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The heterogeneous regional effects of minimum wages in Poland

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Abstract

We evaluate the impact of large minimum wage hikes on employment and wage growth in Poland between 2004 and 2018. We estimate panel data models utilizing the considerable variation in wage levels, and in minimum wage bites, across 73 Polish NUTS 3 regions. We find that minimum wage hikes had a significant positive effect on wage growth and a significant negative effect on employment growth only in regions of Poland that were in the first tercile of the regional wage distribution in 2007. These effects were moderate in size, and appear to be more relevant for wages. Specifically, if the ratio of minimum wage to average wage had remained constant after 2007, by 2018, the average wages in these regions would have been 3.2% lower, while employment would have been 1.2% higher. In the remaining two-thirds of Polish regions, we find no significant effects of minimum wage hikes on average wages or on employment.

KEYWORDS

minimum wage, panel data, spatial heterogeneity

JEL CLASSIFICATIONS

J21; J23; J38

1 | INTRODUCTION AND MOTIVATION

Minimum wages are one of the most popular labour market policies in both the developed and the developing world. Among the 21 EU countries that had statutory national minimum wages in 2008,

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by 2018, Poland had recorded the fourth largest increase in the real value of the minimum wage (58%), and the fourth largest increase in the ratio of the minimum wage to the average wage (7.1 pp).[1] However, between 2002 and 2007, the minimum wage in Poland grew by only 9% in real terms, and declined in relation to the average wage, while wage inequality widened.[2] The government that took power in 2007 increased the minimum wage by 20%, and a series of hikes followed, reflecting a general shift in the attitude towards the minimum wage in Poland. However—and perhaps surprisingly—there has so far been no systematic evaluation of the employment or the wage effects of these minimum wage hikes in Poland. Our paper fills this gap.

In this paper, we evaluate the impact of minimum wage hikes on employment and wage growth in Poland between 2004 and 2018. To do so, we utilize the substantial variation in the minimum wage bite across the 73 NUTS 3 subregions in Poland. We treat the regional minimum wage bite as a continuous treatment variable (Ahlfeldt et al., 2018), and we follow Meer and West (2016) in estimating the effects of changes in the minimum wage on the dynamics of outcome variables. This approach is well-suited to the empirical and institutional setting in Poland. First, as Poland experienced steady increases in employment and wages throughout the period studied, we assume that minimum wage hikes may have affected the growth of these variables rather than their levels. Second, our model is equivalent to a model with the linear trends in variable levels removed, which helps us to isolate the effects of the acceleration of minimum wage hikes that began in 2008.

Our first contribution is to assess the spatially differentiated effects of minimum wage hikes in Poland. By studying labour markets at the subregional level, we are able to grasp much more nuanced differences in the minimum wage effects than previous studies, which analysed labour markets across 16 NUTS 2 regions in Poland (Fialová & Mysíková, 2020; Majchrowska et al., 2016). This is crucial, because the variance of subregional wages within particular NUTS 2 regions in Poland has been substantial: in 2007, 75% of the variance of average wages at the NUTS 3 level could be attributed to the within-NUTS 2 variance, whereas only 25% could be attributed to the between-NUTS 2 variance. Moreover, the within-region variances of wages were significantly different across the NUTS 2 regions. By studying the NUTS 3 subregions, we are able to examine finely disaggregated differences in wages that translate into differences in the minimum wage bite.

Our second contribution is to provide empirical evidence on the spatial heterogeneity of minimum wage effects in Poland. The effects of the minimum wage are likely to differ depending on the level of minimum wage bite (Cengiz et al., 2019; Manning, 2021a). There are several reasons for this. First, higher minimum wage bite is often associated with wage compression close to the minimum wage level, so each following hike can be binding for a larger share of the workforce (Neumark & Wascher, 2008). Higher bite should therefore translate into larger impacts on wages and employment. However, higher bite may also be associated with lower compliance with minimum wage hikes, or with stronger adjustment of other job attributes, such as working hours or noncash compensation (Clemens, 2021), which could attenuate the effects of hikes on wages and employment. Second, employers are likely to have some monopsonistic power in local labour markets, because of search frictions or market concentration (Manning, 2021a). In a monopsonistic labour market, minimum wage hikes would raise employment when the bite is low, but would reduce employment when the bite is

[1]Bulgaria, Latvia and Romania were the only countries with larger increases in the real value of the national minimum wage, whereas Croatia, Latvia and Slovenia were the only countries with larger increases in the minimum-to-average wage ratio. Germany first introduced a national minimum wage in 2015, and is not included in the comparison.

[2]The D9/D1 ratio of annual gross earnings increased from 3.89 in 2002 to 4.32 in 2006, and the D5/D1 ratio increased from 1.99 in 2002 to 2.05 in 2006. The periods used here reflect the data availability in our study.

high (Manning, 2021b). In consequence, the effects of minimum wage can depend on the context, market structure and minimum wage bite (Card & Krueger, 1995). To evaluate this heterogeneity, we allow for the minimum wage effects to differ in subregions that belong to the first tercile of the NUTS 3 wage distribution in 2007; that is, before minimum wage hikes accelerated in Poland.

We find that the effects of minimum wage hikes in Poland were significant only in the subregions that belonged to the bottom 33%, as measured by the average wages before the shift in the minimum wage policy in 2007. In 2007, the minimum wage bite in these subregions was at a level of 48% or higher. In these subregions, the hikes led to significantly higher average wages and lower employment. We also find that employment adjusted to the minimum wage hikes in the same year, whereas wages took longer to adjust, as wages responded to the hikes in both the current and the previous year. In the remaining subregions, no significant effects of minimum wage hikes are found. Our findings are in line with the recent evidence for Germany, which shows that the minimum wage effects have been significant only in areas that had relatively low wages before the introduction of the minimum wage (Ahlfeldt et al., 2018; vom Berge & Frings, 2020).

Importantly, the economic significance of the positive wage effects appears to be larger than that of the negative employment effects. Using our regression results, we find that if the minimum-to-average wage ratio had not changed since 2007, in 2018, the average wage in the low wage subregions of Poland would have been 3.2% lower, whereas employment would have been 1.2% higher. We also provide indicative evidence that the employment effects of minimum wage hikes differed between various groups of workers. In particular, they have been negative among men and among workers in industry, but positive among women and among workers in services. These heterogeneities may indicate that substituting labour with capital has been easier in industry than in services. They may also suggest that higher minimum wages pushed up the labour supply of women, in line with the mechanisms described by Card and Krueger (1994). Unfortunately, due to a lack of data, we cannot assess the sector-specific or gender-specific wage effects.

Our findings shed new light on the effects of minimum wage hikes in Poland, and in Central Eastern Europe economies (CEE) more generally. Majchrowska et al. (2016) argued that minimum wage hikes could limit youth employment growth in less-developed regions in the south-east of Poland. However, our findings show that the small negative employment effects are present in the least developed subregions spread around the country, and have been accompanied by noticeable positive wage effects. Trade-offs between employment and wages have been reported by Baranowska-Rataj and Magda (2015) and Kamińska and Lewandowski (2015). However, those studies were based on annual labour market flows that were constructed using individual Polish labour force survey data. When relying on these data, controlling for unobservable characteristics that may influence both employment status and wage is challenging, which makes it difficult to isolate the effects of minimum wage hikes. Significant, but quantitatively small effects of minimum wage hikes on job separations in Poland were found by Albinowski (2018), who was able to control for time-invariant individual characteristics by using employees' tax return data. Finally, surveys of managers in Central and Eastern European countries have shown that firms often increase wages in response to minimum wage hikes, and that reducing employment tends to be less relevant as an adjustment channel than increasing productivity, reducing non-labour costs and raising product prices (Bodnár et al., 2018). Firms in CEE have also reported that they are more likely to hire fewer new employees than they are to terminate existing employment contracts (Bodnár et al., 2018). These results suggest that studies of worker-level job separations may fail to capture a major component of employment adjustment, while we are able to grasp it by analysing aggregate, subregional employment levels.

In the next section, we describe the setting and the evolution of minimum wage levels in Poland. In Section 3, we outline our methodology and data. In Section 4, we provide descriptive evidence, and in Section 5, we present econometric results. In Section 6, we conclude and discuss the policy implications of our findings.

2 | MINIMUM WAGE IN POLAND

In Poland, minimum wage regulations have been in force since 1956. Until 2002, the level of the minimum wage was set by the Minister responsible for Labour and Social Affairs. Since 2002, it has been set annually by the Tripartite Commission for Social and Economic Affairs (replaced in 2015 by the Social Dialogue Council)[3] based on proposals submitted by the government. If the Commission/Council cannot reach a consensus, the government decides independently. Additionally, since 2003, the minimum wage proposed in a given year cannot be lower than the minimum wage from the previous year, adjusted by the forecasted change of the Consumer Price Index (CPI). Moreover, since 2006, if the minimum wage in a given year is lower than 50% of the average wage in the economy, then in the following year, the minimum wage must be increased by at least two-thirds of the forecasted nominal GDP growth.

The coverage of the minimum wage in Poland is, in essence, uniform: all workers with an employment contract based on the labour code are covered. However, individuals who were self-employed or who were employed through a civil law contract (a contract of mandate or a contract for products) were not covered by minimum wage until 2016. These workers are not included in our sample.

After the law changed in 2002, the level of the minimum wage remained fairly stable until 2007, both in real terms (on average, it remained at 1,079 zloty based on 2015 prices)[4] and in relation to the mean wage (on average, it was 34% of the mean wage, Figures 1 and 2). In relation to the median wage, the minimum wage even declined slightly (Figure 2). However, a series of increases have been implemented since 2007. Between 2007 and 2017, the minimum wage rose from 1,124 zloty to 1,972 zloty, which represented an increase in real terms of 76%. At the same time, the mean wage rose 30% in real terms, which means that the minimum-to-mean wage ratio increased from 33% in 2007 to 44% in 2017. Similarly, the minimum-to-median wage ratio increased from 40% in 2007 to 54% in 2017 (Figure 2). The largest increases in relative terms were implemented in 2008, when the minimum wage rose by 20%; and in 2009, when it increased by 13%.

3 | METHODOLOGY AND DATA

3.1 | Methodology

To assess the effects of minimum wage on outcome y_{it} in subregion i and time t, we take advantage of the substantial variation in the minimum wage bite across the NUTS 3 subregions in Poland. We evaluate the minimum wage bite as a continuous treatment variable (Ahlfeldt et al., 2018). We study two outcome variables: employment and real wages.

Following Meer and West (2016), we estimate the effects of minimum wage changes on the dynamics of the outcome variables. This approach is particularly suitable for Poland, because of the

[3]The Council has a slightly broader mandate than the Commission had, but the process of setting the minimum wage remained intact. Both have included selected ministers, member trade unions and employers' organizations.

[4]Unless stated otherwise, monetary values are given in real terms as per 2015 prices.

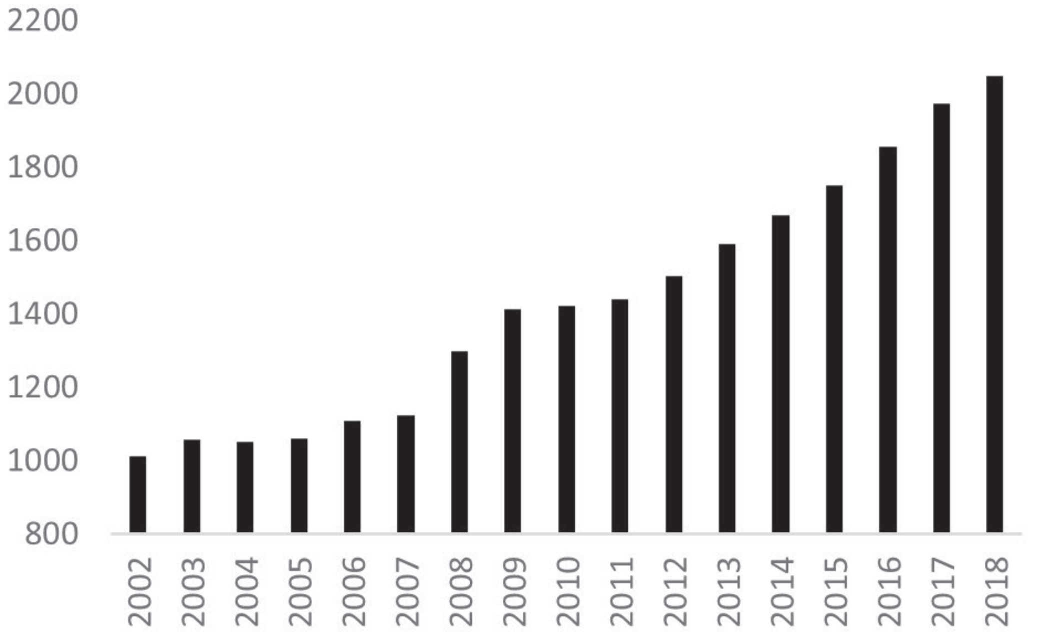


FIGURE 1 Monthly minimum wage in Poland, 2002–2018 (in zloty, 2015 prices)
SOURCE: Own calculations based on the Statistics Poland data.

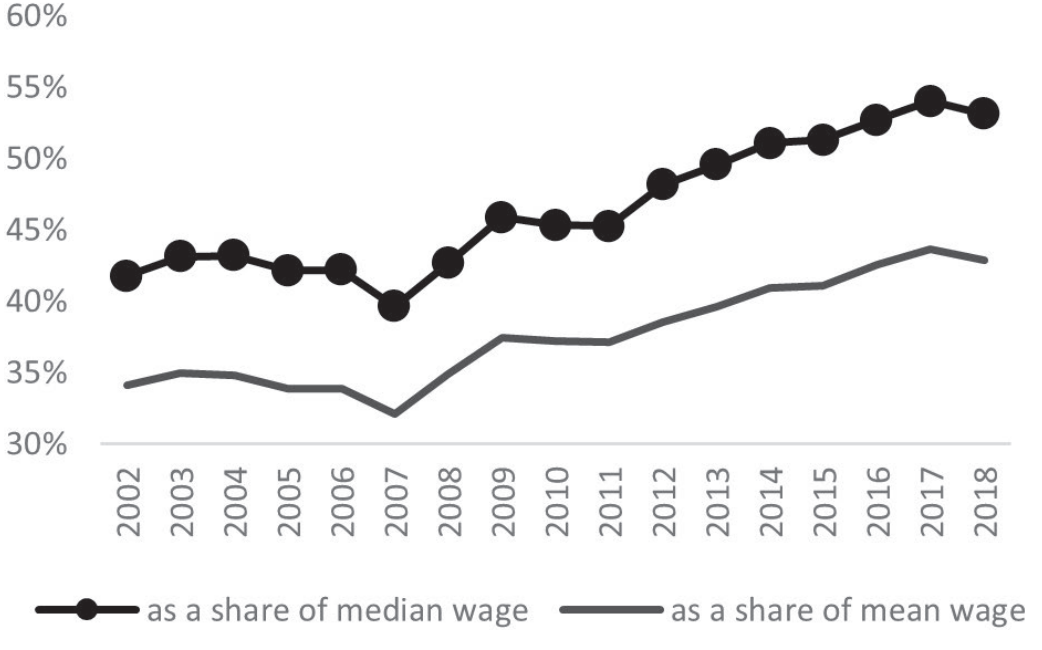


FIGURE 2 Monthly minimum wage as a share of the mean and the median wage in Poland, 2002–2018
SOURCE: Own elaboration based on the OECD Statistics data.

persistent growth in total employment and wages in the country throughout the period covered by our analysis. In other words, outcome variables are non-stationary in levels (formal tests are reported in Table A2). Moreover, our model is equivalent to a model in which the region-specific linear trends in variable levels were removed. This approach is parallel to Ahlfeldt et al. (2018) and Monras (2019) who use variable levels, but control for region-specific linear trends. This makes it suitable for the institutional setting in Poland, where the minimum wage was increased in each year covered by our sample, but the annual hikes were much larger after than before 2008.

Formally, we estimate the following model:

$$\Delta y_{it} = \alpha_i + \beta \times \Delta mw_{it} + \beta_{LW} \times \Delta mw_{it} \times LW_i + \varphi \times \Delta controls_{it} + \varepsilon_{it}. \tag{1}$$

The treatment variable, minimum wage bite mw_{it}, is defined as the ratio of the national minimum wage in year t to the average wage in year $t-1$ in subregion i. LW_i is a dummy variable equal to one if subregion i belongs to the lowest tercile of wage distribution in 2007 (the year before the acceleration of minimum wage hikes). We allow the effects of the minimum wage hikes to differ between the low wage subregions and the rest of the country for two main reasons. First, the subregional differences in wage levels in Poland have been large. In 2007, the coefficient of the variation of average wages across the NUTS 3 subregions amounted to 14.4%, and the ratio of the 90th to the 10th percentile was 1.4. Second, until the policy change went into effect in 2008, the Kaitz index in Poland was low (below 35%). Given that the effects of the minimum wage can be stronger in markets with a larger minimum wage bite (Ahlfeldt et al., 2018; Cengiz et al., 2019), it is possible that minimum wage hikes had a greater impact on subregions with lower average wages before the policy change was implemented. We elaborate on the characteristics of subregions in Section 4.

The vector $control_{it}$ includes key demand- and supply-side factors. The demand-side factors capture aggregate and cyclical fluctuations. They include lagged subregional GDP and labour demand shocks, calculated in line with the Bartik (1991) shift-share method applied to four sectors.[5] The supply-side factors capture changes in the labour supply on local labour markets. They include lagged changes in the domestic labour supply aged 20–64, and in the number of immigrant workers in relation to the population aged 20–64. We control for immigration explicitly, because since 2014, Poland has experienced large inflows of temporary workers, especially from Ukraine, who are not reflected in the population numbers (OECD, 2019).

We also control for subregion fixed effects α_i.[6] Moreover, we account for time-varying macroeconomic trends by using the Bartik instrument for labour demand shocks, and by subtracting the national wage growth from the regional wage growth. The latter also allows us to deal with endogeneity that stems from two features of the data. First, minimum wage changes respond, to some extent, to the overall wage growth. A standard Granger causality test confirms that national real wage growth predicts minimum wage growth, whereas the opposite is not true (see Table A1 in Appendix A). Second, regional wage growth exhibits a strong serial correlation (Table A2 in Appendix A). Wage growth in Poland has been largely driven by the medium-term convergence

[5]For each region, we calculate sectors' shares in employment. Next, we derive employment growth in each sector at the country level, excluding the region of interest. Multiplication of the shares by the growth rates yields the Bartik instrument, which is also expressed as a percentage change. Employment is available in the following four groups of NACE sections: (1) Mining, Manufacturing, Electricity Supply, Utilities, Construction; (2) Trade, Transportation and Storage, Accommodation and Food Service Activities, Information and Communication; (3) Financial and Insurance Activities, Real Estate Activities; and (4) Other Services. The sector structure used reflects data availability at the NUTS 3 level.

[6]We decided against using time fixed effects. Recent developments in econometrics show that the two-way fixed effects estimators often yield unreliable results. De Chaisemartin and D'Haultfœuille (2020) demonstrated that such estimators estimate weighted sums of the treatment effect in each group and period with weights that may be negative. Moreover, Kropko and Kubinec (2020) showed that while a one-way fixed effects estimator is helpful in isolating a particular dimension of variance, a two-way fixed effects estimator provides un-interpretable results. However, as two-way fixed effects estimators are often used in the minimum wage literature, we report results of regressions using them in Appendix B (Tables B2 and B3).

towards the EU average (World Bank Group, 2017). Thus, total wage growth and minimum wage changes are positively correlated with lagged wage growth. Therefore, total wage growth should not be regressed on minimum wage changes and the relevant coefficient cannot be interpreted in a causal sense. We solve this problem by using the demeaned wage growth which is not serially correlated. For transparency, we also report the results with non-transformed regional wage growth as a dependent variable (Table B1 in Appendix B).

We consider three modifications of model (1). First, we estimate a simple model that assumes a uniform effect of the minimum wage across all subregions. Second, we test whether the effects of the minimum wage materialize with a delay, and estimate model (1) with a lagged change in the minimum wage bite, $\Delta mw_{i,j,t-1}$, as an additional regressor. We use only one lag because our time series before the policy change went into effect in 2008 is relatively short. Third, following Meer and West (2016), we use the leading value of the minimum wage hike to validate our research design. If we find significant coefficients pertaining to future minimum wage hikes, the relationship between the dependent variable and minimum wage might be spurious, and could reflect unobserved trends. As in our case $mw_{i,j,t+1}$ contains wages from period t, we instead employ $mw_{i,j,t+2}$, and add it as an additional regressor in Equation 1. In all regressions, standard errors are calculated with the use of the Driscoll and Kraay (1998) estimator that accounts for cross-sectional dependence.[7]

We also perform several robustness checks. We use the Huber–White estimator instead of the Driscoll and Kraay (1998) estimator; we change the definition of low wage subregions to the first quartile of the 2007 distribution of the average NUTS 3 wages; we use unemployment changes as a control variable; and we use growth of real minimum wage instead of changes in minimum wage bite as a treatment variable. We also re-estimate our models on subsamples constructed by omitting NUTS 3 subregions that belong to particular NUTS 2 regions.

Having estimated the models, we assess the economic significance of the minimum wage hikes. We calculate a counterfactual simulation of the evolution of the average wage and employment, while assuming that the ratio of the minimum wage to the average wage has remained constant since 2007; and we compare these simulations with the actual evolution of employment and wages. We also compute the own wage employment elasticity, defined as the semi-elasticity of employment divided by the semi-elasticity of average wage to a given minimum wage hike (Cengiz et al., 2019).

3.2 | Data

Our data cover the period 2003–2018. The Local Data Bank of Statistics Poland (2020) is our main data source. Data on wages and employment are based on non-agricultural entities that employ at least 10 persons (microenterprises do not report the relevant information to Statistics Poland). Employment is measured by the number of employees (neither self-employed individuals nor workers with non-employment contracts, such as civil law contracts, are included in this statistic). As a result, the sample represents, on average, 64% of non-farming employment. There is no information on hours worked, and the employment numbers are not converted into full-time equivalents. Thus, we investigate employment effects at the extensive margin. However, the share of part-time workers in non-farming economy has been at a low level of 7% since 2007 (pre-reform year). Average wages at the NUTS 3 level are based on full-time equivalents. The wage and GDP data are converted to real terms using 2015 prices.

[7]In the case of the wage equation, tests for cross-sectional dependence reject the null hypothesis that error terms are independent across regions.

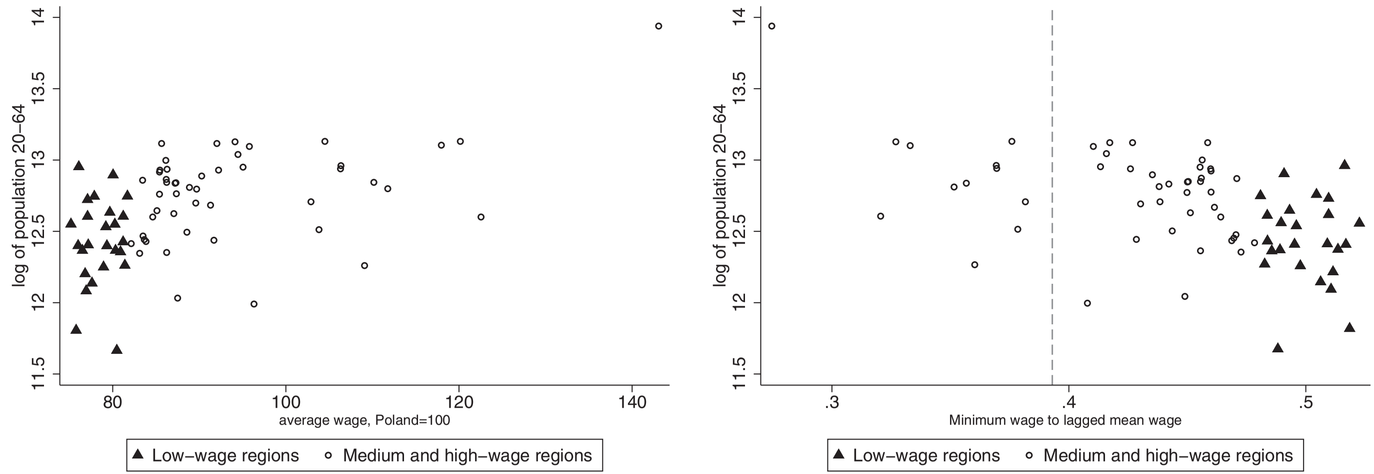


FIGURE 3 Allocation of Polish NUTS 3 subregions into groups based on the 2007 wage level. (a) Average NUTS 3 wages in 2007 (Poland = 100); (b) minimum wage bite in 2008 (2008 minimum wage/2007 NUTS 3 average wage)

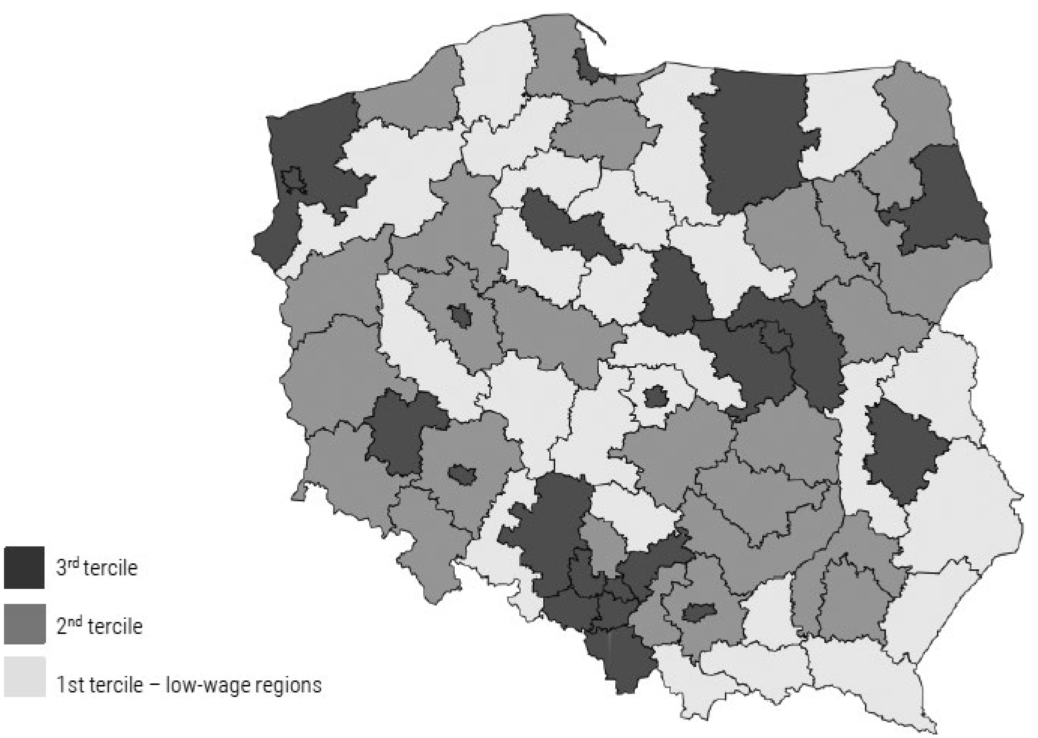
Note: In the right panel, the vertical line denotes the country-average minimum wage bite. Source: Own calculations based on Statistics Poland Local Data Bank data.

To measure immigration inflows, we use data on work permits for foreigners obtained from the Ministry of Family, Labour and Social Policy, which is a common way to measure the spatial allocation of migrants in Poland (Górny & Śleszyński, 2019). These data are available from 2008 onwards, and we assume that in 2002–2008, the numbers of work permits were at the 2008 level. In doing so, we take advantage of the fact that the number of immigrants in Poland was very low until 2015, and has grown rapidly since then (White et al., 2018). The median proportion of temporary immigrant workers to the working-age population in a NUTS 3 subregion was 0.02% in 2008, but exceeded 1% by 2018.

4 | DESCRIPTIVE EVIDENCE

Here, we provide descriptive evidence showing that the low wage subregions and the rest of the country exhibited common trends in the minimum wage bite, employment and wage trends before 2008, but that these trends diverged from 2008 onwards; that is, after the minimum wage hikes accelerated. Thus, our estimate of a continuous treatment variable can be interpreted in terms of the difference-in-difference estimator.

We define the low wage subregions as those in the first tercile of the average NUTS 3 wage distribution in 2007. In the low wage subregions, the average wages in 2007 were below 82% of the national average; and the 2008 minimum wage bite was above 48%, or 8.8 pp above the country-level bite (Figure 3). The low wage subregions were spread across the country (Map 1), and together accounted



MAP 1 Terciles of the 2007 NUTS 3 average wage distribution in Poland
SOURCE: Own elaboration based on the Geostatistics Portal, Statistics Poland.

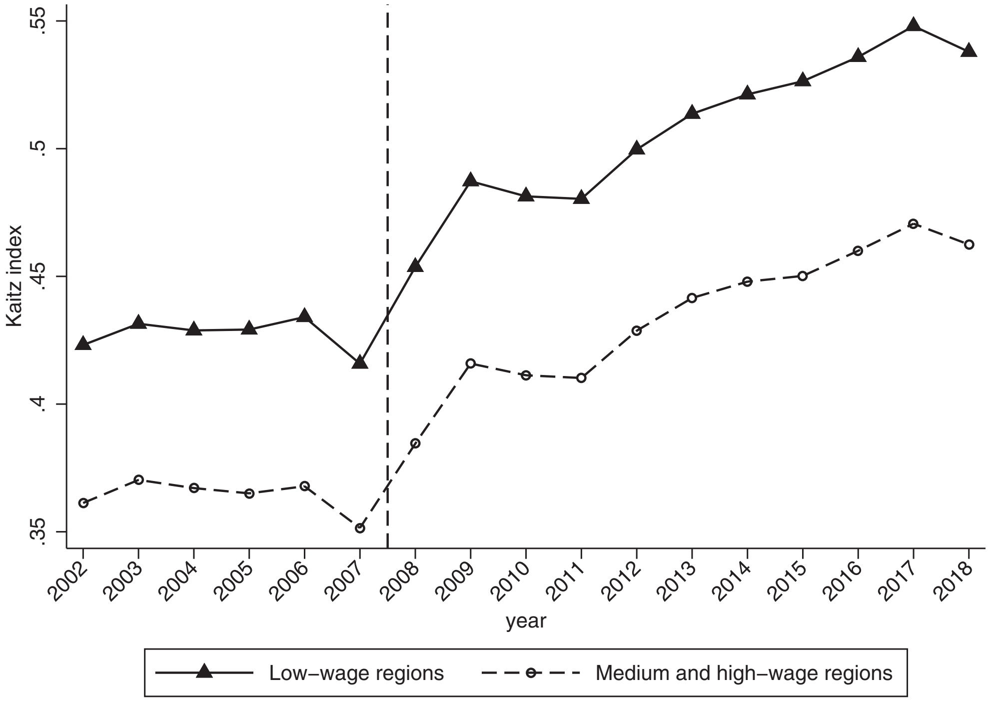


FIGURE 4 Minimum wage bite by groups of NUTS 3 subregions in Poland, 2002–2018
SOURCE: Own calculations based on Statistics Poland Local Data Bank data.

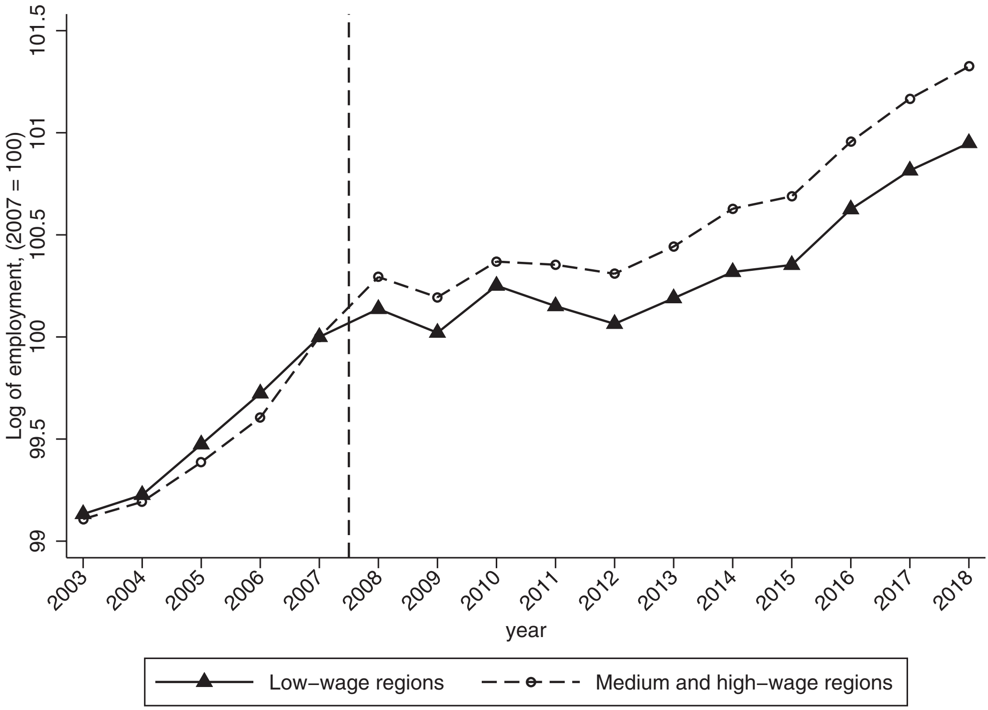


FIGURE 5 Average employment dynamics by groups of NUTS 3 subregions in Poland, 2003–2018 (2007 = 100)
SOURCE: Own calculations based on Statistics Poland Local Data Bank data.

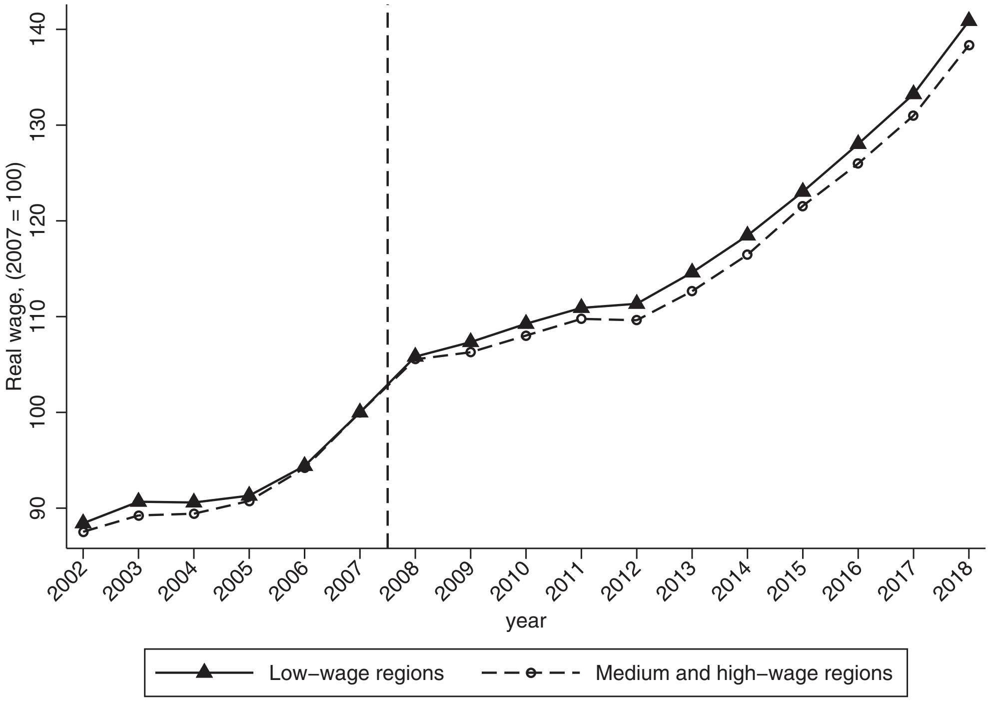


FIGURE 6 Average real wage dynamics by groups of NUTS 3 subregions in Poland, 2002–2018 (2007 = 100)
SOURCE: Own calculations based on Statistics Poland Local Data Bank data.

for 21.2% of employment in 2007. The complete list of NUTS 3 subregions by terciles of the 2007 wage distribution is presented in Table A3 in Appendix A.

Before 2008, the subregional Kaitz indices (the relation of the minimum wage to the subregional mean wage) changed little in both the low wage subregions and the medium and high wage subregions. In 2007, the Kaitz index in the low wage subregions was, on average, 6.4 pp higher than it was in the rest of the country. Since 2008, the Kaitz indices have been increasing in both groups (Figure 4). The employment and wage dynamics were virtually identical in the two groups of subregions in 2002–2007, but have been diverging since 2008. In 2008–2018, average employment growth was 0.9% in the low wage subregions, and was 1.3% in the rest of the country (Figure 5). At the same time, the subregions that initially (2007) had lower average wages experienced stronger growth starting in 2008, with the difference in cumulated wage growth over the 2008–2018 period amounting to 2.5 pp (Figure 6).

Overall, the descriptive evidence suggests that the observed convergence in wages and divergence in employment growth might have been related to the acceleration of minimum wage hikes since 2008. We test this conjecture formally in the next section.

5 | ECONOMETRIC RESULTS

5.1 | Effects of minimum wage hikes on employment and wages

We find that the minimum wage hikes in Poland had a delayed effect on average wage growth and no overall effects on employment in the pooled sample of 73 subregions in the 2004–2018 period. According to our results, the contemporary minimum wage hikes had no effect on demeaned wage growth (column

TABLE 1 Effects of minimum wage hikes on subregional wage growth in Poland

	(1)	(2)	(3)	(4)	(5)
ΔMW	0.003	0.015	−0.034	−0.031	−0.021
	(0.042)	(0.044)	(0.032)	(0.030)	(0.032)
ΔMW × (low wage subregion)			0.093	0.107*	0.090
			(0.059)	(0.058)	(0.089)
ΔMW ($t-1$)		0.099*		0.034	0.038
		(0.047)		(0.036)	(0.032)
ΔMW ($t-1$) × (low wage subregion)				0.153***	0.148***
				(0.047)	(0.048)
ΔMW ($t+2$)		0.021			0.053
		(0.051)			(0.059)
ΔMW ($t+2$) × (low wage subregion)					−0.084
					(0.077)
Observations	1,095	1,022	1,095	1,095	1,022
Within R^2	0.003	0.012	0.005	0.021	0.024
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0			0.435	0.292	0.472
p-val. Coeff. on ΔMW ($t+2$) + Coeff. on ΔMW ($t+2$) × (low wage subregion) = 0					0.644

Notes: The dependent variable is the subregional growth rate of wages decreased by the national growth rate. The MW is defined as the ratio of the national minimum wage to the previous year's average wage in a given subregion. All regressions use NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks. Driscoll–Kraay standard errors in parentheses. We also report the *p*-values for two tests: (i) that the effect of contemporary minimum wage hikes in low wage regions is zero, (ii) that the effect of future minimum wage hikes in low wage regions is zero.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

1 of Table 1), which suggests that there was no wage compression across subregions in the overall sample.[8] The estimated, contemporary effect on employment is negative, albeit small and insignificant (column 1 of Table 2). However, when we account for delayed effects of minimum wage hikes, we find a noticeable effect on wage growth (column 2 of Table 1), but still no effect on overall employment (column 2 of Table 2).

Importantly, we find that the effects of minimum wage hikes on wage changes and employment were much greater in the low wage subregions. These effects are shown by the estimated interaction terms (column 3 of Tables 1 and 2) pertaining to the minimum wage hikes in the subregions that belonged to the lowest tercile of the wage distribution in 2007 (before the change in minimum wage policy). Moreover, the employment effect in the medium and high wage regions is no longer negative, which suggests that the small average effect (column 1 of Table 2) was driven by developments in the low wage subregions. According to the semi-elasticities shown in column (3) of Tables 1 and 2, the increase in the minimum wage bite of 10 pp in the low wage subregions was associated with a 1.9% decline in employment and a 0.6-pp increase in wage growth (though this effect is not statistically

[8]Low-wage regions are characterised by slightly higher growth in the minimum wage bite due to the lower denominator.

TABLE 2 Effects of minimum wage hikes on subregional employment growth in Poland

	(1)	(2)	(3)	(4)	(5)
ΔMW	−0.051	−0.047	0.036	0.038	0.042
	(0.031)	(0.027)	(0.021)	(0.023)	(0.027)
ΔMW × (low wage subregion)			−0.225***	−0.219***	−0.242***
			(0.072)	(0.069)	(0.063)
ΔMW ($t-1$)		0.049		0.024	0.017
		(0.061)		(0.043)	(0.049)
ΔMW ($t-1$) × (low wage subregion)				0.061	0.085
				(0.097)	(0.084)
ΔMW ($t+2$)		0.015			0.017
		(0.031)			(0.027)
ΔMW ($t+2$) × (low wage subregion)					0.012
					(0.085)
Observations	1,095	1,022	1,095	1,095	1,022
Within R^2	0.327	0.335	0.330	0.331	0.341
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0			0.013	0.014	0.003
p-val. Coeff. on ΔMW ($t+2$) + Coeff. on ΔMW ($t+2$) × (low wage subregion) = 0					0.718

Notes: The dependent variable is change in the log of employment in a NUTS 3 subregion. The MW is defined as the ratio of the national minimum wage to the previous year's average wage in a given subregion. All regressions use the NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks; Driscoll–Kraay standard errors in parentheses. We also report the p-values for two tests: (i) that the effect of contemporary minimum wage hikes in low wage regions is zero, (ii) that the effect of future minimum wage hikes in low wage regions is zero.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

significant). The discussed size of the shock is comparable to the actual increase in the minimum wage bite in the low wage subregions between 2007 and 2017 (12.4 pp, on average).

Next, we consider the specification that includes lagged minimum wage hikes and interaction terms with low wage subregion fixed effects. We find a significant and sizeable delayed effect on average wage growth in the low wage subregions (column 4 in Table 1). This finding suggests that the wage effects materialized gradually, possibly because of spillover effects through the wage distribution. At the same time, the results show that the lagged effect on employment was not significant, which suggests that the adjustment of the employment level occurred almost immediately (column 4 in Table 2). Accounting for the lagged minimum wage hikes confirms our conclusion that the minimum wage hikes had no effects in the medium and high wage regions. The resulting own wage employment elasticity in low wage subregions amounts to −0.37, which is a bit larger in absolute terms than the median value of −0.17 reported by Dube (2019) based on a review of the international evidence. However, the elasticity calculated for the whole country (using imprecise estimates for medium and high wage regions) is positive and amounts to 0.43.

A causal interpretation of our findings is supported by the results of the Meer and West (2016) falsification test. The leading values of the minimum wage hikes, both pooled and interacted with low wage subregion fixed effects, are insignificant in both models (columns 2 and 5 in Tables 1 and 2).[9] This result confirms our assumption that the estimated effects of minimum wage hikes were not driven by any unobserved trends.

5.2 | Economic significance of minimum wage increases

To evaluate the economic significance of minimum wage hikes in Poland, we assess the annual and the cumulated economic effects of minimum wage hikes on subregional labour markets using the models that account for lagged minimum wage hikes (column 4 of Tables 1 and 2). We calculate the effects of successive minimum wage hikes on annual employment and wage growth (Figure 7), and perform counterfactual simulations of employment and wage levels under the assumption that the ratio of the minimum to the average wage (Kaitz index) has remained constant since 2007 (Figure 8).

We find that the minimum wage hikes in 2008 and 2016–2017 had the most pronounced effects. In the low wage subregions, about 1 pp of the wage growth in 2009, and about 0.4 pp in 2017–2018 can be attributed to minimum wage increases. In 2008 and 2013–2014, the effects of the minimum wage hikes on wage growth were also positive and quite sizeable (Figure 7). At the same time, the substantial minimum wage hike in 2008 reduced the level of employment growth in low wage subregions by about 0.8 pp, and the hikes in 2016–2017 reduced it by about 0.2–0.3 pp. On the other hand, in 2010 and in 2018, the minimum wage hikes were so small that the minimum wage bite actually declined, and the effects on employment in low wage subregions were positive. The coincidence of the positive effects on wages and employment in low wage subregions in 2018 was due to the delayed effects of 2017 hike on wage growth. In contrast, there were no significant effects on employment or on wage growth in the medium and high wage subregions.

Our counterfactual simulations show that the acceleration of minimum hikes in Poland from 2008 onwards resulted in noticeably higher wages and somewhat lower employment in the low wage subregions (Figure 8). If the minimum-to-average wage ratio had remained constant since 2007, in 2018, the average wage in the low wage subregions would have been 3.2% lower (by 123 zloty per month and 1,475 zloty per year), while employment would have been 1.2% higher (23,000 additional jobs in the population of firms employing at least 10 workers). While the cumulated employment effect was greater in 2017, and amounted to 1.5% of employment (28,000 jobs), a slight decrease in the minimum wage bite in the low wage subregions reduced this effect in 2018.

In the medium and high wage subregions, the minimum wage hikes had no effect on average wages, and had a positive effect on employment (amounting to 51,000 additional jobs, equivalent to 0.7% of total employment). However, employment effects for medium and high wage subregions should be interpreted as insignificant, because they are based on insignificant coefficients and annual effects that are not statistically different from zero (Figure 7).

5.3 | Heterogeneity of employment effects

Next, we examine the heterogeneity of the employment effects of minimum wage hikes by exploring how these hikes affected workers in industry and in services, and men and women. Studying heterogeneous

[9]This specification requires excluding data from 2018 (used to calculate lead values) from the sample. We use $mw_{i,j,t+2}$, because $mw_{i,j,t+1}$ uses wages from period t. However, we verified that coefficients on $mw_{i,j,t+1}$ are insignificant in the employment regression.

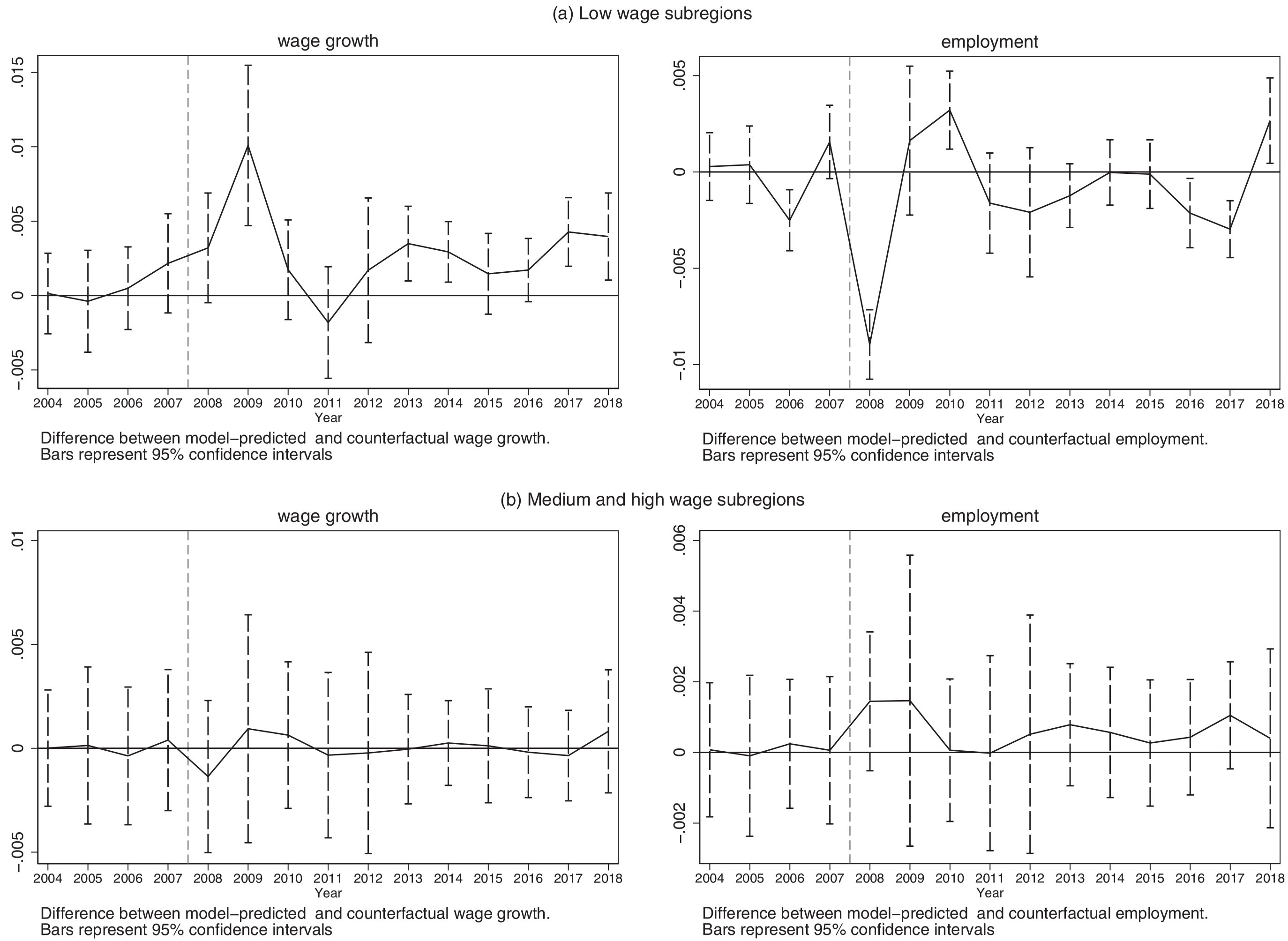


FIGURE 7 The effects of minimum wage hikes on wage changes (left panel) and employment changes (right panel). (a) Low wage subregions. (b) Medium and high wage subregions

SOURCE: Own estimations based on models presented in column (4) of Tables 1 and 2, and on Statistics Poland Local Data Bank data.

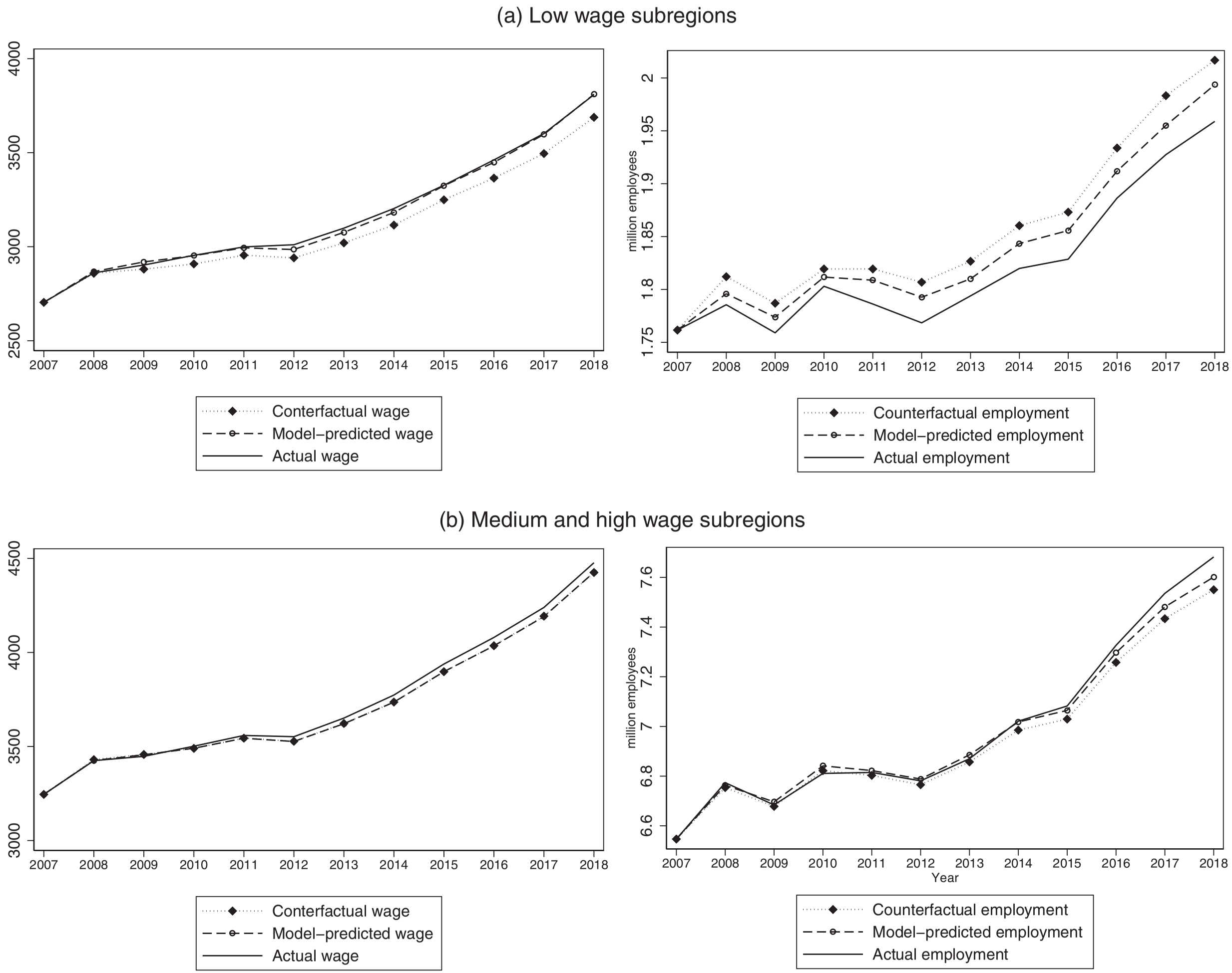


FIGURE 8 Counterfactual scenario of wages (left panel) and employment (right panel) in Poland under the assumption of no changes in the Kaitz index after 2007. (a) Low wage subregions. (b) Medium and high wage subregions

SOURCE: Own estimations based on models presented in column (4) of Tables 1 and 2, and on Statistics Poland Local Data Bank data.

TABLE 3 Employment effects of minimum wage hikes by economic sectors

	Industry		Services	
ΔMW	−0.146*	−0.199**	0.223***	0.250***
	(0.073)	(0.084)	(0.068)	(0.076)
ΔMW × (low wage subregion)	−0.260*	−0.363**	−0.204**	−0.172**
	(0.132)	(0.139)	(0.073)	(0.072)
ΔMW ($t-1$)	−0.089	−0.07	0.100*	0.096
	(0.156)	(0.167)	(0.051)	(0.071)
ΔMW ($t-1$) × (low wage subregion)	−0.129	−0.082	0.232***	0.240***
	(0.336)	(0.309)	(0.068)	(0.075)
ΔMW ($t+2$)		−0.097		0.072
		(0.081)		(0.081)
ΔMW ($t+2$) × (low wage subregion)		−0.083		0.084
		(0.137)		(0.076)
Observations	1,095	1,022	1,095	1,022
Within R^2	0.383	0.397	0.158	0.164
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0	0.046	0.015	0.876	0.538
p-val. Coeff. on ΔMW ($t+2$) + Coeff. on ΔMW ($t+2$) × (low wage subregion) = 0		0.322		0.229

Notes: The dependent variable is change in the log of employment in the industry sector/the services sector in a NUTS 3 subregion. All regressions are based on a specification reported in column (3) of Table 1. However, in columns (1) and (2), we replace the Bartik instrument with growth of industry employment in the rest of the country; and in columns (3) and (4), with analogous growth of services employment. We also report the p-values for two tests: (i) that the effect of contemporary minimum wage hikes in low wage regions is zero, (ii) that the effect of future minimum wage hikes in low wage regions is zero.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

effects on wages is impossible due to the lack of disaggregated wage data. Regression results are presented in Tables 3 and 4, and the cumulated employment effects are shown in Figures A1–A2 in Appendix A.

We find suggestive evidence of differences between particular subgroups of workers. The employment effects among men and among workers in industry are found to be negative in general, and even more so in the low wage subregions. At the same time, the employment effects among workers in services and among women are shown to be positive. The negative effects observed in industry may be related to the fact that the substitution of labour with capital and technology is easier in industry than in services (Alvarez-Cuadrado et al., 2017). The positive effects found among women suggest that minimum wage hikes may have incentivized women—and especially low-skilled women—to increase their labour supply and take up employment, in line with the supply effects described by Card and Krueger (1994).[10] However, the test of the leading minimum wage indicates that the results on women's employment may be driven by unobserved trends; that is, factors other than the minimum wage. Thus, our findings on the gender-specific minimum wage effects in Poland should be interpreted as indicative and taken with

[10]In the studied period, reforms aimed at increasing the labour force participation of older workers, in particular of women aged 55+, were introduced in Poland (Lewandowski and Rutkowski, 2017). We account for the potential effects of these reforms by controlling for changes in the labour force participation of women.

TABLE 4 Employment effects of minimum wage hikes by gender

	Women		Men	
ΔMW	0.132	0.157**	−0.045	−0.069*
	(0.079)	(0.070)	(0.044)	(0.036)
ΔMW × (low wage subregion)	−0.164	−0.079	−0.284***	−0.341***
	(0.115)	(0.109)	(0.095)	(0.096)
ΔMW (t − 1)	0.294***	0.310***	−0.195**	−0.225***
	(0.083)	(0.089)	(0.073)	(0.074)
ΔMW (t − 1) × (low wage subregion)	−0.094	−0.049	0.119	0.14
	(0.078)	(0.096)	(0.192)	(0.192)
ΔMW (t + 2)		0.052		−0.036
		(0.063)		(0.045)
ΔMW (t + 2) × (low wage subregion)		0.258*		−0.048
		(0.121)		(0.112)
Observations	1,095	1,022	1,095	1,022
Within R^2	0.195	0.208	0.319	0.331
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0	0.439	0.118	0.124	0.055
p-val. Coeff. on ΔMW (t + 2) + Coeff. on ΔMW (t + 2) × (low wage subregion) = 0		0.012		0.509

Notes: The dependent variable is change in the log of employment for women/men. All regressions are based on a specification reported in column (3) of Table 1. However, in columns (1) and (2), we replace growth of the active population aged 20–64 with growth of active women aged 20–64; and in columns (3) and (4), analogous growth for men is used. We also report the p-values for two tests: (i) that the effect of contemporary minimum wage hikes in low wage regions is zero, (ii) that the effect of future minimum wage hikes in low wage regions is zero.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

caution. Nevertheless, combined with evidence provided by Majchrowska and Strawiński (2018) that minimum wage hikes have reduced gender pay gap among young workers in Poland, it appears that minimum wage hikes have contributed to improving labour market outcomes of women (Map 1).

5.4 | Robustness analysis

Here, we present a range of robustness checks of our baseline regression (column 4 of Tables 1 and 2). First, we use the Huber–White estimator (instead of the Driscoll and Kraay [1998] estimator) to calculate standard errors (column 1 of Tables 5 and 6). Second, we report results from a regression weighted by subregional employment level as of 2007 (column 2). Third, we add lagged unemployment changes as a control variable (column 3). Fourth, we define the low wage subregions more strictly; namely, as the subregions that belong to the first quartile of the 2007 distribution of average NUTS 3 wages (column 4). Finally, we use alternative explanatory variables. We replace changes in minimum wage bite with the growth of log of real minimum wage (column 5).

Our conclusions are robust to all these modifications. The coefficients pertaining to wage effects are close to those estimated in the baseline specification. The effects in low wage subregions are

TABLE 5 Robustness analysis of wage effects

	(1)	(2)	(3)	(4)	(5)
	Huber–White SE	**Employment weight**	**Added Δunemp ($t-1$)**	**Low wage subregions: quartile**	**Alternative treatment variable**
ΔMW	−0.031	−0.041	0.012	−0.015	−0.023
	(0.043)	(0.039)	(0.030)	(0.031)	(0.018)
MW × (low wage subregion)	0.107*	0.112**	0.098	0.088	0.048***
	(0.058)	(0.054)	(0.066)	(0.071)	(0.009)
ΔMW ($t-1$)	0.034	−0.029	0.049	0.058	0.021
	(0.038)	(0.056)	(0.036)	(0.034)	(0.015)
ΔMW ($t-1$) × (low wage subregion)	0.153***	0.212***	0.151***	0.123*	0.051**
	(0.056)	(0.066)	(0.050)	(0.068)	(0.017)
Observations	1,095	1,095	1,095	1,095	1,095
Within R^2	0.021	0.012	0.031	0.017	0.024
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0	0.069	0.052	0.073	0.388	0.372

Notes: The dependent variable is the subregional growth rate of wages decreased by the national growth rate. All regressions use the NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks; The first column reports the baseline specification with the robust standard errors estimated by the Huber–White estimator. In the second column the regression is weighted by the number of employees in 2007. The third column adds lagged unemployment growth to the baseline specification. The fourth column uses a more strict definition of low wage subregions. The fifth column replaces increase in minimum wage bite with the growth of log of real minimum wage as the MW variable.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

TABLE 6 Robustness analysis of employment effects

	(1)	(2)	(3)	(4)	(5)
	Huber–White SE	**Employment weight**	**Added Δunemp ($t-1$)**	**Low wage subregions: quartile**	**Alternative treatment variable**
ΔMW	0.038	0.029	0.044	0.021	0.017**
	(0.055)	(0.058)	(0.033)	(0.031)	(0.007)
MW × (low wage subregion)	−0.219**	−0.225**	−0.220***	−0.229**	−0.090***
	(0.087)	(0.091)	(0.068)	(0.099)	(0.022)
ΔMW ($t-1$)	0.024	−0.067	0.026	0.035	0.002
	(0.071)	(0.106)	(0.046)	(0.037)	(0.010)
ΔMW ($t-1$) × (low wage subregion)	0.061	0.092	0.061	0.045	0.041
	(0.105)	(0.116)	(0.097)	(0.091)	(0.047)
Observations	1,095	1,095	1,095	1,095	1,095
Within R^2	0.331	0.337	0.331	0.331	0.334
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0	0.007	0.006	0.030	0.030	0.002

Notes: The dependent variable is change in the log of employment in a NUTS 3 subregion. All regressions use the NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks; The first column reports the baseline specification with the robust standard errors estimated by the Huber–White estimator. In the second column the regression is weighted by the number of employees in 2007. The third column adds lagged unemployment growth to the baseline specification. The fourth column uses a more strict definition of low wage subregions. The fifth column replaces increase in minimum wage bite with the growth of log of real minimum wage as the MW variable.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

significant at the 1% level (the lagged wage effect). The only exception is the variant in which the low wage subregions are defined on the basis of the first quartile of the 2007 average wage distribution (column 4 of Table 5): the effect is somewhat smaller in size and is significant at the 10% level, whereas the effect in medium and high wage regions is somewhat stronger than in the baseline specification (column 4 of Table 2). This pattern suggests that the effects in the subregions that belong to the bottom tercile, but do not belong to the bottom quartile of the 2007 average wage distribution, are significant. Thus, our baseline definition of the low wage subregions as those in the first tercile seems to be better suited to capturing which subregions are particularly affected by minimum wage hikes in Poland.

In all specifications, the estimated employment effects in low wage subregions are significant at the 1% or the 5% level, while the estimated employment effects in the medium and high wage subregions are small and, except for column (5), insignificant.

Our next robustness check is to re-estimate the baseline regressions on 16 subsamples that are created by excluding NUTS 3 subregions that belong to particular voivodeships (16 NUTS 2 regions). The main findings are confirmed in all cases: in the low wage subregions, the employment effects are negative, whereas the lagged wage effects are positive and significant at a level of 5% or lower. The small positive employment effect in medium and high wage subregions becomes significant at the 5% level in two of 16 regressions, whereas in 12 regressions it is not significant at the 10% level. These results are available upon request.

Furthermore, in Appendix B we show that our findings are also robust to adding time fixed effects or to interacting all exogenous variables with low wage regions dummy.

6 | SUMMARY AND CONCLUSIONS

We have studied the employment and wage effects of minimum wage hikes in Poland, a country that has introduced some of the largest increases in the minimum wage level in the EU since the 2008 global economic crisis. To examine the effects of these minimum wage hikes, we utilized the large variation in wage levels and in the minimum wage bite across 73 Polish NUTS 3 subregions. Controlling for a range of labour demand and labour supply factors, we found that the minimum wage hikes had a significant positive effect on wage growth and a significant negative effect on employment growth, but only in the subregions that belonged to the first tercile of the subregional wage distribution in 2007. These effects were found to be moderate in size, and to be more relevant for wages: if there had been no minimum wage hikes after 2007, by 2018, the average wage in these subregions would have been 3.2% lower, whereas employment would have been 1.2% higher. On the other hand, we found no significant effects on average wages or on employment in subregions that had medium or high wage levels in 2007. We also found indicative evidence that the effects on employment growth differ between groups of workers: the effects were negative for men and for workers in industry, but were positive for women and for workers in services.

Our study has limitations. Due to a lack of data, we could not account for workers hired under civil law contracts; that is, precarious contracts not covered by the labour code or by minimum wage laws (until October 2016). The number of such workers has been increasing between the early 2000s and the mid-2010s, and these workers were more likely to be paid less than the minimum wage (Goraus-Tańska & Lewandowski, 2019). However, if we had been able to account for these workers, it is likely that the estimated effects would have been smaller in absolute terms. This result would have reinforced our finding that the impact of these workers was insignificant in the medium and high wage subregions. In addition, due to a lack of data, we were only able to analyse employment and wage growth in firms that employed at least 10 workers. We were also unable to account for spillovers between particular subregions or the role of cross-regional commuting, as data on input–output connections or commuting in Poland are not available at such a highly disaggregated spatial level.

Our findings have some important policy implications. First, although the minimum wage has increased substantially in Poland, the effects of this development have been benign. Our findings show that minimum wage hikes have compressed inter-regional wage inequality by accelerating wage growth in the least developed subregions, and that the associated employment losses have been very small. However, there are two caveats when projecting our findings into the future. First, as the minimum-to-average wage ratio has been increased over time, the current elasticities may be larger than the average elasticities that we have estimated. In 2018, in the medium and high wage subregions, the Kaitz index reached the level recorded in the low wage subregions in 2008 (45%), and in the low wage subregions it reached 54%, surpassing the nationwide target set by legislation. This means that the trade-off between wage growth and employment losses may become less beneficial over time if the minimum wage becomes 'too high' (Dube, 2019; Manning, 2021a). However, our analysis does not allow us to determine the optimal minimum wage level in Poland. Second, between 2003 and 2018, economic growth in Poland has been rapid, which has led to strong labour demand (Lewandowski & Magda, 2018; Piątkowski, 2018). An economic slowdown, and especially a recession triggered by the COVID-19 pandemic, may lead to reduced labour demand, which could, in turn, increase the risk of job losses related to minimum wage hikes. The authorities should take such uncertainties into account when setting the minimum wage level. Third, our results also show that in a country with large spatial differences in labour markets, such as Poland, it may be hard to achieve policy goals by applying a nationwide minimum wage level. However, setting a minimum wage at the level of NUTS 2 regions—which in Poland are administrative units (voivodeships)—would not solve this conundrum, and thus should not be pursued. Twelve of the 16 NUTS 2 regions in Poland include both a low wage and a high wage subregion. Therefore, we think that the setting of the nationwide minimum wage should prioritize impacts on the less-developed subregions. In the largest cities with the highest wages, it could be complemented by promoting the living-wage approach and encouraging collective bargaining to set higher levels of wages which account for differences in cost of living.

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APPENDIX A

Additional tables and figures

TABLE A1 Granger causality tests

	Minimum wage	Average wage
Minimum wage ($t-1$)	−0.239	−0.169
	(0.232)	(0.129)
Minimum wage ($t-2$)	−0.252	0.037
	(0.215)	(0.119)
Average wage ($t-1$)	1.255***	0.723***
	(0.471)	(0.262)
Average wage ($t-2$)	0.553	−0.151
	(0.601)	(0.335)
Constant	0.023	0.021**
	(0.016)	(0.009)
Observations	14	14
p-val. Coeff. on Minimum wage ($t-1$) = Coeff. on Minimum wage ($t-2$) = 0		0.385
p-val. Coeff. on Average wage ($t-1$) = Coeff. on Average wage ($t-2$) = 0	0.003	

Note: Table reports vector autoregressive models for growth rate of real minimum wage and growth rate of real national average wage. Standard errors in parentheses.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland data.

TABLE A2 Stationarity and autocorrelation tests of dependent variables

variable	Levin–Lin–Chu unit-root test	Harris–Tzavalis unit-root test	Wooldridge test for autocorrelation	Bias-corrected Born & Breitung test for serial correlation
$\ln(employment_i)$	0.000	1.000	0.000	0.000
$\ln(employment_i)$	0.000	0.000	0.342	0.471
$\ln(wage_i)$	0.998	1.000	0.000	0.000
$\ln(wage_i)$	0.000	0.000	0.000	0.000
$\ln(wage_i) - \ln(wage_{\text{Poland}})$	0.000	0.000	0.181	0.414

Note: Table reports p-values. In the first two columns, the null hypothesis is that panels contain unit root. In column (3), the null hypothesis states that there is no first-order autocorrelation. In column (4), the null hypothesis is that there is no serial correlation up to order $p = 2$.

TABLE A3 List of NUTS 3 subregions by groups distinguished by their 2007 average wage levels

Low wage subregions (first tercile of wage distribution)		Medium wage subregions (second tercile of wage distribution)		High wage subregions (third tercile of wage distribution)	
Subregion	**Mean wage 2007**	**Subregion**	**Mean wage 2007**	**Subregion**	**Mean wage 2007**
Sieradzki	2,155	Łomżyński	2,354	Szczeciński	2,568
Chojnicki	2,172	Koszaliński	2,382	Warszawski wschodni	2,570
Grudziądzki	2,178	Tarnobrzeski	2,392	Olsztyński	2,587
Kaliski	2,181	Pilski	2,394	Białostocki	2,617
Inowrocławski	2,192	Siedlecki	2,397	Tyski	2,628
Nowotarski	2,201	Gorzowski	2,402	Bydgosko-toruński	2,638
Ełcki	2,205	Bytomski	2,426	Bielski	2,643
Elbląski	2,209	Gdański	2,440	Łódź	2,698
Krośnieński	2,210	Wrocławski	2,448	Lubelski	2,708
Łódzki	2,212	Zielonogórski	2,448	Opolski	2,724
Bialski	2,224	Koniński	2,449	Sosnowiecki	2,745
Leszczyński	2,232	Kielecki	2,455	Żyrardowski	2,761
Ciechanowski	2,262	Wałbrzyski	2,468	Gliwicki	2,950
szczecinecko-Pyrzycki	2,270	Radomski	2,470	Szczecin	2,976
Przemyski	2,274	Piotrkowski	2,471	Kraków	2,995
Nowosądecki	2,283	Ostrołęcki	2,471	Rybnicki	3,047
chełmsko-Zamojski	2,294	Krakowski	2,473	Wrocław	3,049
Tarnowski	2,300	Sandomiersko-jędrzejowski	2,495	Płocki	3,126
Skierniewicki	2,302	Rzeszowski	2,500	Poznań	3,157
Świecki	2,306	Jeleniogórski	2,502	Warszawski zachodni	3,203
Włocławski	2,319	Oświęcimski	2,503	Trójmiejski	3,380
Nyski	2,327	Suwalski	2,508	Katowicki	3,444
Puławski	2,328	Starogardzki	2,538	Legnicko-głogowski	3,513
Słupski	2,333	Poznański	2,546	Warszawa	4,100
Częstochowski	2,342				
Average	2,252	Average	2,456	Average	2,951

Source: Own elaboration based on Statistics Poland Local Data Bank data.

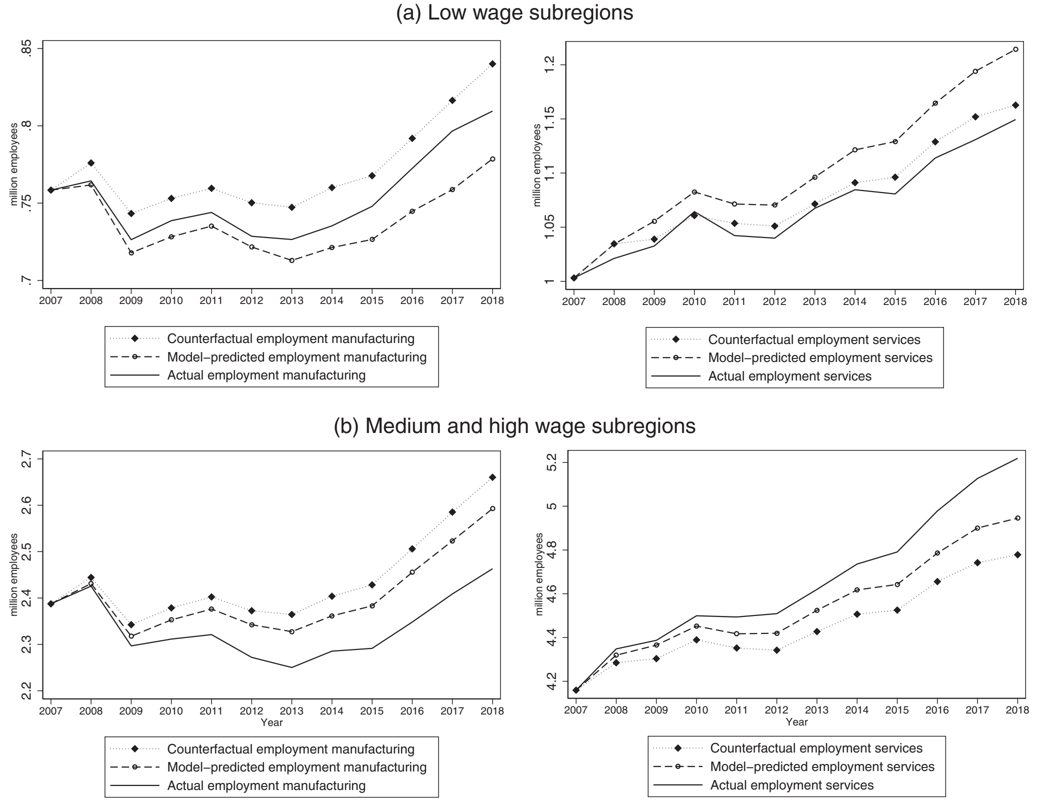

FIGURE A1 Counterfactual scenario of employment in manufacturing (left panel) and in services (right panel) in Poland under the assumption of no changes in the Kaitz index after 2007. (a) Low wage subregions. (b) Medium and high wage subregions

SOURCE: Own estimations based on models presented in Table 3, and on Statistics Poland Local Data Bank data.

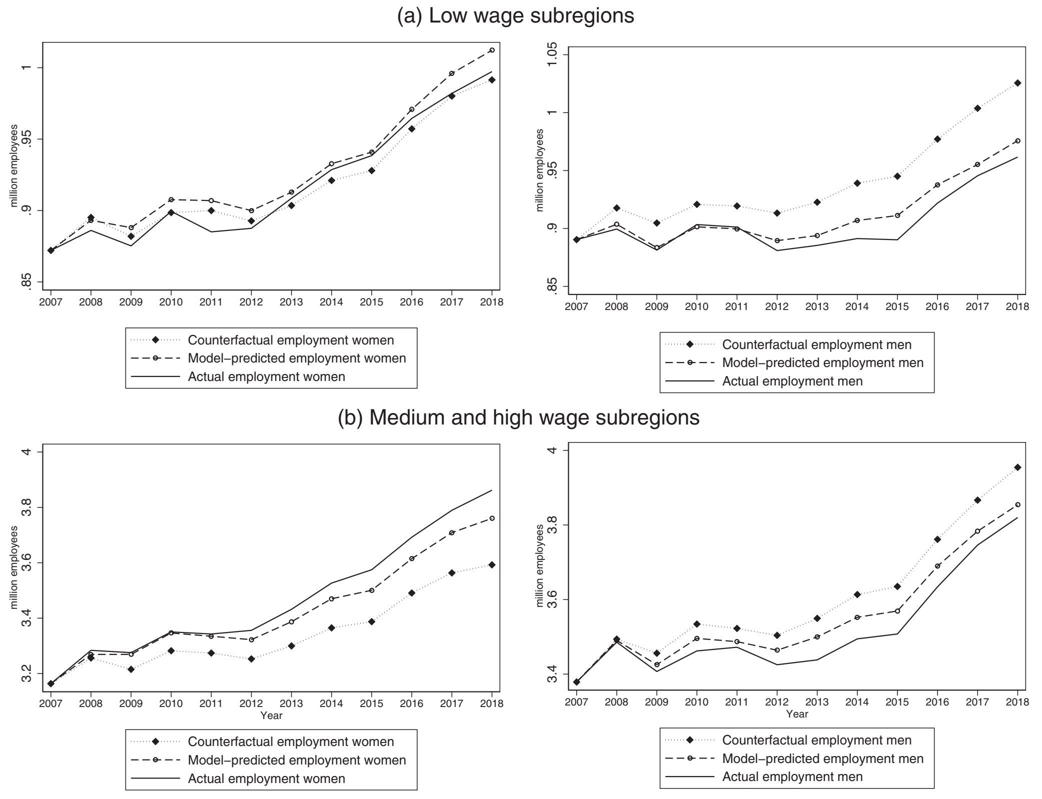

FIGURE A2 Counterfactual scenario of female employment (left panel) and male employment (right panel) in Poland under the assumption of no changes in the Kaitz index after 2007. (a) Low wage subregions. (b) Medium and high wage subregions

SOURCE: Own estimations based on models presented in Table 4, and on Statistics Poland Local Data Bank data.

TABLE A4 Complete estimation results of baseline specifications for employment and wages

	Employment	Wages
ΔMW	0.038	−0.031
	(0.023)	(0.030)
ΔMW × (low wage subregion)	−0.219***	0.107*
	(0.069)	(0.058)
ΔMW (t − 1)	0.024	0.034
	(0.043)	(0.036)
ΔMW (t − 1) × (low wage subregion)	0.061	0.153***
	(0.097)	(0.047)
Bartik instrument	0.987***	0.031
	(0.040)	(0.064)
ΔLN(GDP) (t − 1)	0.023	−0.007
	(0.014)	(0.014)
ΔLN(Population active on labour market) (t − 1)	0.03	0.039
	(0.029)	(0.039)
Δ(Inflow of immigrants/Population aged 20–64) (t − 1)	−0.364*	0.154
	(0.196)	(0.125)
Constant	−0.002*	(0.014)
	(0.001)	−0.001
Observations	1,095	1,095
Within R^2	0.331	0.021

Notes: The dependent variables are change in the log of employment in a NUTS 3 subregion and the subregional growth rate of wages decreased by the national growth rate. The MW is defined as the ratio of the national minimum wage to the previous year's average wage in a given subregion. Both regressions use NUTS 3 fixed effects. Driscoll–Kraay standard errors in parentheses.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

APPENDIX B

Alternative specifications

In our baseline specification, we investigate wage effects using demeaned wage growth. Now we report wage effects from a specification that uses non-transformed regional wage growth as a dependent variable (Table B1). The overall wage effect is stronger than for the baseline dependent variable. However, as discussed in Section 3.1, this specification is affected by the correlation between minimum wage hikes and lagged wage growth. Significant coefficients of leading minimum wage hikes confirm that these results cannot be interpreted in causal terms.

Next, we report results of two modifications to our baseline specification. First, we use specification in which all exogenous variables are included both standalone and interacted with the dummy indicating low wage regions (columns 1 and 2 of Tables B2 and B3). Second, we apply two-way fixed effects models by adding year effects (columns 3 and 4 of Tables B2 and B3). In the fully interacted model, findings for both employment and wage effects are virtually unchanged, with only slight changes in the dynamics of wage effects. Employment effects also do not change in the two-way fixed effects, where the only noticeable difference is larger standard errors pertaining to the changes in minimum wage bite in medium and high wage regions. After adding the time fixed effects, the overall size of the wage effects becomes larger. However, the estimated difference in the wage effects between low wage regions and the rest of the country remains as large as in the baseline specification.

TABLE B1 Non-transformed regional wage growth as dependent variable

	(1)	(2)	(3)	(4)	(5)
ΔMW	0.294	0.521***	0.245	0.262	0.476***
	(0.193)	(0.113)	(0.199)	(0.197)	(0.124)
ΔMW × (low wage subregion)			0.125	0.140*	0.108
			(0.092)	(0.080)	(0.105)
ΔMW ($t-1$)		0.263*		0.199	0.197
		(0.143)		(0.151)	(0.126)
ΔMW ($t-1$) × (low wage subregion)				0.190**	0.161**
				(0.072)	(0.059)
ΔMW ($t+2$)		0.300*			0.360**
		(0.151)			(0.158)
ΔMW ($t+2$) × (low wage subregion)					−0.143***
					(0.047)
Observations	1,095	1,022	1,095	1,095	1,022
Within R^2	0.423	0.494	0.425	0.451	0.501
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0			0.078	0.049	0
p-val. Coeff. on ΔMW ($t+2$) + Coeff. on ΔMW ($t+2$) × (low wage subregion) = 0					0.138

Notes: The dependent variable is the subregional growth rate of wages. The MW is defined as the ratio of the national minimum wage to the previous year's average wage in a given subregion. All regressions use NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks. Driscoll–Kraay standard errors in parentheses. *$p < .1$ **$p < .05$ ***$p < .01$. We also report the p-values for two tests: (i) that the effect of contemporary minimum wage hikes in low wage regions is zero, (ii) that the effect of future minimum wage hikes in low wage regions is zero.

Source: Own estimations based on Statistics Poland Local Data Bank data.

TABLE B2 Wage effects: fully interacted model and time fixed effects

	(1) Fully interacted model	(2) Fully interacted model	(3) Time fixed effects	(4) Time fixed effects
ΔMW	−0.038	−0.034	0.066	0.078
	(0.038)	(0.036)	(0.128)	(0.131)
ΔMW × (low wage subregion)	0.120*	0.129**	0.1	0.112
	(0.059)	(0.058)	(0.069)	(0.067)
ΔMW ($t-1$)		0.044		0.057
		(0.035)		(0.072)
ΔMW ($t-1$) × (low wage subregion)		0.129**		0.162***
		(0.048)		(0.044)
Observations	1,095	1,095	1,095	1,095
Within R^2	0.013	0.026	0.066	0.077
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0	0.259	0.16	0.044	0.038

Notes: The dependent variable is the subregional growth rate of wages decreased by the national growth rate. All regressions use the NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks; In columns (1) and (2), all control variables are also interacted with the dummy indicating low wage regions. In columns (3) and (4), year fixed effects are added. Driscoll–Kraay standard errors in parentheses.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.

TABLE B3 Employment effects: fully interacted model and time fixed effects

	(1) Fully interacted model	(2) Fully interacted model	(3) Time fixed effects	(4) Time fixed effects
ΔMW	0.035	0.037	0.042	0.045
	(0.022)	(0.023)	(0.106)	(0.108)
ΔMW × (low wage subregion)	−0.218***	−0.213***	−0.230***	−0.224***
	(0.071)	(0.071)	(0.072)	(0.069)
ΔMW ($t-1$)		0.024		−0.018
		(0.041)		(0.080)
ΔMW ($t-1$) × (low wage subregion)		0.063		0.065
		(0.137)		(0.099)
Observations	1,095	1,095	1,095	1,095
Within R^2	0.331	0.331	0.343	0.343
p-val. Coeff. on ΔMW + Coeff. on ΔMW × (low wage subregion) = 0	0.013	0.016	0.237	0.244

Notes: The dependent variable is change in the log of employment in a NUTS 3 subregion. All regressions use the NUTS 3 fixed effects and the following controls varying at the NUTS 3 level: lagged growth of the active population aged 20–64; lagged growth in the ratio of the immigrant to the working-age population; lagged growth of GDP; Bartik (1991) labour demand shocks. In columns (1) and (2), all control variables are also interacted with the dummy indicating low wage regions. In columns (3) and (4), year fixed effects are added. Driscoll–Kraay standard errors in parentheses.

*$p < .1$,; **$p < .05$,; ***$p < .01$.

Source: Own estimations based on Statistics Poland Local Data Bank data.