Downward Real Wage Rigidity and Equal Treatment Wage Contracts: Evidence from Germany

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Abstract

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1 Introduction and Overview

There is now a sizeable literature devoted to estimating the cyclicality of real wages using large panel datasets to correct for composition bias (e.g., Solon et al. (1994) for the US, Martins et al. (2012) for Portugal, and Devereux and Hart (2006) for the UK). These papers with few exceptions focus on a single correlation between log of wages and the aggregate unemployment rate — henceforth "semi elasticity" for short. However, theoretical models of downward real wage rigidity (e.g., the real models of Hall (2005) and Snell and Thomas (2010) and a host of DSGE models with nominal rigidities) suggest that wages will be asymmetrically cyclical — rigid in "bad" times when unemployment is relatively high but upwardly flexible during "good" times when unemployment is relatively low. In this paper we use a administrative panel dataset from Germany to establish that such asymmetries are very salient in Germany. Two key empirical features emerge from our study. First, we find that the semi elasticity is close to zero and insignificant when unemployment is above its long term average but that the semi elasticity is large and highly significant when unemployment is below its long term average. Second, we find that the incremental wage cyclicality incurred by new hires and low tenured workers is wholly insignificant. This second result is particularly important. It provides strong support for the existence of equal treatment amongst workers within the firm — a central feature of many contracting models whereby workers in firms are not pay discriminated against according to the date they were hired (see, e.g., Hall, 2005; Gertler and Trigari, 2009; Snell and Thomas, 2010). Equal treatment implies that in recessions unemployed workers cannot price themselves into jobs by undercutting the wages of incumbents. Put another way, involuntary unemployment cannot be eliminated by firm level wage adjustments at the extensive (new hire) margin.

In the second part of the paper we show that a simple three parameter version of Snell and Thomas' (2010) equal treatment contracting model can generate a strong negative semi elasticity when unemployment is below its mean and a semi elasticity close to zero when it is above. In the model, risk neutral firms have an incentive to smooth the wage profiles of risk averse workers by offering a contract that limits the rate at which real wages fall in bad times. The rate of fall is the result of a trade-off for the firm between wanting to avail itself of cheaper new hires on the one hand and not wanting to create too much variability for the incumbents on the other. This results in a maximum rate of wage fall no matter how cheap new hires become. At the same time, the existence of an equal treatment constraint implies that the unemployed cannot price themselves into jobs by offering to undercut incumbent wages. As a result, if the economy suffers a negative and persistent productivity shock, downwardly sluggish wages will generate unemployment that will endure until either productivity recovers or until wages have fallen sufficiently far to clear the labour market. In this scenario, real wages will exhibit little covariation with unemployment when productivity is low. By contrast, wages adjust rapidly (upwards) in good times when productivity is high so that in upswings they will be strongly negatively correlated with unemployment. Workers are ex post mobile so that in good times their outside opportunity has a high value; firms have no option but to raise wages to match this and so wages are upwardly flexible during upswings.

The Snell and Thomas (2010) model uses exogenous productivity — TFP — as an input, but unfortunately TFP data are unavailable for Germany over the 1977–2009 time period of our study.[1] Instead of using real productivity data we feed the model synthetic data generated from a variety of realistic time series processes for TFP. The resulting wage and unemployment series display the

[1]The KLEMS project started publishing sectoral TFP only in 1991 but even using those observations has pitfalls: to solve our model we must start from a year after the peak of a business cycle but Germany suffered a long and protracted recession in the years after 1991.

same asymmetric response patterns we found in the data. In addition, in most of the TFP scenarios, real wages have broadly the same time series properties as those in the panel data. The model does less well in matching the properties of unemployment; it can match its volatility but cannot capture fully its persistence. Extending the model to incorporate frictions such as adjustment costs or job search, something that may have been very prominent after German unification, may be ways to rectify this shortcoming.

2 Related Empirical Literature

In the estimation exercise below we use "downswing" dummies interacted with the unemployment rate to allow and test for differential cyclicality in downswings. Using such interaction terms in this context is not new. In particular Martins (2007), defines upswings (downswings) in terms of falling (rising) unemployment rather than its level. Using this definition he examines the extent of differential real wage cyclicality in the two regimes. However he finds that the semi elasticity actually *increases* (in absolute value) when unemployment is rising.[2] The result that wages are relatively inflexible in times of falling unemployment but highly flexible when it is rising is hard to rationalise. Shin and Shin (2008) also estimate asymmetric wage responses to unemployment. However they focus on asymmetries across workers of different tenure at a point in time rather than asymmetries across different time periods for all workers. The asymmetries they find are consistent with what one would find in a Beaudry and DiNardo (1991) world of implicit contracts under worker mobility.[3] By contrast we find that all workers *regardless of tenure* are subject to the

[2] In addition he looks at each worker's wage growth rather than its level — a classic method to remove various fixed effects but also a limitation as it means that new hires must be omitted from the sample.

[3] In effect they find in favour of "minu" — the minimum unemployment rate during a worker's tenure — which Beaudry and DiNardo (1991) argue determines wages in this scenario and which has found to be significant in a host of other papers.

same wage cyclicality and that the effects are asymmetric across time rather than across tenure. In particular, when we allow a) new hires and b) new hires and low (between one and two years) tenure workers to have differential wage cyclicalities to other workers (incumbents) we find that such differentials are wholly insignificant — incumbents and new hires or workers of low tenure seem to get the same (asymmetric) "cyclical wage deal".

Our finding of equal treatment is not universally supported by other panel data studies. For example, Haefke et al. (2012) find new hire wages in the CPS to be more volatile than those of incumbents. Additionally a plethora of other papers have found evidence to support the existence of bilateral worker-firm implicit contracts (as opposed to equal treatment contracts) in wages (see, e.g., Devereux and Hart (2007), Grant (2003), and the original seminal paper by Beaudry and DiNardo (1991)). But the results of these papers have recently come under attack from several quarters. In particular Gertler and Trigari (2009) and Hagedorn and Manovskii (2013) show that failing to control rigorously for worker-firm match (and within firm job-match) quality[4] can lead to spurious tenure specific cyclicality of the kind consistent with implicit contracts. In this paper we control for worker-firm match quality within very tightly defined occupations. This may be the reason why we find in favour of equal treatment whilst others have not.

Finally, Abbritti and Fahr (2013) document asymmetries in the distribution of real wage growth for several OECD countries. They find that a DSGE model subjected to exogenous downward real wage adjustment costs can generate the skewness they find in the aggregate data. However — and as they themselves note — aggregate data on wage growth is afflicted with composition bias (Solon et al., 1994) something this paper is able to avoid. Additionally, DSGE models impose exogenous

[4]Nearly all of the work on US data uses unmatched panels. Here and in some EU studies also (e.g., Martins, 2007), the usual device to remove worker and firm fixed effects from the panel is to take the first difference of wages. Not only does this preclude analysis of the cyclicality of new hire wages it also assumes a priori that wages are nonstationary. Here we are able to test for the order of integration of the time component of wages.

wage adjustment costs and/or price stickiness at the outset in order to match data on wages and unemployment — in our model real wage rigidity (and involuntary unemployment) arises naturally as a consequence of equal treatment wage contracting.

3 Data and Estimation

3.1 Data

Our panel of firms is drawn from the IAB Beschäftigten-Historik (BeH), the Employee History File of the Institute for Employment Research (IAB) of the German Federal Employment Agency. This file consists of data on the total gainfully employed members of the German population who are covered by the social security system. Not covered are self-employed, family workers assisting in the operation of a family business, civil servants (Beamte) and regular students. The BeH covers roughly 80% of the German workforce. Plausibility checks performed by the social security institutions and the existence of legal sanctions for misreporting guarantee that the earnings data are very reliable.

The data used for the analysis are from 1977–2009 and as we wish to avoid workers in firms that have small time spans we confine our attention to workers in West German firms only. We model time invariant human capital using worker-firm-occupation fixed effects and add tenure, tenure squared, and age squared to pick up the effects of time varying experience.

The BeH is organized by employment spells. A spell is a continuous period of employment within a job within a firm in a particular calendar year. Hence the maximum spell length is 365 days. We collate these spells into annual observations on workers. Unfortunately the BeH only documents total spell earnings and not hours worked in that spell. We therefore only consider full time workers. Nearly all full time workers in Germany work a standard number of hours per week so the average daily spell wage should be very closely related to the spell hourly wage. In Section 3 we analyse the time series properties of two extraneous estimates of the average weekly hours worked by full time employees in Germany. We find cyclicality — in the sense of having a negative correlation with the unemployment rate — to be relatively weak. To calculate the daily real wage (in 2005 prices) we use the Consumer Price Index (CPI). Unfortunately wages are censored at a maximum level equal to the contribution assessment ceiling of the compulsory pension insurance scheme. Earnings spells with wages above or close to (within 4% of) the truncation point are dropped.

In order to test equal treatment we will require a reasonable number of new hires. Our sample consists of all firms who employed at least 10 full-time workers and who hired at least seven new workers per year in each year of the 1977 to 2009 period. This is a minority of the total firms in the BeH but it still includes over 95 million data points gleaned from over 14 million workers of 12,884 firms (see Table A1). Our sample selection criteria imply that the firms in our sample tend to be larger and the dispersion of firm sizes much lower than that of the German economy as a whole. It is important therefore to establish that our results are not sensitive to firm size and this we do in a sensitivity analysis later in the paper.

Since tenure is an important control variable we cannot use employment spells that have missing tenure - we must drop these spells. Unfortunately this means a worker only enters the data when he joins one of the firms in 1976 or in later years.[5] This implies a growing number of new hires

[5] For the analysis we only use the years 1977–2009, but for the identification of firm entrants and the calculation of firm-tenure we use BeH data from 1975 onwards. However, we exclude all spells from 1975 because the tenure could be left censored.

(relative to incumbents) as the years progress. Our estimation of new hire effects — whereby we allow for a completely separate semi elasticity for new hires — should not be affected by this. Again in a sensitivity analysis we show that our results are unaffected when we drop the early years of the sample.

We wish to annihilate wage movements associated with job ladders — i.e. workers getting wage increases/decreases because they switch firms (or jobs within firms) during booms/recessions. We therefore control for worker-firm-occupation match quality via a fixed effects estimator. Our dataset identifies 346 occupations so we would argue that our fixed effects cover a very wide range of specific tasks carried out by specific workers in specific firms.[6]

Aggregate unemployment was measured for West Germany. The implicit assumption here is that East German workers did not compete for jobs in West German firms after 1991 (they obviously did not do so prior to re-unification). In one of our sensitivity analyses we change the definition of unemployment to include all German workers after 1990. We find our results are robust to this.

We have defined upswings and downswings in terms of low (below average) and high (above average) unemployment. This is obviously not a unique and universally accepted definition of the business cycle and nor is it meant to be. But the relationship of wages to unemployment is a natural focus in the context of the economic theories of wage rigidity we are interested in here. The current unemployment rate in wage contract models (including those with search frictions) is a key determinant of the value of the outside option to an incumbent (who may consider quitting the firm) or to a new hire (who may be currently forging a contract with a firm). It is for this reason that the response of wages to unemployment rather than (say) to productivity or output has been

[6] In total we control for 19,211,898 worker-firm-occupation fixed effects. Additionally for workers who change jobs or occupations within a firm we record only their first spell at the firm and delete subsequent spells. Instances of this are extremely rare and dropping them removes only a tiny fraction of the total number of spells.

the prime focus of recent panel studies such as ours. We should add that there is no guarantee our finding of asymmetrical wage adjustment is robust to alternative definitions of upswings and downswings. But we do not investigate the possible existence of and nature of other stylised facts about the cyclicality of wages. We focus here only on the asymmetrical adjustment of wages to unemployment and try and establish that this feature is shared by Snell and Thomas' (2010) equal treatment wage contracting model.

3.2 Estimation

3.2.1 Main Estimates

In a seminal paper Solon et al. (1994) examine wage cyclicality by extracting year effects from a panel using year dummies in a first stage. In the second stage a time series regression of the year effects on unemployment rates is used to estimate common cyclical effects in the panel. This two-step procedure has become the industry standard in such applications. We adapt this procedure to allow for potentially different wage cyclicality of a) new hires b) workers with one year of tenure or less and c) workers of more than one but less than two years of tenure or "low tenured" workers for short. We control for worker-firm-occupation fixed effects (match effects) and the worker's age and tenure. In the first stage therefore we estimate the panel regression

$$ln(w_{ijkt}) = \varpi_{ijk} + \sum_{\tau=1}^{T} \beta_\tau^a \delta_t^\tau + \sum_{\tau=1}^{T} \beta_\tau^0 \delta_{ijt}^0 \delta_t^\tau + \sum_{\tau=1}^{T} \beta_\tau^1 \delta_{ijt}^1 \delta_t^\tau + \phi_2 age_{it}^2 + \lambda_1 ten_{ijt} + \lambda_2 ten_{ijt}^2 + v_{ijkt} \quad (1)$$

where w_{ijkt} is the real wage of worker i in occupation k at firm j during year t, ϖ_{ijk} is her worker-firm-occupation fixed effect and v_{ijkt} is an error term assumed orthogonal to the regressors. There are three summation terms. The first consists of the 33 year dummies δ_t^{τ} (with $\tau = 1, \ldots, 33$) with coefficients β_{τ}^{a}. These coefficients are the year effects that are common to all workers' wages. The second summation interacts year dummies with a "new hire" dummy δ_{ijt}^{0}. This dummy takes the value one if worker i in year t is in her first year of tenure but is zero otherwise. The coefficients on the interaction terms β_{τ}^{0} capture any incremental year effects to wages accruing to new hires. The third summation term interacts year dummies with a "low tenure" dummy δ_{ijt}^{1}. This takes the value one if the worker has worked for more than one but less than two years at the firm. The respective coefficients β_{τ}^{1} pick up incremental year effects accruing to workers of low tenure. The variable ten_{ijt} is worker $i's$ tenure at firm j in year t measured in days and age is her age in years.

In the second stage we treat each of the year effects — the β coefficients — as separate time series and analyse their respective relationships with aggregate unemployment. Following the discussion above we refer to periods where unemployment is below (above) its mean as "upswings" ("downswings"). We re-emphasise that this terminology is designed to make clear the nature of the stylised facts we seek rather than offer a universally acceptable definition of the business cycle. For our purposes then we wish to estimate the following time series relationships

$$\beta_t^k = q_t + \theta_k^u \delta_t^u \widetilde{u}_t + \theta_k^d (1 - \delta_t^u) \widetilde{u}_t + \varepsilon_t \quad \text{with } k = a, 0, 1 \quad \text{and } t = 1, 2, \ldots, 33 \tag{2}$$

where q_t is a quadratic trend, $\widetilde{u}_t$ is demeaned aggregate unemployment rate, and δ_t^u is an upswing dummy taking the value unity if $\widetilde{u}_t < 0$ and zero otherwise. As a result, θ_a^u (θ_a^d) measures the semi-elasticity in upswings (downswings) for all workers whilst θ_0^u (θ_0^d) and θ_1^u (θ_1^d) measure the increment to those semi elasticities for new hires (low tenured workers). The error term ε_t is potentially autocorrelated and heteroscedastic.

Before looking at the main estimates we ran a simple OLS regression of the common wage year effects (β_t^a) on the aggregate unemployment rate to see what the more traditional approach gave. Line (1.1) of Table 1 gives the results including the ADF test for the residual.[7] Our semi-elasticity estimate of -1.23 is of the same order of magnitude as those OLS estimates found elsewhere. In particular using the GSOEP survey data for German workers in employee-employer matches, Peng and Siebert (2007) find semi elasticities between -1 and -2 depending on firm size whilst Anger (2011) using the same dataset finds semi-elasticities between -0.96 to -1.22. For the US Devereux's (2001) finding of a semi elasticity for job stayers of around minus unity is typical.

We now turn to estimation of the asymmetric responses in (1). Estimates of θ_a^u and θ_a^d are given in line (1.2) of Table 1 below[8]. The estimation is subject to first order error autocorrelation[9] and (absolute value) t-ratios are computed from heteroscedasticity robust standard errors. As the upswing and downswing regressors are orthogonal we only need to focus on individual rather than joint significance tests.

The results show that unemployment is an important and highly significant determinant of wages but only in upswings (θ^u) whilst the augmented Dickey-Fuller (ADF) test shows that the residuals are resoundingly stationary. In upswings, the semi elasticity is -1.75. Lines (1.3) and (1.4) give, respectively, the incremental responses for new hires and low tenured workers. All terms in these regressions are wholly insignificant with no p-value lying below 0.6. In further regressions

[7] An ADF test with one augmented term (dropped if insignificant) was performed on OLS residuals from each respective regression here and below. We show later that unemployment is clearly stationary so the standard ADF procedure may be applied to these residuals. The 5% critical value (computed by simulation) was -3.88.

[8] Originally we also included the dummy δ_t^u itself in all specifications to check for upswing/downswing "fixed effects". These effects were quantitatively small and statistically insignificant in all our specifications and adding them made no difference to the main results we present here.

[9] This procedure "whitens" the errors and ensures correct inference. The resulting estimates are conditional semi elasticities. Their (unconditional) counterparts — available on request — derived using OLS were not greatly different.

(not reported here but available on request), the incremental year effects for new hires and low tenures are both found to be unrelated also to detrended GDP and GDP growth (p-values for the significance of the regressions are 0.45 and 0.21 respectively).[10]

In the literature it is more common to allow for incremental cyclicality for new hires *only* rather than for new hires and low tenures separately as we have done. To check that this is not affecting our results we re-estimated (1) with β_t^1 constrained to be zero ($\forall\ t$) in lines (1.5) and (1.6) of Table 1. We see that all of the results are virtually unchanged.

[10] In fact the new hire incremental year effects show no significant trend and are virtually white noise. The low tenure incremental effects also show no trend but do display significant first order autocorrelation ($\rho = .39$).

Table 1

Second Stage Estimates of Semi Elasticities

	LHS Vble.	θ^u	θ^d	*ADF*
(1.1)	β_t^a	-1.23		-3.88
		(5.10)		
(1.2)	β_t^a	-1.75	-0.27	-4.72
		(4.98)	(1.26)	
(1.3)	β_t^0	0.05	0.06	-5.15
		(.20)	(.44)	
(1.4)	β_t^1	0.06	0.04	-5.20
		(0.42)	(0.50)	
(1.5)	β_t^a	-1.82	-0.29	-4.26
		(-6.35)	(1.36)	
(1.6)	β_t^0	$-.03$	0.02	-6.07
		(0.18)	(0.20)	

Robust | t-ratios | in brackets

3.2.2 Robustness Checks

Sample Selection Issues

There are three main sample selection criteria that may influence our results. Firstly, in order to obtain good estimates of the incremental wage cyclicality of new hires and workers of low tenure we had to select firms who had a reasonable number (seven) of new hires in each year. Secondly, in order to obtain good estimates of match effects we required firms to exist for all years 1977–2008. Finally to obtain our crucial tenure variable we had to drop all incumbent workers in 1976 (the first pre-sample year of the BEH). The first two selection criteria select (on average) large firms whilst the third gives a rising profile of the proportion of incumbents in the sample as Table A1 shows. We do a robustness check for the effects of the latter in the next subsection. Here we focus on the consequences of over-sampling large firms by seeing how — if at all — our results are affected by firm size.

Explicitly, we split our sample into "large" and "small" firms according to the average number of workers per year of the firms. Firms with on average equal or fewer than 110.8 workers per year were entered into a new "small" panel dataset and those with on average over 110.8 workers per year into a "large" panel.[11] We then carry out separate analyses on the small and large panels treating them as two distinct panel datasets. Re-estimating (1) using the same method as above but now for small and large firms separately gives

[11] We identify 6443 small firms (average number of full-time workers per year $\leq$ median average number of full-time workers per year = 110.8182) and 6441 large firms.

Table 2

Second Stage Estimates of Semi Elasticities:-

The Effect of Firm Size

		Small Firms			Large Firms		
	LHS Vble.	θ^u	θ^d	ADF	θ^u	θ^d	ADF
(1.1)	β_t^a	-1.36		-3.38	-1.21		-3.94
		(5.13)			(4.94)		
(1.2)	β_t^a	-2.01	-0.25	-4.66	-1.70	-0.28	-4.66
		(5.95)	(1.00)		(4.77)	1.30	
(1.3)	β_t^0	-0.09	-0.08	-5.81	0.08	0.06	-5.02
		(0.48)	(0.76)		(0.31)	(0.43)	
(1.4)	β_t^1	-0.04	-0.01	-5.73	0.08	0.03	-5.15
		(0.26)	(0.14)		(0.52)	(0.40)	

Robust | t-ratios | in brackets

The results for large firms are virtually unchanged from those in Table 1 whilst the only noticeable difference in the small firm estimates is that they seem to display marginally higher upswing cyclicality. Crucially, the incremental cyclicality of new hire and low tenure wages remains insignificant with estimates in the second decimal place and very low p-values. Similarly downswing cyclicality seems as unimportant as before regardless of firm size.

The Leverage of Individual Data Points

Although the first step in the estimation exercise employs millions of worker spells the second stage uses a time series of only 33 data points (year effects). We need to check the results are stable in a time series sense. We do not conduct formal leverage tests such as Cook's D here as they are unlikely to have much power in such a small sample[12]. Instead we drop four observations from the beginning, end and middle of our year effects time series and inspect the extent to which our estimates are changed. As well as allowing us to examine time series structural stability in the usual sense, these three exercises have particular resonance in our context. The first four years (1977–1980) of our panel contained a high preponderance of new hires so omitting them will allow us to gauge the effects — if any — this peculiar feature has on our results. The middle four years (1991–1994) coincide with the first four years of German re-unification — an event which caused considerable macroeconomic upheaval. Finally, the last four years (2006–2009) correspond (roughly) to the introduction of major employment and benefit law changes in Germany known as Hartz IV which may have changed the way labour markets function (in the context of economic theory, it may have decreased the value of being unemployed, for example). They also straddle the start of the Great Recession so all in all were years of great macroeconomic "action" in Germany.

We re-estimated our main specifications (lines 1.2 to 1.4 in Table 1) for our three subsamples (omitting early middle and late years in turn). The results are given in Table 3 below. In each of the three cases we re-compute δ_t^u and $\widetilde{u}_t$ using the relevant subsample only.[13]

[12] For the record the largest Cook's Distance in an OLS estimation was only 0.34 (second largest was 0.1). Although not strictly valid for a number of reasons (e.g. the existence of autocorrelation and heteroscedasticity, low power etc) were we to treat these as test statistics for parameter constancy they would fall well below the relevant critical value of 2.8.

[13] In effect we constrain the long run average unemployment rate to be equal to the average rate for each relevant subsample. But in fact the results are very similar when we use all 33 data points to compute the long run average unemployment rate as we did before. The ADF statistics — which rejected nonstationarity comfortably in each case — have been suppressed.

Table 3

Second Stage Estimates of Semi Elasticities:-

The Effect of Dropping Time Series Data Points

Years Dropped	A:1977-1980		B:1991-1994		C:2006-2009	
LHS Vble.	θ^u	θ^d	θ^u	θ^d	θ^u	θ^d
β_t^a	-1.38	-0.26	-1.68	$-.09$	-1.65	-0.65
	(6.87)	(1.01)	(3.86)	(0.44)	(4.54)	2.48
β_t^0	-0.23	0.10	0.13	0.12	0.18	-0.21
	(1.38)	(0.59)	(0.48)	(0.74)	(0.76)	(1.20)
β_t^1	-0.17	0.04	0.16	0.05	0.10	-0.07
	(1.33)	(0.52)	(0.94)	(0.48)	(0.63)	(0.58)

Robust | t-ratios | in brackets

Only two of the estimates have changed sufficiently relative to their Table 1 counterparts to be worthy of comment; In case A, the upswing cyclicality for all workers has fallen by just over 20% and in case C the downswing cyclicality has doubled and become significant. It is hard to interpret these changes in terms of economic events. For example the latter would seem to imply Hartz IV led to more rather than less wage rigidity. It is possible that the increased downswing cyclicality is purely a consequence of sampling error (the change is less than 1.5 times the coefficient's standard error) and in any event there is still significant asymmetry here (a test of equality of the upswing/downswing coefficients gives a χ_1^2 value of 3.94). Most important of all, the key findings of equal treatment remain wholly intact.

Changing the Unemployment Rate to Include East Germany Post 1991

Our use of the West German Unemployment rate throughout this paper reflects our belief that post re-unification, East Germans did not compete for jobs located in the West. We stand by this view but think it at least prudent to check the sensitivity of our results to using the unemployment rate for all of Germany post 1991. The results are given in Table 4.

Table 4

Second Stage Estimates of Semi Elasticities:-

The Effect of Using the German Unemployment Rate Post 1991

$LHS\ Vble.$	θ^u	θ^d
β_t^a	-1.48	-0.07
	(4.21)	(1.01)
β_t^0	0.07	0.24
	(0.37)	(1.54)
β_t^1	-0.08	0.16
	(0.64)	(2.15)

Robust | t-ratios | in brackets

The key difference compared with Table 1 is that the downswing incremental cyclicality for low tenures has become positive and marginally significant. However the magnitude of the coefficient is small and its sign perverse. It is hard to rationalise why new hires would get the same cyclical wage deal as incumbents on arrival at the firm but during their second year of tenure would experience perverse downswing wage cyclicality whilst incumbents wages remain rigid. We might conclude one of two things: either the perverse nature of this result is evidence against the assumption that East German workers compete for West German jobs or that the result is a consequence of Type I error.

Cyclical Variation in Hours Worked

One drawback of the BeH noted above is that it documents total earnings rather than earnings per hour. It was for this reason that we chose to work with only full time employees because nearly all such workers work a standard length of week in Germany. Whilst it is impossible to show definitively the hours worked by the workers in our panel are unrelated to the unemployment rate there is some extraneous evidence that cyclicality of hours may in fact be quite low.

Annual OECD data for 1979 to 1997 on overtime hours as a percentage of total hours worked per person in Germany shows that hours worked are neither very variable (standard deviation of the measure is below 0.4%) nor are they particularly well synchronised with the business cycle (regressing on unemployment[14] gives a semi elasticity of -0.12 with a robust t-ratio of -1.83).

We also have obtained provisional estimates of the average length of working week (in hours) for full-time workers for each of the years 1977 to 2009.[15] Whilst these data are at present confidential we are able to summarise their time series properties. We measure the log of hours which we henceforth refer to as just "hours". We find that hours[16] have a standard deviation of only 0.6% and are virtually acyclical (estimated semi elasticity is perverse and equal to 0.08 with a robust t-ratio of 1.62).

Overall, whilst there appears to be some variation in hours worked by full time workers in Germany such variation would appear to be small and not particularly well related to the unemployment rate. We should also recall that our "benchmark" semi elasticity estimate for the entire cycle — where we do not allow for upswing/downswing asymmetries — is -1.26. This is in line with

[14] The variable is stationary and as before we estimate subject to a first order autocorrelated error.

[15] Source: working time calculation of the Institute for Employment Research (IAB).

[16] Once again we allow for a quadratic trend — weekly hours have a secular downward trend in Germany. We also allow for a change in mean post 1990 when the series excluded West Berlin.

the results from other German studies where hours *have* been controlled for such as those of Anger (2011) who uses the GSOEP. Again this suggests that bias arising from our inability to control for hours worked is small.

Finally, we note that even if hours were procyclical this would lead us to overestimate the extent of wage cyclicality and yet we find cyclicality in downswings that is close to zero and insignificant. Similarly, with respect to equal treatment, it is hard to see how procyclicality in hours for all workers — per se — could lead to false inferences on this matter.[17]

4 A Simple Equal Treatment Wage Contracting Model

In this section we analyse the extent to which a very simple equal treatment wage contracting model can reproduce the empirical characteristics of within firm wages and aggregate unemployment found above. To be clear, we try and match the properties of the West German unemployment rate and the β_t^a year effects derived above. Whilst focus is on the ability to match the estimates (1.2) in Table 1 we will also need to assess the model's ability to match the time series properties of firm wages (the β_t above which henceforth we call just "wages") and unemployment (u_t).

Given the relatively small number of observations (33 annual data points) we summarise the time series by estimating first-order autoregressive process (AR(1)) models for them. Results for the AR(1) coefficients (ρ_u and ρ_β) and for the unconditional standard deviations (σ_u and σ_β) are given in Table 5 below together with *ADF* test statistics. As before we allow a quadratic trend for wages.

[17] In fact we would require *differential* procyclicality of hours for this to happen. To see this suppose that an upswing (downswing) increases (decreases) new hire wages above (below) that of incumbents. For this to be hidden by the BeH it would be necessary for new hires to work fewer (more) hours than incumbents in the upswing (downswing).

Table 5

Time Series Properties of Wages and Unemployment

	$LHS\ Vble.(y)$	ρ_y	$\sigma_y(\%)$	$ADF\ (p-value)$
(3.1)	β_t	0.81 (3.13)	2.43	−3.75 (.08)
(3.2)	u_t	0.83 (9.45)	1.94	−3.39 (.01)

Note: t-ratios are reported in brackets.

As noted above, aggregate unemployment is highly persistent but as the ADF test shows clearly mean reverting. Wages are slightly less persistent but only reject a unit root at the 8% level rather than at 5%. Interestingly line (1.2) of Table 1 shows that the ADF test on residuals from the regression of wages on our downswing and upswing unemployment terms can reject a unit root with a p-value of less than 0.5% Wages and unemployment appear to have a common persistent component (i.e. they are co-persistent).

We now see the extent to which a simple model of wage contracts under equal treatment can capture the features of unemployment and wages detailed in line (1.2) of Table 1 and in Table 3. We give a brief overview of the model and its parameters here — a full exposition can be found in Snell and Thomas (2010).

There are a large number of identical risk neutral firms producing a homogenous good using only labour as an input in a diminishing returns production process. There are a large number of risk averse workers who choose to work full time or not to work at all, but who cannot borrow or save so that each period consumption of a worker equals her wages (w_{it}). Labour markets are competitive and in equilibrium, all workers wish to work. Each firm must pay its workers the same wage (equal treatment, no human capital). It is assumed that labour turnover is sufficiently high to guarantee that at least one new worker per firm is employed each period. Under this condition, wages will always be equal to the marginal product of labour. The firm has two conflicting objectives: a) to insure its risk averse workforce by arresting the fall of wages in bad times and b) to take advantage of lower outside options of the unemployment in bad times by lowering the wages of new hires. The requirement to pay new hires and incumbents the same results in a compromise where wages do fall in bad times but where the fall does not exceed some maximum rate. To understand the workings of the model consider the consequence of a negative productivity shock to all firms following a period when there was full employment. If the shock is sufficiently large, the maximum-rate-of-wage-fall constraint will bite and wages will not fall far enough to clear the labour market. There will then be involuntary unemployment and this will persist until either productivity recovers or until wages have fallen sufficiently far to clear the labour market — whichever happens first.

The two key equations determining (log) wage (w) and (log) employment (l) are

$$w_{it} = \max\{(A_t + w_{it-1}), a_t\}, \tag{3}$$

$$\text{where } A_t = \frac{\alpha}{\alpha\gamma - 1} \log \lambda - \frac{1}{\alpha\gamma - 1} \Delta a_t \text{ and}$$

$$l_t = \frac{1}{\alpha} a_t - \frac{1}{\alpha} w_t. \tag{4}$$

The first condition (3) determines the current wage (w_{it}) given the previous wage (w_{it-1}) and given current (a_t) and past productivity levels (a_{t-1}), whilst (4), the marginal product equals wage condition, will then determine the level of log employment l_t which, given a fixed labour force of size one, will also tie down the level of unemployment. The parameters α and γ are the inverse of the demand elasticity for labour and the coefficient of relative risk aversion (assumed constant) in per period utility. λ is a parameter directly related to the exogenous separation rates of workers.

For the simulations we use $\alpha = 1$, $\gamma = 4$ and $\lambda = 0.99$. The values for α and γ are in line with the range of estimates obtained in the macroeconomics literature and are very close to the values used in Snell and Thomas (2010). The parameter λ is harder to calibrate: workers have heterogenous quit rates and $1 - \lambda$ is the smallest quit probability across all workers in the firm. It will determine the importance of the downward real wage rigidity constraint. Again our chosen value of 0.99 is close to that used by Snell and Thomas (2010) for US data.

The model uses firm level TFP as an exogenous input. But TFP data are not available to us over the sample at either firm, sector or aggregate level. Therefore we follow the calibration methodology used in representative agent (in this case representative firm) RBC models. We feed the model a variety of synthetically generated productivity data and assess its ability to replicate the joint time series properties of German wages and unemployment found above. We consider the following process for log TFP:[18] $a_t = \rho_p a_{t-1} + \varepsilon_t$, with $var(\varepsilon_t) = \sigma_p^2$.

We consider low ($\rho_p = 0$), medium ($\rho_p = 0.3$), and high ($\rho_p = 0.6$) persistence cases, each with two values of σ_p: 1.5% and 2.0%.[19] We focus on estimates from the simulated data of a) the parameters in wage regression (2), see line (1.2) of Table 1, b) the standard deviation of unemployment (σ_u) and wages ($\sigma_w = \sigma_\beta$), and c) their respective AR(1) coefficients ($\rho_u, \rho_w = \rho_\beta$), see Table 5. Table 6 gives the results from the simulations. The last line in the Table summarises

[18] We add a 1% p.a. deterministic trend — roughly the growth rate of average German wages over the period.

[19] Estimating AR(1) processes for each of the 25 sectoral TFP processes in the KLEMS data for the 1991–2007 period gave estimates of ρ_p in the range 0 to 0.8 with an average value of about 0.3. The average conditional standard error was about 0.03 or 3%.

the estimates from Tables 1 and 3 — the targets of the simulations.[20]

Cases 1 and 2 are not archetypal productivity processes — TFP is more often than not highly persistent whether measured at sector or aggregate level. However, these white noise scenarios are informative as they help us to establish the propagation properties of the model. We see that the model generates persistence in wages — as one would expect from inspection of equations (3) and (4). The same cannot be said of unemployment which has low to zero persistence in these cases. Furthermore, looking at the high TFP persistence cases (5 and 6) we see that persistence in unemployment is always substantially below that of the TFP input. Case 6 comes closest to matching the features of the data but even here the autoregressive coefficient of unemployment is less than one half of what it should be.

The model scores better in other dimensions. With regards to the covariation of wages and unemployment, the model matches the strong cyclicality found in upswings — especially for the cases with high conditional standard derivation. In downswings however there is a perverse but small positive value in all cases. This is a second order effect in the model. We mentioned above that firms faced a conflict in bad times between a) the desire to insure its incumbents and b) the desire to exploit the low outside options of new hires by lowering their wages. However if the firm suffers persistently negative shocks, the number of new hires may become so low that the latter starts to dominate the former. In effect, the desire to insure incumbents starts to outweigh the desire to exploit new hires because there are so few new hires to exploit. Such scenarios rarely occur in any particular time series; but when they do, (small) rises in wages will coincide with further increases in unemployment and it is this that would appear to be generating our small positive semi elasticity in downswings. Tentative results from a two period version of the model subject to

[20]Estimates are based on 5000 time series data points and — as was the case above — estimation of the semi elasticities was subject to a first order error process.

directed search suggest that this second order feature would not be robust under search frictions.[21]

Table 6

Simulation Results

Case	$\rho_p, \sigma_p(\%)$	θ^u	θ^d	ρ_w	ρ_u	$\sigma_w(\%)$	$\sigma_u(\%)$
1	$0, 1.5$	-2.47	0.41	0.47	0.01	1.28	1.20
2	$0, 2.0$	-1.56	0.36	0.57	0.04	1.66	1.87
3	$0.3, 1.5$	-2.8	0.36	0.59	0.13	1.40	1.04
4	$0.3, 2.0$	-1.49	0.33	0.66	0.17	1.81	1.68
5	$0.6, 1.5$	-3.27	0.33	0.74	0.25	1.17	0.95
6	$0.6, 2.0$	-1.43	0.30	0.78	0.32	2.27	1.61
Data Estimates		-1.75	-0.27	0.81	0.84	2.43	1.94

In sum whilst the model can account for many of the features of wages and and co-features of wages and unemployment it cannot match the persistence in the latter. It may be that search frictions and/or adjustment costs (which we do not model) may have been responsible for the persistence in unemployment we see in the German data — effects we might expect were pre-eminent following unification in 1990.

5 Summary and Conclusion

Using a panel from a German administrative dataset (the BeH) we have shown that within firm wages are strongly procyclical in upswings (unemployment above its long term mean) but acyclical

[21] In this prototype model, it seems that if bad productivity persists, the number of potential new entrants arriving at the firm and with whom the firm could form a profitable match, increases. At this point, the second motive of the firm — to exploit new hires — seems to dominate the first (to insure incumbents). As a result, wages fall slightly.

in downswings. Furthermore, in a formal test of equal treatment the incremental cyclicality of new hires and of low tenured workers were found to be wholly insignificant. The implications are that involuntarily unemployed workers cannot price themselves into jobs by undercutting the wages of incumbents. A model of equal treatment insurance wage contracts can explain most of the unemployment/wage properties and co-properties displayed by the panel empirics. It can match the upswing (downswing) procyclicality (acyclicality) of wages and their univariate time series properties but it cannot match the persistence in German unemployment. It is possible that the latter is due to job search frictions which are not allowed for in the model. These frictions may have been particularly pre-eminent in Germany after unification. Extending the model to allow for these frictions is the subject of current research.

Appendix

Table A1
Number of Worker and Number of Employment Spells

Year	Number of workers	Number of spells			
		total	Entrants	Tenure $\geq$ 365 days	Tenure $\geq$ 770 days
1977	810,568	869,669	656,656	159,154	58,344
1978	1,157,780	1,247,651	624,795	502,843	124,318
1979	1,509,287	1,628,570	698,150	803,090	407,205
1980	1,770,961	1,902,761	679,665	1,091,367	700,575
1981	1,866,002	1,986,506	544,270	1,320,175	949,250
1982	1,901,587	2,011,057	411,314	1,499,014	1,150,429
1983	1,977,447	2,089,389	403,080	1,601,676	1,330,682
1984	2,140,097	2,318,946	495,007	1,723,970	1,485,500
1985	2,322,353	2,478,483	538,482	1,823,489	1,535,209
1986	2,538,722	2,698,955	577,171	2,003,206	1,647,679
1987	2,658,468	2,828,842	527,913	2,178,392	1,804,444
1988	2,818,570	3,004,902	557,150	2,337,633	2,013,375
1989	3,017,112	3,212,926	667,832	2,418,369	2,093,367
1990	3,247,368	3,504,928	769,202	2,582,148	2,155,439
1991	3,316,011	3,532,493	692,863	2,682,780	2,200,905
1992	3,282,401	3,488,297	550,927	2,806,739	2,390,438
1993	3,162,871	3,331,879	394,462	2,831,062	2,507,169
1994	3,151,618	3,309,202	422,841	2,799,978	2,529,469
1995	3,155,647	3,317,590	464,970	2,756,054	2,482,550
1996	3,112,065	3,277,341	403,464	2,783,412	2,509,590
1997	3,123,072	3,337,770	428,454	2,796,493	2,557,475
1998	3,163,382	3,363,190	481,051	2,767,321	2,499,262
1999	3,096,053	3,351,333	455,062	2,777,624	2,457,531
2000	3,128,313	3,349,849	508,261	2,740,557	2,456,755
2001	3,108,038	3,476,486	454,748	2,886,325	2,567,608
2002	3,004,282	3,332,497	357,615	2,856,975	2,519,505
2003	3,088,532	3,468,900	301,018	3,063,717	2,772,923
2004	2,998,807	3,295,283	275,865	2,945,038	2,737,292
2005	2,944,527	3,213,403	286,069	2,849,370	2,674,429
2006	2,939,288	3,188,402	306,398	2,804,215	2,584,693
2007	2,968,771	3,314,862	345,247	2,876,518	2,638,431
2008	2,954,916	3,206,159	336,396	2,794,097	2,557,064
2009	2,918,594	3,217,788	265,683	2,870,728	2,641,828
Sum	88,353,510	95,156,309	15,882,081	75,733,529	65,740,733

Table A2
Unemployment Rates

Year	Germany[1]	West Germany
2009	9.1	7.7
2008	8.7	8.2
2007	10.1	8.3
2006	12.0	10.2
2005	13.0	11.0
2004	11.7	9.4
2003	11.6	9.3
2002	10.8	8.5
2001	10.3	8.0
2000	10.7	8.4
1999	11.7	9.6
1998	12.3	10.3
1997	12.7	10.8
1996	11.5	9.9
1995	10.4	9.1
1994	10.6	9.0
1993	9.8	8.0
1992	8.5	6.4
1991	7.3	6.2
1990	7.2	7.2
1989	7.9	7.9
1988	8.7	8.7
1987	8.9	8.9
1986	9.0	9.0
1985	9.3	9.3
1984	9.1	9.1
1983	9.1	9.1
1982	7.5	7.5
1981	5.5	5.5
1980	3.8	3.8
1979	3.8	3.8
1978	4.3	3.8
1977	4.5	4.5

Notes: Unemployment rate of the dependent labor force

Source: Statistik der Bundesagentur für Arbeit, Arbeitslosigkeit im Zeitverlauf.

[1] 1977–1900 former Germany (West Germany including West Berlin), 1991–2009 reunited Germany (West and East Germany).

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