Downward Nominal Rigidity in West-German Earnings 1975-1995*

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Abstract

If downward nominal wage rigidity exists it should affect the distribution of earnings changes. We present a common analytical framework for three distinct and previously unconnected approaches to the analysis of downward nominal rigidity, the skewness-location approach, the symmetry approach and the histogram-location approach. We modify them by dropping the assumption of time-invariant rigidity and apply them to earnings data from the IABS. We find that the distribution of West German log earnings changes is indeed affected by downward nominal rigidity. Our modification of the approaches also allows us to find that the degree of nominal rigidity depends on business cycle conditions, with weaker rigidity in times of rising unemployment. Our findings support the critics of very low inflation targets.

Keywords: Nominal Wage Rigidity; Nominal Wage Stickiness; Inflation; Germany; Micro Data.


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1 Introduction

The European Central Bank has started into monetary union with a fairly ambitious inflation target of „below two percent“. However, there is some concern that very low rates of inflation in Europe might endanger macroeconomic stability. One important argument in that respect is the ‘greasing the wheels of the labor market’ hypothesis of Tobin (1972). The functioning of the labor market requires a multitude of real wage adjustments in both directions, up and down. If there is downward nominal wage rigidity, some of the real wage reductions that require nominal wage cuts do not take place. In that case, the proportion of real wage reductions that are not enacted depends on the trend growth of nominal wages and therefore on the rate of inflation. Recent estimates by Akerlof, Dickens and Perry (1996) put the critical rate of inflation, that allows almost all real wage reductions to be enacted, at three percent.

We are interested in whether and to which extent such downward nominal rigidity exists in Germany. Since Germany is the largest economy in the Euro area, the results are of some importance for the inflation policy of the European Central Bank. In the literature the following main types of empirical analysis have been performed to analyze nominal wage rigidity: studies based on aggregate time series, survey studies of firm wage and personnel policies and studies based on individual compensation data. It is not straightforward to use aggregate time series data for detecting downward nominal rigidity, since the heterogeneity of firm-specific shocks and wage adjustments are not easily taken into account. Of several attempts to analyze wage rigidity in aggregate data, Akerlof, Dickens and Perry (1996) is the only one compatible with the Tobin Hypothesis.1 Some recent firm survey studies which explore firm behavior with respect to wage cuts by asking personnel managers are Bewley (1998), Campbell and Kamlani (1997), Agell and Lundborg (1995) and Blinder and Choi (1990) with different focus on size and type of firm, and geographic region. These studies suffer from the well-known problem that the answers of individuals in such surveys may not conform to the actual economic behavior. The most promising research strategy therefore seems to be the analysis of micro data. Using micro data on actual market outcomes has two compelling advantages. It offers information on the complete distribution of wage changes which is central to the Tobin argument and it reflects actual economic behavior and not mere intent.

For these reasons our study is based on a large micro data set, the 1975 to 1995 version of the IAB-Beschäftigtengstichprobe (IABS). The IABS is a 1% random sample drawn from the German social security accounts which represent one of the most important data sources for labor market research in Germany. The great advantage of the IABS is the high reliability of the earnings data due to plausibility checks that are performed by the social security institutions and legal sanctions for misreporting. Measurement error due to erroneous reporting is therefore not a problem in our analysis, in contrast to studies based on compensation data from household surveys. Furthermore, earnings in the IABS are a broad measure of compensation including most types of fringe benefits and therefore yield a more accurate picture of the rigidity or flexibility of total compensation. Finally, studies based on only high inflation years have been subjected to a version of the Lucas critique, which claims that behavior with respect to

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1 A critical discussion of the time series evidence in Akerlof, Dickens and Perry (1996) can be found in Knoppik (1999).
nominal rigidity would change in times of low inflation. The presence of low inflation years in our data is therefore yet another advantage, since it renders this criticism invalid.

We are not aware of any study of downward nominal wage rigidity based on micro data for Germany. Several approaches to the analysis of micro data have been suggested in the literature and applied to US data: The skewness-location approach by McLaughlin (1994), the symmetry approach by Card and Hyslop (1997) and the histogram-location approach by Kahn (1997). Each approach judges the existence and, except for the first, also the extent of nominal rigidity by analyzing the distributions of relative changes in some compensation variable. Up until now the connections between these three approaches have not been explored. This has led to a situation where the evidence for nominal rigidity from the analysis of micro data is rather mixed. This is quite surprising since most of the evidence is based on very similar data from the PSID. We therefore present these previously unconnected approaches in a common analytical framework, in order to make any implicit assumptions explicit that may be responsible for the inconclusive results. It turns out that one implicit assumption of the approaches is the time-invariance of nominal rigidity. This assumption is strongly contradicted by evidence from interviews with personnel managers, since the firm survey studies uniformly find that nominal wage cuts will be implemented, but only if they can very well be justified by severely adverse market conditions. Since market conditions vary over time, the apparent degree of nominal rigidity will vary correspondingly. Also, if total compensation and not pay-rate variables are used, systematic variation in paid hours over the cycle may have an influence on the degree of rigidity apparent in earnings data.

We make a major modification of the approaches by dropping the assumption of time-invariant rigidity. When we apply the modified approaches to the IABS data, we obtain uniform results for the skewness-location approach, the histogram-location approach, and from descriptive and visual evidence. We find that there is indeed nominal rigidity, which is weaker in times of rising unemployment and more pronounced in times of falling unemployment. Since our compensation variable is annual earnings we can directly only draw conclusions about rigidity in earnings. But we can be confident that it is caused by even more pronounced downward rigidity of wages, since variation in hours will tend to hide some of the effects of wage rigidity in the distribution of earnings changes.

The remainder of the paper is organized as follows. In the next section we present our analytical framework, introduce the three approaches to the analysis of nominal rigidity and the assumptions upon which they are based. In sections three and four we describe our sample from the IABS and present descriptive and visual evidence for downward nominal rigidity, respectively. The following three sections are devoted to a more detailed discussion of the three approaches and the results from their application to our data. The final section summarizes our findings and presents our conclusions.

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2 The earnings-function approach of Altonji and Devereux (1999) also analyzes micro data. The approach is however quite different from the other three, so that we will discuss and apply it in a later paper. An application of the earnings-function approach to Swiss data is contained in Fehr and Goette (1999). Groshen and Schweitzer (1996) and Groshen and Schweitzer (1999) analyze the presence of nominal rigidity in data from the Cleveland salary survey, CSS, a regional survey of the salary structure of firms; their method of analysis is not applicable to employee panel data.

3 We show that the assumption of symmetry is violated and do not report results from the symmetry approach.
2 Nominal Rigidity and the Distribution of Individual Log Earnings Changes

In order to assess the possible impact of nominal rigidity on the distribution of log earnings changes, one has to distinguish the observable distribution from the hypothetical one that would have prevailed in the absence of rigidity. This hypothetical distribution of notional changes of log earnings, \( x^* \), is usually termed the counterfactual or notional distribution; we denote it by \( G(x^*) \). The distribution of actual log changes \( x \) is called actual or factual distribution; we denote it by \( F(x) \). It is the result of the effects of rigidity on the distribution of log earnings changes, which are modeled by a rigidity function \( \rho(x) \). We follow the majority of the literature and do not try to explain the underlying determinants of the counterfactual distributions in this paper, but treat the counterfactual distributions as given instead.

In the first part of this section we present the analytical framework that we have devised to illustrate graphically and mathematically the relationship between counterfactual and factual distribution in the presence of nominal rigidity. The most important types of nominal rigidity are downward rigidity that prevents all or part of the negative changes from being enacted, and menu cost effects that prevent small changes of either sign from being enacted.

In the second part of this section we use our analytical framework to introduce three of the approaches to the analysis of nominal rigidity that have been proposed in the literature, the skewness-location approach, the symmetry approach and the histogram-location approach. It seems warranted to relate these approaches to each other, since in their application to PSID data they have led to quite mixed results. We pinpoint the differences and commonalties of the approaches by explicitly formulating the assumptions that each approach has to make about the three building blocks of the framework. So far, no rigorous description of these assumptions and consequently no comparison of the approaches has been performed in the literature. Finally we modify the approaches by relaxing an assumption that they do adopt, i.e. the assumption of a time-invariant rigidity.

**Counterfactual and factual distributions and the effects of nominal rigidity**

If nominal rigidity is present, part or all changes in real earnings will not be enacted if they require nominal cuts or small nominal changes of either sign. A freeze of nominal earnings is thought to occur instead. This direct effect of nominal rigidity shifts some probability mass directly to zero, either from the left tail of the distribution, or from the immediate neighborhood of zero. This creates a probability mass at zero in the factual distribution, which therefore is not differentiable at zero. Beyond the direct effects there could as well exist indirect effects of rigidity on the distribution of log earnings changes, for instance due to exits into unemployment. We will abstract from these indirect effects in the following exposition and briefly discuss them with a number of other assumptions later in this section. A simple example of a downward nominal rigidity is proportional downward rigidity, where a constant proportion \( \rho(x) = \rho \ (0 < \rho \leq 1) \) of the negative changes is thought to not be enacted. The link between counterfactual and factual distribution in this case is:

\[
F_t(x_t) = \begin{cases} 
(1-\rho)G_t(x_t) & x_t < 0 \\
G_t(x_t) & x_t \geq 0.
\end{cases}
\]
This type of rigidity is illustrated for an arbitrary fictitious counterfactual distribution in FIGURE 1. Panels a) and b) show probability density functions of counterfactual distributions in periods with low and high median, \( t = L, H \). Panel c) illustrates a case of proportional downward rigidity, with the characteristic thinning for negative changes. The area of the arrow above zero represents the probability of freezes that is concentrated at zero.

FIGURE 1: Fictitious Counterfactual Distributions and Corresponding Factual Distributions for Various Types of Nominal Rigidity

Notes: For \( x \) different from zero, \( g(.) \) and \( f(.) \) are the derivatives of \( G(.) \) and \( F(.) \), where the latter is not differentiable at zero. The factual distributions’ probability mass at zero is represented by the arrow-shaped area above zero. \( t = L \) and \( t = H \) denote periods where the counterfactual \( G_t(.) \) has low and high median.

The rigidity could be more complicated than in the proportional downward rigidity example. There might also exist menu costs that prevent small changes of either sign from being enacted. Panel e) of FIGURE 1 illustrates a case where proportional downward rigidity and complete

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4 By a change in a ‘period’ or in a ‘year’, we refer to the change that has occurred between that period or year and the previous one.
symmetrical menu costs coexist. More variants of these types of rigidity have been proposed, but the examples should suffice to get a clear idea of the link between counterfactual and factual distributions. The direct effects of a rather general nominal rigidity can be formulated analogously to equation (1) by introducing a time-variant rigidity function $\rho_t(x)$. This function takes values between zero and one for changes $x$ below the upper bound for menu cost effects $x^m_t$, which is the highest value of $x$ that is affected by the rigidity. Otherwise the rigidity function is equal to zero. Using $g_t(.) = dG_t(.)/dx$, this leads to

$$F_t(x_t) = \begin{cases} \int_{-\infty}^{x_t} (1-\rho_t(z))g_t(z)dz, & x_t < 0 \\ G_t(0) + \int_{0}^{x_t} \rho_t(z)g_t(z)dz + \int_{x_t}^{x_t^m} (1-\rho_t(z))g_t(z)dz, & 0 \leq x_t \leq x_t^m \\ G_t(0), & x_t^m < x_t. \end{cases}$$

The term denoted by $F_t(0)$ in the middle branch of the equation contains the factual’s probability mass from below zero plus the piled-up probability mass from the direct effect of nominal rigidity from below and above zero.

The counterfactual distribution may differ in each year, e.g. due to the effects of inflation and productivity growth on wage formation. Panels a) and b) of FIGURE 1 show two counterfactual distributions that solely differ in location. Clearly such a shift to the right has important implications for the corresponding factual distribution which is illustrated in panel d). Only a smaller part of the distribution’s left tail is affected by thinning and correspondingly less probability mass is shifted to zero. Note that factual and counterfactual distribution have the same median as long as the median of the counterfactual distribution lies above the highest value affected by rigidity, $x^m_t$.

**Approaches to the analysis of nominal rigidity**

There are three approaches to the analysis of nominal rigidity that we use to analyze our data, the skewness-location approach, the histogram-location approach and the symmetry approach. With our conceptual framework in place, linking counterfactual and factual distribution functions by a rigidity function, we are able to sketch the approaches; the formulation of a number of assumptions will enable us to give a concise and comprehensive exposition of the approaches and to clearly indicate where we modify them. All approaches have the following assumption in common.

**A1 Direct Effects of Rigidity**

There are only direct effects of nominal rigidity, i.e. probability mass is taken away from the counterfactual distribution for small or negative changes and added to the probability of a zero change, as formalized in equation (2).

We have defined direct effects of nominal rigidity as those that shift some probability mass of the counterfactual to the probability of a zero change. Correspondingly indirect effects are all other effects of nominal rigidity on the distribution. One important class of indirect effects are those related to additional exits into unemployment that are caused by the rigidity and its effects on real wages. It is however quite unclear which employees will be laid off, exactly those whose wages cannot be cut (which would reduce the pile-up at zero) or a random selection of

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5 Other types of rigidity are proportional downward rigidity plus partial menu cost, as in Card and Hyslop (1997) and the ‘extended downward menu cost’ or ‘downward friction’ of Altonji and Devereux (1999).
other employees. Without explicit modeling of the flows in the labor market and the process of wage formation not much can be said about these and other types of indirect effects. Therefore we had to follow the practical solution of the literature, i.e. to adhere to the assumption that there are only direct effects of rigidity at work.

All approaches also employ assumption two:

**A2 Sufficiently High Median**

The highest value of $x_t$ that is affected by the rigidity, $x_t^m$, is smaller than the median of the counterfactual distribution $m^*_t$:

$$x_t^m < m^*_t \text{ for all } t.$$

Together with the assumption that there are only direct effects of rigidity the assumption that medians are sufficiently high has the important implication that the medians of counterfactual and factual distributions are equal and that only the left tails of the distributions are affected by the rigidity. The assumption is not necessary, but very convenient for the skewness-location approach. Medians seem to be high enough in the sense of assumption A2 in our data.

The skewness-location approach and the histogram-location approach exploit variation in the location of the distributions over the course of years and infer existence and extent of the rigidity from the corresponding variation in measured skewness and variation in the shape of observed histograms, respectively. This makes only sense if the shape of the distributions does not vary for other reasons. Both therefore make assumption three:

**A3 Invariance of Median-Centered Counterfactual**

Counterfactual distributions in different periods $G_t(x^*_t)$ differ only in their location, i.e. the median-centered counterfactual distribution is time invariant:

$$G_t(x^*_t) = G_t(x^*_t - m^*_t) \text{ for all } t,$$

where $m^*_t$ is the median change in $t$.

The assumption of a time-invariant counterfactual may not be literally fulfilled. Since we took great care to create a homogeneous sample, the effects of composition bias on the counterfactual distributions should at least be limited. We are therefore confident that the assumption is a reasonable approximation.

The symmetry approach assumes the symmetry of the counterfactual distribution:

**A4 Symmetric Counterfactual**

The counterfactual distribution $G_t(.)$ is symmetric about its median $m^*_t$.

$$G_t(m^*_t - z) = 1 - G_t(m^*_t + z) \text{ for all } z.$$  

Whether the symmetry assumption is fulfilled in our data will be scrutinized rigorously in section 6.

When applying the approaches enumerated above, we will make the assumptions we have formulated so far. This is not the case with the next assumption, which we only formulate to describe the approaches as used in the literature. We do not make that assumption in our own empirical implementation of the modified approaches. At first glance, the logic of the skewness-location and histogram-location approaches seems to require a time-invariant rigidity,
much as it requires a time-invariant centered counterfactual distribution. In the literature on the two approaches therefore implicitly assumption five is made:

**A5 Invariance of Rigidity**

The type and extent of rigidity are time invariant, i.e. the rigidity function is the same in different periods:

$$\rho_t(z) = \rho(z) \text{ for all } t \text{ and } z.$$  

There are however several arguments why the actual workings of rigidity may change over time. The first argument concerns the circumstances under which wage cuts take place. Studies of firms wage policies uniformly arrive at the conclusion that firms only resort to wage cuts if they can very well justify them with severely adverse market conditions or severe financial distress of the firm. The prevalence of such conditions varies over time, most likely with business cycle conditions, and the apparent degree of wage rigidity should do so as well. A second reason that may make apparent rigidity in earnings depend on business cycle conditions is the variation of paid hours over the business cycle. The third argument is a version of the Lucas critique. In our context it implies that economic agents will give up their aversion against nominal wage or earnings cuts in times of low inflation; again rigidity were not time invariant. We will take into account these arguments by appropriate modifications in the empirical implementation of the approaches.

In the remainder of this section we will introduce the basic idea of the standard versions of each of the approaches and refer a more detailed discussion of the approaches and the modifications to later sections. The essence of the *skewness-location approach* is best understood by comparing the shape of the two distributions in panel c) and d) of FIGURE 1, that arise from counterfactuals with different location. The second factual distribution arising from the counterfactual which is situated more to the right is less asymmetric than the first and has smaller mass at zero. Consequently measured skewness should depend negatively on measures of location of the counterfactual. This notion of a negative skewness-location relationship is the assumption specific to the skewness-location approach; it will be discussed in detail in section 5. If such a functional relationship between measures of asymmetry of the factual distribution and measures of location of the counterfactual distribution can be shown to exist, then the existence of nominal rigidity follows.

The *histogram-location approach* also exploits information revealed by changes in location of the counterfactual distribution. Counterfactual and factual distributions are modeled (nonparametrically) as median-centered histograms. Changes in the shape of the factual distribution can be traced in much more detail than by an overall measure of shape. Also attention can be focused on a small part of the left tail of the distribution. As a specific assumption of the histogram-location approach a functional form of the rigidity function has to be assumed, in addition to the assumptions that it has in common with the skewness-location approach. Due to the explicit modeling of the counterfactual distribution it is possible to directly assess the quantitative impact of the rigidity.

Both, the skewness-location approach and the histogram-location approach make no assumption about the form of the median-centered counterfactual distribution, but about the time invariance of the counterfactual distribution (and in the standard versions of the approaches also about the time invariance of the rigidity function). In a sense the *symmetry approach* follows
an opposite strategy. Its core assumption is that the counterfactual distribution is symmetric around the median (A4), i.e. one about form, whereas time invariance is not assumed, neither of the rigidity nor of the counterfactual distribution. The approach makes more extensive use than the other approaches of the fact that counterfactual and corresponding factual distributions coincide over ranges of changes that are not affected by rigidity (A1, A2). Under these assumptions the complete right tail of the factual distribution is identical to that of the counterfactual distribution. Using the assumed symmetry of the counterfactual distribution, its left tail can be inferred from the right tail of the factual distribution. As in the histogram-location approach factual and counterfactual distributions are modeled explicitly, but this time by kernel density estimates instead of histograms; again, direct quantitative conclusions are possible.

In addition to these relatively sophisticated approaches various types of descriptive analysis have been suggested and performed in the literature. Our analytical framework is also informative on which of these types of analysis are helpful, and which are not, since it is the same set of building blocks and assumptions that is behind such analyses, if only implicitly. From the above discussion it is clear that the presence of rigidity can be identified in two basic ways, by shifts in location, and by knowledge of or assumptions on the form of the counterfactual. The shape or degree of asymmetry of some distribution in themselves do not reveal much about the presence of nominal rigidity, since they may be characteristics of a particular counterfactual distribution. Similar considerations holds for the spike at zero. Without relating it to the other parts of the distribution not much can be inferred from its existence, not even from its changes with location.

In the brief descriptive analysis that we present after introducing our data in the next section, we will limit ourselves to the more informative variants of descriptive analysis which in some ways amount to an eyeballing version of the skewness-location approach. Thereafter we will go on to discuss each approach and its application to our data.

3 Data

Our analysis is based on the second issue of the IAB-Beschäftigtenstichprobe (IABS) covering the years 1975-1995. In August 1999 this dataset has been made available for scientific use by the research institute of the Federal Employment Service (‘Institut für Arbeitsmarkt- und Berufsforschung’, IAB). The IABS is a 1% random sample drawn from the German social security accounts which represents one of the most important data sources for labor market research in Germany. The social insurance procedure introduced in 1973 compels employers to report at least once a year all earnings for those employees who are subject to compulsory social insurance. As shown by comparisons with ‘Mikrozensus’ data, the social security accounts basically cover all dependent employment in the private sector, i.e. almost 80 percent of total employment in Germany. The remaining 20 percent of total employment consist of civil servants, self-employed, unpaid family workers and employees who do not pay social security contributions because their earnings or working-time are too low.

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6 See Bender, Haas and Klose (2000) for a description of the new version of the IABS.
The greatest advantage of the IABS is the high reliability of the earnings data due to plausibility checks that are performed by the social security institutions and legal sanctions for misreporting. Contrary to most other studies based on survey data, measurement error due to false or erroneous reporting is therefore not a problem in our analysis. Furthermore, in US studies based on PSID data it is often regarded as a serious shortcoming that no information on fringe benefits is available. The reason is that the finding of nominal wage rigidity does not exclude the possibility that total compensation and hence employers’ labor costs remain flexible, at least to some degree. Hence it is important to note that the earnings data in the IABS refer to gross earnings excluding employer’s contributions to social security but including all kinds of fringe benefits. However, as explained in appendix A, there are also some drawbacks of the dataset, namely the censoring of the earnings distribution, a structural break in the earnings measure in 1984, the truncation of gross daily earnings to integer values and the lack of information about hours worked.

The focus of our study are year-to-year changes of log earnings that are computed on the individual level for full-time male employees working in the western part of Germany. The data selection is explained in more detail in appendix A and can be summarized as follows: We restrict the analysis to „job stayers“, i.e. employees who remained at the same employer for two consecutive years for the full length of time. For our analysis it is important to construct a subsample of log earnings changes for which the observed individual characteristics of the employees do not change and are therefore not responsible for the observed earnings changes. For this reason we require that the training degree, the profession and the occupational status remain unchanged for two consecutive years. We also request that the marital status and the number of children is the same in adjacent years. We only take skilled and unskilled employees into account, whereas high-skilled employees are excluded. We also restrict the sample to employees being at least 25 and at most 65 years old. Furthermore we consider only workers and salaried employees in manufacturing and services. We deal with the possibility of a structural break in 1984 in different ways: in one variant of the analysis we consider the whole sample and explicitly test for a structural break in 1984, in another variant we exclude all observations before 1984 and consider only the shorter time period from 1984-1995. Due to our data selection we are left with 608,965 observations of year-to-year log earnings changes for the full sample period 1975-1995.

4 Descriptive Evidence

We now proceed to present some descriptive evidence for the existence of downward nominal earnings rigidity. From the theoretical discussion of the last section it follows that the exami-
nation of a single factual distribution does not necessarily reveal much insight about the extent of downward nominal rigidity since it cannot be excluded that the counterfactual distribution of log earnings changes is itself asymmetric. The most convincing evidence for or against the downward nominal rigidity hypothesis is therefore obtained by examining whether the shape of the factual distribution systematically changes with varying locations of the counterfactual distribution. In this section we first inspect histograms of log earnings changes for different years and then examine whether the shape of the factual distribution of log earnings changes as measured by various skewness statistics points to a more asymmetric distribution in periods with more leftward location.11

In FIGURE 2 we consider the distributions of log earnings changes for different years. Due to space limitations only the odd-numbered years from 1979 to 1995 are presented. The distributions are shown separately for workers and salaried employees. The reason is that the counterfactual distributions may differ for these groups since workers are paid on an hourly basis whereas salaried employees receive a monthly salary. FIGURE 2 gives some impression of the range within which the location of the distributions of log earnings changes has shifted over the course of the years. The location of the distributions is influenced by the rate of inflation. For instance, in 1979 inflation was equal to 4.1 percent whereas in 1985 it was equal to 2.2 percent and in 1986 and 1987 it was even negative or roughly equal to zero, respectively. However, the location of the distribution is also affected by other factors, most prominently by productivity growth. For this reason instead of the inflation rate one should use the median as the superior indicator for the location of the distribution.12 For the odd-numbered years presented in FIGURE 2 the medians of the distributions lie between 3.5 and 6.2 percent for salaried employees and between 2.9 and 6.5 percent for workers.

FIGURE 2 provides first evidence that the distribution of log earnings changes of both workers and salaried employees is affected by the presence of downward nominal rigidity. It can be seen that for all years the left tail of the distribution exhibits some „deformation“. There is a step in the distribution at zero and some thinning in the distribution below zero. The important point is that a relationship seems to exist between the location of the distribution and the extent of deformation of the left tail of the distribution. For instance, in 1979 the median for both groups is quite high (6.0 and 6.5 percent, respectively). As a consequence the effects of nominal downward rigidity seem to be nearly negligible. On the other hand in the years 1985 to 1987 the medians are at their respective lower limits of 3.5 and 3.2 percent. For these years a spike at zero and a pronounced thinning of the distribution for negative earnings changes below zero can be found.

11 In contrast to other studies we do not report the proportions of negative and zero changes for different years. The reason is that these figures are always lower when the counterfactual distribution shifts to the right, hence no direct inference about downward nominal rigidity is possible. The report of the zero changes would only make sense if also the changes in the neighborhood of zero are documented.

12 The correlation between inflation and the median of the distribution is 0.49 for both workers and salaried employees. Although the inflation rate is the inferior location indicator we sometimes use it as a rough measure for the location of the distribution.
In FIGURE 2 the distributions for salaried employees seem to exhibit less variation than the distributions for workers. The visual evidence also suggests that negative log earnings changes are somewhat more common for workers than for salaried employees. It is very probable that a higher percentage of negative changes for workers mainly reflects changes in working hours and not changes in hourly wages. A closer look at the histograms of workers reveals that in 1979 there seems to be more mass in the right tail of the distribution than in other years. This leads to a more right-skewed distribution in 1979 than in other years. We have no convincing
explanation for this special shape in 1979 and take account of it by introducing a dummy for the respective year when appropriate.

In FIGURE 3 we look at a histogram of log earnings changes for salaried employees, but this time only over a much smaller range to get a more precise impression of the distribution of small earnings changes. This close-up reveals that there are outright ‘holes’ in the distribution next to the zero bin that could be due to menu cost effects. Unfortunately, the truncation of the earnings variable to integer values in the IABS also leads to ‘holes’ and ‘secondary spikes’ to the left and to the right of the spike at zero. In our analysis it is therefore impossible to reach any conclusions about the existence of menu cost effects.

The presence of downward rigidity should increase the skewness to the right of the factual distribution compared to the counterfactual. For this reason we computed various test statistics for the skewness of the distribution and the respective standard deviations based on McLaughlin (1999). A well-known measure for the skewness of a distribution is the skewness coefficient, computed as the ratio of the third central moment to the cubed standard deviation of the respective variable. The mean-median difference is based on the equality of mean and median in the case of a symmetric distribution. For a distribution skewed to the right the mean (usually) lies to the right of the median, hence the difference between mean and median should be significantly positive. The sign-test statistic counts the observations between mean and median. The higher the (positive) number of observations between mean and median, the more likely is the distribution skewed to the right. To make a comparison of the respective numbers possible, the sign-test statistic is normalized by dividing the statistic by its standard deviation. For the thinness measure, which has been proposed by Lebow et al. (1995), the proportion above twice the median minus the proportion below zero is computed. If the distribution were symmetric about the median the thinness measure would be equal to zero, since zero and twice the median are the same distance from the median. A positive test statistic for the distribution implies that there are ‘too few’ nominal wage cuts. In this sense the thinness measure could be interpreted as a more direct measure for the extent of nominal wage rigidity. This measure has the advantage that the impact of extreme wage observations on the thinness is by definition much lower than on the other skewness test statistics. To assess the impact of extreme obser-

13 Our earnings data are truncated integer values of the daily fraction of annual earnings. In each year the minimum log difference (with respect to the previous year) that is theoretically possible is equal to the log difference between the maximum ‘daily’ earning in the previous year to the next higher integer value. These minimum log differences have changed considerably; for instance, their values for 1976/75, 1985/84, 1995/94 are 1.13%, 0.62% and 0.42%, respectively.
vations on the skewness of the distribution we computed the various measures also for the 35-
percentage-point band around the median.14

**TABLE 1: Skewness Test Statistics for Log Nominal Earnings Changes**
*by Category and Different Periods*

<table>
<thead>
<tr>
<th>Period:</th>
<th>All Employees</th>
<th>Workers</th>
<th>Salaried Employees</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Full band</td>
<td>±35 pt</td>
<td></td>
</tr>
<tr>
<td></td>
<td>Mean Infl. (2)</td>
<td>band (6)</td>
<td></td>
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<tr>
<td></td>
<td>Mean-med. (4)</td>
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<tr>
<td></td>
<td>Sign test</td>
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<td></td>
<td>Thinness (5)</td>
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<tr>
<td></td>
<td>Median (perc.)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1975-95(1)</td>
<td>Skewness</td>
<td>0.36</td>
<td>0.191</td>
</tr>
<tr>
<td>(Mean Infl. (2) = 3.1)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Std. Infl. (4) = 1.6)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Mean-med. (4)</td>
<td>0.261</td>
<td>0.240</td>
</tr>
<tr>
<td></td>
<td>Sign test</td>
<td>40.08</td>
<td>36.94</td>
</tr>
<tr>
<td></td>
<td>Thinness (5)</td>
<td>4.84</td>
<td>4.80</td>
</tr>
<tr>
<td></td>
<td>Median (perc.)</td>
<td>4.51</td>
<td>4.37</td>
</tr>
<tr>
<td>1975-83(1)</td>
<td>Skewness</td>
<td>0.132</td>
<td>0.132</td>
</tr>
<tr>
<td>(Mean Infl. (2) = 4.4)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Std. Infl. (4) = 1.2)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Mean-med. (4)</td>
<td>0.205</td>
<td>0.201</td>
</tr>
<tr>
<td></td>
<td>Sign test</td>
<td>21.95</td>
<td>18.79</td>
</tr>
<tr>
<td></td>
<td>Thinness (5)</td>
<td>3.72</td>
<td>3.70</td>
</tr>
<tr>
<td></td>
<td>Median (perc.)</td>
<td>5.36</td>
<td>5.24</td>
</tr>
<tr>
<td>1984-95(1)</td>
<td>Skewness</td>
<td>0.544</td>
<td>0.242</td>
</tr>
<tr>
<td>(Mean Infl. (2) = 2.2)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Std. Infl. (4) = 1.4)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>Mean-med. (4)</td>
<td>0.37</td>
<td>0.341</td>
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<tr>
<td></td>
<td>Sign test</td>
<td>42.35</td>
<td>39.95</td>
</tr>
<tr>
<td></td>
<td>Thinness (5)</td>
<td>6.26</td>
<td>6.21</td>
</tr>
<tr>
<td></td>
<td>Median (perc.)</td>
<td>3.85</td>
<td>3.80</td>
</tr>
</tbody>
</table>

Note: (1) The subperiods do not contain the changes in log earnings from 1983 to 1984, due to the possibility of a structural break. (2) The inflation rate is based on the consumer price index. (3) Std = standard deviation. (4) Mean-median difference. (5) The coefficients of the skewness test statistics are significant at the 1-percent level; the significance of the thinness measure has not been tested. (6) 35-percentage-point band around the median.

In TABLE 1 the values for the respective skewness measures are reported for the full sample period and for the subperiods 1975-83 and 1984-95 with different average location of the respective distribution. It can be seen that the median of the distribution of log earnings changes for both groups of employees is higher in the first subperiod with high inflation than in the second subperiod with low inflation. The results in TABLE 1 support the downward nominal rigidity hypothesis, since for both groups, workers and salaried employees, all skewness measures point to a more right-skewed distribution in the low-inflation period 1984-95 than in the high-inflation period 1975-83. For instance, from the first period to the second, the skewness of the distribution increases from 0.14 to 0.47 for workers and from 0.25 to 0.88 for salaried employees. Similar results are obtained if the overall distribution comprising both workers and salaried employees is considered. For example the skewness statistic increases from 0.13 to 0.54 and the thinness measure increases from 3.72 to 6.26. From TABLE 1 it is also obvious that the conclusions are the same if the 35-percentage-point band around the median is considered. Overall, the results are compatible with the claimed effects of downward nominal rigidity on the shape of the factual distribution caused by shifts in the counterfactual distribution as explained in the theoretical discussion of section 2.

---

14 A x% point band around the median m is defined to range from m (1 - .01x) to m (1 + .01x).
When interpreting the results in TABLE 1 it must be stressed that a cross-section comparison of skewness measures across different groups of employees is not informative about the relative extent of downward nominal rigidity. The reason is that the shapes of the underlying counterfactual distributions themselves may differ for different groups of employees. Hence the fact that in TABLE 1 the skewness coefficients are higher for salaried employees than for workers does not prove that rigidity is more prevalent for the first group than for the latter. This discussion makes it evident that an examination of the nominal rigidity hypothesis on the basis of skewness measures makes only sense if changes of the respective statistic caused by shifts of the distribution are considered. In principle, the comparison of skewness test statistics for periods with different location as in TABLE 1 is already based on the idea of the skewness-location approach to which we now turn.

5 Skewness-Location Approach

The skewness-location approach was introduced by McLaughlin (1994) and applied by Lebow, Stockton and Washer (1995) and McLaughlin (1999) to US data. If rigidity is present, the shape and the degree of asymmetry of the factual distribution are changed in a characteristic way by changes in location of the underlying counterfactual distribution. According to the approach a functional relationship between measures of skewness and measures of location exists, since skewness to the right of the factual that is caused by the direct effects of nominal rigidity is alleviated by shifts of the counterfactual to the right and exacerbated by shifts to the left. Data on skewness and location over the sample period is used to show the existence or nonexistence of such a skewness-location relationship in econometric estimates. We will first discuss the skewness-location relationship in more detail, then introduce our empirical implementation and finally present and discuss the results for our data.

Skewness-location approach and skewness-location relationship

The notion of a characteristic relationship between skewness and location is specific to the skewness-location approach. It is summarized in the following assumption:

A6 Skewness-location relationship
There is a falling, approximately linear functional relationship between measures of skewness of the factual distribution and measures of location of the shifting underlying counterfactual distribution.

The skewness-location relationship is never analyzed in any detail in the literature but merely based on plausibility arguments. Based on theoretical reasoning and numerical simulations it is shown in Knoppik (2000) that even in ordinary cases (normal counterfactual distribution and constant proportional downward rigidity) the skewness-location relationship is in general much more complicated than assumed in A6. It is nonlinear, non-monotonous, and rigidity may even cause skewness to the left in certain circumstances. Notwithstanding these objections scatter plots of skewness measures against the median of the distribution indicate that for our data assumption A6 is quite reasonable. Two measures of location are used in the literature, inflation and median. Obviously inflation is an important determinant for explaining shifts in the earnings change distribution. But from the conceptual framework in section 2 it has become clear that the shifts in location of the counterfactual are decisive, and that location is best measured by the median. The usual measures of skewness are those that were used in section 4, the skewness coefficient, the mean-median difference, the sign-test statistic and the
thinning measure. As shown in Knoppik (2000) one of the disadvantages of the skewness coefficient is due to the fact that it is centered around the mean, which is itself influenced by the rigidity. Under assumptions A1 and A2 the median of the factual distribution remains unchanged from that of the counterfactual distribution. Therefore we also use a median-centered skewness coefficient, which in the definition of the skewness coefficient replaces the mean with the median. Before we turn to the estimation results we briefly introduce some modifications in the empirical specification that we see as necessary.

**Empirical implementation**

The empirical counterpart of assumption A6 are linear regressions of various measures of skewness on the chosen measure of location with observations for each year in the sample period. Under assumptions A1, A3 and A5 a skewness-location relationship only exists if there is rigidity. Therefore a significant negative coefficient of the measure of location “proves” the existence of nominal rigidity. However, deviations from the assumptions will obscure the skewness-location relationship, at least to some extent. As we have argued in section 2, the potentially most important deviation from the assumptions concerns the assumption of a time-invariant rigidity. It seems quite plausible that changes (and perhaps also the level) of the unemployment rate influence the process of wage formation. Systematic variation in paid hours over the cycle would also have an impact on the degree of rigidity apparent in earnings data. To take account of these arguments we have included the change and the level of the unemployment rate, $\Delta U_t$ and $U_t$, and a trend in our regressions. The regression equation therefore reads:

$$ MS_t = c + \beta_1 \text{MEDIAN}_t + \beta_2 \Delta U_t + \beta_3 U_t + \beta_4 \text{TREND}_t + \mu_t, $$

where $MS_t$ denotes any of the five measures of skewness.

**Results from regressions**

We have estimated equation (3) for various subsamples and in each case step by step removed insignificant regressors. The key coefficient in equation (3) is the location coefficient $\beta_1$, which characterizes the skewness-location relationship. The estimated coefficients from regressions are shown in TABLE 2. To take account of the possibility of a structural break from 1983 to 1984 we not only present the estimation results for the whole sample period but also for the shorter period from 1984 to 1995. The estimations were separately performed for workers and salaried employees. Trend and the level of the rate of unemployment were both found to be insignificant and were consequently removed from the regression. The change in unemployment was found to be always significant or highly significant and remained in the regression equation. In order to preserve space the corresponding coefficient $\beta_2$ is not shown in TABLE 2.

The location coefficient has the right sign in all estimates, i.e. it is negative. The skewness of the factual distributions is in fact reduced by shifts to the right of the underlying counterfactual distribution. For the full sample period all the coefficients are significant or highly significant for both workers and salaried employees. For the shorter period 1984-1995 the coefficient values for salaried employees remain fairly stable, and are somewhat higher for workers. Since there is loss of degrees of freedom if going from the full to the shorter sample period it is not surprising that in the latter case the parameter estimates often are only weakly significant or even insignificant. We performed various stability tests to check whether the estimation results for the whole sample period are reliable. Since we found no evidence for a structural break in 1984 or in other years, our preferred estimation comprises the full sample from 1975 to 1995.
TABLE 2: Skewness-Location Approach: Estimated Location Coefficients

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Workers</td>
<td>Salaried</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>Employees</td>
</tr>
<tr>
<td>Skewness Coefficient</td>
<td>-0.06**</td>
<td>-0.08**</td>
</tr>
<tr>
<td></td>
<td>(-2.73)</td>
<td>(-2.70)</td>
</tr>
<tr>
<td>Median-C. Skewness Coefficient</td>
<td>-0.09**</td>
<td>-0.11***</td>
</tr>
<tr>
<td></td>
<td>(-2.50)</td>
<td>(-2.89)</td>
</tr>
<tr>
<td>Mean-median Difference</td>
<td>-0.06*</td>
<td>-0.07***</td>
</tr>
<tr>
<td></td>
<td>(-1.93)</td>
<td>(-3.06)</td>
</tr>
<tr>
<td>Sign Test</td>
<td>-2.48**</td>
<td>-1.37***</td>
</tr>
<tr>
<td></td>
<td>(-2.35)</td>
<td>(-3.28)</td>
</tr>
<tr>
<td>Thinness</td>
<td>-1.09***</td>
<td>-1.78***</td>
</tr>
<tr>
<td></td>
<td>(-3.32)</td>
<td>(-6.78)</td>
</tr>
</tbody>
</table>

Note: The independent variables in all estimations are a constant, the median and the change in the aggregate unemployment rate. For workers a dummy for 1979 was included (see section 4 for an explanation). The table contains the location coefficient, i.e. the regression coefficient of the median and the respective t-value in parenthesis. *,**,*** denote significance on the 10, 5 and 1 percent level, respectively. The coefficient of the change in the unemployment rate, which is not reported in the table, is significantly negative in all estimates.

Specification and Stability: The Durbin-Watson test statistic is around 2 in nearly all estimates, hence there is no evidence for autocorrelation in the residuals. The $R^2$ varies between 0.41 and 0.78 in the estimates. On the basis of various stability tests (Chow’s breakpoint test, recursive residuals, CUSUM test, one-step and N-step forecast test) the hypothesis of a structural break in 1984 can be rejected. For salaried employees the estimates are stable over the whole sample period. However, for workers there is some evidence for parameter instability in the early years of the sample.

(1) Measures of skewness were computed for observations in a band of ±35 percentage points around the median.

In our estimates the sign of the coefficient for the change in the unemployment rate, $\beta_2$, is always negative, which can be interpreted as evidence that an increase in unemployment reduces the extent of nominal rigidity and therefore leads to a less right-skewed distribution. The ratio $\beta_2 / \beta_1$ is approximately equal to three, which means that a rise in unemployment of one percentage point tends to have three times the effect on measures of skewness as a one percentage point shift of the distribution to the right. It could be argued that our results are biased since the change in aggregate unemployment is itself an endogenous variable being influenced by the wage-setting process which in our estimates is reflected in the shape of the distribution for earnings changes. To deal with the endogeneity problem we performed an instrumental variable estimation where the variable $\Delta U_t$ was replaced by suitable instruments.15 The parameter values in this estimations turned out to be nearly the same as those reported in TABLE 2, hence they are not documented here.

Lebow, Stockton and Washer (1995) have run regressions on the median of the distributions for PSID data. Comparing our results with theirs, from our estimates emerges a surprisingly

---

15 As instruments for $\Delta U_t$, we used the changes in the following variables in period ($t-1$): the unemployment rate, unit labor costs, the structural deficit and the interest rate.
clear evidence for the existence of downward nominal rigidity. In order to be able to make comparisons with studies which have used the rate of inflation as the measure of location we also regressed the various skewness statistics on the inflation rate. The results are documented in Table A.2 in appendix B. When interpreting these results one has to bear in mind that due to our theoretical discussion it is clear that the inflation rate is only an imperfect indicator for the location of the earnings change distribution. Considering the estimation results for the full sample period from 1975 to 1995 all coefficients for the inflation rate have the expected sign, i.e. they are negative, and in most cases they are significant. Since the explanatory power of the inflation regressions is lower than in our preferred specification, it is not very surprising that the estimates for the smaller sample period are all insignificant. However, for the full sample period our estimates with the inflation rate as the independent variable again point more convincingly to the presence of downward nominal rigidity than is the case in comparable estimates for US data.

Although the estimation results of our preferred specification seem to corroborate the downward nominal rigidity hypothesis, the interpretation of our results has to take into account that the skewness measures are measures of the overall shape of the distribution. Hence they are only a crude measure for capturing the effects of downward nominal rigidity. For this reason we will analyze downward nominal rigidity also in the histogram-location approach which limits the analysis to a relatively small part of the left tail of the distribution. It is also in the nature of the skewness-location approach to stop short of the identification of a counterfactual distribution. Therefore it is not possible to directly quantify the impact of nominal rigidity. In order to do just that we discuss two more approaches in the following two sections, the symmetry approach and the histogram-location approach. Both explicitly model the factual and counterfactual distributions and therefore make quantitative conclusions possible.

6 Symmetry Approach

In order to make direct quantitative assessments of the effects of downward nominal rigidity one has to identify the counterfactual distribution. One way to proceed is to follow the symmetry approach that was introduced in section 2. Under the assumptions that the median of the counterfactual is not affected by rigidity (A1, A2) and that the counterfactual distributions are symmetric around the median (A4), the left tail of the counterfactual can be inferred from the right tail of the factual distribution and factual and counterfactual distributions can be compared to assess the effects of downward nominal rigidity.

The most critical assumption of the symmetry approach is that of symmetry of the counterfactual. It may even be violated in Card and Hyslop (1997), the study that introduced the symmetry approach. McLaughlin (1999) has shown for the PSID data that there is considerable asymmetry in the distribution of wage changes in the immediate neighborhood of the median.

---

16 In Lebow, Stockton and Washer (1995) only for the thinness measure a significant negative relationship with the median of the distribution is found.

17 For instance in McLaughlin (1999) the correlations of the various skewness test statistics with inflation for all but the thinness measure show the wrong sign and are statistically insignificant (based on 21 annual observations). Hence McLaughlin [(1999), p. 129] concludes that ‘wage changes in the PSID contain no evidence of less skewed distributions of wage changes in years when inflation was higher’. Similar regression results for PSID data are found by Lebow, Stockton and Washer (1995).
i.e. in a region of the distribution, that should be unaffected by rigidity. His conclusion is that the whole counterfactual distribution is asymmetric and that the symmetry approach should not be used. In order to check the appropriateness of the symmetry assumption we analyze close to median symmetry in two ways, by computing skewness statistics for the observations in a narrow band around the median and by constructing symmetrically-differenced histograms, as in McLaughlin (1999).


<table>
<thead>
<tr>
<th></th>
<th>Workers</th>
<th>Salaried Employees</th>
</tr>
</thead>
<tbody>
<tr>
<td>Frequency in 2% point band</td>
<td>164,949</td>
<td>48,813</td>
</tr>
<tr>
<td>Percent of observations in category</td>
<td>33.84</td>
<td>40.19</td>
</tr>
<tr>
<td>Skewness coefficient</td>
<td>0.044</td>
<td>0.074</td>
</tr>
<tr>
<td>Mean-median difference</td>
<td>0.025</td>
<td>0.028</td>
</tr>
<tr>
<td>Sign-test statistic</td>
<td>745.5</td>
<td>440.5</td>
</tr>
</tbody>
</table>

Note: Skewness measures are computed for 2% point band around the median of respective sample. Each skewness statistic is significant at the 1% level.

(1) A 2% point band is defined to range from \( m(1 - .02) \) to \( m(1 + .02) \).

TABLE 3 shows skewness statistics that were computed for those observations that lie in a 2% point band around the median. If compared with their standard errors the statistics significantly indicate asymmetry close to the median. For further evidence we construct symmetrically-differenced histograms that are a representation of the differences between the relative frequencies of those bins of an ordinary median-centered histogram that are equally distant from the median. Bars for the differences are plotted against the distance from the median. Panel a) of FIGURE 4 is a symmetrically-differenced histogram for the year 1985, which is quite typical for the other years in the sample period 1975-1995. The differences for changes close to the median are clearly negative, implying that changes slightly above the median occur less often than the corresponding changes below the median. This observation corroborates the evidence from skewness statistics that the distributions are remarkably asymmetric in a narrow range around the median. Since this asymmetry is not caused by rigidity, most likely the complete counterfactual distributions are not symmetric, and the symmetry approach is not applicable.

The problems that are caused by the asymmetry of the earnings change distribution around the median are also evident in panel b) of FIGURE 4 where the kernel density estimates of the factual and the corresponding counterfactual distributions are plotted for the year 1985. The counterfactual generally lies above the factual for negative changes and for small changes of either sign.

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18 Finding symmetry close to the median does not allow one to make conclusions about the symmetry of the outer part of the tails of the counterfactual distribution. But while (strictly speaking) evidence for asymmetry from skewness statistics cannot show the asymmetry of the outer parts of the tails of the distributions, the evidence from symmetrically-differenced histograms can. Therefore we agree with McLaughlin (1999) and think that close to median asymmetry does indeed indicate overall asymmetry of the distribution.

19 The counterfactual was constructed using the assumption of symmetry (A4) for purpose of illustration. A normal bandwidth of \( h = .005 \) and an Epanechnikov kernel were used. A smaller bandwidth was used around zero.
These two observations are compatible with the workings of nominal rigidity. However, the factual is also greater than the counterfactual distribution for changes slightly below the median, where according to the assumptions the two distributions should be identical. This again indicates that the central assumption of the approach is not appropriate for our data, since the counterfactual distribution is not symmetric about the median.

7 Histogram-Location Approach

The histogram-location approach, which is due to Kahn (1997), was already introduced in section 2 and placed in our general analytical framework. In the histogram-location approach factual and counterfactual distributions are directly represented as distributions by median-centered histograms. Dealing with a complete model of the distributions and not only with indicators promises the advantage of being able to assess the extent of rigidity and not only to observe its existence. The key idea of the histogram-location approach is to explain bin sizes of the factual median-centered histograms in different years by two determinants, the counterfactual median-centered histogram and the rigidity function. As in the skewness-location approach the shifts of location and the resulting variation of the factual distribution over time play the crucial role. Here in addition they serve to disentangle the two determinants of the factual distribution, which are both estimated in the process.

We will start by presenting the basic idea graphically and then develop the estimation equations. Again we modify the basic approach from the literature to take possible variability in the degree of rigidity into account. Finally we present and discuss the estimation results, which reveal substantial downward nominal rigidity in our data.

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20 The kernel density estimates in Card and Hyslop (1997) reveal the same property upon closer inspection. This supports the criticism of the application of the symmetry-approach to PSID data in McLaughlin (1999).
The effects of nominal rigidity on median-centered histograms

FIGURE 5 and FIGURE 6 illustrate the basic idea of the histogram-location approach in its standard version (with time-invariant rigidity function). The approach uses median-centered histograms for the counterfactual and factual distributions that have the advantage to be directly comparable across time, since the assumption of an invariant median-centered counterfactual distribution (A3) ensures an invariant median-centered counterfactual histogram. Under the assumption of sufficiently high medians (A2) together with the assumptions of only direct effects of rigidity (A1) the medians of counterfactual and factual distribution are identical. This ensures that bins of factual and counterfactual median-centered histograms cover the same intervals of changes. FIGURE 5 shows median-centered histograms that correspond to the counterfactual and factual distributions in panel a) and c) of FIGURE 1 in section 2. Panel a) of FIGURE 5 depicts the counterfactual median-centered histogram. The proportion of observations that fall into a bin $i$ are denoted by $PC_i$ (for ‘proportion counterfactual bin $i$’). Panel b) of FIGURE 5 contains a factual median-centered histogram; its proportions of observations that fall into a bin $i$ are denoted by $PF_i$ (‘proportion factual’). The factual histogram differs from the counterfactual histogram by ‘thinning’ and ‘piling-up’. The ‘piling-up’ of non-enacted cuts is seen in the bin that contains the zero change observations. For brevity this bin is called the ‘zero bin’. The ‘thinning effect’ is seen in the bins containing negative changes to the left of the zero bin, or briefly in the ‘negative bins’; it is caused by the non-enacted cuts. The ‘positive bins’ above the zero bin are unchanged from the counterfactual histogram, unless there are menu cost effects. In that case only positive bins at more than a certain distance from the zero bin are unaffected.

Whenever the location of the counterfactual distribution shifts, the zero bin will be located at different distances from the median. Therefore, the factual distribution and the median-centered factual histogram have different shapes over time, despite the time-invariant median-centered counterfactual distribution and the time-invariant rigidity function. The reason is that the small and negative changes affected by rigidity, will be located at different distances from the median in different years.

![Median-centered histograms](image)

**FIGURE 5: Median-Centered Theoretical Histograms of the Counterfactual and Factual Distributions**

Note: $x$ denotes the changes in earnings, $m$ denotes the median. $PC_i$, $PF_i$ are the proportion of observations contained in bins of the counterfactual and factual histograms, respectively.
It is useful to index the bins in the left tail of the distributions from right to left, i.e. from close to median to far from median. The convention is to start the index with zero. This implies that bin $r$ contains changes that are at least $r$ times, and at most $r + 1$ times the bin width $b$ smaller than the median change. The zero bin in period $t$ is denoted by index $r_t^0$. Only for those bins that are in some years affected by the rigidity and are not affected by it in other years it is possible to distinguish the counterfactual distribution from the effects of rigidity. Therefore only a small part of the histograms in FIGURE 5 is useful for the analysis: The relevant range of bins $r_{\min} \ldots r_{\max}$ covers $R = r_{\max} - r_{\min} + 1$ bins, where $r_{\min}$ and $r_{\max}$ are the smallest and biggest index numbers of zero bins $r_t^0$ over the years.\textsuperscript{21}

\begin{figure}[h]
\centering
\includegraphics[width=\textwidth]{figure6}
\caption{Counterfactual and Factual Median-Centered Histograms and the Role of Location for the Effects of Nominal Rigidity}
\end{figure}

\begin{itemize}
\item[a)] Identical median-centered counterfactual histograms
\item[b)] (counterfactual) zero bin
\item[c)] Factual histogram, low median
\item[d)] Factual histogram, high median
\item[e)] Effects of rigidity, low median
\item[f)] Effects of rigidity, high median
\end{itemize}

Note: Left-hand side panels refer to a period with low median, right-hand side panels to a period with high median. Median-centered counterfactual histograms are time-invariant. The shape of median-centered factual histograms does depend on the median of the period.

The left and right hand side panels of FIGURE 6 show close-ups of histograms over the relevant range of bins for two different periods. The left hand side illustrates a period $L$ with relatively low median and the right hand side a period $H$ with relatively high median. For each period left

\textsuperscript{21} This statement has to be modified slightly if menu-cost effects are modeled.
and right hand side of FIGURE 6 show in vertical order: First, in panels a) and b) the time-invariant and therefore identical median-centered counterfactual histograms with bin sizes $PC_r$. Second, in panels c) and d) the time-variant median-centered factual histograms in the periods with bins sizes $PF_{rt}$. And finally, in panels e) and f) the difference between factual and counterfactual histograms, i.e. the effects of rigidity on different bins, $RIG_{rt}$. Note again, that the effects of rigidity and therefore the factual distributions differ between the two periods with different medians, although rigidity function and median-centered counterfactual distribution do not. Under the assumption that there are only direct effects of nominal rigidity (assumption A1) and in the absence of menu cost effects, the positive bins of the factual histograms equal the corresponding counterfactual bins

$$PF_{rt} = PC_r, \quad r > r_t^0.$$ 

The negative bins and the zero bin of the factual differ by the effect of rigidity from the corresponding counterfactual bins$^{22}$

$$PF_{rt} = PC_r + RIG_{rt}, \quad r \leq r_t^0.$$ 

If in different years there are different medians, some bins are positive bins in years with high median, and negative bins in other years with low median. E.g., in panels c) and d) of FIGURE 6 this can be seen to be the case for bins 11 to 15. Additional information is provided by each period’s zero bin, which only in that particular period contains the piled-up probability of non-enacted wage cuts of that period.$^{23}$ Together these bins provide the essential information for the skewness-location approach.

In short, the time-variant bin sizes of the observed factual $PF_{rt}$ are explained by constant counterfactual bin sizes $PC_r$, and the location-dependent (and therefore time-dependent) effects of rigidity $RIG_{rt}$. An error term $\mu_{rt}$ is added since in the empirical implementation sample histograms instead of theoretical histograms are considered. This summary is equivalent to the following equation,

$$PF_{rt} = PC_r + RIG_{rt} + \mu_{rt},$$

which is a discrete version of equation (2). The next step is to find a version of equation (4) that can be estimated.

### Estimation equations

In order to estimate equation (4) an assumption on the functional form of the rigidity function $\rho(\cdot)$ has to be made. The visual evidence of section 4 together with the remarks on the effects of the truncation of the earnings variable in our data in that section have led us to adopt a proportional downward rigidity without menu cost effects.$^{24}$

#### A7 Functional Form of Rigidity Function

The rigidity function has proportional form

$$\rho(z) = \rho z, \text{ for } z < 0.$$ 

$^{22}$ The index numbers of the zero bins $r_t^0$ are equal to 10 for $t = L$ and equal to 16 for $t = H$ in FIGURE 6.

$^{23}$ E.g., bin 10 is a regular positive bin in period H, i.e. in panel d), but it is the zero bin in period L, i.e. in panel c). Similarly, bin 16 is the zero bin in panel d) and a negative bin in c).

$^{24}$ In an earlier version of this paper we reported estimation results with menu cost effects. They were badly affected by the trend in the truncation effect described in footnote 13.
This amounts to a model with a single constant rigidity coefficient $\rho$ measuring the proportion of non-enacted earnings cuts, as discussed and illustrated in section 2. Besides the assumption about the functional form of rigidity, dummy variables are needed which indicate the location of bins relative to the zero bin for each distance from the median $r$ in each period $t$. The dummy variables are $DNEG_{r,t}$ to indicate negative bin status and $D0_{r,t}$ to indicate zero bin status. Finally, as will be discussed subsequently, for the modeling of the error term it is useful to have $R$ separate equations with $T$ observations, instead of a single equation for $R*T$ observations as in equation (4). Using the functional form and the dummy variables one obtains the following basic specification, which amounts to a ‘lean’ version of Kahn’s (1997) so-called proportional model.

\[
(5) \quad PF_{r,t} = \alpha_r \left( \begin{array}{c}
-p\alpha_r DNEG_{r,t} + \left( \gamma + \rho \sum_{j=\min}^{\max} \alpha_j DNEG_{j,t} \right)D0_{r,t} + \mu_{r,t} \text{ for } r = r_{\min} \ldots r_{\max},
\end{array} \right)
\]

As in equation (4), factual bin size $PF_{r,t}$ is explained by counterfactual bin size $\alpha_r$ (renamed from $PC_{r,t}$) and the two direct effects of rigidity, i.e. thinning and piling-up. Thinning by $\rho$ only takes place for negative bins. Pile-up only takes place in zero bins and consists of two terms: A summation term, that collects the freezes from the negative bins within the range $r_{\min} \ldots r_{\max}$ and second, a constant $\gamma$ for the freezes from the negative bins that are always negative and therefore outside that range. Note that equation (5) is a quite parsimonious specification of equation (4), since only $R$ counterfactual bin sizes $\alpha_r$, the time-invariant rigidity coefficient $\rho$ and the (‘from-outside-’) pile-up constant $\gamma$ have to be estimated.

We consider it important to check whether the assumption of time invariance of the rigidity function is appropriate (A5), for the reasons discussed in section 2. Therefore we modify equation (5) and let the total proportion of non-enacted earnings cuts depend on the change in the rate of unemployment and on a dummy variable $D84$ that marks post 1983 observations.

\[
(6) \quad PF_{r,t} = \alpha_r \left( \begin{array}{c}
-p\alpha_r DNEG_{r,t} + \left( \gamma + \rho \Delta U_t + \rho^{D84} D84_{t} \right)\alpha_r DNEG_{r,t} \right)
\]

\[
\left( \begin{array}{c}
\left( \gamma + (\rho^{AU} \Delta U_t + \rho^{D84} D84_{t}) \right) \sum_{j=1}^{R} \alpha_j DNEG_{j,t} \right)D0_{r,t} + \mu_{r,t} 
\end{array} \right)
\]

for $r = r_{\min} \ldots r_{\max}$.

One would expect $\rho^{AU}$ to be negative, if cyclical downturns, measured as increases in unemployment $\Delta U_t$ relax nominal rigidity. Equally $\rho^{D84}$ is negative, if rigidity is lower in the second part of the sample. We will report unrestricted and restricted estimates of this equation. Note that equation (6) is chosen to contain the basic specification of equation (5) as a special case.25

Bin sizes at different distances from the median, i.e. for different $r$, differ by up to more than a factor of ten. It therefore seems likely that their error terms $\mu_{r,t}$ exhibit different standard errors, i.e. there is heteroskedasticity across equations. This can be taken into account by estimating the system of equations by weighted least squares, WLS. Kahn (1997) has proposed to also take account of the contemporaneous correlations between errors of the different equa-

\[25\text{ We have estimated some alternative specifications that relax the assumption of a time-invariant median-centered counterfactual distribution, but found no compelling reasons to abandon this assumption.}\]
tions of the system by using iterative SUR. This proposal raises two issues. First, it amounts to giving up the assumption of a time-invariant median-centered counterfactual distribution (A3). If one wishes to relax this assumption, probably a more parsimonious specification should be used. Second, there is the question of the small sample efficiency of SUR. Depending on the subsample under consideration, the number of bins \( R \) (and therefore equations) is around ten and the number of observations \( T \) is up to twenty. We have conducted a specification analysis by systematically varying the sample size by limiting the estimation to various subperiods. This has led to strongly varying parameter estimates under SUR, but rather stable estimates under WLS. We conclude that the time dimension of the sample \( T \) is too small to usefully include contemporary correlation. All our subsequent estimation results are therefore based on WLS.

**Estimation results**

The estimation results for the model specified in equation (6) for complete and split sample periods 1975-95, 1975-83, 1984-95 and for the two groups of workers and salaried employees can be found in TABLE 4. In all estimates the rigidity coefficient \( \rho \) has the expected sign, i.e. it is positive and it is clearly significant if judged by the asymptotic \( t \)-values, except for workers in one specification for the earlier subperiod. This means that a substantial percentage of the reductions in real earnings did not take place if they required nominal cuts.

For the complete period 1975-95 for the specification including \( \Delta U \), we find a rigidity coefficient \( \rho \) of more than 20 percent for salaried employees. The rigidity coefficient \( \rho \) is lower for workers, around 10 percent. The difference between salaried employees and workers conforms to the visual impression from section 3. For salaried employees we find a substantial sensitivity to changes in unemployment. An increase of unemployment of one percentage point reduces the skewness coefficient \( \rho \) by eight percentage points (and vice versa), which is compatible with the considerations in section 2. The corresponding coefficient for workers is smaller and not significant. Note that for the full sample period the size of the rigidity coefficients for salaried employees and workers do not depend particularly on the inclusion of the change of the rate of unemployment, as can be seen from columns (5) and (2) of TABLE 4.

If the full sample period is split up into two subperiods, for salaried employees we find a slightly lower rigidity coefficient of around 15 percent in the second part of the sample, but one that is around 30 percent in the first subsample. In both subperiods the sensitivity to changes in unemployment is roughly the same as for the full period. Column (6) of TABLE 4 shows that we obtain comparable results, if we introduce a dummy for post 1983 observations into equation (6) and estimate it for the full sample period. For workers the rigidity coefficients lie around 20 percent in the first and around 15 percent in the second subperiod. Again, this is supported by estimates in a single equation including the post 1983 dummy, as can be seen from column (3) of TABLE 4. The relatively low \( t \)-values may be caused by an increased role of changes in unemployment in the second subperiod, where changes in unemployment are now clearly significant.

---

26 Exclusion of the 1983-84 change for either type of employee or the 1978-79 change for workers from the full sample period did not change the results.

27 In order to obtain the average rigidity coefficient for that period the average change of the rate of unemployment in that period (around half a percentage point) has to be multiplied by \( \rho \Delta U \) and then subtracted from \( \rho \).

### TABLE 4: Estimated Rigidity Coefficients in the Histogram-Location Approach

<table>
<thead>
<tr>
<th>Period</th>
<th>Workers</th>
<th>Salaried Employees</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>1975-95</td>
<td>10.44 (2.69)</td>
<td>8.97 (2.59)</td>
</tr>
<tr>
<td>ρ ΔU</td>
<td>-2.54 (-0.86)</td>
<td>-5.77 (-1.67)</td>
</tr>
<tr>
<td>D84</td>
<td>-11.38 (-1.66)</td>
<td>-8.84 (-1.69)</td>
</tr>
<tr>
<td>1975-83</td>
<td>23.46 (1.33)</td>
<td>19.79 (2.94)</td>
</tr>
<tr>
<td>ρ ΔU</td>
<td>-2.38 (-0.24)</td>
<td>-13.55 (-1.99)</td>
</tr>
<tr>
<td>D84</td>
<td>-11.38 (-1.66)</td>
<td>-8.84 (-1.69)</td>
</tr>
<tr>
<td>1984-95</td>
<td>16.89 (4.68)</td>
<td>13.69 (3.67)</td>
</tr>
<tr>
<td>ρ ΔU</td>
<td>-7.95 (-2.84)</td>
<td>-9.65 (-4.61)</td>
</tr>
</tbody>
</table>

Note: Rigidity coefficients $ρ$, $ρ ΔU$ and $ρ D84$ are reported in percentage points. Dependent variables in the R equations are the proportions of observations contained in the bins of the median-centered factual in each year, $PF r,t$. The bin sizes of the counterfactual median-centered histogram $α$ are all highly significant, but not reported in this table. Bin width $b = .005$. The equations were estimated by iterative weighted least squares (WLS; iteration over weighing matrices and coefficients was simultaneous). Asymptotic $t$-values are reported in parentheses.

Downward nominal rigidity in earnings is clearly present in our data and it is quantitatively important. E.g., based on the estimated rigidity coefficient of approximately 25%, almost 10% of all salaried employees are affected by non-enacted earnings cuts, if one assumes a median nominal wage growth rate of 2%. To put this result into perspective, we compare it with those of Kahn (1997). She is only able to find downward nominal rigidity for salaried employees in a problematic version of the proportional model that includes a time trend and implies rigidity coefficients of the wrong sign in several years, while we find a clear-cut effect of nominal rigidity. There may be behavioral differences between salaried employees in the two countries, but we believe that the quality of our data, our modification of the approach and the alternative estimation procedure also have played a role for our affirmative results for salaried employees. For workers the estimated rigidity coefficient in Kahn (1997) is higher than our estimates, but since it is obtained from data on changes in wage rates it is not directly comparable to our results.

One important aspect of our results is that the rigidity effect is indeed time-variable in two important respects, i.e. it depends on the change of the unemployment rate and it has decreased in the second half of the sample. Concerning the first aspect it is clear that there is an important role of changes in the unemployment rate for nominal earnings rigidity. It is less clear whether it reflects an variable attitude towards cuts in wage or salary rates, or whether it is rather a reflection of cyclical changes in other determinants of total earnings, e.g. paid working hours. Concerning the second aspect, i.e. the decrease in $ρ$, we were not able to decide whether it is
due to a decrease in downward nominal wage rigidity or to an increased variability of non-basic pay components (e.g. increased variability of paid hours), which mask the effects of an unchanged degree of downward nominal wage rigidity. Most likely it has to do with the reduction of weekly working hours in union contracts that has started in the mid 1980s. We are quite certain, that it is not directly caused by changed social security reporting rules, since the one-time major change in reporting rules should only have affected the change in earnings from 1983 to 1984. There is however the possibility that it has led our earnings variable to cover more downward flexible compensation components, e.g. cash bonuses, and thereby to lower apparent downward rigidity. Due to the comprehensive nature of our compensation variable interpretation is not easy, but the central result is still very clear: There is nominal rigidity in earnings and it must have been caused by substantial nominal wage rigidity.

8 Summary and Conclusion

The evidence presented in this paper indicates that in Germany earnings of job stayers are characterized by substantial downward nominal rigidity. Several approaches to the analysis of downward nominal rigidity have been suggested in the literature. Because of the mixed results to which these have led in earlier studies we developed an analytical framework to make explicit their assumptions, differences and commonalities. It turned out that the approaches had to be modified, in order to take into account that the extent of downward nominal rigidity may vary over time. The application of the modified approaches and of various types of descriptive analysis to a sample from the IABS uniformly supported our findings.

The main building blocks of our analytical framework are the counterfactual and factual distribution of relative earnings changes and a rigidity function which transforms the former into the latter. We took up three approaches from the literature, the skewness-location approach, the symmetry approach and the histogram-location approach. Each can be characterized by a set of assumptions about these building blocks, several of which are shared by more than one approach. The skewness-location approach and the histogram-location approach exploit variation in the location of the distributions over the course of years and infer existence and extent of the rigidity from the corresponding variation in measured skewness and variation in the shape of observed histograms, respectively. The symmetry approach assumes the symmetry of the counterfactual and infers its unobserved left tail from the observed right tail. It turned out that the symmetry approach could not be used since the crucial assumption of a symmetric counterfactual distribution of relative earnings changes does not hold in our data. The focus of our analysis was therefore on the other two approaches.

A central assumption shared by skewness-location and histogram-location approach as they have been proposed in the literature is that of a time-invariant rigidity function. However, the survey literature on firm wage policies uniformly reports that there are wage cuts, but only under decidedly adverse market conditions. Since market conditions vary over the business cycle, we concluded that the apparent extent of downward nominal rigidity must depend on business cycle conditions as well. For earnings data, a second reason for such a dependence is variation of paid hours over the business cycle. Therefore we modified skewness-location and histogram-location approach by allowing changes in aggregate unemployment to have an impact on the extent of nominal rigidity. Both modified approaches were applied to our sample
from the IABS, which has several advantages for an analysis of downward nominal rigidity, among which are the reliability of the data, the comprehensiveness of the earnings variable and the existence of data for low inflation years.

The empirical counterpart of the theoretical skewness-location relationship in the skewness-location approach are regressions of various measures of skewness on measures of location and measures of market conditions (according to our modification). The regressions revealed significant coefficients of the expected sign for the median as measure of location and for the change in the rate of unemployment as measure of market conditions. The skewness-location approach therefore clearly supports the hypothesis of downward nominal rigidity in the distribution of earnings changes. The rigidity is lessened in times when unemployment is rising and vice versa. Naturally, a single skewness measure can capture changes in shape, brought about by rigidity, only in a relatively crude manner.

The histogram-location approach is much more specific in this respect, since counterfactual distribution and rigidity function are explicitly modeled and estimated, and the focus is on only that small part of the left tail of the distributions that actually provides information about downward nominal rigidity. Again, the hypothesis of downward nominal rigidity is supported by the estimation results. But not only inferences about the existence, but also those concerning the extent of nominal rigidity can be made. E.g., based on the estimated rigidity coefficient of approximately 25%, almost 10% of all salaried employees are affected by non-enacted earnings cuts, if one assumes a median nominal wage growth rate of 2%.

Our analytical framework also guides the use of descriptive and visual evidence. Appropriate evaluations of such evidence turn out to be informal applications of the three formal approaches. For instance, the measures of skewness for the distributions in the earlier part of the sample with higher median are lower than in the later part of the sample with lower median, which amounts to an eyeballing version of the skewness-location approach.

Each of our approaches to the analysis of nominal rigidity, the skewness location approach, the histogram-location approach as well as our descriptive analysis and visual evidence arrive at the same result. Therefore our conclusion is that there is indeed substantial downward nominal rigidity in German earnings which is dampened in times when unemployment is rising, and stronger in times of decreasing unemployment. We could not directly analyze the extent of nominal wage rigidity since only earnings are reported in the IABS. However, we can be confident that the earnings rigidity we found is caused by even more pronounced downward rigidity of wages, since variation in hours will tend to hide some of the effects of wage rigidity in the distribution of earnings changes.

Our results are of some importance for the monetary policy of the European Central Bank since they indicate that real wage adjustments are indeed hampered with very low rates of inflation and nominal wage growth. Since the ECB for the immediate future is bound to its announced inflation target of „below 2 percent“, it should at least allow inflation to reach that level and not aim at a zero inflation rate.
Appendix A: Data Selection

Whereas the high reliability of the earnings data is a great advantage of the IABS, there are also some drawbacks of the dataset. First, except for the information about part-time or full-time status there is no information about hours worked. It is therefore not possible to compute hourly wages. There is also the problem that changes of part-time to full-time work or vice versa do not lead to a new report of the employer. Hence there is the possibility that a temporary change to part-time work during the course of the year is not documented in the annual report. According to the IAB this leads to some implausible changes in annual earnings for female employees, for whom part-time work is much more common. For that reason we restrict our analysis to full-time employed males working in the western part of Germany.28

Second, earnings are censored to the right due to the contribution assessment ceiling (‘Beitragsbemessungsgrenze’). Earnings are only reported up to this threshold. Therefore, if (monthly) earnings are higher, actual earnings will be unknown.29 Since for employees whose earnings are censored the growth rate of earnings can not be computed correctly, the censored records are removed from the dataset. However, this leads to a substantial change in the skill structure of our sample. Since high-skilled employees are no longer properly represented in the sample, they are completely removed from the dataset. Our analysis is therefore confined to unskilled and skilled male employees. The unskilled are defined as persons without vocational training. These are persons with a lower schooling level and no further occupational qualifications completed; this group includes lower secondary school (Hauptschule) and intermediate secondary school (Realschule) graduates who did not complete an apprenticeship or graduate from a full-time vocational school. The skilled are defined as persons with vocational training. These are persons with an occupational qualification, which might be either a completed apprenticeship or graduation from a vocational school.

Third, earnings as reported in the IABS include fringe benefits that cannot be separated from ‘regular’ earnings. This is of importance since, starting in 1984, one-time payments to the employees have been subject to social security taxation and are therefore included in the earnings report of the employer. Before that date the inclusion of fringe benefits was voluntary. Steiner and Wagner (1996) analyzing the first version of the IABS note that this results in a structural break in the earnings data in 1984 which mainly affects the upper part of the earnings distribution. For that reason Steiner and Wagner restrict their analysis to a subsample starting only in 1984.30 The possibility of a structural break also has to be taken into account in our analysis. A structural break of this type leads to a level effect that affects the 1983-84 log earnings changes. Observations before that date should be valid, since if some employers reported fringe

28 We also excluded apprentices and employees working at home. Furthermore we requested that the information with respect to marital status, number of children, occupational status, training degree, occupation, nationality, industry affiliation and identification number of the respective firm are not missing. We examined whether the exclusion of missing values has led to a biased employment structure but could not find evidence for this hypothesis. We also excluded records with overlapping employment spells, because there is no information about the number of hours worked in each of these jobs.

29 Earnings are also truncated from below due to the lower social security threshold. In our context this point is of no problem since we restrict the analysis to full-time employees who exceed this limit almost with probability one.

30 Fitzenberger (1999) follows an alternative strategy by correcting for the spurious wage growth caused by the structural break.
benefits before 1984 and others did not, it is very likely that employers usually were consistently using a single kind of reporting behavior. All observations after that date derive from correctly reported figures. If the additionally reported compensation components grow with the same rate as the basic pay, the observations on changes after that date should therefore not be affected either. We deal with the problem of a structural break as outlined in section 3: in one variant of the analysis we consider the whole sample and explicitly test for a structural break in 1984, in another variant we exclude all observations before 1984 and consider only the shorter time period from 1984-1995.

A forth problem that is particularly relevant in the context of our analysis is caused by the fact that in the IABS the gross daily earnings variable has been truncated to integer values. This leads to massive changes in the distribution of close to zero earnings changes and could lead to wrong conclusions about the existence of menu costs.

To construct a subsample of wage changes, that is as homogenous as possible, we requested that the training degree, the profession, the occupational status, the marital status and the number of children remain unchanged in two consecutive years. We also restricted the sample to the manufacturing and service sector. For this sample we examined the reason for some implausible high growth rates of (annual) earnings lying between 100 and 400 percent. We found that these unusual growth rates are concentrated in the group of employees being younger than 25 years. We suspect that this is due to a false coding in the variable describing the occupational status. In this variable both the apprenticeship and the full-time/part-time distinction are coded. It is therefore possible that an apprentice is incorrectly classified as a full-time worker. After the apprenticeship the respective person will probably earn more than double the previous income. To guarantee that these effects are not at work in our data, we restrict the sample to employees, who are at least 25 years old. Due to our data selection we are left with 608,965 observations of year-to-year log earnings changes for the full sample period 1975-1995. The number of observations for the subperiods 1975-83 and 1984-95 as well as the number of observations for workers and salaried employees are reported in TABLE A.1.

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31 In the first version of the IABS the decimals had been rounded.
32 Hence employees working in the following sectors were excluded: agriculture, mining, energy, the government sector and private organizations. Furthermore we excluded the employees from the eastern part of Germany who were working in the western part of Germany and the employees who were insured by the so-called „Knappschaft“.
33 The upper limit of 65 years is implied by the IABS.
TABLE A.1: Frequencies of Log Earnings Changes in the Sample for Different Periods by Category

<table>
<thead>
<tr>
<th>Period:</th>
<th>All</th>
<th>Workers</th>
<th>Salaried Employees</th>
</tr>
</thead>
<tbody>
<tr>
<td>1975-95(1)</td>
<td>Frequencies</td>
<td>608,965</td>
<td>487,507</td>
</tr>
<tr>
<td></td>
<td>Row percent</td>
<td>100</td>
<td>80.1</td>
</tr>
<tr>
<td>1975-83(1)</td>
<td>Frequencies</td>
<td>228,470</td>
<td>177,980</td>
</tr>
<tr>
<td></td>
<td>Row percent</td>
<td>100</td>
<td>77.9</td>
</tr>
<tr>
<td>1984-95(1)</td>
<td>Frequencies</td>
<td>346,114</td>
<td>282,167</td>
</tr>
<tr>
<td></td>
<td>Row percent</td>
<td>100</td>
<td>81.5</td>
</tr>
</tbody>
</table>

Note: (1) The subperiods do not contain the changes in log earnings from 1983 to 1984, due to the possibility of a structural break. As a consequence the frequencies in the subperiods do not add to the number of observations for the whole sample period.

In our analysis we also use several aggregate time series. Our inflation variable is based on the consumer price index for all private households. The unemployment rate refers to the total labor force. These series and those used for instrumentation of the change in the rate of unemployment in section 5 are taken from various issues of the annual report of the Sachverständigenrat.

Appendix B: Skewness-Location Approach and Inflation Regressions

TABLE A.2: Skewness-Location Approach: Estimated Location Coefficients for the Regressions on Inflation

<table>
<thead>
<tr>
<th>Explained Skewness measure(1)</th>
<th>1975-1995</th>
<th>1984-1995</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Workers</td>
<td>Salaried Employees</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Skewness Coefficient</td>
<td>-0.07**</td>
<td>-0.07**</td>
</tr>
<tr>
<td></td>
<td>(-2.61)</td>
<td>(-2.27)</td>
</tr>
<tr>
<td>Median-C. Skewness Coefficient</td>
<td>-0.10**</td>
<td>-0.09**</td>
</tr>
<tr>
<td></td>
<td>(-2.38)</td>
<td>(-2.04)</td>
</tr>
<tr>
<td>Mean-Median Difference</td>
<td>-0.07*</td>
<td>-0.04</td>
</tr>
<tr>
<td></td>
<td>(-1.87)</td>
<td>(-1.41)</td>
</tr>
<tr>
<td>Sign Test Statistic</td>
<td>-2.61**</td>
<td>-0.69</td>
</tr>
<tr>
<td></td>
<td>(-2.22)</td>
<td>(-1.55)</td>
</tr>
<tr>
<td>Thinness Measure</td>
<td>-0.87**</td>
<td>-1.05**</td>
</tr>
<tr>
<td></td>
<td>(-2.40)</td>
<td>(-2.72)</td>
</tr>
</tbody>
</table>

Note: The independent variables in all estimations are a constant and the inflation rate (which is based on the consumer price index). For workers a dummy for 1979 was included (see section 4 for an explanation). The table contains the regression coefficient of the inflation rate and the respective t-value in parenthesis. * and ** denote significance on the 10 and 5 percent level, respectively.

The Durbin-Watson test statistic varies between 1.22 and 2.13. Hence in some estimations there is evidence for autocorrelation in the residuals. The $R^2$ varies between 0.10 and 0.46 in the estimations for the full sample period. The explanatory power of some regressions is therefore quite low. For the shorter time period it is still lower. (1) Measures of skewness were computed for observations in a band of +/-35 percentage points around the median.
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