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Minimum Wage and Unemployment: Evidence from Russia

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Abstract: This paper investigates the unemployment effects of the Russian minimum-wage policy. The results suggest that higher minimum wages slightly increase unemployment rates among young workers but do not affect the older workforce. The textbook theory of producer is employed to rationalize the findings, showing that the magnitude of employment responses to minimum wages is associated with the elasticity of capital-labor substitution. Moreover, industries employ more workers informally if they cannot exercise capital-labor substitution easily. In line with the revealed unemployment among youth and informalization of the economy induced by the intervention, the findings show limited income effects of the policy.

Keywords: Minimum wage; Unemployment; Informality; Capital-labor substitution; Russia

JEL: J1; J2; J3; J6

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1. Introduction

The minimum-wage policy remains to be one of the most controversial issues among researchers and policymakers. Proponents of minimum-wage hikes claim that such policies may push the bottom tail of the state's income distribution up, while the implied unemployment effect is either absent or negligible (e.g., Katz and Krueger, 1992; Dube, Lester, and Reich, 2010; Draca, Machin and Van Reenen, 2011; Giuliano, 2013; Ahlfeldt, Roth and Seidel, 2018, Dube, 2019). However, some scholars argue that higher labor cost significantly limits the number of jobs available in the economy, and thus the policy potentially harms rather than protects workers (e.g., Brown, Gilroy and Cohen, 1982; Burkhauser, Kenneth, and Wittenburg, 2000; Neumark and Wascher, 2010; Meyer and West, 2016; Clemens and Winther, 2019).

Given the ongoing debate, the present study pursues two purposes. Our primary goal is to provide credible estimates of the unemployment effects of the Russian minimum-wage policy. Having pinpointed the estimated impact, we investigate whether the existing discrepancies in the unemployment effects that prevail in literature become concurrently noticeable within the Russian economy. We then provide explanations to the revealed pattern by employing the existing theory of producer behavior. We conclude the analysis by assessing the efficiency of the Russian minimum-wage policy as the poverty-alleviating tool.

The institutional arrangement of the Russian minimum-wage policy is comparable to the one in the United States. The government enacts the national-level wage floor, while every region retains the right to establish its unique minimum wage, which cannot be lower than the national level. The minimum wage does not vary across workers within the region-year cell. We thus view the Russian economy as another fruitful ground for conducting an empirical investigation.

We estimate the unemployment effects of minimum wages by exploiting two different empirical designs. First, we fit a canonic two-way panel fixed-effect model of the regional unemployment rate accounting for the potential reverse causality in the two-stage least squares

(2SLS) framework. The Russian institutional context supplies a plausibly valid instrumental variable (IV), the Bodman's Winter Severity Index, which quantifies the human perception of the winter severity (Bodman, 1916). The choice of the instrument stems from the observation that labor unions bargain for higher wages in places where winter is more severe. Our analysis excludes the possibility that the instrument affects regional labor markets through agricultural performance, electricity cost, workers' health, migration, or decisions to participate in the labor force. Apart from that, we consider another IV borrowed from the minimum-wage literature. We show that the results of the exclusion restriction test rule out the non-orthogonality of our instruments to the error term. These results combined reinforce our belief that the mainline proposed instrument (the Bodman Index) allows us to identify the unemployment effects of the Russian minimum-wage policy consistently. The results suggest that a 1% increase in the regional real minimum wage increases the unemployment rate among workers aged below 30 years by 0.05 percentage points. This number belongs to the lower end of the distribution of the current estimates in the literature (e.g., Brown, Gilroy, and Kohen, 1982; Card and Krueger, 1995; Neumark, 2019; Wolfson and Belman, 2019). In line with previous studies, our analysis reveals no impact of the minimum wage across the older groups of workers (e.g., Portugal and Cardoso, 2006; Neumark, Salas, and Wascher, 2014).

Second, we substantiate the above findings by exploiting a sizable increase in the minimum wage in the Kamchatka region, following a long tradition of investigating the unemployment effects of the minimum-wage policy in quasi-experimental frameworks (e.g., Card and Krueger, 1994). The primary motivation behind this empirical exercise is the need to test the unemployment effects of the minimum-wage policy in the absence of reliance on the validity of our instrument. We take advantage of a plausibly unexpected minimum wage change in the Kamchatka region, comparing the outcomes with the ones in a geographically adjacent control region. The highly unusual timing and magnitude of the minimum wage introduction in the "treated" group guide the choice of the pair of regions. The findings suggest that the workers younger than 30 faced a three percentage point higher probability of being unemployed in response to the 63% increase in the regional minimum wage. Differently put, a 1% increase in Kamchatka's minimum wage increased the unemployment rate among youth by approximately 0.05 percentage points, which is strikingly close to the point estimate obtained from the regional panel analysis. Additionally, we show that this point estimate is robust to the inclusion of a rich set of workers' characteristics and flexible regional trends. Finally, our estimation shows no detectable unemployment response to the pseudo–minimum wage assigned in the pre-policy period, which suggests that the observed change in the Kamchatka's employment pattern is attributable to the new wage floor.

If there is an unobserved regional heterogeneity that invalidates our identification, it would have to correlate with our instrument and affect the regional unemployment among young, but not older, workers. This heterogeneity would originate from a channel other than regional economic development, inflation, population size, agricultural performance, workers' health, cost of electricity, or migration. To explain away our findings fully, this unobserved factor should also coincide with the introduction of a sizable minimum wage hike in Kamchatka, which did not occur in the first quarter. A new minimum wage introduction usually occurs in the first quarter in most of the other regions. We posit that such a specific factor is unlikely to exist, so our proposed point estimate can be interpreted as the causal impact of the minimum-wage policy on unemployment.

Next, we examine the cross-industry heterogeneity in the unemployment effects of the minimum wage among youth. One might not expect that the employment effects in one industry will necessarily occur in another, to the extent that industries differ from each other across various dimensions. However, for policymaking, it is important to uncover those differences that drive the heterogeneity in the employment responses to the minimum-wage policy across different industries within the same economy.

Our analysis reveals that younger workers likely experience displacement in those industries that can substitute capital for labor more easily. These findings are consistent with one of the Hicks-Marshall laws of derived demand, which states that wage elasticity of labor demand is likely higher if other production factors can replace workers (Allen, 1938; Hamermesh, 1993). Additionally, we find that those industries that show little or no employment distortions hire more workers off the books. Thus, our findings of this part suggest that employer's ability to (i) substitute capital for labor, and (ii) hire workers off the books, are two revealed culprits behind the revealed cross-industry heterogeneity.

Finally, we assess the capability of the Russian minimum-wage policy to aid the bottom tail of the state's income distribution. In doing so, we replicate the recent study from the United States, applying the same statistical technique to the Russian data. Our results show that the effectiveness of the policy is, at best, limited. Job losses and recruitments off the books are two revealed bottlenecks that hamper the capability of the policy to aid the poor.

The remainder of the paper proceeds as follows. Section 2 introduces the institutional context, describes data and discusses the empirical strategies employed in the study. Section 3 presents the main findings on the unemployment effects of the minimum-wage policy. Section 4 explores the industry-specific youth employment responses to the minimum-wage policy and provides explanations. Section 5 investigates the household income effects of the policy. Section 6 concludes the paper.

2. Institutional Context, Data, and Identification

Russia institutionalized the state-level minimum wage that is regulated by the relevant Federal Law. Federal Subjects (from now on – regions) have the right to enact minima that must not be lower than the nationwide one. Thus, in some regions, nominal minimum wages exceed the federal. However, within each region, the minimum wage does not vary by workers.

We consider total unemployment as well as unemployment by age groups as our mainline outcome variables. It is also worth noting that the existence of the informal economy provides an income opportunity for a substantial fraction of those who are unable to enter labor markets as formally registered workers (La Porta and Shleifer, 2008; Comola and Mello, 2011; Muravyev and Oschepkov, 2016; Mora and Muro, 2017). We hence consider the response of the informal employment to the minimum wage hikes as the additional outcome. We employ two mutually reinforcing empirical designs described below to investigate the relationship between minimum wages and unemployment.

2.1 Evidence from Russian Regions

Our baseline design exploits a region-year panel data set (2009–2012). The key data source is the Russian Federal State Statistics Service, from where we obtain the socioeconomic characteristics of the regions. The data on regional minimum wages was input manually from the regional agreements between local governments and trade unions.

We estimate a classic two-way panel fixed-effect model of the following form:

$$y_{it} = \beta \log MW_{it} + X_{it}'\rho + \tau_i + \theta_t + u_{it}, \qquad (1)$$

where *i* and *t* index region and year, respectively. The outcomes y_{it} are (i) the annual average overall unemployment rate, (ii) the unemployment rate across various age groups (described further in the text), and (iii) the informality rate, defined as a fraction of labor force employed informally. The key regressor is $\log MW_{it}$, the natural log of the regional minimum wage; X_{it}' is a vector of controls, u_{it} is the error term. Following the literature, our preferred estimation controls for regional CPI, log of GDP per capita, and log of the working-age population (Burkhauser, 2000; Orrenius and Zavodny, 2008; Dube et al., 2010; Meer and West, 2016). We allow the intercept (τ_i) in Equation (1) to vary across regions, while including time dummies (θ_t) to control for the common year shocks. All monetary variables are measured in the 2010 prices. Table 1 reports the summary statistics of the variables employed in the analysis. Given the cross-sectional variation in the minimum wages within the given year, we do not expect to consistently estimate β in Equation (1) by ordinary least squares (OLS). The potential problem unresolvable in the OLS setting is the reverse causality: regional legislators are unlikely to push the local minimum wage up if the local unemployment rate is high. We deal with this issue by estimating Equation (1) employing the 2SLS technique.

The main instrument we use for minimum wage is the regional Winter Severity Index. One of the earliest winter severity measurements was developed in 1910 by a Swedish scientist, Goesta Bodman (Bodman, 1916). In his summary of the research expedition to the South Pole in 1901–1903, Bodman attempted to objectively quantify the subjective human perception of the winter severity as follows:

$$S = (1 - 0.04 \times T) \times (1 + 0.27 \times V), \quad (2)$$

where T is the temperature measured on the centigrade scale (°C), and V is the wind speed (measured in meters per second). The higher is the value of S, the worse the human perception of the winter is assumed to be. Equation (2) reflects the Bodman's rationale: the wind speed has a much higher weight than the temperature. For example, if the temperature is –25 °C and the wind speed is 1 meter per second, the Bodman Index is equal to 2.54, while if the temperature is –8 °C and the wind speed is 17 meters per second, the index then takes the value of 7.38. This is explained by the property of wind to create breathing difficulties, skin damage, and mechanical pressure on the human body (Bodman, 1916). The data needed to construct the index are obtained from the Hydrometeorological Centre of Russia.

We argue that the Bodman Index is a good candidate for the valid instrument, as it affects the level of minimum wages (relevance) while having no direct impact on the regional unemployment rates (exclusion restriction). We first discuss the first-stage relationship between the regional Bodman Index and the minimum wage level. We then summarize our effort to justify the exclusion restriction.

Instrument Relevance.— One may notice from the Russian labor market history that in some regions, the minimum wage has always been kept higher than the federal floor. The existence of the so-called Northern Multiplier (NM) explains this phenomenon. In the North and Far East, where winter is much colder, workers are paid higher minimum wages regardless of the labor market conditions. Hence, NM serves as a region-specific constant wage multiplier, which is effectively controlled for by the panel fixed effect estimation. However, the existence of this multiplier in the past rationalizes the modern existence of the first-stage causal relationship between the Bodman Index and regional minimum wages. In the given region, a higher winter severity index endows local labor unions with a higher bargaining power to claim higher minimum wages. This reasoning serves as a rationale behind the instrument's relevance.

Exclusion Restriction.— We now argue that the Bodman Index does not affect local unemployment rates directly. Figure 1 shows the distribution of the OLS residuals of the Bodman Index in 2012 (the most recent cross section in our panel data set) obtained from regressing the Bodman Index on regional and year fixed effects. The visual analysis suggests that, once we partial out regional time-invariant characteristics (such as geographic location, proximity to water bodies) and common year shocks, the remaining variation in the Index is essentially random. Indeed, although the high net levels of the Bodman Index primarily concentrate in the northern part of Russia, changes in the index *within* the given region should not depend upon its location.

Additionally, we think of the list of regional characteristics that might be responsible for the potential correlation between our IV and the outcomes via the channels other than the minimum wage. We consider regional health indicators (population-to-doctors ratio, number of registered diseases per 1,000 people), net migration inflow (per 10,000 people), average household utility expenses, labor force participation rate and performance of the industries potentially affected by winter severity (agricultural productivity index, measured as agricultural output over GDP, and electricity production, measured as billions of kWh). We first show that none of the above indicators is in the systematic relationship with the Bodman Index. However, considering that these factors might jointly belong to Equation (1) while being in the systematic relationship with the Bodman index, we include all of them in the unemployment equations as a robustness check. We show that this exercise does not affect our point estimates.

Apart from that, we borrow the idea of including the mean of the minimum wages in geographically bordering regions as another instrument (Neumark and Wascher 1992; Rybczynski and Sen 2018). The rationale is that local legislators are likely to observe employment responses to the minimum wage in the neighboring regions before enacting the policy on their own. Equation (1) is thus overidentified, which allows us to test the exclusion restrictions formally. The test result cannot serve as proof that Bodman Index is truly excluded from Equation (1). However, we view a failure to detect any systematic relationship between the IVs and the error term in (1), as it is the case in our study, as a good sign.

2.2 Evidence from Kamchatka: A Natural Experiment

In this subsection, we replicate the idea behind many existing minimum-wage studies that rely upon natural experiments (Card and Krueger, 1994; Machin, Manning, and Rahman, 2003; Machin and Wilson, 2004; Baek and Park, 2016, and others). We are primarily motivated to test our findings' robustness without relying upon the validity of our IVs.

Regions often enacted their unique minimum-wage floors, which suggests that every wage hike has the potential to be explored in a quasi-experimental framework. However, given the crucial importance of the parallel trend assumption needed for the identification, we pick the regional case study that satisfies the following criteria:

(i) The minimum-wage increase must be sizable. This criterion excludes those cases in which regions simply adjust local wage floors to inflation.

(ii) The intervention must not occur in the first calendar quarter. Employers will thus less likely anticipate it, as in most cases, new minimum wages get introduced in January or February.

(iii) The initial minimum wage in the treated region should follow the state's minima, as regions with a traditionally higher minimum wage might have labor market dynamics noncomparable to those that follow the state's wage floor.

(iv) The treated region must have an unambiguous geographically neighboring control group that strictly followed the state's minimum wage during the observed period.[2]

One region matches all four criteria: Kamchatka. Kamchatka's minimum wage rose from 4,611 RUB (Federal Minimum Wage) to 7,500 RUB in June 2012. The timing of the intervention is highly unusual, as most of the other regions routinely update their local minimum wages in the first calendar quarter.[3] Also, there exists only one geographically neighboring region that had been complying with the state's minimum wage within the observed period – Chukotka. The latter fact excludes the ambiguity in choosing the control region.

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[2] A growing number of minimum-wage studies utilizes a synthetic control approach to let the control group be generated by the data-driven algorithm. Examples of such include: Sabia, Burkhauser, and Hansen (2012), Allegretto et al. (2017), and Neumark and Wascher (2017). However, we cannot implement this idea in our study, as we do not have sufficiently long regional time-series of the youth employment dynamics to generate a synthetic control group.

[3] It is rather uncommon to enact a higher minimum wage in the periods other than beginning of the year (January) in Russia. The exceptions are those regions that index minimum wages, adjusting them to the inflation. However, few regions index wages, and the resulting changes are negligible in magnitude.

We utilize four waves (years 2010, 2011, 2012, and 2013) of the Russian Labor Force Survey in the two regions mentioned above to conduct our analysis. Past observations suggest that the unemployment response to minimum wages likely varies by workforce age cohorts (Neumark and Wascher, 1992; Currie and Fallick, 1996; Giuliano, 2013; Muravyev and Oschepkov, 2016, and others). We could estimate the policy impact by dividing our sample into young (Y) and older (O) workforce and then fitting the following models:

$$y^Y = \beta_1^Y T + \beta_2^Y Post + \beta_3^Y T \cdot Post + v^Y \tag{3.1}$$

$$y^O = \beta_1^O T + \beta_2^O Post + \beta_3^O T \cdot Post + v^O, \tag{3.2}$$

where y is one of two binary outcomes: (i) *unemployment* – equal to 1 if the individual is unemployed (0 if employed), and (ii) *informal employment* – equals to 1 if the individual is employed informally (0 for formally employed). $Treated$ is a binary indicator of the treated region (Kamchatka), and $Post$ indicates the post-treatment period's dummy (that takes on a value of 1 for the period after June 2012). Equations (3.1) and (3.2) can be recognized as difference-in-differences specifications in their canonic forms. If β_3^Y and β_3^O do capture the policy's impact, then the existence of the difference in unemployment responses across two age cohorts can be detected statistically from exercising the following hypothesis test:

$$H_0: \beta_3^Y - \beta_3^O = 0. \tag{3.3}$$

We explicitly test (3.3) by combining models (3.1) and (3.2) in one equation as follows:

$$y = \delta_1 T + \delta_2 Post + \delta_3 Youth + \delta_4 T \cdot Post + \delta_5 T \cdot Youth + \delta_6 Youth \cdot Post + \delta_7 T \cdot Post \cdot Youth + v, \tag{3.4}$$

where the new variable $Youth$ takes on the value of 1 if the individual's age belongs to the closed interval [15, 29], and 0 otherwise. The primary rationale behind the proposed age interval is to obtain the results comparable to the ones drawn from the regional-level analysis. Further, in the text, we redefine our $Youth$ variable, exploring the behavior of the estimated unemployment effects of the minimum wage as we consider larger age windows.

The parameter of interest is δ_7 in Equation (3.4). For the average outcome in Equation (3.4), $\bar{y}_{rat}$, where *r* indexes region (*T* if it belongs to the treated region, *C* – otherwise), *a* indexes the age group (*Y* if aged between 15 to 29, *O* – otherwise), and *t* indicates the treatment timing (*1* if observed after the policy's introduction, *0* – otherwise), the OLS estimate of δ_7 can be viewed as:

$$\widehat{\delta_7} = (\bar{y}_{TY1} - \bar{y}_{TY0}) - (\bar{y}_{CY1} - \bar{y}_{CY0}) - (\bar{y}_{TO1} - \bar{y}_{TO0}), \quad (4)$$

which is known to be a *difference-in-difference-in-differences* (*DDD*) estimator. In other words, testing that δ_7 in Equation (3.4) is equal to zero is algebraically identical to testing the null hypothesis given by Equation (3.3). Under the assumption that both regions would have followed parallel unemployment trends in the absence of the new minimum wage's introduction, δ_7 from Equation (3.4) identifies the policy impact on the workforce aged below 30. We also consider a regression-adjusted variant of Equation (3.4) by including a rich set of individual controls (age, age squared, marital dummy, gender, educational attainment, urban area's dummy, number of children below 18 years old in the household), region-quarter fixed effect, and the region-specific quartic month-year trend. The latter term partly accounts for possible trend differences between the two regions that are not captured by our regressors. Further in the text, we show that the estimated impact is not sensitive to the inclusion of the proposed set of controls.

3. Unemployment Effects of the Russian Minimum-Wage Policy: Results

This section presents the mainline results of our empirical investigation. Following the sequence of the previous section, we first discuss the results obtained from the regional panel data. We then report the results obtained from our analysis of the new minimum wage introduced in the Kamchatka region. We discuss auxiliary regression results and specification tests where needed.

3.1 Cross-Regional Unemployment Effects

Table 2 reports the IV estimates of Equation (1). Panel A reports the second stage results, while Panel B reports the first-stage regressions of log real minimum wage on the excluded instrument (the Bodman Index) and covariates. We first discuss the results in Panel A.

Baseline Findings.—Column (1) of Table 2, Panel A confirms the findings reported in the recent literature: the estimated minimum wage's coefficient is statistically insignificant and numerically close to zero in the overall unemployment's regression. Columns (2) and (3) suggest that the young workforce is likely to constitute a vulnerable group. According to the results, on average 1 percent increase in the real minimum wage results in 0.046 and 0.056 percentage points increase in the unemployment rates among those aged 15–19 and 20–29, respectively. Column (4) suggests that the older workforce, namely – those aged above 30 years, are not likely to be vulnerable to the minimum-wage policy. The revealed age-pattern is consistent with the existing studies (Brown, Gilroy and Kohen, 1982; Ehrenberg and Marcus, 1982; Swidinsky, 1980; Neumark and Wascher, 1995; Burkhauser, Kenneth, and Wittenburg, 2000; Portugal and Cardoso, 2006, and others). The results in Column (5) suggest an increase in the informality rate in response to the minimum-wage hike. In line with the literature, we view the latter observation as additional evidence of the minimum-wage policy being binding for employers: when facing higher labor costs, companies may hire workers off-the-books to replace formal workers (Comola and Mello, 2011; Muravyev and Oschepkov, 2016; Mora and Muro, 2017).

Validity of the IV.—As we discuss in the methodology section, the key parameter in Equation (1) can only be identified if the Bodman Index affects the regional minimum wage while being excluded from the unemployment function. As the results reported in Panel B of Table 2 show, Bodman Index is strongly related to the minimum wage, which relieves the concern about the strength of the first-stage relationship. Having this result, we assess the possibility of the non-orthogonality of our proposed instrument to the error term in Equation (1) as follows.

We first highlight several regional characteristics that might be related to the Bodman Index, namely: (i) log population-to-doctors ratio, (ii) log number of registered diseases per 1,000 people, (iii) net migration inflow (per 10,000 people), (iv) agricultural output normalized by regional GDP, (v) log electricity production (Billions of kW-hour), (vi) average household utility expenses normalized by average household income, and (vii) labor force participation rate. We then test whether those regional characteristics respond to Bodman Index. As the additional check, we control for the abovementioned variables on the right-hand side of Equation (1) to see whether their inclusion alters the mainline findings.

Tables A1 and A2 report the results of the above two exercises. Table A1 suggests that none of those indicators responds to changes in the Bodman Index. Table A2 shows that the inclusion of those additional indicators does not affect the estimated impact of minimum wages on unemployment rates and informality. These exercises largely exclude the possibility that the Bodman Index affects the regional unemployment rate through health, migration, agricultural productivity, electricity cost, or the regional labor force participation rate.

Moving further, we employ the additional instrumental variable to test the exclusion restriction statistically. As we discuss in the previous section, the idea is to use two instruments for the potentially endogenous minimum-wage variable, which allows us to conduct the test formally. Table 3 reports the results. Panel A of Table 3 suggests that the inclusion of the additional instrument does not change the mainline point estimates much numerically. Panel B displays the first-stage regression results. The coefficients on the additional instrument (log($\overline{MW}$ neighbors)) are positive and statistically significant across all columns, which confirms that regional legislators are likely to monitor those neighboring regions that enact higher minimum wages. The numerical magnitudes of F-statistics of the excluded instruments confirm the strength of the first-stage relationship. The last row of Table 3 reports the *p*-values of the Sargan-Hansen's *J* test. None of them falls below the 10% significance level, which adds to our belief that the Bodman Index is excluded from Equation (1).

We highlight two key findings, to summarize our validity checks. First, the data confirm the existence of the first-stage relationship between the Bodman Index and the regional minimum wage. Second, the analysis reveals no relationship between the Bodman index and a list of potential suspects that might govern the relationship between the instrument and the outcome (unemployment rate by age groups and informality rate). On top of that, a direct statistical test of the exclusion restriction testifies the hypothesis that the regional minimum wage is the only channel through which within-region variation in the Bodman Index affects the outcomes. These combined findings suggest that the revealed second-stage relationship can be interpreted as the causal impact of the minimum-wage policy on unemployment.

3.2 Unemployment Response in Kamchatka

In this subsection, we discuss the results of investigating a natural experiment in the Kamchatka region. Table 4 reports summary statistics for all the variables employed in this analysis. As before, we consider unemployment and informal employment as our outcomes. Individuals' education, gender, age, marital status, number of children, and residence in the urban area constitute the set of individual-level controls.

Table 5 reports the estimates of Equation (3.4) and its variants with various controls. Columns (1)–(3) of Table 5 report the results for unemployment regressions and Columns (4)–(6) for informal employment. Columns (1) and (4) show the results of estimating Equation (3.4) without controls. Columns (2) and (5) show the results obtained after controlling for the individual characteristics. Columns (3) and (6), our preferred specifications, further endow the estimation with the region-specific seasonality and quartic regional time trends. The former allows the regions to follow a region-specific seasonal pattern. The latter flexibly approximate changes in the labor force status resulting from the movements in those unobserved time-variant regional characteristics that are relevant to our outcomes.

Table 5 shows that the key point estimates of the triple interaction term (Post × Treated × Youth) display little variance across all columns. The results in Column (3) suggest that the minimum wage's hike increased unemployment probability among those aged below 30 by 0.028. Column (6) shows that the probability of being employed informally among young workers increased by 0.04 in response to the policy intervention.

Overall, the results in Table 5 suggest that the minimum wage's introduction is associated with roughly 0.03 (0.04) increase in the probability of being unemployed (employed informally) among young workers. Besides, the coefficients of the [Post × Treated] term in Columns (1)–(3) suggest that unemployment among older workers is unlikely to be affected by the minimum-wage policy, which goes in line with the existing studies and our findings in the previous subsection. However, Columns (4)–(6) suggest that informality is likely to increase even among more experienced employees, in line with our findings from the regional panel data.

We test the assumption of the parallel trends to assess the credibility of our results , considering the fact that both regions had been following the state's minimum-wage floor in 2010 and 2011. We hence first define a placebo variable being equal to 1 for the year of 2011 and 0 for 2010, which corresponds to the period with no changes in the minimum-wage policy in the treated region. We then substitute the term *Post* for our placebo treatment in Equation (3.4) and test the resulting model. A failure to detect a systematic difference in the youth employment's response to the placebo treatment across two regions will not prove that the parallel trends' assumption indeed holds. Yet, we believe it might serve as evidence that parallel trends' assumption is plausible.

Table 6 reports the results of this falsification test. We report the coefficients on the key interaction term only (*Placebo* × *Treated* × *Youth*) to save space. Overall, the results suggest that the two regions do not have statistically detectable differences in the employment trends in the pre-treatment periods. This finding reinforces our belief in the credibility of the proposed quasi-experimental design.

Finally, we explore the behavior of the *DDD* estimate under the varying definition of the younger workforce as follows. We first let $Youth$ variable from Equation (3.4) to take on the value of 1 if the individual's age belongs to the half-opened interval $[15, m)$, $m = 24, 25, 26 \ldots, M$, and 0 otherwise. We then run a set of the resulting regressions and trace down the behavior of the estimated parameter on the triple interaction term from Equation (3.4) as we gradually move the upper bound of $Youth$ from 24 (which roughly corresponds to the average age of a fresh university graduate) up by one year.

Figure 2 plots the OLS point estimates and their corresponding 95% confidence intervals of the parameter of the triple interaction term in the unemployment equation by age groups. Panel (a) shows the estimated impact of the actual policy, while Panel (b) shows the outcome's response to the placebo treatment in the pre-treatment period. We maintain the full model specification, controlling for the individual determinants of unemployment, region-specific seasonality, and quartic time trends. Panel (a) suggests that, as we move the definition of the young workforce from the age range $[15, 24)$ to $[15, 44)$, the estimated unemployment effect of Kamchatka's minimum wage diminishes, and becomes no longer significant from $[15, 32)$. These results confirm our previous findings that younger workers are more vulnerable to the minimum-wage policy.

Panel (b) repeats the above exercise, replacing the actual policy by placebo treatment. The point estimates are statistically indistinguishable from zero over the entire age range. The plot does not reveal an age pattern. We conclude that pre-legislation trends in the outcome are likely to be the same across both regions. This finding reinforces our belief in the credibility of our mainline findings in this section.

Our overall findings suggest that the minimum-wage policy does generate a higher unemployment rate. Evidence from Russian regions suggests that a unit percentage change in the minimum wage increases the local unemployment rate among workers aged below 30 years by 0.05 percentage points. A regional case study suggests that a 63 percent increase in Kamchatka's minimum wage resulted in a three percentage point higher probability of being unemployed among workers aged below 30. Put it differently, a 1 percent change increase in Kamchatka's minimum wage increased the share of unemployed among youth by 0.05 percentage points – a substantially comparable result.

4. Industry-Specific Employment Responses to Minimum-Wage Policies

Our findings indicate that youth unemployment rates do increase in response to the minimum-wage hikes, though the estimated impact is numerically small. The estimated impact belongs to a wide range of point estimates in the literature. The next question, then, is whether we can reveal the variation of the employment effect of minimum wages within the same economy. We provide policy-relevant explanations to the revealed pattern after pinpointing the existence of its variability.

We show that the industry-level employment response to the minimum wage follows the Hicks-Marshall law of derived demand, discussed by Allen (1938).[4] Hence, we argue that one structural parameter, which is the elasticity of capital-labor substitution, is the revealed culprit behind the estimated cross-industry heterogeneity of the employment effects of the Russian minimum-wage policy. We provide evidence that employers hire more workers informally if the existing production technology does not permit the substitution of capital for labor.

[4] A more detailed discussion is provided by Hamermesh (1993).

Formally, the key prediction that we aim to test in this section is the following:

Two elastically supplied inputs *K, L* (capital and labor respectively) are mapped into a single output according to the linearly homogenous production function $Y = F(K,L)$. Therefore, the wage elasticity of labor demand is:

$$\frac{\partial L}{\partial w}\frac{w}{L} = -\eta\frac{wL}{Yp} - \sigma\frac{rK}{Yp}, \tag{5}$$

where w, r denote the per-unit cost of labor and capital respectively; η indicates the absolute value of the market's price elasticity of demand for the industry's final good, and σ - elasticity of substitution (see details in Appendix B).

The term $\sigma\frac{rK}{Yp}$, which is central to our empirical analysis, quantifies the shift of the optimum (K^*, L^*) along the isoquant in response to the new input price ratio. As σ approaches infinity, the production technology approaches the "perfect substitutes" case, and hence the industry responds to higher wages by laying off more workers. In the opposite limiting case of σ approaching zero, the capital-labor ratio does not change in the optimum, indicating zero ability of the producer to exercise the input substitution. The term $\eta\frac{wL}{Yp}$ is a so-called scaling effect, or the industry's inputs adjustment resulting from the production isoquant shifting backward. This effect is of interest to those studies that explore the price effects of the minimum wages (Aaronson, 2001; Aaronson and French, 2008; Wadsworth, 2010).[5]

Alternatively, we may rewrite the wage elasticity of labor demand as:

$$\frac{\partial L}{\partial w}\frac{w}{L} = -\eta - \frac{rK}{wL}\epsilon_w^K, \tag{6}$$

[5] We note that our reasoning implies a unit elasticity between the regional minimum wage and industry's wage rate that barely holds in the reality. However, we show that the relationship still holds empirically, consistent with the existing studies, showing positive wage effects of the minimum wage.

where ϵ_w^K denotes the wage elasticity of demand for capital. The above expression will come in handy empirically, as it allows us to test the capital-labor substitution hypothesis without estimating the elasticity of substitution, σ.

Data and Identification.—We first consider the population regression function of the following form:

$$\log Y_{j(i)t} = \alpha \log MW_{it} + \sum_{j=1}^{J} \lambda^j \log MW_{it} \times \psi_j + X_{it}' \varsigma + \psi_j \times \delta_i + \theta_t + \varepsilon_{j(i)t}, \quad (7)$$

where *i, t* index region and year, while *j* stands for the industry (that follows the SIC Division Structure). Following the literature (as well as the findings in the current paper), we focus on young workers under 30. $Y_{j(i)t}$ hence denotes one of the following outcomes: (i) log number of workers aged below 30, (ii) log number of workers employed informally, (iii) log capital investments. MW_{it} and X_{it}' are the real minimum wage and a vector of controls, respectively (identical to the ones employed in Equation (1)). $\psi_j, \delta_i, \theta_t$ indicate industry, region, and year fixed effects, respectively, and $\varepsilon_{j(i)t}$ is the error term. The term $\log MW_{it} \times \psi_j$ is hence the product of $\log MW_{it}$ with the industry *j*'s binary indicator. Therefore, λ^j identifies an industry-specific employment response to the minimum wage increases. We estimate Equation (7) by 2SLS, using the Bodman Index and [Bodman Index × Industry Dummy] as instruments for minimum wages and their interactions with industry dummies. We obtain the data needed to estimate Equation (7) from the Russian Federal State Statistics Service (years 2010–2012). The estimation results in the [*J+1*] first stages are omitted in the paper to save space.

After obtaining a column of the estimators of λ^j, we proceed as follows. If the production technology matters, a set of estimates of λ^j should exhibit a downward-sloping linear relationship with a set of the estimated industry-specific elasticities of substitution, given by σ in Equation (5). The main challenge in estimating σ is data availability. Contemporary macroeconomics literature tends to identify the substitution elasticity relying upon data-intensive approaches, such as estimating the parameters of the first-order conditions of the profit maximization, employing indirect objective function variants (e.g., Berndt, 1976; Klump, McAdam and Willman, 2007; Klump, McAdam, and Willman, 2012), or running non-linear least squares estimations (Kumar and Gapinski, 1974; Henningsen and Henningsen, 2012; Koesler and Schymura, 2015).[6] Given the absence of the regional time series on capital and its rental price, the adoption of any of those empirical strategies is not feasible in our case. However, we may invoke an older strand of literature in an attempt to get an approximate idea on the relative magnitudes of elasticities of substitution across industries. Appendix C summarizes one of the propositions (and its proof) delivered by Arrow, Chenery, Minhas, and Solow (1961; from now on – ACMS). More precisely, we adopt the ACMS idea of utilizing the joint variation of the value-added per worker and the real wages in an attempt to identify σ in every industry as follows:

$$\log y^j{}_{it} = c_i + \sigma^j \log WAGE^j{}_{it} + X'_{it}\rho + \vartheta_t + tc_i + \omega^j{}_{it}, \qquad (8)$$

where $y^j{}_{it}$ and $WAGE^j{}_{it}$ denote real value-added per worker and real wages. X'_{it} denotes controls (log regional real GDP per capita and CPI). The variables *j, i, t* index industry, region, and year, respectively. A superscript *j* indicates that we estimate Equation (8) for every industry (for which we have the necessary data) separately. We allow the intercept to vary by regions while controlling for common shocks (ϑ_t) and region-specific linear trends (tc_i); $\omega^j{}_{it}$ is the error term. All the monetary variables employed are at 2005 prices.

The regional series of per-worker value-added and average wages for the years 2004–2012 are obtained from the Russian Federal State Statistics Service. Before discussing the results, we note that the Panel Fixed-Effects estimation of Equation (8) is unlikely to be consistent. Both the outcome and the main regressor are likely to be determined simultaneously, and hence our proposed approach is inferior to the menu of alternatives discussed by Miguel et al. (2010).

[6] See Miguel, McAdam and Willman (2010) for a detailed overview of the existing approaches.

Yet, we may hope that the resulting estimates will represent the relative magnitude of the substitution elasticities in the population. That is, if the least-square estimation of (8) reveals that the elasticity of substitution in industry A higher than in B, then, we may take it as evidence that this order of magnitudes holds in the population.

Alternatively, we may think of capital investments' response to the minimum wage (that is when we use log capital investments on the left-hand side of Equation (7) representing the industry-specific substitution elasticity, as Equation (6) states. We utilize this simple reasoning as an alternative way to check the existence of the link between the employment effects of the minimum wages and the curvature of the industry's production isoquant.

Results.—Table 7 presents the results. We exclude two non-competitive sectors from our analysis: (i) public administration and defense, and (ii) extraterritorial organizations. Columns (1) and (2) of Table 7 present the estimates of the Equation (7) for the log number of young workers and the log number of workers employed informally, respectively. The omitted industry is "financial and insurance activities." The results suggest that the responses of the two outcomes go in opposite directions. Figure 5 plots the coefficients from column 1 of Table 7 against those in column 2. The resulting pattern suggests that those industries that do not displace many workers in response to the minimum wage hire workers off the books more. We view this pattern as one possible way of industry-level adjustment to the higher labor cost.

Column 3 of Table 7 presents the estimated capital investments' response to the minimum wage. Column 4 reports the results of estimating Equation (8) separately by every sector.[7] To gain insight into the patterns presented in Columns 3 and 4 of Table 7, we visualize the point estimates in Figure 5. Panel (a) plots the point estimates in column (1) against those from column (4). We interpret the resulting downward-sloping line as the estimated cross-industry

[7] We do not have data on financial sector wages and hence leave the corresponding cell empty.

average employment loss driven by the substitution elasticity. Higher substitution elasticity links with higher (in magnitude) employment loss in the industry, consistent with theoretical predictions. However, given the empirical challenges to estimating the substitution elasticity consistently, panel (b) of Figure 5 plots the point estimates in column (1) against those from column (3) of the same table (Table 7). If our instrument, the Bodman Index, is truly excluded from Equation (7), the estimates on both axes are consistent. If one industry is able to invest more capital in response to the minimum wage increase, it is likely to decrease demand for young workers, consistent with the capital-labor substitution hypothesis that we discuss at the beginning of the present section.

Combining the patterns in Figures 4 and 5, we argue that industries have at least two instruments to absorb the minimum-wage shock. If the underlying production technology permits to substitute capital for labor, the industry responds accordingly. This finding is in line with the emerging strand of the literature on the capital investment's response to the minimum-wage capital (Hau, Huang, and Wang, 2016; Geng et al. 2018). However, if easy capital-labor substitution cannot take place, then the industry responds to higher labor costs by employing workers off the books. Finally, column (3) of Table 7 suggests that overall industries invest less capital in response to the higher minimum wage, which goes against the recent evidence from Hungary (Harasztosi and Lindner, 2019). We attribute the difference to the fact that our estimated elasticity of substitution between capital and labor is substantially smaller. Despite the methodological issues related to our attempt to estimate the substitution elasticity, negative capital investments' response is consistent with column (4) of Table 7, suggesting the overall complementarity between capital and labor.

5. Does the Russian Minimum-wage Policy Help the Poor?

We conclude our empirical investigation by assessing the efficiency of the Russian minimum-wage policy as the poverty-alleviating tool. Ultimately, aiding the bottom tail of the state's income distribution is the baseline purpose of the policy. However, potential job losses and imperfect compliance on the employer's side hardly allow predicting that the policy necessarily aids poor households unambiguously.

Our analysis replicates the approach of Dube (2019). Dube estimates the household income effects of minimum-wage policy, fitting the following model using the U.S. data:

$$I_{cist} = \sum_{k=-1}^{3} \alpha_{ck} \log(MW_{s(t-k)}) + X_{ist}\Psi_c + \Theta + \gamma_s t + \epsilon_{cist} \quad (9)$$

The outcome variable is a binary indicator for whether the i-th individual (observed in state *s* at time *t*) comes from the family whose value of equalized income divided by the federal poverty threshold falls below *c*. For example, the observation with $I_{\mathbf{1}ist} = 1$ indicates that the individual's family is below the official poverty threshold. $I_{\mathbf{0.5}ist} = 1$ informs that the equalized income of the i-th individual's household is below 0.5 times the official poverty rate. Following the baseline analysis of the original study from the United States, we consider the values of *c* between 0.5 and 1.75 in increments of 0.25.

The set of key variables of interest includes contemporaneous, one-year leading, and three-year lagged values of the real minimum wage. The original version of the model also contains individual-level characteristics (quartic in age, gender, race, ethnicity, education, family size, number of children and marital status) and state-level macroeconomic indicators (unemployment rate, EITH supplement and income per capita) on its right-hand side. The term Θ includes the state fixed effect, division-year fixed effect, and state fixed effects interacted with the binary indicators for recessionary years (2007-2009); $\gamma_s t$ is the state-specific linear trend, and ϵ_{cist} is the error term.

The existence of the region-year variation of the minimum wage allows us to fit Equation (9) using data on Russian households. For this purpose, we retrieve data from publicly available

"Survey of Income and Program Participation" from the Russian Federal State Statistics Service. The survey results on hand cover 2014-2017 years. Following Dube (2019), we restrict the scope of the analysis to the working-age population, which results in 373,555 individuals in the sample.

The replication exercise conducted in this paper deviates from the original estimation strategy of Dube (2019) in three ways. First, Russia does not have an analog of the Earned Income Tax Credits (EITC) program. Second, microdata on households is not available for the recessionary years. Finally, our data set does not contain information on the ethnicity of respondents. Apart from the forced exclusion of the above information from the econometric investigation, our analysis fully replicates the model employed in the recent U.S. study by Dube (2019).

Having obtained the set of the estimated coefficients of the minimum wage variables from Equation (12), we infer the long-run elasticity of proportion of individuals falling under the income-to-needs cutoff, c, with respect to minimum wage from estimating the following expression:

$$Elasticity = \frac{\alpha_{c0}+\alpha_{c1}+\alpha_{c2}+\alpha_{c3}}{\mu_c}, \tag{10}$$

where μ_c is the share of population below c.

Table 8 shows the resulting elasticities for different income cutoffs. Column (1) shows the multiples of the federal poverty threshold. Column (2) shows the baseline results retrieved from the original U.S. study of Dube (2019).[8] Column (3) shows the results obtained from Russian data. Contrasting the results reported in columns (2)–(3), we note that poverty in Russia is less elastic. For example, the estimated reduction in the shares of individuals falling below 50 and 100 percent of the federal poverty threshold (in response to the unit change in *Log(MW)*) ranges between –0.455 and –0.446 in the United States. However, the estimated range of elasticities in Russia is –0.048 and –0.070. The numerical values of elasticities attenuate in both countries

[8] The estimates have been obtained from Table 3 of the original Dube's paper.

as the higher value of the income-to-needs cutoff, c, is taken into account. However, according to column (3) of Table 8, the estimated effect is significant at 5 percent level only for the share of individuals whose income falls below 100 percent of the federal poverty threshold.

We conclude this section inferring minimal effectiveness of the Russian minimum-wage policy as the poverty-alleviating tool. Job losses among young workers and the growing informalization of the Russian labor market implied by the minimum-wage increases are two revealed culprits that may explain the findings of this section.

6. Summary and Synthesis

This paper estimates the employment effects of minimum-wage policy in Russia. We first show that regional unemployment rates are likely to respond to the rising wage floor. Our results suggest that the overall unemployment rates remain intact, while the potentially affected group is young workers (those aged below 30 years old). The point estimates indicate that a 1% increase in the real minimum wage increases unemployment rates among teenagers (15–19 years old) and young workers (20–29 years old) by 0.046 and 0.056 percentage points, respectively. We also provide evidence on the redistribution of workers toward the informal sector resulting from the enactment of a higher wage floor.

Our estimates are robust to the battery of alternative model specifications, which includes the use of the additional instrumental variable and the inclusion of a rich set of regional controls. We also show that our findings hold even when we employ alternative empirical design by adopting the common natural experimental approach in the Russian setting. The results show that a 1% increase in the nominal minimum wage in Kamchatka is likely to have increased the share of unemployed among youth by 0.05 percentage points. The latter finding also passes its sequence of robustness checks.

Having bracketed the existence of the unemployment effect of minimum wages, we investigate whether the disparity of the estimated employment effects of the minimum-wage policy can be detected within the same economy at the same point in time. We, therefore, invoke an old strand of literature on producer behavior that shows that one structural parameter, the elasticity of substitution, matters for the observed magnitude of the unemployment effect of minimum wages among youth. We show that those industries that are more capable of substituting capital for labor are more likely to experience a steeper youth employment loss. However, if the substitution is not permissible by the underlying production technology, industries are expected to employ more workers informally. We thus infer that employers' ability to replace labor with capital, or formal with informal employment, are two potential factors to reconcile the diverging unemployment effects of minimum wages, which are often estimated from data on different industries, countries, and periods.

We conclude the paper by assessing the Russian minimum-wage policy as the poverty-alleviating tool. In line with the previously revealed job losses among youth and informal recruitments of workers, we find the minimal impact of minimum wages on the workers' income. Our results question the effectiveness of the Russian minimum-wage policy as a tool primarily designed to aid low-income families.

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Table 1—Summary statistics

Variables	Mean (Std.)
Unemployment rate (Outcome):	
Overall	0.0808
	(0.0439)
15–19 years old	0.0802
	(0.0401)
20–29 years old	0.0616
	(0.0349)
30+ years old	0.0537
	(0.0282)
Mainline controls:	
Informality rate	0.113
	(0.0348)
Log (MW)	8.426
	(0.173)
Log (GDP per capita)	12.12
	(0.532)
CPI	-0.0167
	(0.0249)
Log (Working-age population)	13.57
	(0.755)
Additional Controls	
Log (population-to-doctor)	4.494
	(0.151)
Log (Disease per 1,000 people)	6.675
	(0.186)
Net Migration Inflow (per 10,000 people)	-0.000758
	(0.00538)
Agricultural output/ GDP	0.129
	(0.0879)
Log (Production of Electricity)	1.855
	(1.489)
Average utility expense (share of HH Income)	0.0941
	(0.0217)
Instrumental variables:	
Bodman Index	4.278
	(1.414)
Log (Average MW across neighbors)	9.678
	(0.293)

Note: All monetary variables are in 2010 prices.

Table 2—Unemployment effects of minimum wages across Russian regions

	Panel A: The Second Stage				
	Unemployment rate				Informality
	Overall	15–19	20–29	30+	
	(1)	(2)	(3)	(4)	(5)
	Panel A: The second stage				
$\log MW$	0.003	0.046*	0.056**	0.007	0.137**
	(0.013)	(0.022)	(0.012)	(0.011)	(0.014)
	Panel B: The first stage				
	$\log MW$				
Bodman Index	0.172**	0.174**	0.175**	0.172**	0.173**
	(0.021)	(0.022)	(0.022)	(0.021)	(0.021)
Controls	YES	YES	YES	YES	YES
Panel FE	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES
Obs.	321	314	312	319	312
Adjusted R-squared	0.497	0.208	0.058	0.455	0.614

Notes: The table reports the unemployment effects of the minimum wage estimated by 2SLS. The outcome variables are Column (1) – the overall unemployment rate, Columns (2),(3),(4) – the unemployment rates among those aged between 15-19, 20-29, and 30+, respectively. Column (5) reports the informality rate's response to the minimum wage. The unit of observation is region-year. Controls include the log of real GDP per capita, CPI, and log of the working-age population. Standard errors in parentheses are clustered by regions. + p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table 3—Testing overidentifying restrictions

	Panel A: The Second Stage				
	Unemployment rate				Informality
	All	15–19	20–29	30+	
	(1)	(2)	(3)	(4)	(5)
$\log MW$	0.003	0.043*	0.056***	0.008	0.140***
	(0.012)	(0.022)	(0.012)	(0.011)	(0.014)
	Panel B: The First Stage				
	$\log MW$				
Bodman index	0.176***	0.177***	0.178***	0.176***	0.176***
	(0.022)	(0.023)	(0.023)	(0.022)	(0.023)
Log ($\overline{MW}$ neighbors)	0.215+	0.229*	0.214+	0.216+	0.193+
	(0.102)	(0.043)	(0.102)	(0.102)	(0.098)
Controls	Y	Y	Y	Y	Y
Panel FE	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y
F-stat of excluded IVs	32.91	30.12	30.03	32.76	30.79
P-value of Hansen J	0.458	0.384	0.31	0.757	0.34
Observations	320	313	311	318	311

Notes: The table reports the unemployment effects of the minimum wage estimated by 2SLS. The outcome variables are column (1) – the overall unemployment rate, columns (2),(3),(4) – the unemployment rates among those aged between 15–19, 20–29, and 30+, respectively. Column (5) reports the informality rate's response to the minimum wage. The unit of observation is region-year. Controls include the log of real GDP per capita, CPI, and log of the working-age population. Standard errors in parentheses are clustered by regions. + p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table 4—Summary statistics for a case study of Kamchatka

	Mean (Std.)	Mean (Std.)
Outcome variables:		
Unemployed	0.0613	
	(0.240)	
Informally Employed		0.117
		(0.321)
Controls:		
Kamchatka	0.589	0.583
	(0.492)	(0.493)
Grad School	0.00217	0.00225
	(0.0465)	(0.0474)
College	0.268	0.278
	(0.443)	(0.448)
Mid professional	0.237	0.244
	(0.425)	(0.430)
Primary professional	0.195	0.196
	(0.396)	(0.397)
High School	0.234	0.222
	(0.423)	(0.416)
Middle School	0.0591	0.0524
	(0.236)	(0.223)
Male	0.497	0.494
	(0.500)	(0.500)
Youth	0.205	0.194
	(0.404)	(0.395)
Age	41.44	41.76
	(12.33)	(12.19)
Number of children <18 yo	0.571	0.571
	(0.829)	(0.825)
Married	0.625	0.639
	(0.484)	(0.480)
Urban	0.588	0.601
	(0.492)	(0.490)
Observations	36,400	34,170

Table 5—Employment effects in Kamchatka

	(1)	(2)	(3)	(4)	(5)	(6)
	Unemployed			Informal employment		
Post	-0.011*	-0.014**	-0.011	0.006	0.004	-0.004
	(0.004)	(0.004)	(0.011)	(0.007)	(0.007)	(0.018)
Treated	0.015**	0.016**	0.051**	0.070**	0.072**	-0.006
	(0.005)	(0.005)	(0.018)	(0.008)	(0.008)	(0.028)
Post × Treated	-0.002	0.002	0.011	0.057**	0.061**	0.063*
	(0.007)	(0.007)	(0.016)	(0.012)	(0.012)	(0.030)
Youth	0.052**	-0.038**	-0.038**	0.017+	-0.025*	-0.024*
	(0.007)	(0.010)	(0.010)	(0.009)	(0.012)	(0.012)
Post × Youth	-0.011	-0.013	-0.013	-0.025+	-0.027*	-0.027*
	(0.009)	(0.009)	(0.009)	(0.012)	(0.012)	(0.012)
Treated × Youth	0.015	0.029**	0.029**	0.015	0.029*	0.028*
	(0.011)	(0.010)	(0.010)	(0.013)	(0.013)	(0.013)
Post × Treated × Youth	0.032*	0.029*	0.028*	0.042*	0.040*	0.040*
	(0.014)	(0.013)	(0.013)	(0.018)	(0.018)	(0.018)
Controls	NO	YES	YES	NO	YES	YES
Region-Quarters FE	NO	NO	YES	NO	NO	YES
Flexible Trends	NO	NO	YES	NO	NO	YES
Observations	36,400	36,400	36,400	34,170	34,170	34,170
Adjusted R-squared	0.015	0.054	0.056	0.030	0.050	0.052

Notes: The table reports the employment effects of the minimum wage in Kamchatka (the treated region). The control region is Chukotka. The outcome variables are Columns (1)–(3) – unemployment's dummy (equal to 1 if unemployed, equal to 0 if employed); Columns (4)–(6) – informal employment's dummy (equal to 1 if employed informally, equal to 0 if employed formally). Controls include age, age squared, marriage (dummy), educational attainment (a set of dummies), male (dummy), urban residence (dummy), number of kids under 18 years old. "Flexible Trends" refers to region-specific quartic time (month of the year) trends. Standard errors in parentheses are clustered by [month of the year X region]. + p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table 6—Falsification test

	(1)	(2)	(3)	(4)	(5)	(6)
		Unemployed			Informal employment	
Placebo × Treated × Youth	-0.009	-0.018	-0.020	0.017	0.012	0.013
	(0.023)	(0.022)	(0.022)	(0.028)	(0.028)	(0.029)
Controls	NO	YES	YES	NO	YES	YES
Region-Quarters FE	NO	NO	YES	NO	NO	YES
Flexible Trends	NO	NO	YES	NO	NO	YES
Observations	18,274	18,274	18,274	17,078	17,078	17,078
Adjusted R-squared	0.011	0.048	0.051	0.018	0.036	0.039

Notes: The table reports the falsification test. The control region is Chukotka. The outcome variables are Columns (1)–(3) – unemployment's dummy (equal to 1 if unemployed, equal to 0 if employed); Columns (4)–(6) – informal employment's dummy (equal to 1 if employed informally, equal to 0 if employed formally). Controls include age, age squared, marriage (dummy), educational attainment (a set of dummies), male (dummy), urban residence (dummy), number of kids under 18 years old. Flexible Trends refers to region-specific quartic time (month of the year) trends. The variable "Placebo" is a binary indicator of 2011. Standard errors in parentheses are clustered by [month of the year X region]. + $p<0.1$; * $p<0.05$; ** $p<0.01$; *** $p<0.001$.

Table 7—Cross-industry analysis

	Log(Youth)	Log(Informal)	Log(Investments)	σ
	(1)	(2)	(3)	(4)
Finance	0.521*	-1.101*	0.148**	
	(0.249)	(0.457)	(0.050)	
Agriculture & Fishery	-1.212***	0.155	-0.170**	0.309*
	(0.287)	(0.446)	(0.064)	(0.133)
Mining	-1.027	0.332	-0.044	0.717*
	(0.675)	(0.809)	(0.071)	(0.284)
Manufacturing	-0.744**	1.072*	-0.170**	0.574***
	(0.232)	(0.433)	(0.054)	(0.149)
Electricity, Gas and Water	-0.724*	0.375	-0.124*	0.233
	(0.297)	(0.762)	(0.051)	(0.397)
Building and Construction	-0.525*	1.000*	-0.229***	0.387***
	(0.267)	(0.441)	(0.066)	(0.101)
Wholesale and Retail	-0.613*	0.981*	-0.158**	0.094
	(0.240)	(0.418)	(0.056)	(0.088)
Hotels & Restaurants	-0.186	2.453***	-0.215*	-0.093
	(0.343)	(0.533)	(0.096)	(0.257)
Transportation & Communication	-0.597**	0.861*	-0.181***	-0.030
	(0.219)	(0.386)	(0.054)	(0.270)
Real Estate	-0.490+	1.372*	-0.187***	0.328
	(0.279)	(0.617)	(0.051)	(0.224)
Education	-0.863***	-0.126	-0.010	0.274
	(0.238)	(0.489)	(0.050)	(0.193)
Health	-0.844***	-0.278	-0.052	0.723***
	(0.240)	(0.558)	(0.055)	(0.120)
Other	-0.615**	1.279**	-0.247***	0.367*
	(0.230)	(0.435)	(0.056)	(0.177)
Controls	Y	Y	Y	
Panel, Year FE	Y	Y	Y	
Observations	3,131	2,797	2,955	

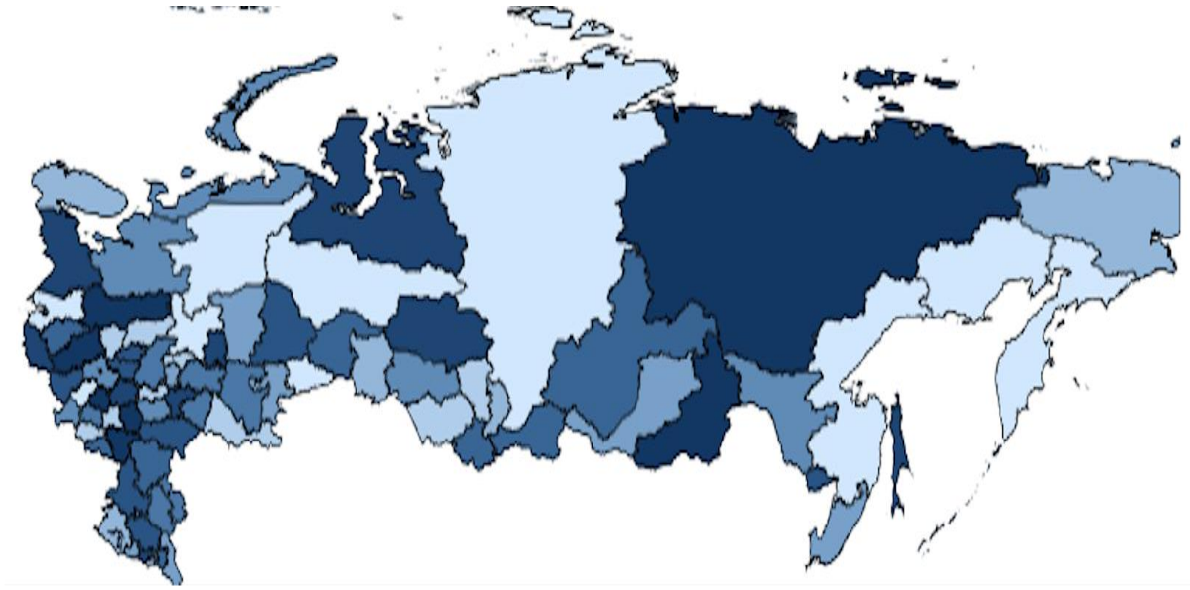
Notes: Columns (1) – (3) report the employment's (measured as log number of workers aged below 30), informal employment's and capital investments' (measured as log investment in the capital) responses to the minimum wages. Controls include the log of real GDP per capita, CPI, and log of the working-age population. Column (4) reports the estimated elasticity of substitution for every industry. All the monetary variables are in 2010 prices. Standard errors in parentheses are clustered by regions. + p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table 8 - Long-run Minimum Wage Elasticities for Share of Individuals with Family Income below Multiples of Federal Poverty Threshold

Family income cutoff	The U.S.	Russia
(1)	(2)	(3)
0.50	-0.455+	-0.048
	(0.247)	(0.058)
0.75	-0.461*	-0.071
	(0.186)	(0.043)
1.00	-0.446**	-0.070*
	(0.137)	(0.033)
1.25	-0.294*	-0.036
	(0.103)	(0.024)
1.50	-0.156*	-0.025
	(0.093)	(0.021)
1.75	-0.167*	-0.022
	(0.084)	(0.018)

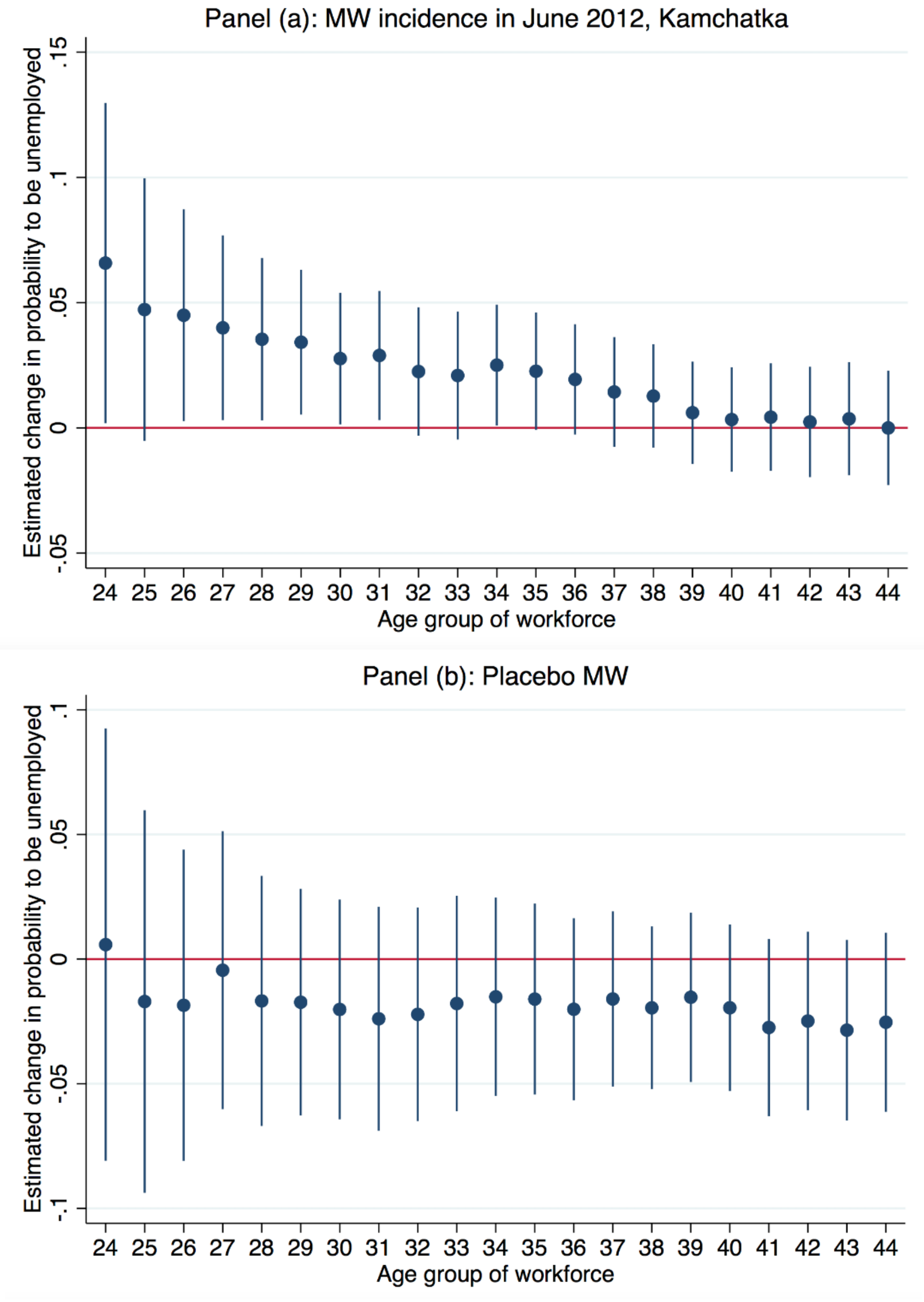
Notes: This table reports the long-run (3 years) elasticities for share with family income under various multiples of the federal poverty threshold. The multiples are reported in column (1). Column (2) shows The U.S. estimates obtained from Dube (2019). Column (3) shows the authors' estimates obtained from the Russian data. Standard errors clustered by states (column 2) or regions (column 3) are in parentheses. + $p<0.1$; * $p<0.05$; ** $p<0.01$.

Figure 1: The geography of Bodman Index



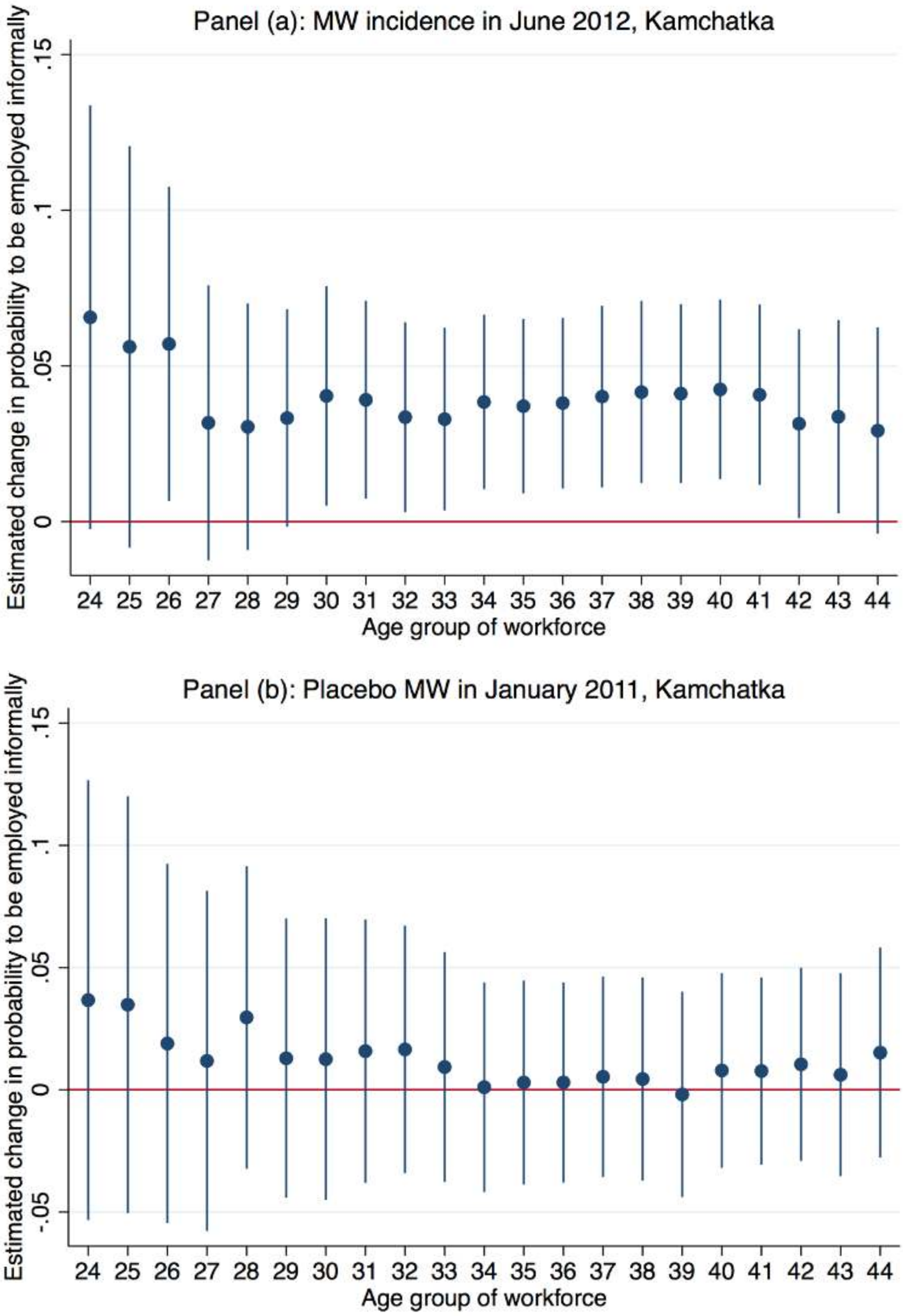
Notes: The figure plots the residuals (in the latest cross-section of the panel employed) resulted from regressing the Bodman index on regional and year fixed effects. Darker zones indicate higher winter severity after controlling for regional and year fixed effects.

Figure 2: The estimated unemployment effects by a varying definition of the young workforce



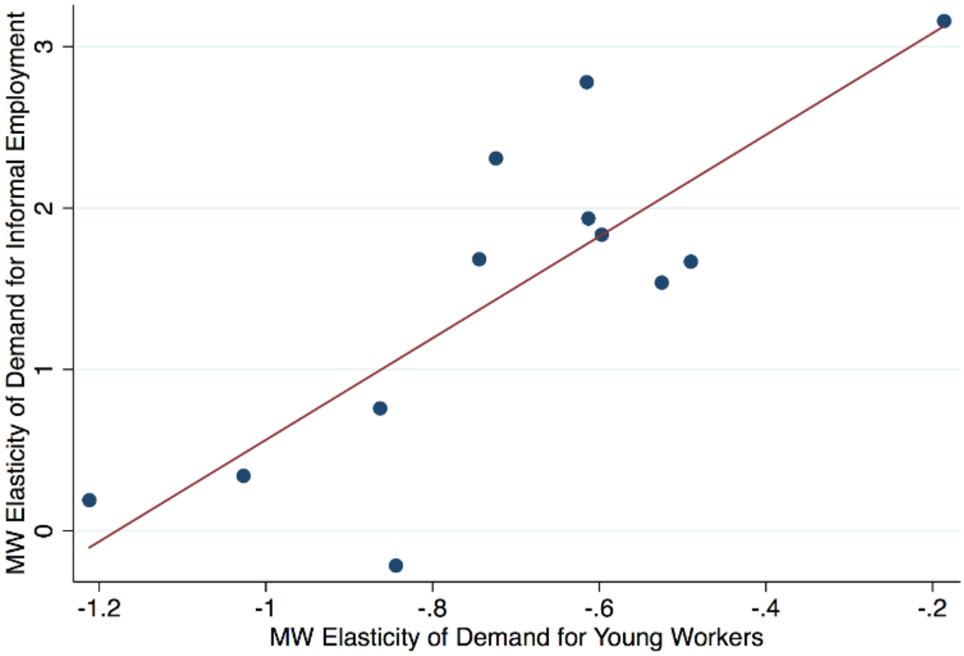
Notes: Bold points in Panel (a) [Panel (b)] indicate the OLS estimates of the parameter on the $Treated \cdot Youth \cdot Post$ [$Treated \cdot Youth \cdot Fals$] term. Vertical solid blue lines are 95% confidence intervals. The variables' definitions are given in the text.

Figure 3: The estimated informality effects by a varying definition of the young workforce



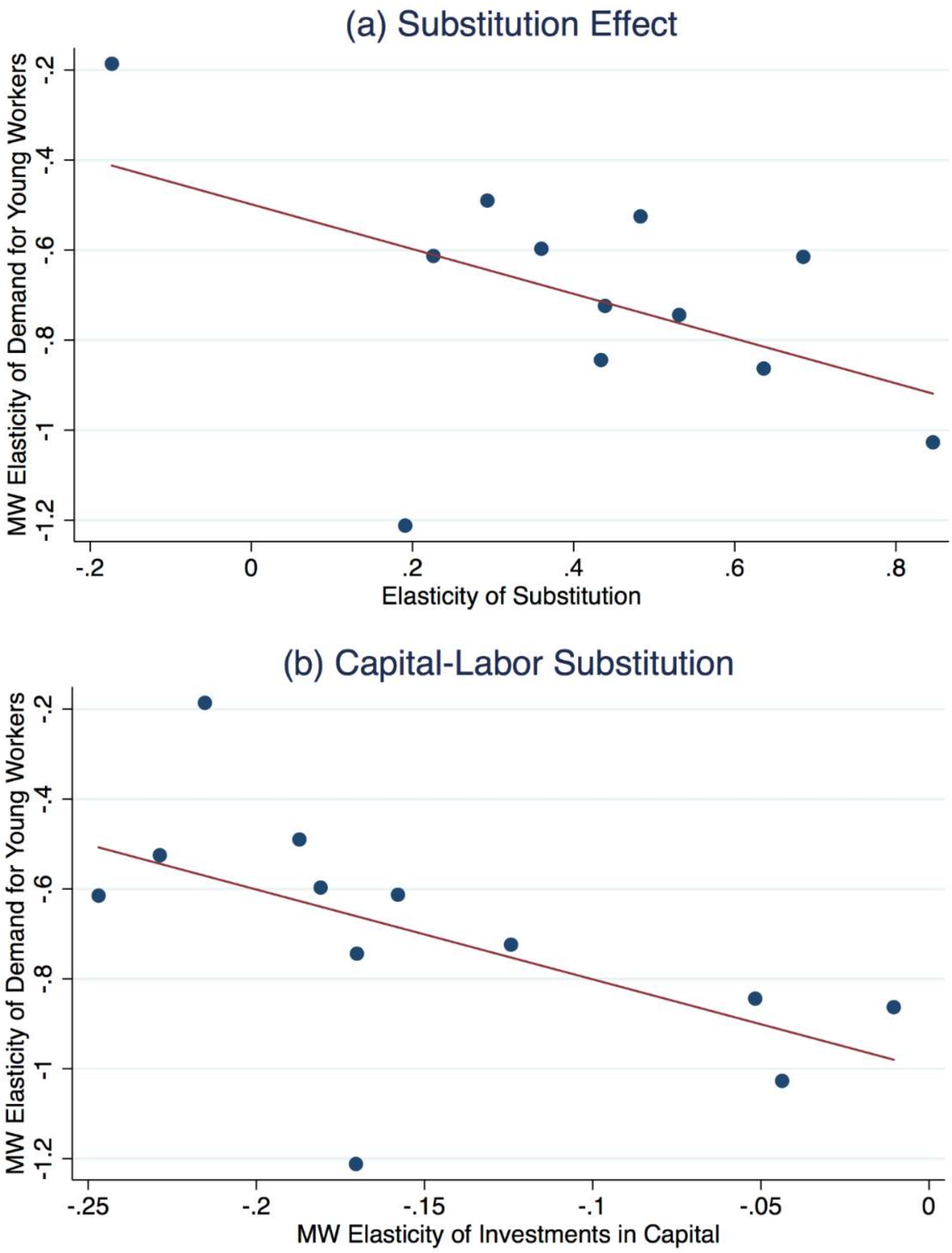
Notes: Bold points in Panel (a) [Panel (b)] indicate the OLS estimates of the parameter on the $Treated \cdot Youth \cdot Post$ [$Treated \cdot Youth \cdot Fals$] term. Vertical solid blue lines are 95% confidence intervals. The variables' definitions are given in the text.

Figure 4: Young workers-informality substitution



Notes: This figure plots the point estimates in Column (2) against those in Column (1) of Table 7, to visually assess whether industries substitute informally–employed workers for young workers in response to the MW policy.

Figure 5: Capital-labor substitution



Notes: This figure visually accesses whether capital-labor substitution exists. Both panels plot columns of the point estimates in Table 7. Panel (a) plots the point estimates in Column (1) against those in Column (4). Panel (b) plots the point estimates in Column (1) against those in Column (3).

Appendix A

A.1. Exclusion Restriction

In this section, we test whether our IV (Bodman Index) is excluded from the unemployment equation. We consider five potential channels through which the Bodman index might affect regional unemployment rates. These are (i) Workers' health, (ii) Migration, (iii) Performance of certain industries, (iv) Regional cost of electricity, and (v) Workers' decision to participate in the labor force.

We hence consider 6 outcomes: (i) log population-to-doctors ratio, (ii) log number of registered diseases per 1,000 people, (iii) net migration inflow (per 10,000 people), (iv) agricultural output to GDP ratio, (v) log electricity production (Billions of kW-hour), (vi) average household utility expenses normalized by average household income, (vii) labor force participation rate. We then test the following reduced-form relationship:

$$y_{it} = \pi Bodman_{it} + X_{it}'\vartheta + \kappa_i + \lambda_t + w_{it}, \tag{A1}$$

where y_{it} is the set of outcomes mentioned before, X_{it}' — covariates included in our model (1), κ_i, λ_t – sets of regional and year effects, w_{it} — the error term. Our main regressor is $Bodman_{it}$.

Table A.1 presents the results. We note that the estimated coefficients are statistically insignificant and numerically close to zero. That is, conditional on covariates, regional and year fixed effect, Bodman Index is unlikely to affect regional health indicators, patterns of migration, agriculture, electricity's production, or utility expenses.

As an additional check, we include all six indicators as a set of additional controls in our estimations of the model (1). Table A.2 reports the results of this exercise that should be contrasted with the ones reported in Table 2 of the main body of the research. We note that the coefficients remained intact. We conclude that conditional on covariates, the minimum wage is the only channel through which Bodman Index affects regional unemployment rates.

Table A.1: The relationship between Bodman Index and regional indicators

Outcome	(1) Doctors	(2) Disease	(3) Migration	(4) Agricultural output	(5) Electricity	(6) Utility	(7) LFP
Bodman Index	0.004	0.000	0.000	-0.003	-0.011	-0.001	-0.003
	(0.006)	(0.008)	(0.001)	(0.003)	(0.036)	(0.002)	(0.003)
Observations	322	318	322	311	302	322	316
Controls	YES	YES	YES	YES	YES	YES	YES
Panel FE	YES	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES	YES
Mean of the dep. var.	4.49	6.68	0	0.13	1.86	0.09	0.86

Notes: The table reports the estimates of equation (A1). The unit of observation is region-year. The controls include the log of real GDP per capita, CPI, and the log of the working-age population. All monetary variables are in 2010 prices. Standard errors in parentheses are clustered by regions. + p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A.2: Unemployment effects of minimum wages across Russian regions: Extended models

Outcome	Unemployment rate				Informality
	All	15–19	20–29	30+	
	(1)	(2)	(3)	(4)	(5)
$\log MW$	0.000	0.044+	0.050**	0.002	0.126**
	(0.015)	(0.023)	(0.012)	(0.015)	(0.012)
Controls	YES	YES	YES	YES	YES
Panel FE	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES
Additional Controls	YES	YES	YES	YES	YES

Notes: The table reports the unemployment effects of the minimum wage estimated by 2SLS. The outcome variables are Column (1) — the overall unemployment rate, Columns (2),(3),(4) – the unemployment rates among those aged between 15-19, 20-29, and 30+, respectively. Column (5) reports the informality rate's response to the minimum wage. The unit of observation is region-year. Controls include the log of real GDP per capita, CPI, and log of the working-age population. Additional controls include the log population-to-doctor ratio, log number of newly registered diseases per 1,000 people, net migration inflow, agricultural output normalized by GDP, log electricity production (kW-hour), labor force participation rate, and average household utility expenses normalized by average household income. All monetary variables are in 2010 prices. Standard errors in parentheses are clustered by regions. + p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Appendix B

The goal of this section is to summarize those properties of CES production functions, that are central to our empirical analysis.
Consider an arbitrary industry operating in a single market. The industry's technology maps two inputs K, L into a single output according to the linearly homogenous function given by $Y = F(K, L)$. The demand for the industry's final good is governed by $Y = \phi(p)$, where p denotes price. Hence, in equilibrium the following holds:

$$Y = F(K, L) = \phi(p) \tag{1}$$

The first order conditions for the profit maximization imply:

$$pF_L = w \;\; or \;\; F_L = \frac{w}{p} \tag{2}$$

$$pF_K = r \;\; or \;\; F_K = \frac{r}{p} \tag{3}$$

The following properties of the production functions will prove to be useful in the further narration:

$$F_{LL} = -\frac{K}{L}F_{LK} \;\; and \;\; F_{KK} = -\frac{L}{K}F_{LK} \tag{4}$$

Additionally, we invoke the following definition of the elasticity of substitution first introduced by Hicks (1932):

$$\sigma = \frac{F_L F_K}{Y F_{LK}} \tag{5}$$

Hence,

$$\frac{1}{\sigma} = \frac{Y F_{LK}}{F_L F_K} \tag{6}$$

and

$$F_{LK} = \frac{F_L F_K}{Y\sigma} \tag{7}$$

Combining (4) and (7), rewrite F_{LL} and F_{KK} as:

$$F_{LL} = -\frac{K}{L}\frac{F_L F_K}{Y\sigma} \;\; and \;\; F_{KK} = -\frac{L}{K}\frac{F_L F_K}{Y\sigma} \tag{8}$$

Next, differentiate (1)-(3) with respect to w to obtain:

$$F_L\frac{\partial L}{\partial w} + F_K\frac{\partial K}{\partial w} = -\eta\frac{\partial p}{\partial w}\frac{Y}{p} \tag{9}$$

$$1 = F_L\frac{\partial p}{\partial w} + pF_{LL}\frac{\partial L}{\partial w} + pF_{LK}\frac{\partial K}{\partial w} \tag{10}$$

$$0 = F_K\frac{\partial p}{\partial w} + pF_{LK}\frac{\partial L}{\partial w} + pF_{KK}\frac{\partial K}{\partial w}, \tag{11}$$

where $\eta \equiv -\frac{p}{Y}\frac{dY}{dp}$ - a negative of the market's price elasticity of demand for the industry's final good.

We now focus on equations (9), (10) and (11) one by one.

We note, that $F_L = \frac{w}{p}$ and $F_K = \frac{r}{p}$, as (2), (3) state. Hence, equation (9) can be rewritten as follows:

$$\eta\frac{\partial p}{\partial w}\frac{Y}{p} + F_L\frac{\partial L}{\partial w} + F_K\frac{\partial K}{\partial w} = \eta Y\frac{\partial p}{\partial w} + w\frac{\partial L}{\partial w} + r\frac{\partial K}{\partial w} = 0 \tag{12}$$

Moving next, we work out equation (10). Multiplying both sides by $\frac{Yp}{w}\sigma$ and rearranging the terms, we get:

$$Y\sigma\frac{\partial p}{\partial w} - r\frac{K}{L}\frac{\partial L}{\partial w} + r\frac{\partial K}{\partial w} = \frac{Yp}{w}\sigma \tag{13}$$

Equation (11), being a "mirror image" of the preceding identity, can be also viewed as:

$$Y\sigma\frac{\partial p}{\partial w} + w\frac{\partial L}{\partial w} - w\frac{L}{K}\frac{\partial K}{\partial w} = 0 \tag{14}$$

Hence, equation (12), (13) and (14) can be viewed as the following system of 3 linear equations:

$$\eta Y\frac{\partial p}{\partial w} + w\frac{\partial L}{\partial w} + r\frac{\partial K}{\partial w} = 0$$

$$Y\sigma\frac{\partial p}{\partial w} - r\frac{K}{L}\frac{\partial L}{\partial w} + r\frac{\partial K}{\partial w} = \frac{Yp}{w}\sigma$$

$$Y\sigma\frac{\partial p}{\partial w} + w\frac{\partial L}{\partial w} - w\frac{L}{K}\frac{\partial K}{\partial w} = 0$$

with 3 unknowns: $[(\frac{\partial p}{\partial w}), (\frac{\partial L}{\partial w}), (\frac{\partial K}{\partial w})]$. The solution for labor demand's slope is given by:

$$\frac{\partial L}{\partial w} = -\frac{L}{w}(\frac{wL}{Yp}\eta + \frac{rK}{Yp}\sigma) \tag{15}$$

Equivalently, we can obtain the solution for $\frac{\partial K}{\partial w}$ as:

$$\frac{\partial K}{\partial w} = \frac{KL}{Yp}(\sigma - \eta) \tag{16}$$

Multiplying both sides of (15) and (16) by $(\frac{w}{L})$, we obtain:

$$\beta = \frac{wL}{Yp}\epsilon_p^Y - \frac{rK}{Yp}\sigma, \tag{17}$$

$$\alpha = \frac{wL}{Yp}(\sigma - \eta) \tag{18}$$

where β, α and ϵ_p^Y are industry's wage elasticity of demand for labor,capital and the market's price elasticity of demand for the final good respectively. The intuition behind the wage elasticity of labor demand given by (17) is this: if the demand for the industry's final good is inelastic, the industry may pass the cost of higher wages to the consumers. Additionally, the employment response of the minimum wages depends upon the underlying production technology. If $\sigma \to 0$, capital and labor serve as perfect complements. Hence, the industry will not be willing to fire workers, despite their higher cost. Alternatively, if $\sigma \to \infty$, the industry's isoquant "linearizes", which implies the presence of the ability to easily substitute capital for workers.

Appendix C

Proposition 1 (ACMS, 1961)**.** *If the production function F(K,L) exhibits constant returns to scale, then the elasticity of substitution (σ) can be derived as:*

$$\sigma = \frac{dy}{dw}\frac{w}{y}, \tag{1}$$

where

$y = \frac{Y}{L}$*, and* w *- wage.*

Proof. If the assumed production function $Y = F(K, L)$ exhibits CRS, then we can say $y = f(k)$, where $y = \frac{Y}{L}$ and $k = \frac{K}{L}$. Then, given the profit-maximizing conditions, the marginal products of capital and labor can be viewed as:

$$y_K = f'(k) = r \tag{2}$$

and

$$y_L = y - kf'(k) = w, \tag{3}$$

where r, w are input prices of capital and labor, respectively.
ACMS depart from the following definition of the elasticity of substitution:

$$\sigma = \frac{-f'(f - kf')}{kff''} \tag{4}$$

Now, differentiate (3) with respect to w to obtain:

$$1 = f'\frac{dk}{dy}\frac{dy}{dw} - kf''\frac{dk}{dy}\frac{dy}{dw} - f'\frac{dk}{dy}\frac{dy}{dw} \tag{5}$$

Normalizing equation (5) by $\frac{dy}{dw} \neq 0$, we obtain

$$\frac{dw}{dy} = \frac{dk}{dy}(f' - kf'' - f') = \frac{dk}{dy}(-kf'') \tag{6}$$

Note, that $\frac{dk}{dy} = \frac{1}{f'}$, so that equation (6) gives raise to the following identity:

$$\frac{dy}{dw} = \frac{-f'}{kf''} \tag{7}$$

The latter, in turns, implies that equation (4), given $y_L = f - kf' = w$, can be viewed as:

$$\sigma = \frac{-f'}{kf''} \cdot \frac{(f - kf')}{f} = \frac{dy}{dw}\frac{w}{y} \tag{8}$$

□