. For doing so, we condition on the sum of the outcomes over time. As the within estimator, the conditional logistic regression does not allow the identification of coefficients on variables that are constant over time. It is possible to test the existence of a bias of unobserved heterogeneity using an Hausman test, based on the comparison of a simple and a conditional logit.

6.2.2 Results: wages react more to a positive than to a negative shock

Table 25 summarises the results of the econometric analysis carried out on the DADS. The discrete case covers the first three columns of Table 25. The first two columns are the results from the same logistic regression à la Mundlak. The terms within and between are used for analogy with the continuous case. They are referred respectively to the coefficients of variables not averaged over time for the same individual, and to the coefficients of variables averaged over time (or stable over time).

The third column corresponds to the conditional logistic regression. It is in general very similar to the within logit à la Mundlak, but not identical. The advantage of the specification à la Mundlak is that it allows to identify the coefficients in the inter-individual dimension. The conditional logistic regression instead allows the estimation only in the inter-individual dimension. It corrects more mechanically for endogeneity biases due to unobserved heterogeneity. We have not reported the results for the random effects logit with autocorrelated residuals following an AR(1) process. The results are only marginally different from those obtained with the logit à la Mundlak and the conditional logit.

The last two columns correspond to the OLS estimator à la Mundlak.

The results obtained are the following. First of all, considering the inter-individual dimension, the probability of receiving an annual earnings cut increases with age, tenure, the average wage, and the local unemployment rate; it decreases instead with the size of the firm, the fact of belonging to the status of manager, of having an intermediate profession, and in a certain measure of being a white collar instead of a blue collar. It decreases also with the average growth of the firm. The results do not vary when regressing hourly earnings instead of total earnings on the same variables. The positive sign of average earnings (in logs) captures in particular the effect of the minimum wage, since earnings close to the SMIC cannot be reduced. Then, regarding the intra-individual dimension, the Mundlak and the conditional logit estimators give very similar results.

Lastly, the probability of an annual or hourly earnings cut for an employee staying with the same firm decreases with respect to the growth in sales. The asymmetry test is not probing for wage cuts: the coefficient corresponding to increases in sales is significantly higher than the one corresponding to decreases only for annual earnings and for the Mundlak estimator.

The second group of regressions, concerning the growth rate of earnings, gives results coherent with those obtained for decreases: the growth rate of earnings decreases with age, tenure, log of initial wage; it is also smaller for women than for men; moreover, earnings grow more for managers, intermediate professions, and in a lower measure for white collars than for blue collars; earnings also increase more in big firms than in small firms. The impact of a change in hours is also confirmed. The coefficient associated to a change in hours is between zero and one. An increase in the number of hours declared induces an increase of annual wages and a decrease of hourly wages, and this can be explained with a certain number of overtime hours that are not paid.

Regressions on the continuous variable allow to establish in the intra-individual dimension a more robust diagnosis concerning the asymmetry of response of earnings to a positive or negative shock in sales. The impact of sales growth rate on earnings dynamics is significantly higher for positive than for negative shocks. Also, the measure in which an increase of sales affects earnings growth is stronger than the measure in which a reduction of sales decreases them, as the sales reduction variable is the absolute magnitude of sales reduction.

In case a firm follows a strategy of constant mark-up, and changes the earnings paid accordingly, we expect a symmetric effect of an increase in sales on earnings. The smaller reaction of earnings when sales decrease seems to show the existence of rigidities, that are compatible with the evidence of an high proportion of wage cuts.

The efficiency wage theory, according to which employees are risk adverse and therefore willing to defend their earnings, gives an argument in favour to the existence of rigidity[38]. The ability of firms to take the risk is based on their privileged access to the financial tools for coping with it. This ability is a priori a prerogative of big firms. It can be easily assumed in fact that firms of big size have a financial structure and an access to credit that allow them to face transitory negative shocks reducing temporarily their benefits without practising wage cuts.

The results obtained for the negative local unemployment rate and the firm size can be interpreted in terms of bargaining power inside the firm. The higher proportion of earnings cuts in firms of small size than in firms of big size can also be explained with the strong presence in the last ones of trade unions, whereas an high local unemployment rate reduces external options of employees.

We can also notice that the comparison of the between and within estimators is relatively reassuring for the growth rate of earnings. Despite

[38] Although if this theory is applied to nominal rigidities it assumes that employees are victim of monetary illusion.

differences sometimes relevant of the estimated coefficients, the between and within coefficients have always the same sign. Therefore an analysis based only on the sign of coefficients can be carried out in the inter-individual dimension without diagnostic bias. This is particularly true for the link between earnings growth and a positive or negative shock of sales at the firm level. Unfortunately this is not true when we try to explain wage decreases more than wage increases.

6.2.3 Exploring the heterogeneity of coefficients

The previous regressions show that earnings react on average more to an increase than to a decrease in sales. Under the assumption of exogeneity of shocks at the firm level, this result can be interpreted as evidence of an asymmetry of the reaction of earnings to positive or negative shocks, corresponding to the definition of rigidity given above.

As it often happens when working with individual data, it is sometimes possible that this degree of asymmetry, as the magnitude of the reaction of earnings upward or downward, changes between individuals. For example, it is clear that an individual whose wage is close to the minimum will not experience an high decrease of her hourly earnings when sales decrease strongly, whereas this can be the case for another employee whose earnings are higher.

We restrict to the continuous case(12). We allow for heterogeneity of coefficients on the firm observed characteristics and coefficients on the sales shocks in $K\left(CA_{j(i)t}, I\right)$ as follows:

$$
\begin{aligned}
Y_{it} &= K\left(CA_{j(i)t}, I_i\right) + \beta^1 X_{it}^1 + \\
&+\Gamma^1 \left[1 + \sum_k \lambda_k \left(V_i^k - \overline{V}^k\right) + \sum_m \mu_m I_i^m\right] Z_{j(i)t}^1 + \varepsilon_{it} + v_{j(i)t}
\end{aligned}
$$

where we have omitted the terms invariant over time, since the estimation has been carried out for within specifications and

$$
\begin{aligned}
K\left(CA_{j(i)t}, I_i\right) &= \left(A^+(1+\beta^{+\prime} I_i) 1_{\Delta \log CA_{j(i)t} \geq 0} + A^-(1+\beta^{-\prime} I_i) 1_{\Delta \log CA_{j(i)t} < 0}\right) * \\
&* \Delta \log CA_{j(i)t}
\end{aligned}
$$

V : individual characteristics, continuous, invariant over time (*e.g.* average age of individual during the time-period) taken in deviations with respect to the average of the data over time

I_i : m-vector of individual characteristics invariant over time, (*e.g.* sex)

β^+, β^- and μ : m-vectors of interaction coefficients

In the absence of individual heterogeneity we have that $\gamma^1 = \Gamma^1$ and $\alpha^+ = A^+$ and $\alpha^- = A^-$. Otherwise, λ, μ, β^+ and $\beta^-$are significant and their sign shows how individuals with similar characteristics differ from the average behaviour, verified by individuals whose V will be equal to the average of the sample, and not having any of the characteristics in I^m. Results for these regressions are given in Table 26[39].

[39]Sometimes employees change their hierarchical position. For facilitating the interpretation of coefficients of variables interacted with the hierarchical position we have restricted our analysis to employees who do not change their position. We have then checked that this restriction does not induce any selection bias.

For our reference-individual (man, blue-collar, employed in firm of more than 200 employees, of age between 35 and 45, whose wage is between 1.3 and 2 times the minimum wage (SMIC), and who works in an area where the unemployment rate is equal to the national average) wages grow as a reaction to a positive shock of sales (with an elasticity of 1,1%), but do not react to a negative shock of sales. As a consequence, wages react asymmetrically to a positive or a negative shock.

Whenever considering individual heterogeneity, ceteris paribus our results are the following:

- The asymmetry is not significantly higher for wages between the SMIC and 1.3 times the SMIC than for wages between 1.3 SMIC and 2 SMIC. On the contrary, wages bigger than 2 SMIC seem to be more downwardly flexible.
- The asymmetry is lower for women than for men (wages less reactive to positive shocks).
- The sensitivity of earnings to positive shocks is higher for managers, whereas the sensitivity to negative shocks is lower. The asymmetry is therefore more important.
- The sensitivity of earnings to positive or negative shocks is lower for services sector than for other industries.
- Lastly, the sensitivity of earnings to negative shocks is higher when the local unemployment rate is higher, whereas the upward sensitivity does not change.

The above results show how important is inter-individual heterogeneity for earnings adjustment to firms shocks, not only in terms of average effect but also in terms of heterogeneity of parameters.

We conclude therefore in favour of the existence of wage rigidities in France, according to our general definition.

6.3 EE and RF

The DADS seem to be the most appropriate source for the analysis of this paper, because they give a measure of wages that is not much distorted by measurement error and because they allow an intra-individuals analysis, cleared from biases linked to unobserved heterogeneity of employees and firms. The EE follows the individuals for three years, and therefore it allows to calculate two growth rates of earnings for each individual and to realise estimations in the intra-individuals dimension, as for the DADS. But the EE gives a measure of wages severely distorted by reporting errors, particularly relevant for wage dynamics analysis. Although aware of introducing likely selection biases, in this section we restrict the analysis only to individuals reporting their earnings for three consecutive years, taking into account of non-responses.

The RF gives a better measure of earnings than the EE, but it covers only the period 1997-1999, without allowing to calculate more than one wage change for each individual over this time-period. The estimation with these data therefore can be carried out only in the inter-individual dimension. The analysis carried out on the DADS shows anyway that qualitative conclusions should not be affected.

In the following section we show the results obtained with the EE and the RF for three reasons: 1) test the robustness of the regressions realised on the DADS; 2) exploit the supplementary information contained in the EE; and 3) analyse the impact of measurement errors on earnings in the EE.

6.3.1 The analysis on RF

First, we consider the growth rate of earnings. Both for annual and hourly earnings, we find in the RF the result of higher sensitivity of earnings to positive than to negative shocks in sales. The asymmetry is significant , and comparable in value to the one obtained with DADS. We find equally the result that the coefficient of hours increases is positive for annual earnings but negative for hourly earnings, despite the measure for hours is potentially different in the two sources. The coefficients are sensibly different. Then, we find that earnings grow less when their initial level is high, when age and tenure increase, and also for women. Earnings also increase more for managers, intermediate professions and in lower extent for white collars than for blue collars. Earnings grow more for big firms. The local unemployment rate is always negative, but more significant. Summing up, these results are coherent with those obtained with DADS.

Among the additional variables that can be found in the EE, employees with an higher degree of education have an higher growth of wages, employees declaring a reduction of annual premia have a smaller growth,as those declaring an improving of their working conditions.

The results obtained for earnings cuts are globally coherent, except for tenure, but also for the asymmetry of shocks. Positive shocks have a negative effect on the probability of receiving a wage cut, and negative shocks are not significant. This result is coherent with the within estimator of the

DADS, but not with the between estimator, for which negative shocks were always significant and bigger than positive shocks. Given that, the fact that the asymmetry in the between estimator is not confirmed in the within estimator indicates the presence of biases, that we should be surprise not to find identical in the RF, but that most of all suggests to study this problem of asymmetry preferably by considering growth rates of earnings.

6.3.2 The analysis on the EE

Table 27 shows the results of estimations carried out on the sample EE matched with the RF, therefore only on the inter-individual dimension. The comparison with Table 26 gives therefore an idea of the impact of measurement errors (including the non-report).

Concerning the coefficients of shocks in sales, they are very similar, but more significant. It seems therefore that reporting errors induce a bias towards zero of this variable. The coefficients for hours are very close. The coefficient for the initial level of earnings has the same sign, but is strongly biased. The same is true for the hierarchical position inside the firm. Moreover, the coefficients for age and tenure are of opposite sign and very significant. The fact that earnings have been reported not personally and the type of contract are also significant, it seems therefore that measurement errors are strongly correlated to the following variables: earnings, age, tenure, hierarchical position inside the firm, the fact that wages are reported not personally and the type of contract. Similar results are found also for earnings cuts.

The analysis carried out on the complete sample of the EE allows to keep an higher number of observations and, more importantly, to carry out an analysis in the intra-individuals dimension. Tab. 28 gives the results of regressions comparable to those given for DADS, but only for annual earnings.

Those regressions confirm the biases of reporting errors shown before for the between estimator (also if the level of earnings is given not personally is no longer significant). The bigger size of the sample allows to find a between coefficient for positive shocks in sales significant and of the same extent than for the RF.

It is most of all extremely clear that reporting errors do not allow an estimation in the intra-individuals dimension for shocks of sales. For the variables bonuses and working conditions, however, these regressions suggest that the coefficient between is very similar to the within. Summing up, the use of the RF seem to be preferable to the EE, also if the analysis has to be restricted to the inter-individual dimension.

6.4 Conclusions

In this chapter we have started from the consideration that a perfectly flexible distribution of wage changes observed does not imply necessarily that wages are flexible in France. Thanks to the construction of an appropriate matched employer-employee data set, that matches three data-set of different source with firms balance sheets (BRN), we have introduced a more general definition of wage rigidity, testing the asymmetry of wage adjustments to positive or negative shocks at firm-level. Shocks are measured with changes in sales.

The result is that wage adjustment is asymmetric and therefore, although wage change distributions are smooth and spikeless in France, we conclude that nominal wages are rigid.

Figure 9: Long-run equilibrium in the AD - AS model.

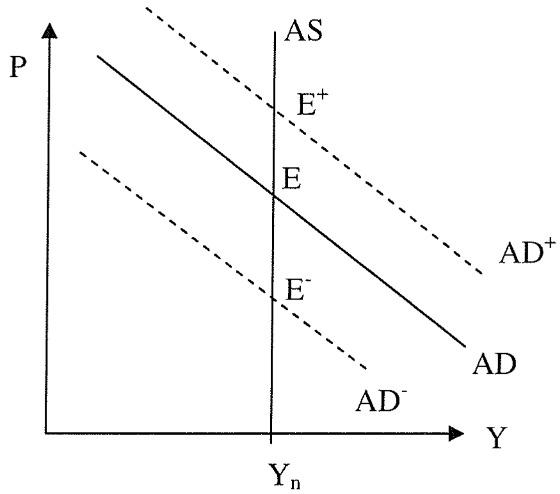


Figure 10: Short and medium-run adjustment in the AD – AS model

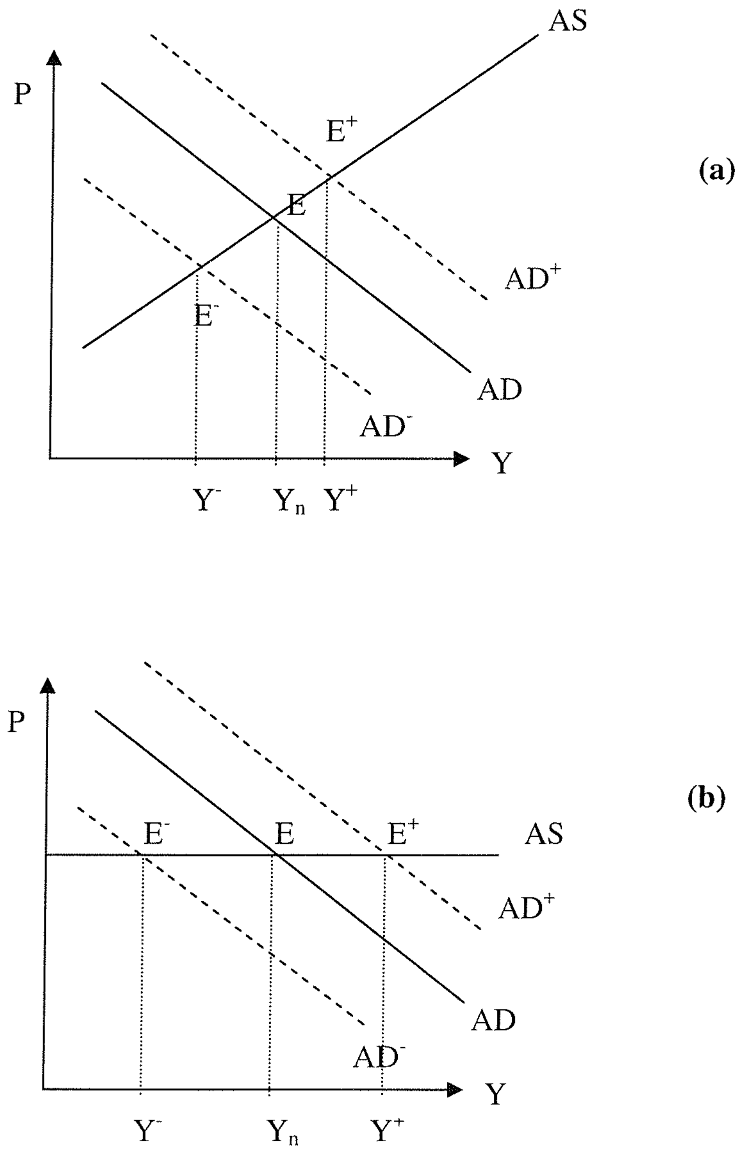


Table 25: Estimation on DADS data, annual wages

	Wage cuts, logit "Mundlak" model		Wage cuts, conditional logit	Wage growth, OLS with the Mundlak method	
	between	within		Between	Within
Sales growth*indicator of increase in sales (1)	-0.621** (0.30)	-0.320** (0.021)	-0.275** (0.022)	0.0347** (0.0009)	0.0110** (0.0006)
Sales growth*indicator of decrease in sales (2)	-0.915** (0.28)	-0.251** (0.018)	-0229** (0.020)	0.0234** (0.0009)	0.0031** (0.0006)
Test of the difference (1)-(2)	0.294**	-0.069*	-0.046	0.0114**	0.0079**
Reduction of hours (qualitative) Growth of hours (continuous)	0.610** (0.010)	0.749** (0.006)	0.741** (0.007)	0.2827** (0.0028)	0.2039** (0.0012)
Log. of initial wage	0.332** (0.010)	13.483** (0.050)	18.403** (0.074)	-0.0057** (0.0003)	-0.6689** (0.0011)
Executives (ref.blue collars)	-0.694** (0.013)	-0.911** (0.031)	-0.786** (0.036)	0.0223** (0.0004)	0.0495** (0.0009)
Intermediate professions (ref.blue collars)	-0.333** (0.008)	-0.353** (0.021)	-0.344** (0.024)	0.0102** (0.0002)	0.0188** (0.0006)
White collars (ref.blue collars)	-0.207** (0.009)	-0.032 (0.026)	-0.039 (0.029)	0.0039** (0.0003)	0.0018 (0.0008)
Women	-0.012 (0.007)			-0.0013** (0.0002)	
Age	0.020** (0.000)			-0.0009** (0.0000)	
Tenure	0.002** (0.000)			-0.0003** (0.0000)	
n.employees<20 (ref. 1000≤employees)	0.186** (0.009)			-0.0032** (0.0003)	
20≤n.employees<200 (ref. 1000≤employees)	0.131** (0.007)			-0.0010** (0.0002)	
200≤n.employees<1000 (ref. 1000≤employees)	0.029** (0.008)			0.0004 (0.0002)	
Unemployment rate in the local market	1.108** (0.076)			-0.0173** (0.0023)	

Source : panel DADS 1994-2000, BRN (size of firm, sector of activity and sales growth) and RP 99 (local unemployment rate) ; coefficients of temporal dummies and their means, as well as coefficients of sector dummies in NAF 16 are not reported.

1 091 002 observations, of which 257 622 annual wage cuts

** and * indicate 1% and 5% of significance, respectively. The other coefficients are not significant at least at 5%.

Notes : in the continuous case, we estimate the following equation:

$\Delta \log W_{it} = \left(\alpha^{+} 1_{\Delta \log CA_{j(i)t} \geq 0} + \alpha^{-} 1_{\Delta \log CA_{j(i),t} < 0}\right) \times \Delta \log CA_{j(i)t} + \lambda_{it}$ where W_{it} represents the annual wage, $CA_{j(i)t}$ indicates firms' sales, and λ_{it} represents the other explicative variables and residuals. Coefficients (1) and (2) correspond respectively to α^{+} and α^{-}. if wage dynamics reflect productivity changes, and that is linked to the activity (productivity cycle), the two coefficients should be positive. Following our definition, α^{+} is bigger than α^{-} if there is downward wage rigidity.

In the discrete case, the dependent variable is no longer wages growth rate, but its propensity to be negative. The two coefficients α^{+} and α^{-} are negative under the previous conditions, and α^{+} is bigger in absolute value than α^{-} in case of downward wage rigidity.

The line « test of the difference (1) - (2) » indicates the value and significance of α^{+} - α^{-}.

Table 26: Heterogeneity of coefficients in the DADS, growth rate of annual wages

Log. of initial wage	-0.691** (0.001)		
Growth rate of hours	0.211** (0.001)		
Growth rate of sales	rises	cuts	Test: rises-cuts
	0.011** (0.002)	-0.001 (0.001)	0.013** (0.002)
INTERACTED WITH:			
SMIC≤average wage of the period<1.3 SMIC	-0.002 (0.002)	-0.002 (0.002)	0.000 (0.003)
1.3 SMIC≤average wage of the period<2 SMIC	Ref.	Ref.	Ref.
2 SMIC≤average wage of the period	-0.002 (0.002)	0.010** (0.002)	-0.008** (0.003)
Women	-0.006** (0.001)	0.002 (0.001)	-0.008** (0.002)
Men	Ref.	Ref.	Ref.
Executives	0.008** (0.002)	-0.011** (0.002)	0.019** (0.004)
Intermediate profession	-0.005** (0.002)	-0.002 (0.002)	-0.003 (0.003)
White collars	0.002 (0.002)	-0.006** (0.002)	0.008** (0.003)
Blue collars	Ref.	Ref.	Ref.
age<35	0.007** (0.002)	0.001 (0.002)	0.006* (0.002)
35≤age<45	Ref.	Ref.	Ref.
45≤age	-0.006** (0.001)	0.004** (0.001)	-0.010** (0.002)
Number of employees<20	0.011** (0.002)	0.018** (0.002)	-0.008** (0.003)
20≤Number of employees<200	0.006** (0.001)	0.009** (0.001)	-0.002 (0.002)
200≤Number of employees	Ref.	Ref.	Ref.
Services	-0.007** (0.001)	-0.006** (0.001)	-0.001 (0.002)
Industry and agriculture	Ref.	Ref.	Ref.
Local unemployment rate (in deviation from the average unemployment rate of the sample)	-0.022 (0.020)	0.077** (0.018)	-0.099** (0.031)

** and * indicate 1% and 5% significance, respectively. The other coefficients are not significant below 5%.

Notes: we estimate the following equation :

$\Delta \log W_{it} = \left[A^{+} \times \left(1+\beta^{+} 1_{sex}\right) \times 1_{\Delta \log CA_{j(i)t} \geq 0} + A^{-} \times \left(1+\beta^{-} 1_{sexn}\right) \times 1_{\Delta \log CA_{j(i)t} < 0}\right] \times \Delta \log CA_{j(i)t} + \lambda_{it}$, where for notational simplicity we have ignored the other terms of interaction different from sex. The coefficients « rises » and « cuts » of the line « growth rate of sales » correspond respectively to α^{+} and α^{-}. The column « test : rises - cuts » of the same line measures the asymmetry α^{+} - α^{-} for the individual of reference : man, blue collar, working in a firm of more than 200 employees, aged between 35 and 45, whose wage is in the interval [1,3 SMIC, 2 SMIC], and who works in an area where the unemployment rate is equal to the national average.

β^{+} et β^{-} correspond to the first two columns of the line « Women ». The third column of the same line measures the asymmetry associated to being woman, ceteris paribus. Coefficients are additive : we can measure the asymmetry for all combinations of individual characteristics (sex, age, etc.) as deviations from the reference individual. Therefore, an individual who is different from the reference person only for sex and for being an executive will have an symmetry of 0,013 - 0,008 + 0,019.

Table 27: Estimation on Revenus fiscaux and Enquête Emploi on the sample of Revenus fiscaux, annual wages

	enquête Revenus fiscaux		enquête Emploi	
	Wage cuts	Wage rises	Wage cuts	Wage rises
Sales increase*indicator of sales rise (1)	-0,305** (0,080)	0,0100** (0,0027)	-0,086 0,094	0,0087 (0,0055)
Sales increase*indicator of sales cuts (2)	-0,154 (0,107)	-0,0026 (0,0042)	-0,024 0,140	-0,0071 (0,0086)
Hours reduction (qualitative) Hours growth (continuous)	-0,068 (0,043)	0,0165** (0,0064)	0,055 (0,054)	0,055** (0,0131)
Log of initial wage	1.245** (0,078)	-0,0703** (0,0029)	1,419** (0,103)	-0,2834** (0,0055)
Executives (ref. blue-collars)	-1,088** (0,095)	0,0647** (0,0034)	-1,248** (0,122)	0,2085** (0,0070)
Intermediate professions (ref. blue-collars)	-0,499** (0,0579	0,0267** (0,0021)	-0,475** (0,072)	0,0794** (0,0043)
White collars (ref. blue-collars)	-0,303** (0,067)	0,0109** (0,0025)	-0,263** (0,085)	0,0264** (0,0051)
Women	0,059 (0,053)	-0,0085** (0,0019)	0,184** (0,066)	-0,0399** (0,0039)
Age	0,005 (0,003)	-0,0004** (0,0001)	0,009 (0,004)	0,0013** (0,0002)
Tenure	-0,007** (0,003)	-0,0004** 80,0001)	-0,009** (0,003)	0,0009** (0,0002)
Number of employees<20 (ref. 1000≤employees)	0,194** (0,066)	-0,0111** (0,0025)	0,392** (0,083)	-0,0198** (0,0050)
20≤Number of employees<200 (ref. 1000≤employees)	0,192** (0,052)	-0,0089** (0,0020)	0,296** (0,068)	-0,0179** (0,0041)
200≥Number of employees (ref. 1000≤employees)	0,045 (0,057)	-0,0005 (0,0021)	0,027 (0,072)	-0,0013 (0,0043)
Local unemployment rate	0,606 (0,587)	-0,0286 (0,0220)	0,180 (0,743)	-0,0465 (0,0451)
Answer given by a third person	-0,083* (0,041)	0,0023 (0,0015)	0,166** (0,052)	-0,0077** (0,0032)
Contract different from permanent	0,592** (0,220)	-0,0063 (0,0091)	0,723** (0,255)	-0,0442** (0,0168)
Secondary school diploma (ref. no diploma)	-0,454** (0,082)	0,0154** (0,0030)	-0,306** (0,101)	0,0708** (0,0061)
Other diploma (ref. no diploma)	-0,126** (0,047)	0,0014 (0,0018)	-0,142** (0,060)	0,0268** (0,0036)
Bonuses cuts	0,159** (0,043)	-0,0051** (0,0016)	2,842** (0,050)	-0,0663** (0,0033)
Jobs conditions	0,037** (0,016)	-0,0019** (0,0006)	0,090** (0,021)	-0,0036** (0,0013)
Difference (1)-(2)	-0,151	0,0126*	-0,062	0,0157

Source : enquête Revenus fiscaux 1996-1999, enquête Emploi 1996-1999 (individuals in Revenus fiscaux), BRN (firm size, sector and sales growth) and RP 99 (local unemployment rate) ; the coefficients of sector indicators, time dummies and their averages are not reported.

Revenus fiscaux : 15 941 observations of which 3 794 annual wage cuts and 4 118 hourly wage cuts.

Enquête Emploi : 13 687 observations of which 3484 annual wage cuts and 3 782 hourly wage cuts.

** and * indicate 1% and 5% significance respectively. The other coefficients are not significant at least at 5%.

Table 28 : Estimation on the enquête Emploi, annual wages

	Wage cuts, logit « Mundlak » method		Wage rises, OLS Mundlak method	
	Between	Within	Betweeen	Within
Sales increase*indicator of sales rise (1)	-0,122* (0,062)	-0,79 (0,112)	0,0140** (0,0032)	0,0038 (0,0058)
Sales increase*indicator of sales cuts (2)	-0,095 (0,084)	0,136 (0,144)	0,0028 (0,0046)	-0,0046 (0,0076)
Hours reduction (qualitative) Hours growth (continuous)	0,176** (0,038)	0,058 (0,052)	0,0669** (0,0086)	0,0023 (0,0094)
Log of initial wage	1,092** (0,055)	20,981** (0,374)	-0,2004** (0,0027)	-1,2456** (0,0078)
Executives (ref. blue-collars)	-0,979** (0,067)	-0,843** (0,259)	0,1565** (0,0034)	0,0411** (0,0134)
Intermediate professions (ref. blue-collars)	-0,460** (0,040)	-0,387* (0,171)	0,0647** (0,0021)	0,0137 (0,0089)
White collars (ref. blue-collars)	-0,238** (0,045)	-0,140 (0,221)	0,0209 (0,0024)	0,0154 (0,0114)
Women	0,072* (0,034)		-0,0286** (0,0018)	
Age	0,010** (0,002)		0,0007** (0,0001)	
Tenure	-0,009** (0,002)		0,0004** (0,0001)	
Number of employees<20 (ref. 1000≤employees)	0,327** (0,044)		-0,0203** (0,0023)	
20≤Number of employees<200 (ref. 1000≤employees)	0,242** (0,036)		-0,0163** (0,0019)	
200≤Number of employees (ref. 1000≤employees)	0,019 (0,039)		-0,0038 (0,0020)	
Local unemployment rate	0,460 (0,389)		-0,0441* (0,0208)	
Answer given by a third person	0,147** (0,031)	0,206** (0,056)	-0,0012 (0,0016)	-0,0054 (0,0029)
Contract different from CDI	0,862** (0,100)	0,774* (0,325)	-0,0280** (0,0058)	-0,0271 (0,0169)
Secondary school diploma (ref. no diploma)	-0,483** (0,054)	-1,450 (2,335)	0,0427 (0,0028)	0,1220 (0,1333)
Other diploma (ref. no diploma)	-0,182** (0,031)	0,668 (1,713)	0,0151 (0,0017)	0,0476 (0,1088)
Bonuses cuts	2,894** (0,033)	2,521** (0,047)	-0,0609** (0,0019)	-0,0379** (0,0025)
Jobs conditions	0,086** (0,015)	0,033 (0,017)	-0,0047* (0,0008)	-0,0020* (0,0009)
Difference (1)-(2)	-0,027	-0,215	0,0111	-0,0046

Source : Enquête Emploi 1994-2000, BRN (firm size, sector and sales growth) and RP 99 (local unemployment rate) ; the coefficients of sector indicators, time dummies and their averages are not reported.

53 816 observations of which 15 119 cuts in annual wage.

** and * indicate 1% and 5% significance respectively. The other coefficients are not significant at least at 5%.

7 Conclusions

Although the state of recent research has led the ECB (2003) to conclude that 'the importance in practice of downward nominal rigidities is highly uncertain and the empirical evidence is not conclusive, particularly for the euro area', we think that the work carried out in this thesis has contributed to trying to improve our knowledge of wage rigidity characteristics in the EU countries. We have analysed DNWR in the EU under different perspectives. Our main results are the following:

- The descriptive analysis of wage change distributions from the ECHP shows that there is quite an high degree of nominal wage rigidity in Europe. However, wages are not completely downwardly flexible.

- The above result is even stronger when we estimate the extent of DNWR using a structural approach. In particular, measurement error explains the almost totality of nominal wage cuts observed, that are instead wage freezes. However, this is true when modelling measurement errors according to the classical assumptions.

- Institutional characteristics of the European labour markets seem to play a role in explaining the extent of DNWR observed and/or estimated. In particular, a robust hump - shaped relationship is found between EPL and DNW flexibility.

- Using the classical assumptions for measuring the extent of DNWR might be very distortionary. This is true in particular for France, were an appropriate validation study carried out on the national LFS shows that a certain flexibility of wages can be hidden by rounding behaviour of individuals.

- Using the classical definition of DNWR, based on some characteristics of the observed wage change distribution, can be restrictive. A more general concept of wage rigidity, testable on an appropriately constructed matched employer-employee data-set, can be introduced. This is based on the asymmetry of wage adjustments to firm-level shocks. We show that, although using the classical definition of DNWR we would conclude that wages are flexible in France, according to our more general definition there is DNWR in France.

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APPENDIX 1

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

Table A1.1 GROSS WAGE DISTRIBUTIONS, ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
1 GERMANY gsoep											
	wave										
	1	3524	3980.392	1966.489	239	40000	2300	2900	3600	4552	5978
	2	3607	4146.084	1877.778	239	41667	2500	3050	3800	4800	6200
	3	3548	4351.143	1950.848	160	25000	2510	3200	4000	5000	6500
	4	2582	4469.778	1993.163	277	22000	2600	3204	4008	5200	6700
	5	2692	4505.005	1976.017	158	21000	2600	3248	4100	5240	6900
	6	3234	4638.449	2068.169	158	23000	2700	3312	4200	5400	7093
	7	2229	4776.075	2199.472	158	29000	2700	3400	4350	5500	7315
	Total	21416	4376.856	2010.91	158	41667	2500	3200	4000	5027	6600
GERMANY echp											
	wave										
	1	2670	4485.885	2221.586	509	30623	2400	3100	4000	5300	7000
	2	2625	4669.872	2219.753	570	30623	2576	3300	4200	5500	7329
	3	2472	4785.555	2161.05	681	26293	2700	3410	4300	5598	7458
	Total	7767	4643.443	2205.041	509	30623	2500	3274	4200	5500	7250
2 FRANCE											
	wave										
	1	3306	13236.34	10141.49	192	250000	6875	8375	10875	15000	21250
	2	3243	11290.72	7071.602	1466	176000	6160	7480	9610	12980	17600
	3	3239	11430.85	7018.594	1258	220000	6380	7700	9900	13200	17662
	4	2980	12701.26	8068.451	1200	240000	7004	8400	10800	14498	19529
	5	2328	13262.5	8471.855	1200	200000	7200	8800	11199	15100	20500
	6	2114	13424.8	7605.933	1738	90000	7440	8997	11400	15576	20897
	7	2068	13513.1	7521.908	1738	86697	7500	9000	11628	15600	21000
	Total	19278	12576.49	8139.763	192	250000	6750	8261	10670	14400	19800

Table A1.1 GROSS WAGE DISTRIBUTIONS, ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
3 UK bhps											
	wave										
	1	2027	1372.941	834.5237	82	15011	650	860	1200	1671	2251
	2	1997	1435.204	798.9011	258	11008	683	908	1275	1769	2333
	3	2080	1498.249	856.4084	173	14511	721	952	1326	1814	2417
	4	1489	1552.831	925.8946	264	15211	736	1000	1356	1871	2500
	5	1513	1599.672	856.4854	277	8676	771	1001	1401	1950	2666
	6	1488	1661.451	892.8996	330	9006	823	1050	1460	2039	2702
	7	1414	1727.405	978.4154	333	12009	833	1083	1501	2101	2819
	Total	12008	1533.367	879.6838	82	15211	733	964	1343	1866	2501
UK echp											
	wave										
	1	2068	1390.609	809.215	13	7500	650	857	1200	1692	2250
	2	2033	1469.375	865.1081	13	10000	693	900	1258	1783	2345
	3	1734	1540.268	884.6413	156	9999	737	961	1333	1842	2416
	Total	5835	1462.527	853.6964	13	10000	693	900	1257	1783	2333
4 ITALY											
	wave										
	1	3336	2411.472	1159.217	218	32000	1600	1864	2200	2600	3300
	2	3277	2469.82	1052.966	218	12500	1670	1910	2207	2700	3400
	3	3441	2573.818	1072.885	600	14100	1750	2000	2338	2800	3500
	4	1363	2702.058	1041.848	494	12500	1850	2100	2500	3000	3600
	5	2948	2786.834	1214.387	449	15833	1900	2150	2520	3000	3848
	6	2836	2858.17	1233.229	680	15833	1900	2200	2600	3100	3900
	7	2720	2940.145	1321.266	850	17000	1957	2300	2700	3200	4000
	Total	19921	2660.32	1178.816	218	32000	1773	2000	2400	2939	3700

Table A1.1 GROSS WAGE DISTRIBUTIONS, ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
5 SPAIN											
	wave										
	1	2543	196160.7	106255.6	18000	1020000	101169	126000	168561	235783	321382
	2	2496	216987.4	117962.9	18000	1300000	113700	140000	186466	262500	351000
	3	2314	227742.9	127216.4	40000	2000000	119651	145833	196927	272425	367713
	4	1222	225855.6	124228.6	40000	1416666	115000	143000	195975	280000	360000
	5	1319	232397.9	132901.8	25000	1869000	120000	149346	200000	284310	375000
	6	1250	242102.7	146355.7	30000	2333333	126000	155000	205000	290000	380650
	7	1243	249594.3	159617	57800	2721667	135000	162036	210857	300000	396372
	Total	12387	223043.3	128590.1	18000	2721667	115000	143061	190000	270040	358333
6 NETHERLANDS											
	wave										
	1	2197	5441.972	5688.697	36	137149	3104	3828	4695	5957	7573
	2	2178	5300.198	2191.676	1179	28761	3269	3924	4887	6085	7713
	3	2265	5366.177	2613.226	86	63822	3231	3913	4898	6205	7830
	4	1626	5543.836	2394.231	722	28761	3327	4081	5055	6394	8320
	5	1654	5756.897	2382.757	241	28761	3513	4307	5242	6680	8546
	6	1688	5764.688	2415.366	900	28761	3413	4277	5245	6700	8639
	7	1350	5854.489	2361.509	1036	31158	3545	4357	5383	6772	8613
	Total	12958	5542.891	3210.591	36	137149	3303	4048	5018	6362	8169

Table A1.1 GROSS WAGE DISTRIBUTIONS, ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
7 BELGIUM											
	wave										
	1	1630	81956.47	37094.88	3334	380000	47500	59000	75000	95000	122500
	2	1600	86178.38	36861.63	3334	335904	51511	62730	78234	98334	134500
	3	1523	88128.15	37080.43	11600	374440	54000	65000	79388	101000	135000
	4	541	90714.19	34536.57	33730	250000	56512	67833	85000	104000	130911
	5	634	95405.03	40091.54	30000	380000	58138	69423	87023	110000	142868
	6	27	86770.11	34928.14	40538	166667	49000	65000	81909	109500	145205
	7	19	108604.2	45576.61	55000	250000	65000	75000	110000	121739	185000
	Total	5974	86987.45	37373.99	3334	380000	52000	64235	79000	100000	134279
8 LUXEMBURG psell											
	wave										
	2	2432	86065.61	45653.35	849	416667	42000	54792	75000	105000	143000
	3	1719	87331.8	42094.57	1018	400000	45000	57000	78000	108000	140000
	4	1657	95146	49403.82	1018	500000	47000	60500	83833	116828	157083
	5	1716	96417.19	49101.84	2397	500000	48000	61667	85000	120000	159000
	6	1619	99275.55	51869.46	1027	508333	50000	62833	87333	121467	163250
	7	1569	103610.7	58553.23	21013	820000	51167	65000	90000	126000	170000
	Total	10712	93898.07	49691.69	849	820000	46400	60000	82083	115000	155000
LUXEMBOURG echp											
	wave										
	1	645	106994.8	54130.23	11223	500000	54000	69364	93009	130000	180000
	2	633	111046.4	55849.94	11223	476873	57090	72018	98000	135394	185000
	3	607	114687.9	57061.37	11223	493381	58000	74000	99950	142623	194000
	Total	1885	110832.7	55723.41	11223	500000	56000	70956	96703	136348	186085

Table A1.1 GROSS WAGE DISTRIBUTIONS, ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
9 IRELAND											
	wave										
	1	1604	1400.188	806.8371	173	10833	661	882	1213	1718	2300
	2	1572	1471.335	791.3946	217	10833	726	953	1300	1799	2394
	3	1389	1525.342	766.0596	217	7257	775	1000	1360	1850	2466
	4	742	1623.013	829.128	303	7272	780	1030	1456	2017	2600
	5	884	1691.344	884.3457	308	7893	853	1083	1495	2076	2770
	6	727	1746.992	897.0342	347	7820	835	1105	1560	2167	2835
	7	585	1909.978	1008.087	282	8820	975	1221	1686	2330	3033
	Total	7503	1567.955	847.5843	173	10833	766	997	1382	1936	2575
10 DENMARK no occ. sect.											
	wave										
	1	2323	19164.5	7596.455	170	90000	12600	15000	17500	21500	27800
	2	1814	20515.12	7918.554	170	98000	14000	16000	18500	23000	29000
	3	1685	21228.78	7775.298	4000	97000	14500	16500	19500	23500	30000
	4	1568	21962.14	7841.939	4000	106000	15000	17000	20000	24450	30800
	5	1467	22945.47	8254.171	4500	110000	15500	18000	21000	25000	33000
	6	1431	23894.62	8549.989	4800	81000	16000	18500	22000	26800	34800
	7	1396	24960.93	9607.527	4400	105000	16800	19000	22500	27500	36000
	Total	11684	21793.94	8382.39	170	110000	14500	16900	20000	24600	31500
12 FINLAND											
	wave										
	3	1741	11795.45	5614.91	656	60000	7000	8493	10000	13500	18000
	4	1711	12198.7	5700.596	656	60000	7500	8600	10600	14000	18500
	5	1628	12612.76	5736.877	3500	68500	8000	9000	11000	14500	19500
	6	674	12966.74	6165.617	4500	69000	7800	9400	11500	15000	20000
	7	585	13317.2	6333.033	4200	67000	8000	9950	12000	15000	20000
	Total	6339	12379.17	5817.056	656	69000	7500	8900	11000	14000	19000

Table A1.1 GROSS WAGE DISTRIBUTIONS, ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
13 AUSTRIA											
	wave										
	2	1788	26988.99	12849.61	1380	150000	15000	19000	24000	32000	42000
	3	1757	25871.11	12305.19	1380	170000	15000	18025	23000	30000	40000
	4	818	26289.89	12755.44	930	130000	15000	18200	23018	30000	40833
	5	1180	27179.67	12161.13	3085	105000	15889	19151	24445	32000	41700
	6	1168	27539.2	12693.44	2084	150000	16000	19674	24600	32000	42000
	7	1110	28443.25	13495.99	1207	132813	16093	20085	25000	32778	43000
	Total	7821	26982.07	12711.65	930	170000	15400	19000	24000	31800	41800
14 PORTUGAL											
	wave										
	1	2492	108571.1	75501.98	11884	900000	52132	64243	84000	120000	199000
	2	2480	114230.8	78435.9	13000	760000	55200	68850	88572	129045	205000
	3	2505	121691.8	86502.9	17000	900000	60000	72100	93000	135000	222000
	4	1662	127938.3	91480.32	13000	900000	64000	75000	96500	142000	240000
	5	1710	132541.4	97727.58	10000	900000	65000	79000	99850	144492	249900
	6	2157	135854	100994.5	2000	1400000	68000	80000	100000	150000	250000
	7	2196	141929.1	101902.8	10000	1350000	71685	85000	106509	155000	260000
	Total	15202	125160	90745.79	2000	1400000	60000	74600	95515	140000	230000
15 GREECE											
	wave										
	1	1588	220598.1	111759.1	16000	1781250	128281	158425	200000	250000	325000
	2	1558	241371.4	111955.4	62765	1500000	145000	175000	220000	275000	360000
	3	1644	265118.4	112769	50000	933333	153000	192622	243008	300000	400000
	4	989	307553.3	140485.9	60000	2189474	175000	220000	285000	360000	450000
	5	1235	327028.9	152751.2	80000	1492857	185000	230000	300000	380000	495000
	6	1324	344127.6	168757.5	93000	2077419	190000	242526	306469	400000	514286
	7	1357	349214.2	180105.1	60000	2500000	197333	250000	310684	409091	510329
	Total	9695	288786.1	148285.9	16000	2500000	155000	198000	255000	345000	450000

Table A.1.2: HOUR DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
1 GERMANY											
	1	3523	42.09197	6.291387	30	96	37	39	40	45	50
	2	3613	42.31525	6.302497	30	90	37	39	40	45	50
	3	3558	42.22288	6.584206	30	90	36	39	40	45	50
	4	2592	42.6115	6.766428	30	96	37	39	40	45	50
	5	2696	42.39206	6.490827	30	80	36	39	40	45	50
	6	3234	42.50618	6.630693	30	96	37	39	40	45	50
	7	2229	42.69448	6.77733	30	96	37	39	40	45	50
	Total	21445	42.37692	6.530525	30	96	37	39	40	45	50
GERMANY echp											
	1	2646	42.27816	7.673914	30	90	37	38	40	45	50
	2	2607	41.15228	6.501861	28	90	37	38	40	40	50
	3	2475	40.86263	6.344378	25	90	36	38	40	40	50
	Total	7728	41.44501	6.904933	25	90	37	38	40	42	50
2 FRANCE											
	1	3306	41.18451	6.530261	30	90	37	39	39	42	50
	2	3243	40.70336	6.036866	25	85	36	39	39	41	50
	3	3242	40.43708	6.96608	10	96	35	39	39	41	50
	4	2980	40.58624	7.188136	10	90	35	39	39	42	50
	5	2330	40.73004	6.497587	30	85	35	39	39	40	50
	6	2122	40.18944	5.980001	30	96	35	38	39	40	48
	7	2074	39.29508	5.779491	30	80	35	35	39	40	45
	Total	19297	40.51832	6.513937	10	96	35	39	39	41	50

Table A.1.2: HOUR DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
3 UK bhps											
	1	2027	43.81894	9.197851	30	96	35	37	41	48	56
	2	1999	43.76138	8.803129	30	96	36	37	41	48	55
	3	2080	43.81731	8.807507	30	96	35	37	42	48	55
	4	1489	43.59772	8.32384	30	96	36	37	41	48	55
	5	1515	43.59736	8.596964	30	92	35	37	41	48	55
	6	1488	43.44422	8.534174	30	96	35	37	41	48	55
	7	1416	43.32062	8.309237	30	96	35	37	41	47	55
	Total	12014	43.64858	8.698097	30	96	35	37	41	48	55
UK echp											
	1	2063	43.39263	9.385327	30	96	35	37	40	48	55
	2	2031	43.37765	9.002171	26	96	36	37	40	47	55
	3	1734	43.39792	8.481113	30	91	36	38	40	48	55
	Total	5828	43.38898	8.988611	26	96	36	37	40	48	55
4 ITALY											
	1	3335	39.94183	5.389138	30	90	36	36	40	40	48
	2	3280	39.92195	5.041094	24	90	36	36	40	40	47
	3	3451	39.4419	5.757493	16	72	36	36	40	40	46
	4	1365	38.91795	5.777828	18	70	36	36	40	40	45
	5	2963	38.80425	5.685384	15	70	36	36	40	40	45
	6	2835	38.52804	6.085793	15	77	36	36	40	40	45
	7	2718	38.52759	6.211978	15	80	35	36	40	40	45
	Total	19947	39.21938	5.719232	15	90	36	36	40	40	45

Table A.1.2: HOUR DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
5 SPAIN											
	1	2539	42.02757	7.4761	30	90	35	40	40	44	50
	2	2502	41.83253	7.266142	25	90	35	40	40	42	50
	3	2322	41.35616	6.53102	23	85	35	40	40	42	50
	4	1224	41.08007	7.342578	24	82	35	37.5	40	42	50
	5	1324	41.24169	7.417914	25	96	35	38	40	40	50
	6	1251	40.66427	6.510483	24	80	35	38	40	40	50
	7	1244	41.18328	6.9767	21	84	35	38	40	40	50
	Total	12406	41.46308	7.111915	21	96	35	39	40	42	50
6 NETHERLANDS											
	1	2197	40.20665	6.32413	30	80	34	38	40	40	48
	2	2178	40.18733	6.294107	21	80	34	38	40	40	48
	3	2263	39.99956	6.291652	25	85	33	38	40	40	50
	4	1625	40.08492	6.690846	25	80	32	36	40	40	50
	5	1654	40.06046	6.829527	23	80	32	36	40	40	50
	6	1689	39.97573	6.845104	26	80	32	36	40	40	50
	7	1353	39.6031	6.514314	30	87	32	36	39	40	50
	Total	12959	40.0402	6.516521	21	87	33	37	40	40	50
BELGIUM											
	1	2206	41.11333	7.074904	30	95	36	38	40	42	50
	2	1725	41.01217	6.420347	25	80	36	38	40	42	50
	3	1677	40.70543	6.048555	20	80	36	38	40	42	50
	4	1620	40.94444	6.581039	20	90	36	38	40	43	50
	5	1516	41.04815	6.423528	21	75	36	38	40	43	50
	6	1434	40.93515	6.581444	20	76	36	38	40	43	50
	7	1381	41.06734	7.05356	20	90	35	38	40	43	50
	Total	11559	40.97924	6.618553	20	95	36	38	40	43	50

Table A.1.2: HOUR DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
8 LUXEMBURG psell											
	2	2432	40.70354	3.933907	30	85	40	40	40	40	40
	3	1722	40.44251	3.103746	30	84	40	40	40	40	40
	4	1657	40.63549	3.725956	30	94	40	40	40	40	40
	5	1719	40.10297	2.415845	30	66	40	40	40	40	40
	6	1626	40.16482	2.441163	30	70	40	40	40	40	40
	7	1571	40.04074	2.487855	30	90	40	40	40	40	40
	Total	10727	40.37615	3.163976	30	94	40	40	40	40	40
LUXEMBURG echp											
	1	645	41.53643	4.853825	30	90	40	40	40	40	48
	2	636	40.80189	3.382455	30	65	40	40	40	40	45
	3	611	40.88871	3.483552	30	65	40	40	40	40	45
	Total	1892	41.08034	3.98608	30	90	40	40	40	40	45
9 IRELAND											
	1	1604	41.8884	7.652179	30	96	35	39	40	44	50
	2	1576	41.43591	7.237887	22	84	35	39	40	42	50
	3	1390	40.49424	7.201484	15	96	35	39	39	40	50
	4	743	40.30956	8.102677	16	84	32	38	39	42	50
	5	886	40.65801	7.81317	16	96	35	39	39	41	50
	6	727	40.18845	7.289676	18	77	33	38	39	40	50
	7	585	39.86325	7.099713	18	72	33	38	39	40	49
	Total	7511	40.91186	7.502522	15	96	35	39	40	42	50

Table A.1.2: HOUR DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
10 DENMARK											
	1	2323	39.57124	7.347677	30	96	35	37	37	40	48
	2	1808	38.42035	5.100568	27	80	35	37	37	37	45
	3	1685	38.44332	4.988576	28	80	35	37	37	38	45
	4	1573	38.45137	5.00226	25	70	35	37	37	37	45
	5	1470	38.65918	5.636375	26	90	35	37	37	38	45
	6	1429	38.49545	5.225665	28	85	35	37	37	39	45
	7	1395	38.65305	5.72192	28	85	35	37	37	39	45
	Total	11683	38.7237	5.756353	25	96	35	37	37	39	45
12 FINLAND											
	3	1741	39.93337	5.771226	25	90	36	38	38	40	45
	4	1714	39.73221	5.119134	25	80	36	38	38	40	45
	5	1628	39.94595	5.373296	25	90	36	38	38	40	45
	6	674	40.30712	5.909348	20	75	36	38	40	40	46
	7	585	40.08205	5.604863	20	76	36	38	38	40	47
	Total	6342	39.93567	5.500409	20	90	36	38	38	40	45
13 AUSTRIA											
	2	1788	41.17841	6.83428	4	85	38	39	40	40	50
	3	1757	41.3506	6.393246	20	80	38	39	40	40	50
	4	819	41.38095	6.253522	20	80	38	39	40	40	50
	5	1180	41.20424	6.127828	20	80	38	39	40	40	50
	6	1171	41.05636	6.151803	20	96	38	39	40	40	48
	7	1114	41.05655	6.032216	20	80	38	39	40	40	48
	Total	7829	41.20654	6.357524	4	96	38	39	40	40	50

Table A.1.2: HOUR DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
14 PORTUGAL											
	1	2478	41.85109	5.945454	30	96	35	40	40	45	48
	2	2469	41.54921	5.191053	25	84	35	40	40	45	45
	3	2505	41.46906	5.241677	18	84	35	40	40	44	45
	4	1661	40.50391	4.903848	16	84	35	40	40	42	45
	5	1713	40.16054	4.887929	16	96	35	39	40	40	45
	6	2168	40.16513	5.108644	15	96	35	39	40	40	45
	7	2199	39.97181	5.517787	15	96	35	39	40	40	45
	Total	15193	40.88857	5.354177	15	96	35	40	40	44	45
15 GREECE											
	1	1586	41.29382	6.491201	30	96	37	40	40	40	48
	2	1556	40.90617	5.51457	21	80	37	38	40	40	48
	3	1644	39.60462	6.499731	15	80	35	38	40	40	46
	4	989	39.49039	7.076731	15	80	35	38	40	40	48
	5	1235	39.72389	6.294298	15	70	35	38	40	40	48
	6	1327	39.69781	7.086506	15	80	32	38	40	40	48
	7	1357	39.9462	7.320773	15	72	32	38	40	40	48
	Total	9694	40.15401	6.625928	15	96	35	38	40	40	48

Table A.1.3: HOURLY GROSS WAGES IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
1 GERMANY gsoep											
	1	3523	23.71177	10.49893	1.244792	166.6667	13.8	17.5	22.1	27.5	35.4
	2	3606	24.48335	10.2343	1.422619	178.5714	14.6	18.4	22.9	28.6	36.1
	3	3548	25.87244	10.74008	0.888889	148.8095	15.1	19.3	24.3	30.1	38.7
	4	2580	26.22702	10.45579	1.73125	100	15.2	19.4	24.7	31.1	39.1
	5	2688	26.69378	10.73105	1.316667	100	15.6	19.8	25	31.4	40.5
	6	3234	27.36437	11.06965	1.128571	104.1667	15.8	20.1	25.6	32.1	40.6
	7	2229	28.02186	11.74244	1.128571	145	16	20.4	26	33.1	41.9
	Total	21408	25.87793	10.83416	0.888889	178.5714	15	19	24.1	30.5	38.7
GERMANY echp											
	1	2646	26.61434	11.7529	3.083333	168.5449	14.4	19	24.4	32.1	40.8
	2	2600	28.25367	12.00524	3.5625	168.5449	15.8	20.5	26.1	33.8	43
	3	2472	29.24224	12.1202	4.365385	172.9803	16.7	21.4	27	34.6	44.4
	Total	7718	28.00828	12.004	3.083333	172.9803	15.6	20.2	25.6	33.3	42.9
2 FRANCE											
	1	3306	79.57508	52.26381	1.371429	1388.889	42.8	52.1	68	91.7	125
	2	3243	68.67285	36.31589	7.33	733.3333	38.8	46.8	60.5	79.1	107
	3	3239	71.0818	37.48501	10.18333	785.7143	39.9	47.9	62.1	83.9	111
	4	2980	78.87498	43.58484	7.2	857.1429	43.3	52.6	68.7	92.3	123
	5	2328	80.37132	39.75493	11.14103	588.2353	45.8	55.1	71.1	94.8	123
	6	2114	82.76996	40.11844	11.14103	450	47.7	57	73.1	97.4	126
	7	2068	84.67496	39.01431	11.14103	433.485	48.6	59.1	75.4	101	129
	Total	19278	77.19943	42.19771	1.371429	1388.889	42.3	52.1	67.3	90.8	121

Table A.1.3: HOURLY GROSS WAGES IN THE ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
3 UK bhps											
	1	2027	7.831004	4.093142	0.581081	48.73701	3.97	5.17	6.94	9.47	12.8
	2	1997	8.185945	4.116623	1.28869	45.86666	4.16	5.41	7.32	9.96	13.1
	3	2080	8.538789	4.370876	1.147727	60.4625	4.33	5.63	7.63	10.3	13.5
	4	1489	8.870458	4.827297	1.304217	76.055	4.55	5.8	7.92	10.6	14.2
	5	1513	9.087485	4.386419	1.683333	39.43636	4.7	6.12	8.06	11.1	14.6
	6	1488	9.515566	4.631194	2.5	47.90425	4.95	6.41	8.4	11.7	15.1
	7	1414	9.950086	5.007668	1.922222	42.88929	5.11	6.61	8.76	12.2	15.7
	Total	12008	8.758121	4.509059	0.581081	76.055	4.44	5.76	7.8	10.6	14.1
UK echp											
	1	2063	7.981593	4.159907	0.083333	42.16071	4.04	5.21	7	9.85	12.8
	2	2028	8.391326	4.315698	0.108333	41.665	4.28	5.54	7.41	10.2	13.2
	3	1732	8.787307	4.390694	1.5625	43.40104	4.47	5.87	7.81	10.9	13.5
	Total	5823	8.363945	4.295429	0.083333	43.40104	4.24	5.47	7.41	10.3	13.1
4 ITALY											
	1	3335	15.13039	6.290163	1.159575	133.3333	10	11.8	13.9	16.9	21.3
	2	3275	15.53048	6.070601	1.09 74.846	10.1	12.2	14.4	17.4	21.7	
	3	3439	16.62454	6.691926	3 81.08334	10.9	12.5	15.3	18.5	23.9	
	4	1362	17.70395	7.170381	2.806818	76.53409	11.5	13.5	16.1	20	25
	5	2948	18.42319	8.494902	2.551136	111.1111	11.7	13.8	16.7	20.3	26.8
	6	2832 19.07	5 113.8889	12.2	14.3	17	20.8	28.7			
	7	2717	19.65673	9.290194	5.902778	118.0556	12.5	14.6	17.5	21.5	29.2
	Total	19908	17.29745	7.732335	1.09	133.3333	10.9	13	15.6	19.3	25

Table A.1.3: HOURLY GROSS WAGES IN THE ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
5 SPAIN											
	1	2539	1196.561	651.2102	142.0455	6785.714	584	750	1029	1467	2063
	2	2492	1327.179	700.403	210.2273	6111.111	650	835	1135	1661	2271
	3	2314	1405.021	760.4957	166.6667	7812.5	693	877	1225	1741	2336
	4	1222	1423.648	799.8512	200	7500	665	866	1212	1834	2456
	5	1319	1448.677	808.0922	156.25	7505.556	684	881	1239	1843	2500
	6	1250	1523.259	882.8628	187.5	9722.221	715	940	1281	1908	2552
	7	1243	1560.762	937.9863	357.1429	12690.36	777	969	1305	1938	2625
	Total	12379	1380.662	779.0667	142.0455	12690.36	656	854	1170	1725	2367
6 NETHERLANDS											
	1	2197	33.79634	33.76038	0.225	857.1813	20.2	24.4	29.6	37.2	45.4
	2	2178	32.86829	12.30995	3.925	205.4357	21.2	25	30.7	37.9	46.9
	3	2263	33.3683	14.327	0.56579	319.11	20.8	25.2	30.6	38.9	48.4
	4	1623	34.58475	13.93802	1.9625	239.675	21.5	25.9	32	40	49.9
	5	1652	36.08712	14.19456	1.9625	239.675	22.3	27.3	33.2	42.3	52.8
	6	1686	36.1614	14.09035	7.98125	199.7292	22.2	27.4	33.4	42.3	52.1
	7	1349	37.04003	13.85746	8.09375	222.5571	23.1	28.2	34.4	42.9	53.7
	Total	12948	34.60243	18.77833	0.225	857.1813	21.4	25.8	31.7	39.9	49.7
BELGIUM											
	1	2206	486.208	204.1226	20.8375	2125.203	280	354	450	572	731
	2	1725	526.4468	207.2794	20.8375	2150.5	321	388	486	614	790
	3	1672	536.8367	205.6091	76.31579	2530	329	406	493	625	786
	4	1614	556.7084	220.9096	208.3333	2745.054	344	411	510	647	822
	5	1516	569.7798	227.0359	160.7143	3128.546	347	427	523	665	833
	6	1432	580.9173	229.4669	152.7778	2810.954	358	428	531	680	868
	7	1381	592.9389	234.588	80.24722	2810.954	367	438	545	688	875
	Total	11546	544.8919	220.0811	20.8375	3128.546	327	403	500	640	813

Table A.1.3: HOURLY GROSS WAGES IN THE ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
8 LUXEMBURG psell no occ. se											
	2	2432	529.4474	277.1506	6.3625	2604.169	260	338	469	656	873
	3	1719	539.5678	254.9007	6.3625	2187.5	281	353	480	669	870
	4	1656	579.47	282.2289	6.3625	2812.5	292	375	512	719	938
	5	1716	601.425	304.4395	19.975	3125	300	381	531	750	988
	6	1619	617.2197	319.7152	6.41875	3177.081	309	389	546	759	1016
	7	1569	646.51	356.6808	131.3313	5125	320	406	563	791	1063
	Total	10711	580.7519	301.4123	6.3625	5125	288	375	510	716	956
LUXEMBURG echp											
	1	645	645.1665	316.1484	46.7625	2287.944	329	422	564	788	1068
	2	633	676.6417	326.6344	43.16539	2384.365	350	438	606	836	1113
	3	607	702.8307	335.9088	43.16539	2118.75	355	446	619	870	1178
	Total	1885	674.3049	326.8087	43.16539	2384.365	344	434	599	833	1125
9 IRELAND											
	1	1604	8.503088	4.849792	1.054878	60.18333	3.92	5.42	7.31	10.5	14.2
	2	1572	9.054979	4.981249	1.205556	57.62234	4.29	5.78	7.84	11	15.4
	3	1389	9.762804	5.430318	1.334615	45.35625	4.71	6.09	8.43	11.9	16.8
	4	742	10.48479	5.973421	1.577273	46.04688	4.82	6.25	8.76	12.9	18.8
	5	884	10.74806	6.028237	1.815	56.81818	5.19	6.6	9.01	13.2	19.4
	6	727	11.32787	6.529724	1.445833	59.24242	5.15	6.89	9.41	14.1	20.4
	7	585	12.3513	6.947276	2.136364	60.71429	6.18	7.8	10.3	14.8	20.9
	Total	7503	9.88615	5.719932	1.054878	60.71429	4.6	6.11	8.37	12	17.5

Table A.1.3: HOURLY GROSS WAGES IN THE ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
10 DENMARK no occ	1										
	2	2323	120.7284	38.16145	1.148649	447.9167	84.5	100	115	135	167
	3	1808	131.931	40.61795	1.0625	480.7692	94.6	108	123	146	181
	4	1684	137.171	42.84048	27.02703	608.1081	98.8	111	128	150	188
	5	1568	142.5943	45.77096	27.02703	716.2162	101	115	135	156	194
	6	1467	147.7298	45.86184	26	550	105	120	139	166	200
	7	1427	154.4384	46.57871	32.43243	425	108	125	145	173	213
	Total	1395	160.6575	52.36629	29.72973	575	113	128	149	176	223
		11672	140.0606	46.02176	1.0625	716.2162	97.3	111	132	156	194
12 FINLAND											
	3	1741	74.0505	33.20782	2.981818	428.5714	46.1	54.1	65.8	84.5	113
	4	1711	76.85574	35.67963	4.1	550	48	55.9	67.7	87.5	117
	5	1627	78.70189	32.8593	22.5	428.5714	50	58.3	69.1	91.2	118
	6	673	80.7361	36.63092	26.11111	492.8571	49.3	59.2	72.5	92	125
	7	585	83.21443	37.24471	27.63158	478.5714	50.6	62.5	75	93.8	125
	Total	6337	77.55814	34.66347	2.981818	550	48	56.3	68.8	88.8	118
13 AUSTRIA											
	2	1788	165.3744	87.6008	8.625	2250	96.2	118	146	194	254
	3	1757	156.3436	66.45313	8.625	607.1429	93.8	113	141	186	239
	4	818	158.8432	68.08136	5.8125	681.25	96.2	113	144	189	244
	5	1180	164.2754	64.42445	14.94231	583.3333	101	119	150	194	249
	6	1168	166.6825	67.52747	10.42	681.8182	101	123	150	193	250
	7	1110	172.4769	69.73419	7.017442	687.5	105	128	156	200	261
	Total	7821	163.7001	72.47731	5.8125	2250	98.7	119	147	192	250

Table A.1.3: HOURLY GROSS WAGES IN THE ECHP, continued

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
14 PORTUGAL											
	1	2478	666.6411	486.5169	66.02222	6000	306	375	500	750	1257
	2	2463	707.0649	523.6415	108.3333	6333.333	325	398	523	807	1346
	3	2505	758.6272	583.5639	118.0556	6683.333	341	421	556	844	1469
	4	1662	810.9132	628.2278	102.0833	7000	378	452	594	889	1630
	5	1710	853.2385	696.0592	62.5	7500	400	482	609	920	1667
	6	2159	870.6705	696.3456	16.66667	8500	419	500	625	938	1694
	7	2195	928.5321	781.13	62.5	9527.027	438	522	664	987	1786
	Total	15172	792.1487	634.2852	16.66667	9527.027	356	444	583	875	1508
15 GREECE											
	1	1586	1346.383	638.2056	114.2857	11066.95	750	969	1250	1563	2000
	2	1556	1487.05	663.4422	312.5	5357.143	844	1063	1351	1750	2244
	3	1642	1723.165	816.4783	312.5	6666.664	938	1190	1563	2000	2708
	4	989	2039.552	1067.871	442.7083	9895.833	1063	1313	1800	2381	3289
	5	1234	2123.459	1067.942	500	9330.356	1102	1389	1875	2500	3600
	6	1324	2268.144	1235.001	625	10819.89	1154	1449	1944	2655	3712
	7	1358	2284.02	1240.465	375	12500	1163	1475	1960	2688	3750
	Total	9689	1859.927	1031.635	114.2857	12500	938	1207	1594	2188	3063

Table A.1.4: GROSS WAGE CHANGES DISTRIBUTIONS IN THE ECHP

COUNTRY N		N	mean	sd	min	max	p10	p25	p50	p75	p90
1 GERMANY gsoep											
	2	3607	0.0464	0.117	-0.396	0.473	-0.0831	0	0.0362	0.103	0.192
	3	3548	0.0508	0.119	-0.336	0.511	-0.077	0	0.0395	0.105	0.194
	4	2582	0.0282	0.119	-0.405	0.461	-0.108	-0.0123	0.0187	0.08	0.167
	5	2692	0.0202	0.107	-0.405	0.399	-0.0959	-0.0161	0.0122	0.0667	0.145
	6	3234	0.0254	0.111	-0.37	0.43	-0.0976	-0.0165	0.0167	0.0727	0.155
	7	2229	0.0404	0.11	-0.357	0.496	-0.0741	0	0.0274	0.0889	0.177
	Total	17892	0.0362	0.115	-0.405	0.511	-0.087	0	0.0256	0.0882	0.177
GERMANY echp											
	2	2625	0.0408	0.119	-0.336	0.693	-0.0741	0	0.0148	0.0841	0.174
	3	2472	0.0394	0.126	-0.381	0.641	-0.08	0	0.00449	0.0834	0.182
	Total	5097	0.0401	0.123	-0.381	0.693	-0.077	0	0.0104	0.0835	0.178
2 FRANCE											
	2	3243	-0.128	0.19	-1.01	0.351	-0.331	-0.187	-0.105	-0.0226	0.0573
	3	3239	0.0244	0.129	-0.511	0.498	-0.12	-0.0339	0.019	0.0902	0.182
	4	2980	0.113	0.124	-0.379	0.629	-0.0245	0.0578	0.109	0.172	0.258
	5	2328	0.0521	0.108	-0.334	0.463	-0.0735	0	0.0435	0.111	0.185
	6	2114	0.0366	0.104	-0.324	0.446	-0.0774	-0.00576	0.0288	0.0792	0.156
	7	2068	0.041	0.117	-0.333	0.588	-0.0728	-0.00676	0.0275	0.0834	0.182
	Total	15972	0.0178	0.157	-1.01	0.629	-0.152	-0.0519	0.0239	0.102	0.187
3 UK bhps											
	2	1997	0.0522	0.138	-0.465	0.534	-0.0945	-0.00272	0.0389	0.115	0.23
	3	2080	0.0542	0.139	-0.39	0.587	-0.108	-0.00564	0.0388	0.118	0.226
	4	1489	0.0571	0.141	-0.47	0.577	-0.0961	0	0.0462	0.127	0.224
	5	1513	0.0578	0.141	-0.414	0.59	-0.105	-0.00265	0.0447	0.125	0.227
	6	1488	0.0588	0.148	-0.407	0.577	-0.116	-0.00529	0.0471	0.129	0.253
	7	1414	0.0551	0.144	-0.466	0.57	-0.102	-0.00676	0.0459	0.118	0.231
	Total	9981	0.0556	0.141	-0.47	0.59	-0.102	-0.00343	0.0427	0.121	0.233

Table A.1.4: GROSS WAGE CHANGES DISTRIBUTIONS IN THE ECHP

COUNTRY N		N	mean	sd	min	max	p10	p25	p50	p75	p90
UK echp											
	2	2033	0.0574	0.14	-0.409	0.607	-0.0815	0	0.0408	0.113	0.229
	3	1734	0.0469	0.141	-0.499	0.511	-0.105	0	0.0408	0.106	0.219
	Total	3767	0.0526	0.141	-0.499	0.607	-0.0896	0	0.0408	0.11	0.223
4 ITALY											
	2	3277	0.0323	0.134	-0.405	0.486	-0.125	-0.0298	0.00844	0.0991	0.201
	3	3441	0.0467	0.141	-0.405	0.575	-0.113	-0.00161	0.0377	0.118	0.223
	4	1363	0.0553	0.134	-0.314	0.539	-0.091	0	0.0408	0.123	0.223
	5	2948	0.0413	0.138	-0.428	0.511	-0.118	0	0.0342	0.105	0.211
	6	2836	0.0357	0.134	-0.405	0.511	-0.113	-0.000593	0.00637	0.0953	0.208
	7	2720	0.0318	0.13	-0.431	0.47	-0.112	0	0	0.0896	0.192
	Total	16585	0.0393	0.136	-0.431	0.575	-0.113	-0.0012	0.0267	0.105	0.211
5 SPAIN											
	2	2496	0.101	0.176	-0.457	0.665	-0.107	0	0.0948	0.203	0.318
	3	2314	0.0463	0.177	-0.512	0.619	-0.172	-0.0599	0.0405	0.149	0.267
	4	1222	0.0222	0.177	-0.462	0.529	-0.206	-0.0834	0.0185	0.124	0.248
	5	1319	0.0568	0.168	-0.419	0.598	-0.147	-0.0377	0.0463	0.156	0.268
	6	1250	0.0478	0.169	-0.439	0.569	-0.158	-0.0492	0.0462	0.149	0.258
	7	1243	0.0561	0.171	-0.492	0.565	-0.153	-0.0465	0.0513	0.157	0.281
	Total	9844	0.0601	0.176	-0.512	0.665	-0.154	-0.0426	0.0562	0.163	0.283
6 NETHERLANDS											
	2	2178	0.0396	0.1	-0.405	0.442	-0.0608	0	0.0337	0.0802	0.159
	3	2265	0.0345	0.0952	-0.288	0.559	-0.0589	0	0.0247	0.0708	0.139
	4	1626	0.0471	0.0983	-0.288	0.624	-0.0465	0	0.0353	0.0823	0.157
	5	1654	0.0467	0.11	-0.511	0.625	-0.0605	0	0.0391	0.09	0.174
	6	1688	0.0294	0.112	-0.341	0.702	-0.087	-0.0214	0.0196	0.0725	0.15
	7	1350	0.0505	0.107	-0.421	0.595	-0.0556	0	0.0402	0.0945	0.167
	Total	10761	0.0405	0.103	-0.511	0.702	-0.0616	0	0.0318	0.0815	0.154

Table A.1.4: GROSS WAGE CHANGES DISTRIBUTIONS IN THE ECHP

COUNTRY N		N	mean	sd	min	max	p10	p25	p50	p75	p90
BELGIUM											
	2	1730	0.062	0.156	-0.368	0.735	-0.087	0	0.029	0.107	0.247
	3	1673	0.0278	0.116	-0.429	0.463	-0.087	0	0.0121	0.0741	0.158
	4	1614	0.0421	0.123	-0.426	0.601	-0.0916	0	0.0267	0.0953	0.182
	5	1516	0.0319	0.118	-0.375	0.453	-0.101	0	0.0205	0.0823	0.167
	6	1432	0.0406	0.122	-0.435	0.463	-0.0852	0	0.0251	0.0953	0.182
	7	1381	0.036	0.117	-0.387	0.464	-0.0976	0	0.0235	0.0844	0.173
	Total	9346	0.0404	0.127	-0.435	0.735	-0.0903	0	0.0228	0.0892	0.182
8 LUXEMBURG psell											
	3	1719	0.00778	0.126	-0.453	0.48	-0.133	-0.0512	0.00171	0.0645	0.158
	4	1657	0.0878	0.132	-0.288	0.785	-0.0426	0.0194	0.0695	0.137	0.24
	5	1716	0.0487	0.124	-0.405	0.53	-0.082	-0.00209	0.0363	0.0989	0.195
	6	1619	0.0406	0.125	-0.38	0.515	-0.0897	-0.0144	0.0248	0.0953	0.185
	7	1569	0.0506	0.14	-0.413	0.606	-0.0832	-0.0113	0.0287	0.106	0.208
	Total	8280	0.0468	0.132	-0.453	0.785	-0.0896	-0.0132	0.032	0.103	0.199
LUXEMBURG echp											
	2	633	0.0377	0.097	-0.288	0.47	-0.0601	0	0.0233	0.0793	0.158
	3	607	0.0322	0.109	-0.281	0.405	-0.0953	-0.0119	0.0169	0.0829	0.16
	Total	1240	0.035	0.103	-0.288	0.47	-0.0767	0	0.0198	0.08	0.16
9 IRELAND											
	2	1572	0.0598	0.202	-0.618	0.807	-0.153	-0.00356	0.0392	0.139	0.318
	3	1389	0.0425	0.166	-0.545	0.595	-0.144	-0.018	0.0343	0.111	0.251
	4	742	0.0685	0.163	-0.462	0.61	-0.12	0	0.0562	0.133	0.28
	5	884	0.0763	0.147	-0.473	0.599	-0.0771	0.00909	0.0643	0.142	0.254
	6	727	0.0616	0.155	-0.472	0.694	-0.106	0	0.0464	0.114	0.246
	7	585	0.101	0.19	-0.578	0.762	-0.124	0	0.087	0.202	0.339
	Total	5899	0.0636	0.175	-0.618	0.807	-0.125	0	0.0479	0.136	0.281

Table A.1.4: GROSS WAGE CHANGES DISTRIBUTIONS IN THE ECHP

COUNTRY N		N	mean	sd	min	max	p10	p25	p50	p75	p90
10 DENMARK											
	2	1814	0.0295	0.0906	-0.288	0.452	-0.0606	0	0.0238	0.0628	0.134
	3	1685	0.0378	0.0882	-0.421	0.346	-0.0488	0	0.0299	0.077	0.143
	4	1568	0.0469	0.0883	-0.226	0.434	-0.0465	0	0.0381	0.0864	0.154
	5	1467	0.0483	0.0997	-0.357	0.507	-0.0541	0	0.0445	0.0912	0.163
	6	1431	0.0412	0.0955	-0.318	0.405	-0.0561	0	0.0357	0.08	0.145
	7	1396	0.0434	0.0958	-0.262	0.431	-0.0541	0	0.0345	0.08	0.163
	Total	9361	0.0407	0.093	-0.421	0.507	-0.0528	0	0.0336	0.0788	0.15
12 FINLAND											
	4	1711	0.0368	0.112	-0.424	0.47	-0.0741	0	0.0264	0.0839	0.166
	5	1628	0.0438	0.12	-0.405	0.5	-0.0953	0	0.0402	0.105	0.182
	6	674	0.0386	0.123	-0.405	0.569	-0.0953	0	0.0253	0.0953	0.182
	7	585	0.0477	0.125	-0.48	0.531	-0.087	0	0.0372	0.0969	0.201
	Total	4598	0.0409	0.118	-0.48	0.569	-0.0839	0	0.0336	0.0953	0.182
13 AUSTRIA											
	3	1757	-0.0401	0.156	-0.506	0.542	-0.241	-0.137	-0.00712	0.0401	0.146
	4	818	0.0209	0.128	-0.533	0.485	-0.118	-0.0186	0	0.0715	0.154
	5	1180	0.0363	0.11	-0.357	0.502	-0.0754	0	0.0217	0.077	0.169
	6	1168	0.0301	0.0973	-0.345	0.425	-0.0606	0	0.0174	0.0645	0.142
	7	1110	0.0386	0.0938	-0.266	0.386	-0.0465	0	0.0163	0.0741	0.154
	Total	6033	0.0112	0.127	-0.533	0.542	-0.147	-0.0328	0	0.0645	0.154
14 PORTUGAL											
	2	2480	0.0602	0.136	-0.405	0.582	-0.0828	0	0.0459	0.115	0.214
	3	2505	0.0682	0.127	-0.386	0.575	-0.0474	0.00942	0.045	0.116	0.223
	4	1662	0.0582	0.113	-0.405	0.56	-0.0313	0.00538	0.0355	0.096	0.197
	5	1710	0.0589	0.11	-0.347	0.504	-0.029	0.00141	0.0379	0.0998	0.202
	6	2157	0.0632	0.101	-0.329	0.511	0	0.0147	0.0392	0.0963	0.186
	7	2196	0.062	0.101	-0.288	0.499	0	0.00494	0.0392	0.0953	0.182
	Total	12710	0.0622	0.116	-0.405	0.582	-0.0335	0	0.0408	0.105	0.203

Table A.1.4: GROSS WAGE CHANGES DISTRIBUTIONS IN THE ECHP

COUNTRY N		N	mean	sd	min	max	p10	p25	p50	p75	p90
15 GREECE											
	2	1558	0.0925	0.199	-0.521	0.788	-0.151	-0.00837	0.0924	0.206	0.329
	3	1644	0.122	0.163	-0.416	0.598	-0.0531	0.0229	0.105	0.223	0.336
	4	989	0.14	0.173	-0.368	0.662	-0.0645	0.0315	0.126	0.251	0.357
	5	1235	0.0896	0.178	-0.496	0.693	-0.113	0	0.0645	0.182	0.336
	6	1324	0.0527	0.149	-0.464	0.55	-0.128	0	0.0368	0.125	0.241
	7	1357	0.0386	0.135	-0.393	0.514	-0.105	0	0	0.0953	0.215
	Total	8107	0.0883	0.171	-0.521	0.788	-0.105	0	0.069	0.182	0.31

Table A.1.5: CHANGES IN HOURS DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
1 GERMANY gsoep											
	2	3613	0.00727	0.085	-0.288	0.318	-0.0931	-0.0253	0	0.0408	0.108
	3	3558	-0.00649	0.0856	-0.288	0.288	-0.11	-0.0465	0	0.0253	0.0953
	4	2592	0.00712	0.0883	-0.288	0.336	-0.0976	-0.0247	0	0.0408	0.115
	5	2696	-0.00362	0.0873	-0.336	0.288	-0.105	-0.0408	0	0.0253	0.105
	6	3234	0.000239	0.0831	-0.265	0.288	-0.105	-0.0392	0	0.026	0.105
	7	2229	0.00128	0.0882	-0.316	0.329	-0.105	-0.0282	0	0.0408	0.108
	Total	17922	0.000865	0.0862	-0.336	0.336	-0.105	-0.0267	0	0.0267	0.105
GERMANY echp											
	2	2607	-0.0204	0.097	-0.405	0.274	-0.143	-0.0488	0	0	0.0645
	3	2475	-0.00578	0.0857	-0.362	0.318	-0.1	-0.026	0	0	0.08
	Total	5082	-0.0133	0.092	-0.405	0.318	-0.118	-0.0267	0	0	0.0741
2 FRANCE											
	2	3243	-0.00818	0.0834	-0.397	0.262	-0.105	0	0	0	0.0741
	3	3242	0.000195	0.0856	-0.368	0.405	-0.08	0	0	0	0.08
	4	2980	0.00166	0.0876	-0.405	0.368	-0.0828	0	0	0	0.0989
	5	2330	-0.00202	0.0822	-0.288	0.432	-0.0953	0	0	0	0.076
	6	2122	-0.00911	0.075	-0.288	0.248	-0.108	-0.0235	0	0	0.0606
	7	2074	-0.0189	0.0801	-0.27	0.288	-0.108	-0.0645	0	0	0.069
	Total	15991	-0.00526	0.0833	-0.405	0.432	-0.105	0	0	0	0.076
3 UK bhps											
	2	1999	0.00355	0.112	-0.405	0.432	-0.124	-0.0328	0	0.0488	0.134
	3	2080	0.000526	0.107	-0.396	0.368	-0.124	-0.0421	0	0.0435	0.127
	4	1489	0.00104	0.105	-0.344	0.357	-0.127	-0.0455	0	0.0435	0.134
	5	1515	0.000253	0.108	-0.405	0.386	-0.13	-0.0473	0	0.0455	0.127
	6	1488	-0.00384	0.108	-0.368	0.359	-0.14	-0.0513	0	0.0465	0.128
	7	1416	-0.0034	0.11	-0.39	0.348	-0.134	-0.0513	0	0.0476	0.13
	Total	9987	-4.13E-05	0.108	-0.405	0.432	-0.128	-0.0455	0	0.0465	0.131

Table A.1.5: CHANGES IN HOURS DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
UK echp											
	2	2031	0.00391	0.107	-0.301	0.405	-0.121	-0.0377	0	0.0455	0.134
	3	1734	0.000414	0.102	-0.336	0.357	-0.118	-0.0392	0	0.0408	0.118
	Total	3765	0.0023	0.105	-0.336	0.405	-0.121	-0.0392	0	0.0435	0.134
4 ITALY											
	2	3280	0.000815	0.0893	-0.336	0.318	-0.105	0	0	0	0.105
	3	3451	0.003	0.0935	-0.3	0.405	-0.105	0	0	0	0.105
	4	1365	0.00373	0.0915	-0.396	0.329	-0.105	0	0	0	0.105
	5	2963	-0.0137	0.0973	-0.405	0.288	-0.134	-0.0187	0	0	0.08
	6	2835 -0.00(	0.0888	-0.357	0.381	-0.105	0	0	0	0.1	
	7	2718	0.00191	0.0815	-0.387	0.354	-0.08	0	0	0	0.0822
	Total	16612	-0.0011	0.0907	-0.405	0.405	-0.105	0	0	0	0.105
5 SPAIN											
		2	2502 -0.0003	0.12	-0.405	0.405	-0.134	-0.0267	0	0.0267	0.141
	3	2322	-0.00319	0.125	-0.442	0.405	-0.154	-0.0299	0	0	0.154
	4	1224	-0.00457	0.132	-0.405	0.47	-0.182	-0.0513	0	0	0.154
	5	1324 -0.00(	0.122	-0.405	0.405	-0.134	-0.0267	0	0.0222	0.154	
	6	1251	-0.0104	0.125	-0.405	0.393	-0.182	-0.0274	0	0	0.134
	7	1244	0.00149	0.126	-0.405	0.434	-0.134	0	0	0.025	0.154
	Total	9867	-0.00255	0.124	-0.442	0.47	-0.154	-0.0267	0	0	0.154
6 NETHERLANDS											
	2	2178	0.00285	0.0834	-0.357	0.322	-0.069	0	0	0	0.0953
	3	2263	-0.00288	0.0783	-0.316	0.341	-0.087	0	0	0	0.069
	4	1625	-0.00749	0.0813	-0.336	0.274	-0.105	-0.0253	0	0	0.0619
	5	1654	-0.00477	0.0849	-0.329	0.318	-0.105	-0.0267	0	0	0.0953
	6	1689	-0.00273	0.0803	-0.318	0.288	-0.105	0	0	0	0.1
	7	1353	-0.00191	0.082	-0.307	0.288	-0.105	0	0	0	0.1
	Total	10762	-0.00256	0.0817	-0.357	0.341	-0.1	0	0	0	0.0896

Table A.1.5: CHANGES IN HOURS DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
7 BELGIUM											
	2	1725	-0.00287	0.0951	-0.446	0.333	-0.118	-0.0282	0	0.026	0.105
	3	1677	-0.00298	0.0939	-0.329	0.405	-0.105	-0.029	0	0.0267	0.105
	4	1620	0.00573	0.0965	-0.288	0.405	-0.105	-0.0267	0	0.0282	0.118
	5	1516	0.00147	0.0921	-0.351	0.329	-0.105	-0.0267	0	0.0267	0.105
	6	1434	-0.00023	0.0962	-0.357	0.405	-0.105	-0.0408	0	0.0282	0.105
	7	1381	-0.00083	0.091	-0.357	0.357	-0.105	-0.0274	0	0.0267	0.105
	Total	9353	7.46E-06	0.0943	-0.446	0.405	-0.105	-0.0274	0	0.0267	0.105
8 LUXEMBURG psell											
	3	1722	-0.00045	0.0554	-0.288	0.288	0	0	0	0	0
	4	1657	0.00256	0.0576	-0.223	0.295	0	0	0	0	0
	5	1719	-0.0123	0.06	-0.354	0.223	-0.0247	0	0	0	0
	6	1626	0.00122	0.0399	-0.223	0.223	0	0	0	0	0
	7	1571	-0.00256	0.0425	-0.288	0.223	0	0	0	0	0
	Total	8295	-0.00239	0.0522	-0.354	0.295	0	0	0	0	0
LUXEMBURG echp											
	2	636	-0.0116	0.0698	-0.318	0.288	-0.0953	0	0	0	0
	3	611	4.23E-05	0.0569	-0.223	0.223	0	0	0	0	0
	Total	1247	-0.00592	0.0641	-0.318	0.288	-0.0513	0	0	0	0
9 IRELAND											
	2	1576	-0.0103	0.113	-0.431	0.368	-0.151	-0.0361	0	0.0253	0.118
	3	1390	-0.00065	0.0987	-0.325	0.357	-0.118	-0.0253	0	0.0253	0.118
	4	743	0.00346	0.115	-0.405	0.431	-0.134	-0.0253	0	0.0247	0.143
	5	886	-0.0001	0.091	-0.329	0.431	-0.0953	0	0	0	0.105
	6	727	0.00237	0.104	-0.357	0.419	-0.105	-0.0247	0	0	0.118
	7	585	-0.0111	0.0991	-0.397	0.318	-0.118	-0.0408	0	0	0.105
	Total	5907	-0.0033	0.104	-0.431	0.431	-0.118	-0.0253	0	0.0183	0.118

Table A.1.5: CHANGES IN HOURS DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
10 DENMARK											
	2	1808	-0.0211	0.0795	-0.336	0.239	-0.127	-0.0267	0	0	0.0267
	3	1685	-0.003	0.071	-0.301	0.248	-0.078	0	0	0	0.078
	4	1573	-0.00087	0.0766	-0.301	0.304	-0.078	0	0	0	0.078
	5	1470	0.0053	0.075	-0.301	0.357	-0.0668	0	0	0	0.078
	6	1429	-9.14E-06	0.0747	-0.357	0.301	-0.069	0	0	0	0.078
	7	1395	0.00289	0.0696	-0.239	0.31	-0.069	0	0	0	0.078
	Total	9360	-0.00351	0.0752	-0.357	0.357	-0.078	0	0	0	0.0775
12 FINLAND											
	4	1714	-0.0003	0.0857	-0.37	0.295	-0.0822	0	0	0	0.0822
	5	1628	0.000454	0.0817	-0.336	0.301	-0.08	0	0	0	0.078
	6	674	-0.0015	0.0837	-0.405	0.318	-0.087	0	0	0	0.076
	7	585	-0.00523	0.0874	-0.318	0.37	-0.105	0	0	0	0.0822
	Total	4601	-0.0008	0.0842	-0.405	0.37	-0.0834	0	0	0	0.08
13 AUSTRIA											
	3	1757	0.00503	0.104	-0.405	0.431	-0.1	0	0	0.026	0.118
	4	819	-0.00333	0.0965	-0.405	0.431	-0.105	0	0	0	0.087
	5	1180	0.00137	0.0794	-0.318	0.336	-0.0723	0	0	0	0.0723
	6	1171	0.00179	0.0815	-0.318	0.405	-0.0513	0	0	0	0.0513
	7	1114	-0.000914	0.0816	-0.405	0.318	-0.0541	0	0	0	0.0513
	Total	6041	0.00146	0.0903	-0.405	0.431	-0.0822	0	0	0	0.0822
14 PORTUGAL											
	2	2469	-0.00434	0.0882	-0.357	0.288	-0.118	0	0	0	0.108
	3	2505	-0.0036	0.0724	-0.288	0.318	-0.0953	0	0	0	0.078
	4	1661	-0.021	0.0798	-0.383	0.333	-0.118	-0.0488	0	0	0.0225
	5	1713	-0.00965	0.0691	-0.336	0.251	-0.0953	-0.0247	0	0	0.0465
	6	2168	0.000225	0.0597	-0.223	0.288	-0.0488	0	0	0	0.0282
	7	2199	-0.00756	0.0666	-0.318	0.288	-0.0953	0	0	0	0
	Total	12715	-0.00686	0.0737	-0.383	0.333	-0.0953	0	0	0	0.0513

Table A.1.5: CHANGES IN HOURS DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
15 GREECE											
	2	1556	-0.00523	0.121	-0.47	0.442	-0.169	-0.0513	0	0.0127	0.134
	3	1644	-0.0218	0.122	-0.511	0.336	-0.182	-0.0488	0	0	0.087
	4	989	0.00802	0.125	-0.405	0.588	-0.128	0	0	0	0.134
	5	1235	0.00448	0.113	-0.419	0.405	-0.118	0	0	0	0.134
	6	1327	-0.00543	0.125	-0.47	0.419	-0.154	0	0	0	0.134
	7	1357	0.00666	0.106	-0.357	0.438	-0.118	0	0	0	0.134
	Total	8108	-0.00355	0.119	-0.511	0.588	-0.136	0	0	0	0.134

Table A.1.6 CHANGES IN HOURLY GROSS WAGES DISTRIBUTIONS IN THE ECHP

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
1 GERMANY gsoep											
	2	3606	0.0386	0.142	0.437	0.533	0.129	0.0325	0.0328	0.114	0.211
	3	3548	0.0582	0.143	0.395	0.598	0.11	0.0157	0.0513	0.134	0.228
	4	2580	0.0202	0.143	0.487	0.524	0.15	0.0479	0.0183	0.0956	0.189
	5	2688	0.0244	0.138	0.444	0.477	0.137	0.046	0.0202	0.095	0.192
	6	3234	0.0246	0.138	0.442	0.518	0.139	0.0451	0.0219	0.0966	0.194
	7	2229	0.0389	0.139	0.411	0.524	0.121	0.0336	0.0306	0.111	0.21
	Total	17885	0.0352	0.141	0.487	0.598	0.131	0.0356	0.0288	0.11	0.208
GERMANY echp											
	2	2600	0.0608	0.152	0.373	0.724	0.108	0.0128	0.0442	0.134	0.246
	3	2472	0.0466	0.15	0.431	0.709	0.121	0.0253	0.026	0.112	0.223
	Total	5072	0.0539	0.151	0.431	0.724	0.115	0.0225	0.0344	0.122	0.236
2 FRANCE											
	2	3243	0.118	0.21	1.01	0.506	0.351	0.195	0.103	0.00638	0.108
	3	3239	0.0235	0.161	0.621	0.654	0.161	0.0524	0.0201	0.103	0.208
	4	2980	0.112	0.161	0.606	0.827	0.0737	0.0331	0.108	0.192	0.294
	5	2328	0.0525	0.144	0.595	0.553	0.113	0.0121	0.0488	0.128	0.225
	6	2114	0.0461	0.132	0.357	0.511	0.108	0.0173	0.0336	0.108	0.207
	7	2068	0.0608	0.141	0.418	0.564	0.0933	0.00809	0.0433	0.134	0.244
	Total	15972	0.0233	0.181	1.01	0.827	0.18	0.0657	0.0282	0.121	0.226
3 UK bhps											
	2	1997	0.0478	0.173	0.608	0.629	0.148	0.0384	0.0399	0.14	0.258
	3	2080	0.0546	0.166	0.581	0.655	0.142	0.0337	0.0464	0.142	0.265
	4	1489	0.0564	0.17	0.564	0.62	0.136	0.0376	0.0464	0.147	0.272
	5	1513	0.0581	0.17	0.49	0.653	0.139	0.0318	0.0503	0.148	0.266
	6	1488	0.0635	0.172	0.486	0.675	0.143	0.0328	0.0513	0.161	0.286
	7	1414	0.0599	0.17	0.505	0.602	0.139	0.0349	0.049	0.154	0.272
	Total	9981	0.0561	0.17	0.608	0.675	0.142	0.035	0.047	0.148	0.268

Table A.1.6 CHANGES IN HOURLY GROSS WAGES DISTRIBUTIONS IN THE ECHP, cont.

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
UK echp											
	2	2028	0.0532	0.159	0.473	0.644	0.125	0.0267	0.0439	0.134	0.252
	3	1732	0.0463	0.155	0.512	0.561	0.134	0.0291	0.0448	0.126	0.229
	Total	3760	0.05	0.157	0.512	0.644	0.129	0.0276	0.0444	0.131	0.241
4 ITALY											
	2	3275	0.0318	0.16	0.462	0.552	0.16	0.0513	0.0202	0.114	0.238
	3	3439	0.0442	0.166	0.488	0.585	0.158	0.0408	0.0408	0.134	0.245
	4	1362	0.0533	0.158	0.389	0.573	0.127	0.039	0.0376	0.129	0.258
	5	2948	0.0572	0.163	0.436	0.575	0.131	0.0256	0.0445	0.14	0.274
	6	2832	0.0361	0.16	0.53	0.536	0.154	0.0359	0.0253	0.113	0.243
	7	2717	0.0294	0.153	0.488	0.511	0.146	0.0364	0.0101	0.1	0.223
	Total	16573	0.041	0.161	0.53	0.585	0.148	0.0392	0.03	0.121	0.245
5 SPAIN											
	2	2492	0.104	0.211	0.511	0.732	0.157	0.0154	0.102	0.228	0.361
	3	2314	0.0514	0.216	0.607	0.712	0.221	0.0799	0.0454	0.182	0.328
	4	1222	0.0262	0.223	0.619	0.69	0.276	0.109	0.026	0.165	0.312
	5	1319	0.0582	0.206	0.545	0.732	0.201	0.0694	0.0529	0.187	0.315
	6	1250	0.0568	0.202	0.5	0.73	0.194	0.069	0.0486	0.179	0.313
	7	1243	0.0534	0.206	0.541	0.716	0.195	0.0717	0.0481	0.173	0.313
	Total	9840	0.0635	0.213	0.619	0.732	0.204	0.0654	0.0594	0.192	0.33
6 NETHERLANDS											
	2	2178	0.0364	0.128	0.454	0.496	0.105	0.0185	0.0334	0.0953	0.19
	3	2263	0.0376	0.123	0.432	0.658	0.0907	0.0174	0.0291	0.09	0.182
	4	1623	0.0542	0.127	0.385	0.624	0.0824	0	0.0441	0.109	0.204
	5	1652	0.0519	0.138	0.485	0.648	0.0953	0.00269	0.0446	0.112	0.209
	6	1686	0.0324	0.136	0.486	0.693	0.105	0.035	0.0241	0.089	0.192
	7	1349	0.053	0.125	0.416	0.774	0.08	0	0.044	0.111	0.197
	Total	10751	0.0432	0.129	0.486	0.774	0.0954	0.0147	0.0357	0.101	0.195

Table A.1.6 CHANGES IN HOURLY GROSS WAGES DISTRIBUTIONS IN THE ECHP, cont.

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
BELGIUM											
	2	1725	0.0674	0.186	0.441	0.77	0.131	0.0253	0.0405	0.143	0.3
	3	1672	0.0302	0.151	0.519	0.533	0.132	0.0377	0.0241	0.109	0.202
	4	1614	0.038	0.159	0.49	0.636	0.141	0.0465	0.0301	0.118	0.235
	5	1516	0.0313	0.149	0.47	0.518	0.149	0.0386	0.0235	0.108	0.22
	6	1432	0.0405	0.153	0.546	0.541	0.134	0.0327	0.0358	0.119	0.223
	7	1381	0.0367	0.146	0.468	0.509	0.125	0.0313	0.029	0.116	0.214
	Total	9340	0.0411	0.159	0.546	0.77	0.136	0.0354	0.0298	0.118	0.23
8 LUXEMBURG psell											
	3	1719	0.00873	0.144	0.516	0.518	0.154	0.0572	0.00249	0.0741	0.182
	4	1656	0.0856	0.145	0.394	0.76	0.0592	0.00968	0.067	0.143	0.256
	5	1716	0.062	0.14	0.395	0.624	0.0871	0	0.044	0.118	0.224
	6	1619	0.0391	0.131	0.405	0.537	0.105	0.0168	0.0251	0.0983	0.19
	7	1569	0.0533	0.149	0.424	0.61	0.0864	0.0113	0.0328	0.11	0.223
	Total	8279	0.0495	0.144	0.516	0.76	0.1	0.016	0.0351	0.109	0.219
8 LUXEMBURG echp											
	2	633	0.0488	0.12	0.341	0.553	0.0745	0	0.0328	0.102	0.188
	3	607	0.0342	0.117	0.262	0.483	0.113	0.0217	0.0241	0.0896	0.172
	Total	1240	0.0416	0.119	0.341	0.553	0.0953	0.0104	0.0271	0.0953	0.182
9 IRELAND											
	2	1572	0.0693	0.222	0.637	0.771	0.173	0.0394	0.0462	0.175	0.355
	3	1389	0.0436	0.19	0.556	0.68	0.189	0.0455	0.0343	0.148	0.275
	4	742	0.0651	0.187	0.588	0.593	0.159	0.023	0.0569	0.164	0.303
	5	884	0.0753	0.172	0.473	0.633	0.127	0.0029	0.0621	0.165	0.287
	6	727	0.0574	0.178	0.529	0.798	0.154	0.0253	0.0495	0.145	0.274
	7	585	0.114	0.199	0.606	0.734	0.126	0	0.105	0.229	0.375
	Total	5899	0.0665	0.197	0.637	0.798	0.16	0.027	0.0522	0.167	0.31

Table A.1.6 CHANGES IN HOURLY GROSS WAGES DISTRIBUTIONS IN THE ECHP, cont.

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
10 DENMARK											
	2	1808	0.0527	0.116	0.352	0.546	0.0685	0	0.0364	0.104	0.197
	3	1684	0.0404	0.108	0.396	0.436	0.078	0	0.0328	0.0951	0.174
	4	1568	0.0468	0.105	0.317	0.439	0.0706	0	0.0403	0.0953	0.177
	5	1467	0.0414	0.121	0.473	0.507	0.0896	0	0.0412	0.0998	0.178
	6	1427	0.0406	0.113	0.331	0.483	0.0809	0	0.0352	0.0852	0.173
	7	1395	0.0408	0.112	0.396	0.472	0.0846	0.00257	0.0339	0.0903	0.175
	Total	9349	0.0441	0.113	0.473	0.546	0.078	0	0.0364	0.0953	0.182
12 FINLAND											
	4	1711	0.0365	0.139	0.424	0.571	0.118	0.0267	0.0266	0.0994	0.205
	5	1627	0.0442	0.138	0.424	0.534	0.118	0.0202	0.0417	0.114	0.215
	6	673	0.0412	0.146	0.439	0.708	0.126	0.0238	0.0307	0.105	0.209
	7	585	0.0526	0.149	0.574	0.575	0.111	0	0.0513	0.118	0.223
	Total	4596	0.042	0.141	0.574	0.708	0.118	0.0224	0.0347	0.108	0.211
13 AUSTRIA											
	3	1757	0.0448	0.186	0.63	0.592	0.274	0.154	0.0302	0.0507	0.173
	4	818	0.0234	0.153	0.511	0.637	0.148	0.0408	0	0.0963	0.194
	5	1180	0.0362	0.13	0.405	0.546	0.113	0.0174	0.0253	0.0953	0.189
	6	1168	0.0291	0.118	0.42	0.511	0.0996	0	0.0189	0.077	0.169
	7	1110	0.0385	0.117	0.377	0.526	0.0822	0	0.015	0.0834	0.182
	Total	6033	0.00992	0.152	0.63	0.637	0.169	0.0513	0	0.0776	0.182
14 PORTUGAL											
	2	2463	0.0662	0.164	0.454	0.634	0.118	0	0.0505	0.145	0.266
	3	2505	0.0717	0.149	0.43	0.588	0.077	0.00242	0.0488	0.14	0.26
	4	1662	0.0797	0.141	0.431	0.654	0.0645	0.0155	0.059	0.143	0.253
	5	1710	0.0699	0.131	0.383	0.548	0.0664	0	0.0488	0.128	0.232
	6	2159	0.0626	0.118	0.357	0.511	0.0476	0.00738	0.0408	0.108	0.21
	7	2195	0.07	0.124	0.327	0.553	0.0469	0	0.0426	0.119	0.229
	Total	12694	0.0696	0.14	0.454	0.654	0.0745	0	0.0487	0.132	0.241

Table A.1.6 CHANGES IN HOURLY GROSS WAGES DISTRIBUTIONS IN THE ECHP, cont.

COUNTRY		N	mean	sd	min	max	p10	p25	p50	p75	p90
15 GREECE											
	2	1556	0.0977	0.231	0.565	0.881	0.182	0.0363	0.0924	0.231	0.388
	3	1642	0.143	0.202	0.449	0.762	0.0914	0.0247	0.121	0.259	0.419
	4	989	0.13	0.211	0.514	0.75	0.133	0	0.121	0.265	0.394
	5	1234	0.0853	0.205	0.531	0.712	0.154	0.017	0.0645	0.2	0.357
	6	1324	0.0584	0.188	0.523	0.613	0.169	0.0266	0.0408	0.154	0.307
	7	1358	0.0336	0.171	0.534	0.56	0.168	0.0247	0	0.105	0.259
	Total	8103	0.0919	0.206	0.565	0.881	0.15	0.00198	0.0695	0.21	0.359

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

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

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

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

to give:

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

So

$$L1_{it}(\vartheta) = \Phi\left[\frac{x_{it}\beta + \left(\frac{\sigma_e^2}{\sigma_e^2+\sigma_m^2}\right)(y_{it} - x_{it}\beta)}{\sqrt{\frac{\sigma_e^2\sigma_m^2}{\sigma_e^2+\sigma_m^2}}}\right] \frac{1}{\sqrt{2\pi}\sqrt{\sigma_e^2+\sigma_m^2}} \exp\left[-\frac{(y_{it} - x_{it}\beta)^2}{2(\sigma_e^2+\sigma_m^2)}\right] \tag{13}$$

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

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

APPENDIX 3

In this appendix we provide all estimation results of the meta-analysis carried out in chapter 4 in the form of a (re-edited) Stata log-file.

REGRESSION FOR "OBSERVED FREQUENCIES"

Table 1: OLS , centr variable

Number of obs =	57
F(17, 39) =	33.79
Prob > F =	0.0000
R-squared =	0.7715
Root MSE =	3.7857

cmhgw	Coef.	Robust Std. Err.	t	P>|t|	[95% Conf.	Interval]
coord	87.87681	68.33832	1.29	0.206	-50.35048	226.1041
coord2	-17.39271	13.93998	-1.25	0.220	-45.58897	10.80356
centr	-23.31119	145.5361	-0.16	0.874	-317.6857	271.0634
centr2	5.259895	32.57863	0.16	0.873	-60.6366	71.15639
epl	124.1266	40.30465	3.08	0.004	42.60277	205.6505
epl2	-25.35058	7.86667	-3.22	0.003	-41.26242	-9.438734
pcov	-818.6494	457.9863	-1.79	0.082	-1745.014	107.7153
pcov2	429.517	258.4894	1.66	0.105	-93.3271	952.3611
uratest	3.421675	.7667482	4.46	0.000	1.870781	4.97257
uratest2	-.1295068	.0407487	-3.18	0.003	-.2119287	-.0470848
cpi	.6646835	.7896929	0.84	0.405	-.9326212	2.261988
labprod	1.102672	.5963884	1.85	0.072	-.1036375	2.308981
cons	165.7876	110.6732	1.50	0.142	-58.07005	389.6453

F-test of joint significance for the time-dummies
F(5, 39) = 2.77
Prob > F = 0.0312

Cameron & Trivedi's decomposition of IM-test

Source	chi2	df	p
Heteroskedasticity	57.00	56	0.4377
Skewness	8.33	17	0.9590
Kurtosis	5.44	1	0.0197
Total	70.77	74	0.5850

Table 2 IV (2SLS), centr variable

Total (centered) SS = 2446.484559		Number of obs =	57
Total (uncentered) SS = 50839.86501		F(17, 13) =	32.81
Residual SS = 560.2116369		Prob > F =	0.0000
		Centered R2 =	0.7710
		Uncentered R2 =	0.9890
		Root MSE =	3.1

cmhgw	Coef.	Robust Std. Err.	z	P>|z|	[95% Conf.	Interval]
uratest	3.357748	.6326118	5.31	0.000	2.117851	4.597644
uratest2	-.1267712	.0328381	-3.86	0.000	-.1911328	-.0624097
cpi	.3712938	.7662844	0.48	0.628	-1.130596	1.873184
labprod	1.143692	.625274	1.83	0.067	-.0818222	2.369207
coord	97.89499	56.76723	1.72	0.085	-13.36675	209.1567
coord2	-19.43507	11.57906	-1.68	0.093	-42.12962	3.259475
centr	-42.66694	126.6184	-0.34	0.736	-290.8344	205.5005
centr2	9.59089	28.37764	0.34	0.735	-46.02827	65.21005
epl	127.7709	33.7531	3.79	0.000	61.61603	193.9258
epl2	-26.04146	6.603987	-3.94	0.000	-38.98504	-13.09788
pcov	-849.7092	364.5987	-2.33	0.020	-1564.309	-135.1089
pcov2	445.9161	205.3533	2.17	0.030	43.43101	848.4011
cons	189.164	100.1408	1.89	0.059	-7.108282	385.4362

Hansen J statistic (overidentification test of all instruments): 30.528

Chi-sq(26) P-val = 0.24632

Summary results for first-stage regressions on the joint significance of excluded exogenous variables:

Variable	Shea Partial R2	Partial R2	F(30, 13)	P-value
uratest	0.9698	0.9558	9.38	0.0001
uratest2	0.9711	0.9463	7.64	0.0002
cpi	0.8854	0.9096	4.36	0.0036
labprod	0.7849	0.8336	2.17	0.0701

Wald Test of joint significance for the time-dummies

chi2(5) = 17.81
Prob > chi2 = 0.0032

Table 3: OLS, centrcd variable

Number of obs =	57
F(17, 39) =	33.93
Prob > F =	0.0000
R-squared =	0.7715
Root MSE =	3.7859

cmhgw	Coef.	Robust Std. Err.	t	P>|t|	[95% Conf.	Interval]
coord	88.90547	67.26917	1.32	0.194	-47.15927	224.9702
coord2	-17.65109	13.93874	-1.27	0.213	-45.84486	10.54268
centrcd	1.116455	6.866018	0.16	0.872	-12.77138	15.00429
centrcd2	-.1846422	1.187633	-0.16	0.877	-2.586856	2.217571
epl	123.9885	34.72349	3.57	0.001	53.75365	194.2234
epl2	-25.3051	6.810077	-3.72	0.001	-39.07978	-11.53042
pcov	-858.6048	399.5123	-2.15	0.038	-1666.695	-50.51503
pcov2	455.4445	231.005	1.97	0.056	-11.8073	922.6963
uratest	3.421222	.781884	4.38	0.000	1.839712	5.002731
uratest2	-.1293925	.0411529	-3.14	0.003	-.212632	-.0461529
cpi	.6675145	.7710456	0.87	0.392	-.8920725	2.227101
labprod	1.110604	.5953666	1.87	0.070	-.0936383	2.314847
_cons	152.8202	56.18277	2.72	0.010	39.17983	266.4606

F-test of joint significance for the time dummies
F(5, 39) = 3.38
Prob > F = 0.0124

Cameron & Trivedi's decomposition of IM-test

Source	chi2	df	p
Heteroskedasticity	57.00	56	0.4377

Skewness	8.19	17	0.9623
Kurtosis	5.45	1	0.0196
Total	70.64	74	0.5892

Table 4: IV (2SLS), centrcd variable

		Number of obs =	57
		F(17, 13) =	33.00
		Prob > F =	0.0000
Total (centered) SS	= 2446.484559	Centered R2 =	0.7710
Total (uncentered) SS	= 50839.86501	Uncentered R2 =	0.9890
Residual SS	= 560.3069902	Root MSE =	3.1

cmhgw	Coef.	Robust Std. Err.	z	P>|z|	[95% Conf.	Interval]
uratest	3.351488	.6430978	5.21	0.000	2.09104	4.611937
uratest2	-.1262875	.0326374	-3.87	0.000	-.1902556	-.0623195
cpi	.3705864	.7442686	0.50	0.619	-1.088153	1.829326
labprod	1.172343	.6242711	1.88	0.060	-.0512064	2.395892
coord	101.0449	60.07221	1.68	0.093	-16.69448	218.7843
coord2	-20.16592	12.45656	-1.62	0.105	-44.58032	4.248484
centrcd	2.123999	5.880177	0.36	0.718	-9.400937	13.64893
centrcd2	-.3555807	1.010567	-0.35	0.725	-2.336255	1.625094
epl	128.1558	31.82701	4.03	0.000	65.776	190.5356
epl2	-26.08273	6.24446	-4.18	0.000	-38.32165	-13.84382
pcov	-931.3104	364.6437	-2.55	0.011	-1645.999	-216.6218
pcov2	498.1992	210.3607	2.37	0.018	85.89976	910.4986
cons	166.6302	51.34788	3.25	0.001	65.99023	267.2702

Hansen J statistic (overidentification test of all instruments): 30.508

Chi-sq(26) P-val = 0.24712

Summary results for first-stage regressions on the joint significance of excluded exogenous variables:

Variable	Shea Partial R2	Partial R2	F(30, 13)	P-value
uratest	0.9759	0.9344	6.17	0.0006
uratest2	0.9696	0.9203	5.00	0.0018
cpi	0.8862	0.9056	4.16	0.0045
labprod	0.7958	0.8333	2.17	0.0707

Wald-test of joint significance for the time dummies

chi2(5) = 22.50

Prob > chi2 = 0.0004

Table 5: OLS, centrln variable

Number of obs =	57
F(17, 39) =	33.97
Prob > F =	0.0000
R-squared =	0.7715
Root MSE =	3.7856

cmhgw	Coef.	Robust Std. Err.	t	P>|t|	[95% Conf.	Interval]
coord	93.69011	205.3086	0.46	0.651	-321.5858	508.966
coord2	-18.54267	41.73313	-0.44	0.659	-102.9559	65.87057
centrln	.0466588	3.810782	0.01	0.990	-7.661375	7.754692
centrln2	-.0055872	.1946785	-0.03	0.977	-.3993617	.3881873
epl	127.752	76.88062	1.66	0.105	-27.7537	283.2578
epl2	-26.09799	14.86021	-1.76	0.087	-56.1556	3.959621
pcov	-876.7226	1154.874	-0.76	0.452	-3212.676	1459.231
pcov2	463.0472	669.4104	0.69	0.493	-890.9631	1817.057
uratest	3.364168	1.016563	3.31	0.002	1.307976	5.42036
uratest2	-.127964	.0443162	-2.89	0.006	-.2176019	-.0383261
cpi	.661487	.7490045	0.88	0.383	-.8535176	2.176492
labprod	1.130707	.5948318	1.90	0.065	-.0724535	2.333868
cons	159.19	145.2975	1.10	0.280	-134.702	453.0821

F-test of joint significance for the time dummies

F(5, 39) = 3.23
Prob > F = 0.0156

Cameron & Trivedi's decomposition of IM-test

Source	chi2	df	p
Heteroskedasticity	57.00	56	0.4377
Skewness	7.84	17	0.9699
Kurtosis	5.43	1	0.0198
Total	70.27	74	0.6015

Table 6: IV (2SLS), centrln variable

Number of obs	=	57
F(17, 13)	=	32.57
Prob > F	=	0.0000
Total (centered) SS	=	2446.484559
Centered R2	=	0.7708
Total (uncentered) SS	=	50839.86501
Uncentered R2	=	0.9890
Residual SS	=	560.6997058
Root MSE	=	3.1

cmhgw	Coef.	Robust Std. Err.	z	P>|z|	[95% Conf. Interval]	
uratest	3.240442	.8987277	3.61	0.000	1.478968	5.001916
uratest2	-.1243851	.0354795	-3.51	0.000	-.1939237	-.0548465
cpi	.3300647	.7037816	0.47	0.639	-1.049322	1.709451
labprod	1.228312	.6478732	1.90	0.058	-.0414966	2.49812
coord	117.8544	182.1144	0.65	0.518	-239.0833	474.7922
coord2	-23.42793	36.98294	-0.63	0.526	-95.91315	49.0573
centrln	.1816974	3.442757	0.05	0.958	-6.565983	6.929378
centrln2	-.0154912	.1756023	-0.09	0.930	-.3596653	.3286829
epl	139.0315	69.14724	2.01	0.044	3.505416	274.5576
epl2	-28.31192	13.37581	-2.12	0.034	-54.52803	-2.095815
pcov	-1010.955	1028.665	-0.98	0.326	-3027.102	1005.192

```
       pcov2 |   538.0795   596.0346     0.90   0.367    -630.1267
1706.286
        cons |   172.7379   130.5966     1.32   0.186    -83.22671
428.7025
------------------------------------------------------------------------------
-----
Hansen J statistic (overidentification test of all instruments):
30.970
                                                   Chi-sq(26) P-val =
0.22940
------------------------------------------------------------------------------
-----
Summary results for first-stage regressions on the joint significance
of excluded exogenous variables:

                Shea
Variable     Partial R2     Partial R2      F( 30,    13)   P-value
uratest        0.9684         0.8464            2.39         0.0494
uratest2       0.9770         0.8408            2.29         0.0579
cpi            0.8836         0.8901            3.51         0.0100
labprod        0.7950         0.8013            1.75         0.1436
------------------------------------------------------------------------------
-----
Wald test of joint significance for the time dummies

          chi2(  5) =    21.05
        Prob > chi2 =     0.0008
```

Table 7: OLS, centr variable

Number of obs =	69
F(18, 50) =	11.68
Prob > F =	0.0000
R-squared =	0.6375
Root MSE =	5.3607

cmhgw	Coef.	Robust Std. Err.	t	P>|t|	[95% Conf. Interval]	
coord	479.0566	77.91242	6.15	0.000	322.5649	635.5483
coord2	-97.41807	16.07642	-6.06	0.000	-129.7085	-65.12764
centr	471.5606	145.6013	3.24	0.002	179.1117	764.0094
centr2	-102.5793	33.03414	-3.11	0.003	-168.9303	-36.22823
epl	271.2955	40.08226	6.77	0.000	190.7879	351.8031
epl2	-52.41504	7.790837	-6.73	0.000	-68.0634	-36.76669
pcov	-3478.288	494.2615	-7.04	0.000	-4471.042	-2485.535
pcov2	1963.82	283.643	6.92	0.000	1394.107	2533.534
uratest	-.7336386	.5687788	-1.29	0.203	-1.876065	.4087873
uratest2	.0009811	.0270035	0.04	0.971	-.053257	.0552193
cpi	-1.402778	.9516591	-1.47	0.147	-3.314242	.5086852
labprod	-.6661562	.588359	-1.13	0.263	-1.84791	.5155976
cons	109.5303	110.9917	0.99	0.328	-113.4032	332.4637

F-test of joint significance for the time dummies

F(6, 50) = 1.15
Prob > F = 0.3473

Cameron & Trivedi's decomposition of IM-test

Source	chi2	df	p
Heteroskedasticity	69.00	68	0.4434
Skewness	31.78	18	0.0234
Kurtosis	2.64	1	0.1045
Total	103.41	87	0.1106

Table 8: IV (2SLS), centr variable

Total (centered) SS	= 3963.588238	Number of obs =	69
Total (uncentered) SS	= 12756.50145	F(18, 24) =	11.39
Residual SS	= 1480.785421	Prob > F =	0.0000
		Centered R2 =	0.6264
		Uncentered R2 =	0.8839
		Root MSE =	4.6

cmhgw	Coef.	Robust Std. Err.	z	P>|z|	[95% Conf.	Interval]
uratest	-.5668467	.523918	-1.08	0.279	-1.593707	.4600136
uratest2	-.0083588	.0234127	-0.36	0.721	-.0542468	.0375292
cpi	-1.614732	.8439872	-1.91	0.056	-3.268916	.0394527
labprod	.1628398	.6434412	0.25	0.800	-1.098282	1.423961
coord	497.1868	68.6733	7.24	0.000	362.5896	631.784
coord2	-101.1806	14.16035	-7.15	0.000	-128.9343	-73.4268
centr	462.9293	124.7866	3.71	0.000	218.3522	707.5065
centr2	-100.8746	28.44949	-3.55	0.000	-156.6346	-45.11459
epl	287.8591	35.48846	8.11	0.000	218.303	357.4152
epl2	-55.67246	6.890033	-8.08	0.000	-69.17667	-42.16824
pcov	-3600.039	431.1595	-8.35	0.000	-4445.096	-2754.982
pcov2	2029.33	247.0551	8.21	0.000	1545.111	2513.55
cons	133.4399	98.02226	1.36	0.173	-58.68017	325.56

Hansen J statistic (overidentification test of all instruments): 19.564
Chi-sq(26) P-val = 0.81174

Summary results for first-stage regressions on the joint significance of excluded exogenous variables:

Variable	Shea Partial R2	Partial R2	F(30, 24)	P-value
uratest	0.9561	0.9886	69.67	0.0000
uratest2	0.9591	0.9874	62.57	0.0000
cpi	0.7582	0.7723	2.71	0.0072
labprod	0.7209	0.7279	2.14	0.0298

Wald test of joint significance for the time dummies

chi2(6) = 9.91
Prob > chi2 = 0.1284

Table 9: OLS, centrcd variable

Number of obs =	69
F(18, 50) =	10.44
Prob > F =	0.0000
R-squared =	0.6277
Root MSE =	5.4327

cmhgw	Coef.	Robust Std. Err.	t	P>|t|	[95% Conf. Interval]	
coord	313.3751	94.4502	3.32	0.002	123.6663	503.0839
coord2	-62.76765	19.57302	-3.21	0.002	-102.0812	-23.45409
centrcd	-25.80382	6.263787	-4.12	0.000	-38.38501	-13.22264
centrcd2	4.709134	1.040018	4.53	0.000	2.620196	6.798072
epl	196.6248	41.29337	4.76	0.000	113.6847	279.565
epl2	-38.24397	8.051187	-4.75	0.000	-54.41526	-22.07269
pcov	-1755.594	513.4467	-3.42	0.001	-2786.882	-724.3056
pcov2	929.8687	296.5938	3.14	0.003	334.1425	1525.595
uratest	-.2539347	.6009264	-0.42	0.674	-1.460931	.9530614
uratest2	-.009563	.0286954	-0.33	0.740	-.0671994	.0480734
cpi	-1.220115	.9606225	-1.27	0.210	-3.149582	.7093525
labprod	-.6629335	.588878	-1.13	0.266	-1.84573	.5198626
cons	245.5388	69.97696	3.51	0.001	104.986	386.0917

F-test of joint significance for the time dummies

F(6, 50) = 0.78
Prob > F = 0.5863

Cameron & Trivedi's decomposition of IM-test

Source	chi2	df	p
Heteroskedasticity	69.00	68	0.4434

Skewness	32.53	18	0.0190
Kurtosis	2.68	1	0.1019
Total	104.21	87	0.1008

Table 10: IV (2SLS), centrcd variable

			Number of obs =	69
			F(18, 24) =	10.58
			Prob > F =	0.0000
Total (centered) SS	=	3963.588238	Centered R2 =	0.6182
Total (uncentered) SS	=	12756.50145	Uncentered R2 =	0.8814
Residual SS	=	1513.160894	Root MSE =	4.7

cmhgw	Coef.	Robust Std. Err.	z	P>|z|	[95% Conf. Interval]	
uratest	-.1740598	.5244106	-0.33	0.740	-1.201886	.8537661
uratest2	-.0163356	.0240597	-0.68	0.497	-.0634916	.0308205
cpi	-1.475031	.8546665	-1.73	0.084	-3.150147	.2000844
labprod	.0984259	.6521655	0.15	0.880	-1.179795	1.376647
coord	347.1079	87.30327	3.98	0.000	175.9966	518.2192
coord2	-69.731	18.08855	-3.85	0.000	-105.1839	-34.2781
centrcd	-24.82165	5.093531	-4.87	0.000	-34.80478	-14.83851
centrcd2	4.502184	.8378357	5.37	0.000	2.860056	6.144312
epl	219.7721	39.95214	5.50	0.000	141.4673	298.0768
epl2	-42.77294	7.800867	-5.48	0.000	-58.06236	-27.48352
pcov	-1983.853	483.7537	-4.10	0.000	-2931.993	-1035.713
pcov2	1057.48	278.2499	3.80	0.000	512.12	1602.84
cons	277.4475	65.78031	4.22	0.000	148.5205	406.3746

Hansen J statistic (overidentification test of all instruments): 19.392

Chi-sq(26) P-val = 0.81946

Summary results for first-stage regressions on the joint significance of excluded exogenous variables:

Shea

Variable	Partial R2	Partial R2	F(30, 24)	P-value
uratest	0.9657	0.9856	54.87	0.0000
uratest2	0.9639	0.9853	53.51	0.0000
cpi	0.7595	0.7693	2.67	0.0080
labprod	0.7250	0.7294	2.16	0.0286

Wald test of joint significance for the time dummies

chi2(6) = 6.88
Prob > chi2 = 0.3323

Table 11: OLS, centrln variable

Number of obs =	69
F(18, 50) =	8.95
Prob > F =	0.0000
R-squared =	0.6075
Root MSE =	5.5783

cmhgw	Coef.	Robust Std. Err.	t	P>|t|	[95% Conf.	Interval]
coord	343.914	297.4963	1.16	0.253	-253.6249	941.4529
coord2	-70.84581	60.56242	-1.17	0.248	-192.489	50.79738
centrln	-1.208709	5.077091	-0.24	0.813	-11.40635	8.988927
centrln2	.1256007	.2563429	0.49	0.626	-.3892792	.6404805
epl	182.0391	108.3226	1.68	0.099	-35.53331	399.6115
epl2	-34.44077	20.80747	-1.66	0.104	-76.2338	7.352262
pcov	-2145.187	1659.377	-1.29	0.202	-5478.144	1187.77
pcov2	1213.405	960.5629	1.26	0.212	-715.9426	3142.752
uratest	.1971928	1.025947	0.19	0.848	-1.863482	2.257867
uratest2	-.0194398	.0367424	-0.53	0.599	-.0932391	.0543596
cpi	-.8985152	1.023029	-0.88	0.384	-2.95333	1.1563
labprod	-.5716181	.5861671	-0.98	0.334	-1.748969	.6057332
cons	318.6113	209.2154	1.52	0.134	-101.6102	

738.8328

F-test of joint significance for the time dummies

F(6, 50) = 0.44
Prob > F = 0.8476

Cameron & Trivedi's decomposition of IM-test

Source	chi2	df	p
Heteroskedasticity	69.00	68	0.4434
Skewness	33.18	18	0.0159
Kurtosis	2.89	1	0.0890
Total	105.08	87	0.0909

Table 12: IV (2SLS), centrln variable

		Number of obs =	69
		F(18, 24) =	9.42
		Prob > F =	0.0000
Total (centered) SS =	3963.588238	Centered R2 =	0.6003
Total (uncentered) SS =	12756.50145	Uncentered R2 =	0.8758
Residual SS =	1584.239805	Root MSE =	4.8

cmhgw	Coef.	Robust Std. Err.	z	P>|z|	[95% Conf. Interval]	
uratest	.0214431	.9378611	0.02	0.982	-1.816731	1.859617
uratest2	-.0190579	.0314165	-0.61	0.544	-.0806331	.0425172
cpi	-1.286427	.9251111	-1.39	0.164	-3.099612	.526757
labprod	.0624387	.6577553	0.09	0.924	-1.226738	1.351615
coord	429.7361	280.603	1.53	0.126	-120.2357	979.7079
coord2	-88.2214	57.08003	-1.55	0.122	-200.0962	23.65339
centrln	-.1094732	4.743351	-0.02	0.982	-9.406271	9.187324
centrln2	.0655103	.2406446	0.27	0.785	-.4061445	.5371651
epl	224.931	105.7176	2.13	0.033	17.72823	432.1338

```
        epl2 |  -42.80975   20.35622    -2.10   0.035    -82.70721
-2.912299
        pcov |  -2654.179   1576.827    -1.68   0.092    -5744.704
436.3453
       pcov2 |   1499.963   910.7118     1.65   0.100    -284.9997
3284.925
        cons |   384.7986   199.3666     1.93   0.054    -5.952832
775.55
------------------------------------------------------------------------
-----
Hansen J statistic (overidentification test of all instruments):
21.583
                                                 Chi-sq(26) P-val =
0.71129
------------------------------------------------------------------------
-----
Summary results for first-stage regressions on the joint significance
of excluded exogenous variables:

                Shea
Variable     Partial R2      Partial R2       F( 30,    24)    P-value
uratest        0.9611          0.9700              25.89        0.0000
uratest2       0.9658          0.9761              32.71        0.0000
cpi            0.7578          0.7653               2.61        0.0093
labprod        0.7270          0.7403               2.28        0.0209
------------------------------------------------------------------------
-----
Wald test of joint significance for the time dummies

           chi2(  6) =    4.12
         Prob > chi2 =    0.6607
```

Appendix 4: Matching the EE and RF, the reference period for measures of wages

For our study of matched data we have restricted the sample to employees whose total annual wages of year n are declared in the RF, and whose wage in March n and March $n+1$ has been reported in the EE. We check that the above employees have been working in the same establishment between January n and March $n+1$, are full-time workers, and declare not to have a secondary activity.

In this framework it is possible to compare the value collected in the RF to the two measures of wages constructed from the two values reported in the EE. The most reasonable assumption we can make is that the RF give the correct measure of wages earned during the year. We can therefore interpret the difference between one of the two values reported in the EE in n or $n+1$ and the RF at n as an approximation of individuals' reporting error in the EE:

$$\begin{aligned} W_{n+1} &= W_n^* + \epsilon_{n+1} + \nu_{n+1} \\ W_n &= W_n^* + \epsilon_n + \nu_n \end{aligned}$$

where:

W_n is the measure of annual wage calculated from the value reported in the March n survey of the EE

W_n^* is the wage received during the year n as reported in the RF

ϵ_n is reporting error of the employee reporting his wage in March n

ν_n represents all the others sources of differences, among which the fact that in the EE wage is measured at one point of time

Since wages are on average increasing over time, ν_{n+1} is a priori positive and ν_n is negative. As a consequence, if we assume that measurement errors characteristics in the EE are constant over time, i.e. ϵ_n follows the same law of ϵ_{n+1} $\forall t$, then such measurement error will be between $(W_{n+1} - W_n^*)$ and $(W_n - W_n^*)$. For this reason we show systematically both distributions. The comparison between them allows to abstract from the second term of error, which is therefore omitted.

An alternative solution consists in considering the average of the two errors. But this method has the inconvenience of hiding many rounded values of wages whereas the type of rounding can be different at the two dates.

We therefore carry out a simple decomposition of variance of the measure of wage in the EE:

$$V(W_n) = V(W_n^*) + V(\epsilon_n) + 2cov(W_n^*, \epsilon_n)$$

In Table 16 we report the variance components of W_n due to W_n^* and ϵ_n together with their correlation coefficient.