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The Pass-through of Minimum Wages into US Retail Prices: Evidence from Supermarket Scanner Data[a]

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Abstract

This paper estimates the pass-through of minimum wage increases into the prices of US grocery and drug stores. We use high-frequency scanner data and leverage a large number of state-level increases in minimum wages between 2001 and 2012. We find that a 10% minimum wage hike translates into a 0.36% increase in the prices of grocery products. This magnitude is consistent with a full pass-through of cost increases into consumer prices. We show that price adjustments occur mostly in the three months following the passage of minimum wage legislation rather than after implementation, suggesting that pricing of groceries is forward-looking. The rise in prices occurs mostly through an increase in the frequency of price increases. Prices rise to the same extent for goods consumed by low-income and high-income households. Our results suggest that consumers rather than firms bear the cost of minimum wage increases in the retail sector.

Keywords: Minimum wages, inflation, retail prices, price dynamics, price pass-through

JEL: E31, J23, J38, L11, L81

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1 Introduction

In recent years, a number of US states and municipalities have increased their minimum wage, in a context of low wage growth and stagnation of the federal minimum wage. Similarly, several European countries have introduced a national minimum wage (e.g., Germany) or hiked their minimum wage (e.g., the United Kingdom). A large body of research in economics shows that moderate increases in the minimum wage have no or limited dis-employment effects (see e.g., Card and Krueger, 1994; Belman and Wolfson, 2014; Cengiz et al., 2019), suggesting that such a policy can raise nominal incomes of low-wage workers. However, there is much less evidence on how changes in the minimum wages affect consumer prices (see Lemos (2008) for a literature review) and therefore *real* incomes. In principle, it is possible that nominal wage increases for low-wage workers may be partly offset by increases in the prices of the goods and services consumed by the poorest households. To assess the economic impact of minimum wage changes on real incomes, it is thus central to understand the pass-through of minimum wage increases into prices.

In this paper, we study the pass-through of minimum wages increases into prices in the US. We exploit a large number of changes in the minimum wage between 2001 and 2012 and leverage scanner-level data from weekly price observations of 2,500 distinct grocery and drug stores. We make three main contributions. First, we provide new evidence on how minimum wages affect prices in the grocery sector, which had not been previously studied in the literature.[1] The grocery sector is especially important because the share of minimum wage labor costs in groceries' marginal cost is sizable, and because groceries make up a large share of consumer expenditure, up to 15% for low-income households. Second, we take advantage of the high frequency of scanner data to study the dynamics of the price response over time. Since minimum wage laws are usually passed several months before implementation and typically set a schedule of increases rather than one-off hikes,

[1]In this article, we use "grocery sector" for grocery and drug stores. Likewise, we sometimes use "grocery stores" for grocery stores *and* drug stores.

firms may increase prices in anticipation of higher future minimum wages. We use a newly collected dataset with legislation dates for every minimum wage increase in our sample period and we find strong evidence for anticipation effects. Third, we use a large consumer panel data linked to the store-level information to investigate how the price response varies across household income groups. This allows us to better understand the implications of minimum wage changes for real incomes.

Our main finding is that there is a full pass-through of minimum wage increases into grocery prices. Our main research design compares monthly price movements across states exploiting time variation in state-level minimum wage hikes. We supplement this approach by using a second identification strategy that exploits within-state variation in the bite of the hikes. We find that a 10% minimum wage hike translates into a 0.36% increase in grocery prices. Importantly, there is no statistically significant difference between the average price elasticity of 0.036 and our estimate of the minimum wage elasticity of groceries' costs, which suggests a full pass-through of minimum wage cost increases into prices. We do not find evidence that the demand for grocery products changes, nor do we find evidence that stores reduce employment. Taken together, these results suggest that consumers, rather than firm-owners or workers, bear the bulk of the burden of minimum wage increases in the grocery sector.

Another important finding of this paper, with implications for macroeconomic models, is that price adjustments occur mostly in the three months following the passage of a minimum wage legislation, rather than after implementation. In other words, grocery and drug stores appear to be forward-looking in their pricing decisions. Using Google Trends data, we show that the legislation of minimum wage increases represents a very salient event in the public. Based on flexible event study regressions tracking prices around the months in which minimum wage hikes are legislated, we find that grocery stores respond to future cost increases by increasing prices months before the minimum wage is actually implemented. This type of forward-looking behavior of firms is qualitatively consistent with the predictions of purely rational models, where firms think about the future as well as the present (i.e. they are not myopic). The rise in prices occurs mostly through an increase in the frequency of price changes.

Last, we quantify the welfare consequences of minimum wage hikes after accounting for our estimated pass-through of minimum wages into prices. We estimate that the price effects of minimum wage increases are similar for goods usually consumed by low-income and high-income households. Low-income households are nevertheless disproportionately affected by the rise in grocery prices since a larger share of their expenditures is on groceries. The rise in grocery store prices following a $1 minimum wage increase reduces real income by about $19 a year for households earning less than $10,000 a year, and by about $63 a year for those earning more than $150,000. The price increases in grocery stores offset only a relatively small part of the gains of minimum wage hikes. Minimum wage policies thus remain a redistributive tool even after accounting for price effects in grocery stores.

This paper contributes to a body of work in labor economics and macroeconomics.

First, this paper provides novel insights into the redistributive effects of minimum wages and into the price effects of minimum wages in low-wage sectors. A small literature studies the product market effects of minimum wage increases. This literature has focused on restaurants (see e.g., Aaronson, 2001; Allegretto and Reich, 2018).[2] Our contribution to this literature is to study the impact of minimum wage changes in a new sector, the grocery sector. This sector employs a high and rising share of workers at or just above the minimum wage: we document that the share of workers earning below 110% of the minimum wage was 12% in 2001-2005; it was up to 25% in 2010-2012. We also document the existence of large spikes around the minimum wage in this sector. Almost 50% of workers earned below 130% of the minimum wage in 2010-2012.[3] Therefore, the effect of minimum wage hikes is potentially large. Moreover, groceries are an important component of households' cost of living, particularly for poor households. Groceries make

[2]Outside of the US, Fougère et al. (2010) analyze the response of restaurant prices to an increase in the French minimum wage. Harasztosi and Lindner (2019) analyze the price response of a large minimum wage increase in Hungary in the manufacturing sector.

[3] See Appendix Table H.1 and Figure B.1.

up 11% of household expenditures, two to three times more than spending on restaurant meals, depending on household income (see Appendix Table B.2).

We break new ground in documenting the price response in the retail sector thanks to the availability of high-quality scanner-level data. The use of this data enables us to overcome certain shortcomings in studies of the price effects of minimum wages. These limitations include classical measurement error (Card and Krueger, 1994; Aaronson, 2001), the use of city-level CPI data that are only available in the largest US metro areas (Aaronson, 2001; Aaronson and French, 2007; Aaronson et al., 2008), and the fact that price and wage changes in restaurants may not be well measured due to tipping and quality changes (e.g., size of portions served). These concerns do not apply to retail scanner data, as products in grocery stores are very standardized and retail workers are not tipped. Compared to official BLS price indexes, our micro-data allow us to compute price changes by income group (as well as price changes conditional on non-zero price adjustments).[4]

Most closely related to our work are the contemporary papers by Leung (2020) and Ganapati and Weaver (2017), who also study the pass-through of minimum wage changes into retail prices. These papers focus on a different period (2006–2015 and 2005–2015, respectively, vs. 2001–2012 in our study), are based on another dataset (the Nielsen data), and use different identification strategies. Ganapati and Weaver (2017) finds a zero pass-through of minimum wage increases into prices, and Leung (2020) more than a full pass-through. In Appendix K, we reconcile our findings with these two studies. The main substantial difference between our work and these two studies is that we document the forward-looking pricing decision of grocery stores, by studying the effect of minimum wage legislation (before implementation) on subsequent price changes. Two other distinctive features of our work are that we study in detail whether our results are consistent with full pass-through of prices into costs, and that we quantify the extent to which the price

[4]The official BLS indexes, although less detailed than micro-data, also have a number of strengths to study the effect of minimum wage changes: they are weighted at all levels of aggregation, rely on case-by-case adjustments for item turnover, and the BLS has established procedures for dealing with missing price observations. See Section 2.1 below for a comparison of the price indices constructed using our data and the official BLS price indices.

increases in grocery stores affect the redistributive effects of minimum wage policies.

Second, our paper contributes to the macroeconomic literature on price-setting. We provide causal evidence of the effect of a rise in labor costs on retail price inflation. This adds to the macro literature that has mainly focused on the effects of rising wholesale costs on pricing decisions.[5] Our detailed micro-data allow us to document a price response to a future cost shock at the time it becomes known and several months before it actually occurs. Because minimum wage changes can be seen as a shock to grocery store activities which is plausibly exogenous, these shocks can help identify the effect of movements in costs on prices. Our results highlights the role of expectations in the propagation of shocks. Forward-looking price setting is a central prediction of state-dependent models (i.e. menu cost models), as well as time-dependent models with nominal frictions. These latter models include the Calvo (1983) model of staggered price setting and models with adjustment costs such as Rotemberg (1982). In the macroeconomics literature, these models have been used as a microeconomic foundation for the New Keynesian Phillips Curve (see, e.g., Galí, 2015).

Finally, we contribute to the research on price rigidity in retail chains. We provide evidence that chains try to maintain uniform prices across grocery stores in the US. We we find that, within interregional chains, a minimum wage hike in one state affects prices in stores within the same chain located in another state. These results suggest that minimum wage hikes can affect consumer welfare in other states. Consistently, we find that grocery prices are more responsive to local minimum wage hikes in regional chains than in national chains. This is consistent with Dellavigna and Gentzkow (2019) who document uniform pricing decisions in the retail sector in response to local economic shocks in general, and to Leung (2020) who documents this behavior in the case of local minimum wage hikes.

The remainder of this paper is organized as follows. The next section details the

[5]For instance, Eichenbaum et al. (2011) find that pass-through is complete but somewhat delayed. Nakamura and Zerom (2010), using variation in the market price of commodity coffee, find that the pass-through into wholesale prices is about one third, but that the increase of wholesale prices is completely passed through to consumers by retail stores.

data we use. Section 3 describes our main identification strategy. Section 4 contains our estimates of the retail price elasticity with respect to the minimum wage, discusses robustness checks and analyzes the heterogeneity of this price response along several dimensions. Section 5 characterizes the anatomy of the price response. Section 6 studies the magnitude of the price pass-through elasticity. Section 7 concludes.

2 Data description

We combine several data sets to conduct our empirical analysis. We begin by describing the construction of our key dependent and explanatory variables before detailing the other data sources we use.

2.1 Retail price data: IRI data

Retail scanner data. Our empirical analysis is based on scanner data provided by the market research firm Symphony IRI. The dataset is described in detail in Bronnenberg et al. (2008) and Kruger and Pagni (2013). It contains weekly prices and quantities for 31 product categories sold at grocery and drug stores between January 2001 and December 2012. The estimation sample covers 2,493 distinct grocery and drug stores and contains their zip code location.[6] It provides information on an average of 60,600 products over this period. Products are identified by Unique Product Codes (UPC). As an example, a 12oz can and a 20oz bottle of Coca Cola Classic are treated as different products in our data. Stores are located in 530 counties, 41 states and belong to one of about 90 retail chains. The data covers 17% of US counties which are home to about 29% of the US population.[7] Most of the included product categories are packaged food products (frozen pizza, cereals, etc.) or beverages (soda, milk, etc.). The data also includes personal care products (deodorants, shampoo, etc.), housekeeping supplies (detergents, paper towels, etc.), alcoholic beverages (beer and some flavored alcoholic beverages) and tobacco.

[6] Grocery stores make up three-quarters of the stores' sample. Drug stores make up one fourth.
[7] Figure A.1 in the appendix shows the regional distribution of stores.

Our key dependent variable is the monthly store-level price inflation, defined as follows:

$$\pi_{jt} = \log I_{jt} \text{ with } I_{jt} = \prod_{c} I_{cjt}^{\omega_{cjy(t)}} \tag{1}$$

where π_{jt} is the inflation rate in store j in month t; I_{jt} is a single Laspeyres price index at the store level that aggregates price changes across product categories c; the weight $\omega_{cjy(t)}$ is the share of product category c in total revenue in store j and month t. We detail in Appendix A how we constructed store-level price indices.

There are several reasons why we use store-level indices instead of more disaggregated product level price data. First, wages are paid at the store and not at the product level, and we thus think that stores are the natural unit of analysis. Second, it is useful to weight products by their importance for stores and consumers, and store-level price indices are a natural way to do so. Third, entry and exit are much less of a concern at the store level than at the product level. Especially low-volume products are frequently introduced and discontinued, and may also go unsold for extended time periods due to stock-outs, seasonality or low demand. This results in frequent gaps in products' price series. By contrast, our panel at the store level is much more balanced. A fully balanced panel is obtained when we conduct our analyses at the state-level, rather than at the store level. We show in section 4.2 and Appendix C that our main results are robust to changing the level of analysis from the store to the state level.

Our approach closely follows methods used in previous articles on retail price movements (see, e.g., Coibion et al., 2015). We show in appendix A that the features of our price index are in line with what other researchers have documented for the IRI Symphony data, and other retail price datasets. We also show that our price index correlates well with inflation measures provided by the Bureau of Labor Statistics (see Appendix Figure A.2). Importantly, we apply a filter suggested by Kehoe and Midrigan (2015) to remove temporary price fluctuations (i.e. sales). The algorithm uses a moving window modal price to determine a "regular price" at any point in time. There are two reasons why we remove sales from our price series: first, we are interested in capturing the *permanent* effect of minimum wages on prices; second, we are interested in studying the *dynamics* of the price response – something that turns out to be empirically infeasible in our demanding specification when sales are inclduded in the price series, because of the multicollinearity issue it introduces.[8] We discuss the two reasons of this choice in more detail in Section 4.2, and show how incorporating sales affects our results.

Consumer panel data. In addition to the retail scanner data, IRI provides a consumer panel dataset with shopping data for about 5,000 households in two local markets: Eau Claire, Wisconsin and Pittsfield, Massachusetts. In general, the shopping data also covers purchases at grocery and drug stores that are not covered by the IRI price data. The panel contains details about household demographic characteristics (e.g., race and education) and most importantly for us, pre-tax household income (in relatively detailed brackets). This is the data we use in section 4.3 to study whether prices of goods consumed by low-income households increase more than prices of goods of high-income households.

2.2 A new minimum wage database

We construct a minimum wage database of federal and state-level minimum wage increases by pulling together data from the Tax Policy Center, the US Department of Labor, and state departments of labor. For each state, we collect the legally binding rate, i.e. the maximum of federal and state rates.[9]

The novelty of our database is that in addition to the implementation dates of minimum wage hikes, we collect information on the time that each minimum wage law is passed, derived from legislative records and media sources. Since most minimum wage

[8]The multicollinearity issue arises because sales lead to a very strong seasonal pattern in month-on-month inflation rates. Including sales impairs our ability to separate seasonal movements in prices from the effects of minimum wages which are often implemented step by step in intervals of 12 months. This makes our specification in monthly first-differences rather sensitive to specification choices

[9]We focus on state-level minimum wage changes in our paper, and not on city or county-level changes, because from 2001 to 2012, only San Francisco, CA, and Santa Fe, NM, had local minimum wage ordinances.

increases are known in advance, firms potentially have ample time to act in anticipation.

In some cases, passage of legislation is preceded by a series of votes and negotiations; in this case, we try to assess from media sources at which point in the process a minimum wage increase became certain. A good example is the "Fair Minimum Wages Act of 2007" that raised the federal minimum wage from \$5.15 an hour to \$7.25 in three steps in July 2007, 2008 and 2009. The act was passed in slightly different versions in January 2007. After a conference committee added tax-cuts for small businesses to the bill, the final version was passed and signed by President Bush in May 2007. Since the passage of the actual minimum wage part of the bill seemed certain already in January, we use January as the month of legislation in our baseline.[10]

An important assumption of our approach is that the legislation dates represent points in time when future minimum wage increases become more salient. We use Google Trends data to assess the plausibility of this assumption. Google trends is available from 2004 onward. We use the search volume for the term "minimum+wage+*statename*" over a month to measure interest in the local minimum wage of a given state.[11] We then estimate the following simple regression using this data:

$$\log search_{s,t} = \delta_s + \gamma_t + \sum_{r=-k}^{k} \beta_r imp_{s,t-r} + \sum_{r=-k}^{k} \alpha_r leg_{s,t-r} + \epsilon_{s,t}. \tag{2}$$

$imp_{s,t-r}$ and $leg_{s,t-r}$ are dummy variables indicating implementation of a higher minimum wage and passage of minimum wage legislation in state s in period $t-r$. The results of this regression are presented in Figure 1a. Both around implementation and around the date of legislation, interest in minimum wages goes up substantially, by about 30% immediately after legislation is passed. There is no elevated interest in minimum wages in the months before legislation is passed. Three months after passage of legislation, search volume is back at the baseline value. These results show that the passage of minimum

[10] We present results using only state-level legislation to show that our conclusions hold more generally and are not driven by this single important event.

[11] Note that we do not measure search requests originating from different states, but from the US as a whole for different search terms.

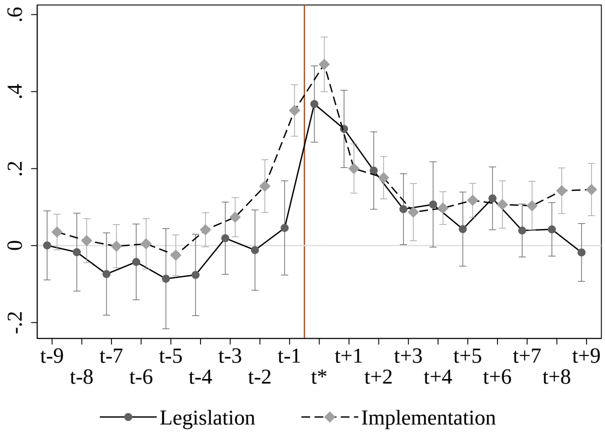
wage legislation is a salient event and that the public takes notice of pending minimum wage increases when they are written in law. The results also validate our coding choices in the collection of legislation dates.

The primary explanatory variables in our analysis are changes in the implemented minimum wage and changes in the "legislated minimum wage." Figure 1b shows how we measure the "legislated minimum wage". It is the highest future binding minimum wage set in current law. The legislated minimum wage increases to the highest future minimum wage at the time the law is passed.

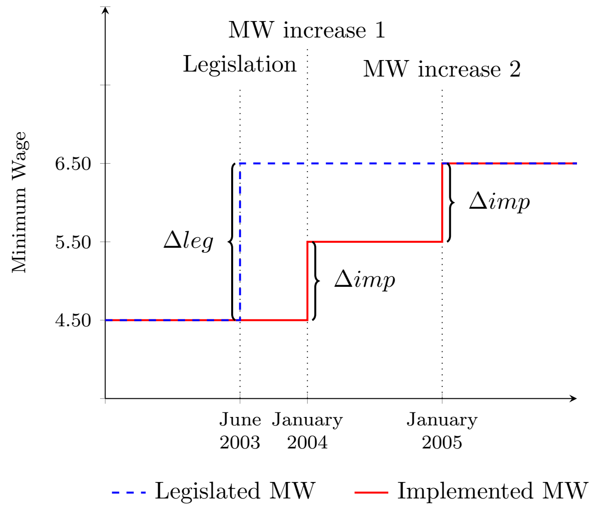
We leverage 166 changes in the implemented minimum wage and 62 changes in the legislated minimum wage. This allows us to exploit variation in minimum wages across states, time and size of hikes. Figure 2 shows the distribution of changes in the implemented and legislated minimum wage over states and time. States in our sample experience between 2 and 11 hikes. Most of the events in our sample occur between 2006 and 2009. The average increase in the binding minimum wage amounts to 8.2% (see Appendix Table B.1). Changes in the legislated minimum wage are larger on average (20%), since they usually encompass several steps. The average interval between passage of legislation and implementation of a first hike is 9 months. Hence, even the first increase in a package is often known long before it is implemented. Moreover, 36% of all increases in the implemented minimum wage and 42% of increases in the legislated minimum wage result from changes at the federal level. 24% of all increases in the implemented minimum wage result from indexation. Minimum wages in states with indexation are pegged to the national development of prices and exhibit small annual increases. We do not assign legislation dates to increases following from indexation.[12]

[12]Indexation is practiced in 10 states at the end of our sample period. The following states in our sample have indexation: Arizona, Colorado, Florida, Missouri, Montana, Nevada, Ohio, Oregon, Vermont, and Washington. Most of these states introduced indexation starting in 2008 after ballots held in November 2006. The exceptions are Florida, Vermont (both began indexation in 2007), Oregon (beginning in 2004) and Washington (beginning in 1999).

Figure 1: Explanatory variables: Changes in implemented and legislated minimum wage



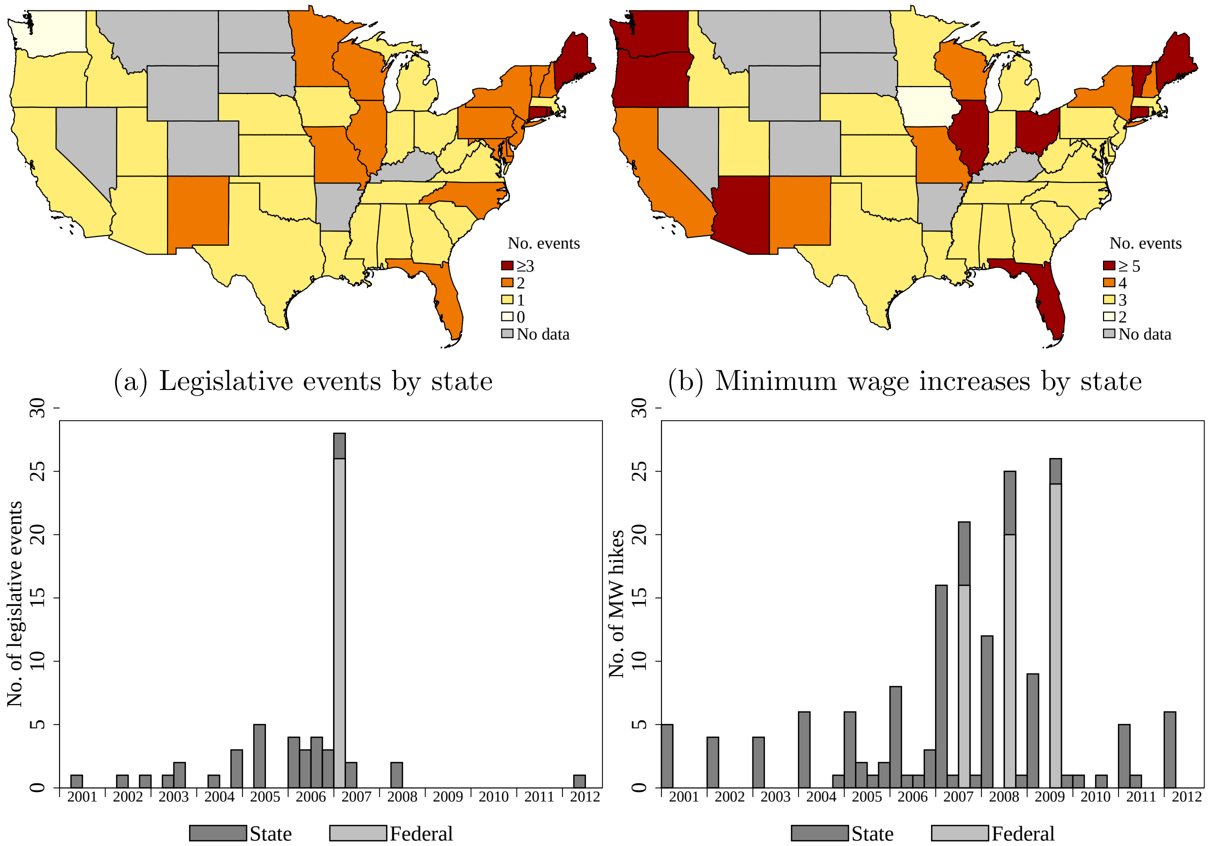
(a) Google search volume for "minimum wage *statename*" around legislation and implementation of minimum wage increases



(b) Example for the measurement of explanatory variables

Notes: Panel (a) shows the log change in monthly Google search volume for the search term "Minimum wage+*statename*" around changes in minimum wage legislation and implementation of higher minimum wages in state *statename*. The coefficients are estimated from equation 2. The effects are relative to state and time fixed effects. Note that the search terms differ between states, but measured search volume is for United States as a whole. Panel (b) illustrates the measurement of changes in the legislated and implemented minimum wage based on an hypothetical minimum wage increase in two steps. In June 2003, legislation is passed that will increase the minimum wage in from an initial value of \$4.50 to \$6.50. The law schedules an increase to 5.50 in January 2004, and to 6.50 in January 2005. Our measure of the legislated minimum wage is equal to 4.50 before June 2003. It increases to 6.50 when the legislation is passed in June 2003. Before June 2003 and after January 2005 the legislated minimum wage is equal to the implemented minimum wage.

Figure 2: Distribution of minimum wage hikes and legislative events over time and states



(a) Legislative events by state

(b) Minimum wage increases by state

(c) Legislative events over time

(d) Minimum wage increases over time

Notes: The figure illustrates the distribution of changes in the implemented minimum wage and changes in the legislated minimum wage over time and states. Overall, we observe 166 increases in the implemented minimum wage and 62 legislative events from 2001 to 2012. 60 changes in the implemented minimum wage and 26 changes in the legislated minimum wage follow from federal minimum wage policy. The remainder follows from state-level policies.

2.3 Other data sources

We rely on several other data sources in our empirical analyses, and detail them in the relevant sections: data on employment and wages come from the Bureau of Labor Statistics (BLS) Quarterly Census of Employment and Wages (QCEW) files (see section 6.2) and the Current Population Survey (CPS) (see Figure B.1 and Appendix B); data on house prices are quarterly state-level series from the Federal Housing Finance Agency interpolated to monthly frequency using monthly division-level indices based on the Denton method; data on the share of labor costs and wholesale costs in grocery stores come from the BLS Annual Retail Trade Survey Consumption data (see section 6.2); consumption data come from the Consumer Expenditure Survey (CES) (section 6.2); and wholesale data from the annual BEA input-ouput tables (see section 6.2).

3 Main identification strategy

We estimate the price response to minimum wage increases by relating month-on-month store-level inflation rates to increases in the binding minimum wage and passage of minimum wage legislation at the state level. The identification strategy is based on the idea that, conditional on a set of controls and fixed effects, inflation in stores in states that did not experience a minimum wage hike or new legislation is a useful counterfactual for stores in states that did. Many papers studying the effects of minimum wages in the US apply variants of this identification strategy (see Allegretto et al., 2017). The high frequency of our price data allows us to estimate detailed temporal patterns of the effects before and after an event. We use a flexible first-differenced specification to capture the dynamics of the price response over time, as, e.g., proposed by Meer and West (2016) in the minimum wage context and similar to the specification commonly used in the international economics literature to study the pass-through of exchange rate variation (for example Gopinath et al., 2010):

$$\pi_{j,t} = \delta_j + \gamma_t + \sum_{r=-k}^{k} \beta_r \Delta imp_{s(j),t-r} + \sum_{r=-k}^{k} \alpha_r \Delta leg_{s(j),t-r} + \psi X_{j,t} + \epsilon_{j,t} \tag{3}$$

In this model, $\pi_{j,t}$ is the month-on-month inflation rate in grocery store j and calendar month t. The main exogenous variables of interest are the change in the logarithm of implemented and legislated minimum wages in the state $s(j)$ in which store j is located, which we denote $\Delta imp_{s(j),t}$ and $\Delta leg_{s(j),t}$, respectively. The coefficients β_r and α_r measure the elasticity of inflation with respect to minimum wage increases or legislation r months ago, or r months in the future in case r is negative. In our baseline estimation we control for time fixed effects γ_t and store fixed effects δ_j. Because our estimation is in first differences, the latter account for *trends* in stores' price levels.

The vector of controls $X_{j,t}$ includes the county-level unemployment rate and state-level house price growth. We include these control variables to absorb variation in grocery prices that is due to business cycles or the boom and bust in house prices (see Stroebel and Vavra, 2019). Yet, we show below that our results are very similar if we omit these controls.

We start by estimating the effects at legislation and implementation separately by omitting all terms related to either $\Delta imp_{s(j),t}$ or $\Delta leg_{s(j),t}$. However, in our preferred specification, we jointly estimate effects at legislation and implementation of minimum wage increases. The reason is the the separate estimates may capture the same variation in prices since legislation is often passed in the 9 months preceding implementation. We cluster our standard errors at the state level. The resulting standard errors are conservative and substantially larger than the standard errors that we would get if we clustered at the store level, for example.

While our estimates of equation 3 are in first differences, the estimates are best illustrated as the effect of minimum wages on the price level. We thus construct cumulative sums of β_r and α_r coefficients in the presentation of our results. We normalize the effect to zero in a baseline period two months before an event, and calculate the cumulative effect as $E_R = \sum_{r=-1}^{R} \beta_r$. We also summarize pre-event coefficients in a similar way. To be consistent with the normalization we calculate them as $P_R = -\sum_{r=2}^{-R-1} \beta_{-r}$. Our baseline measure of overall elasticities is E_4 and thus includes effects one month before to 4 months after an event.[13] We report E_4 separately for implementation of minimum wages and passage of legislation, as well as the sum of both.

An important choice in our estimation is the number of estimated lag and lead coefficients k. One constraint here is that minimum wage hikes generally occur in regular intervals, often within 12 months (see Table B.1). This implies that the event dummies become collinear if k gets larger.[14] A second constraint is that the store panel is not balanced. The more leads and lags we include, the more likely it is that changes in the underlying store sample may affect our estimates. In our baseline estimation, we settle on estimating the effect with $k = 9$. This is sufficient to show the short run impact of minimum wage increases on prices.[15]

The central concern with our estimation and identification strategy is the possibility of reverse causality. States with higher inflation rates could have more frequent and higher nominal minimum wage increases to avoid reductions in the real minimum wage. In this case inflation would cause minimum wage increases, rather than the other way around.[16] Although we view it as unlikely that legislators consider changes in state-level grocery price inflation within the few months relevant for our empirical analyses, we deal with this concern in our estimation in several ways. First, our main specification includes store fixed effects, which absorb differences in trend inflation between states. Second, due

[13] In principle, we could report E_k and include all lag coefficients. However, coefficients beyond 4 months out are typically close to zero and insignificant. In most specifications E_k is not significantly different from E_4 but substantially less precise.

[14] This implies that some observations lie, for instance, 8 months after the last and 4 months before the next minimum wage hike. In principle, we can disentangle the effects of the two events in such cases because many states do not have minimum wage increases before 2005 and after 2009, and because some states increase minimum wages only infrequently. However, our estimation strategy will not work in practice for large k, as the leads and lags become increasingly collinear.

[15] We present results for longer or shorter windows in robustness checks in Appendix Table B.6.

[16] A special case are minimum wage increases following from indexation. All states that practice indexation peg their minimum wage to national inflation rates. Changes in national inflation are absorbed by time fixed effects in our specification.

to the high frequency of our price data and the flexible estimation model, we can closely examine the timing of the effect, and any remaining differences in inflation trends around a minimum wage event would be easily detected in our pre-event coefficients. Third, we present estimates that only use variation due to changes in the federal minimum wage (see section 4.2). We view it as unlikely that federal lawmakers take into account regional inflation differences when setting the federal minimum wage.

4 The price response to minimum wage increases

4.1 Main results

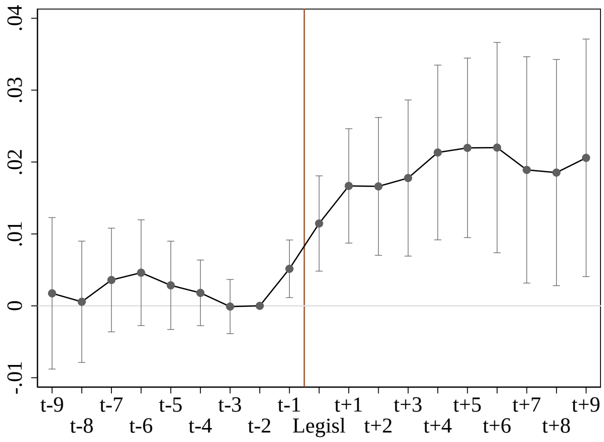
We start by using our main regression model (equation (3)) to estimate the effects of minimum wages on grocery prices at legislation and implementation separately. The dependent variable is the store-level month-on-month inflation rate. Figure 3a presents the estimated price effects at legislation. Reassuringly, the figure provides no evidence for significant movement in grocery prices in the months leading up to passage of minimum wage legislation. In the month that immediately precedes legislation, however, grocery stores start to increase prices. Prices continue to rise for 3 months. Overall, we estimate that the price elasticity of minimum wages at legislation of the hike is 0.021 and statistically highly significant.[17]

Figure 3b presents the results at the time of implementation of minimum wage increases. Our baseline estimate for the elasticity at implementation is comparable in size to the one for legislation. The figure points to a gradual increase in prices in the months leading up to implementation of a minimum wage increase. We show in section 5.2 that these significant pre-trends are driven by minimum wage events that are known long before implementation. They thus capture the effects at legislation for these events shown in Figure 3a. Hence, by the time the minimum wage has actually risen to the level set in

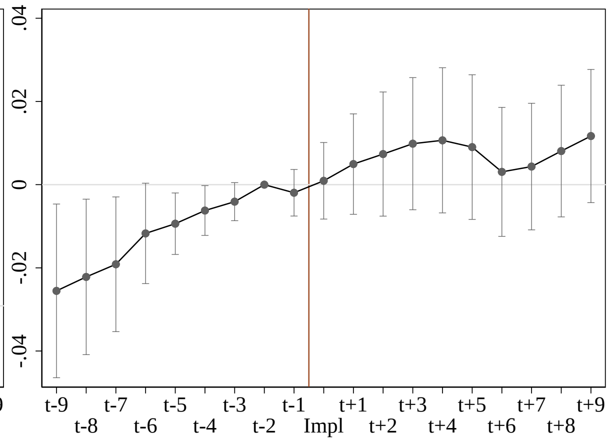
[17] We present our estimates of effects at legislation and implementation separately in Appendix Table B.4, as well as robustness checks that include division- and chain-time fixed effects.

the new legislation, the price adjustment appears to be more or less complete. We return to this evidence for forward-looking price setting of grocery stores in section 5.2.

Figure 3: Cumulative minimum wage elasticities of prices from separate estimation



(a) At legislation



(b) At implementation

Notes: The figures present the cumulative minimum wage elasticity of prices at grocery stores. Effects at legislation and implementation are estimated separately. The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The figures also show 90% confidence intervals of these sums based on SE clustered at the state level.

Column 1 of Table 1 presents the results of our preferred specification that jointly estimates the effects at legislation and implementation.[18] This specification accounts for the fact that the price effects at legislation and preceding implementation may reflect the same variation in prices. Again, the sum of pre-event coefficients ($\sum$ Pre-event) is close to zero and not statistically significant, thus validating our empirical strategy. Our preferred estimate of the price elasticity of minimum wages sums up all coefficients in the five months that follow legislation and implementation ($E_4^{leg} + E_4^{imp}$). This elasticity amounts to 0.036 and is statistically significant at the 5% level. It suggests that the average minimum wage increase in our sample—which we estimate to be +20%[19]—raises prices in grocery stores by 0.72% over three months at the time when legislation is passed. In this example, inflation would more than double during these 3 months relative to the sample average rate of 0.13%.

[18]See the corresponding graphs in Appendix Figure B.2.

[19]See Appendix Table B.1.

4.2 Robustness tests

We first show that these baseline estimates survive an extensive set of robustness checks.

Alternative specifications of our main empirical strategy. We present alternative specifications of our main empirical strategy for the joint estimation in Table 1. Column 2 shows that the estimated effects are similar if we weight each store by the number of products used to construct the stores' price index. Column 3 shows that none of our results depend on the inclusion of controls beyond time fixed effects. The inclusion of controls tend to improve the precision of the estimates. In Column 4, we remove store fixed effects which account for differential trends in stores' price levels in our baseline specification. Controlling for such trends might attenuate estimates of the minimum wage effects if the minimum wage affects the growth rate rather than the level of prices (Meer and West, 2016). Reassuringly, the estimated elasticities in Column 4 are very similar to our baseline results.[20] Column 5 shows that our baseline estimate is robust to the inclusion of state-calendar month fixed effects, which control more restrictively for possible differences in the seasonality of prices increases across states. In Column 6, we winsorize the inflation rates below the 1st and above the 99th percentile of the distribution to show that our results are not driven by outliers. Finally, columns 7 and 8 add census division-time and chain-time fixed effects, respectively. These fixed effects capture changing trend inflation within regions and grocery chains. They also largely control for possible effects of minimum wages on wholesale prices, as these would likely affect stores that are geographically close or belong to the same chain similarly. In those two cases, the price elasticities become indeed smaller, possibly because the fixed effects absorb increases in wholesale costs following minimum wage hikes (see section 6.3 for a discussion).

[20] We find similar overall price elasticities to the (implemented and legislated) minimum wage if we use a version of equation 3 in long first-differences (see Table B.5). Moreover, specifications that are differenced over longer time periods yield larger price elasticities. The incremental increase comes to an end after 9 months. These findings suggest that minimum wages temporarily affect the growth rate of prices, which supports our focus on inflation rates in the months around changes in minimum wages.

Table 1: Main results and robustness checks

	(1) Baseline	(2) Weighted	(3) No Controls	(4) No Store FE	(5) Seasonal	(6) Winsorized	(7) Div.-time	(8) Chain-time	(9) No Sales Filter
E_0^{leg}	0.011***	0.007**	0.011***	0.011***	0.009**	0.009***	0.013***	0.007*	0.018***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)	(0.004)	(0.006)
E_2^{leg}	0.015***	0.013**	0.015***	0.016***	0.015***	0.013***	0.019***	0.011**	0.026***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)	(0.005)	(0.009)
E_4^{leg}	0.019***	0.017***	0.019***	0.021***	0.019***	0.017***	0.020***	0.013**	0.031***
	(0.006)	(0.006)	(0.007)	(0.007)	(0.006)	(0.005)	(0.005)	(0.006)	(0.009)
E_0^{imp}	0.002	0.008	0.002	0.001	0.002	0.003	-0.003	-0.007	-0.004
	(0.006)	(0.007)	(0.006)	(0.007)	(0.006)	(0.006)	(0.006)	(0.006)	(0.008)
E_2^{imp}	0.012	0.016	0.012	0.011	0.011	0.012	0.000	-0.001	0.013
	(0.011)	(0.011)	(0.011)	(0.012)	(0.011)	(0.011)	(0.007)	(0.007)	(0.009)
E_4^{imp}	0.016	0.023*	0.017	0.016	0.018	0.015	0.006	0.002	0.022*
	(0.013)	(0.013)	(0.013)	(0.014)	(0.013)	(0.012)	(0.009)	(0.009)	(0.011)
Estimation Summary									
$E_4^{leg}+E_4^{imp}$	0.036**	0.040***	0.036**	0.036**	0.037**	0.033**	0.026**	0.016	0.053***
	(0.014)	(0.015)	(0.014)	(0.016)	(0.014)	(0.013)	(0.011)	(0.011)	(0.015)
$\sum$ All	0.046*	0.057***	0.046*	0.046	0.045*	0.040*	0.033	0.020	0.041
	(0.024)	(0.020)	(0.024)	(0.028)	(0.024)	(0.021)	(0.024)	(0.018)	(0.026)
$\sum$ Pre-event	0.010	0.016	0.010	0.008	0.008	0.004	-0.006	0.004	-0.004
	(0.016)	(0.013)	(0.016)	(0.018)	(0.016)	(0.014)	(0.019)	(0.013)	(0.018)
N	191,568	191,568	191,641	191,568	191,568	191,568	191,568	181,816	191,568
Controls	YES	YES	NO	YES	NO	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	NO	YES	YES	YES	YES	YES
Weights	NO	Obs	NO	NO	NO	NO	NO	NO	NO
Seasonality	NO	NO	NO	NO	YES	NO	NO	NO	NO
Div. time FE	NO	NO	NO	NO	NO	NO	YES	NO	NO
Chain time FE	NO	NO	NO	NO	NO	NO	NO	YES	NO

Notes: The dependent variable is the store-level inflation rate. Baseline controls are the unemployment rate and house price growth. The table lists cumulative elasticities E_R, R months after legislation or implementation. Column 1 is the result of joint estimation of effects at implementation and legislation in our preferred specification. (2) weights observation by the number of products (UPC) used to construct the store-level price index. (3) does not contain any control variables. (4) does not control for store fixed effects. (5) accounts for state-specific calendar month fixed effects. (6) uses a winsorized outcome (98% winsorization). (7) includes division-time FE, (8) chain-time FE. ((9) does not correct for temporary price changes. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * p < 0.1, ** p < 0.05, *** p < 0.01.

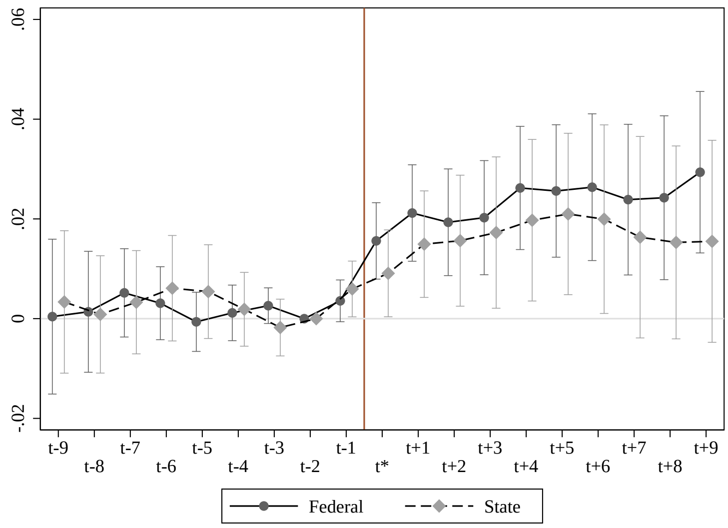
Including sales. Column 9 of Table 1 shows that the price elasticity is larger (0.053) if we use price indices that are not adjusted for temporary price changes. The reason, as we show in section 5.2, is that grocery stores reduce the frequency and the size of sales in the months around the legislation of minimum wage increases. By construction, however, sales represent a temporary deviation variation in prices. These results thus do not necessarily imply that our preferred estimate understates the permanent effect of minimum wages on prices. This would require that higher minimum wages decrease the frequency and size of sales permanently. To check the influence of sales on the price elasticity in the longer term, we present specifications in price levels (cf. column 4 and 6 of Table K.1) and long first-differences (cf. columns 4–7 of Table B.5) using price series that include sales. Both of these specifications are more robust to the large monthly price fluctuations caused by sales than our baseline model. The estimations suggest that our omission of sales does not lead to a downward bias in the estimated minimum wage elasticity of prices. Indeed, our preferred price elasticity of 0.036 is close to the elasticity in column 6 of Table K.1 estimated with prices that include sales.

Robustness to other specification choices. Appendix Table B.6 contains further robustness checks. Our results are robust to using only stores that we observe throughout the whole sample period (and hence are not driven by stores' entry or exit); to controlling for county level trends in the inflation rate; to changing the event window to $k = \pm 6$ or $k = \pm 12$ months; to excluding the Great Recession by focusing on the 2001–2007 period only; and if we only look at the effects of the first minimum wage hike in each state in our sample period, which represents an alternative method to address the fact that all states are treated multiple times in the sample period.

Our results are also robust to changing the level of analysis from the store to the state level (see Appendix C). Advantages of the state-level estimation are that the state panel is balanced and that the estimation can be extended to a longer panel without missing leads and lags due to store entry and exit. Reassuringly, the state-level estimates confirm our baseline estimates. Moreover, we find no evidence for differential trends in state-level prices between states with and without hike in the 15 months leading up to minimum wage legislation. These results speak against the concern that price inflation is the cause for minimum wage hikes rather than vice versa in the short estimation window relevant for our analyses.

Addressing reverse causality. Figure 4 presents a further robustness check that speaks against reverse causality. In particular, we estimate the separate effects for federal and state-level hikes by augmenting our baseline model with separate sets of leads and lags for events following from state and following from federal legislation. The response to new minimum wage legislation is similar in both magnitude and timing for federal and state-level minimum wage changes. While changes in state-level minimum wages could potentially be a response to local price increases, it is arguably very unlikely that price developments in particular states cause adjustments in the federal minimum wage.

Figure 4: Cumulative minimum wage elasticities of prices around federal- and state-level minimum wage legislation



Notes: The figure presents the cumulative minimum wage elasticity of prices at grocery stores around federal and state-level minimum wage legislation. The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The figures also present 90% confidence intervals of these sums based on SE clustered at the state level.

Testing inference and specification. We also conduct a placebo test to test our inference and our regression framework. In particular, we repeatedly match all stores of a state with the minimum wage series of a random state. The match is drawn without replacement from a uniform distribution including the correct match. For each trial, we estimate the cumulative elasticity in the five months after legislation and implementation, $E_4^{leg} + E_4^{imp}$, using equation 3. We present the distribution of 1,000 estimated elasticities in Appendix Figure B.6. Our price elasticity estimate of 0.036 at legislation and implementation lies above *all* the placebo estimates. Furthermore, the placebo estimates are centered around zero. The permutation test suggests that our results are not driven by misspecification or structural breaks in the inflation series that correlate with temporal patterns of minimum wage increases. Moreover, the results suggest that our statistical inference is quite conservative.

Price effects by bindingness of the minimum wage. We show that, as expected, price effects are larger where the earnings effects of the minimum wage are largest. We present two pieces of evidence that this is the case.[21]

First, using our main empirical strategy, we run separate regressions for stores located in "right-to-work" states and all other stores. Right-to-work laws prohibit mandatory union membership for workers in unionized firms, and weaken the position of unions. As a result, wages in grocery stores tend to be lower in those states and substantially more responsive to minimum wage hikes (see Addison et al. (2009), and our own estimates of average earnings elasticities with respect to the minimum wage in Table 4). Consistent with expectations, the price effects of minimum wage hikes at legislation and implementation are substantially larger in right-to-work states (see Appendix Figure D.1 and Table D.1).

Second, we show that our baseline results are robust to an alternative identification strategy that exploits within-state variation in wages. The idea is that a statewide mini-

[21] The full details of our approaches are presented in Appendix D.

mum wage hike affects stores that pay low wages more than stores that pay higher wages. While we cannot observe stores' wages, we can exploit the large geographic variation in average wages of grocery stores across counties within a state. We find that stores in higher wage counties exhibit significantly *lower* inflation than stores in the same state in low wage counties, in the quarter legislation is passed (see Appendix Table D.2).

4.3 Distributional consequences

This section studies the distributional consequences of the price effects of minimum wages in grocery stores. We conduct three analyzes. First, we assess whether grocery prices increase more in cheap compared to expensive stores. Second, we study whether there are differences in the price development of products that differ by their consumers' income. Third, we present estimates of the annual dollar value of price hikes following an increase in the minimum wage by income brackets. [22]

Price effects by store expensiveness. Columns 1–4 of Table 2 present estimates of our joint regression model when splitting the sample into cheap and expensive stores in two ways. We reduce the length of the estimation window to 6 months before and after an event in order to reduce the number of coefficients estimated from these smaller samples. We use a procedure implemented by Coibion et al. (2015) to calculate expensiveness relative to other stores in a state (columns 1 and 2) and county (columns 3 and 4), respectively.[23] While the estimated effects tend to be larger for cheap stores, the difference in the response of the two groups of stores are not statistically significant. This is a first

[22]We have also estimated the price effects separately by product category (see Appendix Figure B.3). We find that the price responses are largest for products for household products (such as laundry detergent, paper towels and facial tissues), alcoholic beverages and certain types of food (such as mayonnaise, yogurt and tomato sauce) – potentially because the demand for those products is less elastic.

[23]We first calculate the mean price during a year for each product and store. For each product, we then calculate the mean price in a county. We then calculate the deviation of each store from this price and aggregate deviations over all products sold in each store, weighted by the dollar revenue of the product. We only use products that are sold in at least 3 stores in a county and drop counties with less than 3 stores. Finally, we label stores that are on average more expensive than other stores in a county as expensive, and the remaining stores as cheap. The results are very similar if we measure expensiveness relative to other stores in a state rather than a county.

piece of evidence that speaks against the fact that the price effects of minimum wage in the grocery sector fall disproportionately on poor households.

Elasticities of income-specific price indices. Do price elasticities differ for products consumed by low- vs. high- income households? Taking advantage of the IRI consumer panel dataset, we construct separate price indices for low-, medium- and high-income households, and run our baseline regression for each index separately.[24] We find that the elasticities for the products consumed by the three types of household are almost identical: 0.030, 0.028 and 0.027 for low-, medium- and high-income households respectively (see Table E.5). This suggests that stores increase product prices across the board. They are also very close in magnitude to our baseline estimate. This is a second piece of evidence that speaks against the fact that the price effects of minimum wage in the grocery sector fall disproportionately on poor households. They also speak against demand shifts as a cause of the price response, a point we discuss in more details in section 6.4.

Magnitude of price hikes across the income distribution. To put the magnitudes of the price hikes along the household income distribution in perspective, we use the IRI consumer panel dataset to estimate the Equivalent Variation of the grocery price caused by a 20% minimum wage hike—which corresponds to the average legislated increase in the minimum wage in our sample (see Table B.1), and to a $1.24 minimum wage increase between 2001 and 2012. The Equivalent Variation is a first order approximation to the welfare cost of a price change, measured in US dollars. It assumes that households maximize utility and abstracts from second order effects reflecting the response to changes in relative prices.

A first-order approximation of the equivalent variation of a price change in the grocery sector j can be written as: $EV_j = E_{hj}\Delta P_j$, where E_{hj} denotes the mean household expenditure for goods sold in sector j for households in income bracket h.[25] We divide

[24]We present the full details of our analysis in Appendix E.

[25]An alternative interpretation of our EV measure is the cost for consumers to maintain consuming the same basket of goods after an x% price change. Our first-order approximation ignores some second-order

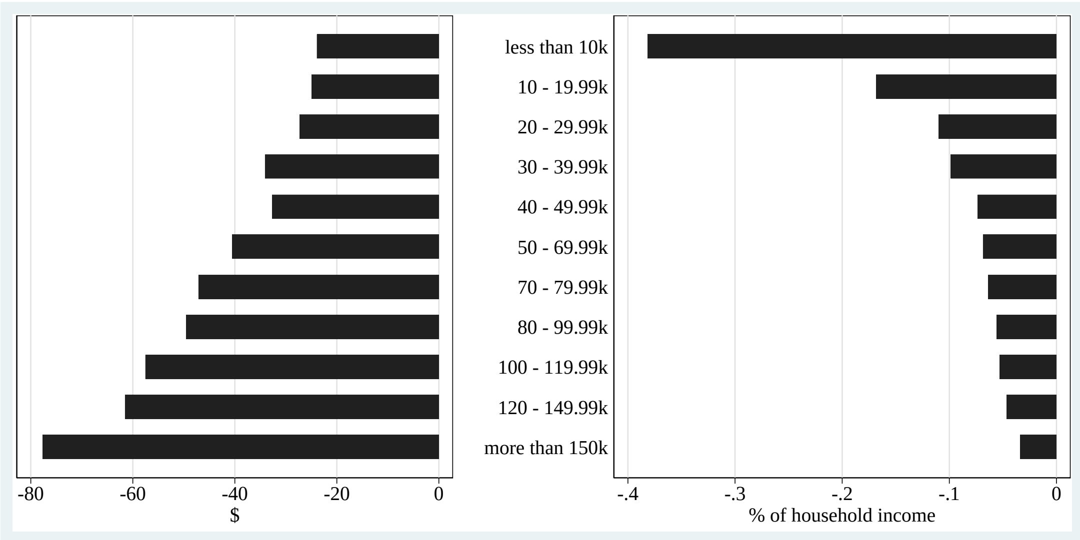
Table 2: Price effects of minimum wage by store characteristics

	(1) Expensive (state)	(2) Cheap (state)	(3) Expensive (county)	(4) Cheap (county)	(5) Regional chain	(6) Interregional chain
E_0^{leg}	0.007**	0.012***	0.005	0.013***	0.015***	0.007*
	(0.003)	(0.004)	(0.005)	(0.004)	(0.005)	(0.003)
E_2^{leg}	0.010**	0.016***	0.011	0.016***	0.018**	0.013**
	(0.005)	(0.006)	(0.007)	(0.006)	(0.008)	(0.005)
E_4^{leg}	0.013**	0.020***	0.014*	0.020***	0.022**	0.014**
	(0.006)	(0.007)	(0.007)	(0.007)	(0.009)	(0.007)
E_0^{imp}	-0.001	0.003	0.003	0.003	0.011	-0.007
	(0.008)	(0.006)	(0.009)	(0.006)	(0.007)	(0.006)
E_2^{imp}	0.002	0.015	0.006	0.016	0.020	0.002
	(0.016)	(0.010)	(0.017)	(0.010)	(0.013)	(0.010)
E_4^{imp}	0.008	0.019*	0.010	0.021*	0.030**	0.002
	(0.018)	(0.011)	(0.021)	(0.011)	(0.014)	(0.012)
Estimation Summary						
$E_4^{leg}+E_4^{imp}$	0.021	0.039***	0.024	0.042***	0.052***	0.017
	(0.019)	(0.014)	(0.022)	(0.014)	(0.017)	(0.014)
$\sum$ All	0.031	0.041*	0.037	0.049**	0.060**	0.013
	(0.024)	(0.022)	(0.026)	(0.023)	(0.029)	(0.021)
$\sum$ Pre-event	0.017	0.008	0.022	0.014	0.012	0.006
	(0.013)	(0.017)	(0.016)	(0.019)	(0.019)	(0.017)
N	47668	146374	30583	119234	111175	82867
Controls	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES

Notes: The table presents cumulative minimum wage elasticities of prices at grocery and drug stores along several heterogeneity dimensions. The dependent variable is the store-level inflation rate. Baseline controls are the unemployment rate and house price growth. The effects shown in the columns are based on the joint estimation (equation 3), estimated separately for each sample indicated in the column header. Columns 1–4 differentiate stores by their price level relative to other nearby stores in the state (columns 1 and 2) and county (columns 3 and 4), as desribed in the text. Columns 5 and 6 split chains into "interregional" and "regional" chains, as described in the text. The table lists cumulative elasticities E_R, R months after legislation or implementation. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * p $<$ 0.1, ** p $<$ 0.05, *** p $<$ 0.01.

the household income distribution into 11 household income brackets (increments of \$10K from 0 to \$100K and coarser categories above). We estimate the mean household expenditure for each of these categories using the expenditure data by income bracket provided in the Consumer Expenditure Survey (CES).[26] ΔP_j denotes the price change in sector j. Since we do not find differences in the price response of products consumed by different household income groups, we use our baseline elasticity (estimated jointly at legislation and implementation) of 0.036 for all household categories.

Figure 5: Equivalent Variation of price increase



Notes: The figure illustrate the Equivalent Variation (EV) of increasing all binding minimum wages in the US by 20%. See section 4.3 and Appendix E for a detailed description of the calculations involved. Figure 5 shows the EV for each income bracket in US dollars (left) and relative to mean household incomes (right).

Figure 5 presents the costs of price increases caused by minimum wage hikes, measured in US dollars and relative to household incomes. The dollar value of costs is increasing in household incomes. Since groceries are not an inferior good, this is to be expected. For households with incomes below \$10,000, the annual costs amounts to \$24 (or, equivalently, to \$19 a year for a \$1 minimum wage increase). The costs increase up to \$78 for households

terms that capture substitution to other products.

[26] We include expenditures for the CES categories Food at Home, Personal Care Products and Services, Household Supplies and Alcoholic Beverages as groceries.

with incomes above $150,000 (or, equivalently, up to $63 a year for a $1 minimum wage increase). Expressing the costs as a percentage of annual household incomes reveals the regressive impact of the price response in grocery stores. The costs make up 0.4% of annual income for households in the poorest bracket, and less than one tenth of that, i.e. 0.03% for households in the richest bracket. The numbers underlying Figure 5 are available in Appendix Table E.3a. We present a full welfare analysis of grocery price increases (and, as an additional exercise grocery and restaurants' price increases taken together) in Appendix E.

5 The anatomy of the price response

This section discusses a number of facts about the effect of the minimum wage on grocery store prices that are of interest to the macroeconomic literature on price setting.

5.1 Uniform pricing of grocery chains

We look at the heterogeneity of the price increase across regional and interregional chains to investigate whether grocery stores apply uniform pricing in response to minimum wage hikes. Regional chains are chains with stores in less than 3 distinct states on average. Interregional chains are those with stores in 3 or more states. We find that prices increase by 5.2% in response to a minimum wage hike in regional chains (Table 2, col. 5). While we find a statistically significant price response in interregional chains around legislation, the estimated sum of the price effects is smaller in interregional chains compared to regional chains (col. 6). This latter finding is consistent with Dellavigna and Gentzkow (2019), who find that US retail chains maintain pricing across stores as uniform as possible, making prices less likely to respond to local economic shocks.

Another implication of uniform pricing within grocery chains is that a minimum wage hike in a specific state may affect prices in stores within the same chain located in another state. We augment our baseline regression model with variables that would capture such spillovers across states within chains (see Table B.7). We find little evidence for spillovers if we estimate the regression using our baseline sample that includes all stores. This is different, however, if we restrict the sample to stores that belong to interregional chains. In these chains, we observe a disproportionate price increase in the quarter of the announcement of the minimum wage increase even in the stores of the chain that did *not* experience the specific minimum wage hike. The estimated spillover effects amount to roughly half of the direct (i.e. within-state) minimum wage effect on prices. There is also some evidence that price spillovers occur at implementation. These results suggest that the price effects of the minimum wage may cause (small) welfare losses for individuals that live in states where the minimum wage does not increase.[27]

5.2 Firms' forward-looking pricing decisions

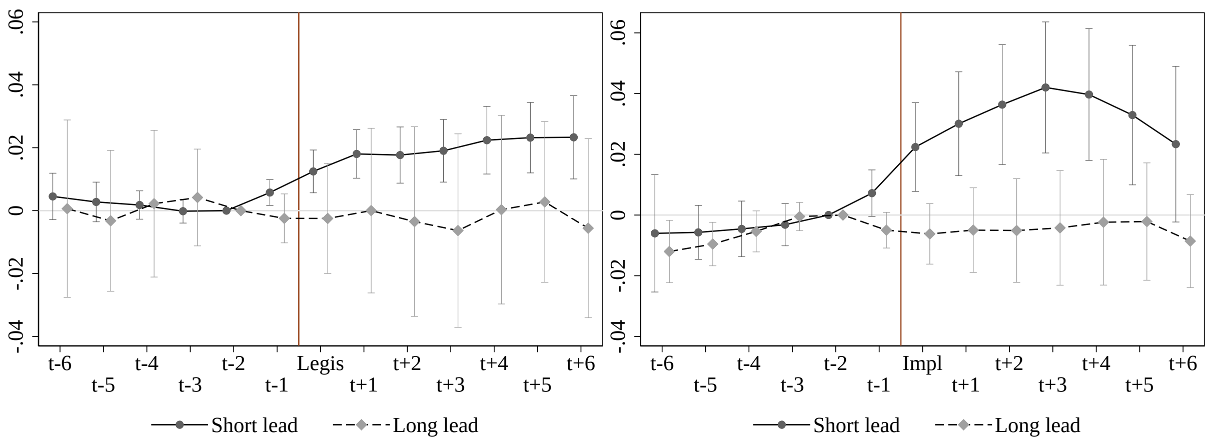
One striking result from our baseline regressions (see section 4.1) is that grocery stores appear to anticipate future cost increases by increasing their prices as soon as the minimum wage hike is announced (i.e. before the hike is implemented). In this section, we provide more details on this result and discuss how it relates to the macroeconomic literature on pricing behavior.

We first establish that retail stores seem to anticipate future cost increases by temporarily raising their markups between announcement and implementation. Using a similar methodology to study the dynamics of the wage response as for prices, we show that the earnings effect of the minimum wage hike is concentrated in the quarter when the hike is implemented. The price response at legislation thus reflects an anticipation of future wage increases, rather than premature compliance with future minimum wage laws (see Appendix Table F.1). Forward-lookingness in pricing decisions is consistent with price-

[27]One might be concerned that these results suggest an issue with our empirical strategy, namely that our control group of stores in states without minimum wage hikes may be partially treated. Reassuringly, however, the implied downward bias in our baseline specification is quantitatively small. We find no spillovers for our baseline sample that includes all stores (column 2 of Table B.7). Moreover, the bias is limited even among stores in interregional chains where the spillovers occur as can be seen by, e.g., comparing the estimated coefficient on $\Delta leg_{s(j),t+0}$ in columns 3 & 4 of Table B.7.

setting models with adjustment frictions, in which firms rationally consider the future as well as the present. These models include time-dependent models in which prices can only be changed in certain periods (see, e.g., Calvo, 1983) and state-dependent models (i.e. menu cost models), in which firms can adjust prices at a cost. Time-dependent models with a low probability of price change can feature a substantial degree of anticipation. In menu cost models, the speed of the price adjustment is more complex to predict[28] and the bulk of adjustment tends to happen close to implementation (see, e.g., Karadi and Reiff, 2019; Hobijn et al., 2006).

Figure 6: Firms' forward-looking pricing decisions: cumulative minimum wage elasticities



(a) Events at different timing, legislation (b) Events at different timing, implementation

Notes: Panel (a) shows the effects at legislation for legislation that is followed by implementation of a first increase in less than a year ("short lead") and legislation that is implemented further in the future ("long lead"). Panel (b) shows the effects at implementation for increases that are preceded by legislation within less than half a year ("short lead") and those whose legislation lies further in the past ("long lead"). The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The figures also show 90% confidence intervals of these sums based on SE clustered at the state level.

To further illustrate these anticipation effects, we look at events with different lead times between legislation and implementation of higher minimum wages. In panel (a) of Figure 6, we split minimum wage laws into those that are followed by a first increase in the minimum wage within less than a year and those with longer time between legislation

[28]The speed of the price adjustment depends on many parameters: the minimum wage increase; the menu cost; and other product-level shocks (see, e.g., Karadi and Reiff, 2019)

and implementation.[29] The figure provides evidence that prices respond at legislation when implementation happens shortly after legislation, but not when implementation is at least a year out. In panel (b) of Figure 6, we split minimum wage laws into those that are followed by an increase in the minimum wage within less than 6 months and those with longer time between legislation and implementation. The figure shows that prices rise at implementation only when there is a short lag (less than 6 months) between legislation and implementation. In contrast, there are no price effects around implementation in the case of minimum wage hikes that are known long in advance. Rather, there is some evidence for an increase in prices in the months longer before the hike. If stores have enough time to anticipate the increase in cost, they appear to increase prices before their labor costs actually increase. Both sets of results are consistent with the predictions of price-setting models with frictions, that adjustment should be quicker for increases that become known shortly before they are implemented.

Finally, Appendix Figure B.5 provides clear evidence in favor of an anticipatory pricing behavior by showing that prices increase 6 (2) months before implementation for events that were legislated exactly 6 (2) months before they are implemented.[30]

Next, Table 3 investigates the channels through which US grocery stores make their forward-looking pricing decisions. The table provides five main insights. First, grocery stores increase the frequency with which they adjust regular (i.e., sales-filtered) prices as a response to minimum wage increases (column 1). The increase in the frequency of price changes happens primarily through an increase in the frequency of price increases (column 2). The point estimates imply that a 10% minimum wage hike raises the weekly frequency of regular price increases by 0.014–0.038 percentage points in the quarters around legislation and implementation, roughly 1.5–3.5% relative to the mean of the frequency.[31] Second, there is no increase in the absolute size of price changes overall

[29] There are 50 legislative events with a "short" and 12 with a "long" lead time between legislation and the first hike. Increases resulting from indexation are excluded from this analysis.

[30] The story is different for hikes legislated 4 months before implementation: there is only a small, if any, price effect at the time of legislation (i.e. at month $t = -4$), but prices increase quite strongly after implementation.

[31] We compute frequencies of price adjustments at the weekly level for each product. We then aggregate

(column 4). These first two results are consistent with menu cost models but not with time-dependent models (see, e.g., Nakamura and Steinsson, 2008; Nakamura et al., 2018). Third, Columns 5 and 6 show that firms increase the size of increases and reduce the size of decreases in regular prices around legislation. Fourth, grocery stores reduce the frequency and the size of sales around legislation (columns 7 and 8). Relative to the mean of the outcome, the effects on sales are smaller than the effects on the frequency of regular price changes. Finally, we find no statistically significant evidence that the pass-through of products whose prices are frequently adjusted occurs closer to the implementation date than for prices with long duration (see Appendix Figure B.4), as would be predicted by standard menu-cost models. Rather, the price effects at legislation appear to be driven both by goods with stickier and less sticky prices.

across products using expenditure weights. The quarterly data represent an average over the weekly frequencies. For instance, the mean of 0.0204 in column 1 means that 2.04% of the regular prices are changed in an average week of a quarter. This implies that regular prices remain unchanged for 49 weeks in the estimation sample. See Appendix A for more details.

Table 3: Effects of minimum wage increases on frequency and size of price changes

	(1) Freq. change	(2) Freq. increase	(3) Freq. decrease	(4) Size change	(5) Size increase	(6) Size decrease	(7) Freq. sales	(8) Size sales
$\Delta leg_{s(j),t-1}$	0.0004	0.0010	-0.0006	0.0001	0.0001	0.0001	-0.0167	-0.0259***
	(0.0013)	(0.0008)	(0.0009)	(0.0006)	(0.0006)	(0.0008)	(0.0119)	(0.0076)
$\Delta leg_{s(j),t+0}$	0.0004	0.0014*	-0.0011	-0.0001	0.0007**	-0.0009**	-0.0187*	-0.0141
	(0.0011)	(0.0008)	(0.0007)	(0.0003)	(0.0003)	(0.0004)	(0.0102)	(0.0099)
$\Delta leg_{s(j),t+1}$	0.0008	0.0013	-0.0005	-0.0004	0.0000	-0.0008**	-0.0210**	-0.0200**
	(0.0014)	(0.0010)	(0.0007)	(0.0004)	(0.0005)	(0.0004)	(0.0102)	(0.0096)
$\Delta imp_{s(j),t-1}$	0.0033***	0.0030**	0.0002	0.0002	0.0006	-0.0003	0.0050	-0.0000
	(0.0011)	(0.0012)	(0.0009)	(0.0006)	(0.0007)	(0.0009)	(0.0186)	(0.0126)
$\Delta imp_{s(j),t+0}$	0.0035*	0.0038**	-0.0003	0.0000	0.0009	-0.0008	0.0092	0.0099
	(0.0020)	(0.0019)	(0.0013)	(0.0007)	(0.0006)	(0.0012)	(0.0201)	(0.0171)
$\Delta imp_{s(j),t+1}$	0.0005	0.0016	-0.0012	-0.0006	-0.0006	-0.0006	0.0091	-0.0143
	(0.0018)	(0.0013)	(0.0012)	(0.0008)	(0.0008)	(0.0009)	(0.0187)	(0.0136)
Observations	75278	75278	75278	75278	75278	75278	75278	75256
Controls	YES	YES	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES	YES	YES
Mean Dependent Variable	0.0204	0.0116	0.0087	0.0064	0.0067	0.0060	0.3440	0.1534

Notes: The table presents estimates of minimum wage effects on the frequency and size of price changes and sales. The estimates are derived from quarterly-level estimations of our joint regression model (equation 3) at the store level. $\Delta imp_{s(j),t}$ and $\Delta leg_{s(j),t}$ denote the percent change in the logarithm of implemented and legislated minimum wages, respectively, in quarter t and state $s(j)$ in which store j is located. The dependent variable in column 1 is the frequency of price increases in regular (i.e. sales-adjusted) prices, computed as the count of price changes between weeks of months at the product level and aggregated to the store level weighting each product equally. Similarly, the dependent variables in columns 2–6 are the frequency of price decreases (column 2), the size of price *increases* in sales-filtered prices (column 3), the size of price *decreases* in sales-filtered prices, the frequency of sales according to the sales filter by Kehoe and Midrigan (2015) (column 5) and the size of sales according to the sales filter. Baseline controls are the unemployment rate and house price growth. SE are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

6 The magnitude of the price pass-through

6.1 Benchmark model of the minimum wage elasticity of groceries' labor costs

In this section, we estimate the impact of minimum wage increases on grocery stores' cost with the aim of quantifying the degree of cost pass-through. We first clarify the assumptions required to estimate the impact of minimum wages on marginal cost.

We describe a general theoretical framework from which we derive our estimation procedure in Appendix G. We assume that grocery stores provide retail services using a production technology $F(L, X)$. F is homogeneous to some degree—including the possibility of non-constant returns to scale. X denotes the quantity of purchased merchandise. L is a composite input defined by a linear homogeneous aggregator over N different types of labor inputs $L_1, L_2, \ldots, L_N$ with wages $w_1, w_2, \ldots w_N$. The wages of these different types of workers may be affected by minimum wages differently[32]. An important implication of these assumptions is that the composition of worker types does not vary with the scale of the firm. Finally, we assume competitive labor markets[33].

Under these assumptions, the minimum wage elasticity of marginal cost at constant output equals:

$$\frac{\partial MC}{\partial MW}\frac{MW}{MC} = \frac{\overline{W}L}{C} \cdot \frac{\partial \overline{W}}{\partial MW}\frac{MW}{\overline{W}} \tag{4}$$

C denotes the total variable cost of a grocery store, and $\overline{W}$ denotes the average wage the

[32]We thus allow for the fact that workers earning wages above the minimum wage may also benefit from minimum wage hikes to some degree (Dube et al., 2015; Autor et al., 2016). For instance, table H.1 in the appendix illustrates that our results on the minimum wage share of groceries' costs would depend substantially on the wage cutoff we use to define minimum wage workers.

[33]We make this assumption because our evidence for positive price effects and no employment effects of minimum wages is generally inconsistent with monopsonistic labor markets (Aaronson et al., 2008). Monopsonistic labor markets have been brought forward as an explanation why minimum wages have limited effects on employment (Card and Krueger, 1995; Stigler, 1946). Our assumptions and our results are compatible with small or no disemployment effects if low-skilled labor is difficult to substitute with other factors—at least in the short run—and full price pass-through has small or no effects on sectoral output.

store pays. We can estimate this elasticity as the product of two factors: (i) the labor share in costs and (ii) the minimum wage elasticity of the average wage $\overline{W}$.[34] We provide an empirical calibration for those two factors in the following section. Note that these assumptions are only necessary to interpret our estimates as marginal cost pass-through. Alternatively our estimates can be interpreted—without any required assumptions—as average cost pass-through.

6.2 Empirical calibration

Labor share in costs. We estimate that the labor cost share of grocery stores is 16%, using the 2007 and 2012 BLS Annual Retail Trade Surveys. This estimate corresponds to the labor cost share in variable cost —which is the one that matters for price setting in the short run (see Appendix Table B.3). Labor costs include salaries, fringe benefits and commission expenses. Variable costs include labor costs, costs of goods sold and some smaller items such as transport and packaging costs.[35] We also note that the most important factor in grocery store costs is the cost of goods sold (83%).

Minimum wage elasticity of the average wage. We estimate the minimum wage elasticity of average earnings of grocery store workers using quarterly county-level data from the Quarterly Census of Employment and Wages (QCEW). QCEW employment and wage measures are the ones reported by employers in their Unemployment Insurance contributions. The QCEW files cover more than 95% of US jobs. We calculate average earnings as the ratio of total earnings of grocery store workers and grocery store employment. We assume that the elasticity of average earnings is equal to the elasticity of the

[34] The fact that the price response is related to the labor share is well-known in the literature (see, e.g., Hamermesh, 1993; Aaronson and French, 2007; Cahuc et al., 2014; Leung, 2020).

[35] Variable costs differ from total costs. In addition to variable costs, total costs include building and equipment costs (such as rents, utilities, depreciation and purchases of equipment), purchased services (such as maintenance, advertisement, etc.) and other operating expenses (such as taxes). Note that our estimate of labor cost share in variable cost does not include purchased services in the denominator. These services make up about 2% of total costs and include some tasks that are likely done by low-skilled workers, for example maintenance work. These costs may depend on minimum wages as well, but it is hard to determine to which extent.

average wage.[36] We restrict the data to the states and time period included in our price regressions. We then estimate standard state-level two-way fixed effects regressions that are often used to estimate minimum wage effects on employment in the US (see Allegretto et al., 2017, for a critical assessment):

$$\log \overline{W}_{c,q} = \gamma_c + \delta_q + \beta \log MW_{c(s),q} + Controls_{c,q} + \epsilon_{c,q} \tag{5}$$

Table 4 shows that we find significant positive effects of minimum wages on average earnings. This is also true if we control for state-specific linear time trends—an important sensitivity check for the two-way fixed effects model in the minimum wage context (Allegretto et al., 2017). Moreover, as we show in appendix H, the elasticity of earnings in grocery stores increases with the bindingness of a minimum wage hike.

Our baseline estimate for the labor cost elasticity in grocery stores is 11%. This is in line with what other papers have found, only slightly smaller than our estimate for the accommodation and food service industry, and larger than for retail trade as a whole (see columns 3–6 of Table 4).[37]

6.3 Implied cost pass-through rates

The combined estimates of the labor cost share and the minimum wage elasticities of the average wage allow us to compute pass-through rates using equation 4. Our baseline point estimate for the elasticity of cost is $0.16 \cdot 0.11 = 0.018$. We compute pass-through rates by dividing the price elasticity by the estimated cost elasticity. The results are shown in Table 5.[38] Our estimate for pass-through based on our baseline specification

[36] The two will be equal if there are no negative effects on employment and hours of low-wage workers. In the case of negative employment effects, the earnings elasticity will underestimate the wage elasticity. However, we do not find evidence for negative employment effects (see Panel B of Table 4, consistent with Addison et al. (2009)'s estimates who also use county-level QCEW data for the period 1990–2005). We also do not find evidence of a negative effect on the number of establishments (see Panel C of Table 4).

[37] Our baseline labor cost elasticities are somewhat smaller than the elasticities for the US retail sector estimated in Sabia (2009) using CPS wage data. They are larger than those estimated in Addison et al. (2009) for the 1990–2005 period. Our estimates are similar to those reported in Leung (2020) and Ganapati and Weaver (2017), who also use QCEW data for a similar time period.

[38] Full details for the calculations made in this section are available in Appendix I.

Table 4: Earnings and employment elasticities to the minimum wage in grocery stores, retail, and restaurants

	Grocery stores		Retail trade		Acc. and food services	
	(1) Baseline	(2) Trend	(3) Baseline	(4) Trend	(5) Baseline	(6) Trend
Panel A: Dep. variable: Labor cost per worker						
log MW	0.108**	0.083***	0.048*	0.038	0.151***	0.147***
	(0.043)	(0.027)	(0.026)	(0.024)	(0.024)	(0.025)
N	80,722	80,759	124,000	124,000	98,056	98,080
Only Right-To-Work states						
log MW	0.165***	0.159***	0.064	0.096	0.246***	0.238***
	(0.056)	(0.050)	(0.070)	(0.063)	(0.062)	(0.070)
N	40,385	40,385	71,583	71,583	56,322	56,322
Panel B: Dep. variable: Employment						
log MW	-0.010	0.089**	-0.002	-0.003	-0.042	-0.046*
	(0.048)	(0.036)	(0.027)	(0.017)	(0.033)	(0.027)
N	80,722	80,759	124,000	124,000	98,056	98,080
Panel C: Dep. variable: Number of establishments						
log MW	-4.30	-1.66	46.57	6.06	-25.51	4.29
	(3.98)	(3.96)	(36.85)	(14.22)	(24.58)	(14.37)
N	114,000	114,000	125,000	125,000	118,000	118,000
Controls	YES	YES	YES	YES	YES	YES
County FE	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES
State linear trends	NO	YES	NO	YES	NO	YES

Notes: The table shows elasticities to state-level minimum wages in the 2001–2012 period by industry, estimated using county-level panel data for 41 states used in our price regressions. The data are based on the QCEW. Retail trade corresponds to NAICS codes 44–45, grocery stores to NAICS code 4451, and accommodation and food services to NAICS code 72. The outcome in panel A is log average earnings by industry. The outcome in Panel B is the log employment in an industry, computed as the average employment in the three months in the respective quarter. The controls are the log of county population and the log of total employment in private industries per county. Standard errors are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

(joint estimation at legislation and implementation, panel A) amounts to 2.026. This number may seem large, but given the large standard errors on the insignificant elasticity estimate at implementation, we cannot reject the hypothesis that pass-through is equal to 1—the p-value on this test is 0.485. The pass-through ratios based on estimates including division-time (1.492, p-value: 0.553) and chain-time fixed effects (0.836, p-value: 0.748) are closer to and also not significantly different from one.

Table 5: Implied cost pass-through for various specifications

	(1) Baseline	(2) Division-time FE	(3) Chain-time FE
	A. Pass-through at legislation and implementation		
Implied cost pass-through	2.026	1.492	0.836
p-value PT=1	0.485	0.553	0.748
	B. Pass-through at legislation and implementation (incl. predicted effects on COGS)		
Implied cost pass-through	0.968	0.784	0.516
p-value PT=1	0.867	0.127	0.003

Notes: The table illustrates the implied cost pass-through. Pass-through at legislation and implementation is the minimum wage elasticity of prices 5 months after legislation E_4^{leg} plus the same elasticity at implementation E_4^{imp} relative to the estimated elasticity of marginal cost. p-values for a test of full pass-through are computed using standard errors for the pass-through ratio calculated using the Delta method.

However, the pass-through rates in panel A do not take into account that minimum wages increases may also increase the cost of goods sold (COGS) in grocery stores. Wholesale prices, in turn, may increase if minimum wage workers are employed in the production of grocery products. Due to the high share of COGS in grocery stores' cost—as shown in Appendix Table B.3, COGS make up about 83% of cost—even a minor increase in wholesale prices could matter for retail prices. Moreover, retail stores have been shown to be very responsive to changes in COGS (Eichenbaum et al., 2011; Nakamura and Zerom, 2010).

We cannot test directly whether minimum wages affect COGS as our data does not include wholesale cost. However, we can calculate an upper bound for this effect using input-output tables by assuming full pass-through of increases in labor costs into prices all along the production chain for each of the sectors producing groceries, similar as in MaCurdy (2015). Assuming that all workers that earn 110% (130%) of the minimum wage are affected by the minimum wage, we predict that a 10% increase in the minimum wage would increase the prices of COGS by 0.016% (0.024%) (see Appendix Table I.1). Hence, under the assumption of full-pass-through, price increases for COGS may indeed affect the marginal costs of grocery stores in a comparable magnitude as the direct effect through higher labor costs in grocery stores themselves.

Importantly, due to our DiD design, the estimated minimum wage elasticities of grocery prices only reflect the effects of higher prices for COGS to the extent that these occur locally. If wholesale groceries are highly tradable, price increases in COGS would affect all stores and pass-through would be absorbed in time fixed effects. We study the origin of groceries sold in different states. Using grocery wholesale-to-retail flows from the Commodity Flow Survey, we find that the majority of groceries sold in a state are delivered by wholesalers located in the same state or census division. As a consequence, it is likely that our estimates partly capture pass-through of increases in COGS.

Panel B of Table 5 thus shows pass-through rates that take into account effects of minimum wage increases on COGS. To calculate these pass-through rates, we assume that the major part of the price effect occurs in the state in which the minimum wage occurs and that the price pass-through along the production chain is the same as in the retail sector. The estimate for pass-through based on our baseline specification falls to 0.97 when we incorporate possible wholesale cost increases. The implied cost pass-through rate is significantly lower than 1 if we incorporate effects on COGS and control for chain-time fixed effects. Note, however, that the division-time and chain-time effects in columns 2 and 3 likely absorb at least part of the price effects of COGS already, so that incorporating this additional cost effect leads to a lower bound on the pass-through rate. For this reason, the estimated pass-through rates in columns 2 and 3 of panel B are possibly biased downward.

6.4 Demand increases: a discussion

So far, we have treated minimum wage increases as a cost shock to grocery stores. However, minimum wages also raise the incomes of low-wage workers, which may affect the demand for groceries. This demand may in turn also elicit a response of grocery prices. This view has been advocated in Leung (2020) and Alonso (2016), who find a positive impact of minimum wages on real grocery store revenues. In contrast, Aaronson et al. (2012) find no evidence for an impact of minimum wages on consumption of non-durables and services in households with minimum wage earners.

Our results also suggest that minimum wages have limited effects on grocery consumption. Appendix Table J.1 shows that we find no effect of minimum wages on quantities sold at or on revenues of grocery stores, neither at legislation nor at implementation. Note, however, that the estimates are not very precise, precluding us from strong conclusions on the magnitude of the effects.

Even if minimum wages affected the grocery demand for households with low-wage workers, there are a priori good reasons to be skeptical that minimum wage hikes lead to a substantial shift in market demand that would have a quantitatively important effect on prices. To see this, note that the role of demand in the price response to minimum wage increases is determined by three factors: First, minimum wages need to have a substantial effect on local *aggregate* incomes. Second, the market demand for groceries has to be responsive to changes in aggregate incomes. Third, grocery stores' prices have to be responsive to changes in local demand.

We expect small effects of minimum wages along on at least the first two dimensions. First, Dube (2019) shows that minimum wages increase incomes of low-income families with an elasticity of up to 0.5 after two years. He finds effects on incomes up to the 15th percentile of family incomes. Yet, in 2011, these families account for less than 2% of total incomes. The elasticity of aggregate incomes to the minimum wage would thus be at the order of $0.5 \cdot 0.02 = 0.01$. Second, the magnitude of the shift in individual demand associated with increasing income depends on the income elasticity of grocery demand. Products sold in grocery stores are typically necessities with income elasticities below one (see, e.g., Banks et al., 1997; Lewbel and Pendakur, 2009; Okrent and Alston, 2012). Any shift in individual demand is thus likely to be smaller than the underlying increase in income. Third, several estimates of grocery stores' supply curve suggest that prices are rather unresponsive to *temporary* changes in demand in the short run, even in the face of very large demand shifts (Chevalier et al., 2003; Gagnon and Lopez-Salido, 2014; Cavallo et al., 2014). More relevant to the study here, Stroebel and Vavra (2019) estimate that the elasticity of retail prices to a *permanent* house-price-induced demand shocks is on the order of 0.1–0.2. This is a relatively sizable effect on prices.[39] However, even taking the upper bound of that third elasticity, along with the upper bounds of the first two elasticities, lead us to estimate that the demand-side price elasticity of minimum wage increases in our study is at most $0.01 * 1 * 0.2 = 0.002$. This is 18 times smaller than our baseline price elasticity of 0.036. We conclude that it is unlikely that our baseline price elasticity is driven by demand side effects.

Finally, we note that the timing of our price response is not consistent with a demand-side effect on prices. Quantity responses would likely occur at the time the minimum wages is implemented and wages of workers actually increase (see Appendix F). The price increases, however, largely happens around legislation, and thus on average several months before household incomes increase.

7 Conclusion

In this paper, we study the effects of minimum wage increases on prices in grocery stores. We use scanner data to analyze the response to 166 minimum wage increases and 62 legislative events in the US from 2001 to 2012.

Our findings can be summarized by three key results. First, the minimum wage elasticity of prices is about 0.036. This means that the average minimum wage increase

[39] Stroebel and Vavra (2019) also document that the price response is mainly driven by markups rather than marginal costs, a result that does not align with ours.

from 2001 to 2012 (+20%) raises prices by 0.72%, i.e. that inflation more than doubles around the minimum wage hike. Our results are consistent with a full pass-through of cost increases to consumers. Second, we find that the response to minimum wage increases happens around the time of passage of legislation, rather than at the time of implementation of hikes. This result suggests that grocery stores set their prices in a forward-looking manner. The price increase occurs mostly through an increase in the frequency of price increases, consistent with menu cost models. Third, we show that prices rise as much for low-, medium- and high-income households. Since groceries make a larger share of low-income households' budgets, low-income households are hit the most by price increases. For households with income below $10,000, the annual costs associated with a $1 minimum wage increase is $19. Overall, consumers rather than firms seem to bear the cost of minimum wage increases in the retail sector.

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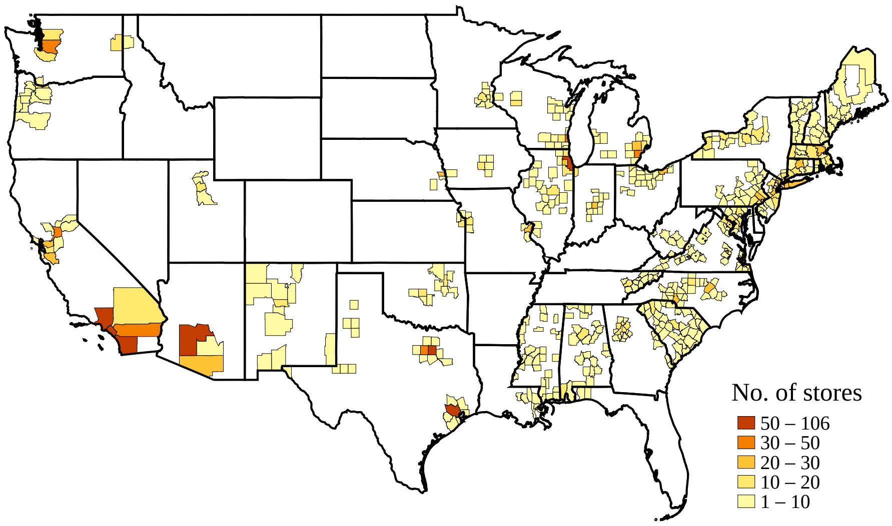
A Construction of store-level price series

Our empirical analysis is based on scanner data provided by the market research firm Symphony IRI. The dataset is described in detail in Bronnenberg et al. (2008). It contains weekly prices and quantities for 31 product categories sold at grocery and drug stores between January 2001 and December 2012. Stores report total revenue (TR) and total sold quantities (TQ) at the level of UPCs for each week. Figure A.1 shows the regional distribution of the stores in our dataset.

Figure A.1: Regional distribution of stores in IRI data across the US



Notes: Geographical distribution of stores in the IRI data. The map shows stores per county. Of the 3142 counties in the US, 530 (17%) are covered with at least one store in the IRI data.

In order to construct store-level price indices, we first calculate the average price of product i in grocery store j and week w from quantities and revenues:

$$P_{ijw} = \frac{TR_{ijw}}{TQ_{ijw}}.$$

We next calculate the average monthly price for each series and construct a geometric index of month to month price changes for each product category c in each store:

$$I_{cjt} = \prod_i \left(\frac{P_{ijt}}{P_{ijt-1}}\right)^{\omega_{ijy(t)}}. \tag{A6}$$

The weight $\omega_{ijy(t)}$ is the share of product i in total revenue of category c in store j during the calendar year of month t.[40] In a second step, we aggregate across different categories to create

[40] Price indices are often constructed using lagged quantity weights. Since product turnover in grocery stores is high, using lagged weights would limit the number of products used in the construction of our

store-level price indices and inflation rates:

$$I_{jt} = \prod_c I_{cjt}^{\omega_{cjy(t)}} \text{ and } \pi_{jt} = \log I_{jt} \tag{A7}$$

Again, the weight $\omega_{cjy(t)}$ is the share of category c in total revenue in store j during the calendar year of month t. Note that this approach does not take into account changes in the price level due to the introduction of new products, or due to reappearance of products at a new price after a stock-out, a feature shared by most price indices.

A common characteristic of retail scanner price data is that many price observations are missing when products are not sold temporarily, or enter or exit the sample. Since our product-level inflation measure is based on monthly averages, we implicitly assume that at the weekly level, the latent price of missing observations is equal to the average monthly price of observations we do observe. Our category level inflation index is a weighted geometric average of all non-missing inflation observations. As a result, when a product is not sold for one month or more, we implicitly assume that its inflation rate is equal to this category-wide average. A special case of missing observations are those before entry or after exit of a product or a store. Price dynamics at entry or exit may be different from those during normal times. However, we treat the timing of these events as independent from the timing of minimum wage increases, and do not attempt to correct the index for entry or exit.

An important characteristic of high frequency retail price data is that prices often change temporarily and return to their original level afterward. These movements, usually due to temporary sales, are large and affect the volatility of inflation rates at a monthly frequency. We thus apply an algorithm developed by Kehoe and Midrigan (2015) to determine "regular prices". Regular prices in our case are "permanent prices". Stores charge this price during long time periods, but often deviate from it during temporary sales. The regular price determined by the algorithm is based on the modal price for a product during a running window. For completeness, we reproduce a slightly edited description of the algorithm given in the web appendix to Kehoe and Midrigan (2015):

1. Choose parameters: $l = 2$ (size of the window: the number of weeks before and after the current period used to compute the modal price), $c = 1/3$ (=cutoff used to determine whether a price is temporary), $a = 0.5$ (=the share of non-missing observations in the window required to compute a modal price).

2. Let p_t be the price in week t and T the length of the price series. Determine the modal price for each time period $t \in (1 + l, T - l)$:

   - If the number of weeks with available data in $(t - l, \ldots, t + l)$ is larger than or equal $2al$, then $p_t^M = mode(p_{t-1}, \ldots, p_{t+l})$ and $f_t =$ the fraction of periods with available data where $p_t = p_t^M$.

index. We thus use contemporaneous weights.

- Else $f_t = .$ and $p_t^M = .$ (missing)

3. Determine the first-pass regular price for $t = 1, ..., T$:

   - Initial value: If $p_{1+l}^M \neq .$, then $p_{1+l}^R = p_{1+l}^M$. Else, set $p_{1+l}^R = p_{1+l}$.
   - For all other $t = l+1, \dots, T$: If $p_t^M \neq .$ and $f_t > c$ and $p_t = p_t^M$, then $p_t^R = p_t^M$. Else: $p_t^R = p_{t-1}^R$.

4. Make sure regular prices are updated at the right times. Repeat the following procedure l times (this adjusts the timing of regular price changes to the first occurrence of a new modal price).

   (a) Let $R = \{t : p_t^R \neq p_{t-1}^R \,\&\, p_{t-1}^R \neq 0 \,\&\, p_t^R \neq 0\}$ be the set of weeks with regular price changes

   (b) Let $C = \{t : p_t^R = p_t \,\&\, p_t^R \neq 0 \,\&\, p_t \neq 0\}$ be the set of weeks in which a store charges the regular price

   (c) Let $P = \{t : p_{t-1}^R = p_{t-1} \,\&\, p_{t-1}^R \neq 0 \,\&\, p_{t-1} \neq 0\}$ be the set of weeks in which a store's last week's price was the regular price

   (d) Set $p_{\{R \cap C\}-1}^R = p_{\{R \cap C\}}$. Set $p_{\{R \cap P\}}^R = p_{\{R \cap P\}-1}$.

Table A.1 reports features of price adjustments for the regular prices that our index is based on. Prices change with a median monthly frequency of 10.3% from 2001 to 2006 and 12.2% from 2007 to 2012. This implies a median duration of a price spell of 9.2 and 7.7 months, respectively. The median size of a price change is about 11.4% during the first half period of the sample, and 10.5% during the second half. The share of price increases in price changes is about 57% during the first half of the sample and 60% during the latter half. Price increases are smaller than price decreases. Finally, monthly inflation rates are lower during the first half of the sample compared with the second half. The monthly rates correspond to annualized inflation rates of 1% in the first and 1.8% in the second half of the sample. Overall, those numbers are in line with what other researchers have documented for our and other retail price datasets.[41]

Finally, following e.g. Stroebel and Vavra (2019), we also report the correlation between the change in prices between 2001 and 2012 using our store-level price index (regular prices) to the change in the metro area food-at-home price indices provided by the BLS for the set of MSAs for which we have overlapping data (BLS produces food-at-home CPIs for 27 metro areas, of which 19 overlap with locations in the IRI data). Appendix Figure A.2, shows there is a strong correlation (0.8) between changes in our price indices and those published by the BLS. In a recent paper, Cooper et al. (2020) look at the effects of minimum wages on prices using

[41]See Nakamura and Steinsson (2008) for CPI data from 1998 to 2005 or Midrigan (2011) for an alternative scanner data set from 1989 to 1997. Stroebel and Vavra (2019) construct state-level indices based on the same data used in our paper and find that inflation rates are lower than CPI inflation from the beginning of the data until 2007.

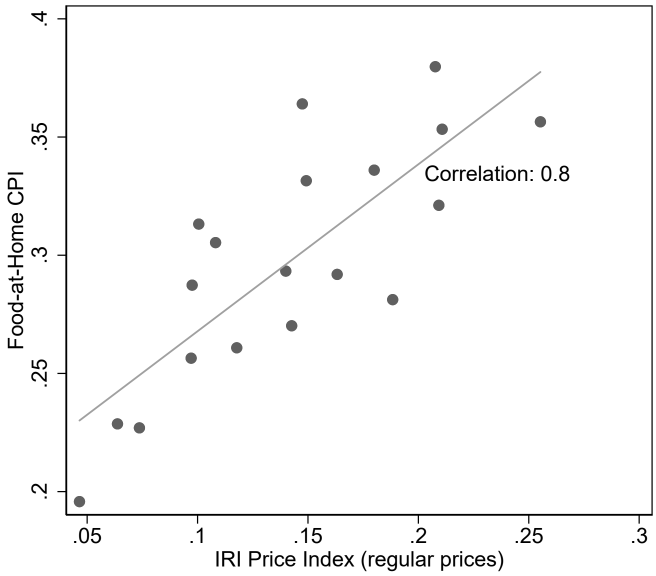
Table A.1: Features of regular prices

	2001-2006		2007-2012	
	Mean	Median	Mean	Median
Frequency of price change	0.117	0.103	0.132	0.122
Implied median duration	8.037	9.200	7.064	7.686
Frequency of price increase	0.067	0.060	0.078	0.074
Frequency of price decrease	0.050	0.040	0.054	0.043
Share of price increases in changes	0.605	0.576	0.623	0.602
Absolute size of price change	0.154	0.114	0.144	0.105
Absolute size of price increase	0.147	0.105	0.140	0.100
Absolute size of price decrease	0.184	0.146	0.166	0.132
SD log price	0.152	0.154	0.150	0.151
Monthly inflation	0.0007	0.0008	0.0016	0.0015

Notes: To construct these measures, we first calculate the frequency and size of price changes for each product in each store separately. For frequencies, we count changes and divide them by the number of observations for which we also observe a lagged price. We also calculate the standard deviation of the logarithm of prices within each state for each unique product. We then construct expenditure weighted means and medians for each category for the periods 2001 to 2006 and 2007 to 2012. Finally, we take expenditure weighted means over all 31 broad product categories. To summarize inflation rates, we take the weighted mean or median of our store-level inflation rates for the same periods.

BLS CPI series. Consistent with our findings, they do not find siginificant price responses for the food-at-home category at implementation.

Figure A.2: Correlation between IRI price index and CPI food at home



Notes: The figure shows a comparison of the change in prices according to our IRI data between 2001 and 2011 to the change in the metro area food-at-home price indices provided by the BLS for the set of MSAs for which we have overlapping data.

B Additional descriptive statistics and regression results

Summary statistics for minimum wage increases. These statistics mainly present the figures underlying Figure 2.

Table B.1: Summary statistics for minimum wage increases and minimum wage legislation

	Changes in implemented MW		Changes in legislation	
	Mean	SD	Mean	SD
Log size of increase	0.0816	(0.0560)	0.201	(0.116)
Events per state	4.049	(1.974)	1.512	(0.746)
Months to last event	13.86	(7.028)	23.32	(16.76)
Months legislation to first hike	15.65	(9.823)	8.742	(8.014)
Share federal hike	0.361	(0.482)	0.419	(0.497)
Share indexed hike	0.235	(0.425)		
Share 2001–2005	0.157	(0.365)	0.242	(0.432)
Share 2006–2008	0.542	(0.500)	0.742	(0.441)
Share 2009–2012	0.301	(0.460)	0.0161	(0.127)
Share January	0.458	(0.500)	0.452	(0.502)
Share July	0.434	(0.497)	0.0484	(0.216)
Number of Events	166		62	

Notes: The table lists descriptive statistics for our two main exogenous variables: Changes in implemented and legislated minimum wages. The legislated minimum wage is the highest future minimum wage set in current law. The data on state-level binding minimum wages is a combination of data from the Tax Policy Center, the US Department of Labor, and state departments of labor. We collected data on legislative events ourselves from media sources and legislative records.

Summary statistics on the importance of the minimum wage in the grocery sector. We present three stylized facts motivating our analysis of the price effects of minimum wages in the grocery sector.

The first fact is that groceries are an important factor in households' cost of living, particularly for poor households. Table B.2 presents the expenditure share of groceries using data from the Consumer Expenditure Survey (CES). We count the categories Food at Home, Household

Supplies, Alcoholic Beverages and Personal Care Products and Services as groceries. Measured this way, groceries make up about 11% of household expenditures on average. For households in the poorest quintile, groceries make up 14 to 15% of expenditures. For households in the richest quintile the share amounts to 9%.

Table B.2: Consumption expenditure shares on grocery stores' products by household income

	All house-holds	1st Quintile lowest	2nd Quintile	3rd Quintile	4th Quintile	5th Quintile highest
2001 - 2005	11.1	15.3	13.6	12.1	11.1	9.1
2006 - 2009	10.7	14.3	12.7	11.4	10.7	9.0
2010 - 2012	11.0	14.4	12.8	11.6	10.8	9.4

Notes: Data are from the Consumer Expenditure Survey. Grocery products include: Food at Home, Household Supplies, Alcoholic Beverages, Personal Care Products and Services. Shares are calculated for each year and quintile of household incomes and then averaged over all years in a period.

Table B.3 presents the second fact that we use to calculate the implied cost pass-through rate (see section 6.3): labor costs are an important part of the overall costs of grocery stores. The table shows the cost shares in total costs, variable costs and revenues of grocery stores (NAICS 4451) in 2007 and 2012. The table is based on a detailed breakdown of the costs of grocery stores available in the BLS Annual Retail Trade Survey for these years. Total costs include all operating expenses plus the Cost of Goods Sold (COGS). Variable costs comprise of labor costs, COGS, transport and packaging costs. The table illustrates that by far the most important factor in grocery store costs are the COGS. According to these data, the labor cost share in variable cost amounts to roughly 16%.[42]

The third fact is that a substantial share of grocery store employees are paid wages close to the minimum wage, and that this share has increased over time. We presented this fact in the Introduction. We provide here more details on how we performed our calculations. Using, data on hourly wages from the NBER files of the CPS MORG, Appendix Figure B.1 plots the distribution of wages in grocery stores relative to the local minimum wage. A large share of grocery store workers are paid wages at or close to the local minimum wage during all three periods. In the period when most of the minimum wage hikes in our sample happen (2006–2009), 21% of grocery store workers earn less than 110% of the minimum wage. Recent literature suggests that even workers with wages above the minimum wage may be affected by "ripple effects" of a hike (Autor et al., 2016; Dube et al., 2015), and as a result a large share of grocery store workers would likely be affected by minimum wage increases. At the end of our sample

[42]These labor shares do not include purchased services. These services make up about 2% of total costs and include some tasks that are likely done by low-skilled workers, for example maintenance work. These costs may depend on minimum wages as well, but it is hard to determine to which extent.

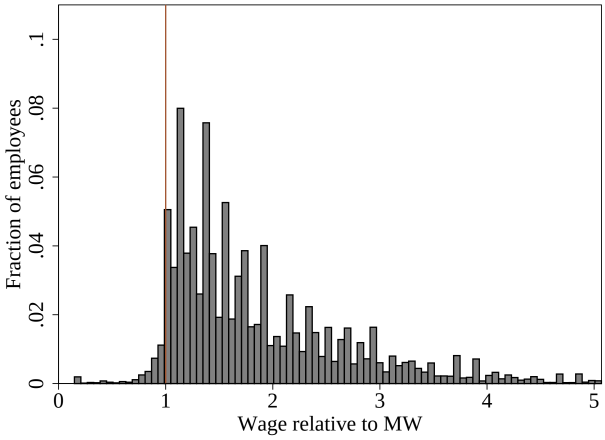
Table B.3: The cost structure of grocery stores

	Variable Cost			Fixed Cost		
	Labor Cost	COGS	Other Variable Cost	Buildings and Equipm.	Purchased Services	Other Operating Exp.
			Share in Total Cost			
2007	14.7	75.1	0.6	5.5	1.9	2.3
2012	14.1	75.4	0.6	5.4	1.8	2.7
			Share in Variable Cost			
2007	16.3	83.1	0.7			
2012	15.6	83.7	0.7			

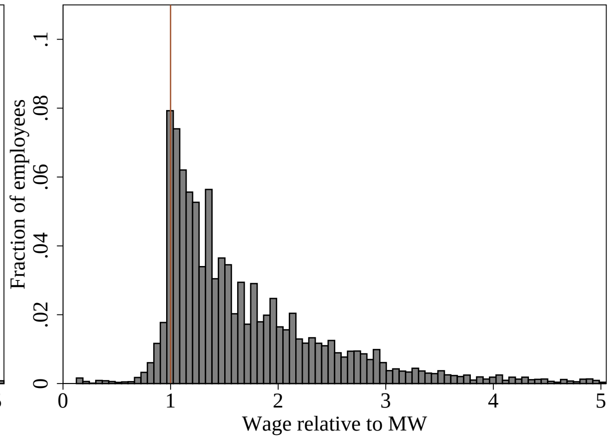
Notes: Data are from the BLS Annual Retail Trade Survey (ARTS). All numbers are in %. A breakdown of operating expenses into categories is published every 5 years. Labor Cost includes salaries, fringe benefits and commission expenses. Cost Of Goods Sold (COGS) is calculated as nominal annual purchases minus nominal year-on-year changes in inventory. Other Variable Cost includes transport and packaging cost. Buildings and Equipment includes rents, purchases of equipment, utilities and depreciation. Purchased Services includes maintenance cost, advertisement, etc. Other Operating Expenses includes taxes and the residual operating expenses category. We illustrate shares in total cost and in Variable Cost (which includes Labor Cost, COGS and Other Variable Cost). Estimates of the shares and SE in parentheses are based on Taylor expansions using the coefficients of variation published in the ARTS.

period, for instance, almost half of all grocery store workers earn less than 130% of the local minimum wage. As shown by Table H.1 in the appendix, the share of these workers in total hours worked in groceries amounts to approximately 40% in this period, and the share in total labor earnings to 25%.

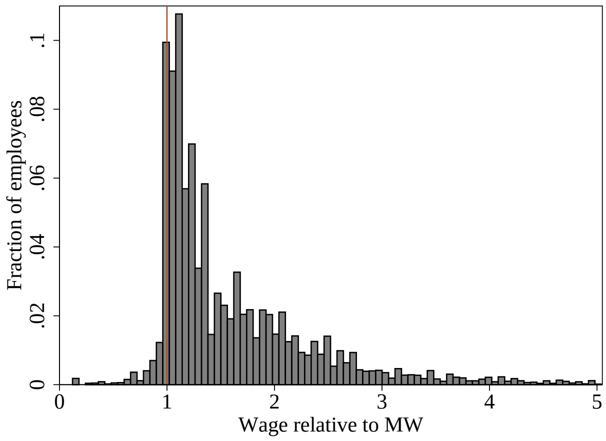
Figure B.1: The wage distribution in grocery stores relative to local minimum wages



(a) 2001–2005



(b) 2006–2009



(c) 2010–2012

Notes: The figure illustrates the wage distribution in grocery stores relative to local minimum wages. It is based on CPS MORG data for the sector "grocery stores" (NAICS 4451). Wages are computed using reported hourly wages for workers paid by the hour, and weekly earnings divided by weekly hours for other workers. All observations are pooled for the indicated periods. Distributions are calculated using CPS earnings weights. Wages below the local minimum may correspond to workers exempted from minimum wage laws (for example full-time students, workers with disabilities) or measurement error in the CPS survey.

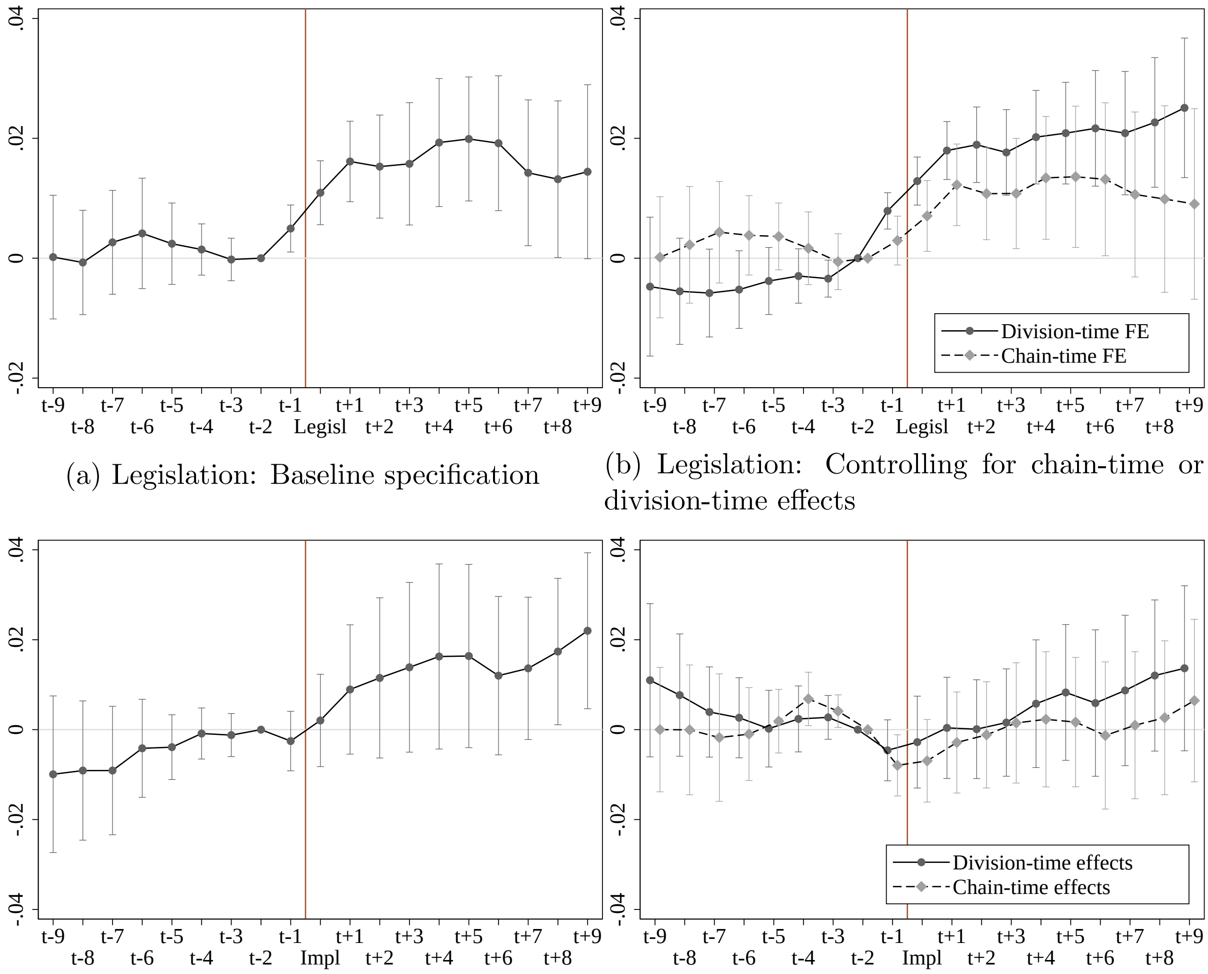
Additional regression results. We present several additional results and robustness checks using our main identification strategy (see section 4.1).

Table B.4: Cumulative elasticities for our baseline estimates

	(1) Baseline	(2) Div.-time	(3) Chain-time	(4) Baseline	(5) Div.-time	(6) Chain-time	(7) Baseline	(8) Div.-time	(9) Chain-time
E_0^{leg}	0.011***	0.009***	0.006				0.011***	0.013***	0.007*
	(0.004)	(0.003)	(0.004)				(0.003)	(0.002)	(0.004)
E_2^{leg}	0.017***	0.013***	0.011***				0.015***	0.019***	0.011**
	(0.006)	(0.004)	(0.004)				(0.005)	(0.004)	(0.005)
E_4^{leg}	0.021***	0.013**	0.013***				0.019***	0.020***	0.013**
	(0.007)	(0.006)	(0.004)				(0.006)	(0.005)	(0.006)
Estimation Summary									
$E_4^{leg}+E_4^{imp}$	0.021***	0.013**	0.013***	0.011	0.004	-0.002	0.036**	0.026**	0.016
	(0.007)	(0.006)	(0.004)	(0.011)	(0.007)	(0.007)	(0.014)	(0.011)	(0.011)
$\sum$ All	0.019	0.018	0.012	0.037**	0.015	0.010	0.046*	0.033	0.020
	(0.016)	(0.016)	(0.014)	(0.015)	(0.013)	(0.013)	(0.024)	(0.024)	(0.018)
$\sum$ Pre-event	-0.002	0.002	0.003	0.025*	0.008	0.011	0.010	-0.006	0.004
	(0.007)	(0.010)	(0.008)	(0.014)	(0.009)	(0.008)	(0.016)	(0.019)	(0.013)
E_0^{imp}				0.001	-0.003	-0.007	0.002	-0.003	-0.007
				(0.006)	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)
E_2^{imp}				0.007	-0.000	-0.004	0.012	0.000	-0.001
				(0.009)	(0.005)	(0.005)	(0.011)	(0.007)	(0.007)
E_4^{imp}				0.011	0.004	-0.002	0.016	0.006	0.002
				(0.011)	(0.007)	(0.007)	(0.013)	(0.009)	(0.009)
N	191568	191568	181816	191568	191568	181816	191568	191568	181816
Controls	YES	YES	YES	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Division time FE	NO	YES	NO	NO	YES	NO	NO	YES	NO
Chain time FE	NO	NO	YES	NO	NO	YES	NO	NO	YES

Notes: The table lists cumulative elasticities E_R, R months after legislation or implementation. The dependent variable is the store-level monthly inflation rate. Baseline controls are the unemployment rate and house price growth. Columns 1–3 show results of separate estimation of effects at legislation. Columns 4–6 show results of separate estimation of effects at implementation. Columns 7–9 show results of joint estimation of effects at implementation and legislation. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * p < 0.1, ** p < 0.05, *** p < 0.01.

Figure B.2: Cumulative minimum wage elasticities of prices from joint estimation



(a) Legislation: Baseline specification

(b) Legislation: Controlling for chain-time or division-time effects

(c) Implementation: Baseline specification

(d) Implementation: Controlling for chain-time or division-time effects

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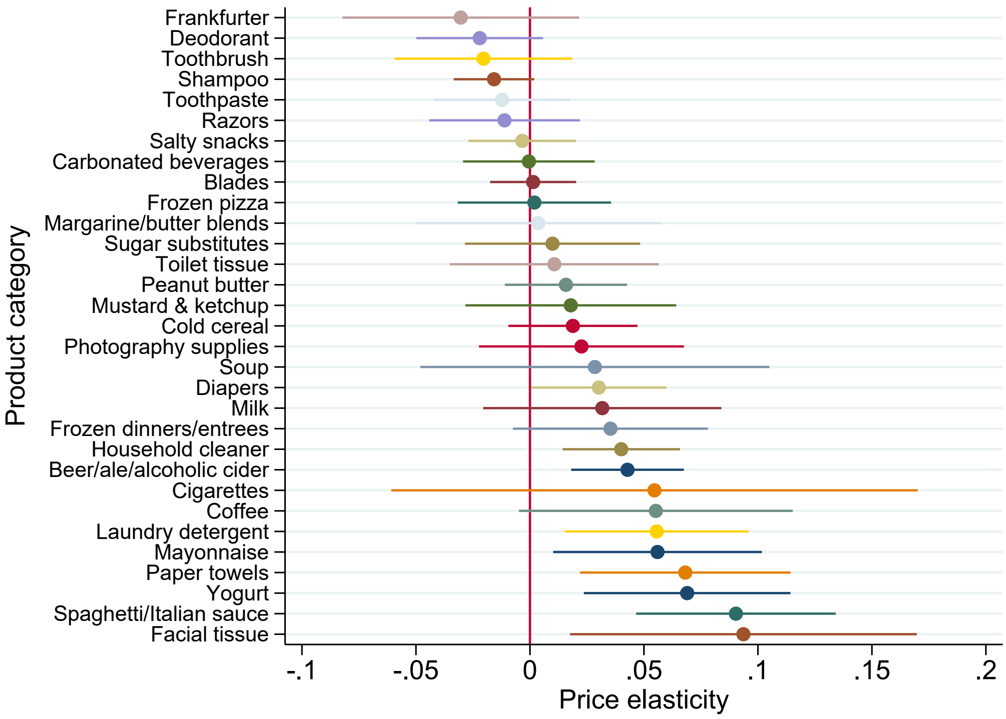
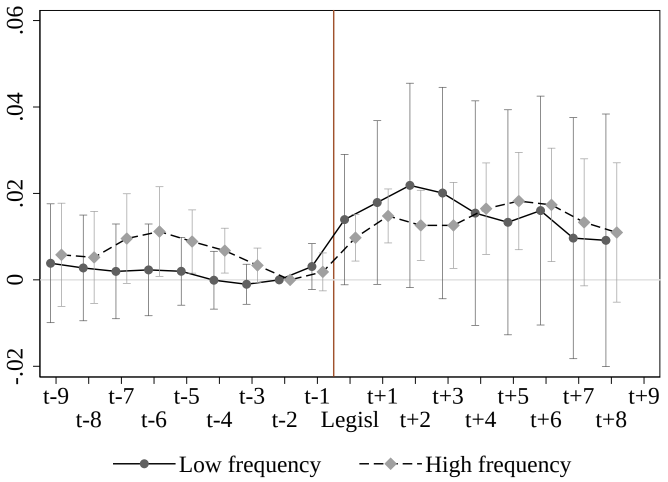
Notes: The figures present the cumulative minimum wage elasticity of prices at grocery stores. For each specification, the effects at legislation and implementation are estimated jointly from equation 3. Panels (a) and (c) show the cumulative elasticities at legislation and implementation estimated from the baseline specification. Panels (b) and (d) show elasticities estimated controlling for chain-time or division-time effects. The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The figures also present 90% confidence intervals of these sums based on SE clustered at the state level.

Figure B.3: Minimum wage elasticity of grocery prices by product group



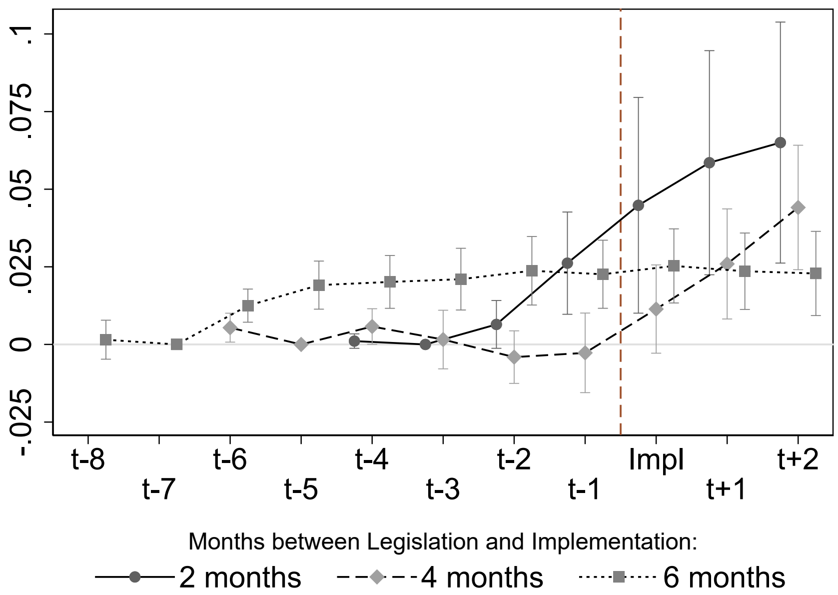
Notes: The figure shows the estimated minimum wage elasticity of grocery prices by product category. The estimates are derived from separate store-level regressions of our baseline joint estimation model (equation 3) for each price series. We focus on the total effect four months after implementation and four months after legislation ($E_4^{leg} + E_4^{imp}$). In each case, the dependent variable is winsorized at the first and 99th percentile to reduce the influence of outliers. Baseline controls are the unemployment rate and house price growth. The horizontal bars represent 90% confidence intervals derived from standard errors clustered at the state level.

Figure B.4: Cumulative minimum wage elasticities of prices at legislation by stickiness of product price



Notes: The figure presents the cumulative minimum wage elasticity of prices at grocery stores for products whose prices are adjusted frequently (prices with above-median frequency of adjustment) and infrequently (below-median frequency). Effects at legislation and implementation are estimated jointly (equation 3), but the figure focuses on the effects at legislation. The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The figure also presents 90% confidence intervals of these sums based on SE clustered at the state level.

Figure B.5: Cumulative minimum wage elasticities of prices at implementation for events with different time lag between legislation and implementation



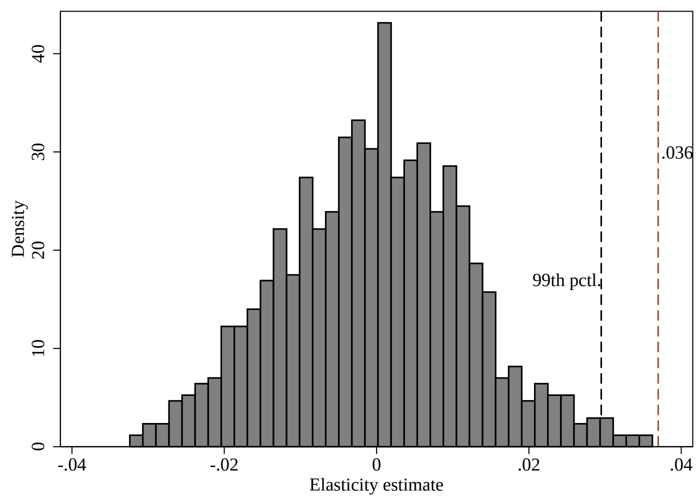
Notes: The figure presents the cumulative minimum wage elasticity of prices around implementation for minimum wage events with 2, 4, and 6 months between legislation and implementation. Effects at implementation for the three groups are estimated jointly using interaction terms between $\Delta imp_{s(j),t-r}$ and indicators for events with 2, 4, and 6 months lead time. The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The graph illustrates that the full price effect of hikes legislated 6 months before implementation occurred at the time of legislation. The results are different for hikes legislated 4 months before implementation: there is only a small, if any, price effect at the time of legislation (i.e. at month $t = -4$), but prices increase quite strongly after implementation. Hikes legislated two months before are an intermediate case. The figure also presents 90% confidence intervals of these sums based on SE clustered at the state level.

Table B.5: Price effects of the minimum wage using a specification in long first-differences

	(1) 2 months	(2) 4 months	(3) 6 months	(4) 10 months	(5) 14 months	(6) 10 months incl. sales	(7) 14 months incl. sales
Δimp	0.004	0.006	0.008	0.012	0.016	-0.003	-0.001
	(0.006)	(0.009)	(0.011)	(0.012)	(0.014)	(0.010)	(0.012)
Δleg	0.011***	0.012**	0.014**	0.014*	0.014	0.020***	0.018*
	(0.003)	(0.005)	(0.006)	(0.008)	(0.011)	(0.007)	(0.010)
Δ Unemp. rate	-0.000	-0.000*	-0.000*	-0.000	-0.000	-0.001**	-0.001*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.001)
Δ House prices	0.003	0.010	0.015	0.020*	0.024*	0.006	0.007
	(0.009)	(0.012)	(0.012)	(0.012)	(0.012)	(0.011)	(0.011)
Observations	230375	225124	220042	209975	200033	209975	200033
Sum of coefficients	0.015**	0.018	0.022	0.026	0.029	0.018	0.017
SE sum	(0.007)	(0.011)	(0.015)	(0.018)	(0.024)	(0.013)	(0.019)
Time FE	YES	YES	YES	YES	YES	YES	YES
Store FE	NO	NO	NO	NO	NO	NO	NO

Notes: The table presents the results of a long-first-difference regression of the form $\Delta p_{j,t} = \gamma_t + \beta \Delta imp_{s(j),t-r} + \alpha \Delta leg_{s(j),t-r} + \psi \Delta X_{j,t} + \epsilon_{j,t}$, where $\Delta p_{j,t} = \pi_{j,t}$ represents the store-level inflation rate excluding temporary sales (columns 1–5) and including temporary sales (columns 6 and 7). The first-difference (Δ) is taken over increasingly long time windows. The window is indicated in the column header. Baseline controls are the unemployment rate and house price growth. The "sum of coefficients" is the sum of Δimp and Δleg, and thus represents an overall estimate of the minimum wage elasticity of grocery prices. The regressions do not control for store FE. The results are very similar if we do but less precise. SE are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Figure B.6: Testing inference and specification: A placebo test



Notes: The figure presents the results of a placebo test in which we match all stores in a state with a random state's minimum wage series. Draws are without replacement and include the correct match. The histogram shows the distribution of elasticity estimates jointly estimated at legislation and implementation over 1,000 randomly matched samples. The mean elasticity estimate is close to zero. Our baseline estimate of the elasticity at legislation and implementation is 0.036 and outside the suggested 99% confidence interval.

Table B.6: Further robustness checks for joint estimation

	(1) Balanced panel	(2) County trends	(3) Short window	(4) Long window	(5) Pre-2008	(6) Only first	(7) Fixed weight
E_0^{leg}	0.008***	0.011***	0.011***	0.011**	0.012***	0.011**	0.018**
	(0.002)	(0.003)	(0.003)	(0.004)	(0.003)	(0.004)	(0.008)
E_2^{leg}	0.010**	0.015***	0.015***	0.017**	0.018***	0.020***	0.018*
	(0.004)	(0.005)	(0.005)	(0.007)	(0.005)	(0.006)	(0.010)
E_4^{leg}	0.013**	0.019***	0.018***	0.022**	0.024***	0.025***	0.024***
	(0.006)	(0.006)	(0.006)	(0.009)	(0.006)	(0.008)	(0.008)
E_0^{imp}	0.002	0.002	-0.000	0.003	0.000	0.002	-0.005
	(0.007)	(0.006)	(0.006)	(0.006)	(0.007)	(0.008)	(0.009)
E_2^{imp}	0.017	0.011	0.005	0.012	0.010	0.010	0.017
	(0.011)	(0.011)	(0.010)	(0.010)	(0.009)	(0.011)	(0.013)
E_4^{imp}	0.024*	0.015	0.008	0.015	0.017	0.017	0.017
	(0.013)	(0.012)	(0.012)	(0.011)	(0.011)	(0.013)	(0.014)
Estimation Summary							
$E_4^{leg}+E_4^{imp}$	0.038***	0.034**	0.026*	0.037**	0.040***	0.042***	0.041**
	(0.014)	(0.014)	(0.014)	(0.014)	(0.015)	(0.014)	(0.017)
$\sum$ All	0.024	0.045*	0.026	0.040	0.044**	0.031	0.032
	(0.024)	(0.022)	(0.021)	(0.039)	(0.021)	(0.021)	(0.027)
$\sum$ Pre-event	-0.003	0.010	0.006	-0.005	0.000	-0.011	-0.014
	(0.016)	(0.016)	(0.012)	(0.022)	(0.014)	(0.014)	(0.019)
N	73646	191568	206477	176822	108217	186151	189923
Controls	YES	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES	YES
County trends	NO	YES	NO	NO	NO	NO	NO

Notes: The dependent variable is the store-level inflation rate. Baseline controls are the unemployment rate and house price growth. The table lists cumulative elasticities E_R, R months after legislation or implementation. Column (1) focuses on stores that are present in all 142 periods of our sample. (2) adds county-specific time trends in the inflation rate. (3) uses an event window of length $k \pm 6$. (4) uses an event window of length $k \pm 12$. (5) restricts to the 2001–2007 periods. (6) computes the price effects only exploiting the first minimum wage hike in each state in the sample period. (7) is based on a price series that uses constant instead of time-varying product weights. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * p $<$ 0.1, ** p $<$ 0.05, *** p $<$ 0.01.

Table B.7: Price spillovers across state border within multi-store chains

	(1) All stores	(2) All stores	(3) Interregional	(4) Interregional
$\Delta leg_{s(j),t-1}$	0.001	-0.000	-0.008	-0.010*
	(0.005)	(0.005)	(0.005)	(0.005)
$\Delta leg_{s(j),t+0}$	0.010***	0.011***	0.013**	0.016***
	(0.004)	(0.003)	(0.005)	(0.005)
$\Delta leg_{s(j),t+1}$	0.003	0.001	0.004	0.001
	(0.002)	(0.002)	(0.003)	(0.003)
$\Delta imp_{s(j),t-1}$	-0.001	0.002	-0.011	-0.011
	(0.006)	(0.007)	(0.008)	(0.009)
$\Delta imp_{s(j),t+0}$	0.008	0.009	0.000	0.004
	(0.009)	(0.008)	(0.008)	(0.007)
$\Delta imp_{s(j),t+1}$	0.002	0.003	0.000	0.004
	(0.005)	(0.005)	(0.005)	(0.007)
Δleg_{t-1}^{chain}		-0.005*		-0.003
		(0.003)		(0.003)
Δleg_{t+0}^{chain}		0.004		0.009**
		(0.003)		(0.004)
Δleg_{t+1}^{chain}		-0.005*		-0.003
		(0.002)		(0.003)
Δimp_{t-1}^{chain}		0.009		-0.001
		(0.007)		(0.007)
Δimp_{t+0}^{chain}		0.006		0.006
		(0.005)		(0.008)
Δimp_{t+1}^{chain}		0.005		0.007*
		(0.004)		(0.004)
Observations	75278	75278	31898	31898
Controls	YES	YES	YES	YES
Time FE	YES	YES	YES	YES
Store FE	YES	YES	YES	YES

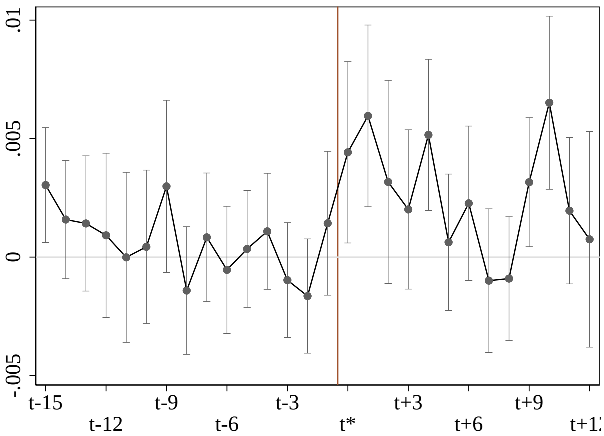
Notes: The table analyzes whether minimum wage hikes in one state affect prices in stores within the same retail chain in another state. The dependent variable is the store-level inflation rate of regular prices aggregated to quarterly frequency. The sample covers 2001–2012. We estimate the effects at implementation and legislation jointly. Column 1 shows the results of our baseline joint regression model (equation 3) for quarterly data. In columns 2 and 4, the model is extended with variables capturing possible across-state spillovers within chain. Δimp_t^{chain} is the average growth rate of the minimum wage stores in the same chain, located in another state, that experience an increase in the minimum wage in quarter t. Minimum wage increases within chains are weighted by the number of stores present in the IRI data within the same chain. Δleg_t^{chain} is an analogous variable for the growth rate in the legislated minimum wage. Columns 3–4 are restricted to stores of "interregional" chains, defined as chains with stores in more than 3 states in the data. Baseline controls are the unemployment rate and house price growth. SE are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

C Results using state-level price series

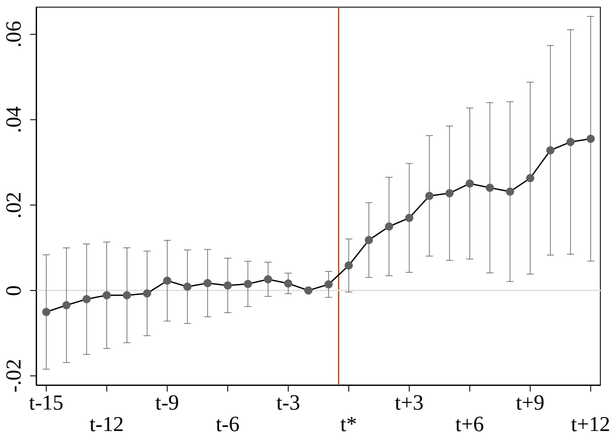
In this section, we conduct an analysis of the response of prices at the state level instead of at the store level. Our construction of state-level price indices largely follows Stroebel and Vavra (2019). One advantage of the state-level compared to our baseline store level estimation is that the state panel is balanced and that we can extend the estimation to a longer panel without missing leads and lags due to store entry and exit.

Table C.1 presents the estimation results for the baseline specifications using the state panel data set. The results confirm our baseline estimates, both in terms of timing and magnitude of the effect. The estimated elasticity, jointly estimated at legislation and implementation is 0.032 (i.e. very close to 0.036 in our preferred specification). The effects at legislation (from a separate regression) amounts to about 0.02 and there are no significant estimates around implementation of hikes. Figure C.1 shows the estimated effect on price inflation (panel a) and on the price level (panel b) if we allow the event window to span more than one year before and after the event, focusing on the effects at legislation. The figures provide no evidence for differential trends in the 15 months leading up to the legislation of a minimum wage hike.

Figure C.1: State level estimates of the price effects of the minimum wage around the time of legislation, using an extended event window



(a) Effect on the inflation rate



(b) Effect on the price level

Notes: The figure presents estimates using state level price indices and an extended event window of $k = -15$ to $k = 12$, focusing on the minimum wage effects at legislation. The dependent variable is the state-level month-on-month inflation rate. The panel on the right presents the estimates of α_r and the left panel their cumulative sum over the 24 month panel. Each panel also shows corresponding 90% confidence intervals based on SE clustered on the state level. The controls included are time and state FE, local unemp. rate and house price growth.

Table C.1: State-level estimations

	(1) Legisl	(2) Legisl	(3) Impl	(4) Impl	(5) Joint	(6) Joint
E_0^{leg}	0.005	0.006*			0.004	0.005
	(0.004)	(0.003)			(0.004)	(0.003)
E_2^{leg}	0.013*	0.014**			0.010	0.012*
	(0.007)	(0.005)			(0.007)	(0.006)
E_4^{leg}	0.019**	0.020***			0.016*	0.016**
	(0.008)	(0.007)			(0.009)	(0.007)
E_0^{imp}			0.008	0.008	0.007	0.008
			(0.007)	(0.006)	(0.007)	(0.006)
E_2^{imp}			0.009	0.011	0.009	0.011
			(0.010)	(0.009)	(0.010)	(0.010)
E_4^{imp}			0.013	0.015	0.013	0.016
			(0.012)	(0.011)	(0.012)	(0.012)
$E_4^{leg}+E_4^{imp}$	0.019**	0.020***	0.013	0.015	0.029*	0.032*
	(0.008)	(0.007)	(0.012)	(0.011)	(0.017)	(0.016)
$\sum$ All	0.018	0.021	0.044**	0.048**	0.051*	0.057**
	(0.015)	(0.013)	(0.022)	(0.019)	(0.026)	(0.023)
$\sum$ Pre-event	-0.003	-0.002	0.024**	0.026**	0.014	0.016
	(0.006)	(0.006)	(0.010)	(0.010)	(0.010)	(0.011)
N	5330	5330	5330	5330	5330	5330
Controls	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES
State FE	YES	YES	YES	YES	YES	YES
Weights	NO	Var	NO	Var	NO	Var

Notes: The dependent variable is the state-level inflation rate. Baseline controls are the state unemployment rate and house price growth. The table lists cumulative elasticities E_R, R months after legislation or implementation. Estimations with "Var" weights use the inverse of the variance of the state-level price series as weight to account for the fact that inflation series in states with few stores are more noisy. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

D Price effects by bindingness of the minimum wage

In this appendix, we present two pieces of evidence that grocery prices are larger where the earnings effects are larger. These pieces of evidence reinforce our main baseline results.

Price effects in Right to work states vs. not Right to work states. We split our sample between states with and without so-called "right-to-work" (RTW) laws. RTW laws prohibit mandatory union membership for workers in unionized firms, and weaken the position of unions. Compared to states without RTW laws, states with such laws exhibit lower unionization rates, laxer labor market regulations in general, and wages in grocery stores tend to be lower. Addison et al. (2009) find that earnings in grocery stores are substantially more responsive to minimum wages in RTW states. Our own earnings regressions, presented in Table 4 also suggest that the minimum wage has more bite in grocery stores located in RTW states. Hence, one may expect grocery prices to be more sensitive to minimum wages in these states.

Our results, presented in Figure D.1, are in line with this expectation. While the effect at legislation is of comparable magnitude in stores in RTW and non-RTW states, prices also increase substantially and statistically significantly at implementation in RTW states. In fact, stores in RTW states are the only subgroup analyzed in which we found evidence for a price effect at implementation, i.e. at the point in time when labor costs actually increase. Taking the effects at legislation and implementation together, the price effects of minimum wage hikes are substantially larger in RTW states.

Price effects in low-wage vs. high-wage counties. Our main identification strategy uses variation in increases in the legislated or implemented minimum wage across states to identify the effect on prices. In this section, we employ an alternative identification strategy which exploits that a statewide minimum wage hike affects stores that pay low wages more than stores that pay higher wages. Similar strategies have been used in the literature studying the employment effects of minimum wages (Card and Krueger, 1994, for example).

To exploit the differences in the bite of a given state-level minimum wage hike across counties, we compute the difference between the actual average quarterly salary in grocery stores and the full-time equivalent minimum wage salary using the QCEW. We then estimate the interaction between local inflation and this relative wage level for different time periods around minimum wage legislation and implementation. The specification for the effects at legislation is presented in equation D8:

$$\pi_{j,q} = \delta_j + \gamma_{t,s(j)} + \sum_{r=-kq}^{kq} \alpha_r \Delta leg_{s(j),q-r} \times wage_{c(j),q-r} + \psi X_{j,t} + \epsilon_{j,t} \tag{D8}$$

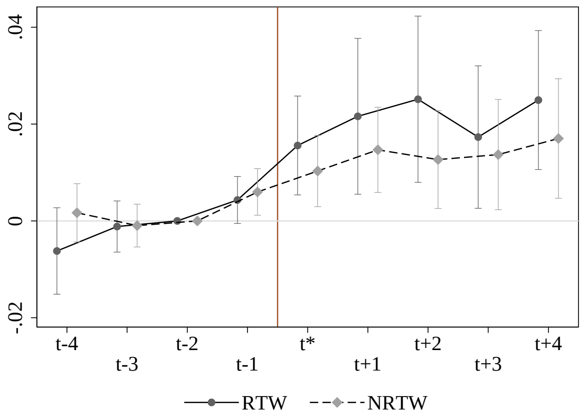
The α_r coefficients in this specification capture the extent to which prices of stores in low-wage counties react more (or less) to a given minimum wage hike than prices of stores in high-wage counties in the quarters around an increase in the minimum wage. In the case of legislation,

Table D.1: Price effects in Right-to-work and non-Right-to-work states

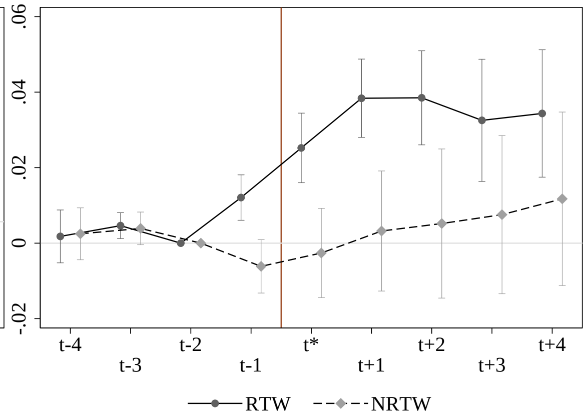
	(1) RTW	(2) No RTW	(3) RTW	(4) No RTW	(5) RTW	(6) No RTW
E_0^{leg}	0.007	0.009*			0.015*	0.010**
	(0.012)	(0.005)			(0.007)	(0.004)
E_2^{leg}	0.020	0.011			0.025*	0.012
	(0.021)	(0.007)			(0.014)	(0.007)
E_4^{leg}	0.028	0.015*			0.023	0.017*
	(0.021)	(0.008)			(0.016)	(0.008)
E_0^{imp}			0.024***	-0.001	0.023*	-0.002
			(0.008)	(0.007)	(0.013)	(0.007)
E_2^{imp}			0.037***	0.007	0.048***	0.008
			(0.008)	(0.011)	(0.015)	(0.012)
E_4^{imp}			0.033**	0.013	0.038**	0.015
			(0.012)	(0.013)	(0.018)	(0.014)
$E_4^{leg}+E_4^{imp}$	0.028	0.015*	0.033**	0.013	0.061**	0.032**
	(0.021)	(0.008)	(0.012)	(0.013)	(0.027)	(0.015)
Sum of coeff	0.059	0.010	0.068***	0.025	0.124**	0.027
	(0.065)	(0.020)	(0.012)	(0.017)	(0.050)	(0.024)
Sum of placebo	0.013	-0.005	0.028	0.011	0.046	-0.006
	(0.037)	(0.010)	(0.016)	(0.015)	(0.035)	(0.019)
N	79891	156329	79798	156329	79798	156329
controls	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES

Notes: (1), (3) and (5) include stores in Right-to-work states. (2), (4) and (6) include only stores in non-Right-to-work states. The dependent variable is the store-level inflation rate. Baseline controls are the unemployment rate and house price growth. The table lists cumulative elasticities E_R, R months after legislation or implementation. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * p < 0.1, ** p < 0.05, *** p < 0.01.

Figure D.1: Effects for stores in RTW and non-RTW states



(a) Elasticity at legislation



(b) Elasticity at implementation

Notes: The figures present the cumulative minimum wage elasticity of prices at in states with or without Right-to-work (RTW) laws. 17 states in our sample have RTW laws. Effects at legislation and implementation are estimated jointly. We estimate the effects for a smaller estimation window and omit controls, because the lower number of states in the split samples limits the number of state clustered standard errors we can estimate. The estimated coefficients are summed up to cumulative elasticities E_R as described in section 3. The figures also show 90% confidence intervals of these sums based on SE clustered at the state level.

we use the wage at the time legislation is passed as the initial wage ($wage_{q-r}$). In the case of implementation, we use the wage two quarters before implementation as the initial wage ($wage_{q-r-2}$) to make sure that the initial wage is not yet affected by minimum wage increases. Because there is variation in wages within a state, we can include state-time fixed effects that absorb all statewide developments that could potentially drive both minimum wage and grocery price increases.

Table D.2 presents the estimation results. We find that stores in higher wage counties exhibit significantly *lower* inflation than stores in the same state in low wage counties in the quarter legislation is passed. We find no significant relationship between inflation and initial wages in any other quarter around legislation, nor in the quarters around implementation of higher minimum wages. Our estimates suggest that a 10% lower initial wage increases inflation in the quarter legislation is passed by about 0.3%. The effects at legislation are robust to the inclusion of chain-time fixed effects. Overall, these results corroborate our main findings.

Table D.2: Interaction between price response and initial wage in a county

Dep. variable: Store inflation	(1) Baseline	(2) Chain-time	(3) Baseline	(4) Chain-time
		Legislation		
$\text{wage}_q \times \Delta\text{leg}_{q-1}$	-0.012	0.005		
	(0.010)	(0.010)		
$\text{wage}_q \times \Delta\text{leg}_q$	-0.028**	-0.031**		
	(0.014)	(0.013)		
$\text{wage}_q \times \Delta\text{leg}_{q+1}$	0.003	0.004		
	(0.013)	(0.010)		
		Implementation		
$\text{wage}_{q-2} \times \Delta\text{imp}_{q-1}$			-0.026	-0.006
			(0.035)	(0.028)
$\text{wage}_{q-2} \times \Delta\text{imp}_q$			0.012	0.036
			(0.035)	(0.023)
$\text{wage}_{q-2} \times \Delta\text{imp}_{q+1}$			-0.016	0.010
			(0.028)	(0.025)
		Estimation Summary		
Observations	84741	84503	84748	84512
Controls	YES	YES	YES	YES
Store FE	YES	YES	YES	YES
State time FE	YES	YES	YES	YES
Chain time FE	NO	YES	NO	YES

Notes: The dependent variable is the store-level inflation rate. This specification is estimated at quarterly frequency. Baseline controls are the unemployment rate and house price growth. wage is the log county-level average weekly wage in grocery stores relative to the state minimum wage. The listed coefficients are the interaction between minimum wage increases and the local wage at legislation or 2 quarters prior to implementation. SE are clustered at the county level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

E The impact of price increases in grocery stores on household welfare

We compare in this appendix the welfare costs of price changes in sector j to the predicted gain in nominal incomes resulting from wage increases. We note that this analysis is partial and does not take into account any other potential costs and benefits of minimum wage hikes, most importantly the price response in other sectors (besides grocery stores first, and besides grocery stores and restaurants) that employ minimum wage workers.[43]

We also note that the core of this welfare analysis is based on predictions of regular (i.e. excluding sales) price increases as opposed to transacted (i.e. including sales) price increases, in order to be consistent with our preferred price elasticity of 0.036 (see Table 1, column 1). However, to fully capture the effect of minimum wages on the cost-of-living, it may be appropriate to take into account increases in actual prices paid by consumers – that include sales (i.e. price elasticity of 0.053, see Table 1, column 9). We report the results of our welfare analysis when using transacted price series in Table E.4. Overall, as we detail below, the two sets of estimates are close to each other.

In what follows, we illustrate static welfare gains and losses based on a hypothetical increase of all binding minimum wages in the US by 20% (i.e. \$1.24 in our sample). Our preferred price elasticity predict that this would trigger a price increase in groceries by 0.4%.[44]

The overall dollar value of the welfare gain of a household can be expressed as:

$$\Delta U_h^{USD} = \Delta Y_h - \sum_j E_{h,j} \Delta P_j \tag{E9}$$

Here, ΔY_h denotes the mean USD increase in household incomes in income bracket h, and the product $E_{hj}\Delta P_j$ represents the Equivalent Variation of a price change in sector j.

The cost of price increases. We provide details on how this cost is calculated in section 4.3 of the paper. Our estimates of the costs of price increases caused by a 20% minimum wage hike, measured in US dollars and relative to household incomes appear in Figure 5, and in Appendix Table E.3a below. In brief, we find that the poorest households – earning less than \$10,000 a year

[43] We also assume in this exercise that: i) there is no statistically significant negative effect of minimum wage hikes on the number of jobs in the grocery stores sector in the short-run – a fact we document in Appendix Table I.2 using QCEW data over 2001-2012. This finding is consistent with the most recent evidence on the employment effects of minimum wage increases in the US (see e.g., Cengiz et al., 2019); ii) we assume minimum wage hikes have no statistically significant negative effect on the number of hours worked; iii) our estimates are presented in the short-run and therefore do not take into account that minimum wage hikes may lead to substantial capital-labor substitution in the longer-run (see e.g., Neumark and Wascher, 2008); iv) finally, our analysis does not take into account that prices increase after announcement while income increases after implementation, and that for a few months consumers pay more without any income gain.

[44] We assume here that the price elasticity is the same for all types of households, along the income distribution. We find empirical evidence that this is the case using the IRI consumer panel data (see Table E.5 below).

– face a \$24 (i.e. 0.4% of their annual income) increase in their grocery expenditures following a \$1.24 minimum wage increase (i.e. a 20% increase). By constrast, the richest households – earning more than \$150,000 a year – face a \$78 (i.e. 0.03% of their annual income) increase in their grocery expenditures. For a \$1 minimum wage increase, we find that the poorest (respectively the richest) households face a loss of \$19 (resp. \$68) after the price increase in gocery products.

When looking at the price response in transacted (i.e. including sales) prices (as opposed to regular (i.e. excluding sales) prices), we show that the poorest (respectively the richest) households face a \$35 (i.e. 0.56% of their annual income) (resp. a \$114 (i.e. 0.05% of their annual income)) increase in their grocery expenditures following a 20% minimum wage increase (see Table E.4).

The benefits of nominal wage increases. We now discuss how the costs of the price response relate to the first order effect of increasing nominal incomes for each household income bracket. We predict the mean increase in household incomes ΔY_h for each income bracket based on the March 2011 joint distribution of wages, hours worked per week, and weeks worked during the last year. Throughout this exercise, we assume that minimum wage increases have no effect on employment. The welfare effects are thus based on an upper bound on the benefit side, and would be lower if employment effects were negative.

We use the wage and weekly hours distribution during March 2011 available for the CPS monthly outgoing rotation group (MORG). We combine the MORG with the CPS Annual Socioeconomic supplement (ASEC) collected each March, which contains information on annual Household incomes and the number of weeks worked during the previous year.[45] For every person i in the MORG, we calculate the distance to the local binding minimum wage $W_i/MW_{s(i)}$. We then construct a counterfactual labor income as follows:

$$\widehat{Y}_i^L = \begin{cases} W_i \cdot 1.2 \cdot hours_i \cdot weeks_i, & \text{if } \frac{W_i}{MW_{s(i)}} \leq 1.1 \\ W_i \cdot \left(1 + 0.2\frac{1.3 - \frac{W_i}{MW_{s(i)}}}{1.3-1.1}\right) \cdot hours_i \cdot weeks_i, & \text{if } 1.1 < \frac{W_i}{MW_{s(i)}} < 1.3 \\ W_i \cdot hours_i \cdot weeks_i, & \text{if } \frac{W_i}{MW_{s(i)}} \geq 1.3 \end{cases} \tag{E10}$$

This calculation assumes that wages below 1.1 times the local minimum wage are increased

[45] We use the ASEC and MORG files provided by the NBER. We first calculate wages for the March 2011 MORG. For workers paid by the hour, we use reported hourly wages. For workers not paid by the hour, we use weekly earnings divided by usual weekly hours to calculate the hourly wage. We then merge the March 2011 ASEC to the March 2011 MORG. For every person i in the MORG, we calculate the distance of the hourly wage to the local binding minimum wage $W_i/MW_{s(i)}$. We then construct a counterfactual labor income as described in equation E10. We set hours and wages to zero for all workers that are not coded as "non-self-employed workers for pay". When the $weeks_i$ variable is missing, but weekly earnings and annual labor income is observed, we impute weeks based on this information and cap it at 52. If we cannot calculate labor income for one household member, we exclude the entire household from the analysis.

Table E.1: Summary statistics for different income brackets

	MORG labor income	ASEC labor income	Labor income share	Wage	Hours	Weeks worked	HH members	MW share	No of HH
less than 10k	2.1	1.5	26.7	12.7	4.7	6.9	1.7	5.8	947.0
10 - 19.99k	6.6	5.5	37.1	11.2	9.1	14.1	1.8	8.4	1764.0
20 - 29.99k	13.3	12.9	52.0	12.4	13.0	20.8	2.1	8.7	1737.0
30 - 39.99k	22.5	22.2	64.4	14.5	16.9	26.2	2.2	7.8	1486.0
40 - 49.99k	29.5	29.7	66.8	15.6	18.9	28.6	2.3	6.4	1287.0
50 - 69.99k	40.9	44.9	76.0	17.0	21.6	32.7	2.6	6.9	2199.0
70 - 79.99k	55.4	59.5	80.2	20.4	24.5	36.3	2.5	6.5	953.0
80 - 99.99k	63.8	72.7	81.7	21.2	25.3	37.1	2.7	6.5	1386.0
100 - 119.99k	79.2	90.4	83.3	24.4	26.2	37.5	2.8	6.1	1053.0
120 - 149.99k	91.2	109.0	82.4	26.8	26.4	39.0	2.8	4.0	892.0
more than 150k	124.5	186.2	82.5	37.3	25.1	38.2	2.9	4.3	1258.0

Notes: MORG labor income is equal to $hours \times wage \times weeks$. Wage and hours are from the MORG, weeks from the ASEC. ASEC labor income is annual labor income reported in the ASEC. The Labor income share is the share of labor in total household income (both from ASEC). Wages, Hours and Weeks worked are unweighted averages over household members, then averaged over households using HH weights.

by 20%, and that wages between 1.1 and 1.3 times the local minimum wage increase by a linearly declining factor. This is in line with ripple effects documented in Dube et al. (2015). We calculate the predicted increase in labor income $\Delta Y_i^L = \widehat{Y}_i^L - Y_i^L$ for each individual. We then sum the increase over all household members. Finally, we calculate the average predicted increase in household income for each income bracket using the ASEC household sampling weights.

In Table E.1, we report some additional statistics in order to cross-check our nominal incomes calculations, and conclude they are reasonably well fitted. We first compare annualized labor earnings based on the March 2011 MORG and reported weeks worked to actual reported labor income in the ASEC. Our calculation fits reported earnings quite well for households earning between $20,000 and $70,0000 a year. The annualized measure is larger than reported labor earnings for poorer and smaller for richer households. Two factors could explain this discrepancy. First, labor market conditions were improving in March 2011 after the through of the recession in 2010. Hours and wages of poor households could thus be higher in March 2011 than during 2010. Furthermore, the discrepancy for rich households could be due to differences in top-coding between the MORG and the ASEC. In addition, we present summary statistics on

Table E.2: Characteristics of minimum wage workers in different income brackets

	Relation					Mean			
	Female HH	Male HH	Child of HH	Other family	Not family	Hours	Weeks worked	Age	Female
less than 10k	37.7	13.3	9.1	1.8	38.1	24.7	28.6	31.0	0.7
10 - 19.99k	34.9	21.1	8.6	5.2	30.2	31.4	41.2	34.5	0.6
20 - 29.99k	34.3	15.0	13.6	7.2	29.9	30.6	42.9	37.4	0.6
30 - 39.99k	33.2	15.3	17.3	5.3	28.9	29.6	40.7	37.4	0.6
40 - 49.99k	31.9	9.5	21.5	10.6	26.5	30.3	37.8	31.8	0.6
50 - 69.99k	27.3	18.6	34.5	5.3	14.3	29.2	42.7	32.8	0.6
70 - 79.99k	27.9	15.1	37.8	3.9	15.2	29.6	40.3	31.5	0.6
80 - 99.99k	21.3	18.2	39.2	7.8	13.6	29.0	40.8	32.6	0.4
100 - 119.99k	21.0	9.7	56.4	8.0	4.9	27.0	37.7	27.7	0.5
120 - 149.99k	12.9	8.4	62.6	3.8	12.2	23.6	36.8	27.0	0.5
more than 150k	15.1	6.8	71.4	3.5	3.2	26.8	40.0	26.0	0.4

Notes: The table breaks down minimum wage workers by relationship to the reference person in their household. Minimum wage workers are all workers earning less than 110% of the local minimum wage. Data from MORG (wages) and ASEC (for income brackets).

wages, hours, weeks worked and the size of households in different brackets. Second, we present in Table E.2 summary statistics of minimum wage workers in the different brackets. There are some important differences between minimum wage workers in poor and rich households. Most importantly, minimum wage workers in richer households tend to be the children of the CPS reference person. In poorer households, minimum wage workers are more likely to be female.

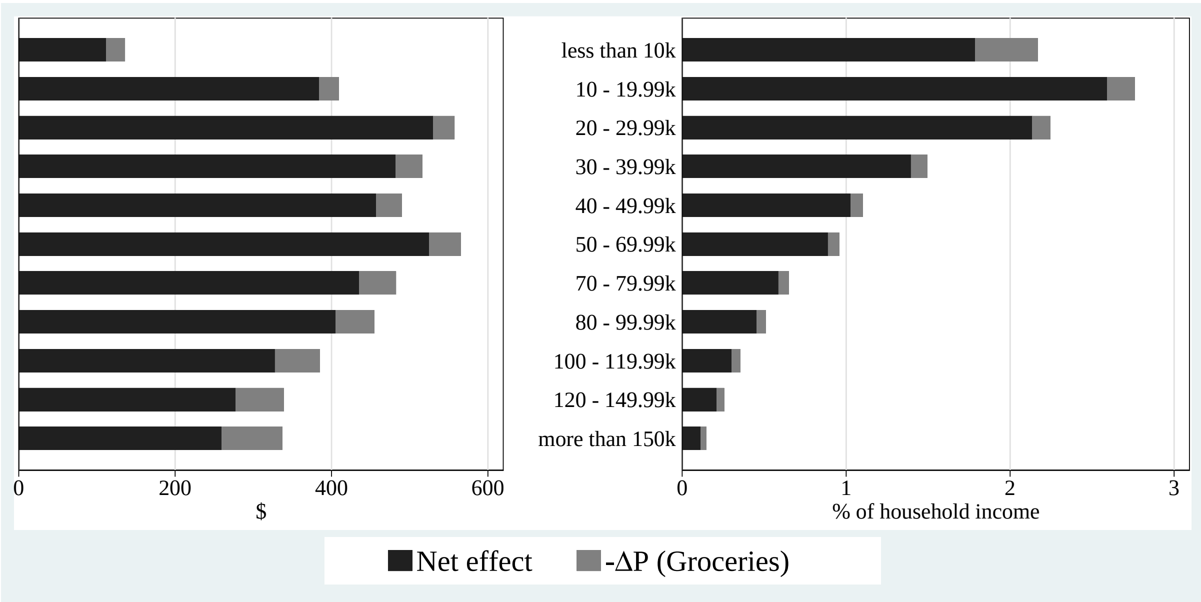
Figure E.1 presents the predicted increase in nominal incomes in US dollars and in percent of household income as the full length of the respective bars. The distribution of the gains expressed in US dollars may seem surprising at first.[46] In absolute terms, the poorest households gain relatively little compared to other brackets. Their annual incomes go up by about $136 and the biggest nominal benefits accrue to middle class households with incomes between $50,000 and $79,000, who gain about $565. This can be explained by low labor supply in the poorest bracket.[47] Second, households in the richest bracket still gain substantially. Minimum wage workers in this bracket differ from those in poorer households in one important aspect. As

[46]**?** estimates the impact of minimum wage increases on family incomes at different percentiles. The range of his reported estimates is quite large and the magnitudes depend on the included controls. He also finds that the poorest families gain less than slightly less poor families. Overall, our predictions for different income brackets are within the range of his estimates.

[47]Table E.1 illustrates that households in the lowest bracket work about 5 hours a week and 7 weeks a year on average, and as a result, labor is a relatively minor source of income.

shown in Table E.2, 71% of minimum wage workers in the richest bracket are children of the CPS household reference person, compared to around 10% in poorer brackets. Relative to household incomes, however, gains are distributed in a more progressive way: the poorest households gain 2.2% of their annual incomes, middle class households 1%, and the richest households gain 0.15%. Figure E.1 also illustrates the part of nominal gains that is offset by the Equivalent variation of price increases, which we discuss next.

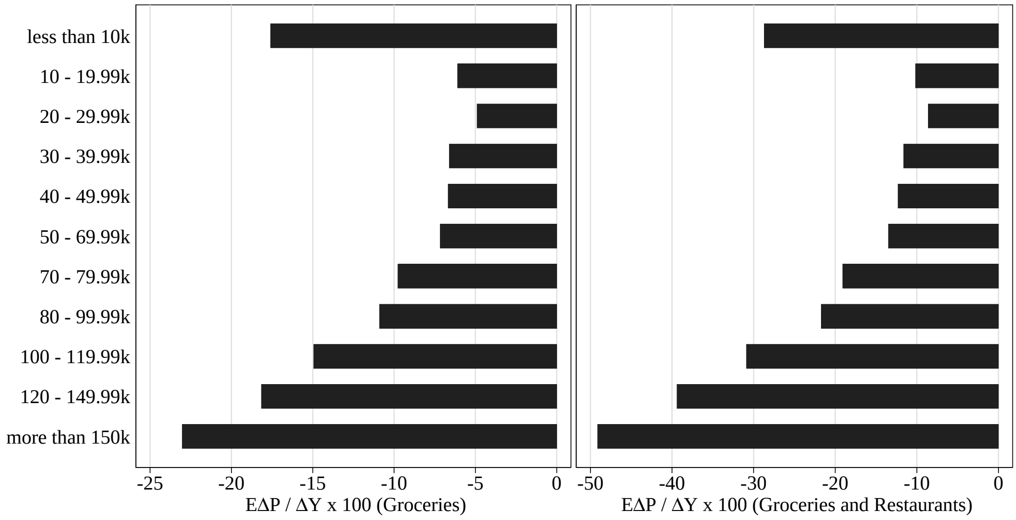
Figure E.1: Nominal gains, Equivalent Variation and net effect of a 20% minimum wage increase, grocery stores only



Notes: The figure shows nominal gains (length of the bar), EV of price increases in grocery stores (gray), and the net effect (black) in US dollars (left) and relative to mean household incomes (right).

Comparing cost and benefits. Figure E.2 and Table E.3a show the Equivalent Variation as a percentage of nominal gains to illustrate how much of the nominal gains are offset by the Equivalent Variation of price increases. For the poorest households, the price response in grocery stores of a 20% minimum wage increase offsets 17.6% of the nominal gains. This is non-negligible. The impact of price increases is very small for slightly less poor households with higher labor supply. In households with annual incomes between $10,000 and $79,000, 6–10% of their nominal gains are offset by the price response. For the richer households the percentage rises again and goes to up to 23.0% for the richest bracket. In the right panel of Figure E.2 and in Table E.3b, we also take into account price increases in restaurants for comparison. We use a minimum wage elasticity of restaurant prices of 0.07 estimated in Aaronson (2001) and expenditures for “Food Away from Home” in the CES to calculate the Equivalent Variation. The calculations suggests that price responses in restaurants matter regarding the gains from minimum wage increases. The effects are largest for the richest households (almost 50%). In the poorest households, the Equivalent Variation now offsets 28.7% of nominal gains.

Figure E.2: Equivalent Variation as percentage of nominal gains



Notes: The figure illustrates the Equivalent Variation (EV) as a percentage of nominal gains. The left panel is based on price increases in grocery stores. The right panel is based on price increases in grocery stores and restaurants.

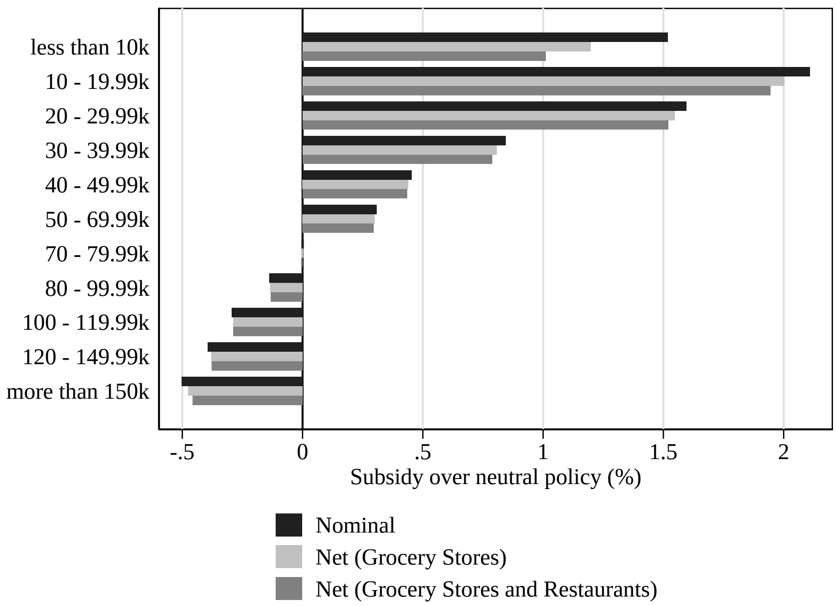
The price response mechanically reduces the nominal gains for all households. Moreover, due to differences in expenditures for groceries, the price response not only affects the level, but also the distribution of gains over different income brackets. To separately analyze the redistributive effect of minimum wage increases, we compare the distribution of gains to an inequality neutral income subsidy. In particular, we decompose gains for each income bracket as follows:

$$\frac{\widehat{Y}_h^L - Y_h^L - E_h \Delta P}{Y_h} = (1 + g + s_h) \tag{E11}$$

In this decomposition, we choose the level of the inequality neutral subsidy g to equal the overall increase in labor incomes, $\sum_i \hat{Y}_h^L - Y_H^L = (1+g)\sum_i Y_i$. We then calculate s_h for each bracket. These bracket-specific subsidies s_h measure the extent to which a minimum wage increase is redistributive. We calculate g and s_h for three measures of gains: for the initial nominal gains, for the gains taking into account price increases in grocery stores, and for the gains taking into account price increases in grocery stores and restaurants.

Figure E.3 presents the bracket specific subsidies. As expected, minimum wages reduce income inequality. The impact on inequality is largest for the purely nominal gains. Taking into account the price response reduces the redistributive impact. In terms of nominal gains, households in the poorest bracket gain an additional 1.5% of household income over an inequality neutral policy. Taking into account the price response in grocery stores reduces the additional gains to 1.20%. Further taking into account restaurants reduces the gains to 1.01%. For less poor households, the price response has a smaller impact on redistribution. Households that earn above $80,000 gain less from a minimum wage increase than they would from an inequality neutral policy. Taken together, these results suggest that price responses in groceries reduce the redistributive effects of minimum wage policies, but they do not offset them.

Figure E.3: Bracket specific income subsidy over inequality neutral policy



Notes: The figure isolates the impact of gains from minimum wage increases on inequality from the level effect. We decompose nominal gains, gains net of price increases in grocery stores, and net of price increases in grocery stores and restaurants into an inequality neutral part and a bracket specific subsidy using equation E11.

Summary of welfare results in tables. Finally, we report the numbers corresponding to Figures 5, E.1, and E.2 in Tables E.3 and E.3b. The Tables do not contain any information not depicted in the Figures, but provide a more readable summary of the results. Finally, Table E.4 present the same estimates as in Table E.3, but for transacted prices (i.e. including sales) as opposed to regular prices (i.e. excluding sales).

Elasticities of income-specific price indices. Are price elasticities different for products consumed by high- vs. low- income households? To study the distributional effects of the price effects more directly, we construct separate price indices for low-, medium- and high-income households using the IRI consumer panel data. The consumer panel data allows us to calculate yearly expenditures for each UPC by household income. Since household income is measured imprecisely,[48] we pool households in three broad brackets of yearly income: less than \$25,000, between \$25,000 and \$74,999 and more than \$75,000. We then use expenditure shares of each UPC for a given bracket as weights to compute a price index for this bracket. Households in the panel are located in two local markets. We pool households in both areas and assume that their expenditure weights are representative for the US overall. Furthermore, we average expenditure shares over all 10 years of data and keep weights constant in our index. Since we only observe expenditures for products bought by households in the panel, the income-specific price indices cover a selected and smaller sample of products.[49]

The inflation rates of the resulting income-specific price indices are highly correlated. Consistent with the findings in Jaravel (2018), the average inflation rate is lower for products consumed by higher income households. In Table E.5, we estimate our baseline specification for each index separately. All estimates are very close to our baseline estimates. The point estimates for the three indices are almost identical, and there are no significant differences between the response of price indices with expenditure weights for different income groups. This suggests that stores increase product prices across the board.

[48]The information on household income is in brackets between 2k and 25k, depending on the income level. More importantly, the income is not updated yearly in the consumer panel. Indeed, in many cases the household income refers to the year that the household entered the consumer panel and remains unchanged for several years.

[49]Many products that are present in the store-level price data are sold to none or few households in our panel. There are two potential reasons for this. First, our sample is much smaller. Second, some products may not be sold in the locations of panel households.

Table E.3: Nominal gains and Equivalent Variation of grocery price increases after a 20% increase in the minimum wage

(a) Taking into account price effects in grocery stores

	($)			(%)			(%)
	ΔY_h^L	$E_h \Delta P$	Net	ΔY_h^L	$E_h \Delta P$	Net	$E_h \Delta P / \Delta Y_h^L$
less than 10k	135.95	−23.92	112.03	2.17	−0.38	1.79	−17.59
10 - 19.99k	409.16	−24.95	384.21	2.76	−0.17	2.59	−6.1
20 - 29.99k	557.47	−27.3	530.17	2.25	−0.11	2.14	−4.9
30 - 39.99k	516.27	−34.09	482.18	1.49	−0.1	1.39	−6.6
40 - 49.99k	489.88	−32.7	457.18	1.1	−0.07	1.03	−6.68
50 - 69.99k	565.3	−40.5	524.8	0.96	−0.07	0.89	−7.16
70 - 79.99k	482.66	−47.13	435.53	0.65	−0.06	0.59	−9.76
80 - 99.99k	454.74	−49.54	405.21	0.51	−0.06	0.46	−10.89
100 - 119.99k	385.13	−57.54	327.6	0.35	−0.05	0.3	−14.94
120 - 149.99k	338.96	−61.54	277.42	0.26	−0.05	0.21	−18.15
more than 150k	337.25	−77.64	259.61	0.15	−0.03	0.11	−23.02

(b) Taking into account price effects in grocery stores and restaurants

	($)			(%)			(%)
	ΔY_h^L	$E_h \Delta P$	Net	ΔY_h^L	$E_h \Delta P$	Net	$E_h \Delta P / \Delta Y_h^L$
less than 10k	135.95	−39.0	96.96	2.17	−0.62	1.55	−28.68
10 - 19.99k	409.16	−41.58	367.58	2.76	−0.28	2.48	−10.16
20 - 29.99k	557.47	−47.92	509.54	2.25	−0.19	2.05	−8.6
30 - 39.99k	516.27	−59.99	456.28	1.49	−0.17	1.32	−11.62
40 - 49.99k	489.88	−60.27	429.61	1.1	−0.14	0.97	−12.3
50 - 69.99k	565.3	−76.21	489.08	0.96	−0.13	0.83	−13.48
70 - 79.99k	482.66	−92.0	390.66	0.65	−0.12	0.53	−19.06
80 - 99.99k	454.74	−98.7	356.04	0.51	−0.11	0.4	−21.71
100 - 119.99k	385.13	−118.93	266.21	0.35	−0.11	0.25	−30.88
120 - 149.99k	338.96	−133.51	205.45	0.26	−0.1	0.16	−39.39
more than 150k	337.25	−165.65	171.6	0.15	−0.07	0.07	−49.12

Notes: The tables show the nominal gains and Equivalent Variation (EV) of price increases in response to increasing all binding minimum wages in the US by 20%. Table E.3a uses Equivalent Variation of price increases in grocery stores. Table E.3b uses Equivalent Variation of price increases in grocery stores and restaurants. We show the mean nominal gains and EV for each income bracket in US dollars and in % of household income. ΔY_h^L is the increase in nominal household incomes. $E_h \Delta P$ is the EV of the predicted increase in prices at grocery stores. Net is the remaining welfare effect. $100 \cdot E_h \Delta P / \Delta Y_h^L$ illustrates the % of nominal income gains that is offset by price increases.

Table E.4: Nominal gains and Equivalent Variation of grocery price increases after a 20% increase in the minimum wage, using transacted prices (i.e. including sales) - price effects in grocery stores only

	($)			(%)			(%)
	ΔY_h^L	$E_h \Delta P$	Net	ΔY_h^L	$E_h \Delta P$	Net	$E_h \Delta P / \Delta Y_h^L$
less than 10k	135.95	−35.21	100.74	2.17	−0.56	1.61	−25.90
10 - 19.99k	409.16	−36.73	372.43	2.76	−0.25	2.51	−8.98
20 - 29.99k	557.47	−40.20	517.27	2.25	−0.16	2.08	−7.21
30 - 39.99k	516.27	−50.19	466.08	1.49	−0.15	1.35	−9.72
40 - 49.99k	489.88	−48.15	441.73	1.1	−0.11	0.99	−9.82
50 - 69.99k	565.3	−59.63	505.67	0.96	−0.10	0.86	−10.55
70 - 79.99k	482.66	−69.39	413.27	0.65	−0.09	0.56	−14.38
80 - 99.99k	454.74	−72.93	381.82	0.51	−0.08	0.43	−16.04
100 - 119.99k	385.13	−84.70	300.43	0.35	−0.08	0.28	−21.99
120 - 149.99k	338.96	−90.60	248.36	0.26	−0.07	0.19	−26.73
more than 150k	337.25	−114.31	222.94	0.15	−0.05	0.10	−33.89

Notes: Theis table show the nominal gains and Equivalent Variation (EV) of price increases in response to increasing all binding minimum wages in the US by 20%. Table E.4 uses Equivalent Variation of price increases in grocery stores. We show the mean nominal gains and EV for each income bracket in US dollars and in % of household income. ΔY_h^L is the increase in nominal household incomes. $E_h \Delta P$ is the EV of the predicted increase in prices at grocery stores. Net is the remaining welfare effect. $100 \cdot E_h \Delta P / \Delta Y_h^L$ illustrates the % of nominal income gains that is offset by price increases.

Table E.5: Effects for income specific price indices

Dep. variable: Store inflation w. different weights	Separate estimation						Joint estimation		
	(1) Low income	(2) Medium income	(3) High income	(4) Low income	(5) Medium income	(6) High income	(7) Low income	(8) Medium income	(9) High income
E_0^{leg}	0.009***	0.007**	0.008***				0.006**	0.005*	0.005
	(0.003)	(0.003)	(0.003)				(0.003)	(0.003)	(0.003)
E_2^{leg}	0.015***	0.013***	0.011***				0.007	0.005	0.004
	(0.004)	(0.004)	(0.004)				(0.005)	(0.006)	(0.006)
E_4^{leg}	0.018***	0.019***	0.015**				0.009	0.008	0.006
	(0.006)	(0.006)	(0.006)				(0.006)	(0.007)	(0.007)
E_0^{imp}				0.007	0.008	0.008	0.010	0.010	0.010
				(0.006)	(0.006)	(0.006)	(0.008)	(0.008)	(0.008)
E_2^{imp}				0.008	0.009	0.010	0.013	0.013	0.015
				(0.010)	(0.010)	(0.010)	(0.014)	(0.013)	(0.014)
E_4^{imp}				0.017	0.016	0.017	0.021	0.020	0.021
				(0.013)	(0.012)	(0.013)	(0.016)	(0.016)	(0.016)
Estimation Summary									
$E_4^{leg}+E_4^{imp}$	0.018***	0.019***	0.015**	0.017	0.016	0.017	0.030*	0.028*	0.027
	(0.006)	(0.006)	(0.006)	(0.013)	(0.012)	(0.013)	(0.016)	(0.016)	(0.016)
$\sum$ All	0.011	0.013	0.008	0.034*	0.036**	0.039**	0.031	0.030	0.033*
	(0.013)	(0.012)	(0.012)	(0.017)	(0.017)	(0.016)	(0.019)	(0.018)	(0.017)
$\sum$ Pre-event	-0.008	-0.008	-0.009	0.028**	0.028**	0.027**	0.017	0.015	0.017
	(0.007)	(0.007)	(0.007)	(0.013)	(0.012)	(0.012)	(0.015)	(0.014)	(0.014)
N	146739	146739	146739	146739	146739	146739	146739	146739	146739
controls	YES	YES	YES	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES	YES	YES	YES

Notes: The dependent variable is the store-level inflation rate with expenditure weights for different HH income brackets. Low: $< 25k$. Medium: $25k - 75k$. High: $> 75k$. Baseline controls are the unemployment rate and house price growth. The table lists cumulative elasticities E_R, R months after legislation or implementation. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $t-2$. SE are clustered at the state level. * p < 0.1, ** p < 0.05, *** p < 0.01.

F Timing of average wage increases

In this section, we look at the dynamics of the wage response around implementation and legislation dates. We show that average wages increase in grocery stores around the time of implementation, but not at the time of legislation. Therefore, firms are not forward-looking when setting wages, in contrast to the behavior we document regarding price setting.

To show this, we start with a similar strategy as the one described in section 6.2. We simply add quarter leads and lags of the logarithm of the implemented ($\log MW_{c(s),q}$) and legislated ($\log MWLeg_{c(s),q}$) minimum wages in state s (and county c) where the grocery store is located. The variable $\log MWLeg_{c(s),q}$ is measured as the logarithm of the highest future binding minimum wage set in current law (see figure 1b). Specifically, we estimate the following model:

$$\log \overline{W}_{c,q} = \gamma_c + \delta_q + \sum_{r=-k}^{k} \beta_r \log MWimp_{c(s),q-r} + \sum_{r=-k}^{k} \alpha_r \log MWLeg_{c(s),q-r} + X_{c,q} + \epsilon_{c,q} \quad \text{(F12)}$$

The coefficients β_r and α_r measure the elasticity of average wages in grocery stores with respect to the minimum wage r quarters ago, or r quarters in the future. We are not able to show the dynamics month after month here since the wage data are only available on a quarterly basis. We first estimate wage elasticities at legislation and implementation separately (i.e. we omit all terms related to $\log MWimp_{c(s),q-r}$ or $\log MWLeg_{c(s),q-r}$). We then jointly estimate wage elasticities at implementation and legislation by estimating equation F12 in full. Our results are displayed in Table F.1. They include 3 quarters before the quarter of implementation or legislation of the minimum wage (i.e. between 9 months and 11 months before implementation or legislation), and 2 quarters after (i.e. between 6 and 8 months after implementation or legislation). Because minimum wage changes often occur during the first month of a quarter (typically in January, at the beginning of the first quarter of the year, or in July, at the beginning of the 3rd quarter), our estimates in Table F.1 represent on average estimates 9 months before the minimum wage increase and 8 months after – a window that is consistent with the window used in all our price regressions.

As shown in the top panel of Table F.1, in columns 1-2, average wage increases in the quarter of the implementation of the minimum wage. A 10% increase in the implemented minimum wage leads to an average wage increase of about 7% (depending on the specification used, with (col. 2) or without (col. 1) state linear trends). Similarly, the second panel shows that there is no statistically significant increase in average wages in the quarter of legislation of the minimum wage. Note that the wage increase happens 2 quarters after the legislation is passed, which on average corresponds to when the minimum wage is implemented in our sample. The pattern of wage increase at implementation and not at legislation is preserved when equation F12 is estimated in full (col. 5 & 6).

Table F.1: Timing of average wage increases in grocery stores

Dep. variable: Labor cost per worker	Separate estimation (1) Baseline	(2) Trend	(3) Baseline	(4) Trend	Joint estimation (5) Baseline	(6) Trend
Implementation						
q-3	0.048	0.021			0.033	0.032
	(0.032)	(0.022)			(0.036)	(0.036)
q-2	-0.037	-0.039			-0.041	-0.043
	(0.037)	(0.037)			(0.042)	(0.040)
q-1	-0.002	-0.008			-0.011	-0.012
	(0.031)	(0.030)			(0.025)	(0.025)
Implementation	0.073**	0.065*			0.081**	0.073**
	(0.036)	(0.034)			(0.034)	(0.034)
q+1	0.010	0.016			0.022	0.020
	(0.042)	(0.042)			(0.040)	(0.040)
q+2	0.034	0.037			0.033	0.033
	(0.034)	(0.028)			(0.036)	(0.029)
Legislation						
q-3			0.015	-0.008	0.023	0.003
			(0.019)	(0.017)	(0.020)	(0.017)
q-2			0.015	0.006	0.004	-0.003
			(0.017)	(0.016)	(0.020)	(0.020)
q-1			-0.018	-0.022	-0.019	-0.021
			(0.020)	(0.019)	(0.021)	(0.020)
Legislation			0.014	0.010	0.026	0.024
			(0.017)	(0.017)	(0.020)	(0.020)
q+1			0.012	0.003	0.009	0.003
			(0.016)	(0.015)	(0.016)	(0.014)
q+2			0.046*	0.033*	-0.018	-0.021
			(0.024)	(0.018)	(0.018)	(0.018)
Estimation Summary						
At impl. or legisl.	0.073**	0.065*	0.014	0.010	0.108***	0.097**
	(0.036)	(0.034)	(0.017)	(0.017)	(0.041)	(0.040)
$\sum$ All	0.053	0.027	0.069	0.012	0.036	-0.008
	(0.061)	(0.052)	(0.049)	(0.029)	(0.067)	(0.056)
$\sum$ Pre-event	0.009	-0.026	0.011	-0.024	-0.009	-0.044
	(0.040)	(0.030)	(0.026)	(0.019)	(0.043)	(0.041)
N	80,722	80,759	80,722	80,759	80,722	80,759
Controls	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES
County FE	YES	YES	YES	YES	YES	YES
State linear trends	NO	YES	NO	YES	NO	YES

Notes: The table shows average wage elasticities wrt state-level minimum wages in the 2001-2012 period in grocery stores (NAICS code 4451), estimated using county-level data for 41 states used in our price regressions. The data are based on the QCEW.The outcome of interest is log average earnings. The table lists the average wage elasticities wrt the minimum wage 3 quarters before and 2 quarters after the quarter of legislation or implementation. Columns 1-2 show results of separate estimation of effects at implementation. Columns 3-4 show results of separate estimation at legislation. Columns 5-6 show results of joint estimation of effects at legislation and implementation. $\sum$ All is the sum of all lead and lag coefficients. $\sum$ Pre-event is the sum of all coefficients up to $q-1$. The controls are the log population and log average employment at the county level. SE are clustered at the state level. * p< 0.1, ** p< 0.5, *** p< 0.01.

G The minimum wage elasticity of marginal cost

In this section, we present a general theoretical model that helps to illustrate the relationship between the minimum wage elasticity of prices and the minimum wage elasticity of marginal cost at constant output.

We assume that the production technology is $Q = F(L, X)$, $L = G(L_1, L_2, \dots, L_N)$, with factor prices $P_x, W_1, W_2, \dots W_N$. F is assumed to be homogeneous of degree h and G is assumed to be linearly homogeneous. We assume competitive labor markets. We derive the elasticity of marginal cost to minimum wages keeping output constant.

Deriving the correct labor cost index

First, we are interested in the correct factor price index $\overline{W}$ that represents the marginal cost of increasing L. The firm minimizes labor cost LC:

$$LC(L, W_1, W_2, \dots, W_N) = \min_{L_1, L_2, \dots, L_N} W_1 L_1 + W_2 L_2 + \dots + W_N L_N$$
$$\text{s.t. } L = G(L_1, L_2, \dots, L_N)$$

The FOC for any L_i is that $\lambda G'_i = W_i$. λ_i is the Lagrange multiplier and equal to marginal labor cost LC_L. Because G is homogeneous of degree one, it follows that:

$$LC(L, w_1, w_2, \dots, w_N) = \lambda \sum_{i=1}^{N} G'_i L_i = \lambda L$$

Since λ is equal to marginal cost of increasing labor inputs, we can plug in $\lambda = LC_L$ and solve the resulting differential equation $LC = LC_L L$ to get that $LC = \overline{W} L$ for some $\overline{W}$ that is constant in L. As a result, marginal cost equals average cost, both are independent of the overall level of L, and $\overline{W} = LC/L$:

$$\overline{W}(W_1, W_2, \dots, W_N) = \sum_{i=1}^{N} \frac{W_i L_i^*}{L}$$

Deriving an expression for the elasticity

We can express the overall cost function as $C(\overline{W}, P_x, Q)$ and the overall marginal cost function as $C_Q(\overline{W}, P_x, Q)$. The derivative of marginal cost w.r.t. minimum wages can be written as:

$$\frac{\partial C_Q}{\partial MW} = \frac{\partial \frac{\partial C}{\partial Q}}{\partial \overline{W}} \frac{\partial \overline{W}}{\partial MW} = \frac{\partial L}{\partial Q} \frac{\partial \overline{W}}{\partial MW}$$

The last step uses Shepard's Lemma. Converting the derivative to an elasticity:

$$\frac{\partial C_Q}{\partial MW}\frac{MW}{C_Q} = \underbrace{\frac{\overline{W}L}{C}}_{(1)}\underbrace{\frac{\partial\overline{W}}{\partial MW}\frac{MW}{\overline{W}}}_{(2)}\underbrace{\frac{AC}{MC}}_{(3)}\underbrace{\frac{\partial L}{\partial Q}\frac{Q}{L}}_{(4)}$$

The minimum wage elasticity of marginal cost is given by the product of: (1) The cost share of labor cost in total variable cost; (2) the minimum wage elasticity of the average wage; (3) the ratio of average to marginal cost; (4) the output elasticity of labor demand.

Final step

We now show that $\frac{AC}{MC}\frac{\partial L}{\partial Q}\frac{Q}{L} = 1$ when F is homogeneous of degree h. If $h = 1$, both $\frac{AC}{MC} = 1$ and $\frac{\partial L}{\partial Q}\frac{Q}{L} = 1$. More generally, if F is homogeneous of degree h, we can write the cost function as $C = Q^{\frac{1}{h}}\omega$, where ω is constant in Q and typically depends on factor prices. As a result:

$$\frac{AC}{MC} = \frac{Q^{\frac{1}{h}-1}\omega}{\frac{1}{h}Q^{\frac{1}{h}-1}\omega} = h$$

and applying Shepard's Lemma:

$$\frac{\partial L}{\partial Q}\frac{Q}{L} = \frac{\partial\frac{\partial C}{\partial \overline{w}}}{\partial Q}\frac{Q}{\frac{\partial C}{\partial \overline{w}}} = \frac{\partial(Q^{\frac{1}{h}}\frac{\partial \omega}{\partial \overline{w}})}{\partial Q}\frac{Q}{Q^{\frac{1}{h}}\frac{\partial \omega}{\partial \overline{w}}} = \frac{1}{h}Q^{\frac{1}{h}-1}Q^{1-\frac{1}{h}} = \frac{1}{h}$$

As a result $\frac{AC}{MC}\frac{\partial L}{\partial Q}\frac{Q}{L} = 1$, and

$$\frac{\partial C_Q}{\partial MW}\frac{MW}{C_Q} = \frac{\overline{W}L}{C}\frac{\partial\overline{W}}{\partial MW}\frac{MW}{\overline{W}}$$

The minimum wage elasticity of marginal cost is equal to the minimum wage elasticity of the average wage, times the labor share in cost.

It is instructive to discuss the three assumptions that are central for this derivation. First, we need to assume that different labor inputs can be aggregated in a linear homogeneous way. This implies that the shares of different types of workers do not depend on the size of a store. Second, we need to assume that grocery stores' overall production technology is homogeneous to some degree. This assumption is much less restrictive and is fulfilled by all commonly used production functions we are aware of. Finally, we derive these predictions assuming constant output. Output does not matter for marginal cost in the case of constant returns to scale. In the case of non-constant returns, any change in output affects marginal cost in a way we do not account for here. We look into the effects on minimum wages on grocery store output in Table J.1 in the appendix and do not find any evidence for a change in grocery stores' output.

H Further evidence on the earnings elasticity in grocery stores

H.1 Minimum wages in the grocery sector

Three stylized facts motivate our analysis of the price effects of minimum wages in the grocery sector.

First, groceries are an important factor in households' cost of living, particularly for poor households. Using data from the Consumer Expenditure Survey (CES), we estimate groceries make up 11% of household expenditures (see Appendix Table B.2).[50] This is more than twice as large as For households in the poorest quintile, groceries make up 14 to 15% of expenditures. For households in the richest quintile, this share amounts to 9%.

Second, labor costs are an important part of the overall costs of grocery stores. Using data from the 2007 and 2012 BLS Annual Retail Trade Surveys, we estimate that the labor cost share in variable cost —which should matter for price setting in the short run—amounts to 16% (see Appendix Table B.3). Labor costs include salaries, fringe benefits and commission expenses. Variable costs include labor costs, costs of goods sold and other variable costs (such as transport and packaging costs).[51] We also note that the most important factor in grocery store costs are the costs of goods sold (83%).

Third, a substantial share of grocery store employees are paid wages close to the minimum wage. Using data on hourly wages from the NBER files of the CPS MORG, Appendix Figure B.1 plots the distribution of wages in grocery stores relative to the local minimum wage. A large share of grocery store workers are paid wages at or close to the local minimum wage during all three periods. In the period when most of the minimum wage hikes in our sample happen (2006–2009), 21% of grocery store workers earn less than 110% of the minimum wage. Recent literature suggests that even workers with wages above the minimum wage may be affected by "ripple effects" of a hike (Autor et al., 2016; Dube et al., 2015), and as a result a large share of grocery store workers would likely be affected by minimum wage increases. At the end of our sample period, for instance, almost half of all grocery store workers earn less than 130% of the local minimum wage. As shown by Table H.1 in the appendix, the share of these workers in total hours worked in groceries amounts to approximately 40% in this period, and the share in total labor earnings to 25%.

We first present some additional statistics on minimum wage employment in the grocery

[50] We define groceries' expenditures as the sum of expenditures in the following categories: Food at Home, Household Supplies, Alcoholic Beverages and Personal Care Products and Services.

[51] Variable costs differ from total costs. In addition to variable costs, total costs include building and equipment costs (such as rents, utilities, depreciation and purchases of equipment), purchased services (such as maintenance, advertisement, etc.) and other operating expenses (such as taxes). Note that our estimate of labor cost share in variable cost does not include purchased services in the denominator. These services make up about 2% of total costs and include some tasks that are likely done by low-skilled workers, for example maintenance work. These costs may depend on minimum wages as well, but it is hard to determine to which extent.

sector. Table H.1 presents the share of workers below 110% and 130% of the local minimum wage in employment, hours and earnings of grocery stores. We also compare the share to other relevant industries. These statistics complement the full wage distribution in grocery store employment presented in Appendix Figure B.1. The shares in hours are lower than in employment—as minimum wage workers are more likely to work part-time—and in earnings, as minimum wage workers have the lowest hourly wages.

Table H.1: Statistics on minimum wage employment in grocery stores and other relevant sectors

	Employment		Hours		Earnings	
	$\leq$ 110%	$\leq$ 130%	$\leq$ 110%	$\leq$ 130%	$\leq$ 110%	$\leq$ 130%
			2001 - 2005			
Grocery Stores	12.1	29.6	9.0	23.0	4.5	13.1
Other Retail Trade	7.6	18.5	5.6	14.1	2.2	6.5
Restaurants	31.7	50.2	26.1	42.0	13.1	25.0
Other sectors	4.0	8.5	3.1	6.8	0.9	2.3
			2006 - 2009			
Grocery Stores	20.7	38.8	16.1	31.4	8.9	19.0
Other Retail Trade	11.6	25.0	8.5	19.3	3.6	9.4
Restaurants	39.5	58.3	32.9	50.1	18.3	31.9
Other sectors	5.2	11.1	4.1	9.0	1.2	3.2
			2010 - 2012			
Grocery Stores	25.1	48.8	19.2	40.3	11.1	25.4
Other Retail Trade	15.9	34.8	11.8	27.4	5.3	13.9
Restaurants	45.2	66.5	37.9	58.1	22.5	39.4
Other sectors	6.5	14.7	5.1	12.0	1.6	4.4

Notes: Based on CPS ORG data. Retail trade corresponds to NAICS 44–45, grocery stores to NAICS 4451, and restaurants to NAICS 722. Wages are computed using reported hourly wages for workers paid by the hour, and weekly earnings divided by weekly hours for other workers. Shares are calculated first for each state and year and subsequently averaged over all states and years in a period. All statistics are calculated using the CPS earnings weight.

In the main part of the paper (Section 6) we report regressions that show that earnings in grocery stores are strongly affected by minimum wage hikes. This section discusses several extensions to this result. In Table H.2 we first look into the dynamics of the wage effects by including leads and lags of the minimum wage to the regression. We find that the earnings effect of the minimum wage hike are concentrated in the period when the hike is implemented. The leads and lags are generally not statistically significant. Second, the table also reports the results of specifications that account for Census-division period fixed effects (col. 2, 5, 8) and that weight the regressions with county average total employment (col. 3, 6, 9). The results are similar as in our baseline table reported in the main part of the paper.

In Table H.3, we study how the estimated earnings elasticity varies with the bindingness of the minimum wage in a county. We expect larger earnings effects in counties where the difference between the new minimum wage and the initial prevailing wage is larger. For each county, industry and each minimum wage hike, we thus compute the difference between the new minimum wage after the hike and the prevailing average wage in the respective industry four quarters before the hike.[52] For each county and industry, we then average these differences over all hikes in a county. We use this average difference to assign counties into four groups in terms of the bindingness of the minimum wage, based on the county's position in the distribution of differences between prevailing wage and new minimum wage. If it belongs to the first quartile of this distribution, the county is considered a county where the minimum wage has low bindingness in the respective sector. If it belongs to the top quartile of the distribution, the minimum wage is considered to be strongly binding. Table H.3 reports separate earnings elasticities for the four categories of counties. In the case of grocery stores, the earnings elasticity is larger than our baseline elasticity in counties in which the minimum wage is strongly binding. We find no differences within the remaining three groups of counties. In each of them, the elasticity is slightly lower than our baseline estimate.

[52] The difference is estimated by computing a rough measure for the quarterly earnings of a full-time minimum wage worker. We do this by multiplying the hourly minimum wage by eight hours and $22 * 3$ days per quarter.

Table H.2: Robustness: Earnings and employment elasticities to the minimum wage

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
		Retail trade			Grocery stores			Acc. and food services	
Panel A: Earnings									
t-4	0.011			0.004			-0.019		
	(0.020)			(0.035)			(0.027)		
t-3	0.022			0.043			0.037*		
	(0.015)			(0.039)			(0.019)		
t-2	-0.021*			-0.024			-0.042		
	(0.012)			(0.037)			(0.026)		
t-1	-0.003			-0.001			0.057*		
	(0.010)			(0.030)			(0.030)		
t	0.039**	0.075***	0.048*	0.056*	0.062*	0.108**	0.046*	0.171***	0.151***
	(0.015)	(0.020)	(0.026)	(0.029)	(0.036)	(0.043)	(0.027)	(0.027)	(0.024)
t+1	0.011			0.011			0.073***		
	(0.018)			(0.037)			(0.022)		
t+2	0.009			0.021			-0.011		
	(0.013)			(0.034)			(0.030)		
t+3	-0.003			-0.008			0.057*		
	(0.010)			(0.027)			(0.029)		
t+4	-0.013			0.036			-0.029		
	(0.015)			(0.034)			(0.024)		
Obs	124,000	124,000	124,000	80,722	80,722	80,722	98,056	98,056	98,056
Panel B: Employment									
t-4	0.026			-0.076			-0.035		
	(0.021)			(0.050)			(0.034)		
t-3	-0.018**			0.010			0.054**		
	(0.007)			(0.016)			(0.021)		
t-2	0.021*			0.005			0.031		
	(0.011)			(0.014)			(0.022)		
t-1	0.007			0.002			-0.001		
	(0.010)			(0.010)			(0.024)		
t	-0.017*	0.029*	-0.002	-0.018	0.072**	-0.010	-0.055**	-0.008	-0.042
	(0.009)	(0.016)	(0.027)	(0.018)	(0.036)	(0.048)	(0.024)	(0.026)	(0.033)
t+1	-0.012			0.018			0.022		
	(0.007)			(0.014)			(0.021)		
t+2	0.017			0.005			0.013		
	(0.011)			(0.014)			(0.023)		
t+3	0.014			0.003			-0.019		
	(0.011)			(0.014)			(0.028)		
t+4	-0.033*			0.034			-0.069**		
	(0.016)			(0.041)			(0.026)		
Obs	124,000	124,000	124,000	80,722	80,722	80,722	98,056	98,056	98,056
Controls	YES	YES	YES	YES	YES	YES	YES	YES	YES
County FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Div.-time FE	NO	YES	NO	NO	YES	NO	NO	YES	NO
Weights	NO	NO	YES	NO	NO	YES	NO	NO	YES

Notes: The table shows elasticities to state-level minimum wages in the 2001–2012 period by industry, estimated using county-level panel data from the QCEW for the 41 states used in our price regression. Retail trade corresponds to NAICS 44–45, grocery stores to NAICS 4451, and accommodation and food services to NAICS 72. The outcome in panel A is log the average earnings by industry. The outcome in Panel B is the log of the number of workers by industry, computed as the average employment of the three months in the respective quarter. Controls are log of county population and the log of total employment in private industries per county. SE are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table H.3: Earnings elasticities by bindingness of the minimum wage

	(1) Strongly binding	(2) Moderately binding	(3) Weakly binding	(4) Very weakly binding
		Grocery stores		
log MW	0.155***	0.081**	0.083*	0.079
	(0.045)	(0.033)	(0.042)	(0.067)
N	16,567	19,200	21,406	19,851
		Retail trade		
log MW	0.081***	0.026	0.010	0.007
	(0.024)	(0.025)	(0.026)	(0.033)
N	28,606	30,840	32,216	32,139
		Accomodation and food services		
log MW	0.168***	0.183***	0.184***	0.079***
	(0.026)	(0.022)	(0.033)	(0.028)
N	21,242	23,724	25,076	25,880

Notes: The table shows elasticities to state-level minimum wages in the 2001–2012 period by industry, estimated using county-level panel data from the QCEW for the 41 states used in our price regression. Retail trade corresponds to NAICS 44–45, grocery stores to NAICS 4451, and accommodation and food services to NAICS 72. The outcome is log the average earnings by industry. Controls are log of county population and the log of total employment in private industries per county. SE are clustered at the state level. The minimum wage bindingness is the average county-level difference between the industry-specific wage (4 quarters before a subsequent hike) and the new minimum wage, averaged across all hikes in a county. The four categories correspond to quartiles of the distribution of this gap.

I Minimum wages and prices of COGS

In this section, we discuss the possibility that minimum wages increase the cost of goods sold (COGS). As shown in Table B.3, COGS make up about 83% of grocery stores' variable cost. Moreover, retail stores have been shown to be very responsive to changes in COGS (see, e.g., Eichenbaum et al., 2011; Nakamura and Zerom, 2010). If minimum wage workers are employed in the production of grocery products, producers may also increase their prices. Due to the high cost share of COGS in retailers' cost, even a relatively minor increase in producer prices could affect retail prices.

Our price data does not include any measure of wholesale cost, and we cannot estimate the impact of minimum wages on the wholesale cost of grocery products directly. We thus follow MaCurdy (2015) and use input-output tables to calculate a prediction of the elasticity of prices for sectors producing groceries under the assumption of full pass-through all along the production chain.

The input-output tables of the national accounts cover sectoral labor shares[53] s_j^L, which we use as the labor cost elasticity of prices for this sector. We use minimum wage shares in sectoral earnings s_j^{mw} computed from the CPS as the elasticity of labor cost to minimum wages. Finally, we compute the value of output of industry j used to product one dollar of output in industry i from the input-output tables.[54] We denote this coefficient $\alpha_{i,j}$. We can then predict the minimum wage elasticity of producer prices in each sector as:

$$\frac{\partial P_i}{\partial MW}\frac{MW}{P_i} = \sum_j \alpha_{i,j} \cdot s_j^L \cdot s_j^{mw} \tag{I13}$$

We present the predicted elasticity of producer prices based on equation I13 in Table I.1. We use the domestic requirements table for 389 disaggregated industries provided by the BEA. We predict the elasticity for 26 manufacturing industries that are relevant for grocery stores. All calculations are based on data from 2007. Columns 2 and 3 present the direct cost elasticity, which captures the impact of minimum wage workers employed in the sector itself. These elasticities are quite small. Columns 4 and 5 present the final elasticities, which also capture predicted price increases of inputs. These elasticities are substantially larger. The difference is driven by low wages in the sectors that deliver primary inputs to food manufacturing sectors. We present both measures for minimum wage shares based on workers earning below 110% and 130% of the local minimum wage.

Overall, the elasticities reported in Table I.1 are of similar magnitude as the direct impact of increases in labor cost on retail marginal cost. The output-weighted average predicted elasticity of producer prices in grocery manufacturing industries amounts to 0.016 when we use 110% of the minimum wage to define minimum wage workers and 0.024 when we use 130%. Full pass-through in manufacturing industries could thus affect the marginal cost of grocery stores

[53] These labor shares are in revenues, not cost.

[54] This corresponds to the i, j entry of the domestic requirements matrix in the input-output tables.

Table I.1: Predicted MW elasticities of producer prices in grocery manufacturing

Manufacturing Sector	Direct cost elasticity		Final cost elasticity	
MW worker definition:	< 110%	< 130%	< 110%	< 130%
Breakfast cereal	0.003	0.005	0.011	0.017
Sugar and confectionery	0.006	0.009	0.019	0.029
Frozen food	0.006	0.009	0.020	0.030
Fruit and vegetable canning, pickling, and drying	0.005	0.008	0.019	0.029
Fluid milk and butter	0.003	0.005	0.018	0.026
Cheese	0.002	0.003	0.018	0.026
Dry, condensed, and evaporated dairy	0.002	0.003	0.017	0.024
Ice cream and frozen dessert	0.004	0.005	0.015	0.022
Animal slaughtering, rendering, and processing	0.005	0.008	0.018	0.026
Poultry processing	0.007	0.012	0.022	0.033
Seafood product preparation and packaging	0.009	0.013	0.020	0.030
Bread and bakery	0.012	0.018	0.020	0.031
Cookie, cracker, pasta, and tortilla	0.008	0.012	0.017	0.026
Snack food	0.008	0.012	0.019	0.029
Coffee and tea	0.008	0.012	0.029	0.042
Flavoring syrup and concentrate	0.013	0.019	0.017	0.025
Seasoning and dressing	0.011	0.016	0.024	0.035
All other food	0.011	0.016	0.026	0.039
Soft drink and ice	0.002	0.003	0.011	0.018
Breweries	0.001	0.002	0.008	0.012
Wineries	0.002	0.004	0.013	0.021
Distilleries	0.001	0.002	0.004	0.007
Tobacco	0.001	0.001	0.004	0.006
Sanitary paper	0.004	0.005	0.011	0.016
Soap and cleaning compound	0.002	0.003	0.006	0.010
Toilet preparation	0.002	0.003	0.006	0.010
Output weighted average	0.005	0.008	0.016	0.024
Average	0.005	0.008	0.016	0.024

to a comparable extent as the direct effect of increasing labor costs.

The extent to which increases in COGS affect our estimates depends on whether they are passed through, but also on whether they occur locally. If wholesale groceries are perfectly tradeable, a minimum wage hike would increase COGS equally for stores everywhere, and any pass-through of this cost increase would be absorbed in time fixed effects in our baseline estimation.

Our data does not contain information about where a particular product is produced. However, we can study the origin composition of groceries sold in a state using grocery wholesale-to-retail flows reported in the Commodity Flow Survey.[55] This dataset covers sales of manufacturing companies, but also intermediaries such as merchant wholesalers or warehouses. As a result, we cannot identify the location of production with certainty.

[55]The Commodity Flow Survey has been used to document home bias in intra-national trade in the US by Wolf (2000). We refer the reader to his paper for a detailed description of the data.

We analyze the origin composition of products sold in retail using data on intrastate trade flows from wholesalers for groceries, farm products, alcoholic beverages and drugs (subsequently summed up under the term "groceries") provided in the 2007 Commodity Flow Survey. The Commodity Flow Survey data are subject to some important limitations. First, it counts flows originating from manufacturers, but also from distribution centers and similar establishments. The latter may not be produced locally. Second, the flows capture all flows originating from merchant wholesalers, irrespective of the destination industry. Merchant wholesalers are defined by selling to retail establishments, but the flows in the CFS capture not just flows to grocery stores but also other retail establishments. The numbers we calculate here should be interpreted as very suggestive evidence.

Let Y_{ij} be the flow of groceries from state i to state j. We calculate "production" of state s valued at wholesale prices as the sum of all flows originating in state s, $\sum_j Y_{sj}$. We can calculate "consumption" of state s as all flows with destination in state s, $\sum_i Y_{is}$. The share of locally produced products in grocery consumption of state s is then given by $Y_{ss}/\sum_i Y_{is}$. The exposure of state s to cost increases in another state S can be calculated as $Y_{Ss}/\sum_i Y_{is}$.

Our results suggest that the share of local products in grocery consumption is higher than the state's share in national grocery production. For example, California has a 14% share in the national production of groceries and 91% of groceries consumed in California are produced locally. Vermont accounts for a mere 0.1% of US grocery production, yet 30% of groceries consumed in Vermont are produced locally. This suggests a substantial home bias in US grocery consumption, a fact that has been documented for interstate trade as a whole in Wolf (2000). We document this relationship in Figure I.1.

Figure I.1: Home bias in grocery wholesale-to-retail flows

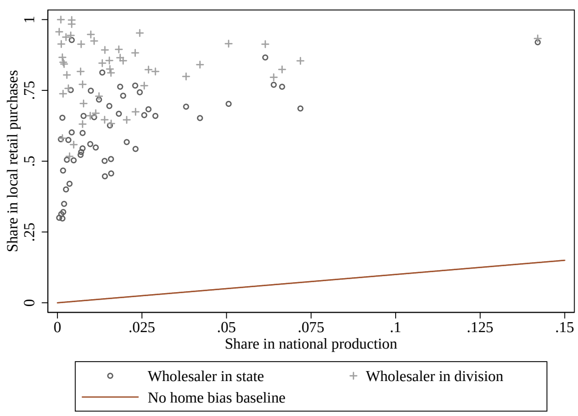


Table I.2 documents trade flows for all states. The share of local grocery products in consumption (Destination) is systematically higher than the share of states' products in national production (Origin). Flows from other states are small on average. Even in small states like Delaware or Rhode Island, the average flow from other states amounts to less than 1.5% of consumption.

Overall, these results suggest that a disproportionate share of grocery products are delivered by wholesalers located in the same state or census division as the retailers they supply. We interpret this as evidence for some home bias in grocery consumption. Consequently, it is possible that some effects of local wholesale price changes would be captured in our estimation, especially if we do not account for chain-time fixed effects.

Because a disproportionate share of grocery products are delivered by wholesalers located in the same state or census division as the retailers they supply, we assume that the major part of the price effect occurs in the state in which the minimum wage occurs when calculating pass-through rates in the lower part of Table 5 . We further assume that the price pass-through is the same in the retail sector and in COGS. In order to do so, we proceed in several steps. We first calculate the implied cost pass-through assuming a full price pass-through in COGS, assuming that a 10% increase in the minimum wage would increase the prices of COGS by 0.024% (if minimum wage workers earnings up to 130% of the minimum wage are affected by the hike). This gives us a 'first round' estimate of the implied cost pass-through that includes predicted effects on COGS. Second, using those pass-through estimates in the retail sector, we equate this pass-through to the price pass-through in the COGS and calculate a new marginal cost elasticity with respect to the minimum wage. This leads us to 'second round' estimates of the implied cost pass-through that includes predicted effects on COGS. We repeat this procedure until the price pass-through in the retail sector and in the COGS converge. Convergence happened in the fifth round. We therefore report our fifth round estimates in the lower part of Table 5. For the estimates of pass-through at legislation and implementation we predicted that a 10% in the minimum wage would increase the prices of COGS by 0.023% (column 1), 0.019% (column 2), and 0.013% (column 3).

Table I.2: Summary of wholesale-to-retail flows between states

	Share in National Total		Share in Consumption		Flows from other states		
	Consumption	Production	Production in State	Production in Division	Mean Consumption Share	Max Consumption Share	Max Origin
California	0.134	0.142	0.92	0.933	0.002	0.013	New Jersey
Florida	0.06	0.061	0.866	0.913	0.003	0.044	Georgia
Texas	0.067	0.064	0.77	0.796	0.005	0.039	Tennessee
Washington	0.021	0.023	0.767	0.883	0.005	0.063	Oregon
Minnesota	0.019	0.019	0.763	0.865	0.005	0.061	Illinois
Illinois	0.052	0.066	0.763	0.824	0.005	0.093	Missouri
Nebraska	0.008	0.01	0.749	0.948	0.005	0.094	Kansas
Michigan	0.029	0.024	0.743	0.953	0.005	0.083	Ohio
North Carolina	0.02	0.019	0.731	0.855	0.005	0.047	Georgia
Arizona	0.014	0.012	0.717	0.73	0.006	0.2	California
New Jersey	0.039	0.051	0.702	0.915	0.006	0.145	New York
Iowa	0.016	0.015	0.695	0.856	0.006	0.065	Illinois
Ohio	0.039	0.038	0.692	0.799	0.006	0.092	Pennsylvania
New York	0.074	0.072	0.686	0.854	0.006	0.135	New Jersey
Massachusetts	0.026	0.027	0.683	0.823	0.006	0.1	New York
Wisconsin	0.021	0.018	0.668	0.895	0.007	0.215	Illinois
Tennessee	0.017	0.026	0.663	0.767	0.007	0.094	Kentucky
Missouri	0.023	0.029	0.66	0.817	0.007	0.137	Kansas
Utah	0.007	0.008	0.66	0.704	0.007	0.112	Arkansas
Oregon	0.011	0.011	0.655	0.925	0.007	0.179	Washington
Vermont	0.001	0.001	0.653	0.867	0.007	0.14	New Hampshire
Pennsylvania	0.041	0.042	0.652	0.841	0.007	0.104	New Jersey
Kansas	0.015	0.016	0.626	0.825	0.007	0.15	Missouri
Oklahoma	0.009	0.007	0.6	0.771	0.008	0.152	Texas
New Mexico	0.004	0.003	0.575	0.757	0.009	0.17	Texas
Louisiana	0.033	0.02	0.568	0.646	0.009	0.107	Illinois
Alabama	0.012	0.01	0.56	0.661	0.009	0.131	Georgia
Georgia	0.025	0.023	0.543	0.674	0.009	0.147	Tennessee
South Carolina	0.01	0.007	0.532	0.913	0.009	0.189	Georgia
Mississippi	0.007	0.007	0.522	0.817	0.01	0.147	Tennessee
Virginia	0.021	0.016	0.508	0.812	0.01	0.205	Maryland
Idaho	0.004	0.003	0.505	0.805	0.01	0.262	Utah
Connecticut	0.012	0.014	0.501	0.646	0.01	0.188	New York
Maryland	0.016	0.016	0.457	0.633	0.011	0.139	Pennsylvania
Indiana	0.016	0.014	0.447	0.893	0.011	0.266	Illinois
West Virginia	0.004	0.004	0.42	0.517	0.012	0.222	Pennsylvania
Maine	0.005	0.003	0.4	0.938	0.012	0.442	Massachusetts
New Hampshire	0.003	0.002	0.349	0.843	0.013	0.285	Massachusetts
Rhode Island	0.003	0.002	0.32	0.849	0.014	0.364	Massachusetts
DC	0.002	0.001	0.313	0.914	0.014	0.384	Maryland
Delaware	0.002	0.001	0.298	0.581	0.014	0.284	Maryland
Mean	0.023	0.023	0.607	0.811	0.008	0.158	

J Effects of minimum wages on output

Table J.1 presents the results of equation 3 estimated with quantities and revenues as dependent variables. Quantity indices are constructed the same way as the price index. Log revenues are total store revenues. Both outcome variables have a higher variance than price indices. To increase the precision of the estimates, we thus aggregate the data to the quarterly frequency. Additionally, we weight the regressions with the inverse of the store-level variance of each outcome to account for the unequal precision/variability of each store-level series. These two adjustments reduce the standard errors of the estimates substantially but have very limited influence on the point estimates of the coefficients. For completeness, the last row of the table reports price regressions with quarterly price inflation constructed analogously to the quantity index and the revenues.

Overall, the table provides no evidence that minimum wages affect quantities and revenues, neither at legislation nor implementation. In fact, the point estimates at implementation, where demand effects are most likely to occur, are negative. Overall, the estimates in column 3, for instance, rule out elasticities of quantities with respect to the minimum that are larger than 0.02 at implementation. Note that the gap between quantity indices and revenues is insignificant but largely consistent with the price response that we estimate (presented in column 7). The price and quantity responses do not exactly equal the revenue responses because of the slightly different weighting of stores across outcomes and because prices and quantities are constructed as geometric averages of product price changes while revenues are simply total store-level revenues.

Table J.1: Effects of minimum wages on output, revenues and prices at quarterly level

	(1) Quantities	(2) Revenues	(3) Quantities	(4) Revenues	(5) Quantities	(6) Revenues	(7) Prices
$\Delta leg_{s(j),t-1}$	0.012	0.007			0.010	0.004	-0.004
	(0.037)	(0.029)			(0.038)	(0.029)	(0.004)
$\Delta leg_{s(j),t+0}$	-0.034	0.004			-0.031	0.004	0.015***
	(0.032)	(0.026)			(0.031)	(0.024)	(0.003)
$\Delta leg_{s(j),t+1}$	-0.019	-0.027			-0.009	-0.023	0.003
	(0.023)	(0.021)			(0.024)	(0.019)	(0.004)
$\Delta leg_{s(j),t+2}$	-0.021	-0.009			-0.012	-0.002	0.003
	(0.028)	(0.026)			(0.032)	(0.032)	(0.003)
$\Delta imp_{s(j),t-1}$			-0.038	-0.042	-0.026	-0.032	-0.004
			(0.042)	(0.037)	(0.047)	(0.046)	(0.006)
$\Delta imp_{s(j),t+0}$			-0.027	-0.014	-0.024	-0.008	0.006
			(0.024)	(0.027)	(0.027)	(0.032)	(0.008)
$\Delta imp_{s(j),t+1}$			0.011	-0.035	0.010	-0.036	-0.008
			(0.045)	(0.050)	(0.045)	(0.048)	(0.006)
$\Delta imp_{s(j),t+2}$			0.021	-0.002	0.019	-0.004	-0.001
			(0.036)	(0.039)	(0.036)	(0.039)	(0.007)
Observations	76711	76711	76711	76711	76711	76711	76711
Controls	YES	YES	YES	YES	YES	YES	YES
Time FE	YES	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES	YES
Weights	Var	Var	Var	Var	Var	Var	Var

Notes: The dependent variable is the quarterly growth rate in the store-level quantity index, in store-level log revenues, and price inflation based on quarterly versions of equation 3. We do not use sales-filtered prices for comparability with the quantity index (where sales are included as well). The sample covers the 2001–2012 period. Columns 1–2 show results of separate estimation of effects at legislation. Columns 3–4 show results of separate estimation of effects at implementation. Columns 5–7 show results of joint estimation of effects at implementation and legislation. To increase precision, the estimations are weighted using the inverse of the store-level variance of the outcome as weights. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

K Comparison to Ganapati and Weaver (2017) and Leung (2020)

Two closely related contemporaneous papers also study the effects of minimum wages on prices in grocery stores. Ganapati and Weaver (2017) and Leung (2020) both use scanner data provided by Nielsen that covers a somewhat different time period. Despite using the same data, the two papers reach different conclusions: Leung (2020) finds somewhat larger elasticities than we do, while Ganapati and Weaver (2017) find no effects of the minimum wage on grocery prices.

Our empirical analyses differ slightly from those of Leung (2020) and Ganapati and Weaver (2017). First, these papers study the effects of minimum wage increases at the time of implementation, while our main effects occur at the time legislation is passed, in line with Aaronson (2001). Second, we estimate pass-through regressions in first differences, which relate inflation to changes in the minimum wage and include fixed effects that control for differential inflation trends. The leads and lags in our pass-through regressions allow us to study the timing of the effect in detail. Both other papers estimate level regressions and control for a slightly different set of fixed effects.

Leung (2020) finds minimum wage elasticities of prices that are somewhat larger than ours, i.e. around 0.06 (Table 3 in the paper). Because the baseline estimates in Leung are quite imprecise, they are statistically indistinguishable from ours. A likely explanation for the differences in the point estimates is the fact that Leung's data cover the 2006–2015 period, whereas our data spans the 2001–2012 period. As we show in Appendix Figure B.1, minimum wages are considerably more binding toward the end of our sample period. For the years in which the two samples overlap, Leung's estimates are remarkably in line with our preferred elasticity.[56]

Table K.1 provides evidence that the differences in terms of specifications between our paper and Leung (2020) do not lead to major differences in results. As in Leung (2020), the table presents regressions of the effect of the implemented log minimum wage on the log price *level* of grocery stores. The first column shows the estimates for our usual sample and including our standard controls. The estimated price elasticity is similar to our preferred elasticity but remarkably less precise—in fact statistically insignificant—suggesting that, at least in our data, the first-difference specifications is more efficient than the level specification. Column (2) adds two control variables that Leung includes in his baseline specification of Leung (2020): the log of the county population and the average county-level wage (at quarterly frequency, taken from the QCEW). As in Leung (2020), these controls barely affect the estimated price elasticity. Column 3 adds linear, store-specific time trends. These trends are the equivalent to the store FE contained in our baseline first-differenced model. Again, their inclusion matters little for the estimated price elasticity (but increases precision). Columns 4–6 include the *legislated* minimum wage into the regressions. As is the case in our main regressions, we find a positive effect of the

[56] In fact, if one focuses on the 2006–2012 period and weights the estimates presented in Table 7 in Leung (2020) with the number of minimum wage events per period, the implied price response is remarkably close to our baseline price elasticity of 0.036.

minimum wage on grocery prices at the time of legislation, independent of whether we control for store trends or not or whether we use regular prices or prices unadjusted for temporary price changes (e.g. sales). In the latest version, Leung (2020) finds that the price effect of the minimum wage hike is concentrated around legislation in the Nielsen data, too.

Table K.1: Main regressions using a regression model in price levels

	(1) Price lvl	(2) Price lvl	(3) Price lvl	(4) Price lvl	(5) Price lvl	(6) No sales filter
mw	0.039	0.039	0.018	0.022	0.026	0.000
	(0.039)	(0.038)	(0.019)	(0.029)	(0.026)	(0.021)
leg				0.031	0.032*	0.036**
				(0.020)	(0.018)	(0.015)
Unemployment rate	0.000	-0.000	-0.000	0.000	0.000	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
House prices (log)	0.061***	0.065***	0.014**	0.061***	0.055***	0.027
	(0.020)	(0.020)	(0.007)	(0.020)	(0.018)	(0.019)
County population (log)		-0.071				
		(0.048)				
Avg. county wage (log)		-0.026				
		(0.027)				
Observations	222166	222046	222166	222166	222166	222166
Time FE	YES	YES	YES	YES	YES	YES
Store FE	YES	YES	YES	YES	YES	YES
Store trends	NO	NO	YES	NO	YES	YES

Notes: The dependent variable is the store-level price level. *mw* is the binding minimum wage. *leg* is the highest future minimum wage set in current law (the legislated minimum wage). The sample period is 2002–2012. The baseline controls are the county unemployment rate and state-level house prices. Column 2 additionally controls for log county-level population and the quarterly county-level average wage as in Leung (2020). Column 6 is based on a price series that does not correct for temporary price changes (e.g. sales). SE are clustered at the state level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Our results are harder to compare to those of Ganapati and Weaver (2017), as their empirical approach differs from ours and Leung (2020). Instead of constructing store-level price indices, they draw a 1% sample of 5,000 unique products from their data, collapse prices to the county-product level, and estimate the effects with county-product combinations as their unit of observation.[57] Overall, however, our results are not inconsistent with theirs although they

[57] Ganapati and Weaver (2017) and Leung (2020) discuss the advantages and disadvantages of using indices versus product-county level prices in the appendices of their papers. There are two main reasons why we chose to estimate our models at the store rather than at the product level. First, we view it as ex ante desirable to weight products by their importance to both consumers and grocery stores. Second, entry and exit rates at the product-store level are very high in retail, since low-volume products are frequently introduced and discontinued, and may also go unsold in for extended time periods due to stock-outs, seasonality or low demand. Using revenue weights in index construction assigns low weights to products that are likely to exit or to have frequent gaps in their price series. In general, entry and exit is much less pronounced at the store level.

find no impact of minimum wages on prices. As shown in columns 1–3 of Table K.1, we also find no statistically significant effect of the level of the minimum wage on the level of grocery prices if we focus solely on the effects at implementation. Moreover, all of their specifications include product-time fixed effects. As many grocery products are chain- or region- specific, their baseline specification thus likely absorb some variation in prices caused by increases in costs of goods sold. Hence, their baseline results are probably best compared to our estimates that largely absorb this variation. These are the specifications that control for division-time and/or chain-time FE. The estimated price elasticities are smaller in these specifications in our data (see the discussion in section 6.3).

In addition to the differences in specification and estimation strategy, we document the channels through which grocery and drug stores adopt a forward-looking pricing decision, and provide a set of explanations for this behavior (see section 5). We view this piece of work as the main substantial difference between our work and these two studies.