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The regional effects of Germany's national minimum wage*

Abstract: We show that the minimum wage introduced in Germany in 2015 led to spatial wage convergence, in particular in the left tail of the distribution, without reducing relative employment in low-wage regions within the first two years.

Key words: Difference-in-differences, employment, Germany, minimum wage
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JEL: J31, J58, R12

1 Introduction

While there is a vast and controversial literature about the implications of minimum wages for employment and the distribution of wages, little is known about the spatial implications of such a policy. With productivity and, hence, wage differences across locations, the introduction of a national minimum wage affects regions to different extents. While the policy bites hard in poor places, there is only a small fraction of workers earning less than the minimum in rich places.

We follow this idea when exploring the wage, employment, and migration effects of the federal minimum wage that was introduced in Germany in 2015. Since then, German employers have to pay at least 8.50 euros per hour corresponding to 48 percent of the median salary of full-time workers. This level is high compared to the US (36 percent) and because no similar regulation preceded the statutory wage floor, it represented a potentially significant shock to regions in the left tail of the regional wage distribution. [1]

To identify the differential effects across locations, we exploit the variation in the fraction of workers who earned less than the minimum in 2014 across German counties. We compare counties subject to different intensities of treatment in a difference-in-differences (DD) strategy that accounts

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1 The level is comparable to many other developed economies, see https://stats.oecd.org.

for heterogeneity in pre-treatment outcome trends. In doing so, we exploit a micro data set covering the universe of employment and unemployment in Germany from 2011 to 2016.

We show that the minimum wage policy raised the wages of low-wage workers without affecting employment. Unemployment even shrinks in regions with a high minimum-wage bite in 2015 relative to low-bite locations owed to a temporary reduction in in-migration, but these effects already vanish in 2016. The policy's primary effect thus far has been to transfer producer surplus to workers in low-wage regions, indicating that low-wage employees were paid below their marginal value product (Machin, Manning, and Woodland, 1993, Machin and Manning, 2004). Hence, the competitive labour market model has to be rejected.

This paper contributes to the literature on the labour market implications of minimum wages that largely builds on experience in the US. Our evidence is novel in that it is based on the largest European economy, focuses on the *regional* implications of a national minimum wage, and covers the effects on regional migration.[2]

2 Data

The empirical analysis is based on the Employment Histories (BeH) and the Integrated Employment Biographies (IEB) provided by the Institute of Employment Research (IAB) which contain individual-level data on the universe of labour market participants in Germany. Despite their comprehensiveness, the data do not include information about the number of hours worked. We therefore impute average working hours separately for full-time and part-time workers from an auxiliary regression that accounts for sector of employment, federal state of employment, and various socio-demographic attributes and uses a 1% sample from the 2012 census (for details, see section 5 of the online appendix). We find that full-time employees work approximately 40 hours per week while the number is lower for regularly employed (21 hours) and for marginally employed part-time workers (10 hours). Combining working hours with average daily earnings delivers hourly wages from which we compute the 2014 (the year prior to the policy change) share of workers (at the workplace) below the minimum wage for each of the 401 German counties (NUTS3 regions). Since labour markets are integrated across county borders, we define the minimum-wage bite as the average of the shares of below-minimum-wage workers at all counties, weighted by the bilateral commuting flows from the year 2010. Table 1 provides an overview of the key variables.

[2] See Brown (1999) and Neumark and Wascher (2008) for reviews and Dube, Lester, Reich (2010), Baek and Park (2016) and Caliendo et al. (2017) for more recent evidence.

Tab. 1. Summary statistics

VARIABLES	(1) mean	(2) sd	(3) min	(4) p10	(5) p25	(6) p75	(7) p90	(8) max
2014 minimum wage bite	14.84	3.06	7.10	11.26	12.46	16.75	19.33	25.43
Ln hourly wage at the 10th percentile	1.94	0.12	1.50	1.77	1.86	2.03	2.09	2.24
Ln hourly wage at the 25th percentile	2.31	0.10	2.00	2.17	2.24	2.38	2.44	2.62
Ln hourly wage at the 50th percentile	2.72	0.12	2.34	2.54	2.65	2.80	2.86	3.11
Ln labour force	11.07	0.66	9.49	10.29	10.66	11.48	11.86	14.18
Ln employment	10.96	0.66	9.38	10.19	10.56	11.36	11.74	14.00
Unemployment rate (percentage points)	9.85	4.32	2.19	4.90	6.51	12.38	16.17	26.59

Notes: Unit of observation is county-year. 401 counties are repeatedly observed over 2011-2016.

3 Empirical strategy

To evaluate the effects of the minimum wage policy on an outcome $y_{c,g,t}$ in county c in region g at time t, we use a difference-in-difference specification with a continuous treatment variable (Ahlfeldt et al., 2017). It allows for treatment effects on both the level and the trend of an outcome (Ahlfeldt and Feddersen, 2018) and controls for county-specific time trends. In particular, we have

$$y_{c,g,t} = \beta_1 T_c \times I(t \geq 2015) + \beta_2 T_c \times I(t \geq 2015) \times (t - 2015) + \mu_c + \vartheta_{g,t} + (\eta_c \times t) + \epsilon_{c,g,t}, \quad (1)$$

where T_c is the treatment variable (the minimum wage bite) that interacts with time through an indicator variable $I(.)$ that takes the value of one if the observation refers to years 2015 or 2016, and zero otherwise. Further, the inclusion of the second term allows us to identify time-specific treatment effects. μ_c are county effects, $\vartheta_{g,t}$ denote region (East Germany, West Germany) effects interacted with year effects and $\epsilon_{c,g,t}$ is a random error. We also control for county-specific effects that interact linearly with time t, $(\eta_c \times t)$, to absorb unobserved spatio-temporal heterogeneity that could induce a non-parallel-trends problem. The time-specific treatment effect we estimate is $\frac{\partial y_{c,g,t,I=1}}{\partial T_c} - \frac{\partial y_{c,g,t,I=0}}{\partial T_c} = \hat{\beta}_1 + \hat{\beta}_2(t - 2015)$, where hats indicate estimated values.

To depict the temporal pattern of the treatment effect without imposing parametric constraints, we use an intervention-study design of the following form:

$$y_{c,g,t} = \sum\nolimits_{Z \neq 2014} \beta_Z T_c \times I(t = Z) + \mu_c + \vartheta_{g,t} + \epsilon_{c,g,t}. \quad (2)$$

The estimated time-varying effects $\hat{\beta}_Z$ capture the effect of the treatment on the outcome $\frac{\partial y_{c,g,t=Z}}{\partial T_c} - \frac{\partial y_{c,g,t=2014}}{\partial T_c}$ and the effects of a time-trend that interacts with unobserved county-specific effects, η_c in (1). To control for a confounding effect if $cov(\eta_c, T_c) \neq 0$, we compute the treatment effect at time t=Z as the difference between $\hat{\beta}_Z$ and a linear extrapolation of the trend in $\hat{\beta}_Z$ during the pre-treatment period. The counterfactual is then the same as in specification (1). The treatment effects for Z>2014 are identical in both specifications in this setting with two post-intervention periods.[3] We report clustered standard errors (by county) as they turn out to be more conservative than a panel-derivative of Conley's (1999) standard errors.[4] We acknowledge that T_c incorporates hours worked, which are measured with error at the individual level. Within each county, however, we aggregate over a large number of workers (≈150k on average), thus the county-level mean and variance of the error is likely near zero.

A precisely estimated zero effect of the minimum wage bite on employment will have important policy implications. However, given that there is suggestive evidence for some employers paying less than €8.50 per hour after 2015 (e.g. Mindestlohnkommission 2016), a zero-employment effect could be driven by non-compliance if the (unobserved) compliance rate and the minimum wage bite were spatially correlated. To rule out that an economically and statistically insignificant employment effect is driven by non-compliance, we show that the bite has a significantly positive effect on wages, i.e. there is at least imperfect compliance. Further, we compute the minimum wage bite using wage and employment data from 2014 (before the policy was implemented) to ensure that the compliance rate is not a component of the bite measure. We develop the above argument formally in section 4 of the appendix.

4 Results

In line with the spatial distribution of the minimum wage bite (see Figure A1 in the online appendix), the minimum wage appears to have had a stronger bite in the economically still weaker eastern states. At the 10th percentile of the distribution within counties, hourly wages increased from 2014 to 2016 by about €1.25 in the eastern states, compared to less than €1 in the western states. We note that we hold the (imputed) hours worked constant, so hourly wages in our data cannot increase due to reductions in working hours.

In Figure 1, we use our baseline empirical specification (2) to more formally evaluate the effects of the minimum wage. Panel A shows that the minimum wage policy helped low-wage workers (10th

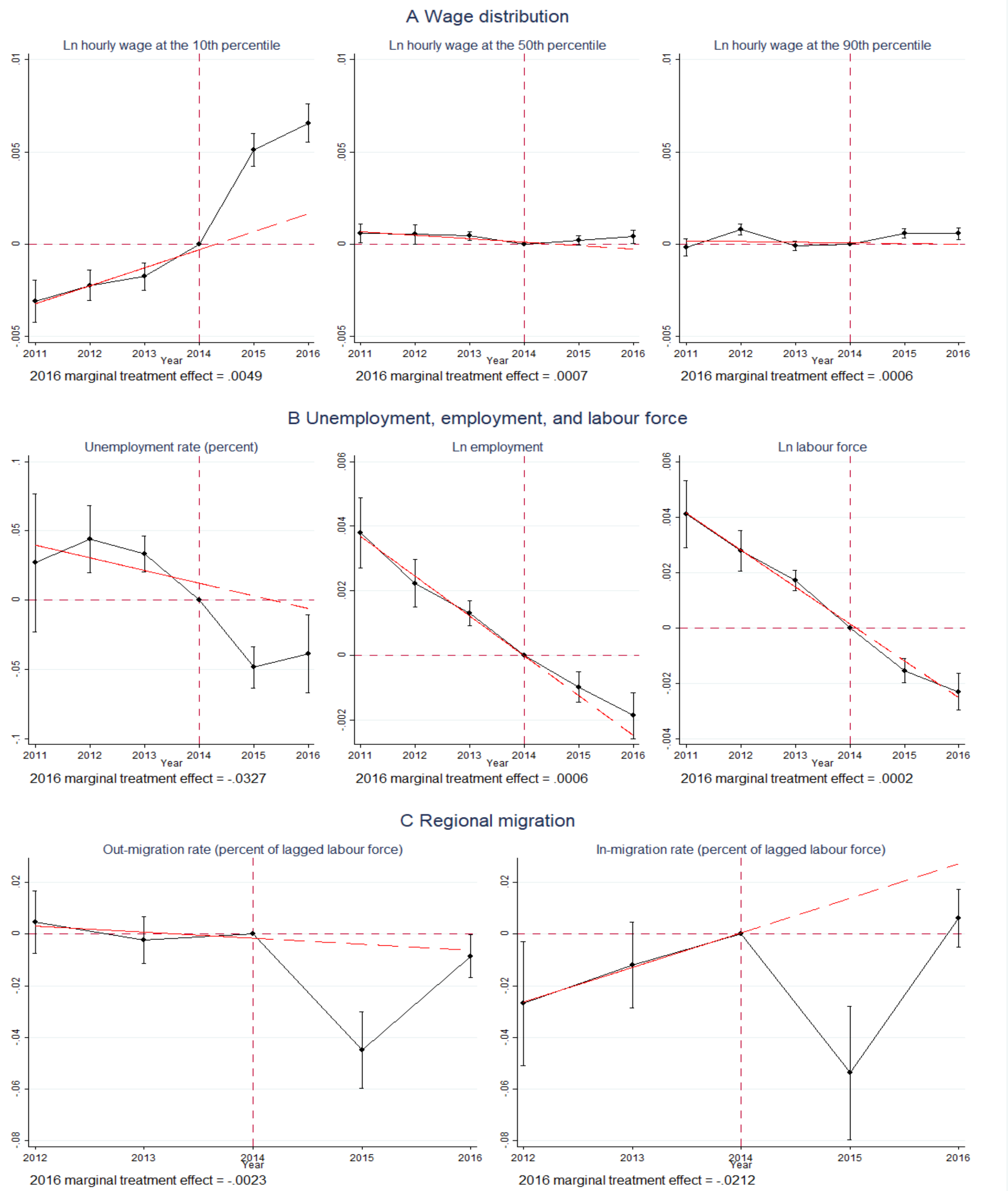
3 Notice that we do not add $(\eta_c \times t)$ to specification (2) because this means we have to drop another $\beta_Z T_c \times I(.)$ interaction term and the point estimates are no longer the same.

4 We use the Stata module Conley spatial HAC for models with fixed effects by Thimo Fetzer with cutoffs of 100 kilometers and one year to address a correlation of errors cross space and time.

percentile) to increase their wage relatively more in counties with a higher bite. The treatment effect (gap between the 2016 dot and the dashed line) implies that an increase in the minimum wage bite by one percentage point is associated with a 0.5% larger increase in the low wage. The lower bound of the 95% confidence interval (indicated by the error bar) of the 2016 treatment clearly exceeds the counterfactual trend (dashed line), so the effect can be considered statistically significant. The low wage in a county at the 90th percentile of the minimum wage bite distribution (henceforth high-bite county) compared to one at the 10th percentile (henceforth low-bite county) increased by some additional 4.0%$(= \exp(0.0049 \times (19.33 - 11.26)) - 1)$. The respective effects at the 50th and 90th percentiles of the wage distribution are economically small as expected.

These results imply a spatial wage convergence that was intended by the policy, but the results could be mechanically driven by reduced employment rates of low-wage workers. In Panel B we therefore replicate the analysis using the unemployment rate, employment, and labour force as outcome variables. We observe that a one-percentage point increase in the minimum-wage bite significantly reduces the unemployment rate by approximately 0.05 percentage points in 2015 and 2016. These changes appear initially to be driven by a combination of a higher employment level and a lower labour force in high-bite counties, while for the year 2016 the primary explanation is an increase in employment. To further assess the importance of these channels we use the standard error estimates from specification (1). The results show that the treatment effect on employment is 0.06% in 2016 which is significant at the 5 percent level, while the corresponding effect on the labour force in 2015 is -0.04% which just fails to be significant at the 10 percent level (see Table A3 in the appendix for details). A possible explanation for the decrease in the local labour force in high-bite counties in 2015 are changes in migration. Panel C shows that in-migration as well as out-migration rates drop sharply in 2015, but the former effect is considerably larger in magnitude. While the effect on the in-migration rate continues to be negative in 2016, there is no significant effect on the out-migration rate. While this combination may have led to a further reduction in the labour force, it appears that the increase in employment levels is sufficiently high to outweigh this effect.

Fig. 1. Effects of the minimum wage



Notes: Each panel illustrates the results of separate county-year-level DD regressions of an outcome against treatment-year interactions (excluding the 2014 base year), county effects and Year x East Germany effects. Treatment variable is the 2014 minimum wage bite (commuting-flow weighted average of shares of below-minimum-wage workers of surrounding counties). Dots are the estimated treatment-year effects and vertical error bars are the corresponding 95% confidence intervals. The red solid line is the linear fit into treatment-year effects up until 2014 and the dashed line is the linear extrapolation. The treatment effect is the 2016 difference between the point estimate and the dashed line.

5 Conclusion

Our analysis reveals that the introduction of the federal minimum wage in Germany in 2015 led to spatial wage convergence. As expected, wages in low-wage counties increased more rapidly than in high-wage counties, especially so for workers in the left tail of the wage distribution. This shift in the spatial distribution of wages did not come at the expense of significant job loss in low-wage regions (relative to high wage regions). In contrast, we find that locations with a higher share of low-wage workers experienced lower unemployment rates in 2015 and 2016, although these effects are economically marginal. While these changes appear to be initially driven by a reduction in the size of the labour force in high-bite counties, increases in employment levels are the primary driver in the year 2016.

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The regional effects of a national minimum wage: Appendix

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1 Introduction

This appendix adds to the main paper by providing complementary information, results and analyses. It is not designed to stand alone or replace the reading of the main paper.

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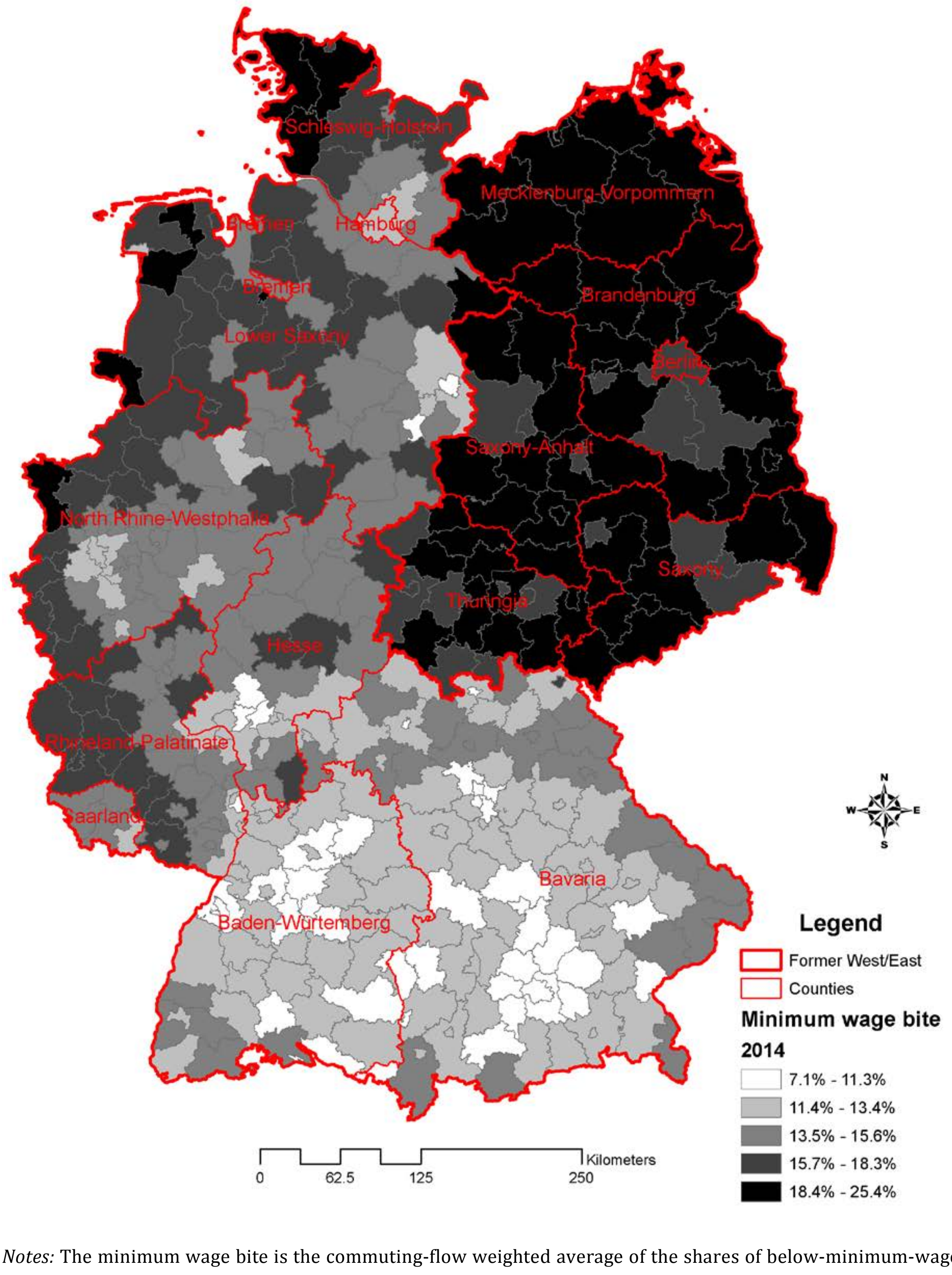
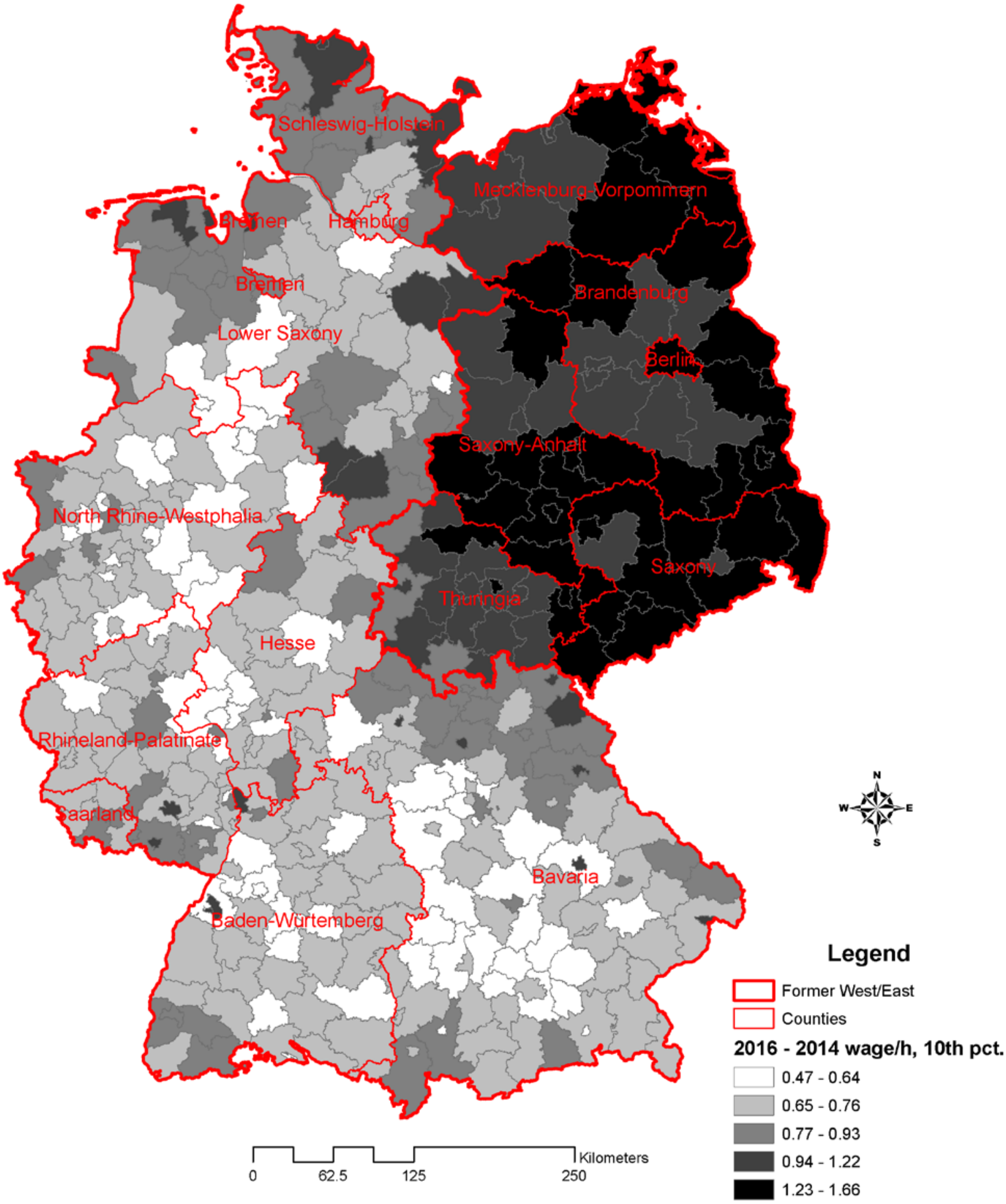
2 Spatial minimum wage bite and changes in low wages

Fig A1. Minimum wage bite



Notes: The minimum wage bite is the commuting-flow weighted average of the shares of below-minimum-wage workers (at workplace) of surrounding counties.

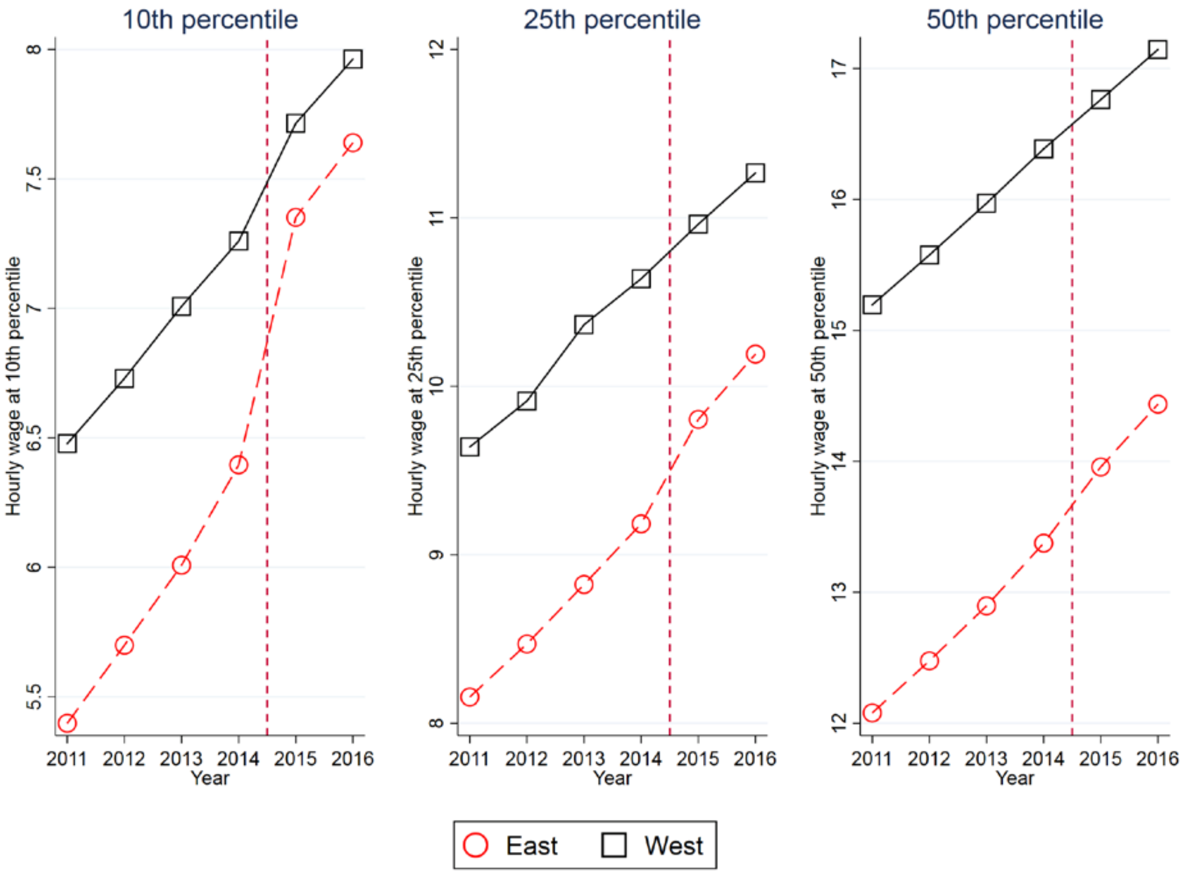
Fig A2. Change in low wages from 2014 to 2016



Notes: The low wage is defined as the 10th percentile in the distribution of hourly wages (in euros) in a county.

3 Descriptive evidence

Fig A3. Trends in wages by region and percentiles



Notes: East indicates six federal states within the territory of former East-Germany, including Berlin.

Figure A3 shows trends in wages, separated by East and West Germany, for the 10th, the 25th and the 50th percentile of the respective distributions. While the impact on the 10th percentile wage is visible, in particular in the East, it still falls short of the statutory minimum of 8.50 euros. This observation is in line with survey evidence from the DIW Berlin according to which an estimated 1.8 million workers in Germany still earn less than the minimum wage[1] and with evidence from the minimum wage commission (Mindestlohnkommission, 2016, p. 50). Appendix 4 details how we address this issue econometrically.

[1] See https://www.diw.de/en/diw_01.c.572687.en/topics_news/in_germany_approximately_1_8_million_workers_eligible_for_the_minimum_wage_are_earning_less.html for details.

4 Dealing with non-compliance

In section 3 of the main paper, we argue that because of the potential non-compliance effects, it is insufficient to evaluate the policy based on quantity effects alone. To illustrate the problem and the empirical strategy to address it, we consider that a labour outcome such as total employment in a region depends on the effective wage bite t and some other factors X,

$$y = f(X, t).$$

We do not observe the actual bite t, but the potential bite T. We further assume that t=T C, where C is the compliance rate, i.e. the share of workers to whom the bite is binding.

The total differential of the above equation is:

$$\mathrm{d}y = \frac{\partial f}{\partial X}\mathrm{d}X + \frac{\partial f}{\partial t}\mathrm{d}t.$$

Our empirical task is to estimate $\frac{\partial f}{\partial t}$. Substituting in TC for t, and rearranging, we get

$$\frac{dy}{dT} = \frac{\partial f}{\partial X}\frac{\mathrm{d}X}{dT} + \frac{\partial f}{\partial t} + \frac{\partial f}{\partial t}\frac{\mathrm{d}C}{dT}$$

Our empirical strategy is carefully designed to control for unobserved factors X (we control for arbitrary level effects and smooth individual trends in a difference-in-differences setup). Since we are confident in claiming that we hold the effect X on the outcome constant in our empirical specifications, we can drop the first term on the right-hand side of the above equation to get:

$$\frac{dy}{dT} = \frac{\partial f}{\partial t} + \frac{\partial f}{\partial t}\frac{\mathrm{d}C}{\mathrm{d}T}.$$

Obviously, our estimate of $\partial f/\partial t$ is unbiased only if $\mathrm{d}C/\mathrm{d}T = 0$, i.e. non-compliance is uncorrelated with the potential treatment measure. With non-compliance, we have $\mathrm{d}C/\mathrm{d}T < 0$. For $\mathrm{d}C/\mathrm{d}T = -1$, an estimate of $\mathrm{d}y/\mathrm{d}T = 0$ could simply imply that there is no policy effect because there is no compliance. So, even a precisely estimated zero for a labour quantity effect does not carry much meaning in a world with potential non-compliance.

To address this problem, we also consider the treatment effect on wages. In perfect analogy to the above, we consider the following wage equation:

$$w = g(X, t)$$

to get

$$\frac{dw}{dT} = \frac{\partial g}{\partial t} + \frac{\partial g}{\partial t}\frac{\mathrm{d}C}{\mathrm{d}T}.$$

If, empirically, we can show that $\mathrm{d}w/\mathrm{d}T > 0$, this will necessarily imply that $\mathrm{d}C/\mathrm{d}T > -1$. For $\mathrm{d}C/\mathrm{d}T > -1$, however, an estimate of $\mathrm{d}C/\mathrm{d}T = 0$ cannot be driven by non-compliance alone.

5 Estimating the number of hours worked per week

The Employment Histories (BeH) contain an employee's average daily earnings, but no information on the number of hours worked. In order to estimate an hourly wage variable and to determine whether an individual earns above or below the minimum-wage threshold, we utilize information from the 2012 version of the German census. This dataset is derived from a representative household survey that is conducted by the statistical offices of the federal states. It contains detailed individual- and household-level information on approximately 1% of households in Germany.

The main variable used in the analysis is the number of hours regularly worked per week. In order to control for differences in working hours between different groups in the population, we regress this variable on a set of indicators for gender, part-time status, place of employment at the level of the federal state and sector of employment. We use the 21 sectors based on the 2008 version of the *Klassifikation der Wirtschaftszweige*. In addition, we control for mean adjusted individual- and household-level characteristics (age, German nationality, tertiary education, marital status, personal income, household size, number of children and household income) as shown in Equation S1:

$$\begin{aligned}\ln[h_i] = \alpha_0 + \alpha_1 fem_i + \alpha_2 pt_i + \alpha_3 fem_i pt_i + \sum_{j=2}^{16} \beta_j D_i(state_i = j) \\ + \sum_{k=2}^{21} \gamma_k D_i(sector_i = k) + \boldsymbol{\delta_i}'\boldsymbol{x_i} + u_i\end{aligned} \tag{A.1}$$

We estimate (A.1) separately for regular and marginal employees. Setting individual- and household-level characteristics to their sample means, we next compute the predicted number of hours worked for each cell defined by type of employment, gender, part-time status, place of employment and sector of employment. Since this set of variables is also part of the BeH data, we are able to assign the corresponding predicted number of hours to all individuals within the corresponding cells. Table A.1 provides an overview of the predicted average number of hours for different cells.

Tab. A1. Predicted weekly working hours

Gender	Part-time status	Hours (regular)	Hours(marginal)
Female	Full-time	39.43	-
Female	Part-time	21.24	9.98
Male	Full-time	41.22	-
Male	Part-time	20.71	10.43

Notes: Mean values across federal states and sectors.

6 Parametric treatment effects

In this section, we summarize estimates for various outcomes according to (1). Table A2 reports treatment effects on wages for five distinct percentile levels. In Table A3, we show treatment effects on (un)employment and migration rates.

Tab. A2. Parametric treatment effect on wages

	(1) Ln hourly wage at the 10th percentile	(2) Ln hourly wage at the 25th percentile	(3) Ln hourly wage at the 50th percentile	(4) Ln hourly wage at the 75th percentile	(5) Ln hourly wage at the 90th percentile
T x (year >= 2015)	0.0044***	0.0009***	0.0003**	-0.0002	0.0005***
	(0.0004)	(0.0003)	(0.0001)	(0.0002)	(0.0002)
T x (year >= 2015) x (year - 2015)	0.0005	-0.0003	0.0004***	0.0003	0.0000
	(0.0003)	(0.0002)	(0.0001)	(0.0002)	(0.0001)
2016 treatment effect	0.0049***	0.0006*	0.0007***	0.0001	0.0006**
	(0.0006)	(0.0003)	(0.0002)	(0.0004)	(0.0002)
County effects	Yes	Yes	Yes	Yes	Yes
Year effects	Yes	Yes	Yes	Yes	Yes
County*trend effects	Yes	Yes	Yes	Yes	Yes
East*year effects	Yes	Yes	Yes	Yes	Yes
R^2	0.993	0.995	0.997	0.995	0.995
Obs.	2,406	2,406	2,406	2,406	2,406

Notes: Standard errors (in parentheses) clustered on counties. Berlin assigned to East-German states (denoted by the binary variable East). Treatment (T) is the percentage of workers below the minimum wage in 2014 (the year before the policy was introduced). (year >= 2015) is a dummy variable taking the value of one if the condition is true. County*trend effects are county-specific linear trends. * p < 0.1, ** p < 0.05, *** p < 0.01.

Tab. A3. Parametric treatment effect on employment and migration rates

	(1) Unemployment rate (percent)	(2) Ln employment	(3) Ln labour force	(4) Out-migration rate (percent of lagged labour force)	(5) In-migration rate (percent of lagged labour force)
T x (year >= 2015)	-0.0517***	0.0003	-0.0004	-0.0410***	-0.0677***
	(0.0158)	(0.0002)	(0.0002)	(0.0083)	(0.0153)
T x (year >= 2015) x (year - 2015)	0.0190	0.0003**	0.0006***	0.0387***	0.0464***
	(0.0137)	(0.0002)	(0.0002)	(0.0087)	(0.0145)
2016 treatment effect	-0.0327	0.0006**	0.0002	-0.0023	-0.0212***
	(0.0269)	(0.0003)	(0.0004)	(0.0082)	(0.016)
County effects	Yes	Yes	Yes	Yes	Yes
Year effects	Yes	Yes	Yes	Yes	Yes
County*trend effects	Yes	Yes	Yes	Yes	Yes
East * year effects	Yes	Yes	Yes	Yes	Yes
R^2	0.803	0.98	0.975	0.691	0.548
Obs.	2,406	2,406	2,406	2,005	2,005

Notes: Standard errors (in parentheses) clustered on counties. Berlin assigned to East-German states (denoted by the binary variable East). Treatment (T) is the share of workers below the minimum wage in 2014 (the year before the policy was introduced). (year >= 2015) is a dummy variable taking the value of one if the condition is true. County*trend effects are county-specific linear trends. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.