Institut für Arbeitsmarkt- und Berufsforschung

Die Forschungseinrichtung der Bundesagentur für Arbeit

IAB-Discussion Paper 11/2008

Beiträge zum wissenschaftlichen Dialog aus dem Institut für Arbeitsmarkt- und Berufsforschung

Wage Cyclicality and the Wage Curve under the Microscope

Johannes Ludsteck

Wage Cyclicality and the Wage Curve under the Microscope*

preliminary version – do not cite without permission

Johannes Ludsteck†

Revision : 1.2

July 31, 2008

Abstract

Using large data sets from the German employment and unemployment register 1985-2004, we investigate aggregate wage cyclicality and the wage curve for establishment stayers and movers. We find that movers' wage responses to aggregate unemployment rate changes exceed these of stayers by about 30-40 percent. A new finding is that the increments of movers over stayer responses to regional unemployment shocks are considerably greater and amount to about 150 percnet. This difference in differences (responses to regional compared with aggregate cycles and responses of movers compared with stayers) can be explained by the importance of centralized wage bargaining in Germany.

JEL codes: J31, J41, E32

Keywords: wage setting, hiring standards, wage rigidity, wage curve, matching, censoring,

*I would like to thank Ekkehart Schlicht, Uwe Blien, Stefan Bender, Peter Jacobebbinghaus, and Katja Wolf for helpful conversations and comments. Of course, all remaining errors are my own.

†Institute for Employment Research (IAB) Nuremberg, Regensburger Str. 104, 90478 Nuremberg, Email: johannes.ludsteck@iab.de

1 Introduction

Cyclicality of wages is an important driving force for the adjustment of economies to macroeconomic shocks and has therefore always been a central issue of macroeconomic theory. It has been studied extensively at the aggregate level (see Abraham & Haltiwanger, 1995) and at the regional level in the related wage curve literature (see e.g. Blanchflower and Oswald (2005)). Current empirical literature has shifted the focus to isolating the sources of wage adjustment (or rigidity) in more detail. This is achieved by estimating wage regressions separately for particularly interesting subsamples (gender, education, age groups ...). The distinction between establishment stayers and movers has proven to be very important in two respects. It delivers sizeable and significant differences between these groups, and sheds light on the theory of the recruitment and wage policy of firms.

Devereux (2001) estimates standard wage regressions (explaining wage changes by aggregate unemployment changes and controls) for job stayers in the US and finds only weak evidence for wage cyclicality. Devereux and Hart (2006) compare reponses of job-stayers, within-company movers and between-company movers for UK data. The responses are significant for all groups but more pronounced for between-company movers (which exceed the stayer coefficients by about 30-50percent). Aggregate level evidence for Germany is unclear. Anger (2007) is very accurate to distinguish between hourly paid and salaried workers and exerts some effort to account bonus payments and overtime markup in the computation of hourly wages. Whereas monthly earnings respond significantly to the business cycle for several subsamples, she finds a significant impact on *hourly* wages only for the (rather small[1]) group of salaried workers with unpaid overtime. The paper by Siebert and Fei (2007) compares real wage cyclicality of movers and stayers for Germany and the UK and finds significant responses for stayers. This is somewhat irritating since the analysis related to German workers is based on the same data set (GSOEP) in both studies.

We provide aggregate level evidence on Germany using a different data set and complement it with regional level estimates (using district-specific unemployment rates). In all specifications we compare wage responses of several subgroups: establishment stayers[2], establishment movers, job-to-job

[1] The size of salaried workers with unpaid overtime is 3941 compared to 37999 observations in the full sample.

[2] Establishment stayers are defined as persons remaining in the same establishment between cutoff dates of successive years. Correspondingly, movers change the establishment

establishment movers, and involuntary establishment movers. Our aggregate level results are similar to the existing literature. A new and noteworthy aspect of the regional level analysis is that the increments of mover over stayer responses to the business cycle are much more pronounced at the regional level. Mover responses to aggregate cyclical shocks exceed stayer responses by 30-40 percent. The corresponding rates for regional shocks are about 150 percent. A self-suggesting interpretation is based on centralized collective wage bargaining ('Flächentarifvertrag') which is conducted at the industry and regional level in Germany. Though there is also a regional dimension in the bargaining process, the relevant regions are fairly large in most cases and agreements negotiated in important regions are customarily extended to other regions without further dispute. Centralized wage agreements are well suited for wage adjustment to aggregate business shocks (if conducted at a sufficiently high frequency). They imply, however, significant restrictions for adjustment to regional shocks. If firms want to deviate from centralized agreements, this can best be done for new hires, since job changes coincide always with changes of the working environment and frequently with a change of performed tasks. This creates scope to adjust wages by putting workers into better or lower paid slots of the firm-specific remuneration scheme. Downard wage adjustment for stayers is more difficult since continuity of tasks and the working environment provokes opposition and discontent of the workers. On the other hand, firms in regions affected by *positive* regional business shocks may wish to confine wage increases as an recruitment device to new hires while keeping incumbents' wages constant. (From the theoretical point of view, the wage required to attract new hires deviates from the wage required to retain incumbent workers.) This is feasible if regional business cycles increase worker fluctuation less than aggregate ones. Empirical studies based on personnel data (starting with Baker, Gibbs, & Holmstrom, 1994) show that such strategies are relevant in practice.

Though the size of our data set allows precise estimates for 'special' samples as involuntary establishment movers at the regional (district) level, it suffers from one – possibly crucial – drawback. It contains – for every worker spell – reliable and precise information on the number of days worked and gross earnings related to this spell, allowing to compute daily wages. But information on worked hours per day or overtime hours is missing. Therefore our estimates of wage responses to the business cycle may be biased and cannot be interpreted as hourly wage changes if overtime hours are correlated strongly with the business cycle. In the extreme case, our estimates may reflect pure overtime hours adjustment at constant hourly wages. This

between successive years. See the data section for the details.

problem is discussed in the wage curve literature (see. e.g. Black & FitzRoy, 2000; Card, 1995; Anger, 2007) and may be present in all wage curve studies based on the German register data (e.g. Baltagi & Blien, 1998; Bellmann & Blien, 2001; Baltagi, Blien, & Wolf, 2007). Our estimates, delivering large differences in responses between stayer and mover wages to the business cycle suggest that the a good deal of the estimated daily wage responses translates into hourly wage effects – at least under the required but plausible assumption that overtime hours movements are similar for stayers and movers.

The plan of the paper is as follows. In the next section we provide a short description of our data sets and data problems, and give detailed information on data processing steps (sample selection etc). This is followed by a section laying out the econometric model. Then we report our results and discuss them in a concluding section. Descriptive statistics and further results (from robustness checks) are moved to an appendix.

2 Data and Definition of Estimation Samples

Data Sources
Our data sets are drawn from the employment and unemployment register of the German Federal Employment Office (Bundesagentur für Arbeit and relate to the period 1985-2004.[3] The register covers nearly 80 percent of the German workforce, excluding only the self-employed, civil servants, individuals in (compulsory) military services, and individuals in so-called 'marginal jobs' (jobs with at most 15 hours per week or temporary jobs that last no longer than 6 weeks or jobs with monthly earnings below the social security threshold). It comprises important personal characteristics (sex, age, number of days worked, number of days in registered unemployment, education, working time categories) and identificators for establishments and districts. An important advantage of the data is that mis-reporting of earnings is subject to severe penalties, making them highly reliable. In principle, the register contains complete worker histories coded as spells. To simplify processing, we consider only spells at reference data of 30th June of each year.

Needless to say that our data suffer from several problems as most other real data sets.

[3]The register data currently cover the period 1975-2004. We drop the years 1975-1984 because of a structural break in 1984 related to earnings reporting. (Bonus payments were not reported before 1985.)

Censoring and Missing Information
About 10 percent of all workers are top-coded at the social security threshold on average (the limit changes slightly between years). Censoring rates exceed even 50 percent for high-skilled men (technical college or college). To check whether this causes significant bias in our panel data analysis, we implemented the consistent Honoré (1992) fixed effects GMM estimator and run it for the stayer sample. Since differences between the Honoré estimates and conventional OLS turned out to be neglectible, we proceeded with the computationally less demanding OLS.[4]

Remuneration and working hours
The most crucial problem of our data set (mentioned already in the introduction) concerns imprecise hours information and requires special treatment. Working time is reported only in three classes: full time, part time with at least 50 percent of full time working hours, and part time with less than 50 percent. Since standard working hours of the full-time employed vary between 35 and 40 hours (variation depends mainly on industry and time period) whereas working hours of part-timers are almost arbitrary, we restrict our sample to full-time employed West-German workers.

Identification of establishment changes
Missing hours information generates a further – possibly severe – problem for the interpretation of wage curves. Responses of daily wages to unemployment may reflect mainly overtime hours changes. However, as detailed in the introduction, differential responses between job stayers and job movers can be identified as true wage rate effects if responses of overtime hours are supposed to be similar for both groups. To this aim, we split our sample into establishment stayers and movers. Stayers are defined as employees remaining in the same establishment between reference dates (30th of June) of two successive years. Correspondingly, movers change establishments between reference dates of two succesive years. Further distinction between voluntary and involuntary movers is based on the duration of (registered) associated with an establishment change. We declare movers staying longer than 31 days in unemployment as involuntary movers. The required information is obtained from the German unemployment register (LEH).

To obtain stayer and mover samples of similar size, we keep all movers from the register but draw a random 10 percent sample of all stayers.[5] Since

[4]The implementation (*Mathematica* Code) and the results are available on request from the author.

[5]Note that the mover and stayer samples differ somewhat with respect to important

the resulting sample is still larger than required, we draw again a (random) 10 percent sample.

3 Empirical Model

Previous literature (cf. Solon, Barsky, & Parker, 1994; Abraham & Haltiwanger, 1995) has identified composition bias as the main obstacle to consistent estimation of wage cyclicality. It arises since unskilled workers are more vulnerable to layoffs in downswings than the rest of the work force. This thins out the the lower part of the wage distribution during downswings and induces (ceteris paribus) *countercyclical* behavior of aggregate wages. The problem cannot be tackled simply by controlling for qualification since it remains present even within narrow qualification groups where again low wage workers bear higher risks of becoming unemployed in downswings.

A simple, powerful and intuitive approach to eliminate composition bias and individual heterogeneity is to estimate wage equations in first differences.[6] Basically, our model for regional unemployment has the form

$$\Delta w_{i,r,t} = \Delta u_{r,t}\,\beta + x_{i,r,t}\,\gamma + \delta_t + \theta_{q,q'} + \Delta\epsilon_{i,r,t}$$

where Δ denotes the difference operator $\Delta y_{i,r,t} \equiv y_{i,r,t} - y_{i,r,t-1}$, $w_{i,r,t}$ denotes the log wage of individual i in district r and year t, $u_{r,t}$ the (log) unemployment rate, δ_t is a fixed year effect and $\theta_{q,q'}$ a dummy for migration between region type q and q'.[7] $x_{i,r,t}$ contains second order polynomial of the change of establishment size (which turns out to be an important control for movers)

characteristics (see appendix table 4). Movers are younger, work in smaller establishments and receive lower wages on average. In principle, the size of our data set would allow to implement a matching-like approach (i.e. use stratified sampling to obtain a stayer sample with similar age structure). This seems to be, however, not necessary since relevant characteristics are included as control variables in our model. Furthermore, stayers are not only relevant as a control group but interesting *per se.*

[6] The treatment of composition bias by differencing is not perfect since it does not account for selection into work: We observe wage changes only for workers employed at *both* cutoff dates of two successive years. In principle, explicit modelling of the selection problem is possible but requires exclusion restrictions (variables affecting employment but *not* wages) which are not fullfilled in most applications. Exclusion restrictions proposed in the literature (e.g. presence of a wife as in Keane, Moffit, & Runkle, 1988) appear to be rather weak and are not available in our data.

[7] To obtain the migration dummies, all 326 West-German districts are grouped into 9 region types. Dummy $\theta_{q,q'}$ takes on value 1 if person i migrates from region type q to q' in period $[t-1,t]$ and 0 otherwise. All combinations of sources and destinations amount

and a second order polynomal of age (in levels). The corresponding specification for the aggregate unemployment rate if obtained by dropping the region index r at some positions. We repeat it here for convenience.[8]

$$\Delta w_{i,r,t} = \Delta u_t \, \beta + x_{i,t} \, \gamma + \delta_t + \theta_{q,q'} + \Delta \epsilon_{i,r,t}$$

A minor technical issue in our model comes from the fact that the unemployment rate is aggregated (i.e. is the same for all workers in a year in the aggregate specifications and the same for all workers in a district in one year in the regional one). To obtain consistent standard errors the covariance matrix of the coefficients is clustered by year cells for the aggregate and district $\times$ year cells for the regional level estimates.

The specification above is the final one for the sample of job stayers. Additional heterogeneity in the sample of job movers, however, requires further differentiation. An stablishment change may be associated with intermediate unemployment spells which may affect wage prospects. Therefore movers are grouped further by duration of intermediate unemployment. They are put into the job-to-job group if length of intermediate unemployment is less than or equal to 31 days and in the complementary group (labelled 'involuntary job movers') otherwise.[9] Additionally, all regressions for samples including involuntary movers include dummies for the duration of intermediate unemployment (corresponding to the intervals 31-60, 61-90, 91-120, 121-150, 151-180, >180 days).

A last issue discussed in the literature concerns the dynamic specification of our model. First differences are standard in the wage cyclicality literature but incompatible with the Phillips curve and differ somewhat from the standard wage curve. The recent wage curve literature has inspected the dynamics more thoroughly using autoregressive models of the form

$$w_{i,r,t} = w_{i,r,t-1}\alpha + u_t \, \beta + \text{controls} + \epsilon_{i,r,t}$$

to $9 \times 9 = 72$ dummies. The region types are constructed by the German Bundesamt für Bauwesen und Raumordnung (BBR) mainly as agglomeration measures (ranging from rural regions to metropolitan areas and cities) and are therefore well suited to capture regional wage and price level effects.

[8] Note that we keep the region type migration dummies. Of course, the terms $\theta_{q,q'}$ with $q \neq q'$ are identically zero for the stayers samples.

[9] The limit was set to a value significantly above zero for two reasons. First, workers may intentionally become unemployed even if they have found a new job before being dismissed from the current one. Second, short unemployment spells may have a neglectible impact on hiring decisions of firms and acceptance decisons of workers. Then it is sensible to treat persons with short intermediate unemployment like job-to-job movers.

and found estimates of α between 0.3 and 0.5 for Germany (see Baltagi et al., 2007). Taken seriously, this implies that the constraint implied by the standard differences model (fixing α to unity) is wrong. We nevertheless retain it here for two reasons. First, to keep comparability with other papers from the wage cyclicality literature. And second (and more importantly), comparison of our estimates with the results in Baltagi et al. reveals that they appear to be good approximations. Note that the Phillips curve specification is rejected, however, unambiguously from the data. This is demonstrated in an appendix section.

4 Results

The model described above is estimated for several subsamples of stayers and movers separately. All results relate to West German full-time working prime age (20-60 years) men or women for the years 1985-2004. The regional analysis is based on 326 West German districts.

To start with, consider the right hand side of panel A in table 1 comprising responses of men to changes of the aggregate unemployment rate. Stayers' wages decrease by roughly one percent if the unemployment rate rises by one percentage point. Movers' coefficients denote incremental effects, compared to stayers. E.g. the value -0.408 (all movers) in column 'All Movers' of panel A implies a mover response of $-1.046 - 0.408 = -1.456$.[10] Movers exceed the stayers effects by almost 50 percent for the undifferentiated sample (blue + white collar workers). For blue collar workers, the incremental mover effect is somewhat lower and insignificant. For the white collar workers it is considerably stronger (about 75 percent). Note however, that all increments are at most marginally significant. If the movers sample is split further into job-to-job and involuntary movers, the incremental effect of involuntary movers almost doubles the job-to-job movers effects and becomes significant for the pooled and the white collar sample.

Effects for female are somewhat smaller in absolute size but similar with the only exception that involuntary movers do not deviate from the stayers sample. As mentioned above, our results may be biased due to missing hours and overtime hours information in our data. To assess this bias we compare them with two studies, Anger (2007) and Siebert and Fei (2007).

[10]Computation of the incremental effects is performed by pooling mover and stayer samples and interacting *all* regressors with an indicator (dummy variable) for movers.

Table 1: Response of log wage changes to changes of the West German unemployment rate. Effects for stayers and *inremental* effects for movers.

	Men				Women			
	Stayers	Incremental Effect of Movers			Stayers	Incremental Effect of Movers		
		All	JTJ	Invol		All	JTJ	Invol
				Panel A: All				
β_u	-1.058	-0.421	-0.381	-0.566	-0.807	-0.320	-0.343	-0.029
se	0.276	0.181	0.180	0.254	0.281	0.228	0.223	0.361
SUC	0.852	0.861	0.843	0.970	0.970	0.966	0.963	0.988
				Panel B: Blue Collar				
β_u	-1.147	-0.350	-0.326	-0.422	-0.868	-0.285	-0.333	-0.012
se	0.281	0.193	0.194	0.276	0.313	0.214	0.224	0.262
SUC	0.978	0.989	0.987	0.997	0.999	0.999	0.999	1.000
				Panel C: White Collar				
β_u	-0.826	-0.487	-0.460	-0.744	-0.782	-0.324	-0.344	-0.042
se	0.276	0.187	0.181	0.382	0.275	0.244	0.231	0.463
SUC	0.646	0.638	0.627	0.823	0.959	0.954	0.952	0.981

Standard errors (se) are computed by clustering the covariance matrix by year.

SUC: share of uncensored observations. JTJ: job-to-job movers. Invol: involuntary movers (associated duration of unemployment greater than 31 days)

Observation numbers are not reported to save space. They are above 30 000 for all subsamples and can be read from the corresponding blocks in the tables below.

Both studies employ the GSOEP, a large representative household survey containing detailed information on working hours and compensation schemes (salaries, hourly paid workers).

Anger (2007) computes several different earnings and wage measures (monthly earnings with and without overtime and extra payments, standard hourly wages, average hourly wages, and effective wages) and conducts separate regressions for several subsamples (salaried and hourly paid workers, all sectors, private sector workers only). Anger reports only estimates on pooled samples (men and women). For the following comparison with Anger's study, we have computed results from our data using the pooled (men + women) sample. These results are, however, not reported in the tables as they deliver less additional information and fit not well in the table structure. Her coefficient estimate for *monthly* earnings (including overtime and extra payments) of all stayers (men and women) is -0.450 with standard error 0.394. This is considerably lower than our estimates for a comparable sample (men and women, not reported in tables) which amount to -1.011 with standar error 0.278. Her corresponding estimate for *hourly* paid workers is -1.158 with standard error 0.492. This should be roughly comparable to our blue collar estimates amounting to about -1.148 with standard error 0.248. Her salaried workers results are, again, much lower in absolute size than our white collar results She finds a highly insignificant coefficient of -0.174 for monthly earnigs of all salaried workers, our corresponding value is 0.825 (with standard erro 0.276).

An important result of Anger's study is that responses shrink considerably and become insignificant if hourly wages are considered instead of monthly earnings. Her full sample estimate shrinks from -0.450 to -0.265 (with standard error) 0.473 if monthly earnings are replaced by effective[11] wages. The self-suggesting conclusion is that cyclical variation in overtime hours explains a large share of significant observed earnings responses. Taken seriously, this implies that the cyclical component of hourly wages is overestimated by about 100 percent if it is proxied with daily wages. On the other hand, Anger's estimates are considerably smaller than ours also for a comparable wage measure (monthly earnings) and quite imprecise. Therefore it is unclear what to learn from this comparison.

A comparison with the paper Siebert and Fei (2007) changes the big picture significantly. The authors apply a broad wage definition based on standard wages, overtime wages and bonus payments and run separate regressions by the public sector[12] and the private sector where they additionally distinguish between four firm size groups ($1-20, 20-200, 200-2000, > 2000$). Whereas the coefficient ist small (-0.54) and insignificant in the public sector, coefficients for the private sector are significant and similar for the four firm size groups: -1.53 , -1.40, -1.83, and -0.97. Taking into account that the public sector is small compared to the private sector, the weighted average over their samples (sectors and firm size) should be well above -1 in size. Unfortunately, we cannot figure out the reason for the differences with Anger's study, but conclude that our results roughly 'average' the survey data studies.

[11] Effective wages are calculated by averaging earnings over all (standard, paid overtime and unpaid overtime) working hours.

[12] We have run separate regressions for the public and private sector at early stages of our investigation. The results are similar to the evidence known from the literature the but not reported since they contribute nothing to the main aspects and arguments of the paper.

Table 2: Response of log wage changes to changes of national and regional West German **log** unemployment rates, **men**. Effects for stayers and *incremental* effects for movers.

Men

	National Unemployment		Regional Unemployment			
		Increm. Effect of		Incremental Effect of Movers		
	Stayers	Movers	Stayers	All	JTJ	Invol
			Panel A: All			
β_u	-0.076	-0.027	-0.012	-0.017	-0.016	-0.021
se	0.018	0.013	0.002	0.003	0.003	0.004
SUC	0.852	0.861	0.852	0.861	0.843	0.970
obs	1487650	1622345	1487650	1622345	1365602	256743
			Panel B: Blue Collar			
β_u	-0.083	-0.022	-0.014	-0.018	-0.018	-0.018
se	0.018	0.014	0.002	0.003	0.004	0.004
SUC	0.978	0.989	0.978	0.989	0.987	0.997
obs	1050383	1082383	1050383	1082383	876709	205674
			Panel C: White Collar			
β_u	-0.058	-0.033	-0.007	-0.017	-0.016	-0.023
se	0.019	0.014	0.002	0.003	0.003	0.007
SUC	0.646	0.638	0.646	0.638	0.627	0.823
obs	430029	432262	430029	432262	400207	32055

Standard errors (se) are computed by clustering the covariance matrix by year for aggregate level and by district for regional level regressions.
SUC: share of uncensored observations. JTJ: job-to-job movers. Invol: involuntary movers (associated duration of unemployment greater than 31 days)

The main and new contribution of our paper regards differences between incremental mover effects at the aggregate and regional level. They can be found in table 2. All results in this table relate to changes of *log* unemployment for two reasons. First, to make the results comparable with standard wage curve estimates. Second, results for aggregate and regional level estimates are directly comparable as the coefficients are elasticities. Comparison of the first two columns with the corresponding results in table 1 shows that taking logs essentially amounts to a rescaling of coefficients, leaving the relative differences between stayer and mover subsamples unaffected.

Note that the interpretation of regional level estimates is different from aggregate ones. Regional level regressions include a full set of time dummies, which filter out *all* aggregate shocks (including aggregate unemployment shocks). Thus the unemployment coefficients measure wage responses to *purely regional* business cycle movements. First consider panel A. Regional wage responses of stayers amount to about a fifth of the aggregate effects. This is a striking difference to U.S. data where regional and aggregate responses are roughly similar. A self-suggesting explanation is that centralized industry-wide collective wage agreements shrink local adjustment in Germany by restricting firm's possibilities to adjust to *local* fluctuations. Centralized bargaining seems also to be responsible for the large size of mover increments at the regional level. Mover responses for the full men sample are $-0.012 - 0.017 = -0.029$, i.e. exceed that stayer responses by almost 150 percent. For the women sample (see table 3) the difference is even greater (increment is -0.02, to be compared with a stayer effect of-0.011).
footnote The huge difference cannot be explained by bias through regional price differences since the regional level regressions include 72 dummies for migration between 9 region types. Furthermore incremental mover effects are similar if the regressions are run for a subsample of job movers remaining in the same district. The results (not reported in the tables) are available from the author on request. If the agreements apply to industries and large regional units as is the case in Germany (the so-called 'Flächentarifvertrag'), they allow adjustment to aggregate and industry-wide shocks *but not to purely regional ones.* As was laid out in the introduction, firms may want to treat new hires (movers) and incumbents (stayers) differently if purely regional shocks increase worker fluctuation less than regional ones.

Finally we note that – considering the large differences between stayer and mover effects – it seems difficult to attribute the responses mainly to overtime hours adjustment. This were plausible only if overtime hours of movers showed much higher correlation with the unemployment rate than that of stayers.

5 Conclusion

This paper compares wage adjustment in West Germany at the aggregate and regional level and extends the regional level analysis to establishment

Table 3: Response of log wage changes to changes of national and regional West German **log** unemployment rates, **women**. Effects for stayers and *incremental* effects for movers.

Women

	National Unemployment		Regional Unemployment			
		Increm. Effect of		Incremental Effect of Movers		
	Stayers	Movers	Stayers	All	JTJ	Invol
			Panel A: All			
β_u	-0.057	-0.020	-0.011	-0.020	-0.019	-0.028
se	0.020	0.016	0.001	0.003	0.003	0.005
SUC	0.970	0.966	0.970	0.966	0.963	0.988
obs	752641	789381	752641	789381	695464	93917
			Panel B: Blue Collar			
β_u	-0.062	-0.019	-0.014	-0.025	-0.024	-0.027
se	0.021	0.016	0.002	0.006	0.006	0.009
SUC	0.999	0.999	0.999	0.999	0.999	1.000
obs	202042	159415	202042	159415	132159	27256
			Panel C: White Collar			
β_u	-0.055	-0.020	-0.009	-0.019	-0.018	-0.027
se	0.019	0.017	0.002	0.003	0.003	0.005
SUC	0.959	0.954	0.959	0.954	0.952	0.981
obs	548307	574405	548307	574405	519856	54549

Standard errors (se) are computed by clustering the covariance matrix by year for aggregate level and by district for regional level regressions.
SUC: share of uncensored observations. JTJ: job-to-job movers. Invol: involuntary movers (associated duration of unemployment greater than 31 days)

movers. We find that regional wage adjustment amounts to about a fifth of the aggregate value. The relation between stayer and mover adjustment at the aggregate level is similar to estimates found for UK. The central and new finding of our paper is that the difference in cyclical wage adjustment between stayers and movers is much greater for regional unemployment shocks than for aggregate ones. The aggregate elasticity for male movers is -0.103 and exceeds the stayers value of -0.076 by about 30 percent. The corresponding regional (male) elasticity is -0.029 and exceeds the stayer value of -0.012 by almost 150 percent. The difference can be explained by the relevance of centralized collective wage bargaining in Germany. Whereas industry-wide agreements are well suited to adjust to aggregate cyclical movements, they put tight restrictions on adjustment to purely regional shocks. This urges firms to focus on movers (new hires) when they attempt to adjust wages. (These restrictions are, of course not binding for updward adjustment in booms, but may nevertheless have an extenuating effect on stayers' wages – this may be wellcome to the firm.)

Currently, a possibly important extension of the analysis is missing, however. Splitting the sample further in boom and slowdown periods would deliver a more detailed picture of the wage adjustment process within firms. Since a sensible allocation of time periods to boom and slowdwon phases appears to be tricky especially for the regional level analysis, this step is postponed to future versions of the paper.

References

Abraham, K., & Haltiwanger, J. (1995). Real Wages and the Business Cycle. *Journal of Economic Literature*, *33*, 1215–1264.

Anger, S. (2007). *The Cyclicality of Effective Wages within Employer-Emploee Matches – Evidence from the German Panel Data* (Discussion Paper No. 719). DIW Discussion.

Baker, G., Gibbs, M., & Holmstrom, B. (1994). The Internal Economics of the Firm: Evidence from Personnel Data. *Quarterly Journal of Economics*, *109(4)*, 881–919.

Baltagi, B., & Blien, U. (1998). The German Wage Curve: Evidence from the IAB Employment Sample. *Economics Letters*, *61*, 135–142.

Baltagi, B., Blien, U., & Wolf, K. (2007). *Phillips Curve or Wage Curve? Evidence from West Germany: 1980–2004* (Discussion Paper No. 14/2007). IAB Nuremberg and Syracuse University.

Bellmann, L., & Blien, U. (2001). Wage Curve Analysis of Establishment Data from Western Germany. *Industrial and Labor Relations Review*, *54*, 851–863.

Black, A., & FitzRoy, F. (2000). Earnings Curves and Wage Curves. *Scottish Journal of Political Economy*, *47(5)*, 471–486.

Blanchflower, D., & Oswald, A. (2005). *The Wage Curve Reloaded* (working paper). NBER.

Card, D. (1995). The Wage Curve: A Review. *Journal of Economic Literature*, *33*, 785–799.

Card, D., & Hyslop, D. (1997). Does Inflation Grease the Wheels of the Labor Market? In C. Romer & D. Romer (Eds.), *Reducing Inflation: Motivation and Strategy* (Vol. 30, pp. 71–121). National Bureau of Economics Research Studies in Business Cycle.

Devereux, P. (2001). The Cyclicality of Real Wages Within Employer-Employee Matches. *Industrial and Labor Relations Review*, *54(4)*, 835–850.

Devereux, P., & Hart, R. (2006). Real Wage Cyclicality of Job Stayers, Within-Company Job Movers and Between-Company Job Movers. *Industrial and Labor Relations Review*, *60(1)*, 105–119.

Honoré, B. (1992). Trimmed LAD and Least Squares Estimation of Truncated and Censored Regression Models with Fixed Effects. *Econometrica*, *60(3)*, 533–565.

Keane, M., Moffit, R., & Runkle, D. (1988). Real Wages over the Business Cycle: Estimating the Impact of Heterogeneity with Micro Data. *Journal of Political Economy*, *96*, 1232–1266.

Siebert, S., & Fei, P. (2007). *Real Wage Cyclicality in Germany and the UK: New Results Using Panel Data* (Discussion Paper No. 2688). IZA.

Solon, G., Barsky, R., & Parker, J. (1994). Measuring the Cyclicality of Real Wages: How Important is Composition Bias? *Quarterly Journal of Economics*, *109(1)*, 1–25.

A Appendix

A.1 Further information regarding the data sets and details of data processing

Data sources
All computations in this papers are based on extracts out of the German employment register (Beschäftigungsstatistik) of the Federal Agency of Labor, Bundesagentur für Arbeit). It contains all spells crossing cutoff dates 30th of June for the years 1985 to 2004. Employment information is complemented with (registered) unemployment duration data (extracted from the German unemployment register, Leistungsempfängerhistorik.)

The register data are generated from the reporting procedure of the compulsory German health, pension and unemployment insurance system. Employers have to submit a report for each employee (if she is subject to the social insurance system) to the social insurance office either at the end of every year or if a significant change of employment status takes place (e.g. part-time to full-time). The report comprises information on important demographic characteristics, the number of days worked (begin and end of employment spell), and the yearly average of gross earnings (including bonusses like holiday and christmas payments).

Definition of estimation samples
Our sample is restricted to full-time prime-age (20-60 years) West-German employees. According to our definition, 'West-German' means that an observation (difference of current and lagged wage) is dropped from the estimation sample if he worked in East Germany either in the current or previous year.[13] East-German wage spells are dropped since wages are considerably lower in East-Germany and human capital may be lower *ceteris paribus* due to usage of outdated equipment in the former communist economy. 'skilled' means all workers with *completed* apprenticeship training or technical school. Apprenticeship training is subject to tight standardization and regulation in Germany. Duration depends on the education level of apprentices (Hauptschulabschluss, Realschulabschluss or Abitur) but lasts at least two years and at most three years.[14] To avoid bias due to erroneous working hours information, workers with earnings lower than three times the lower social

[13]Note that this definition does not exclude workers with East-German employment spells in $t-2, t-3, \ldots$ or $t+1, t+2, \ldots$.

[14]This base training can be complemented with additional qualification courses for masters and technicians.

contribution threshold are dropped from the sample. This selection affects less than one percent of all workers in every year.

Identification of establishment movers

We split our sample into stayers and movers by checking whether workers remain in the same establishment between cutoff dates (30th of june) of successive years. Establishments are identified using a unique identificator ('Betriebsnummer') which is assigned by the German Federal Employment Office assigns ato every German establishment. This identificator is re-assigned, however, if the legal form of the firm changes, if it is sold to a new owner, or in case of spin-off or merger. Thus our strategy to identify establishment movers by change of the establishment ID may overrate the true number. Fortunately our argumentation is not weakened by this. We estimate our regressions separately for job stayers and movers to show that a good deal of wage responses to unemployment cannot be explained by hours adjustment. Large deviations in wage responses between stayers and movers (located at the same wage quantile) are in favour of our story. Putting job stayers falsely into the mover group shrinks the difference between estimated responses of stayers and movers (if response of movers is greater). Thus we are on the save side if the estimated differences are nevertheless significant.

A.2 Descriptive Statistics

Table 4: Descriptive statistics for important variables

		Men			Women		
		Stayers	Movers		Stayers	Movers	
Var	Stat.		JTJ	Invol		JTJ	Invol
W	$Q_{0.10}$	4.062	3.973	3.812	3.694	3.705	3.569
	$Q_{0.50}$	4.367	4.309	4.128	4.159	4.130	3.937
	$Q_{0.90}$	4.699	4.688	4.439	4.566	4.560	4.341
AGE	mean	39.610	35.436	33.894	37.498	32.452	31.710
	sd	10.389	9.770	9.653	10.966	9.865	9.407
	$Q_{0.10}$	26.000	24.000	23.000	24.000	22.000	22.000
	$Q_{0.50}$	39.000	34.000	32.000	36.000	29.000	29.000
	$Q_{0.90}$	54.000	50.000	49.000	53.000	48.000	46.000
ES	mean	1696	643	385	872	485	317
	sd	6033	2942	2691	3753	2148	1967
	$Q_{0.10}$	8	5	3	4	3	3
	$Q_{0.50}$	133	62	27	92	49	31
	$Q_{0.90}$	2892	1092	416	1482	910	501
DES	mean	-13	50	131	-2	7	74
	sd	474	3606	3134	355	2756	2410
	$Q_{0.10}$	-56	-491	-213	-25	-432	-253
	$Q_{0.50}$	0	-0	0	0	0	1
	$Q_{0.90}$	35	525	293	28	473	360
DUE	mean	-	-	111.186	-	-	110.321
	sd	-	-	62.907	-	-	64.953
	$Q_{0.10}$	-	-	43.000	-	-	43.000
	$Q_{0.50}$	-	-	95.000	-	-	92.000
	$Q_{0.90}$	-	-	201.000	-	-	206.000
Observations		1487650	1365237	256652	752641	695429	93897

Legend: W: log real wage, AGE: age, ES: establishment size, DES: first (time) difference of establishment size, DUE: mean duration of unemployment (movers only), Q_x : Quantile x.

A.3 A Simple Test of the Phillips Curve against the Wage Curve Specification

Our regression model is indentical to the one customarily used in the wage cyclicality literature and similar to the wage curve but inconsistent with the Phillips curve where wage (price) growth depends on the *level* of unemployment. Card and Hyslop (1997) propose to test that by replacing the change of (log) unemployment rate by the levels of current und lagged (log) unemployment.

$$\Delta w_{i,t} = \beta_0 \, u_t + \beta_{-1} \, u_{r,t-1} + \ldots$$

If the Phillips-curve specification were the correct one, β_{-1} should be insignificant. If the standard difference specification were correct, β_0 and β_{-1} should be equal in absolute value but show different signs, implying $\beta_0 + \beta_{-1} = 0$. The following table reports coefficients of current and lagged (log) unemployment on (log) wage changes. The specifications used here are otherwise identical to the ones in the main part.

As can be seen from the results listed in the following table, the Phillips-curve specification is rejected for movers and stayers (lagged unemployment coefficients deviate significantly from zero), whereas the standard difference specification is rejected at the five percent confidence level only for stayers (the sum $\beta_0 + \beta_{-1}$ differs significantly from zero). Nevertheless the difference appears to be a good approximation to the true model, as $\beta_0 + \beta_{-1}$ is small compared with the absolute size of β_0 and β_{-1}.

Table 5: Coefficients of current and lagged (log) unemployment rates on (log) wages

statistic	Stayers		Movers	
	aggregate	regional	aggregate	regional
		Men		
β_0	-1.151	-0.122	-1.780	-0.289
$se(\beta_0)$	0.302	0.034	0.319	0.040
β_{-1}	0.992	0.103	1.242	0.328
$se(\beta_{-1})$	0.274	0.034	0.296	0.028
obs	1487650	1487650	1621889	1621889
$\beta_0 + \beta_{-1}$	-0.160	-0.019	-0.538	0.039
$PV(\beta_0 + \beta_{-1})$	0.253	0.001	0.004	0.088
		Women		
β_0	-1.037	-0.091	-1.707	-0.306
$se(\beta_0)$	0.292	0.019	0.286	0.037
β_{-1}	0.645	0.100	0.718	0.340
$se(\beta_{-1})$	0.267	0.018	0.254	0.025
obs	752641	752641	789326	789326
$\beta_0 + \beta_{-1}$	-0.391	0.008	-0.989	0.034
$PV(\beta_0 + \beta_{-1})$	0.008	0.096	0.000	0.139

Legend: $se(\beta_i)$: standard error of β_i. $PV(\beta_0 + \beta_{-1})$ denotes the P-value of a test $H_0 : \beta_0 + \beta_{-1} = 0$ against $H_0 : \beta_0 + \beta_{-1} \neq 0$.

Recently published

No.	Author(s)	Title	Date
23/2007	Bender, S. Koch, S. Meßmann, S. Walwei, U.	Was muten sich Arbeitslose zu? : Lohnkonzessionen von ALG-II-Empfängern published in: Sozialer Fortschritt, Jg. 57, H. 3 (2008), S. 75-85	7/07
24/2007	Bruckmeier, K. Schnitzlein, D.	Was wurde aus den Arbeitslosenhilfeempfängern? : Eine empirische Analyse des Übergangs und Verbleibs von Arbeitslosenhilfeempfängern nach der Hartz-IV-Reform	9/07
25/2007	Büttner, T.	Ankündigungseffekt oder Maßnahmewirkung? : Eine Evaluation von Trainingsmaßnahmen zur Überprüfung der Verfügbarkeit	10/07
26/2007	Brücker, H. Defoort, C.	Inequality and the (self-)selection of international migrants: Theory and novel evidence	10/07
27/2007	Brücker, H. Schröder, P.J.H.	International migration with heterogeneous agents: Theory and evidence	10/07
28/2007	Krug, G.	In-work benefits for low wage jobs: Can additional income hinder labor market integration?	11/07
29/2007	Wolff, J. Jozwiak, E.	Does short-term training activate means-tested unemployment benefit recipients in Germany?	11/07
30/2007	König, M. Möller, J.	Mindestlohneffekte des Entsendegesetzes? : Eine Mikrodatenanalyse für die deutsche Bauwirtschaft	11/07
32/2007	Hohmeyer, K. Wolff, J.	A fistful of Euros: Does One-Euro-Job participation lead means-tested benefit recipients into regular jobs and out of unemployment benefit II receipt?	12/07
1/2008	Schnitzlein, D.	Verbunden über Generationen: Struktur und Ausmaß der intergenerationalen Einkommensmobilität in Deutschland	1/08
2/2008	Bauer, T. K. Schmucker, A. Vorell, M.	KMU und Arbeitsplatzdynamik: Eine Analyse auf Basis der Beschäftigten-Historik-Datei	1/08
3/2008	Baas, T. Brücker, H.	Macroeconomic consequences of migration diversion: A CGE simulation for Germany and the UK	1/08
4/2008	Amend, E. Herbst, P.	Labor market pooling and human capital investment decisions	1/08
5/2008	Bernhard, S. Wolff, J.	Contracting out placement services in Germany: Is assignment to private providers effective for needy job-seekers?	2/08
6/2008	Lehmer, F. Ludsteck, J.	The returns to job mobility and inter-regional migration	2/08
7/2008	Klinger, S. Wolf, K.	What explains changes in full-time and part-time employment in Western Germany? : A new method on an old question	2/08
8/2008	Hohmeyer, K. Jozwiak, E.	Who is targeted by One-Euro-Jobs? : A selectivity analysis	2/08
9/2008	Schnabel, C. Kohaut, S. Brixy, U.	Employment stability of entrants in newly founded firms: A matching approach using linked employer-employee data from Germany	2/08
10/2008	Konle-Seidl, R.	Changes in the governance of employment services in Germany since 2003	2/08

As per: 27.02.2008

For a full list, consult the IAB website
http://www.iab.de/de/publikationen/discussionpaper.aspx

Imprint

IAB-Discussion Paper 11/2008

Editorial address
Institute for Employment Research
of the Federal Employment Agency
Regensburger Str. 104
D-90478 Nuremberg

Editorial staff
Regina Stoll, Jutta Palm-Nowak

Technical completion
Jutta Sebald

Website
http://www.iab.de

Download of this Discussion Paper
http://doku.iab.de/discussionpapers/2008/dp1108.pdf

For further inquiries contact the author:

Johannes Ludsteck
Phone +49.911.179 4810
E-mail johannes.ludsteck@iab.de