
MINIMUM WAGE LAWS AND TOP INCOME SHARES: EVIDENCE FROM A LARGE PANEL OF STATES

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ABSTRACT

Over the three decades prior to the onset of the Great Recession in 2007, the income share of the top 1% in the United States rose by over 150% (from 9.3% to 23.5%). During the same period, the real federal minimum wage fell by about 35% (from \$8.92 to \$5.76 in 2011 dollars). This paper uses a comprehensive panel of U.S. states to explore the effect of changes in the real minimum wage on top income shares. Our findings indicate that the relationship between is negative in nature, but not robust to small changes in the econometric specification or in the measurement of inequality.

JEL Classifications: D3, J3

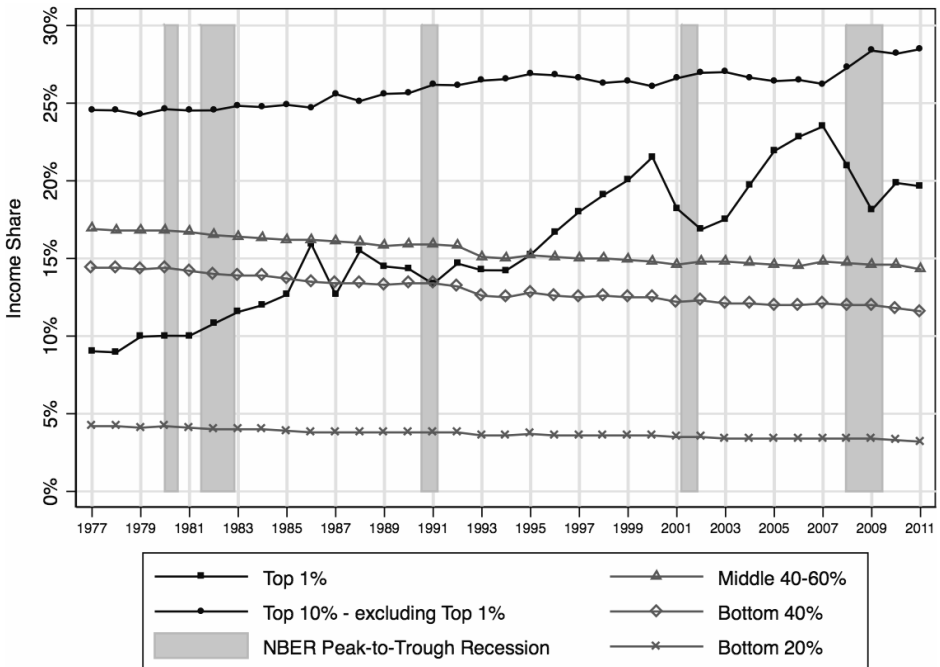
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INTRODUCTION

Few issues in economics have received as much attention as the recent rise in income inequality, and in particular, the rise in top income shares. The seminal work by Piketty and Saez (2003) shows that the rise in income inequality experienced in the U.S. since the late-1970s has been primarily driven by income changes in the upper-end of the distribution (see also Burkhauser, et. al. 2012, Piketty and Saez 2014). Figure 1 illustrates this point by presenting recent trends in income shares at various points of the income distribution.¹ When compared to the top 1%, the changes in shares from other parts of the distribution appear minor. Even the increase in the income share of the top 10% is relatively pedestrian once the top 1% is removed (rising from 24.5% in 1977 to 28.5% in 2011). The top 1%'s income share, by contrast, rose from 9.3% in 1977 to 23.5% on the eve of the Great Recession in 2007.

¹ Data on the top 1% and top 10% income shares are from Piketty and Saez (2003), and updated from the webpage of Emmanuel Saez: <http://eml.berkeley.edu/~saez/>. Data on the income shares of the bottom 60%, 40% and 20% are from the Census Bureau's Historical Income Tables: <http://www.census.gov/hhes/www/income/data/historical/household/>.

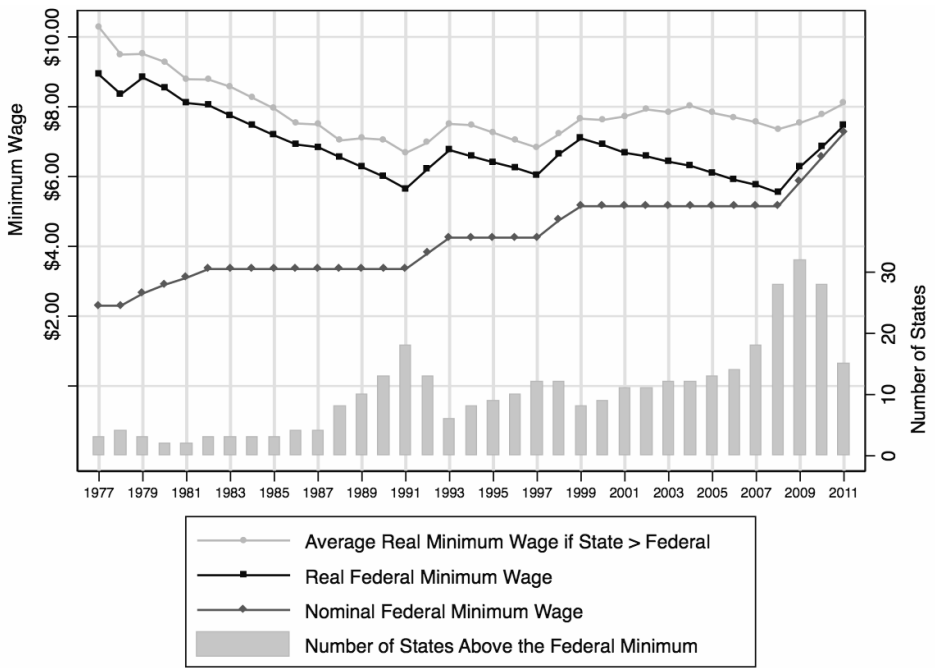
Figure 1
U.S. Income Shares



Policy reactions to this rise in top income shares have ranged from taxing capital (see Piketty 2014), to calls for greater investment in human capital (see Goldin and Katz 2008), to measures aimed at shifting the balance of power in the workplace, such as strengthening worker unions or increasing the minimum wage (see Stiglitz 2012). In this paper, we look at the later of these policy options: using minimum wage laws to redress the increase in top income shares.

Federal minimum wage laws in the U.S. are an interesting case because while the nominal minimum wage has increased, the real federal minimum wage fell about 35% over the three-decade period prior to the onset of the Great Recession in 2007 (see Figure 2). Notably, this fall in the real minimum wage correlates inversely with the rise in the top 1% income share shown in Figure 1. Many states, however, compensated for the real decline in the federal minimum wage by raising their own state minimum wage above the federal mandate (also shown in Figure 2). In this paper, we use this variability unique to the state-level to help identify the income inequality effects from real minimum wage changes. We argue that given the frequency of changes at the state-level, states are, consequently, the proper unit-of-analysis for investigating these effects.

Figure2
Federal and State Minimum Wage Levels



For changes in the real minimum wage to have a meaningful role in accounting for the rise in top income shares, these changes must go beyond a purely mechanical effect on wages. Instead, they must have spillovers onto wages beyond those at (or near) the minimum. The previous literature is sparse in this regard, with only about a half dozen articles broadly investigating the impact of minimum wage laws on income inequality. Most of this prior work has focused on spillover effects in the lower-half of the income distribution (see DiNardo, Fortin, and Lemieux 1996, Lee 1999, Card and DiNardo 2002, Dickens and Manning 2004, and Autor, Katz and Kearney 2008). A notable exception, however, is Autor, Manning, and Smith (2016); they find some evidence of spillovers into the upper-half of the income distribution (e.g., the 90/10 ratio). They do not consider spillovers into the upper-tail, however. As Figure 1 illustrates, this is a significant omission given that increases in U.S. income inequality over recent years have been driven by income increases in the upper-tail (i.e., the top 1%).

To our knowledge, this is the first paper to evaluate the impact of minimum wage changes on the rise in top income shares at the state-level. Prior research has focused on other parts of the income distribution, and as a result, used state-level

distributional data available from the March Current Population Survey (CPS). March CPS data does not, however, account for much of the rise in the top 1% since it omits relevant information on capital gains, stock options, and bonuses (see Burkhauser, et. al. 2012). Thus, we depart from the use of March CPS data, and instead use the IRS-based top income shares constructed by Frank (2009).

We base our econometric approach on the voluminous literature testing the role of minimum wage laws on employment. Much of this research is based on the seminal work of Card and Krueger (1994), and relies on the difference-in-differences estimation approach (see Neumark and Wascher 2007). Although this literature is focused on questions of employment and employment growth, its methodology is a good candidate because it is well vetted in the state-level panel context, and is able to exploit the unique state-level variations present in minimum wage laws (as shown in Figure 2). Moreover, the contexts of employment and income distribution are necessarily related, as distributional changes are a consequence of changes in earnings and employment.

While the initial results are promising, we ultimately conclude that the relationship between the minimum wage and top income shares is not particularly robust. Though consistently negative in sign, we find that small changes in the econometric specification, as well as the measurement of income inequality, lead to statistically insignificant associations. Hence, we conclude that changes in the minimum wage do not appear to have meaningful spillovers into the upper-tail of the income distribution, and fail to find support for the claim that increases in the minimum wage might lead to a lowering of top income shares.

In the following section, we introduce our state-level panel. Section III then presents the difference-in-differences estimation and results from our empirical investigation. Section IV offers a brief set of conclusions.

DATA

Our sample consists of annual state-level data for the period 1977-2011. We include the District of Columbia in the sample, meaning we have 51 cross-sections and 35 years of data (for a total of 1785 observations). Descriptive statistics for all the variables used in the analysis are shown in Table 1. The top income share measures are taken from the comprehensive panel of income inequality measures constructed by Frank (2009).² The state-level measures of Frank are based on income data reported by the IRS using individual tax returns, and follow the construction

² Available online at: http://www.shsu.edu/eco_mwf/inequality.html.

methodology of the national panel of top income shares constructed by Piketty and Saez (2003). Though the focus of this paper is on top income shares, later in the analysis we extend our estimations to include the Gini coefficient, a broad measure of income inequality.

Table 1
Summary Statistics

	Mean	Std. Dev.	Median
State minimum wage (\$)	4.40	1.360	4.25
State minimum wage (\$real)	7.09	0.916	6.89
<i>Income Inequality Measures:</i>			
Top 1% Income Share	0.14	0.0444	0.13
Top 1% Growth Rate	0.025	0.0887	0.029
Top 10% Income Share	0.38	0.0529	0.38
Gini Coefficient	0.56	0.0507	0.56
<i>Additional Controls:</i>			
GSP/capita (\$real)	41,591.6	16,309.7	38,447.1
Unemployment Rate	6.03	2.101	5.70
Population (thousands)	5,160.6	5,725.6	3,513.4
Pop Share Aged 15-59	61.7	1.957	61.7
Observations	1785		

Notes: The panel includes each state plus Washington D.C. ($n = 51$) annually over the period 1977-2011 ($t = 35$).

For the state minimum wage data, we use information provided by the U.S. Department of Labor.³ We also compared our data to that of Meer and West (2013), and verified any differences with state-level sources. To assure proper casual ordering between the variables, we use the previous year's value of the minimum wage (as of March 12th) in our panel estimations. For states that use a multi-track menu of minimum wages, we use the maximum wage in this menu, as is common in the literature. State-level price indexes are not available; hence we use the national CPI-U available from the Bureau of Labor Statistics to the data into constant 2011 dollars.

The other control variables we use in the analysis are widely available. The state-level real gross domestic product per capita is taken from the Bureau of Economic Analysis's Regional Economic Accounts.⁴ State unemployment rates are taken from the Bureau of Labor Statistic's Local Area Unemployment Statistics division.⁵ Total state population and the share of population aged 15-59 are from the Census Bureau's Population Estimates division.⁶

ESTIMATION

The use of state-level panels to explore the effects of minimum wages on top income shares is surprisingly sparse. Without a direct literature to base our econometric approach upon, we take the view that the voluminous literature on minimum wages and employment is a strong substitute. This literature is based largely on the seminal work of Card and Krueger (1994), which uses the difference-in-differences estimation approach (for an overview, see Neumark and Wascher 2007). While this literature is narrowly focused on the effect of minimum wages on employment and employment growth, its methodology is a good candidate because it is well vetted in the state-level panel context, and because income distributional changes are in part a consequence of the employer/employee relationship. Hence, in the respect that changes in the minimum wage result in higher or lower earnings among employees, or less employment overall, then changes in the minimum wage may have an meaningful impact on the distribution of income within a population.

The difference-in-differences approach contrasts a state's level of income inequality before and after a change in its minimum wage, relative to the counterfactual change in the other states' level of income inequality. The appropriate counterfactual

³ Available online at: <http://www.dol.gov/whd/state/stateMinWageHis.htm>.

⁴ Available online at: <http://www.bea.gov/regional/index.htm>.

⁵ Available online at: <http://www.bls.gov/lau/home.htm>.

⁶ Available online at <http://www.census.gov/popest/data/historical/index.html>.

state is important here, as states are necessarily heterogeneous in their adjustment of the minimum wage and other socioeconomic factors. To improve the quality of the counterfactuals, we follow the minimum wage and employment literature in testing a variety of space and time effects, as well as business cycle controls (see Meer and West 2013).

Accordingly, the first specification we use is the classic panel difference-in-differences estimator:

$$\text{Inequality}_{s,t} = \beta \cdot \ln(\text{MinWage})_{s,t} + \mu_s + \tau_t + \varepsilon_{s,t}, \quad (1)$$

where μ_s is the time-invariant fixed effect for state s , τ_t is the state-invariant time effect for time t , and ε is the idiosyncratic, time and state-varying error term.

Different regions within the country, however, may face heterogeneous economic shocks correlated with changes in the minimum wage. Hence, in Specification (2), we allow for time fixed effects that vary across the four census regions ($\tau_{r,t}$):

$$\text{Inequality}_{s,t} = \beta \cdot \ln(\text{MinWage})_{s,t} + \mu_s + \tau_{r,t} + \varepsilon_{s,t}. \quad (2)$$

Furthermore, to appropriately capture changes in inequality that might be correlated with changes in the state minimum wage, Specification (3) adds state linear time trends ($\eta_s \cdot t$):

$$\text{Inequality}_{s,t} = \beta \cdot \ln(\text{MinWage})_{s,t} + \mu_s + \tau_{r,t} + \eta_s \cdot t + \varepsilon_{s,t}. \quad (3)$$

Finally, Specification (4) adds additional control variables to capture variations in state-level economic climates:

$$\text{Inequality}_{s,t} = \beta \cdot \ln(\text{MinWage})_{s,t} + \mu_s + \tau_{r,t} + \eta_s \cdot t + \alpha \cdot X_{s,t} + \varepsilon_{s,t}. \quad (4)$$

In this specification, $X_{s,t}$ is a vector of explanatory variables that includes real gross state product per capita, state unemployment rates, total state population, and the share of the state population aged 15-59.

Column [1] of Table 2 shows the effect of minimum wages on the income share of the top 1% across the four specifications defined above. The results are small but fairly consistent across the econometric specifications, though statistically significant in only in the initial specification (Row 1). Taken together, the estimates in Column [1] imply an elasticity between -0.16 to -0.14. That is, a real minimum wage increase of 10% reduces the income share of the top 1% by about 1.5 percentage points.

Table 2
Effect of the Minimum Wage on Top Income Shares

Model Specification:	Dependent Variable:				Obs.
	[1] Top 1%	[2] Top 1% Growth	[3] Top 10%	[4] Gini	
(1) μ_s, τ_t	-0.0212** (0.0083)	-0.0457** (0.0216)	0.0230 (0.0167)	-0.0203 (0.0151)	1785
(2) $\mu_s, \tau_{r,t}$	-0.0214 (0.0132)	-0.0338 (0.0234)	0.0053 (0.0180)	-0.0150 (0.0164)	1785
(3) $\mu_s, \tau_{r,t}, \eta_s \cdot t$	-0.0218 (0.0167)	-0.0713 (0.0441)	-0.0123 (0.0100)	-0.0355* (0.0186)	1785
(4) $\mu_s, \tau_{r,t}, \eta_s \cdot t, X_{s,t}$	-0.0198 (0.0160)	-0.0776 (0.0464)	-0.0069 (0.0083)	-0.0329* (0.0171)	1785

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are reported in parenthesis. Sample uses annual state data from 1977-2011. Model Specification: μ_s is the time-invariant fixed effect for state s , τ_t is the state-invariant time effect for time t , $\tau_{r,t}$ is the census region time fixed effects, $\eta_s \cdot t$ is the state linear time trends, and $X_{s,t}$ is a vector of explanatory variables that includes real gross state product per capita, state unemployment rates, total state population, and the share of population aged 15-59.

It is possible, however, that the dynamics of the top 1%'s income share are more appropriate to look at, given that transitions in the levels of inequality may be slow. With this in mind, Column [2] evaluates the growth rate of the top 1%. The results are more volatile in magnitude, as one would expect given the higher standard deviation of the top 1% growth rate, but are statistically significant again in only the initial specification. Recall that Row (2) reflects changes relative to each region's linear time trend, while Rows (3) and (4) reflect changes relative to each state's linear time trend. The statistically significant results for the top 1% from Row (1) are then likely an artifact of ignoring state or regional time trends.

Column [3] of Table 2 extends this exploration by instead using the income share of the top decile. Here, the results are quite volatile (changing in both magnitude and sign), and none are statistically significant.

Finally, Column [4] tests a fourth measure, the Gini coefficient. Unlike top income shares, the Gini coefficient is a broad measure of income dispersion, capturing changes throughout the income distribution. The results here appear consistent in sign and magnitude, and are statistically significant when state linear

time trends are included (Rows 3 and 4). The overall effect from these two cases, however, is quite small, with and implied elasticity of around -0.06.

In Table 3 we present several alternative specifications to assess the robustness of the results. As a reference, Row (1) from Table 3 reproduces the baseline results for each of the four inequality measures reported in Row (4) of Table 2.

Table 3
Robustness Checks for the Effect of the Minimum Wage
on Top Income Shares

Model Specification:	Dependent Variable:				Obs.
	[1] Top 1%	[2] Top 1% Growth	[3] Top 10%	[4] Gini	
(1) $\mu_s, \tau_{r,t}, \eta_s \cdot t, X_{s,t}$	-0.0198 (0.016)	-0.0776 (0.046)	-0.0069 (0.008)	-0.0329* (0.017)	1785
(2) Census Division	-0.0128 (0.017)	-0.0803 (0.053)	0.0014 (0.011)	-0.0129 (0.019)	1785
(3) Quadratic Trend	-0.0245 (0.016)	-0.0909* (0.053)	-0.0073 (0.006)	-0.0348** (0.017)	1785
(4) Non-Indexing	-0.0163 (0.016)	-0.0772 (0.048)	-0.0073 (0.009)	-0.0360** (0.017)	1740
(5) Pre-2008 Only	-0.0170 (0.021)	-0.0607 (0.063)	-0.0153 (0.011)	-0.0335 (0.020)	1581
<i>Minimum Wage Alterations:</i>					
(6) MW_t	-0.0224 (0.020)	-0.0511 (0.053)	-0.0107 (0.010)	-0.0401* (0.022)	1785
ΔMW	0.0008 (0.002)	-0.0084 (0.008)	0.0012 (0.001)	0.0023 (0.002)	
(7) MW_t	-0.0169 (0.016)	-0.0565 (0.054)	-0.0128 (0.008)	-0.0240 (0.016)	1734
MW_{t+1}	0.0040 (0.010)	-0.0188 (0.041)	0.0076 (0.007)	-0.0092 (0.010)	
(8) MW_t	-0.0109 (0.012)	-0.0816* (0.048)	0.0027 (0.007)	-0.0207 (0.013)	1734
MW_{t-1}	-0.0142 (0.016)	0.0060 (0.047)	-0.0140** (0.006)	-0.0167 (0.016)	
(9) MW_t	-0.0184 (0.014)	-0.0981** (0.044)	-0.0026 (0.008)	-0.0302* (0.016)	1683
MW_{t-2}	-0.0046 (0.019)	0.0581 (0.058)	-0.0141* (0.007)	-0.0063 (0.020)	

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are reported in parenthesis. Sample uses annual state data from 1977-2011. Row (1) replicates row (4) from Table 2.

In Row (2) we replace the four census regions time effects (τ_{rt}) with time effects from the nine census divisions. With this specification change, we find that none of the estimated minimum wage coefficients are statistically significant, including that from the Gini coefficient in Column [4].

Quadratic state-time trends are added to the baseline specification in Row (3). Allowing for nonlinear time effect appears beneficial as both the top decile and Gini coefficient become statistically significant. The top 1% income share and growth rate, however, remain insignificant (as they are in the baseline results).

In Row (4) we drop all observations from states that shifted to indexing their minimum wage for inflation. This leads to the dropping of only 45 observations, as inflation indexing is a fairly new phenomenon. This does not appear to be a meaningful alteration, as the estimated results in Row (4) are quite similar to the baseline results in Row (1).

In Row (5), we instead use only the pre-Great Recession observations, as this period is marked by substantial volatility among the income inequality measures. With this specification change, we find that none of the estimated minimum wage coefficients are statistically significant.

Finally, Rows (6)–(9) test several alterations of the minimum wage variable: Row (6) adds a dummy variable for periods when the nominal minimum wage is changed, Row (7) adds a one period lead of the minimum wage, and Rows (8) and (9) include a one- and two-period lag of the minimum wage. Lags are useful in picking up delayed effects that might result from a slow adjustment process following a minimum wage change. A lead, on the other hand, is useful if the adjustment process begins before the implementation of a minimum wage hike. Given that there is often a delay between the passage of minimum wage legislation and its implementation, it is plausible that firms anticipate these changes and begin their adjustment process before legislative enactment.

None of these alterations in the minimum wage appear to add much power to the original minimum wage coefficients. All previously insignificant minimum wage coefficients from the baseline specification (Row 1) remain statistically insignificant. In addition, the estimates from the Gini coefficient measure (Column 4) become statistically insignificant when either a one-period lead or lag of the minimum wage is included. Moreover, the alternative minimum wage variables are also generally statistically insignificant across the estimations. The only exception is with the top decile measure in Column [3] when using either a one- and two-period lag (Rows 8 and 9).

CONCLUSION

This paper has employed a state-level panel to evaluate the relationship between minimum wage laws and top income shares. Our results indicate that the relationship is not particularly robust. While consistently negative in sign, we find that small changes in the econometric specification, as well as the measurement of income inequality, lead to statistically insignificant associations. In total, we report twelve alternative model specifications using three different measures of top income shares; of these 36 estimated minimum wage coefficients reported, only *two* are statistically significant at the $p < 0.05$ level (see Table 2, Row 1, Columns 1 and 2). Two is also approximately the same number one would expect to find falsely significant at a 95% confidence level, given the number of estimations conducted.

However, the results are marginally more robust when the Gini coefficient is used instead of a top income share measure (see Column 4 in Tables 2 and 3). This is less surprising given that prior empirical research has found positive effects from the minimum wage occurring within the lower-half of the income distribution (see DiNardo, Fortin, and Lemieux 1996, Lee 1999, Card and DiNardo 2000, Dickens and Manning 2004, and Autor, Katz and Kearney 2008). One would expect such a result if minimum wage changes have spillover effects upon workers earning at and near the minimum wage. Spillovers unto those in the upper-half of the income distribution have also been found (see Autor, Manning, and Smith 2016). Whether spillovers extend into the upper-tail (e.g., the top 1%) has not, to our knowledge, been considered prior to this paper.

The lack of statistical robustness from our results thus casts doubt on whether increases in the minimum wage can help mitigate the rise in top income shares. In a broader sense, investing in education and skills may be a better *long-term* remedy for reducing income disparities (see Goldin and Katz 2008). The public's appetite for long-term solutions, however, is famously fickle. This short-term focus may in fact be driving parts of the public discussion with respect to the minimum wage. The evidence presented here, nonetheless, would suggest caution is warranted in taking this view.

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