DO MINIMUM WAGES LEAD TO JOB LOSSES? EVIDENCE FROM OECD COUNTRIES ON LOW-SKILLED AND YOUTH EMPLOYMENT

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The author investigates effects of minimum wage rates on low-skilled, female low-skilled, and youth employment. The sample consists of 19 Organisation for Economic Co-operation and Development (OECD) countries from 1997 to 2013 for low-skilled workers and from 1983 to 2013 for young workers. Six different static or dynamic estimation approaches are applied on different versions of the specifications, controlling for up to quadratic time trends. The author further investigates the effects over the long run and over the business cycle as well as the effects of high minimum wages and of institutional complementarities. The findings provide little evidence of substantial disemployment effects for low-skilled, female low-skilled, or young workers. The estimated employment elasticities are small and statistically indistinguishable from zero. The author then considers why his results on youth employment differ from those of Neumark and Wascher (2004), showing that they overstate precision and that small changes in their specifications lead to minimum wage effects close to zero.

This article investigates whether cross-country time-series evidence occurs for what Neumark and Wascher (2004: 224) called the "consensus view" that minimum wages reduce employment among lower-skilled workers. This consensus view, however, has been challenged by new evidence. Several recent studies focusing on the United States suggest modest or no minimum wage effects on the employability of restaurant workers and of young workers (see Allegretto, Dube, Reich, and Zipperer 2017 for a

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detailed discussion). Also the introduction of a statutory minimum wage in the United Kingdom (Metcalf 2008) and Ireland (O'Neill, Nolan, and Williams 2006) in the late 1990s has led to tiny or no disemployment effects. The same seems to be true for Germany, which introduced a minimum wage in 2015 (vom Berge et al. 2016). Recent case studies investigating huge and persistent increases in the minimum wage in Hungary (Harasztosi and Lindner 2015) and the Czech and Slovak Republics (Eriksson and Pytlikova 2004) in the early 2000s found only very small disemployment effects. Similarly, a strong increase in the youth minimum wage in Portugal in the mid-1980s had small effects on youth employment (Portugal and Cardoso 2006).

Although national studies are able to pin down minimum wage effects very precisely, they are severely limited in other dimensions. For example, they typically observe a small variation in the treatment variable, and treatment occurs largely in institutionally similar local labor markets. Thus, it is unclear to what extent findings for some Organisation for Economic Co-operation and Development (OECD) countries are generalizable to others that may differ in terms of the minimum wage and the institutional structures of the labor market. Cross-country studies allow answering such questions by investigating the effects of a large variation in the minimum wage, ranging from approximately 30% of the median wage in some countries to 70% in others and to test for complementarities between the minimum wage and other labor market institutions.

To date, relatively few cross-country studies on disemployment effects of minimum wages are available, and they focus on only a small set of OECD countries. Elmeskov, Martin, and Scarpetta (1998) and Bassanini and Duval (2006) investigated the impact of minimum wages on total unemployment, Neumark and Wascher (2004) and Bassanini and Duval (2006) explained youth employment, and Addison and Ozturk (2012) focused on female employment. OECD (1998) investigated each of the above groups individually. The results of this literature are inconclusive: Neumark and Wascher (2004) and Addison and Ozturk (2012) found that minimum wages led to lower employment of youth and women, Elmeskov et al. (1998) and Bassanini and Duval (2006) found no adverse impact of the minimum wage on total unemployment and youth employment, and OECD (1998) presented evidence that minimum wages decreased teen employment, with no robust effect on any of the other groups.

None of these previous studies investigated the impact of minimum wages on low-skilled workers directly. Low educational attainment is a dominant characteristic of minimum wage earners in OECD countries (OECD 1998, 2015a; Rycx and Kampelmann 2013). Recently, OECD started compiling data on adult employment by level of educational attainment. Data are also available for female low-skilled workers, for whom the bite of minimum wages is especially strong, as their incidence of earning wages around the legal minimum occurs more frequently than it does for men (OECD 1998, 2015a; Rycx and Kampelmann 2013). These data can address an important gap in the existing literature. Data are further available for various Eastern European countries not included in previous studies. I apply a large set of estimation techniques, such as two-way fixed effects, a data-driven approach for selecting relevant covariates, an instrumental variable approach to address the potential endogeneity of the minimum wage variable, two dynamic estimators, and a specification in first differences. I address heterogeneous trends across countries by including up to quadratic country-specific time trends and test for heterogeneity of the treatment effect by minimum wage levels, by the phase of the business cycle, and by other labor market institutions. Finally, I assess long-run effects of minimum wage hikes.

Because of their low level of experience, a large share of young workers earn wages at or only slightly above the minimum wage in many countries (OECD 1998, 2015a; Rycx and Kampelmann 2013). Thus, young workers often serve as a proxy for low-skilled workers in empirical studies. I reassess the robustness of the results of previous studies regarding the impact of minimum wages on youth employment by extending the analysis over a more recent time period, by including additional OECD countries, and by applying a larger variety of statistical techniques. I also directly reassess the evidence presented by Neumark and Wascher (2004) with their original data, as well as newly updated data. Finally, I present results on youth employment including a smaller set of controls for a larger sample of 24 OECD countries, spanning the period from 1970 to 2013.

A Survey of the Cross-Country Literature

Starting with the publication of the well-known textbook *Unemployment* by Layard, Nickell, and Jackman (1991) and Lazear's (1990) seminal empirical study, a rich literature has emerged explaining cross-country determinants of unemployment with structural labor market characteristics (see, e.g., Scarpetta 1996; Nickell 1997; Elmeskov et al. 1998; Blanchard and Wolfers 2000; Baccaro and Rei 2007; Bassanini and Duval 2009).

OECD (1998) applied this approach to test for the impact of minimum wages on employment-to-population ratios of teenagers, young adults, and prime-age adults in nine OECD countries over the period 1975 to 1996 for both men and women, jointly and separately. Employment-to-population ratios were explained by 1) the minimum wage as a share of the median wage;[1] 2) institutional control variables—such as unemployment benefits, union density, and the tax wedge; 3) country and time dummies and country-specific time trends; and 4) either the prime-age male unemployment rate or the output gap as a control for cyclical fluctuations. OECD (1998) calculated

[1]In all cross-country studies cited in this article, the minimum wage as a share of the median wage is applied as the minimum wage variable to "measure the extent to which the minimum wage cuts into the wage distribution, and to capture variation in the relative prices of less-skilled and more-skilled labor induced by minimum wages" (Neumark and Wascher 2004: 226).

the minimum-to-median wage ratio separately for all relevant demographic subgroups, using the corresponding average wage for each group as well as taking account of subminimum wages for young workers where they exist. They presented serial correlation and heteroscedasticity robust results, as tests suggest the presence of serial correlation in the residuals.

OECD (1998) found employment elasticities of minimum wages mostly close to zero for adults, and small negative but mostly insignificant elasticities, ranging from −0.14 to −0.03, for young workers who were 20 to 24 years old. For teenagers, they found largely negatively significant employment elasticities between −0.58 and −0.27. These employment elasticities tend to be noticeably smaller and very close to zero if Spain and Portugal are not omitted from the sample, and they are larger for females. OECD (1998) stressed that their results on teenage employment were sensitive to the inclusion of time trends and argued that this might be an indication for missing variables. According to OECD (1998), an obvious candidate for a missing variable was school enrollment, as "[i]n many countries, the employment-to-population ratios for teenagers and youth have declined substantially while their participation in education has been rising" (ibid.: 70). They further found little evidence that disemployment effects were larger in high minimum-wage countries than they were in low minimum-wage countries. They concluded tentatively that minimum wages might decrease employment of teenagers.

Elmeskov et al. (1998) explained unemployment with institutional labor market characteristics for a set of OECD countries from 1983 to 1995. In one specification they included the statutory minimum wage as a share of the median wage as an explaining factor for unemployment in nine OECD countries. They estimated a random effects model and controlled for the output gap, unemployment benefits, union density, employment protection legislation, the tax wedge, and several characteristics of the wage-bargaining regime, such as wage-bargaining coordination and centralization. The minimum wage variable was found to have a statistically insignificant coefficient virtually equal to zero.

Neumark and Wascher (2004) estimated the impact of minimum wages on youth and teen employment-to-population ratios for an unbalanced panel of 17 OECD countries for the period 1976 to 2000. Contrary to the previous studies, they also included countries with bargained wage minima in their sample. They presented ordinary least squares (OLS), fixed effects, and difference general methods of moments (GMM) estimates and reported consistently and mostly statistically significant negative coefficients of the minimum wage variable. Neumark and Wascher concluded that "[i]n general, our results provide evidence that minimum wages tend to reduce employment rates among the youth population" (2004: 243). Their estimated employment elasticities of minimum wage hikes lie between −0.28 and −0.13 in their baseline models.

One potential criticism of Neumark and Wascher's study is that they overstated the precision of their estimates, as they do not correct standard errors for heteroscedasticity or serial correlation. Bertrand, Duflo, and Mullainathan (2004) illustrated the pitfalls of ignoring serial correlation in panel data. Their simulation results suggested that standard errors are heavily downwardly biased, especially for persistent and long time series, as in the case of Neumark and Wascher (2004).

In my replication of Neumark and Wascher (2004) with the original data (see online Appendix C), I thus cluster standard errors, which is the standard procedure in the literature to correct for serial correlation (e.g., Cameron, Gelbach, and Miller 2008). This method increases the standard errors by 100 to 200%, depending on the specification, and renders the minimum wage effects highly insignificant.

Bassanini and Duval (2006) followed an approach similar to Elmeskov et al. (1998). They investigated the impact of various labor market institutions on unemployment and included some specifications in which they tested for the impact of minimum wages on unemployment. Their sample for the latter covered the period 1982 to 2003 for 10 OECD countries. They estimated fixed-effects specifications with robust standard errors and included control variables for unemployment benefits, the tax wedge, union density, employment protection legislation, a dummy variable for a corporatist wage-bargaining regime, an indicator for product market regulation, the output gap, and time dummies. They did not find a statistically significant impact of the minimum wage on the unemployment rate, with estimates pointing to small positive or negative elasticities. Furthermore, using the same sample, Bassanini and Duval (2006) investigated whether youth employment rates were affected by the minimum wage. Depending on the specification, they controlled for adult employment, or various labor market institutions and the output gap. They clustered standard errors at the country level to account for serial correlation and heteroscedasticity. Perhaps surprisingly, they found "that minimum wage hikes significantly increase youth employment rates" (Bassanini and Duval 2006: 47).

Addison and Ozturk (2012) investigated the prime-age employment rate and labor force participation rate for females for a sample of 16 OECD countries for the period 1970 to 2008. They considered negotiated wage floors as minimum wages. They controlled for the male unemployment rate and also included the fertility rate and the gender wage gap, and an interaction term between the minimum wage and these variables as explaining factors in most of their specifications.[2] In auxiliary regressions, they included further variables, such as employment protection, union density, unemployment benefits, active labor market policy, a dummy for the presence of youth subminimum wages or bargained minimum wages, as well as interaction terms between the minimum wage variable and these regressors to test for institutional complementarities. They presented OLS, fixed effects, and

[2]They reasoned that the higher the gender wage gap, the more female workers are concentrated in low-paying jobs, and the stronger is the impact of the minimum wage.

difference GMM specifications, and they mostly found statistically significant negative effects of the minimum wage. The employment elasticities of minimum wages in their preferred baseline model were approximately -0.35 for the OLS specifications, -0.10 for the fixed-effects estimator, -0.04 for the difference GMM estimator, and more than double that if institutional complementarities were accounted for.[3]

As summarized in Table 1, no evidence supports that minimum wage policies affect total employment or unemployment. For young workers, the findings are mixed, with some estimates showing large negative, others no or even large positive employment effects. For females, the cross-country evidence points to moderate disemployment effects. In addition to different samples and in parts control variables, the studies differ in the definition of the minimum wage; OECD (1998) and Bassanini and Duval (2006) focused on statutory minimum wage policies, whereas Neumark and Wascher (2004) and Addison and Ozturk (2012) also included bargained minimum wages in their analysis.[4] Neumark and Wascher (2004) performed augmented specifications, including a dummy variable for bargained minimum wages and an interaction term with the minimum wage. Their results indicated that unlike statutory minimum wages, bargained minima showed much lower, or mostly even no, disemployment effects. Thus, assorted definitions of the minimum wage variable can be ruled out as explanations for the differences in outcomes for youth employment. Addison and Ozturk (2012) also included bargained minima and showed that effects for legal and bargained minimum wages differ. Contrary to Neumark and Wascher (2004), however, they found that disemployment effects of bargained minima were much higher, about double in size, than were legal ones.

Empirical Approach

To test for the impact of minimum wages on low-skilled or youth employment, I follow the seminal approaches of OECD (1998) and Neumark and Wascher (2004), and expand on their analysis. I specify the following reduced-form unemployment equation:

$$\ln(E_{i,t}) = \beta \ln(MWR_{i,t-1}) + \gamma \ln(X_{i,t}) + \alpha_i + \tau_t + \varepsilon_{i,t} \quad (1)$$
$$i = 1, \ldots, \mathrm{N};\ t = 1, \ldots, \mathrm{T}$$

[3]Addison and Ozturk (2012) noted that they robustify their standard errors against heteroscedasticity. For their dynamic specifications, which they reported as their preferred estimation approach, they did not provide any details. It is, for example, unclear if their instruments passed standard tests of instrument validity, if serial correlation in the residuals was present, and if, or how, they corrected their standard errors (see, e.g., Arellano and Bond 1991; Windmeijer 2005).

[4]See also Bertola, Blau, and Kahn (2007) on the effects of collective bargaining on employment of young or female workers.

Table 1. Summary Results of Cross-Country/Time-Series Minimum Wage Studies

Dependent variable	*Author(s)*	*Year*	*Sample*	*Estimator*	*Disemployment effect of minimum wage*
Youth employment	OECD	1998	9 OECD countries, 1975–1996	FE	Tiny and insignificant for youth; strong and significant for teenagers if Spain and Portugal are excluded, small and mostly insignificant if not
	Neumark & Wascher	2004	17 OECD countries, 1976–2000	OLS, FE, difference GMM	Moderate to strong for youth; tiny to strong for teenagers
	Bassanini & Duval	2006	10 OECD countries, 1982–2003	OLS, FE, SURE	OLS: moderate and insignificant; SURE & FE: strong positive significant
Female employment	OECD	1998	9 OECD countries, 1975–1996	FE	Tiny and insignificant to moderate and significant if Spain and Portugal are excluded (results with Spain and Portugal not reported)
	Addison & Ozturk	2012	16 OECD countries, 1970–2008	OLS, FE, difference GMM	OLS: strong; FE & diff. GMM: tiny to moderate; mostly significant
Total (un-)employment	OECD	1998	9 OECD countries, 1975–1996	FE	Insignificant and close to zero
	Elmeskov et al.	1998	9 OECD countries, 1983–1995	RE	Insignificant and close to zero
	Bassanini & Duval	2006	10 OECD countries, 1982–2003	FE	Insignificant and close to zero

Notes: FE, fixed effects; OLS, ordinary least squares; GMM, generalized method of moments; SURE, seemingly unrelated regression; RE, random effects.

$E_{i,t}$ is the employment-to-population ratio of a sociodemographic group with a high incidence of working at the minimum wage, that is, low-skilled, female low-skilled, or young workers, respectively, in country i and year t. $MWR_{i,t-1}$ is the ratio of the statutory minimum wage of a full-time worker to the median full-time wage, lagged by one period in the static specifications as in Neumark and Wascher (2004) and Addison and Ozturk (2012). $X_{i,t}$ is a vector of controls. Following standard practice, I include the employment rate of a corresponding group with no or marginal policy treatment ($EC_{i,t}$), that is, high-skilled, female high-skilled, or adults, to control for general labor market conditions. In addition, I include controls for institutional labor market characteristics across countries and over time unrelated to minimum wage policies, specifically the unemployment benefit replacement rate ($UB_{i,t}$), employment protection legislation ($EPL_{i,t}$), government spending on active labor market policies ($ALMP_{i,t}$), and union density ($UD_{i,t}$). When explaining youth employment, I follow Neumark and Wascher (2004) and further include the relative cohort size ($cohort_{i,t}$) as regressor. For female low-skilled workers, I follow Addison and Ozturk (2012) and include the fertility rate ($fertility_{i,t}$).[5] All variables are expressed in natural logarithms (ln), which allows interpreting the coefficients as elasticities. Country fixed effects (α_i) are included to purge time-invariant heterogeneity between countries, and time fixed effects (τ_t) to control for common time shocks. Depending on the specification, I also include linear ($\theta_{i1}t$) and quadratic ($\theta_{i2}t^2$) country-specific time trends (see below). To account for heteroscedasticity and within-group serial correlation, I cluster standard errors at the panel level.[6]

I apply a total of six different estimators on each of these samples of low-skilled, female low-skilled, or young workers.[7] The first specification is estimated by applying the canonical two-way fixed-effects (FE) estimator. This is also the baseline approach, which serves as a starting point for the other specifications and for various robustness checks.

[5]Addison and Ozturk (2012) also included the gender wage gap and its interaction with the minimum wage as regressors in many specifications. If the minimum wage increases wages of female workers, which moreover causes some of them to lose their jobs, whereas others might increase their labor supply because of higher wages, thereby altering the composition of the female work force, the wage gap is affected by minimum wage policies and is thus endogenous.

[6]Serial correlation tests indicate that in most of the static models the null hypothesis of no serial correlation is strongly rejected. Simulation results of Bertrand et al. (2004), Hansen (2007), and Cameron et al. (2008) showed that with a cross-sectional dimension of 15-30, clustering standard errors at the panel level might be an appropriate procedure to account for serial correlation. Some studies, however, found relevant over-rejection rates for few clusters, especially in the case of unbalanced data (Cameron and Miller 2015). I conclude that, if anything, my standard errors might be moderately downward biased.

[7]Unlike Neumark and Wascher (2004) and Addison and Ozturk (2012), I do not show OLS results as they are biased if unobserved heterogeneity between countries is correlated with the explaining variables. For comparability, I note, however, that I obtain negative significant minimum wage effects in most OLS specifications explaining low-skilled employment, and positive significant ones when explaining youth employment.

In the second specification, I aim at better capturing time-varying heterogeneity between countries, in line with the argument of Neumark and Wascher concerning "the importance of accounting for institutional and other policy-related differences when using data for different countries" (2004: 244). Therefore I add all bi- and trivariate interactions of the control variables $X_{i,t}$ to the list of controls. I then apply a data-driven selection procedure to pick a sparse list of relevant controls of this set of about 60 characteristics. Specifically, I apply the double selection least absolute shrinkage and selection operator (LASSO) estimation approach described in Belloni, Chernozhukov, and Hansen (2014). This approach chooses covariates to minimize the sum of the squared residuals plus a term that penalizes the size of the model. The latter term guards against overfitting and ensures feasibility of estimation by returning a small set of relevant regressors. When β is causally interpreted, shrinkage or omitted variable bias is a concern, as LASSO drops any variable highly correlated with the treatment variable. To avoid this, Belloni et al. (2014) suggested the double selection or post-LASSO approach, in which LASSO is applied to predict the outcome and the treatment variable separately. In a third step, these double-selected controls are included in a simple regression of the outcome on the treatment variable.[8]

In the third specification, I address the potential endogeneity of the median wage in the minimum-to-median wage ratio (e.g., Card, Katz, and Krueger 1993) by applying a two-stage least squares instrumental variable (2SLS IV) approach, in which $MWR_{i,t}$ is instrumented with the real minimum wage in year t as a share of the average real median wage in t–1 to t–4.[9] Thus, the median wage component in the so-called Kaitz index is instrumented with lagged values. If the results of the FE specifications are driven by variations in median wages, for example, because median wages pick up effects of unaccounted macroeconomic shocks, the 2SLS IV regressions return results different from the two-way FE specifications.

The fourth and fifth specifications include the lagged dependent variable as an additional regressor. I present results of two different dynamic estimation approaches. First, I apply the difference GMM (Arellano and Bond 1991), and second, the system GMM estimator (Arellano and Bover 1995; Blundell and Bond 1998). Both allow expunging endogenous and predetermined components of regressors by utilizing internal instruments. Whereas the difference GMM estimator purges fixed effects by first differencing the equation and uses lagged levels as instruments, the system GMM estimator allows for a larger instrument set after introducing additional restrictions.

[8]The underlying code file "lassoshooting.ado" for the post-LASSO approach is available at Christian Hansen's homepage: http://faculty.chicagobooth.edu/christian.hansen/research/. I use Belloni et al.'s default rule to optimally choose parameters as penalty level. All LASSO and post-LASSO regressions include country and time fixed effects, and depending on the specification, also time trends (which are partialed out in the LASSO specifications using the "controls()" option).

[9]The instruments are highly correlated with the endogenous variables and easily pass standard tests for weak instruments in all 2SLS models (test results not reported).

System GMM is found to perform better for persistent time series in which lagged levels possess little explanatory power over future changes. I apply the one-step procedure with cluster-robust standard errors. To guard against overfitting I limit the instrument count by collapsing the instrument matrix.[10]

Finally, Meer and West (2016) argued that minimum wages in the United States mainly affect employment growth rather than employment levels. Moreover, in some of the system GMM specifications, the lagged dependent variable has a coefficient close to unity. I therefore also estimate a specification in first differences (FD).

The first set of specifications does not include time trends. If minimum wages reduce employment trends of minimum wage workers, the inclusion of such trends may overcontrol and hide the true minimum wage effect (e.g., Dube, Lester, and Reich 2010). Following standard practice (e.g., OECD 1998; Neumark and Wascher 2004; Addison and Ozturk 2012), the same set of specifications is repeated with country-specific linear time trends ($\theta_{i1}t$) added to capture heterogeneous employment trends across countries. Finally, I present results including linear ($\theta_{i1}t$) and quadratic trends ($\theta_{i2}t^2$), following the advice of Wolfers (2006) to include quadratic trends as a robustness check in short panels. The LASSO-specifications with linear and quadratic trends additionally allow for country-specific trends up to 4th order polynomials, to address the argument of Neumark, Salas, and Wascher (2014) on the importance of including high-dimensional time trends. However, 3rd or 4th order polynomials are never picked by the selection process.

I further address issues of heterogeneity and nonlinearity in minimum wage effects. First, I investigate if the effects of minimum wage increases depend on the state of the economy. Specifically, it might be that increases in the minimum wage are much easier for firms to deal with in a period of high aggregate demand. Therefore, I re-estimate the baseline fixed-effects specifications and allow for a separate coefficient of the minimum wage when output is below its potential by adding a recession dummy and the term $\ln(MWR_{i,t-1})^{*} Recession\text{-}Dummy_{i,t}$ to the regression. Second, I test if the effect of the minimum wage depends on other labor market characteristics. I thus include interaction terms between the minimum wage and the other labor market institutions to assess the role of institutional complementarities. Following standard practice, I subtract the sample mean of the variables before creating the interaction terms.[11] Third, I test if high minimum wages have stronger effects than do low ones, as was found for the United States in Zipperer (2014). Thus, I estimate a spline function, which allows

[10]This approach is implemented in the "xtabond2" routine for STATA; see Roodman (2009). I treat all variables as endogenous. The results of the Hansen test of overidentifying restrictions never reject the null and provide support for the validity of the instruments; no regressions reject the null of no second order autocorrelation according to the Arellano-Bond serial correlation test (results not reported).

[11]Therefore, the coefficient of the minimum variable can be readily interpreted as the marginal effect on the outcome variable at its sample mean, when all other covariates are kept constant at their sample means (see, e.g., Baccaro and Rei 2007; Bassanini and Duval 2009).

for lower effects of low minimum wages and higher effects for high minimum wages. For low minimum wages, I allow for a threshold at 40% of the minimum wage as a share of the median wage, by adding $\ln(MWR_{i,t-1}) * MWR < 40\%Dummy_{i,t-1}$ as an explaining variable, which corresponds broadly to the minimum wage level in the Czech Republic, Japan, Spain, and the United States in 2010. Further, for high minimum wages, I allow for a threshold at 55% by adding $\ln(MWR_{i,t-1}) * MWR > 55\%Dummy_{i,t-1}$, which corresponds to the level of Australia, Portugal, and New Zealand in 2010.

Finally, I follow Dube et al. (2010) and Allegretto et al. (2017) to investigate the existence of long-run effects with distributed lags by estimating:

$$(2) \qquad \ln(E_{i,t}) = \sum\nolimits_{r=-2}^{3} \beta_r \ln(MWR_{i,t-r}) + \gamma \ln(X_{i,t}) + \alpha_i + \tau_t + \varepsilon_{i,t}$$
$$i = 1, \ldots, \mathrm{N};\ t = 1, \ldots, \mathrm{T}$$

Thus, I re-estimate the baseline fixed-effects specifications including two leads to capture pre-existing trends (for the years $t + 1$ to $t + 2$), as well as the contemporaneous value and up to three lags (for the years t to t–3) to capture long-run effects over a four-year window.

Data

The sample is an unbalanced panel and covers 19 OECD countries over the maximum time period of 1997 to 2013 in the case of low-skilled labor market outcomes, and 1983 to 2013 in the case of youth employment (see online Appendix A for summary statistics and additional information). Sample size is limited by the availability of employment rates by educational attainment, which are available from the late 1990s onward, and by labor market institutions data, which are available from the early 1980s onward for most variables.[12]

In line with Elmeskov et al. (1998), OECD (1998), and Bassanini and Duval (2006) but in contrast to Neumark and Wascher (2004) and Addison and Ozturk (2012), I do not consider negotiated wage floors as minimum wages, as they "can vary substantially across sectors and often depend on workers' age, experience and qualifications. Such detailed information is rarely available and, in any event, is inherently hard to summarise in a single, cross-country comparable indicator" (Bassanini and Duval 2006: 28). When explaining youth employment, however, I also present results including bargained minimum wages (based on measures constructed by Dolado et al. 1996) for a sample of 24 OECD countries from 1970 to 2013 in online Appendix A.

[12]The countries in the sample are Australia, Belgium, Canada, Czech Republic, France, Greece, Hungary, Ireland, Japan, Korea, Luxembourg, Netherlands, New Zealand, Poland, Portugal, Slovak Republic, Spain, United Kingdom, and United States. For countries that introduced a minimum wage during the period under investigation, Ireland and the United Kingdom, I include only observations where a minimum wage is in place (following Neumark and Wascher 2004).

Data for employment and the minimum wage are from various OECD sources. Low- and high-skilled employment data for those 25 to 64 years old are from the OECD *Education at a Glance* series, which provides this information for three skill groups: below upper secondary (low skilled), upper secondary or postsecondary non-tertiary (medium skilled), and tertiary education (high skilled).[13] I choose high-skilled employment as a control when explaining low-skilled employment, as it is more likely to be unaffected by policy treatment than is medium-skilled employment. For low-skilled female employment, I control for high-skilled female employment to capture general labor market conditions for women.[14] Youth (15–24 years old) and adult (25–64 years old) employment data,[15] as well as data on minimum and median wages, are from http://stats.oecd.org or are constructed from information provided there. I follow OECD (1998) and Neumark and Wascher (2004) and include adults as a control.

I also control for a set of labor market institutional variables—including information on unemployment benefit replacement rates, strictness of employment protection legislation, active labor market policy spending as a share of GDP, and union density—that are also provided by OECD and described in detail in Bassanini and Duval (2006).[16] Relative cohort size and the fertility rate are from the World Bank Health Nutrition and Population Statistics. The output gap is from the OECD.[17]

For most countries with minimum wage policies in place, the youth employment-to-population ratio shows a decreasing trend over the sample

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[13]Data for the years 1997 to 2005 are from OECD (2012), for 2006 to 2013 from OECD (2015b). For the countries of interest, both editions show basically identical values for the overlapping years of 2005–2008. In a few cases, for example, Spain, minor differences occur in levels but not in growth rates for these overlapping years. To adjust for this gap in levels, I chain the two series.

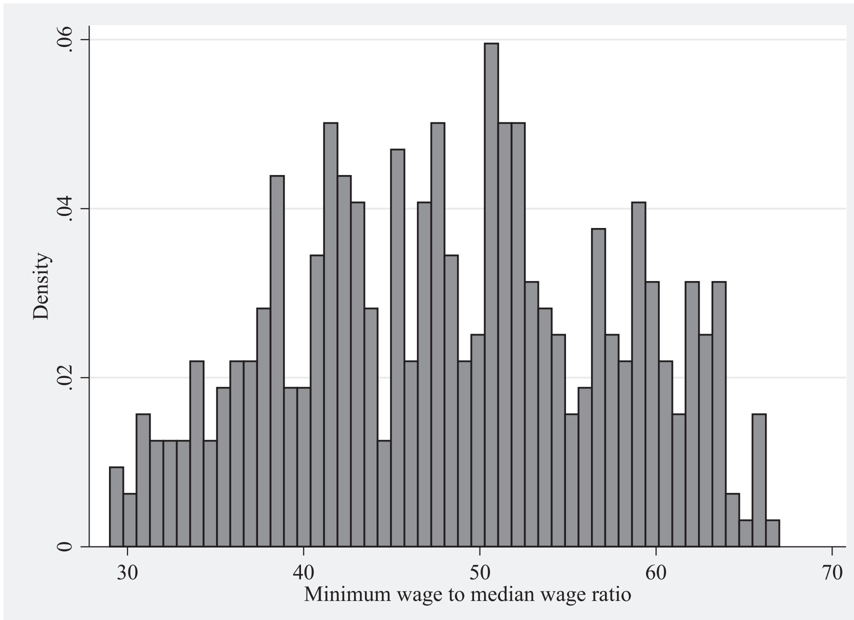
[14]This control is preferred for female employment, as male employment patterns follow very different trends. Results do not change substantially, however, if male high-skilled employment is chosen as the control.

[15]The underlying Labor Force Survey shows several statistical breaks in many countries. Whereas most of the previous literature seems to ignore potential implications, Neumark and Wascher (2004) included a step-dummy for two major breaks, in the Netherlands in 1987 and in Germany after reunification (according to e-mail correspondence with William Wascher). Unclear, however, is why such a dummy should be included for only a few among several significant breaks. I apply the following arbitrary procedure to test for the robustness of my results regarding such statistical breaks: I include a step-dummy for major breaks, which I define as being indicated by OECD in the legend as significant, and in which the youth employment rate changes by more than 5% over the previous year. I end up including dummies for Belgium (1984, 1993, 2000), Spain (1999, 2005), Greece (1998), Hungary (2003), Korea (2000), Luxembourg (1993, 2003), Netherlands (1987), Poland (1999, 2000, 2001), Portugal (1983, 1992, 1998), and Slovak Republic (1999). Including these dummies has little impact on the overall results.

[16]OECD calculated unemployment benefits as a share of the average production worker wage until 2005, and as a share of the average worker wage from 2001 onward, including more sectors in the latter measure. For the years 2001 to 2005, the levels of both measures are extremely similar in all cases but Italy, while growth rates are basically identical in all cases. I thus chain the two measures with growth rates to construct a series covering the whole sample period.

[17]The output gap for Greece and Luxembourg is from the AMECO database (the annual macroeconomic database of the European Commission's Directorate General for Economic and Financial Affairs).

Figure 1. Histogram of Minimum-to-Median Wage Ratios, 1983–2013



Source: OECD, author's presentation.
Notes: Observations are aggregated into 50 bins.

period, but there are also countries with an increase or no trend (see online Appendix A). The Great Recession, starting around 2008, led to a noticeable decrease in employment ratios in several countries. Youth employment ratios (41.1%) are on average much lower than those of adults (69.6%). This difference, of course, has to do with the fact that a significant share of young cohorts is enrolled in schools and universities and not participating in the labor market. The overall picture is similar for low-skilled employment. Many countries show a trend decline over time, but equally many show an increase or no trend change. The employment-to-population ratios for the skill groups differ considerably, with an average of 55.1% for low-skilled workers and 83.2% for high-skilled workers. The minimum-to-median wage ratio falls moderately in some countries during the sample period but increases or remains unchanged in several others. For the period 1983 to 2013, it ranges from 28.8 to 68.3% of the median wage, with an average value of 48.4% across countries (see Figure 1).

Regression Results

Low-Skilled and Female Low-Skilled Employment

Table 2 presents the regression results explaining low-skilled employment, estimating a static or dynamic equation using a total of six different approaches (see section above on Empirical Approach). Every result is presented in three different versions, without country-specific trends (panel A), with linear trends (panel B), and with linear and quadratic trends (panel C). I control for high-skilled employment, various institutional characteristics of the labor market, and country and time fixed effects in all specifications, with the exception of the LASSO specifications, which include controls picked by the data-driven selection process. Because of space considerations, I present only the minimum wage effects, and for the dynamic specifications also the lagged dependent variable, and the corresponding standard errors.

The canonical fixed-effects (FE) estimator yields economically very small and statistically insignificant negative or positive employment elasticities of minimum wage shocks, ranging between −0.04 and + 0.02, depending on whether, and which time trends are controlled for. The post-LASSO results are very similar to the FE results, with point estimates ranging from −0.02 to −0.01. The two-stage least squares instrumental variable (2SLS IV) estimates are also in this order of magnitude, although their standard errors are a little larger. Point estimates range from −0.05 to + 0.03. The results of the dynamic estimators are similarly small, ranging from −0.03 to + 0.03 in the case of the difference GMM, and −0.05 to + 0.02 for the system GMM estimator. The first difference estimates range from −0.01 to + 0.06. In all cases, the estimates are statistically insignificant.[18]

In sum, the estimated employment elasticities in Table 2 are always close to zero and insignificant. The average value of the estimated employment elasticity across all specifications is −0.01, with −0.05 as the most negative point estimate and + 0.06 as the most positive one. The most negative or imprecise estimate presented in Table 2 allows ruling out an employment elasticity more negative than −0.19 at the 95% confidence level. Thus, the bulk of evidence points to disemployment effects very close to zero, and even the most negative or imprecise estimate allows ruling out large negative employment elasticities. The magnitude of these results seems in line with recent evidence from the United States: Allegretto et al. (2017) presented results from several different estimation approaches, such as state-panel fixed effects and post-LASSO with time trends or division-time fixed effects, a border-discontinuity design, and a pooled synthetic controls approach. They report employment elasticities of minimum wage increases for teens and restaurant workers of −0.06 or smaller after accounting for spatial heterogeneity in some form.

I repeat the analysis of Table 2 for *female low-skilled workers* only, a group that shows an especially high incidence of earning minimum wages (OECD

[18]I present robustness checks of the results in Table B.2 in online Appendix B. First, in Table B.1, I illustrate that the results do not depend on the functional form by presenting estimates in levels. Evaluated at the mean, the elasticities are very similar to the logarithmic form. Second, if minimum wages raise unemployment of low-skilled or young workers and governments respond to it by spending more on active labor market policy (ALMP), controlling for ALMP might pick up some of the minimum wage effects. To address this possibility, I present results omitting ALMP from the list of controls in online Appendix Table B.2. This had virtually no effect on the minimum wage elasticities.

Table 2. Effects of Log Minimum Wage Ratio (MWR) on Log Low-Skilled Employment for Different Estimators and Specifications

Variable	*FE*	*Post-LASSO*	*2SLS IV*	*Difference GMM*	*System GMM*	*FD*
			Panel A: without trends			
	(A1)	(A2)	(A3)	(A4)	(A5)	(A6)
ln(MWR)	0.019	−0.019	0.026	0.027	0.020	−0.005
	(0.062)	(0.063)	(0.050)	(0.048)	(0.023)	(0.033)
ln(lagged dep. variable)				0.565***	0.984***	
				(0.083)	(0.022)	
Adjusted R^2	0.661	0.677	0.629			0.531
			Panel B: with linear trends			
	(B1)	(B2)	(B3)	(B4)	(B5)	(B6)
ln(MWR)	−0.039	−0.011	−0.052	−0.023	−0.012	0.004
	(0.036)	(0.036)	(0.057)	(0.038)	(0.036)	(0.035)
ln(lagged dep. variable)				0.317***	0.984***	
				(0.085)	(0.030)	
Adjusted R^2	0.825	0.842	0.802			0.583
			Panel C: with linear and quadratic trends			
	(C1)	(C2)	(C3)	(C4)	(C5)	(C6)
ln(MWR)	−0.001	−0.021	−0.021	−0.025	−0.054	0.064
	(0.054)	(0.050)	(0.063)	(0.051)	(0.066)	(0.064)
ln(lagged dep. variable)				0.244**	0.948***	
				(0.098)	(0.078)	
Adjusted R^2	0.870	0.876	0.852			0.647
Observations	269	269	260	241	260	250
No. of countries	19	19	19	19	19	19

Source: OECD, author's calculations.
Notes: All specifications except post-LASSO include high-skilled employment (EC), active labor market policy (ALMP), employment protection legislation (EPL), unemployment benefits (UB), and union density (UD) as controls, and purge country and time fixed effects. The post-LASSO specifications include controls picked by the selection process, that is, *EC, UD, UB,* $EPL*EC^2$, *EC*UB*ALMP*, and $EC*ALMP^2$ in specification A2; *EC, UB,* UD^2, $UB*EPL^2$, *EC*UD*EPL, EC*UB*EPL, UD*UB*EPL, ALMP*UB*EPL*, and *EC*ALMP*EPL* in specification B2; and *EC, UD, UB, EC*UB, EC*ALMP*UB,* $UB*EPL^2$, *EC*ALMP*EPL, EC*UB*EPL,* $EC*ALMP^2$, and *EC*UD*EPL* in specification C2. Cluster-robust standard errors in parentheses. FE, fixed effect; LASSO, least absolute shrinkage and selection operator; 2SLS IV, two-stage least squares instrumental variable; GMM, generalized method of moments; FD, first differences.
*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

1998; Rycx and Kampelmann 2013). The set of controls includes female high-skilled workers, various labor market institutions, and in line with the previous literature, also the fertility rate. Results are presented in Table 3. They largely resemble the results of Table 2 for both genders, with estimated employment elasticities generally close to zero and insignificant. However, the minimum wage elasticities are less consistently estimated compared to Table 2, with some of the results indicating statistically insignificant but noticeable negative or positive employment effects. On average, across all specifications, the employment elasticities lie at 0.00, which is basically similar to the average elasticity of all low-skilled workers. The most negative point estimate is −0.12, and + 0.11 is the most positive one. Overall, little evidence supports the premise that minimum wages cause significant job losses for female low-skilled workers, or that disemployment effects are more negative for female low-skilled workers than for all low-skilled workers. In the following robustness analysis, I thus focus on all low-skilled workers.

I continue by testing for *heterogeneity* and *nonlinearity* (Table 4) for the sample of all low-skilled workers. The first panel in Table 4 (specifications G1–G3) addresses the possible dependence of minimum wage effects on the state of the economy; specifically, that minimum wages might be more harmful when output is below its potential. Thus, I add a dummy variable equaling 1 if the output gap is negative and an interaction term between this dummy and the minimum wage to the FE specifications of Table 2. The results suggest that little difference in minimum wage effects occurs between booms and recessions, with no strongly or statistically significant negative estimate in any of the specifications. In the second panel (specifications H1–H3), I test for institutional complementarities between minimum wages and other labor market institutions. The FE specifications of Table 2 are therefore augmented with de-meaned interaction terms between the minimum wage and each of the labor market institutional controls. Including these interaction terms has little effect on the minimum wage coefficient, with all three specifications returning low positive and insignificant elasticities. One of the interaction terms is found to have a consistent and mostly statistically significant effect. It suggests that a high minimum wage can ease the negative effects of unemployment benefits on low-skilled employment.[19] Thus, minimum wage increases seem to incentivize unemployed workers to look for jobs. These results, however, do not indicate any negative employment effects of minimum wages.

Finally, I test if high minimum wages lead to more negative employment outcomes than do low ones (specifications I1–I3). I include an interaction term between the minimum wage and a dummy equal to 1 if the minimum wage is below 40% of the median wage, as well as an interaction term between the minimum wage and a dummy equal to 1 if the minimum wage is above 55% of the median wage.[20] Thus, the specification allows for less-negative effects of low minimum wages below 40% of the median, and more-negative effects of high minimum wages above 55% of the median. Results show that low minimum wages indeed lead to less negative (or more positive) outcomes than do those above the threshold; however, the difference in the estimated employment elasticities (+ 0.01) is economically

[19]The coefficients and standard errors (in parentheses) of unemployment benefits in the specification without trends, with linear trends, and with quadratic trends are −0.042 (0.024), −0.037 (0.047), and −0.066 (0.026), respectively.

[20]Forty-seven observations are below and 43 are above the threshold of 40% and 55% of the median wage, respectively.

Table 3. Effects of Log Minimum Wage Ratio (MWR) on Log Female Low-Skilled Employment for Different Estimators and Specifications

Variable	*FE*	*Post-LASSO*	*2SLS IV*	*Difference GMM*	*System GMM*	*FD*
				Panel D: without trends		
	(D1)	(D2)	(D3)	(D4)	(D5)	(D6)
ln(MWR)	−0.122	−0.022	−0.055	−0.027	0.018	0.002
	(0.102)	(0.069)	(0.069)	(0.061)	(0.017)	(0.051)
ln(lagged dep. variable)				0.646***	0.979***	
				(0.076)	(0.017)	
Adjusted R^2	0.528	0.627	0.512			0.329
				Panel E: with linear trends		
	(E1)	(E2)	(E3)	(E4)	(E5)	(E6)
ln(MWR)	0.017	0.023	−0.014	−0.003	−0.036	0.020
	(0.043)	(0.057)	(0.077)	(0.066)	(0.026)	(0.059)
ln(lagged dep. variable)				0.290***	0.933***	
				(0.084)	(0.046)	
Adjusted R^2	0.813	0.821	0.789			0.395
				Panel F: with linear and quadratic trends		
	(F1)	(F2)	(F3)	(F4)	(F5)	(F6)
ln(MWR)	0.080	0.079	0.107	−0.032	−0.093	0.015
	(0.076)	(0.065)	(0.092)	(0.087)	(0.058)	(0.066)
ln(lagged dep. variable)				0.147	0.845***	
				(0.122)	(0.088)	
Adjusted R^2	0.850	0.852	0.833			0.448
Observations	269	269	260	241	260	250
No. of countries	19	19	19	19	19	19

Source: OECD, World Bank, author's calculations.
Notes: All specifications except post-LASSO include female high-skilled employment (EC), active labor market policy (ALMP), employment protection legislation (EPL), unemployment benefits (UB), union density (UD), and fertility rate (fertility) as controls, and purge country and time fixed effects. The post-LASSO specifications include controls picked by the selection process, that is, *EC, EC^2, $fertility^2$, EPL*EC^2, ALMP*EC^2, fertility*EPL^2, EC*UB*ALMP, fertility*ALMP, UB*EC^2, EC*EPL*ALMP,* and *fertility*EPL*EC* for specification D2; *EC, UB, $fertility^2$, UD*UB*fertility, UB*EPL^2, fertility*EPL, fertility*EPL*EC, EC*$ALMP^2$*, and *UD*$fertility^2$* for specification E2; and *EC, UB, $fertility^2$, EC*UB*ALMP, EC*$ALMP^2$, UB*EPL^2, ALMP*EC^2, EC*EPL*ALMP,* and *UD*$fertility^2$* for specification F2. Cluster-robust standard errors in parentheses. FE, fixed effect; LASSO, least absolute shrinkage and selection operator; 2SLS IV, two-stage least squares instrumental variable; GMM, generalized method of moments; FD, first differences.
*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

marginal. High minimum wages result in virtually similar elasticities as those below this threshold. Thus, somewhat surprisingly but in line with OECD (1998), I do not find support for strong heterogeneous effects of low or high minimum wages in the data.

I continue the analysis by investigating the presence of pre-existing trends and long-run effects (see Equation 2). I re-estimate the FE specifications of

Table 4. Heterogeneous and Nonlinear Effects of Log Minimum Wage Ratio (MWR) on Log Low-Skilled Employment

	MWR interacted with recession dummy			*MWR interacted with institutions*			*MWR threshold at 40% and 55%*		
Variable	*(G1)*	*(G2)*	*(G3)*	*(H1)*	*(H2)*	*(H3)*	*(I1)*	*(I2)*	*(I3)*
ln(MWR)	0.016	−0.052	−0.017	0.048	0.054	0.065	0.030	−0.006	0.041
	(0.068)	(0.042)	(0.061)	(0.044)	(0.046)	(0.046)	(0.083)	(0.040)	(0.056)
ln(MWR)*Recession dummy	−0.019	0.022	0.029						
	(0.030)	(0.025)	(0.024)						
ln(MWR)*ln(Union density)				0.042	−0.003	−0.120			
				(0.099)	(0.113)	(0.146)			
ln(MWR)*ln(Empl. protection legislation)				−0.054	0.071	−0.048			
				(0.060)	(0.058)	(0.048)			
ln(MWR)*ln(Unempl. benefits)				0.067	0.254**	0.315***			
				(0.069)	(0.089)	(0.079)			
ln(MWR)*ln(Active labor market policy)				0.107	−0.101	−0.052			
				(0.086)	(0.074)	(0.075)			
ln(MWR)*MWR<40% dummy							0.006	0.006	0.008***
							(0.004)	(0.003)	(0.003)
ln(MWR)*MWR>55% dummy							0.001	−0.001	−0.001
							(0.003)	(0.002)	(0.002)
Linear trends	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Quadratic trends	No	No	Yes	No	No	Yes	No	No	Yes
Adjusted R^2	0.669	0.828	0.871	0.676	0.835	0.880	0.663	0.827	0.874
Observations	269	269	269	269	269	269	269	269	269
No. of countries	19	19	19	19	19	19	19	19	19

Source: OECD, author's calculations.

Notes: Two-way fixed-effects specifications including high-skilled employment, active labor market policy, employment protection legislation, unemployment benefits, and union density as controls. In Panel H, interaction terms are constructed by subtracting the variables' respective sample means. Cluster-robust standard errors in parentheses.

*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

Table 2, including two years of leads, the contemporaneous value, and three years of lags of the minimum wage variable. The findings are presented numerically in Table 5 and graphically in Figure 2, where the cumulative response of minimum wage increases is plotted over time, set to zero the year before treatment occurs. Results of the distributed lag model suggest that, first, some negative pre-existing trends are present in all specifications, and especially strong and highly significant ones in the specification with linear trends. Second, the size of the long-run effect is maximally less than a third of the size of the pre-existing trends across specifications, which potentially means that low-skilled employment was falling during the years prior to minimum wage increases, and it continued to fall in the following years, independent of minimum wage policies. Third, the contemporaneous effect is noticeably, though insignificantly, positive in all specifications, contrary to what one expects to see if minimum wage hikes cause unemployment. Either way, there is no evidence for statistically or economically significant negative long-run effects in the specifications, with long-run elasticities ranging from −0.06 to + 0.01. These long-run estimates are very similar to the contemporaneous effects of the FE specifications in Table 2.

Overall, no evidence supports significant job losses in response to minimum wage hikes in these samples of low-skilled and female low-skilled workers. Most estimated employment elasticities are very close to zero. There is also no evidence that minimum wages in combination with other labor market institutions or in recessions harm employment. Some statistically significant evidence suggests that low minimum wages yield less adverse effects than do higher minimum wages; however, the difference in elasticities is economically tiny, as is the employment effect of minimum wages above this threshold. Finally, little evidence validates statistically or economically significant long-run effects.

Youth Employment

As in the previous subsection for low-skilled workers, I assess the impact of minimum wages on job prospects for young workers by estimating static and dynamic specifications, applying six different estimators on three different versions of the specifications, without controlling for time trends, with linear trends, and with up to quadratic trends (see the section above titled Empirical Approach for details). These results are presented in Table 6.

For the fixed-effects specifications, the results vary from small insignificantly negative (–0.01) to sizably positive (+ 0.16). The post-LASSO estimates are very closely clustered around zero, ranging from −0.04 to + 0.04. The 2SLS IV results range from −0.04 to + 0.22. The dynamic estimators return elasticities between −0.04 and + 0.11. Finally, the FD estimates are all positive, between + 0.01 and + 0.10.[21]

[21]Robustness tests of these results are shown in online Appendix B. First, the elasticities remain very similar if the specifications are estimated in levels and evaluated at the mean (Table B.3). Second, the results are robust to omitting ALMP from the list of controls (Table B.4).

Table 5. Pre-existing Trends and Long-Run Effects of Log Minimum Wage Ratio (MWR) on Log Low-Skilled Employment

Variable	*(J1)*	*(J2)*	*(J3)*
$\ln(\text{MWR}_{t+2})$	0.036	−0.112*	−0.126*
	(0.059)	(0.061)	(0.070)
$\ln(\text{MWR}_{t+1})$	−0.107	−0.169*	−0.057
	(0.075)	(0.096)	(0.090)
$\ln(\text{MWR}_{t})$	0.101	0.103	0.056
	(0.095)	(0.067)	(0.068)
$\ln(\text{MWR}_{t-1})$	−0.039	−0.056	−0.029
	(0.048)	(0.043)	(0.048)
$\ln(\text{MWR}_{t-2})$	−0.035	−0.019	−0.058
	(0.054)	(0.062)	(0.056)
$\ln(\text{MWR}_{t-3})$	−0.014	−0.085	−0.028
	(0.050)	(0.068)	(0.061)
Sum of leads	−0.070	−0.281***	−0.183
from $t+2$ to $t+1$	(0.079)	(0.083)	(0.137)
Sum of lags	0.013	−0.056	−0.060
from t to t–3	(0.078)	(0.080)	(0.133)
Linear trends	No	Yes	Yes
Quadratic trends	No	No	Yes
Adjusted R^2	0.658	0.799	0.859
Observations	231	231	231
No. of countries	19	19	19

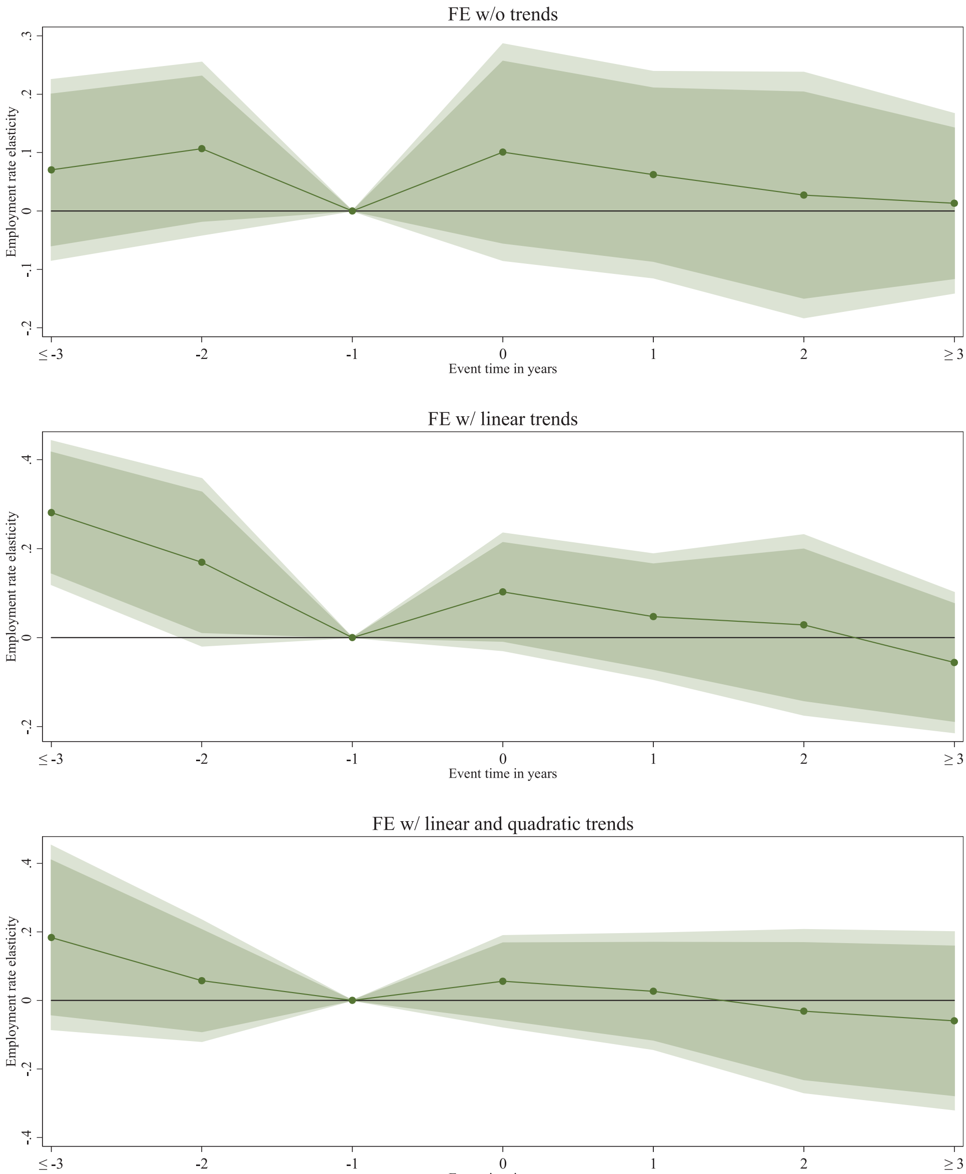
Source: OECD, author's calculations.
Notes: Two-way fixed-effects specifications including high-skilled employment, active labor market policy (ALMP), employment protection legislation (EPL), unemployment benefits (UB), and union density (UD) as controls. Cluster-robust standard errors in parentheses.
*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

In sum, the employment elasticities in Table 6 point to small negative or positive elasticities ranging from −0.06 to + 0.22, with an average value of + 0.04. The most negative or imprecise estimate in Table 6 allows ruling out an employment elasticity more negative than −0.35 at the 95% confidence level. Thus, no evidence for strong negative effects is found, and most of the specifications allow ruling out large negative employment effects, as they were found in some of the previous cross-country studies. Specifically, the estimated youth employment elasticities in Table 6 are moderately less negative than in OECD (1998), significantly less negative than in Neumark and Wascher (2004), and less positive than in Bassanini and Duval (2006).

In online Appendix Table B.5, I show that the results of Table 6 are robust for a larger sample, spanning from 1970 to 2013 and including 24 OECD countries. This sample covers the period investigated by Neumark and Wascher (2004) as well as in this section, but has fewer controls available to capture differences in labor market institutions.[22] It also includes

[22]Further, the information necessary to construct the instruments for the 2SLS IV estimation is not available.

Figure 2. Cumulative Response of the Log Low-Skilled Employment Rate over Time to a Log Point Increase in the Minimum Wage Ratio



Notes: Graphical representation of the results in Table 5. The time path is calculated by summing up their joint effects and setting the year before treatment to 0. The dark shaded area represents 90% and the light shaded area 95% cluster-robust confidence intervals. FE, fixed effects.

countries with bargained minimum wages as a share of the mean wage from Neumark and Wascher's (2004) data set (based on Dolado et al. 1996). The average employment elasticity across specifications for this sample is also close to zero (–0.06), with no estimate being statistically significant.

Table 6. Effects of Log Minimum Wage Ratio (MWR) on Log Youth Employment for Different Estimators and Specifications

Variable	*FE*	*Post-LASSO*	*2SLS IV*	*Difference GMM*	*System GMM*	*FD*
				Panel K: without trends		
	(K1)	(K2)	(K3)	(K4)	(K5)	(K6)
ln(MWR)	0.161	0.035	0.221	0.109	−0.036	0.013
	(0.169)	(0.099)	(0.165)	(0.090)	(0.045)	(0.038)
ln(lagged dep. variable)				0.708***	0.958***	
				(0.049)	(0.027)	
Adjusted R^2	0.697	0.822	0.682			0.618
				Panel L: with linear trends		
	(L1)	(L2)	(L3)	(L4)	(L5)	(L6)
ln(MWR)	−0.011	−0.039	−0.038	0.098	−0.023	0.057
	(0.089)	(0.102)	(0.159)	(0.090)	(0.044)	(0.044)
ln(lagged dep. variable)				0.475***	0.884***	
				(0.073)	(0.050)	
Adjusted R^2	0.899	0.905	0.893			0.661
				Panel M: with linear and quadratic trends		
	(M1)	(M2)	(M3)	(M4)	(M5)	(M6)
ln(MWR)	0.046	0.031	0.094	−0.025	−0.010	0.096
	(0.075)	(0.067)	(0.107)	(0.095)	(0.059)	(0.064)
ln(lagged dep. variable)				0.329***	0.702***	
				(0.068)	(0.058)	
Adjusted R^2	0.943	0.947	0.939			0.701
Observations	420	420	407	401	420	401
No. of countries	19	19	19	19	19	19

Source: OECD, World Bank, author's calculations.
Notes: All specifications except post-LASSO include adult employment (EC), active labor market policy (ALMP), employment protection legislation (EPL), unemployment benefits (UB), union density (UD), and cohort size (cohort) as controls, and purge country and time fixed effects. The post-LASSO specifications include controls picked by the selection process, that is, *EC*, $cohort^2$, $EC*UD^2$, $UD*EPL^2$, $EC*ALMP^2$, *cohort*EC*, *cohort*EC*ALMP*, *UD*ALMP*, *UB*EC*EPL*, *EC*EPL*ALMP*, *EC*UB*, EPL^3, *EPL*cohort*, $EC*UB^2$, and $UB*cohort^2$ for specification K2; *EC*, *UD*, $UD*EPL^2$, *EPL*UB*, *EPL*ALMP*cohort*, *UB*EC*EPL*, *EC*UB*ALMP*, *EC*cohort*, $EC*cohort^2$, $UD*EC^2$, $EPL*UD^2$, $cohort*UD^2$, $cohort^2$, and $cohort^3$ for specification L2; *EC*, *UB*, *UD*EPL*, $UB*EPL^2$, *EPL*ALMP*cohort*, $UD*EC^2$, *EC*UB*, $EC*UB^2$, *EC*UB*cohort*, $cohort*UD^2$, $cohort^2$ for specification M2. Cluster-robust standard errors in parentheses. FE, fixed effect; LASSO, least absolute shrinkage and selection operator; 2SLS IV, two-stage least squares instrumental variable; GMM, generalized method of moments; FD, first differences.
*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

Table 7 shows the test results assessing heterogeneity and nonlinearity. In the first panel (specifications N1–N3), I investigate if minimum wage effects are stronger in recessions. This inquiry is supported by the results, but only in the specification without trends is the difference significant. Furthermore, the total effect of minimum wages is positive or very close to zero in all specifications. Taken at face value, the results indicate that minimum wages have an insignificant positive effect on youth employment in boom times, and basically no effect in recessions.

Table 7. Heterogeneous and Nonlinear Effects of Log Minimum Wage Ratio (MWR) on Log Youth Employment

	MWR interacted with recession dummy			*MWR interacted with institutions*			*MWR threshold at 40% and 55%*		
Variable	*(N1)*	*(N2)*	*(N3)*	*(O1)*	*(O2)*	*(O3)*	*(P1)*	*(P2)*	*(P3)*
ln(MWR)	0.205	0.004	0.057	0.025	0.017	0.047	0.201	−0.064	0.087
	(0.168)	(0.084)	(0.072)	(0.132)	(0.086)	(0.093)	(0.220)	(0.103)	(0.092)
ln(MWR)*Recession dummy	−0.095*	−0.033	−0.037						
	(0.045)	(0.028)	(0.027)						
ln(MWR)*ln(Union density)				0.443**	0.093	−0.204			
				(0.203)	(0.250)	(0.235)			
ln(MWR)*ln(Empl. protection legislation)				0.058	−0.071	−0.060			
				(0.145)	(0.100)	(0.076)			
ln(MWR)*ln(Unempl. benefits)				−0.151	0.239**	0.170			
				(0.202)	(0.107)	(0.178)			
ln(MWR)*ln(Active labor market policy)				−0.142	−0.033	0.029			
				(0.126)	(0.094)	(0.072)			
ln(MWR)*MWR<40% dummy							0.014	0.006	0.003
							(0.018)	(0.008)	(0.004)
ln(MWR)*MWR>55% dummy							0.001	0.009	−0.003
							(0.013)	(0.005)	(0.004)
Linear trends	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Quadratic trends	No	No	Yes	No	No	Yes	No	No	Yes
Adjusted R^2	0.700	0.899	0.943	0.739	0.900	0.943	0.700	0.901	0.943
Observations	420	420	420	420	420	420	420	420	420
No. of countries	19	19	19	19	19	19	19	19	19

Source: OECD, World Bank, author's calculations.

Notes: Two-way fixed effects specifications including adult employment, active labor market policy (ALMP), employment protection legislation (EPL), unemployment benefits (UB), union density (UD), and cohort size as controls. In Panel O, interaction terms are constructed by subtracting the variables' respective sample means. Cluster-robust standard errors in parentheses.

*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

In the second panel (specifications O1–O3), institutional complementarities of minimum wages with other labor market institutions are addressed. Little consistent evidence across specifications occurs for such complementarities. In the specification without trends, I find that high minimum wages might alter the negative (though insignificant) effects of union density on youth employment.[23] This result, however, is not robust against the inclusion of trends. In the specification with linear trends, the findings suggest that high minimum wages in combination with unemployment benefits increase youth employment.[24] Such a positive complementarity between minimum wages and unemployment benefits is also present for low-skilled workers (see the earlier section titled Low-Skilled and Female Low-Skilled Employment), as well as for female workers (see Addison and Ozturk 2012). This result is significant only in the specification including linear trends, however, and should thus be interpreted cautiously. The crucial point is that no evidence in this sample supports that significant adverse effects of minimum wages occur in combination with other labor market institutions; if anything, these complementarities go in the opposite direction. Neumark and Wascher found that strict employment protection legislation reduces the adverse effects of rising minimum wages, and they argued that "stricter employment protection regulations [. . .] offset the negative employment consequences of a wage floor, perhaps because it is more costly to dismiss workers in countries with such regulations" (2004: 241). No such effects are present in my sample.

Last, the effects of low and high minimum wages are assessed (specifications P1–P3).[25] I find little evidence that low or high minimum wages result in less or more adverse youth job market perspectives, in line with OECD (1998).

In the next specifications, I test for pre-existing trends and long-run effects with distributed lags (see Equation 2), including two leads, the contemporaneous value, and three lags of the minimum wage ratio in the specification. Results are shown in Table 8 and Figure 3, on which the cumulative response of minimum wage increases is plotted. Some insignificant but noticeable positive effects of the sum of the leads are seen in specification Q1, suggesting that it might be polluted by pre-existing trends. The long-run employment elasticities range from -0.08 to $+0.05$. Although these long-run results are somewhat more negative than the contemporaneous effects in Table 6, they are neither statistically significant nor economically large.

Overall, no statistically significant disemployment effects of minimum wages are present in this sample of employment-to-population ratios of young workers. The minimum wage elasticities are very small in most cases.

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[23]The coefficient (standard error) of union density is -0.077 (0.117).

[24]Unemployment benefits themselves are found to be statistically significantly positive in this specification. The coefficient (standard error) of unemployment benefits is $+0.133$ (0.047).

[25]Eighty-seven observations are below the 40% threshold, 114 observations are above the 55% threshold.

Table 8. Pre-existing Trends and Long-Run Effects of Log Minimum Wage Ratio (MWR) on Log Youth Employment

Variable	*(Q1)*	*(Q2)*	*(Q3)*
$\ln(\text{MWR}_{t+2})$	0.034	0.045	0.102
	(0.227)	(0.134)	(0.126)
$\ln(\text{MWR}_{t+1})$	0.108	0.006	−0.171
	(0.156)	(0.109)	(0.099)
$\ln(\text{MWR}_{t})$	0.196	0.009	−0.078
	(0.222)	(0.147)	(0.141)
$\ln(\text{MWR}_{t-1})$	−0.003	0.020	0.105
	(0.151)	(0.093)	(0.073)
$\ln(\text{MWR}_{t-2})$	0.052	−0.041	−0.017
	(0.078)	(0.066)	(0.056)
$\ln(\text{MWR}_{t-3})$	−0.298***	−0.065	0.041
	(0.102)	(0.096)	(0.063)
Sum of leads	0.142	0.051	−0.07
from $t+2$ to $t+1$	(0.186)	(0.164)	(0.137)
Sum of lags	−0.054	−0.077	0.051
from t to t–3	(0.119)	(0.125)	(0.174)
Linear trends	No	Yes	Yes
Quadratic trends	No	No	Yes
Adjusted R^2	0.630	0.866	0.927
Observations	378	378	378
No. of countries	19	19	19

Sources: OECD, World Bank, author's calculations.
Notes: Two-way fixed-effects specifications including adult employment, active labor market policy (ALMP), employment protection legislation (EPL), unemployment benefits (UB), union density (UD), and cohort size as controls. Cluster-robust standard errors in parentheses.
*, **, and *** means significant at the 10%, 5%, and 1% level, respectively.

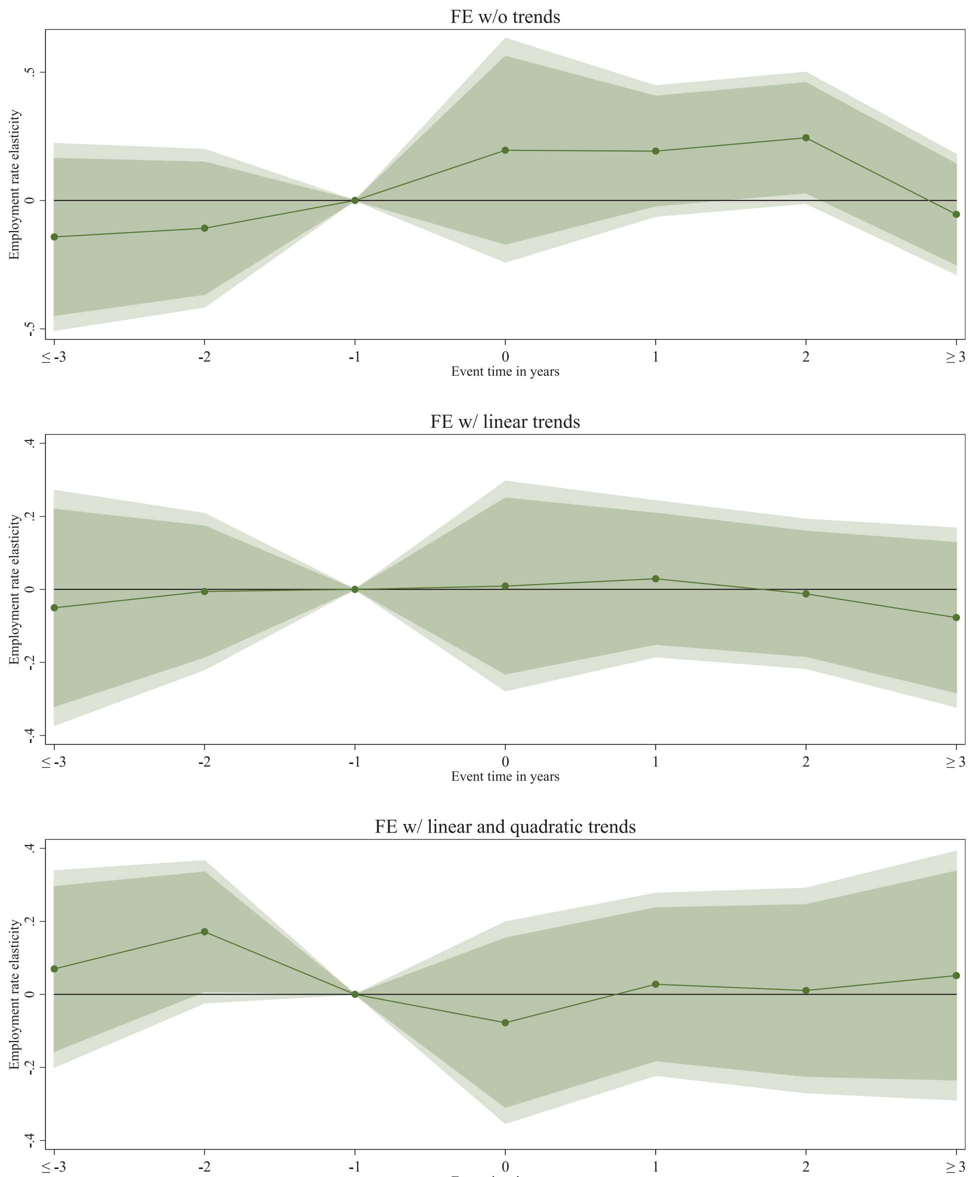
Further, no significant evidence occurs for adverse effects of minimum wages in recessions, in combination with other labor market institutions, or of high minimum wages. Finally, I do not find evidence for strong adverse effects of minimum wages in the long run.

My results on youth employment differ from some of the previous findings, most notably Neumark and Wascher (2004), who reported significant negative minimum wage effects (see Table 1). In online Appendix C, I investigate the reasons for these differences. When reassessing Neumark and Wascher's (2004) evidence with their original data as well as with newly updated data, I find that their results are not statistically significant once standard errors are clustered. Furthermore, their point estimates become very close to zero once variables are transformed into logarithms or when adult employment instead of unemployment is controlled for. Overall, I find no evidence in their sample for significant disemployment effects of minimum wages.

Conclusion

In this study, I investigate the impact of minimum wages on low-skilled employment in a sample of 19 OECD countries, and I reassess its impact on

Figure 3. Cumulative Response of the Log Youth Employment Rate over Time to a Log Point Increase in the Minimum Wage Ratio



Notes: Graphical representation of the results in Table 8. The time path is calculated by summing up their joint effects and setting the year before treatment to 0. The dark shaded area represents 90%, and the light shaded area 95% cluster-robust confidence intervals. FE, fixed effects.

youth labor market outcomes. I present the results of several different static and dynamic estimation approaches, such as two-way fixed effects, the post-double selection least absolute shrinkage and selection operator, two-stage least squares instrumental variables, difference and system general methods of moments, and first differences. Results are shown in three different versions: without time trends, with linear ones, and with linear and quadratic trends, amounting in total to 18 different specifications.

The findings suggest little evidence for substantial disemployment effects of minimum wages in the samples of low-skilled and youth labor market outcomes. The average estimated employment elasticity across specifications is −0.01 for low-skilled workers, 0.00 for female low-skilled workers, and + 0.04 for young workers. The most negative point estimate for low-skilled workers amounts to −0.05, −0.12 for female low-skilled workers, and −0.04 for young workers. Most of these estimates rule out large negative employment elasticities of minimum wage hikes at conventional levels of significance. For youth employment, the absence of strong minimum wage effects is also confirmed for a larger sample of 24 countries. Thus, the evidence points to employment elasticities in response to minimum wage hikes close to zero and rules out strong disemployment effects with certainty; however, moderate negative effects cannot be ruled out.

I apply several robustness checks and extensions to investigate issues of heterogeneity and nonlinearity. First, I test if minimum wages are more harmful when output is below its potential. Although this seems to be the case, the difference is economically marginal. Second, I investigate institutional complementarities between minimum wages and other labor market institutions. No evidence confirms that minimum wages in combination with other institutions are harmful to employment of low-skilled or young workers. Third, I investigate nonlinear minimum wage effects by allowing for a kink in the minimum wage function at 40% and 55% of the median wage. There is some significant evidence that employment elasticities of low minimum wages show less adverse effects on low-skilled and young workers, but the size of this difference is small, as are the effects of high minimum wages. Fourth, I assess long-run effects. The long-run effects are somewhat more negative than the contemporaneous effects, yet nothing suggests significant or sizeable long-run effects.

Finally, I investigate why my results on youth employment differ from the findings of Neumark and Wascher (2004) by reassessing their results with the original data, as well as with newly updated data. I show that Neumark and Wascher overstated the precision of their estimates by not correcting standard errors for bias caused by serial correlation in the residuals. Once this shortcoming is addressed, no statistically significant effects are present in their sample. Further, I show that the estimated minimum wage effects become close to zero once a logarithmic specification is estimated, or when employment instead of unemployment of adults is controlled for.

I conclude that while some recent cross-country evidence suggests that minimum wages narrow wage and income inequality (Koeniger, Leonardi, and Nunziata 2007; Jaumotte and Buitron Osorio 2015), for the existing range of minimum wages they are unlikely an economically relevant determinant of job losses for low-skilled, female low-skilled, or young workers in OECD countries in recent decades.

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