
Properties of the Histogram Location Approach and the Extent and Change of Downward Nominal Wage Rigidity in the EU

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Abstract

The histogram location approach has been proposed by Kahn (1997) to estimate the fraction of wage cuts prevented by downward nominal wage rigidity. In this paper, we analyze the validity of the approach by means of a simulation study which yielded evidence of unbiasedness but also of potential underestimation of rigidity parameter uncertainty and therefore of potential anticonservative inference.

We apply the histogram location approach to estimate the extent of downward nominal wage rigidity across the EU for 1995-2001. Our data base is the User Data Base (UDB) of the European Community Household Panel (ECHP). The results show wide variation in the fraction of wage cuts prevented by nominal wage rigidity across the EU. The lowest rigidity parameters are found for the UK, Spain and Ireland, the largest for Portugal and Italy. Analyzing the change of rigidity between sub periods 1995-1997 and 1999-2001 even shows a widening of the differences in nominal wage rigidity. Due to the finding of large differences across the EU, the results imply that the costs of low inflation policies across the EU differ substantially.

JEL Classification: J30

Keywords: Wage rigidity, Inflation, Unemployment

1. Introduction

Of the existing approaches to testing for nominal wage rigidity, the histogram location approach seems promising. Contrarily to most approaches, the histogram location approach proposed by Kahn (1997) allows an explicit estimation of the share ρ of persons potentially subject to nominal rigidity, referring to the highest effective share in the case of substantial deflation. The share ρ multiplied by the share of persons who would face negative wage changes in the event of zero nominal wage rigidity yields the share of currently effected employees. Hence, the share of ultimately effected employees differs according to the location of the distribution of wage changes. The more the distribution of nominal wage changes is shifted to the left, the higher the share of ultimately effected employees, who according to the neoclassic paradigm in the following will face a strongly increased probability of becoming unemployed. Hence, if a significant and relevant share of employees is found to be subject to nominal wage rigidity, low inflation may cause unemployment, which, in this situation could be avoided with a policy leading to higher nominal wage changes at unchanged real wage changes due to higher inflation. Therefore, empirical evidence of relevant nominal wage rigidity would be a strong argument for allowing moderate inflation rates above zero inflation.³

¹ I would like to thank the editors and two anonymous referees for helpful comments

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³ Akerlof, Dickens and Perry (1996) provide a thorough discussion of the effects of inflation on the labour

The histogram location approach suggested by Kahn (1997) has been subsequently applied to different data sets by Beissinger and Knoppik (2001) and Knoppik and Beissinger (2003, 2005) who found strong evidence of downward nominal wage rigidity in Germany and the EU. Christofides and Leung (2003) applied the approach to Canadian contract data and they found strong nominal wage rigidity which might be due partly to the use of union contract data. Holden and Wulfsberg (2004, 2005) apply a non-parametric approach by means of bootstrap and simulation methods which allows the use of much less suitable data at the industry level. Despite the averaging effect of wage changes within industries, their analysis also yielded strong evidence of nominal wage rigidity in the EU. Their estimates are comparable to estimates obtained from micro data. Crawford and Harrison (1997) analyse macro data on wage union settlements 1952-1996 and find the share of constant or almost constant wages to decrease systematically with inflation rates. Using additional micro data sources they find wage cuts to be more common for employees of smaller firms and non-unionised employees.

Smith (2000) discusses such data issues as rounding and measurement error in some detail, based on the British Household Panel Study which could potentially lead to overestimation of the extent of rigidity measures based on survey data. Nickell and Quintini (2003) provide evidence on nominal wage rigidity in the UK, based on the New Earnings Survey and using a different approach based on two truncated normal distributions with different dispersions below and above 0 wage changes. The effect of downward nominal wage rigidity on unemployment is discussed controversially. In a somewhat different context, Iara and Traistaru (2004) analyze wage flexibility in EU accession countries using a Phillips curve approach and find only moderate unemployment elasticities of wages for most accession countries. Kuroda and Yamamoto (2003) using a simple threshold model derive an increase of up to 1.8% in the unemployment rate for Japan caused by downward wage rigidity. But their estimates strongly depend on elasticity estimates of labor demand used in the simulations.

In a survey, asking employees directly about their attitude towards wage cuts Agell and Lundborg (2003) find strong resistance to wage cuts even in high-employment and relatively low-inflation environments. Therefore, Agell and Lundborg provide strong evidence of underlying psychological reasons at the individual level, causing the observed phenomenon of downward wage rigidity.

This paper contributes to the analysis of nominal wage rigidity in two respects. Firstly, we analyze in detail, by means of a Monte Carlo simulation study, the properties of the histogram location approach. Secondly, by analyzing sub periods of the complete time span (1995-2001), we provide empirical results of the change in extent of downward nominal wage rigidity within ten European countries.

Our findings suggest that the rigidity parameter in question can be estimated quite accurately, but that the estimated standard error depends on the binwidth and potentially understates the true estimation uncertainty. Therefore the estimated standard errors can lead to anti-conservative inference. Additionally, we give evidence of the sensitivity towards the different estimators on hand, and find no notable improvement when allowing for heteroscedasticity and correlation of error terms. Because the estimation approach as suggested by Kahn (1997) has been formulated as a

market. Using macro data Altavilla and Marani (2005) provide evidence on costs of price stability and budget consolidation in the EU.

regression approach, we suspect that the strong violations of the usual error term assumptions cause the inference to be misleading.

With data from the User Data Base (UDB) of the European Community Household Panel (ECHP), we apply the approach empirically. Even taking into account the potential underestimation of standard errors, nominal wage rigidity is found to be significant according to conventional significance levels in all ten countries included in the analysis. However, taking into account the potential underestimation of the standard deviation of the rigidity parameter, the significance of the rigidity parameter for the UK, Ireland and Spain is unclear when analyzing sub periods.

The results show large differences in wage rigidity across the EU, with parameter estimates ranging from low, about 4–6% for the UK, Spain and Ireland, and as high as 33% for Portugal. Analyzing the change of nominal wage rigidity in two sub periods reveals that differences in nominal wage rigidity across the EU even widened between 1994 and 2001. Due to the finding of large differences across the EU, the results imply widely differing costs imposed by low-inflation policies across the EU. But there is only weak evidence that effects of high fractions of wage cuts prevented (FWCP) have been reduced by means of inflation.

The paper is structured as follows. In Section 2 we outline the basic method used throughout the paper to estimate the share of potentially effected employees by wage rigidity. Simulation evidence to assess the validity of the histogram location approach is also provided in Section 2. Section 3 contains a very brief description of our data source and data selection criteria. Section 4 contains the empirical results and Section 5 concludes.

2. The histogram location approach

1. The estimation procedure

Throughout the paper, we denote the nominal hourly wage rate by X and relative wage change by x :

$$\begin{aligned} x_t &= \log(X_t) - \log(X_{t-1}) \\ &\approx \frac{X_t}{X_{t-1}} - 1 \end{aligned}$$

The histogram location approach starts with the notion of a counterfactual distribution of wage changes $F(x)$ which would prevail in the absence of nominal wage rigidity. The empirical distribution of wage changes $G(x)$ differs from the hypothetical distribution $F(x)$ below the point of 0 wage changes due to nominal wage rigidity, as the share ρ of employees hypothetically facing wage cuts experiences a wage change of 0 instead, due to nominal wage rigidity. ρ is also referred to as the 'fraction of wage cuts prevented' (FWCP). Hence, we find $F(x) > G(x)$ for $x < 0$. Let $f(x)$ denote the counterfactual relative frequency at $x=0$ and $g(x)$ the empirical relative frequency.

This basic assumption that notional and empirical distributions are identical for positive wage changes is made in the majority of empirical analyses. Based on a

theoretical model Elsbey (2004) argues that the empirical distribution might be compressed even for positive wage changes. This compression results in the model from two different sources: Employers anticipate difficulties in cutting wages in the future. Therefore, they lower their offered wage increases today ('active compression'). Additionally, due to the overall wage increasing effect of DNWR, employers have to raise wages less under DNWR to reach 'optimal' wage levels ('latent compression'). Whether the empirical content of these proposed effects is relevant has to be examined in future research.

Using the indicator function

$$I(x) = \begin{cases} 1 & \text{if } x < 0 \\ 0 & \text{else} \end{cases}$$

we have the following condition

$$I(x) \rho F(x) = g(x) - f(x) = \eta$$

The missing area for x values below 0 in the counterfactual distribution is shifted towards the empirical relative frequency at $x = 0$.

Figure 1: Observed and counterfactual wage change distributions

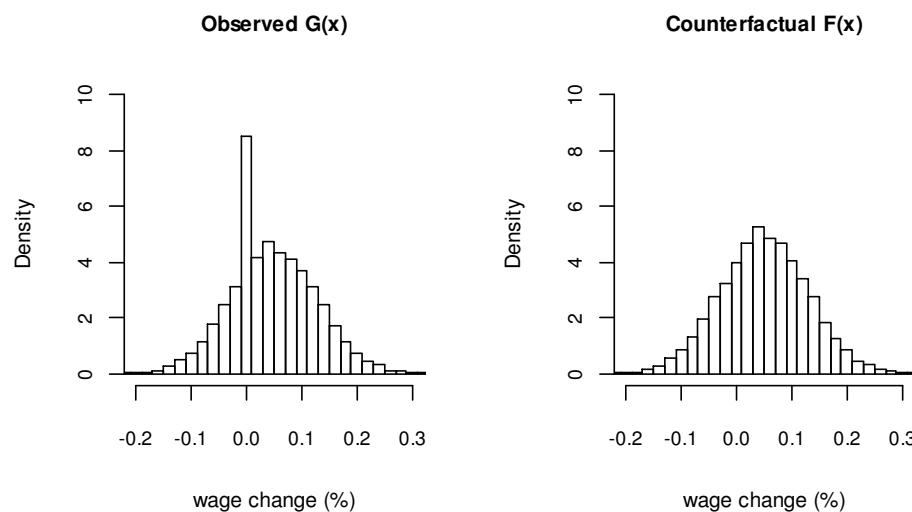


Figure 1 visualizes the effect of nominal wage rigidity, as the bin including the value $x = 0$ contains additionally the relative frequency η . The left histogram in Figure 1 shows the observed distribution $G(x)$ in the case of nominal wage rigidity, while the counterfactual distribution $F(x)$ is displayed in the right histogram. Using the index r ($r = 1, \dots, R$) for bins and r^0 to indicate the bin containing the value $x = 0$ and f_r and g_r the counterfactual and empirical densities of the bins r , we find that

$$\begin{aligned}
 f_r &< g_r \text{ for } r < r^0 \\
 f_r &= g_r \text{ for } r > r^0 \\
 f_{r^0} &+ \sum_{r=1}^{r^0-1} (g_r - f_r) \\
 &= f_{r^0} + \rho \sum_{r=1}^{r^0-1} f_r = g_{r^0}
 \end{aligned}$$

Note that the fraction ρ is assumed to be constant throughout the bins $r = 1$ to $r = r^0 - 1$.

To set up the estimation equation, we introduce two indicator variables dn_r and dz_r :

$$\begin{aligned}
 dn_r &= \begin{cases} 1 & \text{if } r < r^0 \\ 0 & \text{else} \end{cases} \\
 dz_r &= \begin{cases} 1 & \text{if } r = r^0 \\ 0 & \text{else} \end{cases}
 \end{aligned}$$

The estimation equation expresses the empirical bin densities g_r as a function of counterfactual bin sizes f_r and the rigidity parameter ρ :

$$g_r = f_r - \rho f_r dn_r + dz_r \rho \sum_{r=1}^{r^0-1} f_r dn_r + \varepsilon_{rt} \quad \text{for } r = 1, \dots, R$$

ε_{rt} denoting the error term. Considering only one cross section, this equation is not estimable as there are $r^0 + 1$ parameters to estimate, but only r^0 equations. Equations with $r > r^0$ do not provide any information because for these bins $f_r = g_r$.

Therefore, at least two cross sections indexed by t ($t = 1, \dots, T$) are needed to estimate the model. This necessitates the assumption of constancy of the rigidity parameter ρ and constant counterfactual bin frequencies $f_{rt} = f_{rs}$ throughout time.

Because empirical distributions of wage changes clearly shift over time and yield an altering dispersion, these distributions have to be standardized for location and dispersion in order to meet this assumption at least approximately. Due to the shifting of empirical wage distributions over time, r_t^0 may differ for different t . We denote the highest r_t^0 as r^{\max} .

Additionally, a considerable number of bins at the very left of the distribution will contain only negative x -values in all cross sections. These bins can be summed to reduce the number of parameters to be estimated. Let r^{\min} denote the lowest bin experiencing the inclusion of the x -values of 0 at least once, then the sum of relative frequencies of the bins to the left of r^{\min} can be expressed as

$$\sum_{r=1}^{r^{\min}-1} f_{rt} = 1 - \sum_{r=r^{\min}}^{r^{\max}} f_{rt} - \sum_{r=r^{\max}+1}^R g_{rt}$$

using the restriction

$$\sum_{r=1}^R f_r = \sum_{r=1}^R g_r = 1$$

and

$$\sum_{r=r^{\max}+1}^R f_r = \sum_{r=r^{\max}+1}^R g_r$$

The estimation equation can then be expressed as

$$g_{rt} = f_{rt} - \rho f_{rt} dn_{rt} + dz_{rt} \rho \left(\sum_{r=r^{\min}}^{r^{\max}-1} f_{rt} dn_{rt} + \sum_{r=1}^{r^{\min}-1} f_{rt} \right)$$

In Kahn's (1997) seminal paper a standardization of location using the median without any correction for changing dispersion was suggested. To standardize the distribution for location by median subtraction as well as for dispersion, we use the percentile difference $\tilde{x}_{high} - \tilde{x}_{low}$. The lower quintile (e.g. $\tilde{x}_{low} = \tilde{x}_{0.5}$) must be chosen so that in all cross sections $\tilde{x}_{low} > 0$, because otherwise, that quintile value will be effected differently by rigidity for different t due to different locations of the wage-change distribution. In our estimation we use $\tilde{x}_{low} = 0.5$ and $\tilde{x}_{high} = 0.9$.

One drawback of standardization is that the value 0 will be located at different points in bin r_t^0 for different t . Ideally, the value of 0 would be located centrally in bin r_t^0 in all standardized distributions which is not feasible in practice. This effect of non-centrality of 0 in r_t^0 will be less problematic as T , the number of available cross sections, increases due to averaging of left and rightward location in r_t^0 .

2. Simulation results

To analyze the estimation procedure of the histogram location approach, we perform a Monte Carlo simulation. We are especially interested in whether the approach allows an unbiased estimation of the rigidity parameter. We also wish to assess the adequacy of the estimated standard error and to analyze the sensitivity to the choice of bin width.

We investigate three different estimators, simple Least Squares (LSR), iteratively weighted Least Squares allowing for heteroscedasticity (IWLSI) under and iteratively weighted Least Squares to account for heteroscedasticity as well as negative correlation between bins (IWLSII). As explained in the last subsection in each estimation procedure, we restrict the sum of estimated bin densities to 1. Additionally, we restrict the number of bins used in the estimation to a maximum of 16, to prevent extremely small bin densities strongly influencing the numerical estimation procedure.

The simulation set up is as follows and is aimed to mimic empirical wage change distributions:

$$\begin{aligned}
 & x_{it}^{unr} N(\mu_t, \sigma_t) \\
 & \mu_t U(0.02, 0.10) \\
 & \sigma_t U(0.2, 0.3) \\
 & x^{obs} = \begin{cases} x^{unr} & \text{if } x^{unr} \geq 0 \\ I \cdot x^{unr} & \text{if } x^{unr} < 0 \end{cases} \\
 & I = \begin{cases} 1 & \text{with Pr} = 1 - \rho \\ 0 & \text{with Pr} = \rho \end{cases} \\
 & \rho = 0.2 \\
 & i = 1, \dots, n_t; t = 1, \dots, 7 \\
 & n_t = 1000
 \end{aligned}$$

For each simulation, we draw seven cross-sections ($T = 7$) which will be used for estimation. We analyze four different binwidth. We use i ($i = 1, \dots, n_t$) for indexing the individual wage changes.

The estimation results are given in Table 1.⁴ For each binwidth, we give in the first line the average of estimated ρ . Below that, we provide the average estimated standard deviation $\hat{\sigma}_\rho$ and in the third line, the "true" standard deviation σ_ρ calculated as the standard deviation of estimated ρ coefficients.⁵

Table 1: Simulation results for the histogram location approach

binwidth		LS	IWLSI	IWLSII
0.005	ρ	0.1990	0.1992	0.1988
	$\hat{\sigma}_\rho$	0.0053	0.0090	0.0090
	σ_ρ	0.0115	0.0104	0.0104
0.01	ρ	0.1990	0.1996	0.1991
	$\hat{\sigma}_\rho$	0.0083	0.0070	0.0070
	σ_ρ	0.0102	0.0091	0.0091
0.015	ρ	0.1990	0.1998	0.1993
	$\hat{\sigma}_\rho$	0.0112	0.0063	0.0062
	σ_ρ	0.0100	0.0085	0.0086
0.02	ρ	0.1991	0.1998	0.1994
	$\hat{\sigma}_\rho$	0.0143	0.0061	0.0060
	σ_ρ	0.0093	0.0087	0.0087

⁴We present only simulation results for $n = 1000$ because the results were almost identical when using $n = 500$ or $n = 2000$.

⁵Number of simulation runs has been 250.

The results show that no estimator contains a considerable bias. The corrections for heteroscedasticity and heteroscedasticity and cross correlation do not lead to a notably improved efficiency of the Least Squares estimator. But one critical drawback is obvious: the estimated standard deviations are no reliable estimates for the true standard errors because of their dependence on the binwidth. For small binwidth we find that estimates understate the true standard deviations considerably, leading to anticonservative inference in the case of the simple restricted Least Squares estimator. Thus the null hypothesis of no downward nominal wage rigidity will be rejected too often. The underestimation is about 54% for the Least Squares estimator for the smallest binwidth. When increasing binwidth the underestimation reduces and turns towards overestimation for the largest binwidth in the case of the simple restricted Least Squares estimator. The corrected estimators underestimate the true standard deviation at all bin widths. The simulation results imply that familiar critical t -values can not be used for reliable inference. To prevent anticonservative inference, we use the simple restricted Least Squares estimator and binwidth 0.02 in the empirical estimations.

3. The data source

The ECHP is a longitudinal survey of households and individuals in countries of the European Union (EU). Due to its far-reaching harmonization, the ECHP facilitates cross-country comparisons within the EU in many different aspects of economic and social life.

Comparability across countries is the main objective of the ECHP. While great effort has been devoted to harmonizing the surveys, they are still not completely standardized. The differences occur mainly in sampling procedures at the start of the ECHP, the panel's follow-up rules and field operations.

Peracchi (2002) provided a comprehensive description of the ECHP and detailed information about the organization of the survey. The first wave of the ECHP in 1994 covered about 130,000 individuals older than 16 years resident in about 60,000 households. Twelve countries participated in the first wave: Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal Spain and UK. While Austria took part from the second wave onwards in 1995, Finland started its participation in 1996.

In most of the participating countries, the survey had been newly commenced, whereas a few countries used already existing panel surveys. In Belgium and the Netherlands, already ongoing panels were used for data collection, while in three countries, Germany, Luxembourg and the UK, a unique situation emerged, because ongoing panels ran parallel with the new ECHP national sub samples for three years. In 1997, these three national sub samples were terminated and from that year onwards, the data for the ECHP are derived from existing national panels. These are the German Social Economic Panel (GSOEP), the Luxembourg's Social Economic Panel (PSELL) and the British Household Panel Survey (BHPS). The ECHP User Data Base covers only the ECHP survey in Luxembourg. Since our objective is to analyze wage growth between 1994 and 2001, we consider only national surveys in Germany (GSOEP) and the UK (BHPS) in our analysis by country.

This analysis is based on the 2004 version of the ECHP User Data Base which contains all available 8 waves 1994 to 2001. In our analysis, we include ten countries which took part in all 8 waves: Germany, Denmark, Belgium, France, UK, Ireland, Italy,

Greece, Spain and Portugal.⁶ Peracchi (2002), Watson (2003) and Behr, Bellgardt and Rendtel (2005) provide analyses of panel participation and attrition in the ECHP. The effect of panel attrition on different measures of income inequality and income mobility has been analyzed by Behr, Bellgardt and Rendtel (2004), finding only very moderate effects of panel attrition.

Contrary to previous studies (e.g. Kahn 1997), we do not restrict the sample to "on the job stayers" firstly, because nominal wage rigidity could also prevent job movers from accepting wages below previous wage levels and secondly, we want to assess overall nominal wage rigidity irrespective of possible influences such as low or high shares of job movers.

In our sample finally used for estimation we include all employees aged between 18 and 65, working at least twenty hours a week. Allowing for changes in work hours from year to year, we use hourly nominal wages to calculate yearly wage changes. Because we regard observations with relative wage changes below -75% and above 300% being subject to grave suspicion, observations outside this range will not be regarded in the analysis. Because the approach is purely nonparametric, and even the statistics for standardizing the distribution are almost unaffected by this procedure, this data cleaning will not influence the results of the analysis.

4. Empirical evidence of nominal wage rigidity in the EU

1. Evidence of nominal wage rigidity 1995-2001

In this section, we analyze in detail the distribution of wage changes in 2001 for ten countries. Table 2 contains some descriptive statistics characterizing the wage change distribution 2001. Figure 2 displays histograms of the wage change distributions 2001 for all countries. A vertical line indicates the median and the bin containing 0 is black.

Table 2: Descriptive statistics of wage change distributions 2001

	n	median	IQR	$x \leq 0$ (%)	$x = 0$ (%)
Germany	4260	0.032	0.172	0.403	0.066
Denmark	1830	0.042	0.111	0.31	0.086
Belgium	1634	0.045	0.163	0.367	0.061
France	3108	0.044	0.183	0.356	0.04
UK	3320	0.069	0.227	0.322	0.025
Ireland	1340	0.108	0.273	0.251	0.011
Italy	3573	0.029	0.167	0.448	0.161
Greece	1919	0.037	0.143	0.39	0.164
Spain	3209	0.056	0.309	0.381	0.009
Portugal	3757	0.053	0.131	0.234	0.091

⁶Due to severe data problems we did not include the Netherlands in our analysis.

The histograms indicate strong wage rigidity in Denmark, Greece and Portugal, while there is no clear sign of rigidity for Ireland and Spain. For Italy and Greece we find an extremely high density for the bin containing 0 wage changes of about 16% in 2001.

To standardize the T wage change distributions individually for each country, we calculate for each year t standardise wage changes according to

$$x_t^s = \frac{x_t - \tilde{x}_{0.5,t}}{\tilde{x}_{0.9,t} - \tilde{x}_{0.5,t}}$$

Therefore, it holds for all standardised wage distributions that $\tilde{x}_{0.9,t} - \tilde{x}_{0.5,t} = 1$. Table 3 contains the estimated wage rigidity parameter using the restricted Least Squares estimator discussed in Section 2 and binwidth 0.02.

2. Analyzing the change in nominal wage rigidity based on two sub periods

After having estimated the average extent of wage rigidity for the complete available time span in the previous subsection, we now estimate the same model for the two sub periods 1995-1997 and 1999-2001. Comparing the two estimated rigidity parameters allows an evaluation of whether there is increasing or decreasing nominal wage rigidity. The estimation results are given in Table 4.

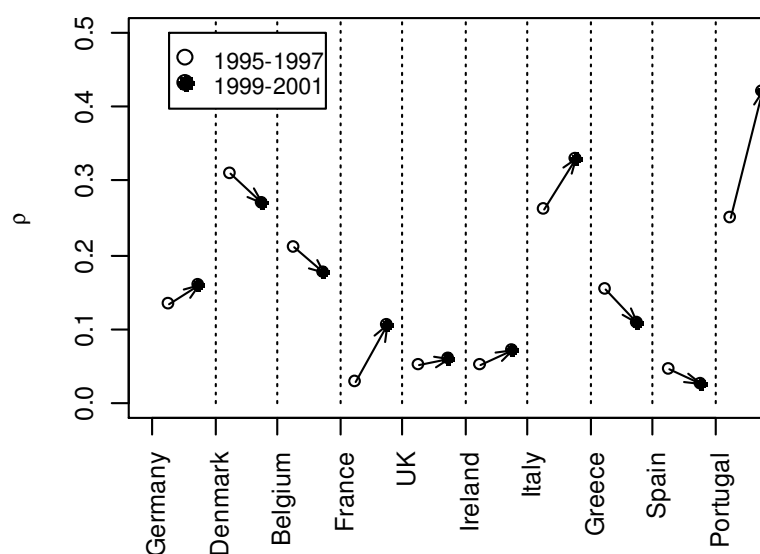
Table 4: Estimated rigidity, 1995-1997 and 1999-2001

	period	binwidth	ρ	sd(ρ)	t(ρ)
Germany	1995-1997	0.02	0.133	0.024	5.52
	1999-2001	0.02	0.159	0.028	5.68
Denmark	1995-1997	0.02	0.309	0.029	10.74
	1999-2001	0.02	0.27	0.036	7.43
Belgium	1995-1997	0.02	0.209	0.026	8.07
	1999-2001	0.02	0.175	0.024	7.14
France	1995-1997	0.02	0.028	0.019	1.44
	1999-2001	0.02	0.105	0.033	3.18
UK	1995-1997	0.02	0.052	0.035	1.49
	1999-2001	0.02	0.058	0.029	2.02
Ireland	1995-1997	0.02	0.051	0.029	1.77
	1999-2001	0.02	0.071	0.042	1.71
Italy	1995-1997	0.02	0.261	0.022	11.66
	1999-2001	0.02	0.331	0.025	13.37
Greece	1995-1997	0.02	0.152	0.032	4.74
	1999-2001	0.02	0.109	0.025	4.32
Spain	1995-1997	0.02	0.046	0.025	1.81
	1999-2001	0.02	0.025	0.031	0.82
Portugal	1995-1997	0.02	0.25	0.039	6.45
	1999-2001	0.02	0.422	0.037	11.5

To ease the interpretation of results, we display the estimation results graphically for the two sub periods, which allows an easy assessment of both the extent and direction of change of nominal wage rigidity across countries.

Figure 3 makes it evident that rigidity is highest in Portugal, Italy and Denmark and lowest in the UK, Ireland and Spain. A strong increase is found for France, Italy and Portugal when comparing the two sub periods, while in Greece rigidity declined considerably. Both the level of downward nominal wage rigidity, and the extent of changes reveal a very heterogeneous picture for the ten EU member states under analysis.

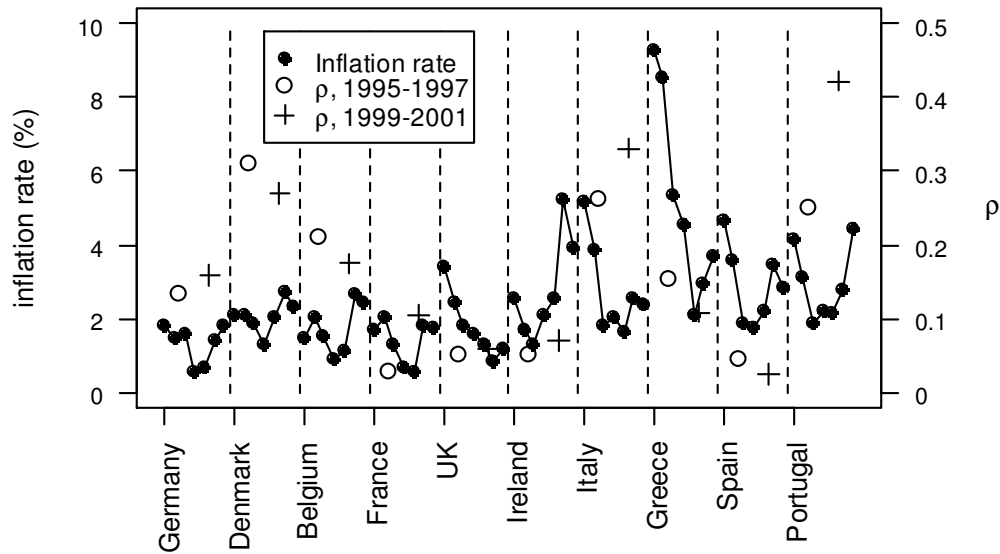
Figure 3: Downward nominal wage rigidity in two sub periods



3. Downward nominal wage rigidity and inflation

When interpreting the empirical results, it has to be kept in mind that the actual share of employees subject to nominal wage rigidity is the product of the unobservable share of employees facing wage cuts in the absence of rigidity, multiplied by the fraction of wage cuts prevented (ρ). Therefore, countries with high FWCP will experience stronger effective downward nominal wage rigidity, the further the wage change distribution is located to the left. Except in the presence of complete nominal money illusion, the location of the wage change distribution depends on inflation rates. Because higher inflation rates shift the wage change distribution to the right, inflation is a means to render high FWCP ineffective. Figure 4 displays inflation rates across countries for the time interval 1995-2001 and the estimated FWCP for the two sub periods 1995-1997 and 1999-2001.

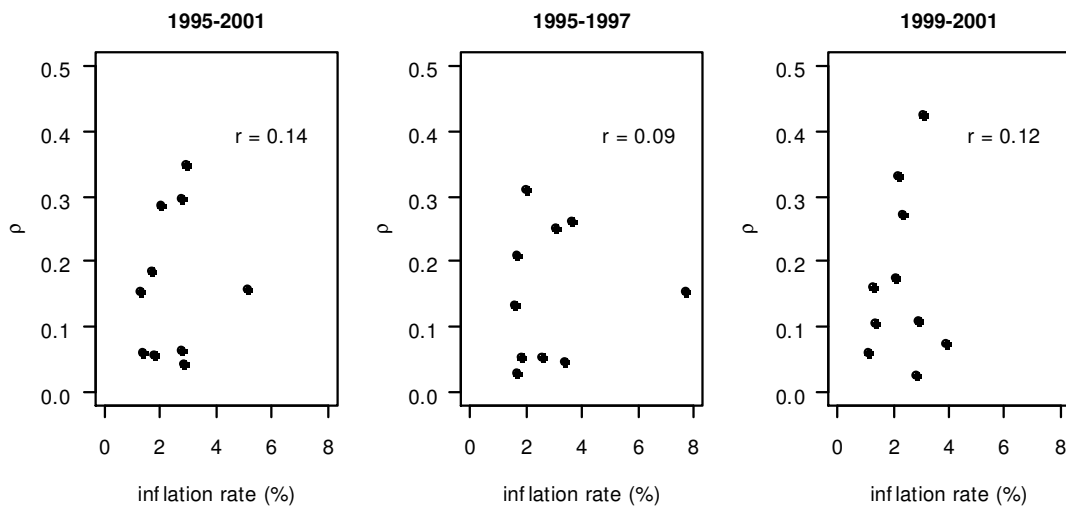
Figure 4: Inflation and downward wage rigidity



We find a very inhomogeneous development of inflation rates across countries. Most noteworthy is the strong decline in inflation rates in Greece.

In Figure 5 we display the relation between estimated FWCP and average inflation rates for the complete time period and for two sub periods. We find a small positive correlation of 0.09 for the sub period 1995-1997. The right diagram in Figure 5 displays the analogue relation for the sub period 1999-2001 with a correlation which has slightly increased to 0.12. While the dispersion of inflation rates in the sub period 1999-2001 is smaller due to the strong reduction of inflation in Greece, the estimated ρ parameters display higher variation.

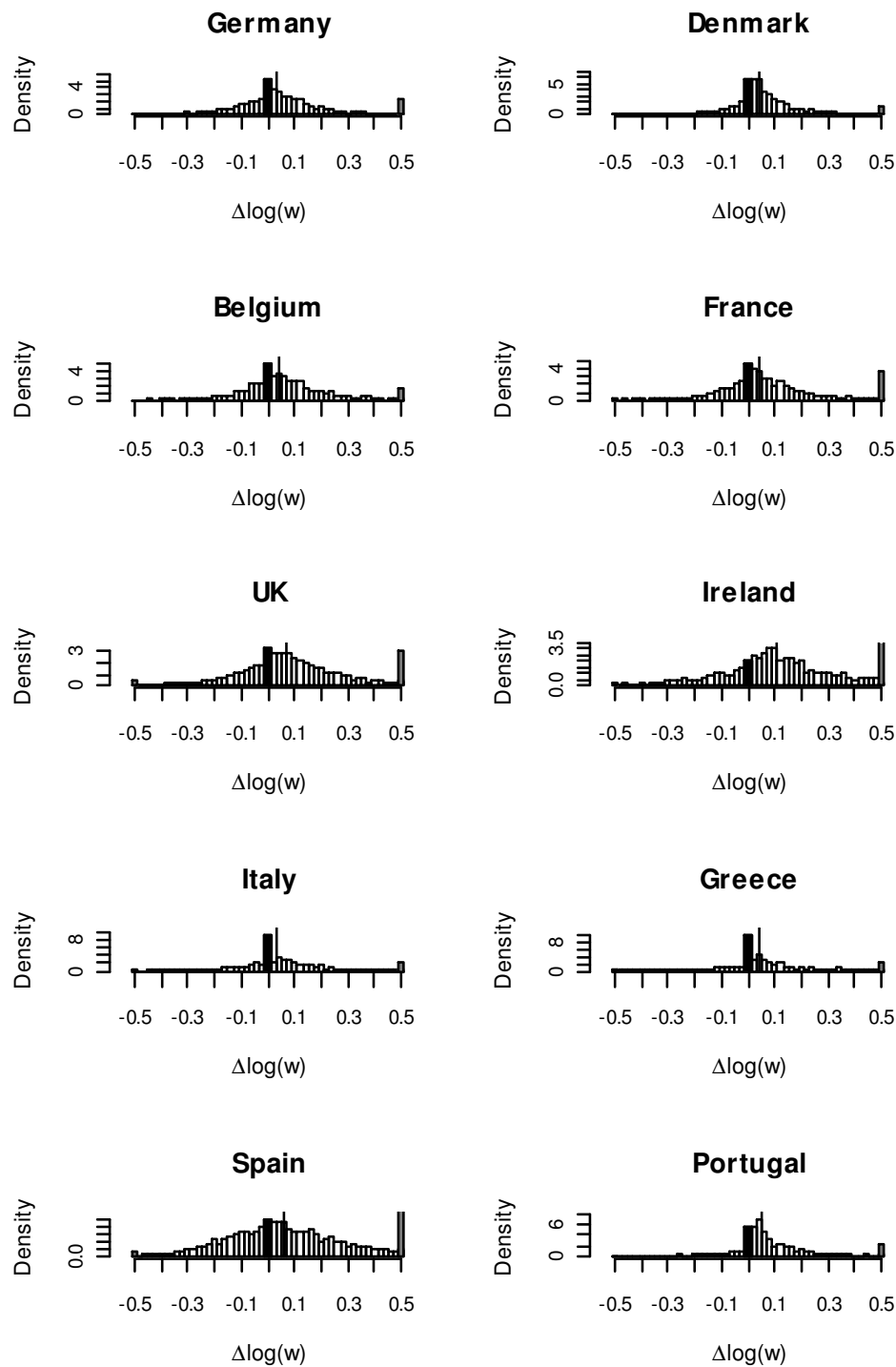
Figure 5: Correlation of inflation and rigidity



Due to the European Monetary Union, only moderate differences in national inflation rates can be expected in the current rather low inflation environment throughout the EU. This precludes countries found to have high rigidity parameters from reducing the distorting effects of downward rigid nominal wages through

moderate inflation policy. But according to the low correlation between rigidity parameters ρ and inflation rates there is only weak evidence for such policy in the period 1995 to 2001.

Figure 2: Histograms of wage change distributions 1995-2001



We find wage rigidity to be especially strong in Portugal, Italy and Denmark. Rigidity is found to be lowest in Spain and the UK. These estimation results strongly confirm the graphical evidence given in Figure 2. Despite the wide variety of rigidity measures obtained, all estimates are statistically significant according to conventional t -values. When taking into account the potential underestimation of "true" standard errors using $\hat{\sigma}_\rho$, the significance of ρ for Spain, Ireland and the UK has to be regarded as insecure.

Table 3: Estimated rigidity, 1995-2001

	binwidth	ρ	sd(ρ)	t(ρ)
Germany	0.02	0.151	0.015	9.85
Denmark	0.02	0.286	0.019	15.27
Belgium	0.02	0.185	0.014	13.16
France	0.02	0.059	0.013	4.57
UK	0.02	0.054	0.019	2.83
Ireland	0.02	0.061	0.019	3.26
Italy	0.02	0.294	0.013	22.00
Greece	0.02	0.155	0.015	9.99
Spain	0.02	0.041	0.018	2.25
Portugal	0.02	0.347	0.023	15.35

5. Conclusion

We first analysed the histogram location approach by means of a Monte Carlo simulation. The results showed this approach to allow for an unbiased estimation of the rigidity parameter of interest and to be robust with respect to differing choices of the sample size. But estimated standard errors of the wage rigidity parameter are found to vary with binwidth and can potentially underestimate true uncertainty for small binwidth. Therefore, inference is potentially anticonservative and conventional t -statistics are unreliable.

Our empirical analysis for ten European countries revealed strong differences in the extent of wage rigidity, ranging from about 4-6% for the UK, Ireland and Spain up to about 33% for Portugal. When splitting the complete time span into two sub periods, 1995-1997 and 1999-2001, wage rigidity was found to have decreased considerably in Greece, while increased rigidity was found for France, Italy and Portugal. These findings do not permit a clear conclusion to be drawn regarding a general trend in the extent of downward nominal wage rigidity. While the strong differences in the fraction of wage cuts prevented imply differing distorting effects of downward nominal wage rigidity in low inflation environment, there is only weak evidence that inflationary policies have been applied to lessen this effect in the time period between 1995 and 2001.

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