

# Minimum Wages and the Rigid-Wage Channel of Monetary Policy<sup>1</sup>

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**Abstract:** We argue that wage rigidity induced by the minimum wage is an important channel for the transmission of monetary policy. Intuitively, while expansionary monetary policy leads flexible prices to increase, the nominal minimum wage is held fixed by legislation and therefore falls in real terms. We present empirical evidence showing that the peak effect on employment of a 1% federal-funds rate shock is 2.7 percentage-points higher in states where the cost share of workers affected by the minimum wage is at its 90<sup>th</sup>-percentile value compared to its 10<sup>th</sup>-percentile value. This result is robust to a wide variety of econometric specifications and, in our preferred specification, accounts for 39% of the total effect of monetary policy during the Volcker era. Verifying the mechanism, we find substantially larger effects of monetary policy on near-minimum-wage employment relative to higher-wage employment. We conclude that minimum wages, and rigid wages more generally, are a crucial channel through which monetary policy operates.

## 1 Introduction

The core source for efficacy of monetary policy is the existence of price and wage rigidities, and the minimum wage is an important legislatively-imposed wage rigidity in modern economies. It may thus come as a surprise that little attention has been paid to the intersection of the minimum wage and monetary policy literatures, and no systematic empirical investigation of the role that minimum wages play in mediating monetary policy efficacy has been undertaken as of yet. We

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aim to fill this gap in the literature.

We argue that the minimum wage is a compelling context in which to study the importance of the rigid-wage channel of monetary policy. The empirical literature on nominal wage rigidity has yielded mixed evidence about the extent to which wages are downwardly rigid. For example, recent evidence from administrative data suggests that only 7-8% of job stayers experience no year-to-year wage changes during normal times, a lower fraction than the share of total U.S. employment near the minimum wage in the late 1970s and early 1980s, and up to 30% of job stayers experience wage cuts during recessions (Kurmann and McEntarfer 2019). Thus, focusing on the minimum wage – a wage which is known by definition to be nominally-rigid and is binding for a non-trivial fraction of the population – allows for a directed examination of the wage-rigidity channel of monetary policy.

Our analysis exploits variation across states and time in the share of workers affected by the minimum wage. We find robust evidence that a higher share of minimum-wage labor leads to stronger effects of monetary policy. In particular, for our baseline regression specification, we obtain data on monetary policy shocks from Coibion et al. (2017), who expanded the original Romer and Romer (2004) narrative shock series beyond 1996. We obtain monthly data on state-level employment from the Quarterly Census of Employment and Wages (QCEW). And we compute the labor share of workers affected by the minimum wage in each state and year using the Current Population Survey (CPS) and GDP and employee compensation data from the Bureau of Economic Analysis (BEA). Combining these sources, we run a regression specification that has much in common with canonical national-level monetary policy regressions, albeit adapted for a state panel setting by featuring two-way clustering on both the time variable (year and month) and the state variable in order to robustly account for complex autocorrelation structures. That is,

$$\Delta E_{s,t} = \alpha + \sum_{i=0}^{48} \beta_i Shock_{t-i} + \gamma MWShare_{s,t} + \sum_{i=0}^{48} \delta_i Shock_{t-i} \cdot MWShare_{s,t} + \sum_{i=1}^{48} \eta_i \Delta E_{s,t-i} + \varepsilon_{s,t},$$

where  $\Delta E_{s,t}$  represents month-over-month employment growth and  $MWShare_{s,t}$  represents the labor share of workers affected by the minimum wage. From this specification, we find evidence that the short-run employment fluctuation induced by monetary policy is significantly higher in states where the share of the minimum wage workers is higher. The peak effect on employment

of a 1 percentage-point Romer and Romer Federal Funds Rate shock on employment growth is approximately 2.7 percentage points stronger where the *MWShare* is at its 90<sup>th</sup>-percentile value compared to its 10<sup>th</sup>-percentile value.

To illustrate the variation we exploit, Figure 1 shows the federal real minimum wage over time and the distribution of state shares of wage and salaried employment affected by the minimum wage in each year. We follow the finding of Autor et al. (2016) that minimum wages have a statistically-significant effect on the earnings of workers up to the 15<sup>th</sup> percentile of the national wage distribution during their sample period, noting that this implies spillovers of the minimum wage on wages up to 28% above the legislated minimum. Figure 1 displays a boxplot summarizing state variation over time in the share of wage and salaried workers earning wages less than 128% of the minimum wage in the state. Figure 1 reveals a substantial fraction of workers are affected by the minimum wage, particularly around 1980 – with much variation across states. This suggests we should be able to detect whether the rigid-wage channel of monetary policy is operative using state-level employment data. In our literature review, we argue that a threshold of 28% above the legislated minimum is perhaps a conservative number, given other estimates in Autor et al. (2016) and earlier research by Lee (1999), but we show that our empirical results remain strongly statistically-significant and similar when focusing on either a narrow band of workers earning wages very near the legislated minimum or a broader set making up to 40% above the legislated minimum.

We apply a battery of other robustness checks to our finding as well. We run a regression analogous to a difference-in-differences specification, adding state and year fixed-effects to the baseline regression to focus on cross-sectional variation exploiting differences in MW-affected shares across states at a given moment in time. Observing that changes in the share of minimum wage workers can be driven either by plausibly-exogenous factors such as minimum wage changes or by more endogenous factors such as changes in a state’s industry composition, we construct a Bartik-type variable that accounts for the latter effect and add it to our baseline regression. In an alternative approach to isolating the plausibly-exogenous variation, we run an IV specification instrumenting the state MW-affected share with the state minimum wage. Our main result is robust

to all these alternative specifications.

Additionally, we replace the Romer and Romer narrative shocks in our main specification with VAR shocks, and the result is not much changed. We run the baseline specification in the Canadian context – using Canadian data on provincial minimum wages, employment, and monetary shocks to exploit variation in MW-affected shares across provinces – again finding the same significant relationship. Next, we focus specifically on within-state time-series variation in the MW-affected share by interacting the Romer and Romer shock series with state-fixed effects and, separately, with a time trend variable. Using this approach, we again confirm our findings. Comparing these findings to the total effect size of monetary policy during the Volcker era, we find that the rigid-minimum-wage channel accounts for 39% of monetary policy’s total effect on employment.

To better interpret our results and assess the realism of the magnitudes we find, we set up a model of monetary policy in which minimum wages are the only source of non-neutrality. The key assumption is that minimum wages in each state are binding: low skill workers would like to supply more labor than firms demand. Formally, low skill workers take their labor supply as given and, because the minimum wage is exogenous, firm demand determines the quantity of low skill labor in equilibrium. Expansionary monetary policy increases capital rental rates, endogenous wages, and prices, leading to reductions in the real cost of low skill labor for firms. Factor price changes induce both substitution and scale effects: under expansionary policy, firms substitute towards more use of low skill labor and also scale up their operations. Because our mechanism is fundamentally a supply shock, the model predicts larger effects on tradable employment than non-tradable employment; intuitively, more production shifts towards the places where it has become relatively cheaper when that production can be consumed nationally rather than just locally. The tradable sector is also marked by large capital shares, which, combined with the large elasticity of substitution between capital and minimum wage labor, contribute to minimum wage labor demand being relatively more elastic in tradable sectors.

To test the mechanism suggested by the model more clearly and confirm that the effects are indeed driven by minimum wage workers, we run an individual-level specification that compares

the effects of monetary policy shocks on near-minimum-wage workers relative to higher-wage workers. We find that, in response to a contractionary (expansionary) shock, workers who are initially near the minimum wage are significantly more (less) likely to become unemployed – relative to workers who are initially well above the minimum wage. Finally, we separately examine the effects on employment in tradable versus non-tradable sectors, finding that the effect is somewhat larger among tradable sectors. This result is consistent with our channel – which is fundamentally a supply shock – and inconsistent with the competing explanation that all effects we measure are driven by differences in the MPC across states, as local demand shocks tend to boost *non-tradable* employment.

We conclude that minimum wages are an overlooked but important factor in determining the efficacy of monetary policy, confirming the more general hypothesis that wage rigidity is a key contributor to monetary non-neutrality. Indeed, our empirical magnitudes suggest that a sizeable fraction of monetary policy’s effectiveness is filtered through precisely this channel.

## **2 Literature Review**

Our findings contribute directly to the literature on the rigid nominal wage channel of monetary policy. While models generating non-neutrality of monetary policy through nominal wage rigidity are common in the literature, there are no empirical tests of this channel in settings where the extent of wage rigidity is not in question. This is important because the empirical evidence on the extent to which nominal wages are rigid is quite mixed. We compile a brief history of the wage rigidity literature here, including comparisons between the magnitude of the minimum wage share of employment and the fraction of employment subject to rigid (non-minimum) wages, before discussing additional literatures to which our paper contributes.

Early microdata evidence on downward nominal wage rigidity from the PSID, which contains individual-level wage changes, was relatively unfavorable. Fallick, Villar, and Wascher (2020) describe this evidence: McLaughlin (1994) and Lebow, Stockton, and Wascher (1995) do not find strong evidence of downward nominal rigidity, though Kahn (1997) finds some evidence for hourly wage workers. Later work, e.g. Altonji and Devereux (2000), found that the mixed evidence

on downward nominal wage rigidity in the PSID might be due to measurement error in reported wages.

Moving beyond the PSID, the evidence on downward nominal wage rigidity in small employer surveys and case studies is mixed but somewhat more favorable. While Wilson (1999) and Altonji and Devereux (2000) find supporting evidence, Blinder and Choi (1990) find that five of the nineteen interviewed firms had recently cut wages despite the booming economy.

Studies using the CPS – e.g. Daly, Hobijn, and Lucking (2012) and Daly and Hobijn (2014, 2015) – find an increase between 2007 and 2011 in the fraction of workers in the same job (hereafter, “job stayers”) who report no change in their wage relative to the previous year. These studies are reassuring for the rigid nominal wage hypothesis, since the ORG component of the CPS, like the PSID, contains reported hourly wages, where we may be most likely to find rigidity. One issue with these studies is they focus on the fraction of workers with no wage change rather than focusing on the fraction of workers who receive wage cuts.

More recent studies turn to large surveys of employers that are less likely to suffer from measurement error. An early example is Lebow, Saks, and Wilson (2003). The authors use microdata from the BLS’s Employment Cost Index (ECI) program, which collects information on compensation for thousands of jobs across thousands of establishments. They find stronger evidence of downward nominal wage rigidity than was typically found in panel data on individual wages: from 1981 to 1999, about 14.5% percent of year-to-year wage and salary changes were negative, and about 18.5% were zero. Fallick, Villar, and Wascher (2020) turn again to this data and find increased downward nominal wage rigidity during and after the Great Recession.

Administrative data point to the importance of analyzing wage cuts and wage freezes separately. Kurmann and McEntarfer (2019) use data collected by the unemployment insurance office in Washington state, which covers over 95% of private-sector employment in the state. They find that, during the Great Recession, the fraction of job stayers who are paid the same wage as a year earlier increases from 7-8% to 16% and then gradually returns to its pre-recession average. The fraction of job stayers who experience wage cuts increases during the recession from 20% to 30%, and the fraction of stayers who experience declines in annual earnings increases to 39%,

suggesting some role for composition effects in hours. Using the same data, Jardim, Solon, and Vigdor (2019) find that for every quarter of year-to-year wage changes they in their data, at least 20% of job stayers experienced nominal wage reductions.

Elsby and Solon (2019) survey evidence from employers' payroll records and pay slips in multiple countries, which includes the research from Kurmann and McEntarfer (2019) and Jardim, Solon, and Vigdor (2019) cited above. They find that, except during periods of high inflation or when nominal wage cuts are legally-prohibited, an average of 15-25% of job stayers receive nominal wage cuts from one year to the next.

Comparing these measurements of wage rigidity with our estimates of the MW-affected share highlights why low wage employment constitutes a valuable setting in which to study the wage rigidity channel of monetary policy. Recall from Figure 1 that the share of employment affected by the minimum wage in the median state was around 25% in 1980. This is a larger number than the fraction of job stayers paid the same wage as a year earlier during the Great Recession, as found by Kurmann and McEntarfer (2019), and is over three times the magnitude of the fraction of stayers they found receive no year-over-year wage changes during normal times. It is also above the fraction of job wages and salaries that received no year-over-year wage changes from 1981-1999, as found by Lebow, Saks, and Wilson (2003). 25% is also the upper bound on the fraction of job stayers who receive nominal wage cuts from one year to the next, as found by Elsby and Solon (2019) in more recent data. It is 10 percentage points above the fraction of job wages and salaries that experienced reductions from 1981-1999, as found by Lebow, Saks, and Wilson (2003).

We also contribute to the literature that develops tests of underlying economic mechanisms relying on differential effects of shocks on tradable and non-tradable employment. Our channel – which is fundamentally a supply shock – suggests larger effects on tradable employment than non-tradable employment. If, on the other hand, the MW-affected share in a region is correlated with the marginal propensity to consume (MPC) of a region, and this MPC channel is the true underlying mechanism, we might expect that monetary policy leads to larger demand shocks in these regions. Research shows that local demand shocks often lead to larger effects on non-tradable

employment than on tradable employment, the opposite of what we would expect from our minimum wage channel. In two papers, Mian, Rao, and Sufi (2013) and Mian and Sufi (2014) develop local demand shocks using changes in housing market wealth and argue these shocks have effects on non-tradable employment but no effects on tradable employment. Chodorow-Reich et al. (2021) similarly argue that local demand shocks generated from changes in stock market wealth affect non-tradable employment but not tradable employment. We think our work further validates the usefulness of analyzing tradable and non-tradable employment when testing underlying economic mechanisms.

Additionally, our research relates to an extensive literature on the effects of minimum wage changes on employment and inequality. There is limited consensus in this literature on the effect of minimum wage changes on employment (Neumark 2017). It is beyond the scope of this paper to summarize this literature, but we will point to some key research. Well-known papers such as Card and Krueger (1994) and Dube, Lester, and Reich (2010) find no adverse effects of minimum wage increases on employment. More recent evidence includes Cengiz et al. (2019), which also finds no evidence of negative effects on overall employment but does find some effect on employment in tradable sectors. Neumark and Wascher (1992), on the other hand, find that a 10% increase in the minimum wage causes a 1-2% decline in employment among target groups such as teenagers and young adults. More recent work by Clemens and Wither (2019) finds that a 9% minimum wage increase reduces employment by as much as 9% in a key target group. Reich, Allegretto, and Godoy (2017) analyze Seattle's 2015-16 minimum wage increase from \$9.47 to \$11 and find it led to no disemployment effects on the food services industry (argued to have a high share of minimum-wage workers). Conversely, Jardim, Solon, and Vigdor (2019) use administrative data beyond the food-services sector to study the same minimum wage increase, finding the data points to an elasticity of -0.9, and the subsequent increase to \$13 point to large disemployment effects, an elasticity of -2.6.

The disagreements between these papers seem partially due to researchers focusing on different groups of minimum wage workers, using different kinds of data, and studying different time periods. The finding that minimum wage changes have fairly mild effects on restaurant and

teenage employment seems to be the most robust of the findings in this literature. Studies focusing on all minimum wage workers show substantially more disagreement, and studies using longitudinal data (e.g., Clemens and Wither 2019) to study this group rather than repeated cross-sections (e.g., Cengiz et al. 2019) sometimes find much larger elasticities (e.g., the -1 values cited above).

Of course, our calculation of MW-affected shares uses the literature examining the effects of minimum wage changes on wage inequality. We've drawn primarily from Autor et al. (2016), which shows that increasing minimum wages has a statistically-significant effect on the earnings of workers up to the 15<sup>th</sup> percentile of the national wage distribution. We noted before that these estimates may be conservative. Autor et al. (2016) find statistically-significant effects up to the 15<sup>th</sup> percentile, but point estimates remain suggestive of spillovers up to the 25<sup>th</sup> percentile, and significance is retained at some percentiles between the 15<sup>th</sup>-25<sup>th</sup>. Earlier research by Lee (1999) suggests that minimum wage increases may spill over even further, up to the median of the wage distribution. Conversely, though, Autor et al. (2016) argue that these estimates are upward biased and argue their estimates correct this bias. Indeed, Autor et al. (2016) argue that their results could be affected by measurement error in a way that results in overestimates of spillovers, so on this basis, our use of the 15<sup>th</sup> percentile of the wage distribution as the limit on minimum wage spillovers may *not* be conservative. This is why we verify that our results hold for adjusted MW-affected shares only focusing on workers whose wages are within a narrow band of the legislated minimum. But the magnitude of our results, we think, lends support to the hypothesis that wage rigidity induced by the minimum wage extends somewhat up the wage distribution.

Finally, we draw on an extensive literature, with a diverse methodological history, devoted to studying the effects of monetary policy on economic outcomes. A key division in the literature on the effects of monetary policy exists between papers using vector autoregression (VAR) frameworks, narrative approaches, and high-frequency identification approaches. While these literatures are impressive in depth, defying a systematic listing here, key examples of VAR papers include Bernanke and Blinder (1992), Leeper, Sims, and Zha (1996), Bernanke and Mihov (1998), Christiano, Eichenbaum, and Evans (1999), Uhlig (2005), and Bernanke, Boivin, and Elias

(2005). Key examples of narrative papers include Romer and Romer (1989), Romer and Romer (2004), and Coibion et al. (2017). These branches of the literature find significant effects of monetary policy on real outcomes – but the effects found in the narrative literature are typically much larger. Ultimately, our regressions will interact monetary policy shocks derived in these literatures with MW-affected shares that we compute in the data. The high-frequency identification literature includes contributions from Cook and Hahn (1989), Kuttner (2001), Cochrane and Piazzesi (2002), Gürkaynak, Sack, and Swanson (2005), and, more recently, Gertler and Karadi (2015) and Nakamura and Steinsson (2018). We will not use the monetary policy shocks from these papers, since the data used to form these shocks is unavailable for the time periods when the share of minimum wage employment was high.

### **3 Empirical Framework**

Is the efficacy of monetary policy mediated, in part, through the minimum wage? To answer that question, we begin with some empirical analysis. We use standard data sources for employment – the QCEW and the CPS – and an adapted version of a very simple and standard specification from the narrative monetary policy literature. We subsequently branch out from this specification and run a broad variety of robustness checks intended to encapsulate many potential critiques of the baseline specification.

#### **3.1 Data**

We obtain data on narrative monetary policy shocks from Coibion et al. (2017). Coibion et al. follow the technique devised by Romer and Romer (2004), who obtained narrative records of the Federal Reserve’s intentions for the federal funds rate around FOMC meetings and regressed this series on internal Fed forecasts “to derive a [monthly] measure free of systematic responses to information about future developments.” The Romer and Romer (2004) series of monetary shocks has become one of the canonical sources of exogenous variation used in the monetary policy literature. Because the series initially terminated in 1996, however, Coibion et al. extended it through 2015. We also obtain an alternative VAR shocks series from Coibion (2012), a paper

dedicated in part to explaining why the Romer and Romer (2004) shocks generate such large effects of monetary policy. This VAR series yields effects of monetary policy close to those found in Christiano, Eichenbaum, and Evans (1999), somewhat smaller than those found in Leeper, Sims, and Zha (1996) and somewhat larger than those found in Bernanke and Blinder (1992); in particular, these shocks lead to output effects that are roughly six times smaller than those found in Romer and Romer (2004). Further, the VAR series is much less sensitive to the inclusion of monetary policy episodes in 1980 that drive the estimated Romer and Romer (2004) shock effects to be large.

For two key reasons, we will not use monetary policy shocks derived from the high-frequency identification literature, which includes Cook and Hahn (1989), Kuttner (2001), Cochrane and Piazzesi (2002), Gürkaynak, Sack, and Swanson (2005), and, more recently, Gertler and Karadi (2015) and Nakamura and Steinsson (2018). First, we have the most power to detect our results in the late 1970s and early 1980s, when there was a relatively high share of minimum wage workers. The earliest futures data used in Nakamura and Steinsson (2018) begins in 1995, and their measurements of real interest rates require TIPS data. TIPS were issued beginning in 1997. We cannot use the futures rate surprises from Gertler and Karadi (2015) because the key data is available from 1991 to 2012, leaving out the core period during which we want strong, exogenous monetary policy variation. Second, we question the power of these shocks in our context because some of these series are able to detect effects of monetary policy on financial variables but not real variables such as output and employment.

For data on employment, we turn to the Quarterly Census of Employment and Wages (QCEW), which has collected population data on employment by county by industry in the United States since 1937. Despite the name of the dataset, employment data is available at the monthly level<sup>4</sup>. Digitized data from January 1975 onward is readily available for download on the Bureau of Labor Statistics (BLS) website. For our baseline regression specification, we use the state-level figures aggregated across all industries, but for certain alternative specifications – such as our

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<sup>4</sup> The reason for the apparent discrepancy is that the census is conducted quarterly but asks employers how many workers were on their payroll at the end of each of the three months of the preceding quarter.

specifications with Bartik-style controls – we make use of the underlying industry-specific data.

We use the CPS Outgoing Rotation Groups (CPS-ORG) to compute the share of minimum wage workers as a proportion of all workers by state and year. Households in the CPS sample respond to the questionnaire for four months in a row; they are then out of the sample for eight months; finally, they return to the sample for another four months. It is possible to exploit this panel structure following the procedure for longitudinal matching of CPS respondents by Madrian and Lefgren (2000), which we do in our individual-level analysis. At the end of each of the two four-month blocks during which a household is present in the CPS sample, they are asked a specific set of questions not asked in other months. These questions – which include amongst them an explicit question about what the respondent’s hourly wage is – make up the Outgoing Rotation Groups questionnaire. Because the monthly size of the CPS is 60,000, this means that the monthly size of the ORG is (approximately) 15,000 – an annual sample size of 180,000.

By merging this data with Vaghul and Zipperer’s (2016) dataset on historical state and federal minimum wages, we can identify minimum-wage-affected workers as any wage or salaried worker whose computed hourly wage is below 128% of their state’s binding minimum wage. Our results remain significant, though with different magnitudes, if we instead use a tighter percentage band—defining near minimum wage workers as those within 10% of the minimum wage—and they also remain significant, although again with changed magnitudes, if we instead define MW-affected workers as having wages below 140% of their state’s binding minimum wage. The 140% threshold is associated with minimum wage changes affecting wages up to the 25<sup>th</sup> percentile of the wage distribution, in accord with the less conservative estimates of spillovers associated with the Autor et al. (2016) analysis discussed in our introduction and literature review. We can then compute the MW-affected share of payroll in a state as the total payroll to minimum wage workers in the state divided by total payroll in the state.

To compute the MW-affected share in the state, we multiply the MW-affected payroll share computed above by the labor share in the state. The Bureau of Labor Statistics (BLS) describes how they compute the national labor share, and we implement this procedure at the state level. We divide total compensation in a state by total GDP in that state, using data published by the Bureau

of Economic Analysis (BEA). In the BLS calculations, this share is then adjusted upwards for proprietors' incomes due to their own work at their businesses, which is not included in total compensation as measured by the BEA. We elect not to include this adjustment to keep our shares conservative; this means we are assuming that business proprietors whose wages fall within our definition of MW-affected wages are not affected by our channel<sup>5</sup>. We show how our computed labor share without the proprietor's adjustment at the national level compares to the BLS's labor share in the appendix.

It is worth noting that an alternative approach to using the CPS-ORG is to use the Current Population Survey's Annual Social and Economic Supplement (CPS-ASEC) to compute the share of minimum wage workers. The ASEC commenced in 1962 and since its inception has asked respondents their total wage income, weeks worked, and hours worked per week over the last year. From this, it is possible to compute each individual's hourly wage. However, prior to 1977, the aforementioned variables were binned, so weeks and hours worked – and thus hourly wage – can only be approximately known. And compared to the ORG's annual sample size of 180,000, the ASEC has an annual size of 60,000. As a consequence of its lower sample size, the approximation implicit necessary as a part of the preceding process, and the fact that the QCEW data is only available from 1975 onward anyway, we use the ORG instead of the ASEC. Having said this, our findings remain statistically-significant and qualitatively similar if we use the ASEC data.

We also obtain data on some additional control variables for robustness checks. We obtain data on per-capita bank deposits by county from the FDIC and data on personal income per-capita by county from the Bureau of Economic Analysis (BEA). We obtain a Canadian narrative monetary shocks series (constructed analogously to the Romer and Romer shocks) from Champagne and Sekkel (2018), and we obtain Canadian data on monthly employment and the share of minimum wage workers by province from Statistics Canada's Monthly Labour Force Survey Public Use Microdata File (PUMF).

### **3.2 Identification**

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<sup>5</sup> Including proprietors does not have a major effect on the magnitude or statistical significance of our results.

Our baseline specification is an adapted version of a standard autoregressive distributed-lag (ARDL) monetary policy regression with narrative shocks. We add interaction effects between the shock terms and the MW-affected share, and we additionally two-way cluster our standard errors at the state and time level in order to allow for complex correlation structures induced by the fact that our dataset is a panel dataset but the monetary shock series is state-invariant.

$$\Delta E_{s,t} = \alpha + \sum_{j=0}^{48} \beta_j Shock_{t-j} + \gamma MWShare_{s,t} + \sum_{j=0}^{48} \delta_j Shock_{t-j} \cdot MWShare_{s,t} + \sum_{j=1}^{48} \eta_j \Delta E_{s,t-j} + \varepsilon_{s,t}, \quad (1)$$

where  $\Delta E_{s,t}$  denotes employment growth in state  $s$  at time  $t$ ,  $MWShare_{s,t}$  denotes the minimum-wage labor share of costs in state  $s$  at time  $t$ , and  $Shock_t$  denotes the (nationwide) monetary policy shock at time  $t$ . The  $\delta_i$ 's are the main coefficients of interest, as they reveal the interaction effects between the monetary policy shocks and the MW-affected share – i.e., whether the effects of monetary policy are larger where the MW-affected share is higher.

In various robustness checks, we enhance this specification with additional control variables and/or different approaches to identification. We add state and time fixed-effects to the specification to account for all time-varying, state-invariant and state-varying, time-invariant confounds. Separately, we add controls for the interaction effect of a couple of other variables with the shock series: bank deposits per-capita and per-capita income (proxying for the marginal propensity to consume). The idea is that there may remain some crucial variables correlated with MW-affected share that could plausibly be the true channel for monetary policy efficacy, rather than the MW-affected share itself. Additionally, we simultaneously interact state fixed-effects with the shock series and, separately, interact a linear time trend with the shock series. This allows for monetary policy efficacy to vary across states and be changing over time for reasons unrelated to the MW-affected share.

Observing that changes in the MW-affected share in a state can be driven either by plausibly-exogenous factors such as minimum wage changes or by more endogenous factors such as changes in the share of each industry in that state's employment, we construct a Bartik-type variable that controls for the latter effect and add it to our baseline regression. In particular, we construct the variable by computing

$$\Delta S_{s,t} = \sum_j Shift_{-s,j,t} \cdot Share_{s,j,t-1}, \quad (2)$$

where  $Shift_{-s,j,t}$  represents the national-level growth of employment in industry  $j$  over time period  $t$  (calculated as a leave-one-out average) and  $Share_{s,j,t-1}$  represents the employment share of industry  $j$  in total state- $s$  employment in the preceding time period  $t-1$ . This shift-share isolates the national-level, non-idiosyncratic component of growth in employment that stems from broader trends. Adding this control to the specification should help ensure that the effect we are finding is not driven by that more endogenous source of minimum-wage-share variation.

As an alternative approach to isolating the plausibly-exogenous variation, we run an IV specification instrumenting the state MW-affected share with the state minimum wage itself. In particular, for our first-stage, we instrument the direct effect and the interaction effects involving the MW-affected share with the corresponding minimum wage variables as follows,

$$MWShare_{s,t} = \omega + \rho MinWage_{s,t} + \sum_{j=0}^{48} \theta_j Shock_{t-j} \cdot MinWage_{s,t} + u_{s,t} \quad (3)$$

$$Shock_{t-i} \cdot MWShare_{s,t} = \chi_i + \pi MinWage_{s,t} + \sum_{j=0}^{48} \phi_j Shock_{t-j} \cdot MinWage_{s,t} + v_{s,t} \quad (4)$$

and the second-stage constitutes placing the predicted values of the left-hand-side variables from these first-stage regressions back into our baseline specification.

## 4 Results

### 4.1 Main Results

Beginning with the baseline specification given by Equation (1), Figure 2 depicts its results in the form of an impulse response function cumulating the interaction effect over time. The error bands represent 90% confidence intervals. For this impulse response function and all others, we obtain standard errors on the cumulated effects using Monte Carlo methods. We take 1000 draws from a multivariate normal distribution with mean and variance-covariance matrix corresponding to the point-estimates and variance-covariance matrix of the regression coefficients. For each draw from the distribution, we compute the implied response of output to a 1pp shock after  $m$  months, for each  $m \in (0, 48)$ . We then obtain the standard error for the response after  $m$  months by taking the standard deviation over the 1000 aforementioned implied responses after  $m$  months.

Note that the magnitude of the interaction effect peaks at approximately -0.6. Thus the figure can be interpreted as follows: a 1 percentage-point *higher* MW-affected share corresponds to 0.6 percentage-point *lower* employment growth (at peak) resulting from a 1 percentage-point Romer and Romer contractionary monetary policy shock (i.e., a 1 percentage-point unexpected increase in the Fed Funds Rate). Phrased more intuitively, in response to a 1 pp contractionary shock, a state at the 90<sup>th</sup> percentile of the MW-affected share will experience a decline in employment that is approximately 2.7 percentage-points greater (at peak) than a state at the 10<sup>th</sup> percentile of the MW-affected share.

A major concern is whether that the industries which have the highest share of minimum wage workers might just be the industries that are most affected by monetary policy – for reasons unrelated to the MW-affected share itself. If these industries are concentrated in specific states, that could be driving our results. To deal with this concern, Figure 3 turns to the Bartik controls, adding them to the baseline specification. The idea here is that controlling for the Bartik instrument purges the component of employment growth driven by broad industrial trends; the remaining unexplained left-hand-side variation in economic growth is not a consequence of which industries happen to be concentrated in which states. As can be seen in Figure 3, the effect survives and, indeed, is little changed from baseline.

Now, suppose some states or some years just have persistent differences in employment growth from others. Given that MW-affected shares also vary across places and time, this could potentially produce a spurious result. We address these potential concerns by adding state and time fixed-effects to the baseline specification. The resulting impulse response function is plotted in the left panel of Figure 4. Notably, the effect not only survives – it is no less strongly significant than in the baseline specification. The magnitude, however, is (non-significantly) smaller by a factor of approximately one-half. The right panel of Figure 4 adds *both* the Bartik control and the state and time FEs to our baseline specification, combining the desirable characteristics of both robustness checks. Here, too, the effect retains significance with a similar magnitude.

A broader concern is whether certain states are more responsive to monetary policy for reasons unrelated to – but potentially correlated with – the minimum wage. For example, poorer

states are likely to have a higher average marginal propensity to consume, which should boost monetary policy efficacy through more traditional channels. Similarly, suppose the efficacy of monetary policy is declining over time for, again, reasons unrelated to the minimum wage. Because the share of minimum wage workers is also declining over time, this could plausibly pollute the coefficients we estimate in the baseline specification. Consequently, we interact the Romer and Romer shock series (and its 48 lags) with state fixed-effects and, separately, with a time trend variable, thereby eliminating persistent differences across states in monetary policy efficacy and national differences over time in monetary policy efficacy as potential confounds. Effectively, in this setting, our identifying variation is time-series variation in the MW-affected share within states. Figure 5 shows that the results remain significant, and the magnitude scarcely changes.

Jorda (2005) pioneered the local projection approach estimating impulse responses, which has become a well-recognized alternative to distributed-lag and VAR approaches. While Choi and Chudik (2019) find that – in contexts such as our own where the shock series is observed – distributed-lag specifications similar to those we use for the majority of our analysis perform best in terms of mean-squared error, the local projection approach is nonetheless a valuable alternative. In Figure 6, we display the results of a local projection estimation. It yields the same qualitative result as the rest of our specifications: a higher MW-affected share translates into significantly greater monetary policy efficacy.

In Figure 7, we interact a variety of different control variables with the Romer and Romer shock series (and its 48 lags); we add these terms to the baseline specification to help demonstrate that the effect we find is not due to correlation of the MW-affected share with other important variables that mediate the effects of monetary policy. We can see that doing this with bank deposits per-capita and per-capita income – two variables which are particularly likely to correlate with MPC – do not materially change our result from baseline. Nor does using the cost share of workers covered by union contracts or the cost share of government workers – other variables relating to nominal wage rigidities. Nor, finally, does using the ratio of state and local government debt to revenue or the real value of debt per capita – measures relating to the stock of the government's

nominal obligations outstanding.

Variation in the state MW-affected share may come from a variety of sources – including changing industry shares within the state. Again, this source of variation may be somewhat endogenous. As a result, a somewhat different approach from the Bartik control of factoring out this industry-correlate-driven variation in the MW-affected share is to instrument for the MW-affected share with the legislated state minimum wage. This isolates variation driven by political decisions on the part of the state legislature, plausibly a more exogenous source of variation than changing industry shares. The left panel of Figure 8 turns to this IV strategy. Again, the result survives; the peak magnitude of the employment response, however, increases to double that of the baseline specification, though our previous results remain mostly within the standard error bars of this point estimate.

Continuing on the theme of alternate sources of variation, the specification represented in the right panel of Figure 8 makes use of the VAR monetary policy shocks from Coibion (2012) instead of the Romer and Romer shocks. The monetary policy literature has proceeded along two main strands – one pursuing narrative approaches and the other pursuing more structural VAR-based approaches. We aim to show that our result endures regardless of which shock series we use; it's not an artefact of one approach or the other. While the shape of the impulse response function is somewhat different – with the effect peaking some two years later than in the baseline specification – the takeaway is the same: monetary policy is significantly more effective where the share of minimum wage workers is higher. The peak magnitude of the effect on employment using VAR shocks is slightly less than half that of the baseline specification using Romer and Romer shocks.

Another shock series that can be used is the series of raw Federal Funds Rate (FFR) shocks itself. Due to endogeneity concerns, this approach is not typically favored for estimating the aggregate effects of monetary policy. Federal Funds Rate changes are often made due to national economic considerations. However, in a setting such as ours concerned with cross-state comparisons, these concerns are less salient – especially if one includes time fixed-effects in the regression specification. The left panel of Figure 9 displays the results of a specification identical to the baseline specification, albeit with FFR shocks instead of Romer and Romer shocks on the

right-hand-side. The right panel of Figure 9 adds state and time fixed-effects to this specification. In both cases, the result that monetary policy is significantly more effective in the context of a higher MW-affected share endures; however, the magnitude of this effect is noticeably smaller than in the baseline specification – by a factor of four in the latter case.

Figure 10 plots the results of running the baseline specification on the province-level Canadian data. As in the case of the VAR robustness check, we do not necessarily have a reason to believe that the US data is inferior to the Canadian data (or vice versa) – we merely regard it as a second laboratory in which to test our hypothesis and provide evidence of its generality. Again, despite a somewhat modified shape of the impulse response function and a larger magnitude relative to the baseline specification, the evidence remains that a higher share of minimum wage workers significantly boosts monetary policy efficacy.

All of the above specifications have relied on the MW-affected share as computed using the finding from Autor et al (2016) that effects of minimum wages spill over to workers higher up in the wage distribution, with significant positive effects found on workers wages' up to the 15<sup>th</sup> percentile. However, it is important to verify that the effects are strongest for the minimum-wage workers themselves. Conversely, Autor finds *non-significant* positive effects on workers' wages as high as the 25<sup>th</sup> percentile of the wage distribution. Consequently, to show that our results are not driven by the specific definition of the MW-affected share that we choose, we compute alternate versions of the MW-affected share – one narrower and one broader than our baseline definition. Firstly, we compute a version that focuses specifically on minimum-wage workers and (virtually) no-one else – individuals earning wages in a 10% band around the minimum wage. Secondly, we compute a version that focuses on workers with wages up to the 25<sup>th</sup> percentile of the wage distribution. Figure 11 shows the results of versions of the baseline specification using these two alternatively-defined MW-affected shares. The result remains significant in both cases. It is worth noting that, as expected, the effect is strongest for the workers nearest the minimum wage. However, statistical significance is not lost as these additional workers are included. Indeed, the fact that standard errors are narrower in the specification with the more broadly-defined MW-affected share based on workers with wages up to the 25<sup>th</sup> percentile of the wage distribution

suggests that including these additional workers does not merely add noise to our results.

While we primarily focus on the minimum wage cost share as a proxy for wage rigidity due to its nature as a legislatively-set, legally-enforceable nominal rigidity, it is not the only identifiable source of nominal wage rigidities in the economy. Union wages and the wages paid to government workers are other examples of downwardly-rigid wages. For this reason, the cost share of workers covered by union contracts and the cost share of government workers were included as control variables in Figure 7 – to demonstrate that the MW-affected share is important in itself and not merely due to potential correlation with these alternative shares. However, these shares are of independent interest; the same intuition about nominal rigidities that applies to the MW-affected share also implies that higher values of these shares should also induce stronger efficacy of monetary policy. Figure 12 plots the results of running the baseline specification, albeit with the minimum wage cost share replaced with the cost share of workers covered by union contracts and the cost share of government workers. A similar result – though smaller in magnitude – can be observed in these contexts. In response to a 1 pp contractionary shock, a state at the 90th percentile of the union cost share will experience a decline employment that is approximately 2.1 percentage-points greater (at peak) than a state at the 10th percentile of the union cost share. In response to a 1 pp contractionary shock, a state at the 90th percentile of the government employment cost share will experience a decline employment that is approximately 0.8 percentage-points greater (at peak) than a state at the 10th percentile of the government employment cost share.

The QCEW also includes quarterly data on the number of establishments by state. Consequently, it is possible to examine whether there are analogous effects on business creation/destruction mirroring the effects on employment. To the extent that many businesses rely on minimum-wage labor as an important input, varying the cost of that input may have important impacts on business formation and closure. Figure 13 examines this conjecture, displaying the results of a quarterly version of the baseline specification in its left panel and the specification with state and time fixed-effects in its right panel. Contractionary monetary policy does appear to lead to disproportionate business shutdown in places with a higher share of minimum wage workers. In response to a 1 pp contractionary shock, a state at the 90<sup>th</sup> percentile of the MW-affected share

will experience a decline in the number of existing business establishments that is approximately 1.2 percentage-points greater (at peak) than a state at the 10<sup>th</sup> percentile of the MW-affected share. Notably, this suggests that while near-minimum-wage workers should still disproportionately account for employment contraction (growth) resulting from contractionary (expansionary) monetary policy shocks, the minimum-wage channel of monetary policy is still likely to be associated with contraction (growth) in non-minimum-wage workers. Even establishments that rely on minimum-wage employment as an important input also tend to employ some higher-skilled labor.

#### 4.2 Comparison of Effect Magnitudes: Minimum-Wage Channel vs. Overall

So far, we have provided substantial evidence that minimum wages – as one of the key sources of wage rigidity in the modern macroeconomy – are an important channel through which monetary policy is operationalized. We have shown evidence that monetary policy is significantly more effective where the MW-affected share is higher. An important question remains: what fraction of monetary policy's total effect is due to the minimum wage channel? In this section, we provide an answer to this question.

To do so, it is first necessary for us to obtain an estimate of the total effect of monetary policy. An obvious choice is to run a specification directly in line with that of Romer and Romer (2004) and other papers which have utilized narrative monetary shock series:<sup>6</sup>

$$\Delta E_{s,t} = \alpha + \sum_{i=0}^{48} \beta_i Shock_{t-i} + \sum_{i=1}^{48} \eta_i \Delta E_{s,t-i} + \varepsilon_{s,t} \quad (5)$$

That is, we regress the growth in employment on its lags and the Romer and Romer monetary shock series in order to measure the overall effect of monetary policy on employment. We weight the regression by employment in this context in order to get an accurate measure of the national-level effect.<sup>7</sup> We can then use our data on the (nationwide) minimum-wage labor share along with the results from our regression specifications in Section 5.1 to compute the component of monetary

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<sup>6</sup> Note also that this specification is equivalent to a version of our baseline specification with the minimum-wage share and interaction effects removed.

<sup>7</sup> Another approach is to simply run this regression after aggregating the variables to the national level; this produces a nearly identical result.

policy's effect originating from the minimum wage channel in a given year. Dividing the latter by the former, we obtain estimates of the fraction of monetary policy's total effect that is due to the minimum wage channel.

There is an issue with using this specification for the purpose of measuring the fraction of monetary policy's total effect that is due to the minimum wage channel. Estimates resulting from this regression treat the effect size of monetary policy as a fixed object, one which cannot change over time. In reality, there is some evidence that monetary policy has been less effective since the 1990s. Nakamura and Steinsson (2018) find evidence of a relatively flat Phillips curve over this period. Similarly, Mavroeidis, Plagborg-Møller, and Stock (2014) summarize later papers as typically finding statistically insignificant estimates of the Phillips curve slope; further, they mention how slope estimates in the well-known analysis of Galí and Gertler (1999) fall to insignificance when their sample is extended to include later data. And, indeed, we find the same. Figure 14 displays the results of running the above specification over two non-overlapping samples: 1975-1990 and 1990-2007. As can be seen, monetary policy has a strongly significant effect on employment in the former time period and no discernible effect in the latter. Consequently we focus on the former period.<sup>8</sup> We call 1975-1990 the “Volcker era,” despite Volcker’s tenure extending from 1979 to 1987, since our regressions use four years of lags. Thus, our regression sample contains outcomes starting in 1979, with lagged regressors extending back to 1975, and concludes with 1990 outcomes using regressors extending back to 1987.

As seen in Figure 14, the peak effect of a 1 percentage-point federal-funds rate shock during the 1975-1990 period is a 4 percentage-point reduction in employment. As seen in Figures 2 and 4, the peak interaction effect of a federal funds rate shock with the minimum-wage labor share is approximately -0.6 in the baseline specification and as small as -0.15 in the specification with state and time fixed-effects. Our preferred specification is the one depicted in Figure 5, which interacts the Romer and Romer shock series with state fixed-effects and, separately, with a time trend variable, thereby eliminating persistent differences across states in monetary policy efficacy and

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<sup>8</sup> Because there appears to be no statistically-significant effect of monetary policy in the latter period, attempting to obtain an estimate of the share of the effect driven by the minimum wage channel would constitute dividing by a (statistical) zero.

national differences over time in monetary policy efficacy as potential confounds. In this specification, the peak interaction effect is  $-0.35$ . Since the average minimum-wage labor share over this period is 4.4 percent, this implies that the minimum-wage channel of monetary policy is responsible for a reduction in employment of 1.54 percentage points in response to a 1pp federal-funds rate shock, depending on the specification. That is, over the 1975-1990 period, our estimates suggest that the minimum-wage channel of monetary policy is responsible for approximately 39% of monetary policy's total effect, a non-trivial share.

This finding also suggests a further point. As discussed in a handbook chapter by Ramey (2016), much recent work on the effects of monetary policy has found smaller effect sizes than older work studying earlier periods. This is consistent with our preceding result. And because we find that a substantial fraction of monetary policy efficacy was due to the minimum wage channel, our evidence suggests that declining real minimum wages – which induce a declining minimum wage labor share – are one important factor behind the reduced efficacy of monetary policy over time.

## 5 Model

Since minimum wage workers make up a relatively small fraction of employment, how large should the effects of our channel of monetary policy be? Further, how much heterogeneity across states should monetary policy generate through our channel? We address these points formally in the model, which provides quantitative estimates of how large an effect monetary policy should generate through the minimum wage channel alone. Throughout this section, the “MW-affected share” in a sector refers to the total payroll of workers affected by the minimum wage in that sector divided by total cost in that sector.<sup>9</sup> Monetary policy must enter our model in an empirically calibrated way if we hope to match our empirical results; as is well-known and documented in Coibion et al. (2012), the estimated effects of monetary policy differ substantially with whether one's empirical analysis uses VAR or Romer and Romer (2004) shocks. Thus, we will set up a rule for how interest rate movements affect nominal GDP that can be empirically calibrated to the

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<sup>9</sup> We provide full details on how we compute these shares from CPS and BEA data in the data section.

VAR and Romer and Romer (2004) results.

The share of minimum wage workers is correlated with numerous other variables that may lead to differential effects of monetary policy across regions and time, so the model also provides an opportunity for us to generate the unique implications of our channel relative to competitor explanations. The model focuses on one confound in particular: the share of minimum wage workers may be high in regions where a higher share of households is credit constrained. In this case, any effects we attribute to monetary policy relaxing the minimum wage may be due instead to monetary policy alleviating or exacerbating credit constraints. More generally, higher MW-affected share regions may be regions where there is a higher marginal propensity to consume (MPC). We address this concern by analyzing tradable and non-tradable sectors in the model. As discussed in our literature review, shocks going through the MPC channel should lead to larger effects on employment in non-tradable sectors than on employment in tradable sectors. The minimum wage channel should lead, in contrast, to larger effects in tradable employment, a result we will confirm in the model and in our empirical analysis.

## 5.1 Households and Money Supply

The household in each state  $s = 1, \dots, S$  purchases tradables and non-tradables to produce a commodity, which can be invested or consumed. So, though there are two types of goods available in each state, there is only one type of capital, produced out of both non-tradables and tradables, in each state. Households cost minimize over tradable and non-tradable inputs when producing the commodity, yielding the expenditure function

$$E_s(P_t^T, P_{s,t}^{NT}, Y_{s,t}) = \min_{Y_{s,t}^T, Y_{s,t}^{NT}} P_t^T Y_{s,t}^T + P_{s,t}^{NT} Y_{s,t}^{NT} \quad s. t. \quad F_s(Y_{s,t}^T, Y_{s,t}^{NT}) = Y_{s,t}$$

Where  $F_s$  is assumed to be constant returns to scale for each  $s$ . The commodity price is an ideal price index given by

$$P_{s,t} = E_s(P_t^T, P_{s,t}^{NT}, 1).$$

We assume the steady state elasticity of substitution between tradables and non-tradables – denoted by  $\sigma_{NT,T}$  – does not vary by state, an assumption driven by the absence of measurements of the parameter at this level of granularity. The elasticity  $\sigma_{NT,T}$  is a parameter if  $F_s$  has a constant

elasticity of substitution form with elasticity of substitution  $\sigma_{NT,T}$ ; otherwise, it is a steady state object.

We use lower-case variables to refer to the natural log of their upper-case variants, e.g.  $k_{s,t} = \ln K_{s,t}$ . A hat denotes a variable's deviation from its steady state value (the existence of which we show in the appendix). Finally, a boldface variable refers to a vector or matrix.

The cost minimization conditions can be log-linearized as

$$\begin{aligned}\hat{\mathbf{y}}_t^{NT} - \hat{\mathbf{y}}_t^T &\approx -\sigma_{NT,T}(\hat{\mathbf{p}}_t^{NT} - \mathbf{1} \hat{p}_t^T) \\ \hat{\mathbf{y}}_t &\approx \boldsymbol{\eta}_{NT} \hat{\mathbf{y}}_t^{NT} + \boldsymbol{\eta}_T \hat{\mathbf{y}}_t^T.\end{aligned}$$

Where endogenous boldface variables are  $S \times 1$  vectors. The steady state cost shares  $\boldsymbol{\eta}_{NT}$  and  $\boldsymbol{\eta}_T$  are diagonal  $S \times S$  matrices giving the share of a state's GDP in the state's non-tradable and tradable sectors, respectively. The first equation captures the standard notion that relative input use varies with relative prices mediated by the elasticity of substitution (indeed, it is sometimes used to define the elasticity of substitution), and the second decomposes changes in commodity demand into changes in demand for the components producing the commodity. Both equations follow from the cost minimization solutions, which for the tradable input takes the form

$$Y_t^T = Y_t^T(P_t^T, P_{s,t}^{NT}, Y_{s,t}),$$

so that

$$\begin{aligned}\hat{y}_t^T &= \frac{\partial \ln Y_t^T}{\partial \ln P_t^T} \hat{p}_t^T + \frac{\partial \ln Y_t^T}{\partial \ln P_{s,t}^{NT}} \hat{p}_{s,t}^{NT} + \hat{y}_{s,t} = \frac{\partial \ln Y_t^T}{\partial \ln P_{s,t}^{NT}} (\hat{p}_{s,t}^{NT} - \hat{p}_t^T) + \hat{y}_{s,t} \\ &= \eta_s^{NT} \sigma_{NT,T} (\hat{p}_{s,t}^{NT} - \hat{p}_t^T) + \hat{y}_{s,t},\end{aligned}$$

Where the first equality used constant returns to scale, the second used homogeneity of degree 0 of the conditional factor demand function in prices, and the final equality used the definition of the (symmetric) elasticity of substitution in terms of a cost share and the demand elasticity (sometimes called a Hicks-Marshall law of derived demand).

The price index defined as above can be log-linearized as

$$\hat{\mathbf{p}}_t \approx \boldsymbol{\eta}_{NT} \hat{\mathbf{p}}_t^{NT} + \boldsymbol{\eta}_T \mathbf{1} \hat{p}_t^T,$$

and this is just a straightforward application of the envelope theorem on the cost function. We can now define the household's dynamic problem at the commodity level, abstracting from tradables

and non-tradables. First, we turn to how monetary policy will operate in our model.

The national money supply  $M_t$ , which is the quantity of money times its velocity, is the numeraire for the economy. It satisfies

$$M_t = \sum_s P_{s,t}^{NT} Y_{s,t}^{NT} + P_t^T Y_{s,t}^T = \sum_s E_s(P_{s,t}^{NT}, P_t^T, Y_{s,t}) = \sum_s P_{s,t} Y_{s,t} \equiv GDP_t$$

Where  $GDP_t$  is nominal gross domestic product. Under a truly unexpected shock to the interest rate in period  $t$ , we will use the convention that a 1 log point increase in the nominal interest rate relative to steady state leads to a  $\beta_m^i$  reduction in the money supply relative to steady state. The money process follows  $\hat{m}_t = -\rho_m \hat{m}_t$ , with  $0 < \rho_m < 1$ . This means the money supply ultimately returns to its steady state path, though we will be mainly be interested in the case where  $\rho_m$  is very small, so that the interest rate shock has a persistent effect on the money supply. We can calibrate  $\beta_m^i$  using Romer and Romer's (2004) estimates for how shocks to the federal funds rate affect nominal GDP, or we could take this number from VAR estimates, depending on which shocks we are using in the reduced form empirical results we want the model to match.

The household supplies capital  $K$  and two types of labor to the production side of the economy. Labor type  $L$  is subject to a binding wage floor  $\bar{W}$ , and labor type  $H$  is paid an endogenous wage  $W$ . We will refer to labor type  $L$  as “less skilled” or “minimum wage” labor and to labor type  $H$  as “skilled” or “endogenous wage” labor. The model contains no uncertainty. The results relevant for our empirics do not meaningfully change if we permit two households in each state, (1) a less skilled household that consumes hand-to-mouth and faces a binding minimum wage and (2) a skilled household that performs all investment in the state and has an endogenous wage. The budget constraint is

$$P_{s,t} Y_{s,t} = P_{s,t} (C_{s,t} + I_{s,t}) = \bar{W}_{s,t} L_{s,t} + W_{s,t} H_{s,t} + R_{s,t} K_{s,t},$$

Where the first equality links the budget constraint to the previously described cost minimization component of the household problem. The law of motion for capital is

$$\dot{K}_{s,t} = I_{s,t} - \delta K_{s,t},$$

Where a dot refers to the time derivative of a variable, and  $\delta$  is the depreciation rate of capital. The utility function is separable and does not vary by region or time:

$$U(C_{s,t}) - V(H_{s,t}) - V_L(L_{s,t}).$$

Our key assumption is that the wage floor is binding in each state. Thus, the household in each state would like to choose a higher value of  $L_{s,t}$  than the state can support. In the maximization problem, the household therefore takes  $L_{s,t}$  as given. The minimum wage will be set by the government, rather than by supply and demand, and the quantity of minimum wage labor will be determined by firm demand only, not household supply. This is a simple application of the disequilibrium framework of Barro and Grossman (1971). Changes in the interest rate will affect endogenous nominal variables but not the exogenous minimum wage, leading to non-neutrality of monetary policy.

We use the budget constraint to unconstrain the maximization problem, which we write as

$$\max_{\{\dot{K}_{s,t}\}_{t=0}^{\infty}, \{H_{s,t}\}_{t=0}^{\infty}} \int_0^{\infty} e^{-\rho t} \left( U \left( \frac{\bar{W}_{s,t}}{P_{s,t}} L_{s,t} + \frac{W_{s,t}}{P_{s,t}} H_{s,t} + \left( \frac{R_{s,t}}{P_{s,t}} - \delta \right) K_{s,t} - \dot{K}_{s,t} \right) - V(H_{s,t}) \right) dt$$

Where the initial capital stock is given in each state. Though the utility function does not vary by state, the inverse elasticity of intertemporal substitution, denoted  $\gamma$ , and the Frisch-elasticity of labor supply, denoted  $\epsilon$ , may still vary by state, since they depend on the level of consumption and skilled labor supply, respectively. We assume they do not vary by state, which could easily be micro-founded using CRRA forms for consumption utility and skilled labor disutility. Optimization yields a standard intratemporal and intertemporal Euler equation in each state, which we can log-linearize, respectively, as

$$-\gamma \hat{c}_t + \hat{w}_t - \hat{p}_t \approx \frac{1}{\epsilon} \hat{h}_t$$

$$\dot{c}_t \approx \frac{\rho + \delta}{\gamma} (\hat{r}_t - \hat{p}_t)$$

All of the boldface objects are endogenous  $S \times 1$  vectors. The first equation governs the relative choice of consumption and skilled hours as a function of the real skilled wage, while the second governs the growth rate of consumption as a function of the real rate of return on capital. We note again that minimum wage hours do not appear in these equations. The household takes those hours as given in the maximization problem, and they are separable from consumption and skilled hours

in the utility function.

## 5.2 Non-tradable Sector

There is a firm in each state which produces non-tradables for use in that state. The sector first solves the cost minimization problem

$$E_s^{NT}(R_{s,t}, W_{s,t}, \bar{W}_{s,t}, Y_{s,t}^{NT}) \\ = \min_{K_{s,t}^{NT}, H_{s,t}^{NT}, L_{s,t}^{NT}} R_{s,t} K_{s,t}^{NT} + W_{s,t} H_{s,t}^{NT} + \bar{W}_{s,t} L_{s,t}^{NT} \quad s.t. \quad F_s^{NT}(K_{s,t}^{NT}, H_{s,t}^{NT}, L_{s,t}^{NT}) = Y_{s,t}^{NT}.$$

Where  $F_s^{NT}$  is assumed to be constant returns to scale for each  $s$ . Profit maximization then yields non-tradable prices in each state,

$$P_{s,t}^{NT} = E_s^{NT}(R_{s,t}, W_{s,t}, \bar{W}_{s,t}, 1).$$

We assume the elasticities of substitution in steady state, denoted by  $\sigma_{HL}^{NT}$ ,  $\sigma_{LK}^{NT}$ , and  $\sigma_{HK}^{NT}$ , do not vary by state. The non-tradable firms' cost minimization conditions can be log-linearized (following the same procedure outlined in the household section) and stacked as

$$\hat{\mathbf{i}}_t^{NT} \approx \sigma_{HL}^{NT} \boldsymbol{\eta}_H^{NT} (\hat{\mathbf{w}}_t - \widehat{\mathbf{w}}_t) + \sigma_{LK}^{NT} \boldsymbol{\eta}_K^{NT} (\hat{\mathbf{r}}_t - \widehat{\mathbf{w}}_t) + \hat{\mathbf{y}}_t^{NT} \\ \hat{\mathbf{h}}_t^{NT} \approx \sigma_{HL}^{NT} \boldsymbol{\eta}_L^{NT} (\widehat{\mathbf{w}}_t - \hat{\mathbf{w}}_t) + \sigma_{HK}^{NT} \boldsymbol{\eta}_K^{NT} (\hat{\mathbf{r}}_t - \hat{\mathbf{w}}_t) + \hat{\mathbf{y}}_t^{NT} \\ \hat{\mathbf{k}}_t^{NT} \approx \sigma_{LK}^{NT} \boldsymbol{\eta}_L^{NT} (\widehat{\mathbf{w}}_t - \hat{\mathbf{r}}_t) + \sigma_{HK}^{NT} \boldsymbol{\eta}_H^{NT} (\widehat{\mathbf{w}}_t - \hat{\mathbf{r}}_t) + \hat{\mathbf{y}}_t^{NT}$$

The endogenous variables and minimum wage variable are again  $S \times 1$  vectors, and the  $\boldsymbol{\eta}_i^{NT}$  are diagonal  $S \times S$  matrices with entries given by the cost share of input  $i$  in non-tradable production in the relevant state.

Note that these equations are the standard Slutsky equations for the firm. The first two terms denote substitution effects, and the final term denotes the scale effect. If production were Leontief, then each  $\sigma$  would be 0, and we would be left only with the scale effect.

## 5.3 Tradable Sector

There is one national tradable firm, which produces in all states. We find this setup more realistic than permitting distinct tradable sectors in each state that produce a homogeneous output, since we will be able to allow parsimoniously for differences in state-level tradable output. The

sector operates by producing a commodity in each state and then combining these commodities to produce final tradable output. In the first stage, it cost minimizes over production in each state:

$$\begin{aligned} E_s^T(R_{s,t}, W_{s,t}, \bar{W}_{s,t}, Y_{s,t}^T) \\ = \min_{K_{s,t}^T, H_{s,t}^T, L_{s,t}^T} R_{s,t}K_{s,t}^T + W_{s,t}H_{s,t}^T + \bar{W}_{s,t}L_{s,t}^T \quad s. t. \quad F_s^T(K_{s,t}^T, H_{s,t}^T, L_{s,t}^T) = Y_{s,t}^T. \end{aligned}$$

The production function  $F_s^T$  exhibits constant returns to scale in each  $s$ . This generates an ideal price index for the price of the state commodities required in production of the national tradable:

$$P_{s,t}^T = E_s^T(R_{s,t}, W_{s,t}, \bar{W}_{s,t}, 1).$$

The firm then minimizes national-level costs:

$$E^T(P_{1,t}^T, \dots, P_{S,t}^T, Y_t^T) = \min_{\{Y_{s,t}^T\}_{s=1}^S} \sum_{s=1}^S P_{s,t}^T Y_{s,t}^T \quad s. t. \quad F^T(Y_{1,t}^T, \dots, Y_{S,t}^T) = Y_t^T$$

Finally, profit maximization yields the national tradable price,

$$P_t^T = E^T(P_{1,t}^T, \dots, P_{S,t}^T, 1).$$

We can combine and stack the cost minimization conditions. This is a more complicated procedure than in the non-tradable sector, but we will give intuition after defining the relevant objects. Assume the elasticities of substitution between the state commodities in producing the national tradable are all equal and given by  $\sigma_s$ . This could be micro-founded by assuming  $F^T$  has a CES form with a single elasticity of substitution,  $\sigma_s$ . Note that, had we modeled the tradable sector with distinct tradable sectors in each state that produce homogeneous output, we would implicitly be letting  $\sigma_s \rightarrow \infty$ , the case of perfect substitutes. Further, define  $\eta_s^T$  as the cost share of the state  $s$  commodity in producing the national tradable, measurable by tradable GDP in that state divided by tradable GDP in the U.S. Denote the  $S \times S$  diagonal matrix of these shares by  $\boldsymbol{\eta}^T$ . The diagonal  $S \times S$  matrices of cost shares  $\boldsymbol{\eta}_i^T$  are analogous to those defined in the non-tradable sector: their entries are given by the cost share of input  $i$  in tradable production in the relevant state. We define the *diag* operator, which retrieves the diagonal entries of a matrix as a column vector, and the  $*$  operator, which performs elementwise multiplication between two objects of the same dimension.

First, we define the scale effect

$$\mathbf{s}_t = \mathbf{S} * (\boldsymbol{\eta}_L^{T*} \widehat{\mathbf{w}}_t + \boldsymbol{\eta}_H^{T*} \widehat{\mathbf{w}}_t + \boldsymbol{\eta}_K^{T*} \widehat{\mathbf{r}}_t) + \mathbf{1} \hat{y}_t^T,$$

where

$$\mathbf{S} = \sigma_s \begin{pmatrix} -\frac{1 - \eta_1^T}{\eta_1^T} & 1 & \dots & 1 \\ 1 & -\frac{1 - \eta_2^T}{\eta_2^T} & \dots & 1 \\ \vdots & 1 & \ddots & \vdots \\ 1 & 1 & \dots & -\frac{1 - \eta_S^T}{\eta_S^T} \end{pmatrix}, \quad \boldsymbol{\eta}_i^{T*} = \begin{pmatrix} \text{diag}(\boldsymbol{\eta}^T \boldsymbol{\eta}_i^T)' \\ \text{diag}(\boldsymbol{\eta}^T \boldsymbol{\eta}_i^T)' \\ \vdots \\ \text{diag}(\boldsymbol{\eta}^T \boldsymbol{\eta}_i^T)' \end{pmatrix}$$

And we note that elementwise multiplication of  $\mathbf{S}$  with the  $\boldsymbol{\eta}^*$  matrices must occur before the  $\boldsymbol{\eta}^*$  matrices multiply the factor prices. Then it follows that

$$\begin{aligned} \hat{\mathbf{l}}_t^T &\approx \sigma_{HL}^T \boldsymbol{\eta}_H^T (\widehat{\mathbf{w}}_t - \widehat{\mathbf{w}}_t) + \sigma_{LK}^T \boldsymbol{\eta}_K^T (\widehat{\mathbf{r}}_t - \widehat{\mathbf{w}}_t) + \mathbf{s}_t \\ \hat{\mathbf{h}}_t^T &\approx \sigma_{HL}^T \boldsymbol{\eta}_L^T (\widehat{\mathbf{w}}_t - \widehat{\mathbf{w}}_t) + \sigma_{HK}^T \boldsymbol{\eta}_K^T (\widehat{\mathbf{r}}_t - \widehat{\mathbf{w}}_t) + \mathbf{s}_t \\ \hat{\mathbf{k}}_t^T &\approx \sigma_{LK}^T \boldsymbol{\eta}_L^T (\widehat{\mathbf{w}}_t - \widehat{\mathbf{r}}_t) + \sigma_{HK}^T \boldsymbol{\eta}_H^T (\widehat{\mathbf{w}}_t - \widehat{\mathbf{r}}_t) + \mathbf{s}_t. \end{aligned}$$

As in the non-tradable sector, these equations all represent Slutsky equations for the tradable sector, except now there are two substitution effects. On the one hand, when a particular input in a state becomes more expensive, the first substitution effect, and the same one we saw in the non-tradable sector, drives substitution to the cheaper inputs in that state. The second substitution effect, contained in  $\mathbf{s}$ , drives the tradable firm to substitute away from the commodity in the state that has seen the factor price increase. Thus, for tradable production in the state where the factor price has increased, the substitution effect in  $\mathbf{s}$  is a scale effect, whereas for tradable production construed nationally, it is just another substitution effect away from the more expensive input, which in  $\mathbf{s}$  is the state-level commodity. Finally,  $\mathbf{s}$  also contains the standard output scale effect we also saw in the Slutsky equation for the non-tradable sector.

Note that if production at the state level and national level were both Leontief, all elasticities of substitution would be 0, and we would have  $\hat{\mathbf{l}}_t^T \approx \mathbf{1} \hat{y}_t^T$ ,  $\hat{\mathbf{h}}_t^T \approx \mathbf{1} \hat{y}_t^T$ , and  $\hat{\mathbf{k}}_t^T \approx \mathbf{1} \hat{y}_t^T$ . Note that this scale effect, unlike the scale effect in the non-tradable sector, is the same for all states. In this case, the model predicts no heterogeneity in log employment effects by state in the tradable sector. If production at the state level were Leontief, but we kept the general form for national production,

we would still have  $\hat{\mathbf{l}}_t^T \approx \mathbf{s}_t$  and be unable to simplify  $\mathbf{s}_t$  to just  $\mathbf{1} \hat{y}_t^T$  (and similarly for the other inputs).

For the shocks we consider in our model, increasing  $\sigma_s$  from arbitrarily close to 0 will have only a modest effect. This is because our national shocks will change all states' real minimum wages at the same time. Recall that  $\sigma_s$  provides a measurement of how the national firm's relative use of state commodities varies with the relative price of those commodities. When the minimum wage increases by the same amount in all states, relative commodity price changes are governed by differential cost shares of minimum wage workers across states. This generates less dramatic variation than the case where minimum wages do not increase in all states simultaneously, particularly when minimum wage cost shares are not too close to 0 and minimum wage changes are large.

#### 5.4 Equilibrium, Calibration, and Solution

Goods markets and labor markets clear. It is worth mentioning that labor market clearing in the less skilled labor market means that the quantity of less skilled labor taken as given by households in each state equals the combined tradable and non-tradable demand for less skilled labor in that state. We will solve a log-linearized version of the model and so will only be concerned with local versions of the transversality conditions.

We show in the appendix that a steady state in which all nominal variables grow at the same rate as the money supply exists. A key feature here is that nominal minimum wages in each state all grow at the same rate, the same rate as the money supply. Without this feature, the real minimum wage may change, leading to shifts in real variables.

We can simplify the log-linearized equations described above to a  $(3S + 1) \times (3S + 1)$  system of differential equations given by

$$\begin{pmatrix} \dot{\mathbf{c}}_t \\ \dot{\mathbf{k}}_t \\ \dot{\widehat{\mathbf{w}}}_t \\ \dot{\mathbf{m}}_t \end{pmatrix} \approx A \begin{pmatrix} \widehat{\mathbf{c}}_t \\ \widehat{\mathbf{k}}_t \\ \widehat{\widehat{\mathbf{w}}}_t \\ \widehat{\mathbf{m}}_t \end{pmatrix}.$$

It is not difficult to solve this system numerically. Now the matrix  $A$  must be calibrated. We use

standard parameter values where possible:

<i>Parameter</i>	<i>Description</i>	<i>Value</i>	<i>Notes</i>
$\gamma$	Risk aversion in steady state	2	Upper bound suggested by Chetty (2006)
$\rho$	Discount rate	.01	Quarterly
$\delta$	Depreciation rate	.025	Quarterly
$\epsilon$	Frisch elasticity of high skill labor supply in steady state	.4	Whalen and Reichling (2016)
$\sigma_{HL}$	High/low skill labor elasticity of substitution in steady state	1.41	Katz and Murphy (1992)
$\sigma_{HK}$	High skill labor/capital elasticity of substitution in steady state	.5	Oberfield and Raval (2020)
$\sigma_{LK}$	Low skill labor/capital elasticity of substitution in steady state	1.67	Krusell et al. (2000)

Note that we use the same elasticities of substitution in the table for both the tradable and non-tradable sectors. Our code can handle setting these separately. Now the elasticity  $\sigma_s$  is somewhat nonstandard. We use the Cobb-Douglas case,  $\sigma_s = 1$ , as our baseline. In the appendix, we show that the results do not change much when reducing  $\sigma_s$  to 0.1, a calibration close to Leontief that minimizes the ability of substitutions across state commodities in the tradable sector to drive the result that monetary policy has a larger effect on tradable employment than non-tradable employment through the minimum wage channel.

We also set  $\sigma_{NT,T}$ , the elasticity of substitution between tradables and non-tradables in

consumption, to a value consistent with the Cobb-Douglas case, or  $\sigma_{NT,T} = 1$ . This is used by Mian and Sufi (2014), which motivated our analysis of tradable and non-tradable employment in the first place.

The only way our calibration will differ with the time period we analyze is in the cost shares  $\eta$ . We show our calibrations for the minimum wage cost shares in Figure 15. These are computed as the MW-affected share of total payroll of wage and salaried workers in a given state, multiplied by the labor share in the state—more details are provided in the data section. It is clear in the figure that minimum wage cost shares in the tradable sector are typically lower than those in the non-tradable sector. The bulk of cost shares in both sectors are below 0.05 across states and time. We wish to highlight how much higher MW-affected shares are in the early part of the sample: MW-affected shares in the non-tradable sector commonly exceed 0.075, and occasionally the shares exceed the 0.125 range. Thus, in this period in certain states, more than 12.5% of production cost is affected by the price floor.

A key calibration is the magnitude of the shock to the effective money supply we feed into the model. We recall that  $M_t = GDP_t$ , which can be measured in the data as  $P_t Y_t$ . Romer and Romer (2004) studied the effects of shocks to the federal funds rate on prices and output, separately. Thus, the cumulative effects on a shock to  $\ln M_t$  can be measured as the sum of the cumulative effects on  $\ln P_t$  and  $\ln Y_t$ . The effect of a 1 percentage point increase in the federal funds rate, measured in this way, accumulates to a 4% decline in  $M$  over two years (a 4% decline in output and 0% decline in prices) and a 7% decline in  $M$  over four years (a 1% decline in output and a 6% decline in prices). To be conservative, since shocks other than those from Romer and Romer (2004) usually lead to smaller effects, and because we are not particularly interested in the impulse response functions resulting from this model, we will calibrate a 1 percentage point shock to the federal funds rate as a completely unanticipated and permanent 4% shock to the money supply. We will then compare our model's results to the reduced form empirics using the Romer and Romer (2004) shock series.

## 5.5 Model Results

We vary our cost share calibrations by year and include all 50 states in the model, so it is infeasible to present impulse response functions to summarize our results. Instead, for each year of calibration and each outcome variable of interest, we compute the impulse response functions over a 4-year horizon and take the largest magnitude effect achieved over that 4-year horizon for each state. We summarize these maximal results in each year using a boxplot. Figure 16 shows these graphs for various outcomes of interest in two panels.

We wish to highlight several outcomes of interest, particularly in the late 1970s and early 1980s. In these years, the minimum wage channel of monetary policy contributes an overall employment decline of about -1.25% in response to a 1 percentage point increase in the federal funds rate. This effect is driven by low wage employment, which declines by more than 6%. With some outlier exceptions, high skill employment actually increases slightly, by around 0.15%. These results are sensitive to the elasticities of substitution between inputs that we use to calibrate the model; a higher degree of complementarity can cause high-skill employment to fall as well. We have selected the elasticity calibrations above due to the strong correlation of our employment groups with education and the extensive literature estimating elasticities of substitution between capital and various education groups.

Returning to our results, we see that capital and consumption both fall somewhat, by -0.075% and -0.35%, respectively. Prices fall by less than 4%, the magnitude of the shock to the money supply, highlighting that our channel is fundamentally a supply shock. Put differently, contractionary monetary policy raises the real cost of production by increasing the nominal minimum wage relative to other prices, and so prices fall by less than what would be predicted under monetary neutrality. The heterogeneity in these effects over time is less than the heterogeneity in employment, however.

We also want to highlight how our results look much closer to monetary neutrality from the late 1980s onward. This is the period where the MW-affected share has become relatively small, meaning less economic cost is at the binding wage floor. We think the fact that a declining proportion of economic cost has been at the binding minimum wage over time contributes to findings that the effects of monetary policy may have fallen over time.

Panel 2 shows that median states experience smaller declines in tradable employment than non-tradable employment, but the minimum wage cost shares in tradable employment are often much smaller; it is also apparent that heterogeneity across states in tradable employment is much larger than in non-tradable employment. To highlight how our employment effects depend on the MW-affected share in a state, see Figure 17. Panel 1 shows how our overall employment effects increase in magnitude with the MW-affected share in the state. Panel 2 shows how our tradable employment effects increase in magnitude more quickly with the tradable MW-affected share than the non-tradable employment effects do with the non-tradable MW-affected share. As we discussed above, this is not driven by cross-state substitution in the tradable sector as much as it is by large capital shares, combined with the large elasticity of substitution between capital and labor affected by the wage floor.

## **5.6 Comparison of Effect Magnitudes: Model vs. Empirics**

At this point, we have presented many different empirical specifications, some yielding differing magnitudes for how the effect of monetary policy on employment varies with the MW-affected share. While most of these magnitudes are statistically indistinguishable, we wish to discuss them in-depth here and compare them to the magnitudes from the model.

The effect size from our model can be seen in Figure 17, Panel 1, which plots the peak employment effect over 4 years in a state as a function of the minimum-wage labor share of total costs in the state. Clearly, the effect on employment of a 1 percentage point increase in the federal funds rate is decreasing and convex in the MW-affected share. To estimate our model's effect size, we can regress the employment data in that plot on the MW-affected share and its square. This yields the following maximal effect: in response to a 1pp increase in the federal funds rate, a state with a 0.01 MW-affected share will experience effect a .24 percentage point larger change in employment than a hypothetical state with no MW-affected workers.

This maximal effect size is between two and three times smaller than the effect size from our baseline and Bartik control regressions, displayed in Figures 2 and 3, roughly two times larger than the effect size from our specifications with time fixed effects in Figure 4, and slightly smaller than

the effect size in our specification controlling for state- and linear-time-trend-varying effects of monetary policy in Figure 5. It is important to note that the model is measuring the change in employment *hours*, whereas our empirics using the QCEW data are analyzing employment *counts*. This would cause magnitudes to differ to the extent that minimum wage workers work fewer hours on average than higher wage workers.

## 6 The Mechanism: Testing Model Implications

Are the effects we uncover driven by near-minimum-wage workers themselves, as implied by our model? In order to confirm this, we make use of the longitudinal nature of the CPS-ORG, whereby households are observed twice, one year apart. This structure enables a more direct test of our mechanism, allowing us to confirm that our effects are indeed driven by near-minimum-wage workers themselves – as opposed to simply being driven by other workers in states that happen to have a high share of minimum wage workers. That is, we can examine the employment response of workers initially employed as minimum-wage workers to the employment response of those initially employed as non-minimum-wage workers:

$$\begin{aligned} \Delta E_{i,t} = & \alpha + \sum_{j=0}^3 \beta_j AccumShock_{t-j} + \gamma \mathbf{1}\{NearMinWage_i\} \\ & + \sum_{j=0}^3 \delta_j AccumShock_{t-j} \cdot \mathbf{1}\{NearMinWage_i\} + \varepsilon_{i,t}, \end{aligned} \quad (6)$$

where  $\Delta E_{i,t}$  denotes the change in employment status (relative to one year ago) of individual  $i$  at time  $t$ .  $AccumShock_t$  denotes the cumulated Romer-and-Romer monetary policy shocks over the past year, as of time  $t$ . Note that this is an individual-level regression equation. Because the CPS-ORG contains specific data on individual employment status and wages for each individual  $i$ , it is not necessary to perform this analysis at a more aggregated level. This specification omits lags of the left-hand-side variable because they cannot be constructed using longitudinal CPS data; a household is only observed twice.

We plot the results of this regression specification in the left panel of Figure 18. In response to a contractionary (expansionary) shock, workers who are initially near the minimum wage are significantly more (less) likely to become unemployed – relative to workers who are initially well

above the minimum wage. Specifically, these workers have a 1 percentage-point higher probability of becoming unemployed as a result of a 1 percentage-point Romer and Romer monetary policy shock. The right panel of Figure 18 plots the (nearly identical) results of a regression that adds state-by-time fixed-effects to Equation (6).

In order to further validate the mechanism of the effect, we decompose employment growth into employment growth in tradables and employment growth in non-tradables<sup>10</sup>. We run a version of the baseline specification that interacts the Romer and Romer shocks and their 48 lags with the tradable MW-affected share and, separately, a version of the baseline specification with the non-tradable MW-affected share as the interaction term. As we show in Figure 19, a higher MW-affected share significantly boosts monetary policy efficacy in both tradables and non-tradables. The effect is not driven by non-tradables. Indeed, if anything, the effect is stronger on *tradable* employment, precisely the opposite of what the MPC channel would suggest.

One remaining concern is that the effect on tradables is only larger (or the same size) as the effect on non-tradables because of business-stealing effects. That is, since tradables can more easily be produced in one state and then transported/sold in another state, a state with a low minimum-wage labor share may simply siphon business from states with high minimum-wage labor shares in response to a contractionary monetary policy shock. Because the minimum wage is (mostly) non-binding in the former state and binding in the latter, an increase in the real minimum wage reduces the relative cost of business in the former state. If this dynamic drives the effects we find on tradables, we might remain concerned that any overall, non-zero-sum stimulus effect is entirely a consequence of the non-tradable sector and therefore the MPC channel. To determine whether this is the case, we compare the effect of a monetary policy shock on tradable versus non-tradable employment *in a 0% minimum-wage share state*. As discussed above, business-stealing would imply positive effects of a contractionary shock on tradable employment relative to non-tradable employment in this setting. To this end, we utilize a within-state

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<sup>10</sup> We define “Tradables” as the Agriculture, Mining, Manufacturing, and Finance sectors. We define “Non-tradables” as the Construction, Transportation, Communications, Utilities, Retail Trade, Wholesale Trade, Services, and Public Administration sectors. There is no one authoritative definition of these two terms, and some classifications omit Finance from the Tradables category and/or omit Wholesale Trade and parts of Services from the Non-tradables category. Omitting some or all of these sectors from our classification does not materially change our results.

specification comparing employment growth in tradable versus non-tradable sectors within-state:

$$\begin{aligned} & \alpha + \sum_{i=0}^{48} \beta_i Shock_{t-i} \cdot \mathbf{1}\{Tradable\} \\ \Delta E_{s,k,t} = & + \sum_{i=0}^{48} \theta_i Shock_{t-i} \cdot MWShare_{s,k,t} + \gamma MWShare_{s,k,t} \cdot \mathbf{1}\{Tradable\} + \rho MWShare_{s,k,t} \quad (7) \\ & + \sum_{i=0}^{48} \delta_i Shock_{t-i} \cdot \mathbf{1}\{Tradable\} \cdot MWShare_{s,k,t} + \sum_{i=1}^{48} \eta_i \Delta E_{s,k,t-i} + \omega_{s,t} + \theta_{s,k} + \varepsilon_{s,t} \end{aligned}$$

where  $\omega_{s,t}$  denotes state-by-time fixed effects and  $\theta_{s,k}$  denotes state-by-tradability fixed-effects. Intuitively, this specification roughly corresponds to a triple-differences approach. Note that in this setting the coefficients  $\delta_i$  record the interaction effect of monetary policy shocks on tradable employment relative to non-tradable employment as a function of the minimum-wage labor share, and the “level effect” coefficients  $\beta_i$  measure the effect of monetary policy shocks on tradable employment relative to non-tradable employment independent of the MW-affected share.

As seen in Figure 20, there is no statistically-significant effect on tradable employment relative to non-tradable employment when the MW-affected share is 0%. This suggests that business-stealing effects are of minimal importance here and do not drive our results. This finding agrees with the results of our model, which suggest that stronger within-state input substitution drives larger effects in the tradable sector. Intuitively, there is a somewhat limited role for cross-state substitution when the shock changes all states’ real minimum wages simultaneously.

## 7 Conclusion

We observe that the standard theoretical and empirical understanding of monetary policy suggests that it should erode real minimum wages. Our model establishes this point formally, providing quantitative predictions about how differences in the minimum wage share across states and time generates heterogeneity in the effects of monetary policy. The model also predicts that our channel of monetary policy should lead to larger changes in low-wage and tradable employment compared to high-wage and non-tradable employment, respectively. To test empirically whether the minimum wage channel is indeed an important channel through which monetary policy operates, we turn to the data. Using QCEW data on employment growth, CPS and BEA data on the share of minimum wage workers by state, and Romer and Romer narrative monetary policy shocks, we find that, indeed, this channel is crucial. This result is robust to a

variety of different identification strategies and the inclusion of a variety of controls. The relationship also manifests itself using VAR shocks instead of Romer and Romer narrative shocks. It is present in the Canadian data as well. An individual-level specification comparing near-minimum-wage to higher-wage workers reveals that in response to a contractionary (expansionary) shock, workers who are initially near the minimum wage are significantly more (less) likely to become unemployed – relative to workers who are initially well above the minimum wage. And analysis of tradable and non-tradable employment shows that our channel of monetary policy goes more strongly through the tradable sector, as predicted by our model. Our evidence suggests that this rigid-minimum-wage channel of monetary policy accounts for approximately 39% of monetary policy’s total effect in our preferred specification.

Taken as a whole, these findings reveal that minimum wages are an overlooked but important factor in determining the efficacy of monetary policy. Our results imply that monetary policy is less effective in the absence of minimum wages. This suggests that, on the one hand, higher minimum wages function as an additional dimension of “policy space” that boosts the ability of monetary policy to stabilize the economy as desired by policymakers. On the other hand, the Fed – typically thought of as an independent agency – is actually relaxing legislated policies and thereby dependent on the political process. In any case, these findings suggest that the interaction effects of monetary policy and wage rigidity is an understudied topic with room for future exploration.

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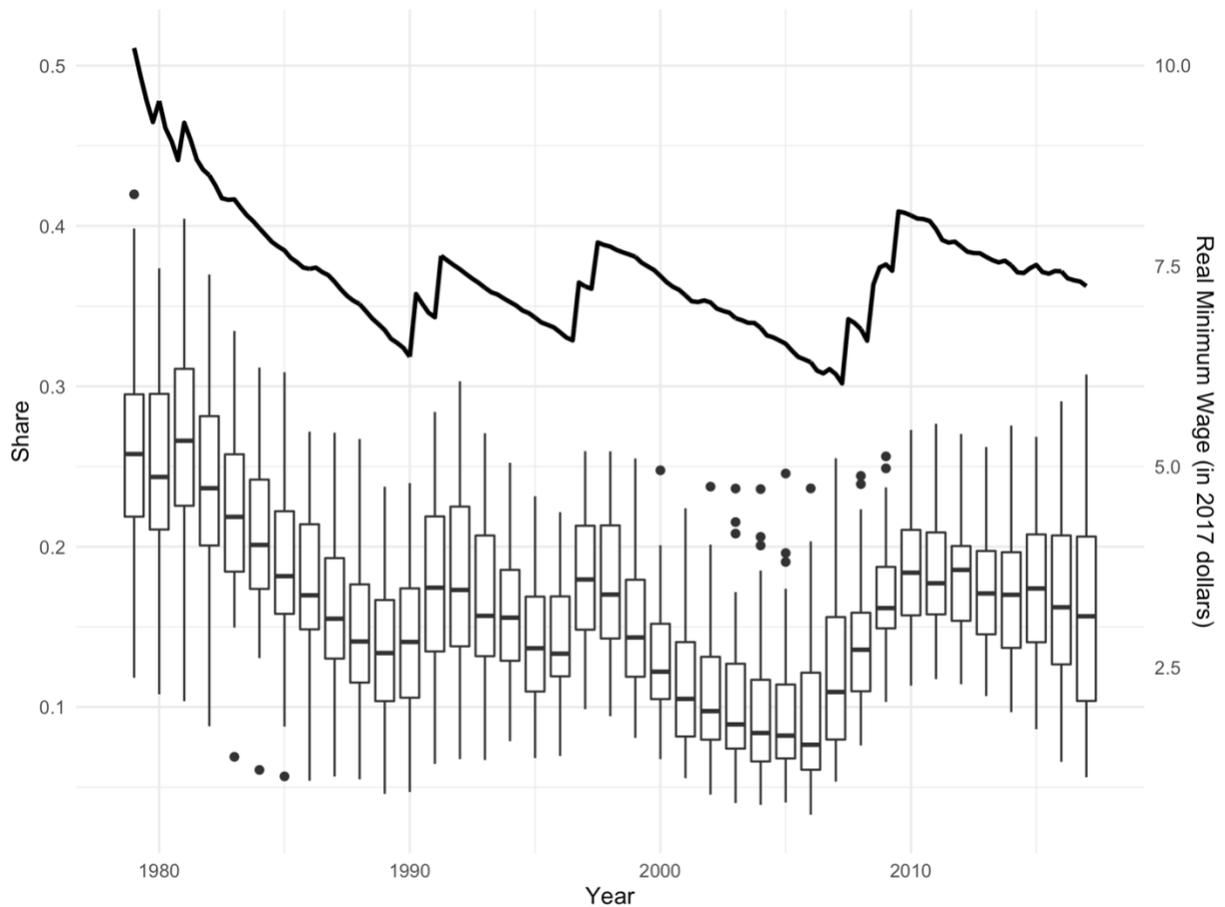
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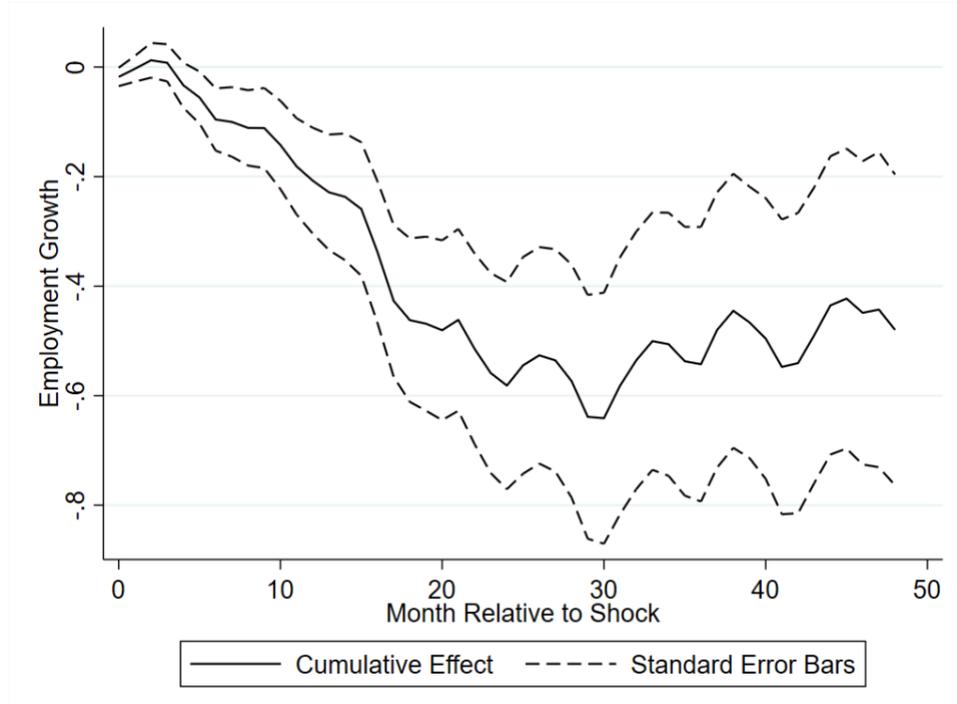
# Figures

Figure 1: Real Minimum Wages and MW-affected Employment Shares



Note: The solid line and right axis show the real federal minimum wage in 2017 dollars. The boxplot and left axis show heterogeneity across states in the share of employment affected by the minimum wage.

Figure 2: Baseline Specification – Effects of a 1pp MP Shock on Employment Interaction Coefficient between Shock and MW-affected share



Note: This figure plots the cumulated coefficients on the interaction effect between the cost share of minimum wage employment and the Romer and Romer monetary policy shocks. In other words, it displays the differential effect of monetary policy in places with a higher share of minimum wage employment.

Figure 3: Baseline with Bartik Control – Effects of a 1pp MP Shock on Employment Interaction Coefficient between Shock and MW-affected share

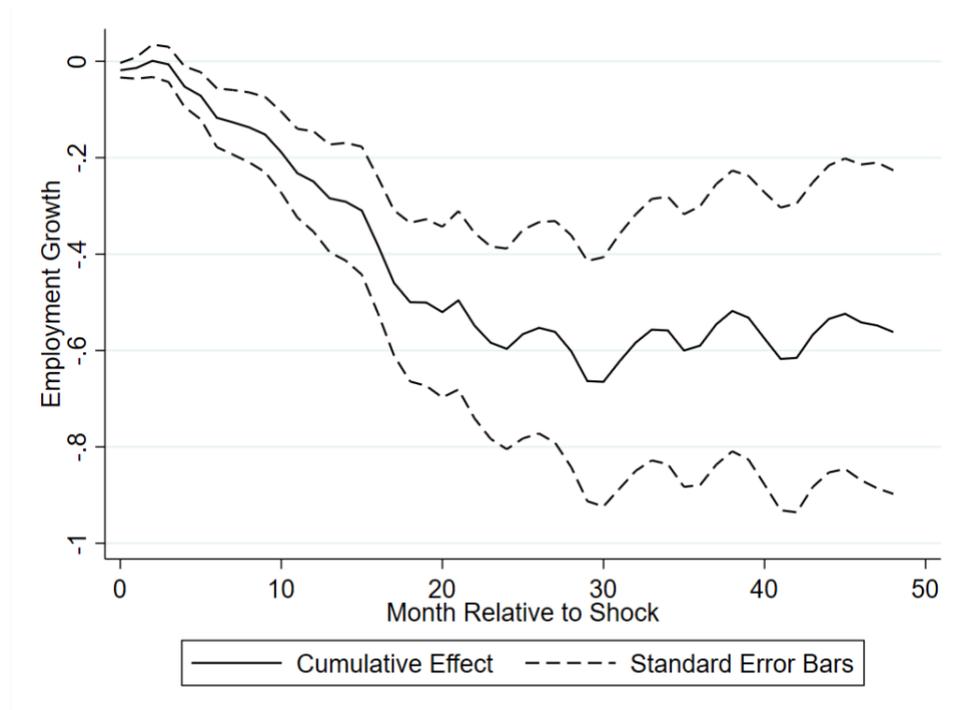
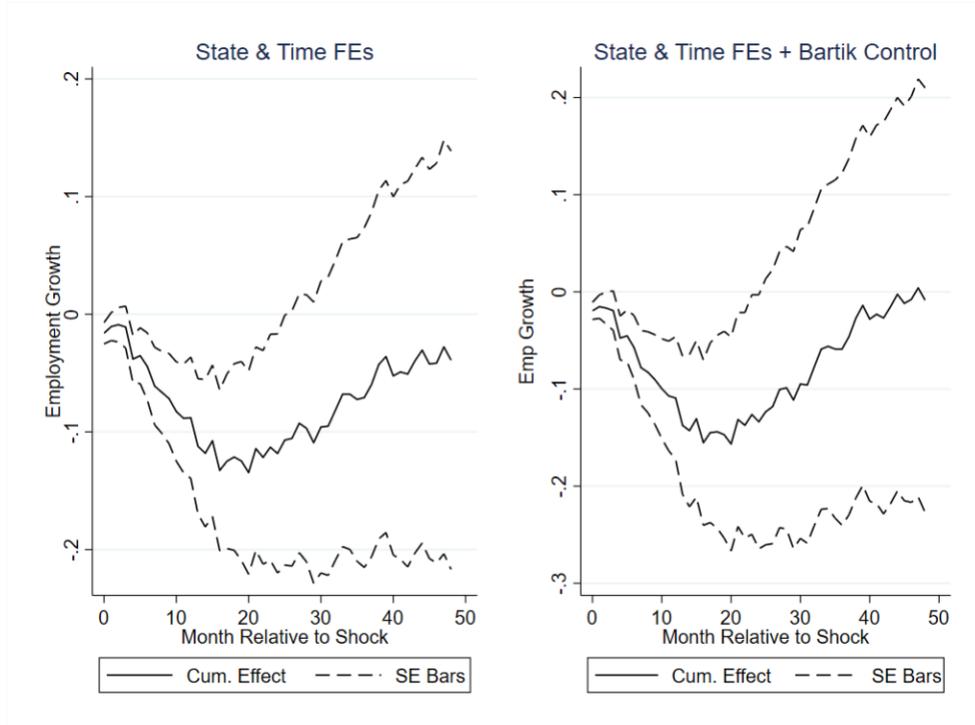
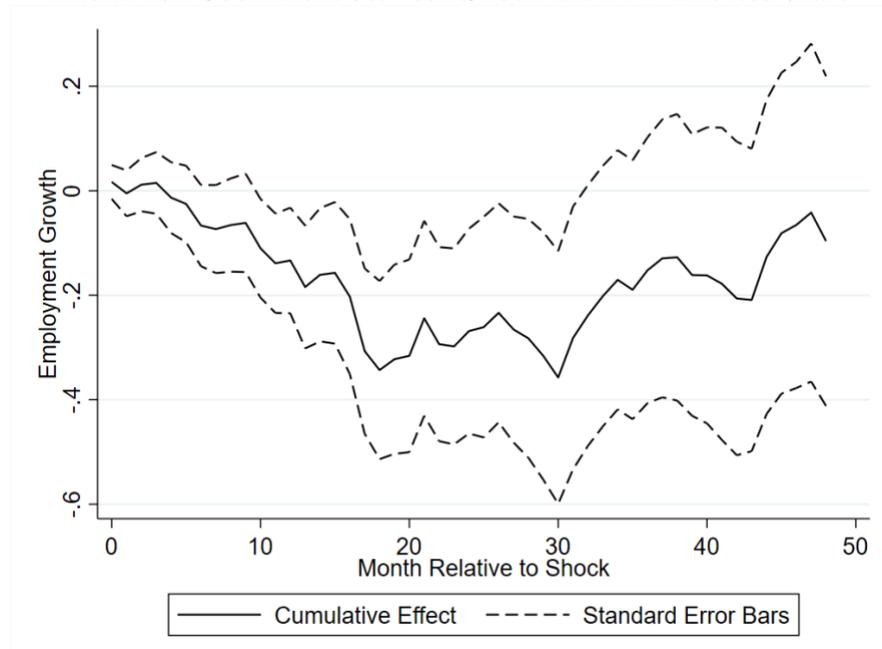


Figure 4: State and Time FE Specification – Effects of a 1pp MP Shock on Employment Interaction Coefficient between Shock and MW-affected share



Note: These specifications add state and time fixed-effects to the baseline specification and the specification with a Bartik control variable.

Figure 5: State-Specific & Time-Varying MP Efficacy – Effects of a 1pp MP Shock on Emp. Interaction Coefficient between Shock and MW-affected share



Note: This specification interacts the monetary policy shock series with state fixed-effects and, separately, with a time trend variable to control for the possibility that the efficacy of monetary policy varies by state for reasons unrelated to the MW-affected share and that the (national) efficacy of monetary policy is varying over time.

Figure 6: Local Projection Approach – Effects of a 1pp MP Shock on Emp.  
Interaction Coefficient between Shock and MW-affected share

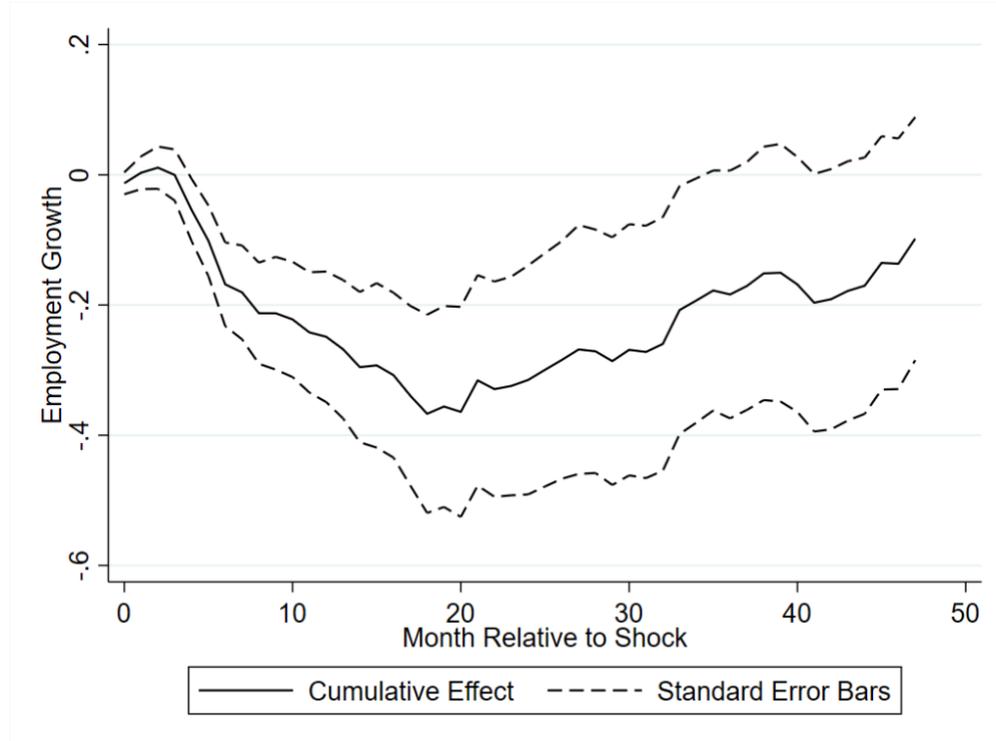
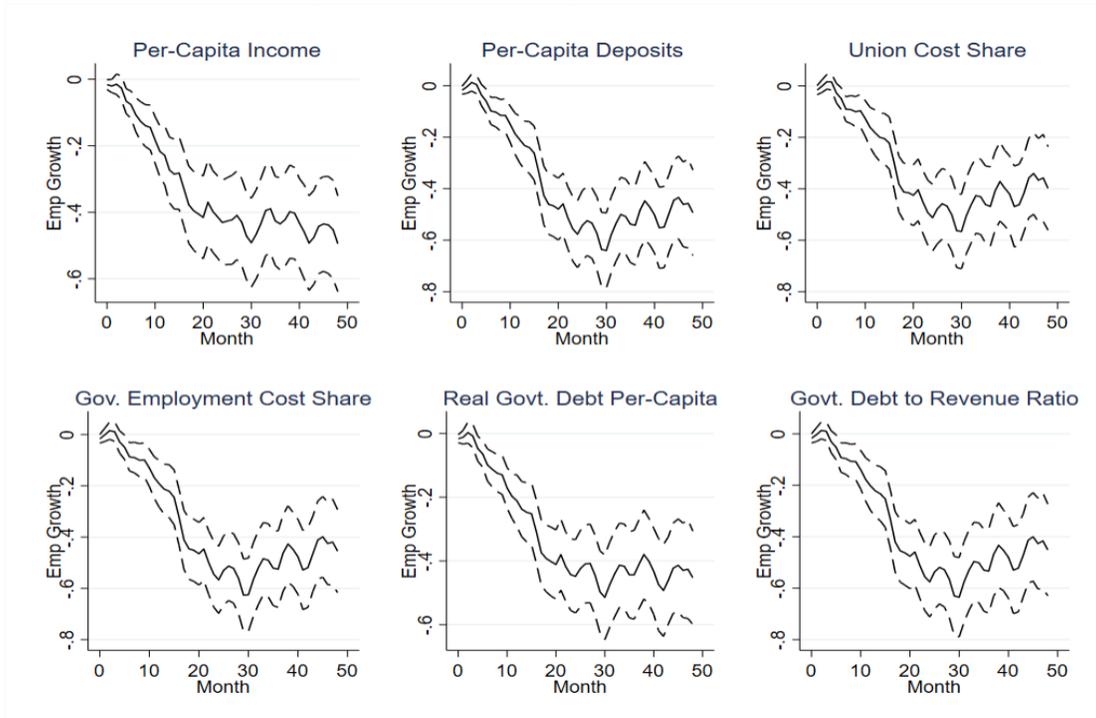
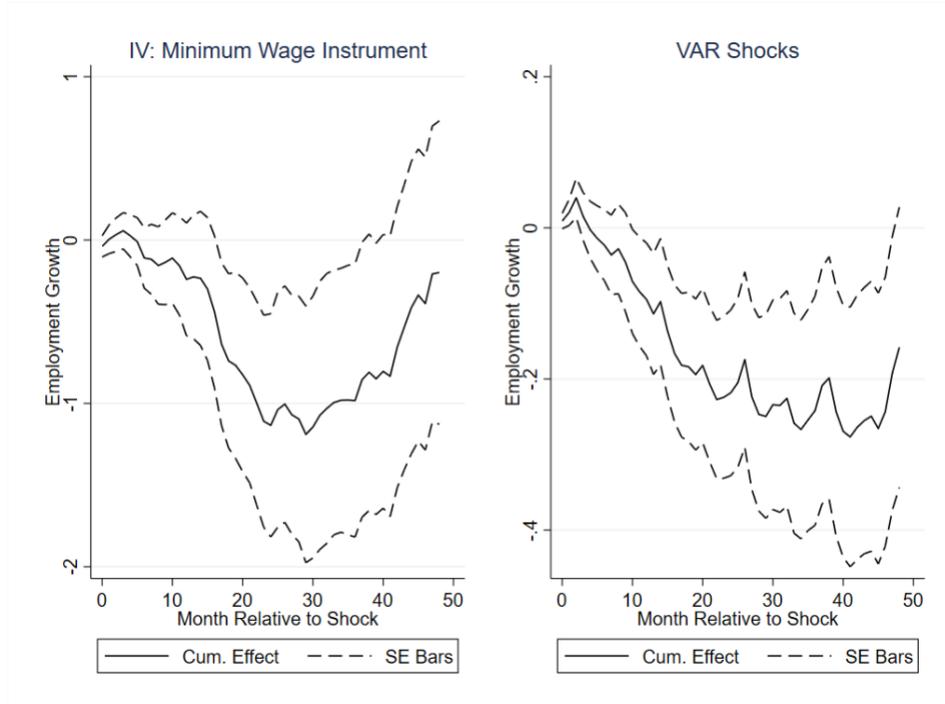


Figure 7: Baseline with Various Controls – Effects of a 1pp MP Shock on Employment  
Interaction Coefficient between Shock and MW-affected share



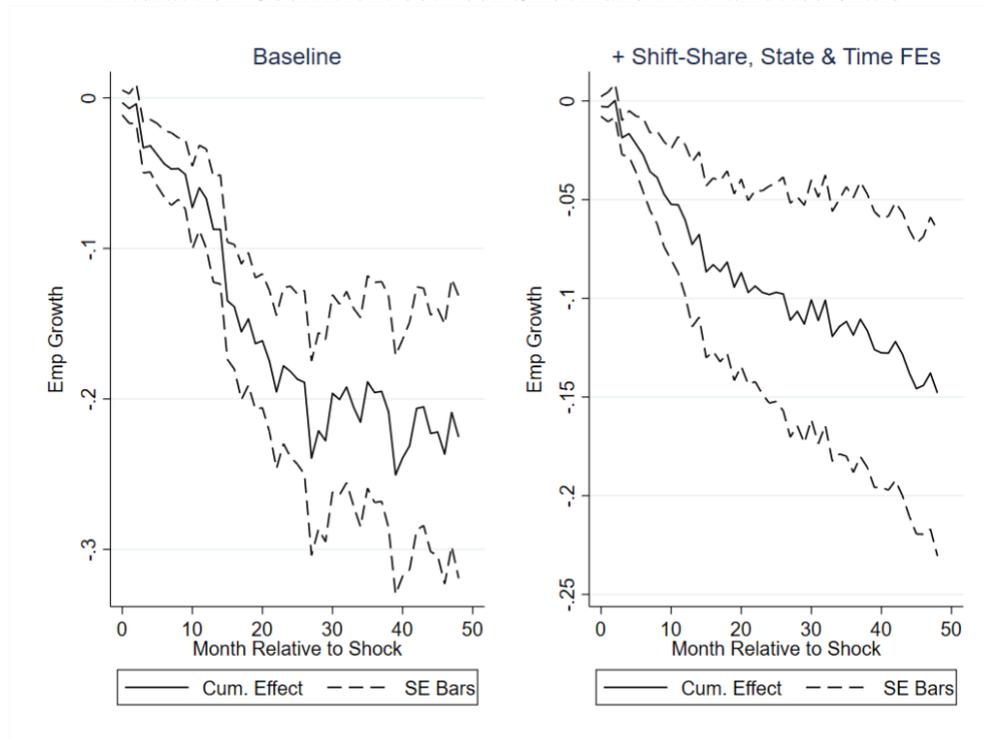
Note: Each specification depicted here controls for the titular variable *interacted* with the Romer and Romer monetary policy shock series in order to ensure the main result is not driven by omission of these variables that are potentially correlated with the MW-affected share.

Figure 8: IV & VAR Shock Specifications – Effects of a 1pp MP Shock on Employment Interaction Coefficient between Shock and MW-affected share



Note: The IV specification instruments the state MW-affected share with the level of the minimum wage in the state. The VAR shocks specification replaces Romer and Romer shocks with VAR shocks from Coibion (2012).

Figure 9: FFR Specifications – Effects of a 1pp MP Shock on Employment Interaction Coefficient between Shock and MW-affected share



Note: These specifications replace Romer and Romer shocks with raw Federal Funds Rate shocks.

Figure 10: Canada Specification – Effects of a 1pp MP Shock on Canadian Employment  
Interaction Coefficient between Shock and MW-affected share

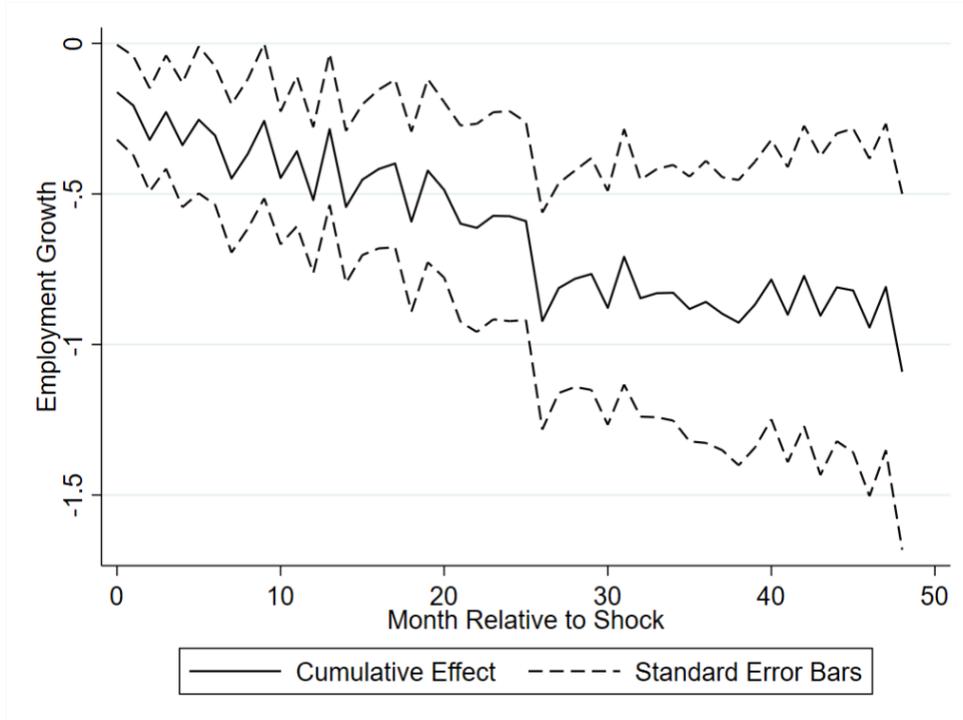
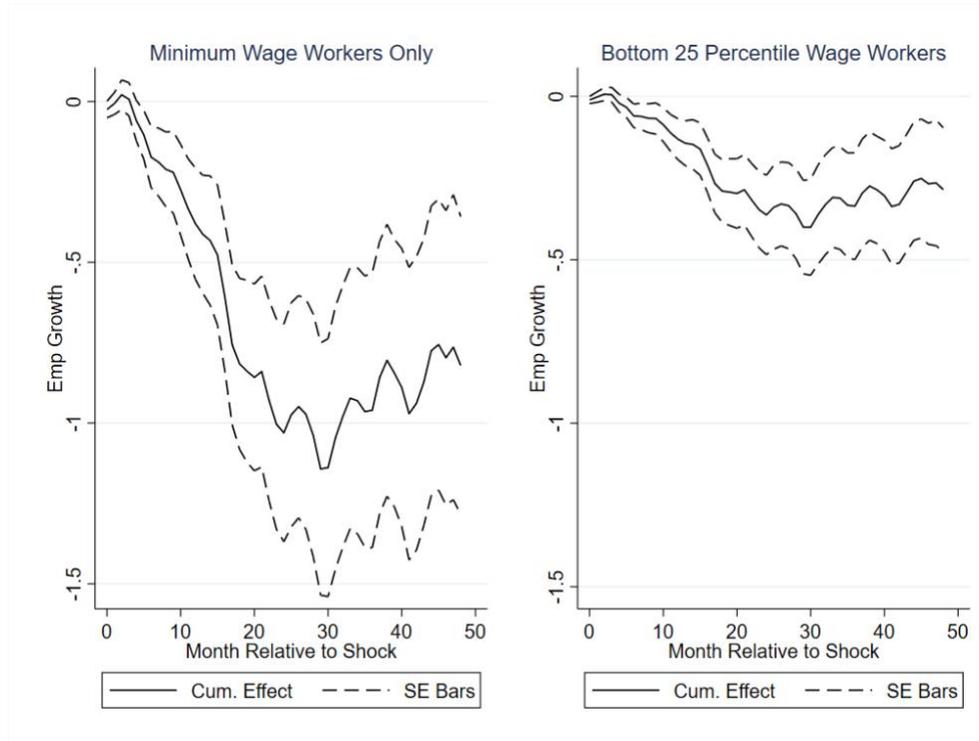
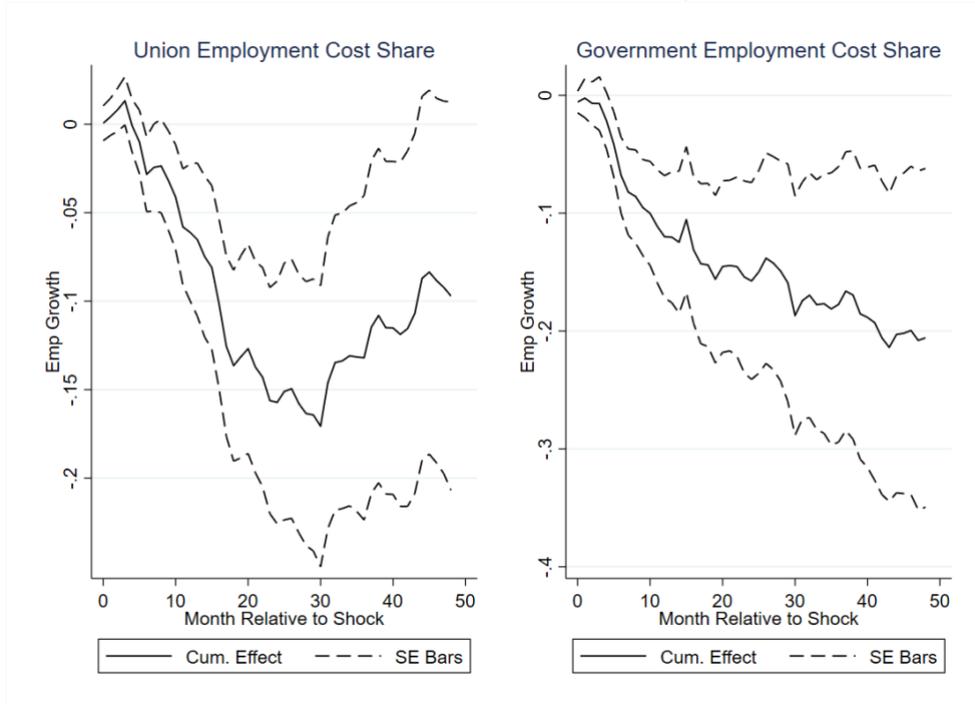


Figure 11: Baseline with Alternative MW-Affected Shares – Effects of a 1pp MP Shock on Emp.  
Interaction Coefficient between Shock and MW-affected share



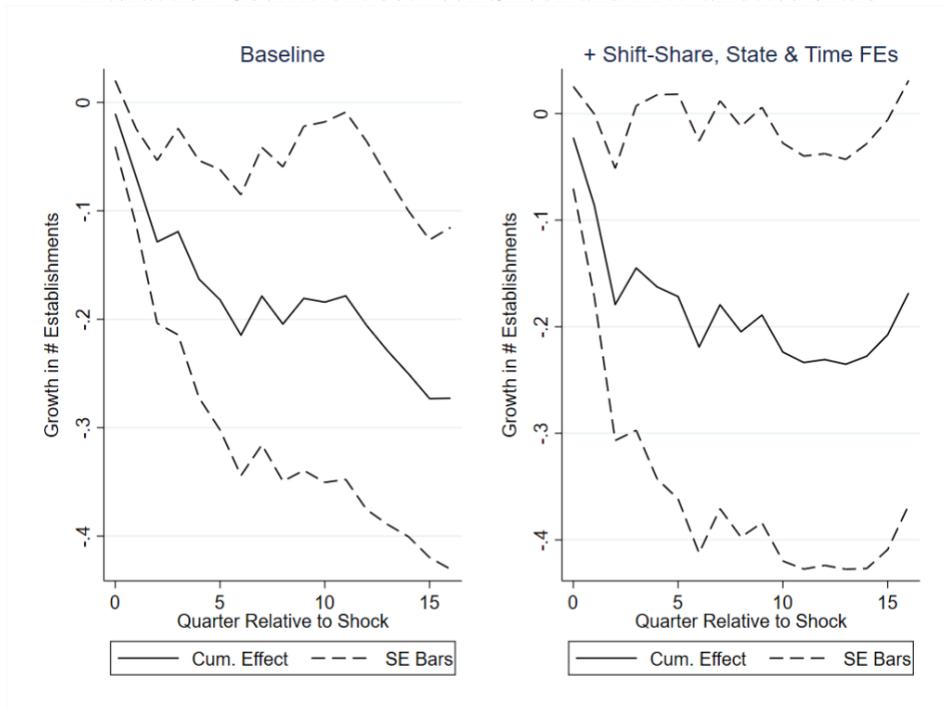
Note: These specifications duplicate the baseline specification but with alternatively-defined MW-affected shares. The share in the left panel includes only workers in a 10% band around the minimum wage. The share in the right panel includes workers up to the 25<sup>th</sup> percentile of the wage distribution.

Figure 12: Union and Govt. Share Specifications – Effects of a 1pp MP Shock on Employment Interaction Coefficient between Shock and Union, Government Share



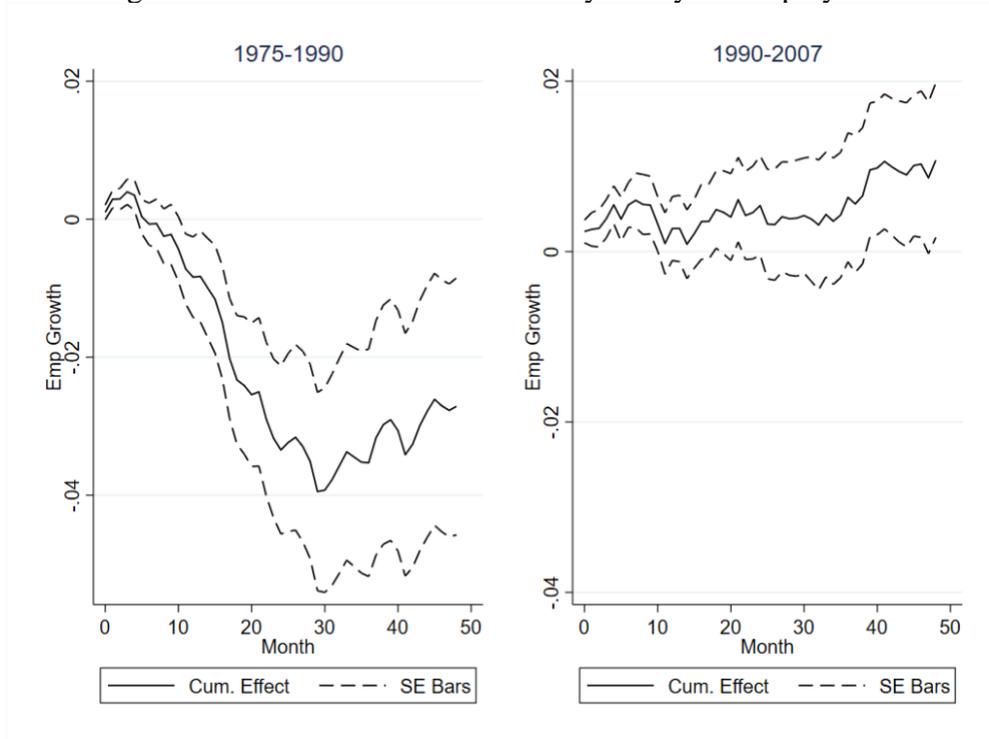
Note: These specifications duplicate the baseline specification but replace the MW-affected share with the cost share of employment covered by union contracts (left) and the cost share of government employment (right).

Figure 13: Establishments – Effects of a 1pp MP Shock on Number of Establishments Interaction Coefficient between Shock and MW-affected share



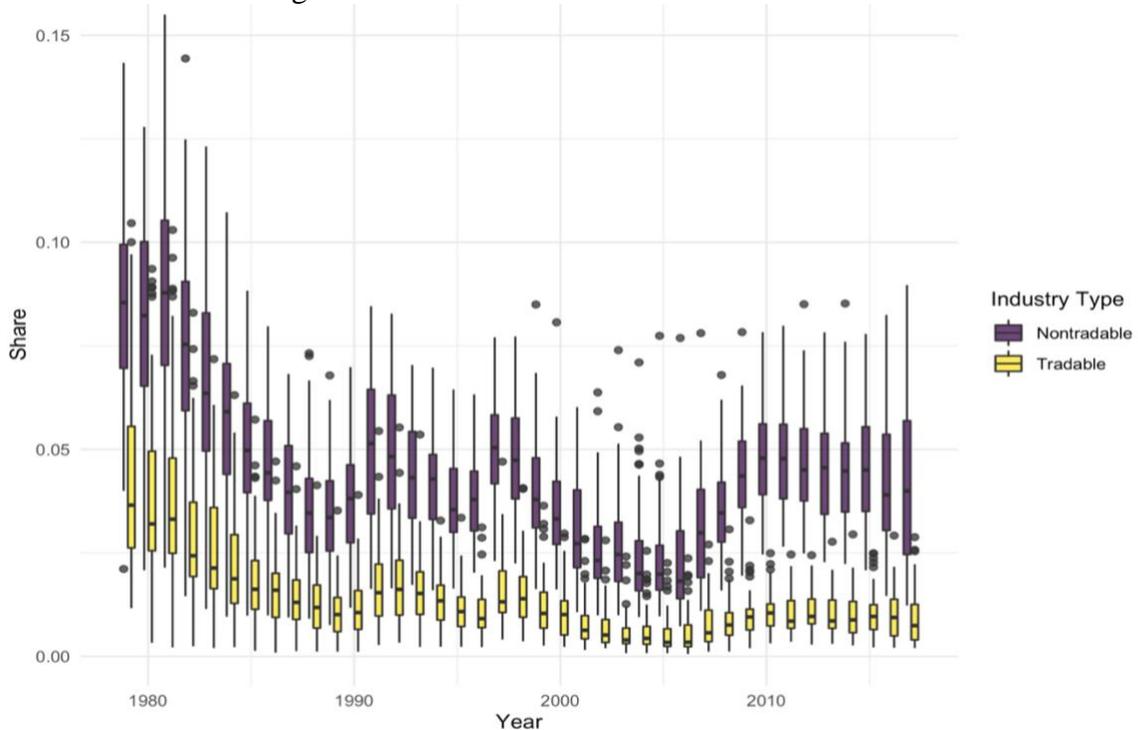
Note: These plots repeat the baseline specification, albeit using QCEW data on the number of establishments as the outcome.

Figure 14: Overall Effects of Monetary Policy on Employment



Note: These specifications are regressions of employment growth on Romer and Romer shocks (and their lags) with no interaction terms.

Figure 15: MW-affected Cost Share Calibrations



Note: The boxplot shows heterogeneity across states in the share of total sectoral cost in labor affected by minimum wages. This is much smaller than the share of sectoral employment affected by minimum wages, as (1) less skilled payroll per person is lower than skilled payroll per person, and (2) total sectoral cost also includes payments to capital.

Figure 16: Model Outcomes

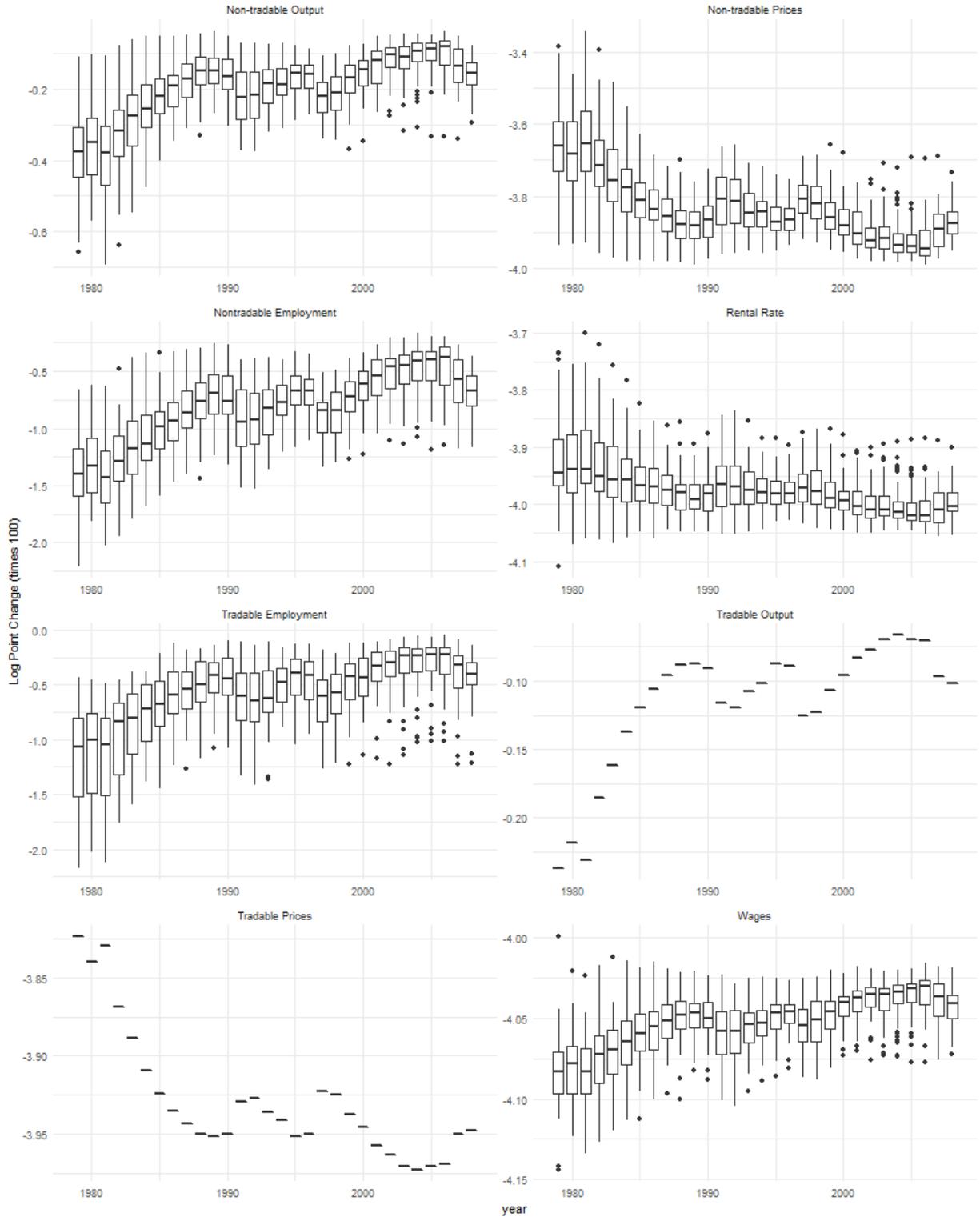
Panel 1

Peak Effect over 4 Years of a 1pp Unexpected Increase in the Federal Funds Rate  
when the shock occurs in the x-axis denominated year



Panel 2

Peak Effect over 4 Years of a 1pp Unexpected Increase in the Federal Funds Rate when the shock occurs in the x-axis denominated year

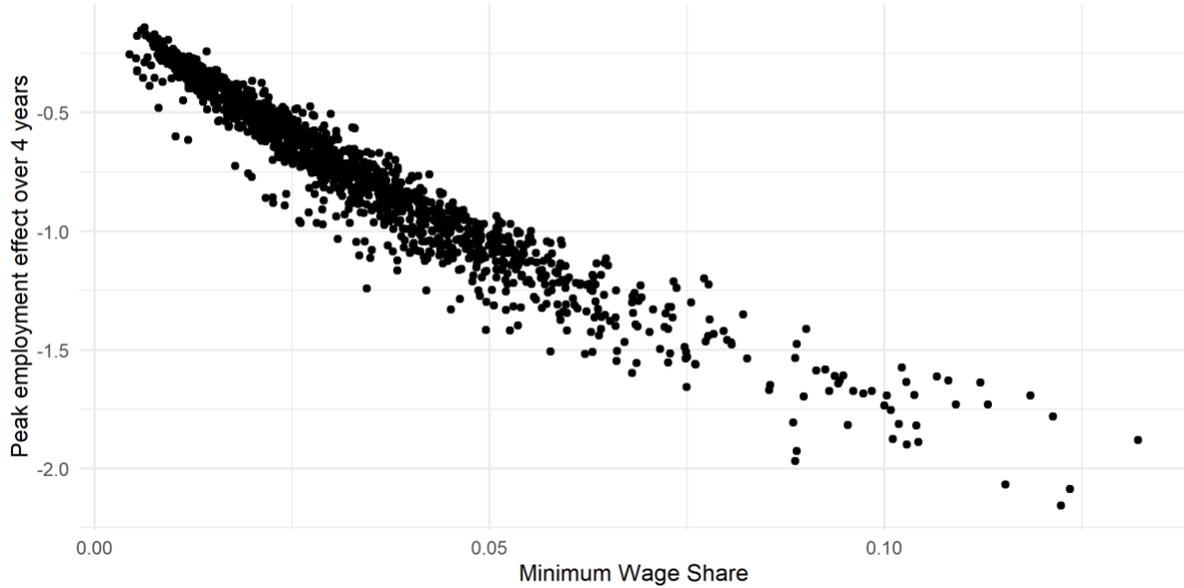


Note: Boxplots show the heterogeneity in the effects of monetary policy across states, when the monetary shock occurs in the year designated on the x-axis. The cost shares of inputs vary across years in the model calibration.

Figure 17: How Employment Effects Vary with the MW-affected share

Panel 1

Peak Effect of a 1pp Federal Funds Rate Shock  
As the minimum wage share varies



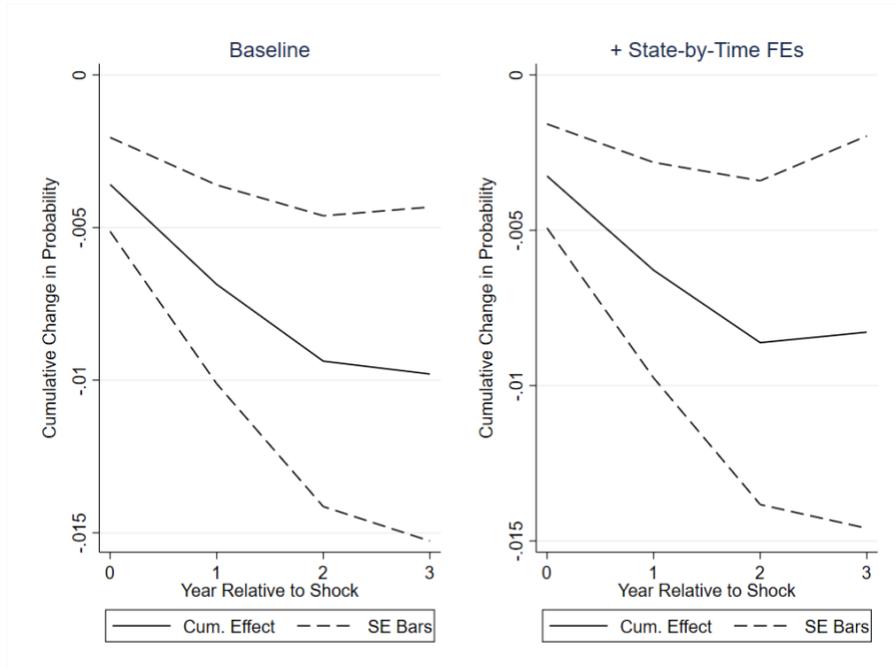
Panel 2

Peak Effect of a 1pp Federal Funds Rate Shock  
As the sectoral minimum wage shares vary



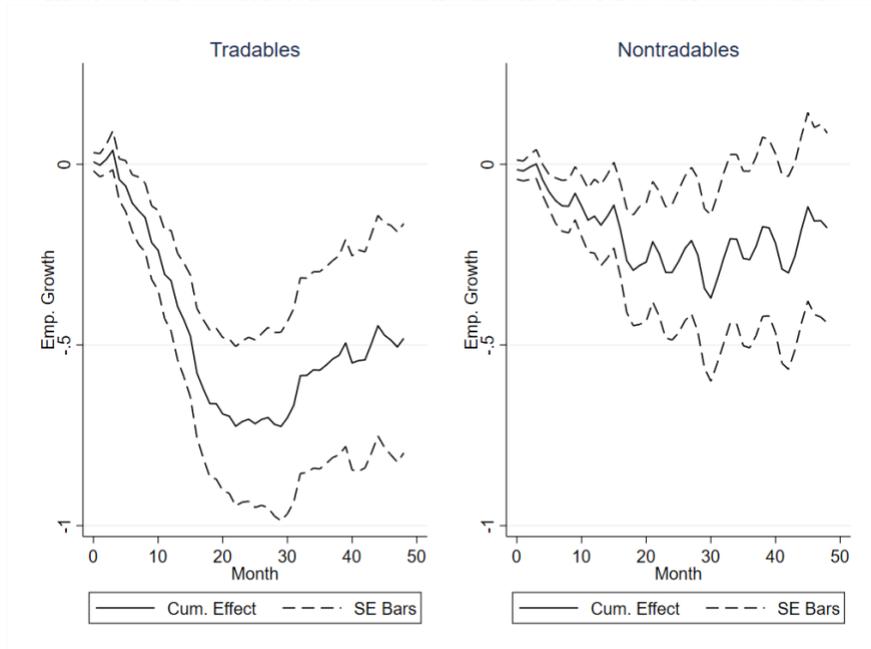
Note: The figure shows how the state-level employment effects of monetary policy generated by the model vary with state-level MW-affected shares. In panel 2, the effects are split further into state-by-sector level employment effects as a function of state-by-sector level MW-affected shares.

Figure 18: Longitudinal CPS – Effects of a 1pp MP Shock on Probability of Being Fired for Minimum Wage Workers Relative to non-Minimum Wage Workers



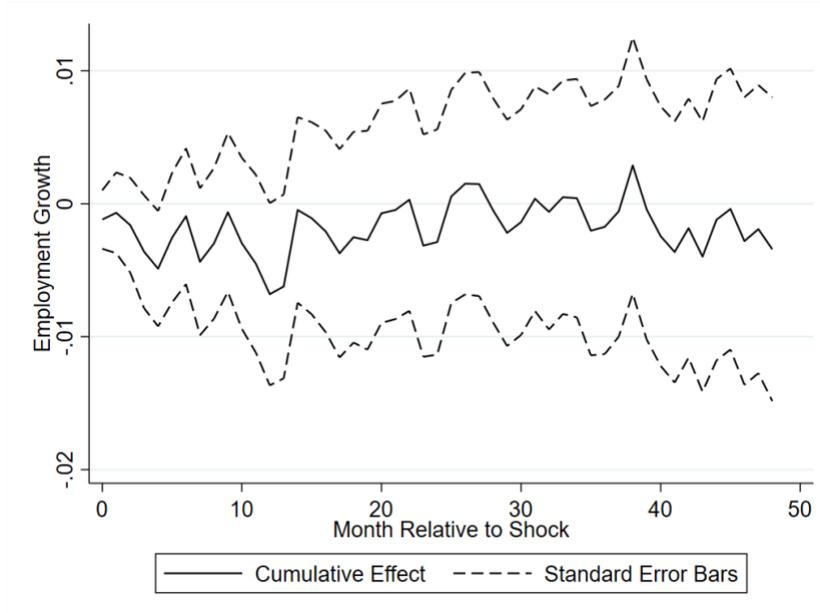
Note: These plots show results of the individual-level specification given by Equation (6), which uses longitudinal CPS-ORG data to study the effects of monetary shocks on near-minimum-wage relative to higher-wage workers.

Figure 19: Baseline Specification for Tradable and Non-Tradable Employment Interaction Coefficient between Shock and MW-affected share



Note: Sectors classified as tradable are Agriculture, Mining, Manufacturing, Financial Services, and Wholesale Trade. Sectors classified as non-tradable are Construction, Transportation, Communications, Utilities, Retail Trade, Non-Financial Services, and Public Administration. The result of a stronger effect on tradables is robust to omitting or re-classifying Financial Services, Wholesale Trade, and Retail Trade.

Figure 20: Effect of a 1pp MP Shock on Tradables Employment Relative to Non-Tradables  
 Level Effect with a 0% MW-affected share



Note: Because production of tradables can readily be moved across states and then exported, the lack of a differential effect of monetary policy on tradable employment relative to non-tradable employment in low minimum-wage-share states suggests the absence of business-stealing effects.

# Appendix

## A.1 Existence of Steady State

Steady state in our model requires constant real variables. We show that there exists a steady state with a constant growth rate of the money supply. This steady state requires that the exogenous minimum wages in each state grow at the same rate as the money supply. We start with the intertemporal Euler equation, which says

$$\dot{c}_{s,t} = \frac{1}{\gamma} \left( \frac{R_{s,t}}{P_{s,t}} - (\delta + \rho) \right).$$

For all  $s$  and for all  $t$  in steady state, it must therefore hold that

$$\frac{R_{s,t}}{P_{s,t}} = \delta + \rho.$$

Thus, in each state  $s$ ,  $R$  and  $P$  must change at the same rate in steady state. The intratemporal Euler equation is

$$U'(C_{s,t}) \frac{W_{s,t}}{P_{s,t}} = V'(H_{s,t})$$

This tells us that in each state  $s$ ,  $W$  and  $P$  must change at the same rate in steady state. Note now that the consumer substitution equation in steady state,

$$\dot{y}_{s,t}^{NT} - \dot{y}_{s,t}^T = 0 = -\sigma_{NT,T} (\dot{p}_{s,t}^{NT} - \dot{p}_t^T),$$

tells us that all non-tradable prices must grow at the same rate as the national tradable price in steady state. This in turn tells us that the state price indices grow at the same rate in all states, since

$$\dot{p}_{s,t} = \frac{GDP_s^{NT}}{GDP_s} \dot{p}_{s,t}^{NT} + \frac{GDP_s^T}{GDP_s} \dot{p}_t^T.$$

Our above analysis then yields that  $R$  and  $W$  also grow at the same rate in all states, the same rate as the state price indices. This rate of price growth is given by the rate of money growth, since

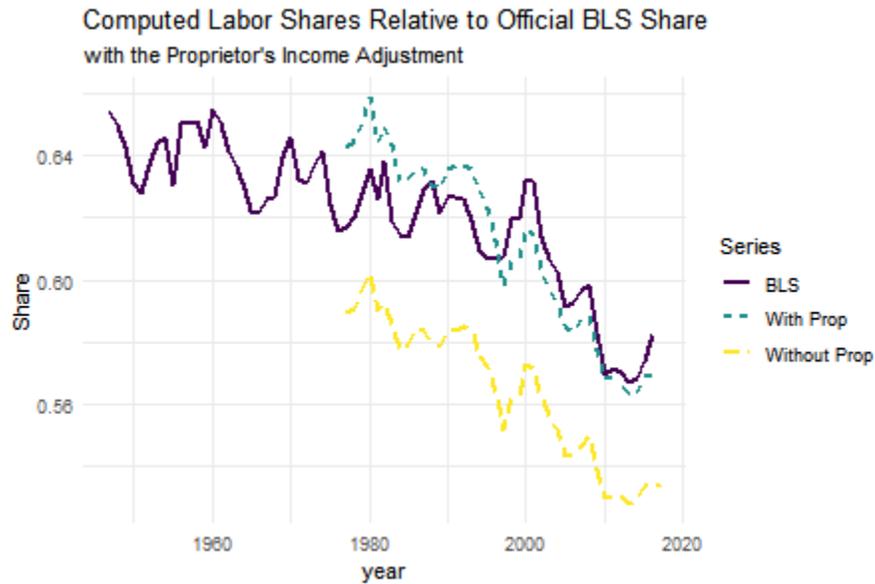
$$\dot{m}_t = \frac{GDP'_s}{GDP_t} \dot{p}_t.$$

Now the profit maximization equations of the tradable and non-tradable sectors tell us that it must be that minimum wages grow at the same rate in all states, a rate that is given by the rate of money

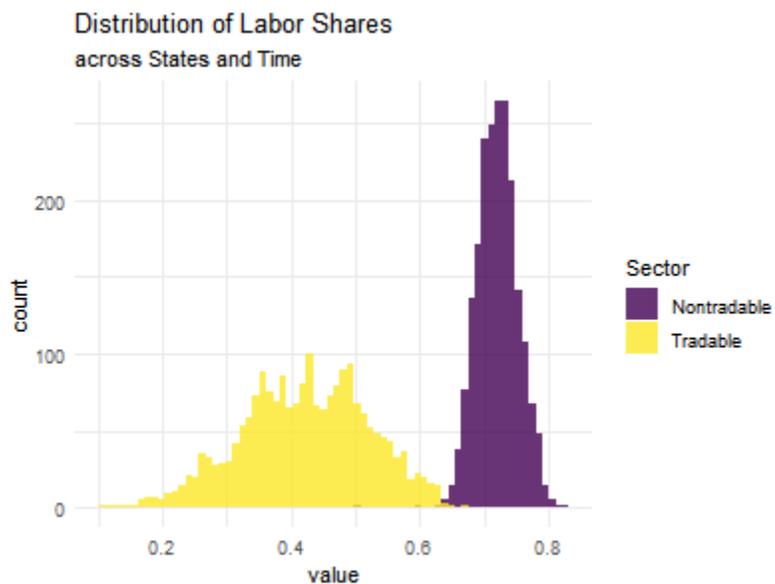
growth.

## A.2 Additional Plots for Calibrations

Below we show how our computed labor share at the national level compares to the BLS share. Adjusting for proprietor's income as the BLS suggests brings our calculation much closer to theirs.

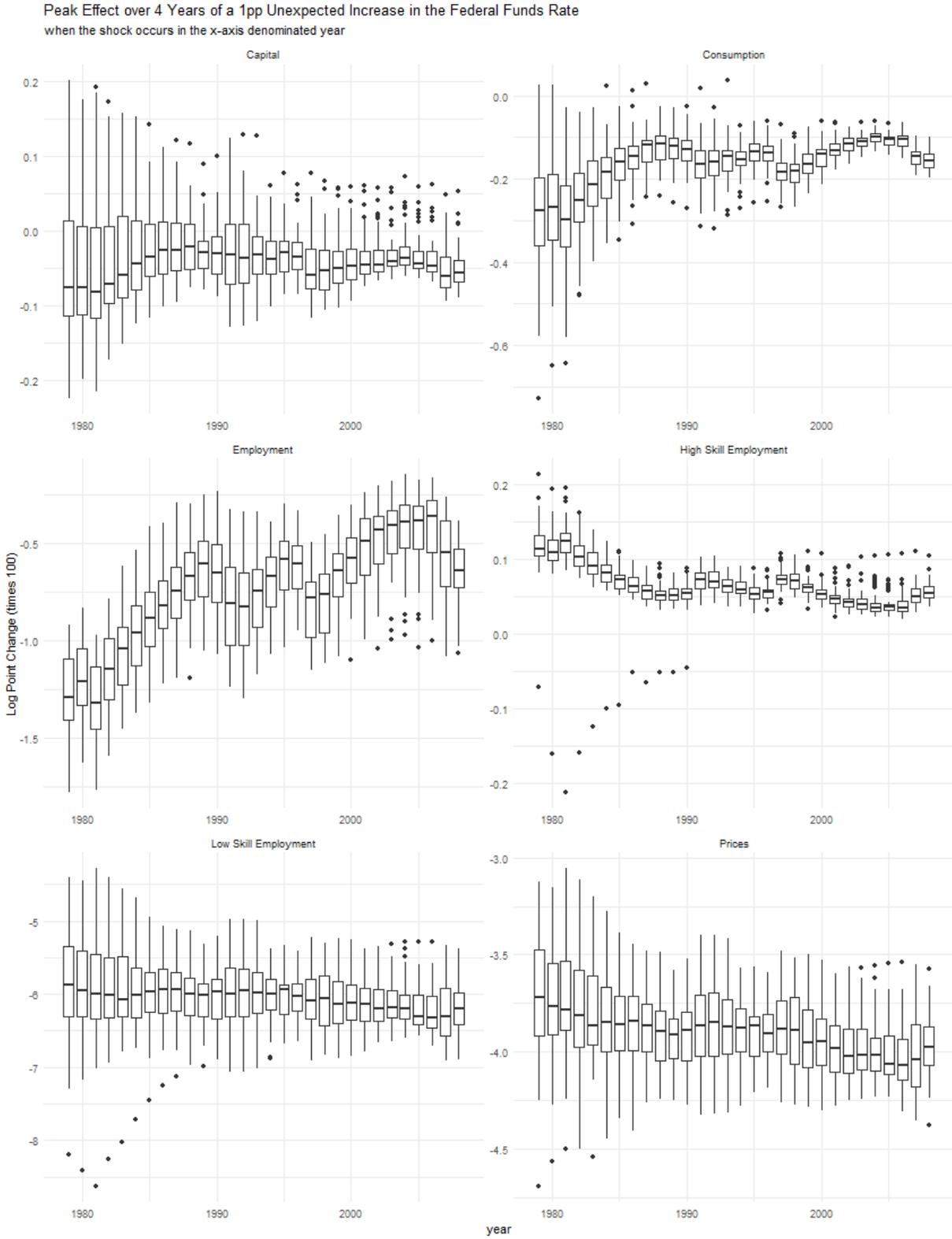


Below, we compute our labor shares using the BLS methodology at the state level, separately for tradables and non-tradables. Shares are much lower in the tradable sector.

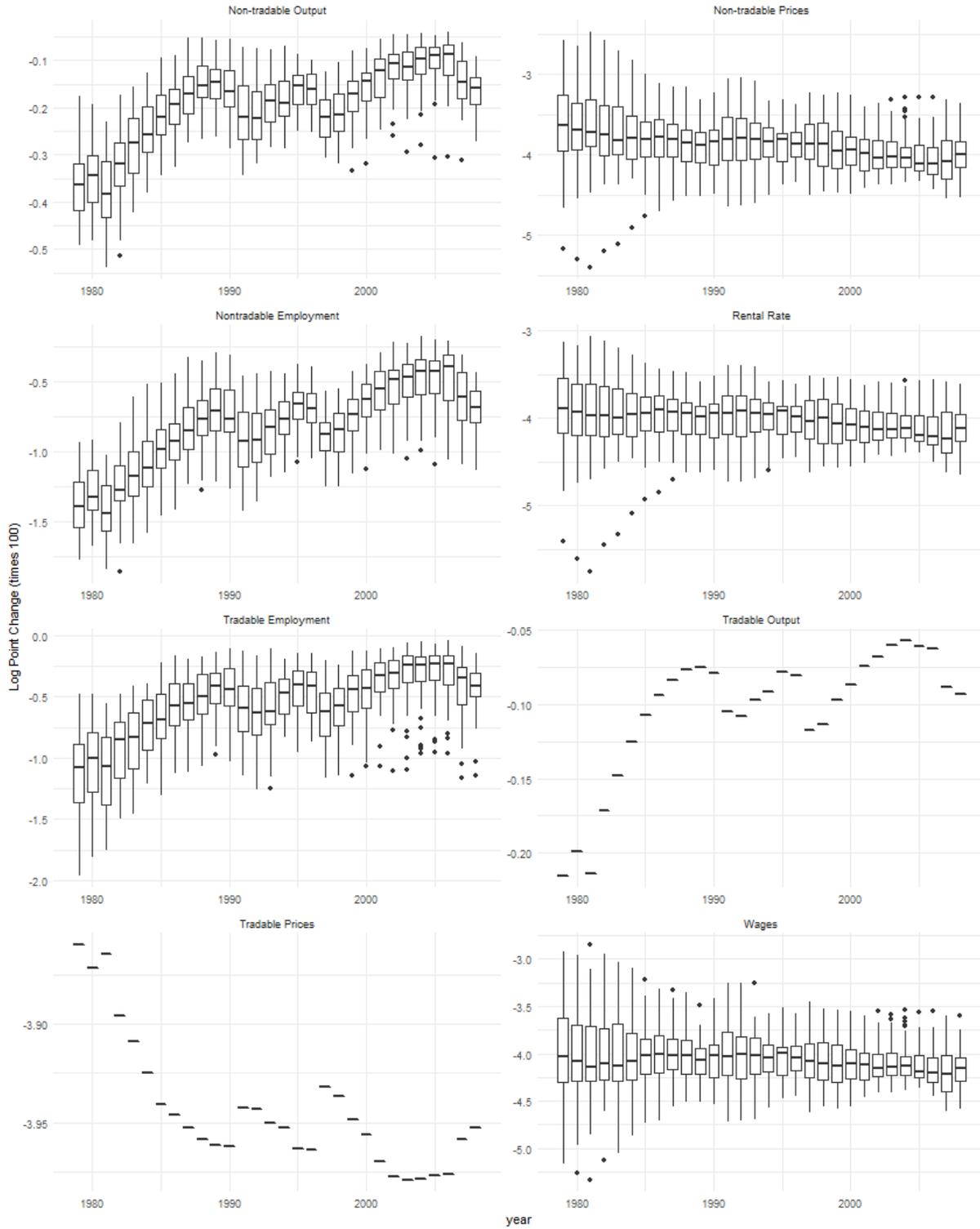


### A.3 Additional Model Outcomes

Below, we show all model plots in the paper for an alternative specification,  $\sigma_s = .1$ :

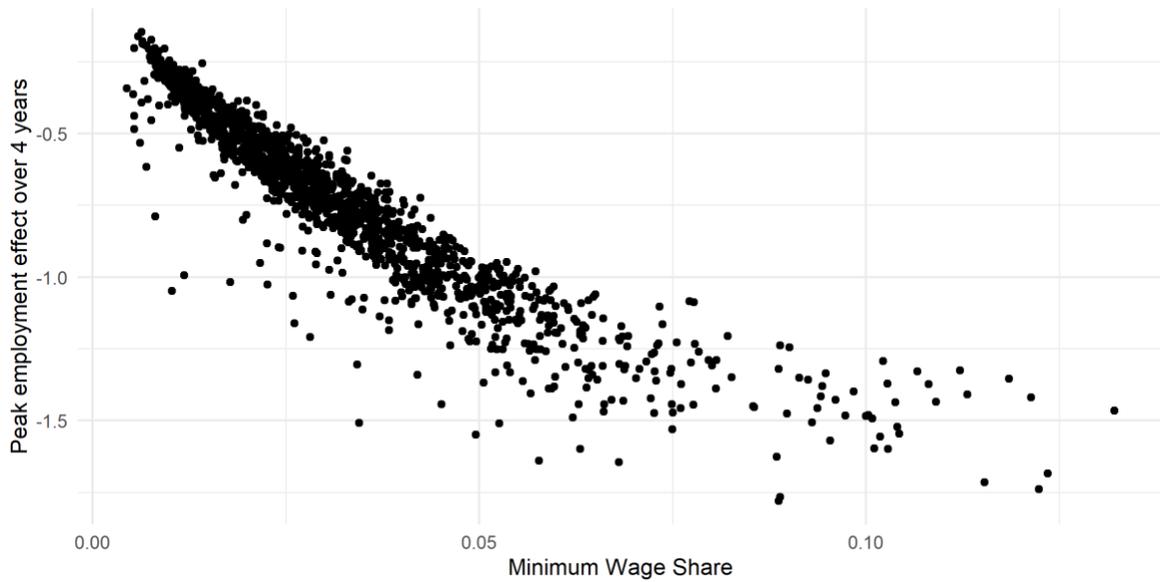


Peak Effect over 4 Years of a 1pp Unexpected Increase in the Federal Funds Rate  
when the shock occurs in the x-axis denominated year

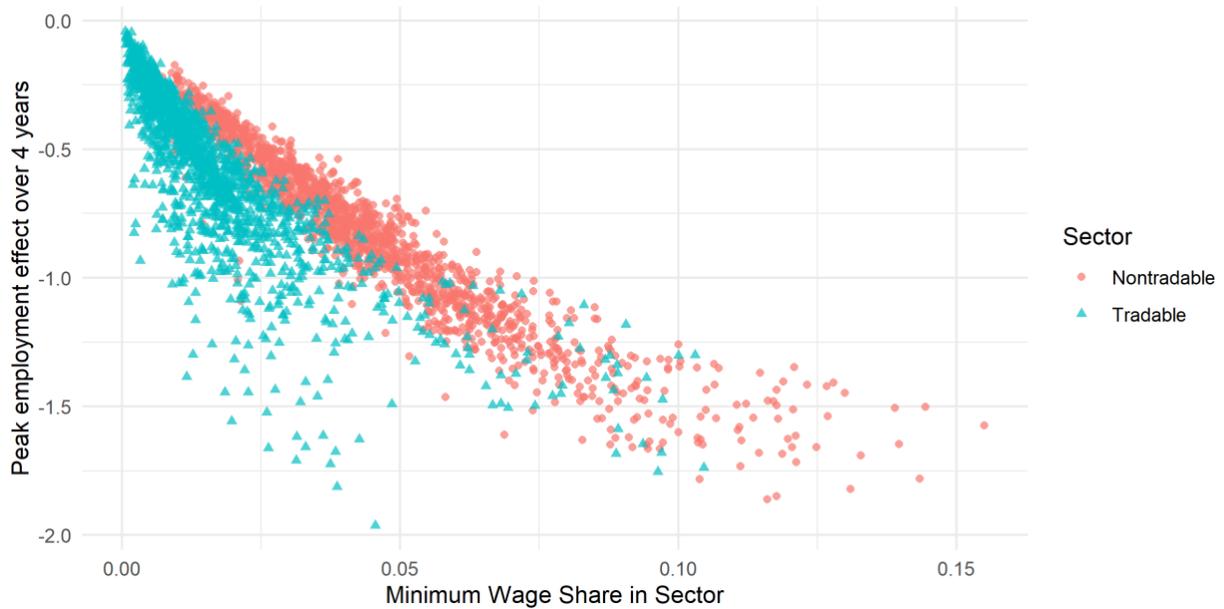


Note: Boxplots show the heterogeneity in the effects of monetary policy across states, when the monetary shock occurs in the year designated on the x-axis. The cost shares of inputs vary across years in the model calibration.

Peak Effect of a 1pp Federal Funds Rate Shock  
As the minimum wage share varies



Peak Effect of a 1pp Federal Funds Rate Shock  
As the sectoral minimum wage shares vary



Note: The figure shows how the state-level employment effects of monetary policy generated by the model vary with state-level MW-affected shares. In panel 2, the effects are split further into state-by-sector level employment effects as a function of state-by-sector level MW-affected shares.