

# Minimum wage effects on employment, substitution, and the teenage labor supply: Evidence from personnel data

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*Abstract:* Using personnel data from a large U.S. retail firm with more than 700 stores nationwide, this study exploits geographic variation in initial wage levels to estimate the employment effects of the 1996 federal minimum wage increase. It focuses on the differential effects for teenagers and adults. The main findings are, first, that increases in average wages had negative, but small and statistically insignificant effects on the full-time equivalent level of employment. Second, in stores where the legislation led to larger increases in the relative wages of teenagers, there was a relative *increase* in the rate at which teenagers were employed, and a relative decline in the employment of young adults (ages 20-22). While small, these positive effects are statistically significant and fairly robust. Third, the explanation for these opposing results is that the overall employment effect can be decomposed into two countervailing effects. To the extent that the legislation increased the wages of all employees without compressing their distribution, there is a negative effect on the overall employment of both teenagers and adults. But to the extent that the legislation increased the relative wages of teenagers, it significantly increased the level of teenage employment, and thus the level of overall employment.

The results show the importance of distinguishing between subgroups of low-wage workers, and contradict the neoclassical prediction that firms respond to increases in the minimum wage by substituting away from workers whose wages increase the most. In markets for teenage workers where the relative wages of teenagers are initially low, the results fit better with more recent models that suggest positive employment effects can result from increased labor market participation—e.g., models incorporating search costs or informational asymmetries. Further analysis supports the idea that higher relative wages caused more teenagers to enter the market. First, the increase in teenage employment is driven largely by a group that is likely to have had relatively less to gain from employment—those living in high-income ZIP codes. Second, an analysis of new hires shows the same pattern of compositional changes found in employment rates. Further analysis also suggests that informational asymmetries are a factor in the low-wage, teenage labor market. An analysis of dismissals and sales shows that the high-income teenagers who are driving the increases in teenage employment are indeed more productive than low-income teenagers who are paid similar wages.

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How do minimum wages affect employment? The current theoretical work on this question is ambiguous. The standard neoclassical theory predicts that a firm will respond to a rise in the minimum wage in two ways—it will cut employment, and it will substitute high-skilled labor for the labor of less-skilled workers whose wages have risen. However, recent models have suggested that minimum wages can lead firms to increase employment. These models incorporate a variety of market frictions, including search costs (e.g., Ahn, Arcidiacono and Wessles, 2005; Flinn 2006), informational asymmetries (Drazen, 1986), and efficiency wages (Rebitzer and Taylor, 1995). In general, these models suggest that employment effects may depend on the group of workers affected and on the specifics of particular labor markets.<sup>1</sup>

Given the nature of the theoretical ambiguities, there is a need for empirical studies that contain enough detail both to analyze effects at the establishment level and to test for heterogeneous effects on different groups of workers. Prior studies have lacked such detail. On one hand, establishment-level studies typically have not had enough information on employees to estimate separate effects for different types of workers. On the other hand, studies of specific groups of low-wage workers (mainly teenagers and sub-groups of teenagers) have relied on household survey data, and have been unable to examine changes in relative wages and employment within firms.<sup>2</sup>

Using personnel data from a large U.S. retail firm with more than 700 stores nationwide, the present study estimates the employment effects of the 1996 federal minimum wage increase. To identify these effects, the research design exploits geographic variation in initial wage levels. Because teenagers at this firm earn less than adults in the same job, the study focuses on the

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<sup>1</sup> For a survey of the theoretical literature on employment effects of the minimum wage, see Neumark and Wascher (2008).

<sup>2</sup> See the review of empirical work in the next section.

differences between teenagers and adults in wage and employment effects. The key advantage of the dataset is that it allows precise measures of wage and employment changes both for a store's workforce as a whole and for different groups of workers within a store. A second advantage is that because the data begins eight months before the change in the minimum wage and because changes in employment are measured daily, this study can assess whether variation in post-legislation employment growth is related to pre-existing trends.

The results show the importance of distinguishing between subgroups of low-wage workers, and suggest standard neoclassical predictions may not hold for teenage workers in markets where the relative wages of teenagers are initially low. In such markets, the results fit better with newer models that link labor demand to labor market participation.

First, legislation-induced increases in average wages had negative, but statistically insignificant effects on the full-time equivalent level of employment. Estimates imply labor demand elasticities ranging from -0.09 to -0.79. Though the estimates are imprecise, their sign is consistent with standard neoclassical predictions.

Second, however, in stores where the legislation led to larger increases in the relative wages of teenagers, there was a relative *increase* in the rate at which teenagers were employed, and a relative decline in the employment of young adults (ages 20-22). The estimated effects are small, but are statistically significant and robust to various model specifications. They imply that a one percent increase in the relative wage of teenagers led to 0.6 to 0.9 percentage point increase in the teenage share of employment. This finding contradicts the prediction of neoclassical theory that firms respond to increases in the minimum wage by substituting away from workers whose wages increase the most.

Third, the explanation for these contradictory results is that the overall employment effect

can be decomposed into two countervailing effects. To the extent that the legislation increased the wages of all employees without compressing their distribution, there is a negative effect on the overall employment of both teenagers and adults. But to the extent that the legislation increased the relative wages of teenagers, it significantly increased the level of teenage employment, and thus the level of overall employment as well.

Because the standard model cannot explain why a rise in the relative wages of teenagers can cause an increase in teenage employment, the last section of this paper considers two types of models that could explain this finding. In both types, positive employment effects are an outcome of an increase in labor market participation. Hence both models could explain the results of the present study if in markets where initial wages for teenagers are relatively low, the rise in relative wages of teenagers caused more teenagers to enter the labor market.

First, positive employment effects for teenagers could be explained by a set of models that incorporate search costs for both job-seekers and employers (e.g., Ahn, Arcidiacono and Wessels, 2005; Flinn 2006). In these models, minimum wages induce individuals to enter the labor market and search for jobs. This in turn reduces the cost of filling vacancies for firms and may induce firms to create more job openings. Second, an alternate explanation is provided by an adverse selection model (Drazen, 1986) in which informational asymmetries prevent employers both from conditioning wage offers on productivity and from attracting higher-quality applicants with higher wage offers. Here, a minimum wage may increase labor demand because it induces labor market entry of relatively productive workers, and this raises the average productivity of the applicant pool.

Further analysis provides support for the hypothesis that the minimum wage caused more teenagers to enter the labor market. First, the increase in teenage employment was driven mainly

by a group that is likely to have had relatively high opportunity costs and low benefits from employment—those living in high-income ZIP codes. The assumption that this group gains less from employment is supported by the fact high-income teenagers receive starting wages similar to low-income teens, but are more likely to quit their jobs in order to return to school. Second, an analysis of new hires shows the same pattern of compositional changes that is seen in employments rates. This suggests the compositional changes in employment are driven not simply by differential changes in turnover rates, but by changes in hiring patterns that could reflect changes in the labor pool.

While both the search and adverse selection models assume an increase in labor market participation, a key assumption that distinguishes the latter model is that the new labor market participants are relatively productive. In an analysis of dismissals and sales, I find evidence that the high-income teenagers who are driving the increases in teenage employment are indeed more productive than low-income teenagers who are paid similar wages. Hence, while acknowledging the potential importance of search costs, I conclude that informational asymmetries are also likely to be a factor in the market for low-wage, teenage labor.

## **1. Related Empirical Literature**

This paper builds on and extends existing research in a number of ways. Because it employs establishment-level data within one low-wage industry, it is similar to the studies of fast-food establishments conducted by Katz and Krueger (1992) and Card and Krueger (CK, 1994). Katz and Krueger (1992) use surveys of restaurants in Texas before and after the federal minimum wage increase and exploit geographic variation in the law's impact on wages. Card and Krueger (1994) use surveys of restaurants in New Jersey and Pennsylvania before and after New Jersey's 1992 minimum wage increase. They identify the law's effects both by using stores

in Pennsylvania as a control group and by exploiting variation in the law's impact among New Jersey stores. Both studies find positive but often insignificant employment effects.

These controversial results prompted several follow-up studies of restaurant employment. Neumark and Wascher (NW, 2000), and Card and Krueger (CK, 2000) both revisit CK's (1994) analysis with new employer-reported data. They reach different conclusions, with NW finding small negative employment effects and CK finding effects that are sometimes positive but typically small and insignificant.<sup>3</sup> Two recent studies (Dube, Lester, and Reich, 2008 and Addison, Blackburn, and Cotti, 2008) analyze nation-wide, county-level panels of earnings and employment in the restaurant industry, using various methods of controlling for unobserved geographic heterogeneity in employment trends. In both studies, the preferred specifications show small, negative but statistically insignificant employment effects.

The present study differs from the restaurant studies in that it focuses on a different low-wage retail industry.<sup>4</sup> But more importantly, the personnel data employed here allow a detailed examination of compositional changes in employment that has not been possible with the data used in earlier studies.<sup>5</sup> This is important because, as often noted, small changes in overall employment levels might mask offsetting changes for different groups of workers. This is especially likely if there is sufficient substitutability between more productive, high-wage workers (e.g. adults) and less-skilled, low-wage workers (e.g. teenagers).

Because teenagers comprise a low-skilled group whose relative wages are likely to be affected by minimum wages, several studies have focused on the employment affects for

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<sup>3</sup> NW (2000) use payroll data collected by themselves and the EPI and attribute the difference between their own findings and those of CK (1994) to advantages of their payroll data over CK's survey data. CK (2000) reanalyze NW's data and repeat their own analysis using employer-reported data from the Bureau of Labor Statistics (BLS). In both analyses, CK again find no evidence that the wage minimum reduced employment, and they attribute NW's negative estimates to the fact that NW's sample of Pennsylvania stores is not representative.

<sup>4</sup> I cannot identify the specific industry for reasons of confidentiality.

<sup>5</sup> Card & Krueger (1994) did examine changes in the fraction of employees that is fulltime, but their results are inconclusive due to large standard errors.

teenagers. But in contrast to the studies of restaurant employees, studies of teenagers have generally relied data from household surveys such as the Current Population Survey (CPS). As a result, they cannot examine effects on the relative wages and the composition of employment within firms.

As with the previous set of literature, studies of teenagers have reached mixed conclusions, and again early studies by CK (1992, 1994) and NW (1992, 1994) fed opposite sides of the debate. In these studies, both sets of authors use state panels of aggregated CPS data and exploit changes in state minimum wages to identify the effects on teenage-to-population employment rates. But they differ in their model specifications and in their methods for measuring certain key variables, including the impact of the minimum wage increase on the relative price of teenage labor.<sup>6</sup> As a result, NW find small negative employment effects that are consistent with substitution away from teenagers, but CK find positive or statistically insignificant effects. Results from later studies have also been mixed. Lang and Kahn (1998) find evidence that the 1991 federal minimum wage increase caused employment shifts away from adults and *toward* teenagers in food-service occupations. Recent studies that use longer panels find small negative employment effects for teenagers that become statistically insignificant after controlling for unobserved heterogeneity in regional trends (Neumark and Wascher, 2007; Allegretto, Dube and Reich, 2008).

A few studies have examined the possibility that the absence of large employment effects for teenagers as a whole may mask more refined compositional shifts *among* teenagers. These studies provide some evidence of heterogeneous effects. Neumark and Wascher's (1996) analysis of matched CPS surveys shows that an increase in the minimum wage increases the

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<sup>6</sup> Another key disagreement between CK and NW revolves around how best to measure school enrollment rates and the appropriateness of controlling for this variable.

probability that older teenagers leave school for employment and that younger, previously employed teenagers become unemployed. Ahn, Arcidiacono and Wessels (2008) use CPS data to estimate a structural model and their findings are similar, suggesting substitution away from teenagers who live in poorer, less educated households and toward more affluent teens. On the other hand, Neumark and Wascher (2007) examine minimum wage effects for various subgroups of teenagers and young adults, and find significant disemployment effects for only minority young adults.

Only one study I am aware of examines the effect of a wage floor on compositional changes in employment within establishments. This study, by Fairris and Bujanda (2008), estimates the effects of the 1997 Los Angeles Living Wage Ordinance on changes in employee characteristics at a sample of affected establishments. Using worker surveys, they compare the characteristics of incumbent workers to those of new (post-ordinance) hires, and find evidence of substitution toward groups who commanded higher pre-ordinance market wages. In contrast to the present study, Fairris and Bujanda do not look explicitly at teenagers (presumably because of the age composition of their sample) and are also unable to estimate effects on overall levels of employment.

## **2. Data and Setting**

The data set is constructed from the personnel records of a large national retail employer from February 1, 1996, through July 31, 1998.<sup>7</sup> The analysis sample consists of more than 700 stores located throughout the United States. This sample includes all retail stores that had been open for at least four months at the start of the sample period and that had an average at least five

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<sup>7</sup> I have been given permission to use the data on the condition that I do not disclose any information (such as exact sample sizes) that may allow the firm to be identified.



employees, including at least one adult and one teenager, during the first six months.<sup>8</sup> Though geographically diverse, these stores are all part of a national chain and hence in many ways they are very similar. The products they sell are highly uniform, and roughly 90 percent of their employees occupy the same frontline, entry-level positions.

This study focuses on employment in these entry-level jobs. These positions all have the same job description: employees in these jobs all rotate through several tasks that involve both dealing with customers and doing support work. These jobs require only basic skills and employees receive little training. As is common in this sector, employees have very high rates of turnover. The median spell in a store for an entry-level employee is 91 days, and roughly 80 percent of employee spells end within a year.

At each store, a single, overall store manager is responsible for all personnel decisions. Manager compensation is tied to performance mainly through bonuses that are based on store sales. Managers may adjust wages and wage offers as they see necessary to retain and attract qualified employees. However they must also manage to a year-end wage budget, which is based partly on market wages in their store's region, and they receive small bonuses for meeting goals for the ratio of payroll to sales.<sup>9</sup> There are 41 company-defined regions in the sample, with an average of 23 stores per region, and the regression analysis below includes tests for sensitivity to the inclusion of region fixed effects.

The personnel records contain information on every individual employed at one of the company's retail stores during the 30-month sample period. This information includes

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<sup>8</sup> Sufficient employment during the first six months is needed to construct reasonably precise measures of the legislation's impact on wages. New stores are excluded both because they tend to have few employees on record during the first few months and because employment growth during these months is highly variable and may depend on many unobserved factors. I also do not analyze the handful of stores that close during the sample period because there are too few of these to allow for a meaningful analysis of store closings.

<sup>9</sup> Unfortunately, I do not have access to the formulas used in determining each store's wage budget.

employment status (full-time vs. part-time), wage, age, race, gender, residential ZIP code, store of employment, and the date and description of each personnel action taken. These records are used to construct daily store-level employment variables, including a “full-time equivalent” measure of employment, the fraction of employees who work part-time, and the fraction who are teenagers. Because there is no information on hours worked, “full-time equivalent” employment is calculated under the assumption that part-time employees work half as many hours as full-time employees.<sup>10</sup> Additional store-level variables, obtained from company records, include the store’s size (square feet), its ZIP code, city and state, and its location “type” (indoor mall, open mall, street, or strip). The data was also merged with 1990 Census-based variables describing the population within a two-mile radius of each store’s ZIP code, and with local unemployment rates averaged over the first six months of the sample period (Feb.-July 1996).<sup>11</sup>

Table 1 shows sample statistics for the store-level variables during the first six months of the sample period, including store-level employment and average employee characteristics. During this period, a typical store has about 27 employees. On average, 94 percent of these employees are part-time and average full-time equivalent employment is 14.7. The workforce is young: more than 40 percent of employees are teenagers, and more than 80 percent are less than 30 years old. And the workforce is mainly female (77 percent) and white (72 percent vs. 11 percent black and 9 percent Hispanic).

### **3. The 1996 Minimum Wage Legislation and Measures of its Impact on Wages**

#### *3.1. Legislation*

The analysis focuses on the federal minimum wage law that was enacted in August 1996.

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<sup>10</sup> Full-time status required a minimum of 30 hours per week. Part-time employees were required to work a minimum of eight hours per week and a maximum of 29 hours.

<sup>11</sup> Unemployment rates are from the Local Area Unemployment Statistics of the Bureau of Labor Statistics, and are based on metropolitan areas as defined in that data set.

For employers, the first key date regarding the law was likely July 10, 1996. Though the original bill was passed by the House of Representatives on May 24, its fate remained uncertain until July 10 when the Senate passed an amended bill that was supported by President Clinton. The final bill was passed by the House on Aug. 2, and became law on Aug. 20.

The law mandated a 21 percent increase in the minimum wage—from \$4.25 to \$5.15—that was to be implemented in two steps. First, the minimum would rise to \$4.75 on Oct. 1, 1996; it would then rise to \$5.15 on Sept. 1, 1997. The law also included a “training wage” provision that kept the minimum at \$4.25 for teenagers during their first 90 days on the job.

While the federal law is the focus of this study, several state minimum wage laws are also relevant to the analysis—either because they contribute to the geographic variation in initial wage levels or because they result in additional wage increases during the sample period. Table 2 shows the effective minimum wage in each month of the sample (Feb. 1996-July 1998) for states where the state minimum exceeded the federal minimum during some or all of the sample period.<sup>12</sup> These states are grouped into three categories. First, there are states where existing state law required that the state minimum stay above the federal minimum, and hence mandated that the state minimum rise in tandem with any federal increases. This category accounts for only three percent of the sample stores, and is comprised of Connecticut, Alaska, and Washington, D.C.<sup>13</sup>

Second, there are five states where the state minimum exceeded the federal minimum of \$4.25 at the start of the sample period, but where there was no independent increase in the state minimum. Stores in these “high initial minimum” states account for 8.6 percent of the sample,

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<sup>12</sup> States may set their own minimum wages; however, the federal minimum is legally binding in the event that the state minimum is lower than the federal.

<sup>13</sup> In Connecticut, the state minimum was kept at only ½ percent above the federal level. The state minimum was kept much higher in Alaska and Washington, D.C. (at \$.50 and \$1.00 above the federal rate, respectively). However, Alaska and D.C. are unimportant in the analysis because they have very few stores in the sample.

and are comprised of Rhode Island, Iowa, New Jersey, Washington, and Hawaii. In Rhode Island and Iowa, the new federal minimum surpassed the state level and thus became binding in Oct. 1996 when it rose to \$4.75. In New Jersey and Washington, the federal law became binding only in Sept. 1997 when it rose to \$5.15. And in Hawaii, where the initial minimum was \$5.25, the federal minimum was never binding during the sample period.

Finally, five states—California, Delaware, Massachusetts, Oregon, and Vermont—increased their own minimum wage on a schedule that differed from that of the federal legislation.<sup>14</sup> Stores in these states account for 20 percent of the sample—with roughly 14 percent being in California alone. Notably, California had the largest total increase of any state during the sample period, with the minimum rising 35 percent from \$4.25 to \$5.75. Because of differences in the timing of wage increases in these states, much of the analysis excludes stores in these states. However, all results are checked for robustness to their inclusion in the estimation sample.

### *3.2. Compliance with the Law*

The personnel records indicate full compliance with the law at this company. The lowest wage on record rises from \$4.25 to \$4.75 on October 1, 1996, and from \$4.75 to \$5.15 on September 1, 1997. In states with higher minimums, the lowest wages on record are also consistent with the laws. Interestingly, the fact that \$4.75 was the company's lowest wage as of October 1, 1996 indicates that it did not utilize the "training wage" provision of the federal legislation.<sup>15</sup>

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<sup>14</sup> In all these states except Vermont, the relevant legislation was passed by the end of 1996; hence all the minimum wage increases scheduled for 1997 and 1998 were anticipated. The Vermont law (as of May 1996) approved minimum wages of \$5.00 and \$5.15 effective January 1997 and January 1998. The law was amended in April 1997, changing the schedule to that shown in Table 2.

<sup>15</sup> Card & Krueger (1994), pp. 166-68, provide several sources of evidence showing that employers in the early 1990s rarely utilized the subminimum wage allowed by the 1989 amendment to the Fair Labor Standards Act.

An examination of the wage adjustments made on October 1, 1996, and on September 1, 1997, also reveals evidence of “spillover” effects on higher-wage employees. First, employees whose wages were below the new minimum but above the old one received adjustments that were greater than necessary to meet the new minimum. Second, some employees whose wages exceeded the new minimum also received adjustments. In particular, wage adjustments were given on October 1, 1996 to all employees earning less than \$5.45, and adjustments were given on September 1, 1997 to employees earning less than \$5.65. Further, these adjustments were made even in stores that had no employees earning less than the new minimum, including those in states with high initial minimum wages.

### *3.3. Measures of the Legislation’s Impact on Wages*

The research design exploits the fact while the stores in my sample are all part of the same national firm, there is significant variation across stores in the impact that the minimum wage legislation has on entry-level wages. There are two basic sources of this variation. First, the impact of both the federal and state-legislated increases varies across stores, because the level of wages paid *prior* to the legislation varies across stores. Second, the effective increase in the minimum wage varies across states because of the differences in state laws (summarized above).

Table 3 summarizes store average hourly wages in the pre-legislation period from Feb. 1, 1996 to July 31, 1996.<sup>16</sup> The mean of this variable is \$5.65 for stores in states where the federal minimum wage was the effective minimum in all months of the sample period (column 1).

There is substantial variation within this sub-sample (std. dev. = \$0.51). Further, average wages are higher in states with high initial minimums and in California (columns 2-4).

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<sup>16</sup> The data set contains each employee’s dates of employment, wages paid, and dates of wage changes, but not hours worked. To calculate the average hourly wage, I assumed that full-time employees worked twice the hours of part-time employees each day, and that hours were distributed evenly across all days that the employee was on the payroll.

Average wages are also constructed for three subgroups of employees within each store: part-time employees, teenagers, and adults. Full-time workers earn roughly 30 percent more than part-time workers in the same store; hence average wages tend to be slightly lower when the sample is restricted to part-time employees.<sup>17</sup> There is also a significant difference between the wages of teenagers and those of adults (defined here as anyone who is at least 20 years old). In the pre-legislation period, teenagers earn an average of 90 cents for every dollar earned by adults.

Table 3 also summarizes the variables used in the analysis to measure the impact that the federal and state legislation had on wages. These variables are constructed using the wages paid in the first six months (Feb. 1-July 31, 1996) and the ultimate minimum wage mandated by the 1996 laws. With the exception of stores in Vermont, California, and Oregon, this ultimate new minimum is the minimum effective as of September 1997—that is, \$5.15 in states that are bound by the federal minimum, \$5.18 in CT, \$5.25 in MA and HI, \$5.65 in AK, and \$6.15 in DC. For stores in VT, the ultimate minimum is \$5.25 (effective Oct. 1, 1997); for stores in CA, it is \$5.75 (effective March 1, 1998); and for stores in OR it is \$6.00 (effective Jan. 1, 1998).<sup>18</sup>

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<sup>17</sup> I do not report or analyze full-time wages separately because many stores have no full-time employees in the entry-level positions.

<sup>18</sup> An alternate approach—at least in the case of the federal legislation—would be to construct separate impact measures for the October 1996 increase and the September 1997 increase. The problem with treating these as two separate increases is that they were both anticipated as of August 1996, and hence stores may have adjusted both wages and employment gradually after October 1996 in anticipation of the second increase. Indeed, the fact that “spillovers” in October 1996 were relatively large suggests that stores got a head start on the second increase. For this reason, my analysis focuses on changes from the period before August 1996 to the period after September 1997.

There are two other issues that one might debate regarding how to define the effective new minimum wage. One is whether to acknowledge the training wage provision for teenagers employed less than 90 days. I chose to ignore this provision both because the data indicates that the company did not take advantage of it, and because previous literature suggests that this company’s behavior is common (e.g. see Card & Krueger, 1994, pp.166-68). The other issue is how to treat spillover effects on higher-wage employees. The fact that the company raised wages of those earning below \$5.65 might suggest an effective new minimum above \$5.15 for some employees. However, although the company did raise the wages of higher-wage employees who were present on the days that the minimum wage increases took effect, it was not compelled by law to maintain the wage distribution that resulted from this initial spillover effect. For this reason, I do not incorporate observed spillover effects when constructing the measure of the legislation’s impact on wages. Hence, with the caveat that I ignore the training wage, the impact measures capture the cost to the company of complying with the law, and nothing more.

In the analysis of average entry-level wages and employment, the legislation's impact on average wages is measured using the "store wage gap." This variable is defined as the average proportional increase necessary to bring all wages up to the new minimum.<sup>19</sup> To construct it, I first define the individual wage gap for employee  $i$  in store  $j$  as the proportional increase in the employee's wage ( $w_{ij}$ ) necessary to meet the ultimate new minimum in store  $j$ 's state (*minimum wage<sub>j</sub>*). That is:

$$\begin{aligned} \text{wage gap}_{ij} &= (\text{minimum wage}_j - w_{ij}) / w_{ij} && \text{if } w_{ij} < \text{minimum wage}_j \\ &= 0 && \text{if } w_{ij} \geq \text{minimum wage}_j \end{aligned}$$

The store wage gap is then calculated as the average wage gap for wages paid to entry-level employees in a store between Feb. 1 and July 31, 1996.<sup>20</sup> Because wages differ significantly by full-time vs. part-time status and by age, separate store wage gap measures are also constructed for part-time employees, teenagers, and adults.

The summary statistics in Table 3 show, first, that while the direct impact on average wages at a typical store is modest, there is substantial variation across stores. Among stores where the federal minimum was always binding (column 1), the mean store wage gap is 2.5 percent but the standard deviation is 3.1 percent. Indeed, stores in the bottom quartile of this group have store wage gaps of less than half of one percent, while the mean among the top quartile is 7.0 percent. Further, the store wage gap is also substantially smaller in states with high initial minimums where the mean is 0.2 percent, and is much higher in California stores where the mean is 5.2 percent.<sup>21</sup>

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<sup>19</sup> This measure is similar to that used by Katz & Krueger (1992) and by Card & Krueger (1994).

<sup>20</sup> The average is calculated in a fashion similar to that described in footnote 14 above; full-time wages receive twice as much weight as part-time wages, and hours are assumed to be distributed evenly across all days that an employee was on the payroll at a given wage.

<sup>21</sup> A similar pattern emerges if one looks at an alternate measure of the legislation's impact on wages: the fraction of a store's employees whose wages are affected by the legislation. For the sample as a whole, the average "fraction affected" is roughly 48 percent, and the standard deviation is 21 percentage points. The fraction affected is highest

As expected, the store wage gap tends to be slightly higher when the focus is restricted to part-time workers. But because part-time workers make up the vast majority of the workforce, the part-time measures are not much different from the overall impact measures. Also as expected, teenagers have larger wage gaps than adults. Among the stores bound by the federal minimum, for example, the average store teenage wage gap is 3.7 percent while the average adult wage gap is 1.6 percent.

Because the minimum wage legislation raised teenage wages more than it raised adult wages, it caused an increase in the relative wage of teenagers in most stores. To measure the legislation’s impact on *relative* wages, I construct a variable called the “store relative wage gap.” This variable is again based on wages paid between Feb. 1 and July 31, 1996, and is defined as the proportional change in the relative wage that results from bringing all wages up to the ultimate new minimum. The store relative wage gap is constructed from the store wage gaps for teenagers and adults as:<sup>22</sup>

$$\text{store relative wage gap}_j = \frac{(\text{store teenage wage gap}_j - \text{store adult wage gap}_j)}{(1 + \text{store adult wage gap}_j)}$$

In the last row of Table 3, we see that compliance with the law caused on average a two percent increase in the relative wage of teenagers. However, the size of the relative wage gap varies significantly across stores, with a standard deviation of 2.7 percent. Also, it is close to zero in states with high initial minimums and largest in stores located in California.

#### 4. Methods of Analysis

The goal of the analysis is to determine the extent to which differences across stores in

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in California (59 percent) and lowest in states with high initial minimums (27 percent). It is also higher among part-time employees (49 percent) and teenagers (53 percent) compared to fulltime employees and adults (43 percent).

<sup>22</sup> The relative wage in store  $j$  prior to the legislation is  $\text{store relative wage}_{j1} = (\text{store average teenage wage}_j) / (\text{store average adult wage}_j)$ . After wages are adjusted to comply with the new minimum, the relative wage becomes:  $\text{store relative wage}_{j2} = (\text{store average teenage wage}_j)(1 + \text{store teenage wage gap}_j) / (\text{store average adult wage}_j)(1 + \text{store adult wage gap}_j)$ . The  $\text{store relative wage gap}_j$  is then:  $(\text{store relative wage}_{j2} - \text{store relative wage}_{j1}) / (\text{store relative wage}_{j1})$ , which can be expressed more simply by the above expression.



post-legislation employment growth and changes in the relative employment of teenagers can be attributed to differences in the legislation's impact on average and relative wages. As in most minimum wage studies, a key challenge to identifying the legislation's causal effect lies in controlling for differences in initial market conditions that could be correlated with both the difference between current wages and the new minimum and with subsequent changes in wages and/or employment. I approach this problem by supplementing regression analysis with two complementary graphical analyses that aid in interpreting the regression estimates.

The basic estimates are obtained from store-level regressions relating the legislation's impact on wages, as measured by the "wage gap" variables defined above, to changes in the outcome variables of interest. The regression equations take the form:

$$(1) \quad \Delta Y_j = \alpha + \beta \cdot X_j + \gamma \cdot \text{wage gap}_j + \varepsilon_j$$

where  $\Delta Y_j$  is the change in the wage or employment outcome of interest for store  $j$ ,  $X_j$  is a set of characteristics of store  $j$  or its location, and  $\text{wage gap}_j$  is either the store average wage gap or store relative wage gap. Because the sample period begins six months before the legislation was passed, changes are calculated using the first six months (Feb. 1, 1996-July 31, 1996) and last six months (Feb. 1, 1998-July 31, 1998) of the sample period. I examine the sensitivity of these regression estimates to the inclusion of an increasingly detailed set of control variables, including the age initial distribution of employees (whose wages are used to construct the wage gap variables), region fixed effects (based on regions defined by the company), and the other store characteristics described in Table 1. I also test for robustness to the inclusion in the estimation sample of states where state minimum wage increases are not synchronized with the federally mandated increases.

There are two important limitations of this regression approach. First, it cannot rule out

endogeneity of the legislation's impact on wages. Of particular concern is the possibility that differences in initial wage levels (and hence in the wage gaps) are driven by unobserved, pre-legislative market conditions, and that these conditions, in turn, caused wage and/or employment trends to differ even before the legislation was passed. Second, the regression estimates reveal nothing about the timing of the adjustments. To address both concerns, I perform two complementary graphical analyses. For these analyses I construct a panel data set containing store averages of the relevant variables for each of the 30 months in the sample period. These analyses exclude states where the legislation is not synchronized with the federal law.

In the first graphical analysis, the sample is divided into "high-impact" and "low-impact" stores; these categories are defined using the median of the relevant wage gap variable as a cutoff. Then monthly averages of the relevant wage and employment variables are plotted for each group of stores. The resulting graphs show the overall time-series patterns of the wage and employment variables of interest and allows for a crude comparison between high and low-impact stores.

The second graphical analysis is based on estimates from store-level regressions. For each month of the sample period, I estimate equations of the form:

$$(2) \quad Y_j = \alpha + \beta \cdot X_j + \gamma \cdot \text{wage gap}_j + \varepsilon_j$$

where  $Y_j$  is the level of the wage or employment outcome of interest for store  $j$ ,  $X_j$  is a complete set of store characteristics including region dummies, and  $\text{wage gap}_j$  is either the store average wage gap or store relative wage gap. I then plot, by month, the wage gap coefficients (i.e. the estimates of  $\gamma$ ) and their 95 percent confidence intervals. These coefficients are multiplied by .01 so that they measure, at each point in time, the difference in the outcome variable associated with a one percentage point increase in the wage gap.

The resulting graphs provide a way to assess both the exogeneity of the wage gaps and the magnitude and timing of any responses to the legislation. First, if the wage gap is indeed exogenous to unobserved determinants of wage and employment growth, then the estimates of  $\gamma$  should be zero—or, at least, shown no trend—during the six months before the legislation was passed. Second, responses to the legislation are reflected in the time series pattern of these coefficients during the months after the legislation was passed.

## 5. Main Results

### 5.1. Changes in Average Wages

Table 4 shows the regression estimates of the legislation's effect on store average wages. The estimates indicate that a one percentage point increase in the store wage gap corresponds to roughly a .77 percentage point increase in the growth rate of wages over two years. The estimate varies only slightly across different model specifications and estimation samples, ranging from .75 to .78. Robustness to the inclusion of numerous control variables suggests that these estimates reflect the legislation's impact on wages and not differences in market-driven wage trends.

The graphical analyses of average wages are shown in Figures 1a and 1b. Figure 1a shows the overall time-series pattern of wages for high impact stores (which have an average wage gap of 0.042) and low-impact stores (with an average wage gap of 0.003). Both groups show similar upward trends over the sample period, except for the two discrete jumps in October 1996 and September 1997.<sup>23</sup> These jumps, which coincide with the dates of the federally mandated increases, are substantially larger in the high-impact stores, and overall wage growth is higher in the first group as a result of these jumps.

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<sup>23</sup> Figure 1a also shows a small amount of seasonal variation in wages. Wages dip in December and peak in July, and there is a visible increase between June and July of each year. The June-to-July increases are due largely to merit raises, which are made during the last week in June in all stores. For more on merit raises, see footnote 26.

In Figure 1a, the universal upward trend explains why the legislation's estimated impact on wages after two years is roughly 25 percent less than what is implied by the average wage gap. In short, the legislation's impact was weakened somewhat by wage growth that would have occurred in the absence of the law. Further, the fact that trends in the two groups are similar (except at the times of the two legislated increases) supports the interpretation that the regression estimates from Table 4 reflect the legislation's impact and not differences in pre-existing trends.

This interpretation is given further support by Figure 1b, which plots the wage gap coefficients from monthly regressions of average wages that control for all the store characteristics used in Table 4, column 4. Here again, there is no evidence that wages would have grown more quickly in high-impact stores without the legislation. First, the trend in the wage gap coefficients is flat during the pre-legislation period, indicating similar pre-existing trends in high and low-impact stores. Second, there is a slight negative trend after each legislated wage increase, suggesting that wages in higher-impact stores grew a bit more slowly than lower-impact stores following the legislation.

## *5.2. Changes in Full-Time Equivalent Employment*

Table 5 shows the regression estimates of the legislation's effect on two-year changes in full-time equivalent employment. The estimated effect of a one percentage point increase in the store wage gap on the change in employment ranges from -0.01 to -0.09 and becomes increasingly negative in specifications with more controls; but in no case is it statistically significant. The corresponding labor demand elasticities range from -0.09 to -0.79 when evaluated at the sample mean of 14.7 full-time equivalent employees.<sup>24</sup> The fact that the estimates become increasingly negative with more controls in the regression suggests that they

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<sup>24</sup> These elasticities are calculated with respect to the wage increases induced by the legislation using the estimates in Table 4.

may be biased toward zero due to unobserved market differences in pre-existing employment trends. However, the graphical analysis discussed below shows no evidence that this is the case.<sup>25</sup>

Another potential concern with these estimates is the possibility of measurement error in the full-time equivalent employment variable due to the (possibly inaccurate) assumption that part-time employees work half as many hours as full-time employees. This type of measurement error would be especially problematic if the fraction of part-time employees varies across stores or over time. The specification shown in column (5) addresses this concern by controlling for changes in the fraction of employees that is part-time. The coefficient on the store wage gap is very similar to that in column (4), suggesting that differences in the relative growth of part-time employment are not a significant source of bias.<sup>26,27</sup>

Finally, in column (6), the estimation sample is expanded to include stores in states with “unsynchronized” minimum wage increases. Again, there is little change in the estimated coefficient.

The graphical analyses of employment are shown in Figures 1c and 1d. Figure 1c plots by month the average levels full-time equivalent employment for both high-impact and low-impact stores. While high-impact stores have lower initial employment levels, both groups show

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<sup>25</sup> Additional sensitivity tests show that if employment growth is expressed as a *proportional* change in employment rather than the change in levels, the estimated effect of the minimum wage is small and *positive*, but again statistically insignificant. The discrepancy in sign is explained by the fact that higher-impact stores have lower initial levels of employment than lower-impact stores, but similar upward trends in employment levels. As a result, proportional employment growth has a slightly more positive trend in higher-impact stores even before the legislation is passed.

<sup>26</sup> A more in depth analysis of changes in the fraction part-time reveals a small negative correlation between the store wage gap and changes in the fraction part-time. However, this correlation is due entirely to the slight growth in part-time employment in less-impacted stores. In high-impact stores (as defined above), the fraction part-time is initially high—at roughly 96 percent—and remains high throughout the sample period. Hence, it does not appear that the legislation had a significant impact on the ratio of part-time to full-time employees. Complete results are available from the author.

<sup>27</sup> A related concern is that the store may adjust hours worked by each individual instead of adjusting the number of employees. Unfortunately, like many minimum wage studies, the present study cannot address this concern.

an upward trend in employment, and it is difficult to detect any difference in employment growth between the two groups. Instead, what stands out in this graph are the similarly large seasonal employment swings in both groups of stores that dwarf any relative decline in high-impact stores. For example, despite an overall upward trend, employment falls by roughly seven full-time equivalents between December of each year and the following May.

The relationship between the wage gap and the time-series pattern in employment is easier to discern in Figure 1d, which plots the wage gap coefficients from monthly regressions of employment that control for store characteristics as in Table 5, column (4). Here, two patterns are noteworthy. First, the wage gap coefficient is nearly constant throughout the pre-legislation period, suggesting that the wage gap is not correlated with pre-existing trends in employment. This allays the concern that the regression estimates in Table 5 are biased due to unobserved geographic heterogeneity in employment trends. Second, consistent with the regression results, employment growth during the post-legislation period appears to be slightly lower in stores with higher wage gaps.

### *5.3. Changes in Relative Wages*

Because teenagers were typically paid less than adults in the pre-legislation period, the minimum wage legislation tended to increase the relative wage of teenagers. However, because of variation in both the initial level and the initial distribution of wages, the legislation's impact on relative wages varied across stores. This section exploits variation in the *store relative wage gap* to estimate the legislation's effect on the relative wages of teenagers; I then estimate the effect on teenagers' relative employment.

Table 6 shows the estimates from the regression of two-year changes in the teenage relative wage on the store relative wage gap. The estimate from column (1), with no controls,

suggests that a one percentage point increase in the relative wage gap resulted in roughly a one percentage point increase in the relative wage. When region dummies and store characteristics are controlled for in the regression, the estimate increases to around 1.3, suggesting that the impact of the legislation on relative wages may have been greater than what is implied by the relative wage gap. Sensitivity of the estimate to the inclusion of controls also suggests that the relative wage gap is correlated with market differences in the time series pattern of relative wages. These differences are investigated in Figures 2a-2f.

Figures 2a-2c show the grouped analysis of teenage relative wages. Here, “high-impact” stores have a store relative wage gap of at least 0.01, with an average of 0.036, and in “low-impact” stores the relative gap is less than 0.01 with an average of 0.002. Figures 2a and 2b plot the average teenage wage and average adult wage by month for the high and low-impact groups. Teenage wages (Fig. 2a) are flat over the sample period with the exception of jumps in October 1996 and September 1997. The average teenage wage is initially about \$.60 lower in the high-impact group, but due to the minimum wage increases, the difference between the two groups cut in half by the end of the sample period.

Adult wages (Fig. 2b) exhibit upward trends over the sample period in both groups of stores, but there are two notable differences in the time series patterns between these two groups. First, initial wages are lower in high-impact stores, and as a result, these stores have larger jumps in wages in October 1996 and September 1997 and also have slightly higher overall wage growth. Second, there are noticeable wage increases between June and July of each year and these increases are more pronounced in the low-impact stores, especially in 1996. This latter set of wage increases is due to merit raises, which are given during the last week in June. An analysis of merit raises reveals that difference between high and low-impact stores is due

primarily due to the fact that higher-impact stores have higher turnover rates, and thus lower rates or eligibility for merit raises.<sup>28</sup>

Figure 2c plots the relative wage of teenagers (the average teenage wage divided by the average adult wage) for both groups of stores. There is a noticeable seasonal pattern to relative wages in both groups; the relative wage peaks in December, due to a dip in adult wages (see Fig. 2b), and dips in July after merit raises are given (mainly to adults). The merit raise effect is especially large in July 1996 and as anticipated by the pattern in Figure 2b, it causes an especially large drop in the relative wage in low-impact stores. The result is that relative wages in the two groups converge somewhat even before the minimum wage legislation. Nevertheless, as the rest of the graph makes clear, the legislation still had a substantial effect on relative wages. Indeed, while the relative wage began substantially lower in the high-impact stores, it ends up higher in this group after October of 1997.

Figure 2d plots the coefficients from regressions of the teenage relative wage on the store relative wage gap, controlling for the store-level variables used in Table 6 column (4). Here again, we see the effect of the June 1996 merit raises on the teenage relative wage cause a jump in the coefficient. However, the pattern in the five months prior to that shows no evidence that relative wages were following different trends prior to the legislation. Moreover, the pattern in the following 18 months shows large jumps at the dates of the minimum wage increases. In sum, while the graphical analysis cannot rule out the possibility of some bias in the regression estimates, it does confirm that the minimum wage legislation had a substantial impact on relative wages in some stores and that the relative wage gap is a reasonable proxy for the size of the

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<sup>28</sup> Employees must be employed for at least 90 consecutive days to be eligible for a merit raise. The average merit raise is about 2.2 percent, and approximately 80 percent of eligible employees receive one. The average merit raise for eligible adults does not differ significantly between high and low-impact stores. Teenagers in both groups are much less likely than adults to be eligible for a merit raise, they are less likely to receive raises when they are eligible, and they also receive smaller raises on average. Complete results are available from the author.



impact.

#### 5.4. Changes in the Teenage Share of Employment

We now ask whether legislation-induced changes in the relative wage paid to the teenagers led to changes in the employment of teenagers relative to adults. The regression results, shown in Table 7, suggest that the answer is yes; but contrary to the conventional theory, increases in the relative wages of teenagers led to significant *increases* in their relative employment. The estimated coefficients on the relative wage gap range from 0.62 to 0.93. These are small effects, implying that a one percentage point increase in the relative wage gap leads to less than a one percentage point increase in the teenage share of employment. However, the estimates are all statistically significant at a one percent significance level.

The estimated effect varies somewhat across model specifications; it increase with controls for the store's initial age distribution and region dummies (columns 2 and 3) and declines again with the inclusion of other time-invariant store-level variables (column 4). Columns 5 and 6 control for changes in percent part-time and percent single female, respectively—both of which are correlated with changes in the teenage employment share. The variable “change in percent single female” is included to control for compositional changes that may have resulted from the August 1996 welfare reform legislation, which thrust many unemployed single mothers into the labor force.<sup>29</sup> These controls have little effect on the estimated coefficient. There is also very little change in the coefficient when the sample is expanded to include stores in all states (column 7).

The remaining columns in Table 7 show in more detail an increase in the relative wage gap affects the age distribution of employees. Here, we see that two-thirds of the increase in the teenage share of employment is due to an increase in the share of 16-17 year-olds, and that the

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<sup>29</sup> See Blank (2002) for a review of the literature on the effects of welfare reform in the 1990s.

offsetting decline in the adult employment share is driven mainly by young adults, ages 20-22.

The graphical analysis in Figures 2e and 2f confirm that the positive relationship between the relative wage gap and growth in the teenage share of employment is not driven by differences in pre-existing trends. The grouped analysis (Figure 2e) shows that during the first half of the sample period, both high and low-impact stores shown a negative trend in the teenage employment share, but that that teenage employment is falling more quickly in the high-impact stores. After the first minimum wage increase, however, this pattern is reversed. Between the first half of 1997 and the first half of 1998, the teenage employment share increases in the high-impact stores, while appearing to level off (aside from seasonal swings) in the low-impact group.

Figure 2f shows that after controlling for observed differences across stores there is no apparent correlation between the relative wage gap and pre-legislative trends in the teenage share of employment. Specifically, this graph plots the coefficients from monthly regressions of the teenage employment share on the relative wage gap with the control variables used in Table 7, column 4. In the months before the first wage increase, the coefficient shows some slight seasonal variation but no apparent trend. But after the first minimum wage increase in October 1996, the coefficient shows a significant holiday-season spike (in December 1997) and then an upward climb beginning again in June of 1997. Now, the holiday and summer seasons both tend to see increases in teenage employment at all stores (Figure 2e). However, Figure 2f suggests that these surges in teenage employment are especially large where teenagers' relative wages have risen most as a result of the new minimum. Finally, the upward trend in the relative wage gap coefficient continues after the second minimum wage increase in September 1997.

### *5.5. Countervailing Effects of Average and Relative Wage Increase on Employment Levels*

The results thus far indicate that the minimum wage increase had two very different

effects on employment at the stores in our sample. First, by increasing average wages, the legislation had a negative effect on stores' average employment levels. Though the estimated effect is statistically insignificant, the direction of the effect is at least consistent with the prediction of the conventional model. But second, by increasing the relative wage of teenagers, the legislation had a *positive* effect on the relative employment of teenagers. This is inconsistent with the conventional prediction that the largest employment declines should occur among those whose wages increase most. Instead, it suggests that increases in the relative wage may have led to efficiency gains and to a reduction in the effective cost of employing teenagers.

Taken together, these results imply that the estimated employment effects from Table 5 may mask the presence of two competing effects, and that the net effect on employment depends on whether the legislation affected all employees or mainly teenagers. Table 8 shows the results from employment regressions that include the legislation's impact on both average and relative wages. These results confirm that the employment effect can be decomposed into two countervailing effects. First, to the extent that the legislation increased wages of all employees without compressing their distribution, there is a negative effect on the employment of both teenagers and adults (columns 3, 6 and 9). Indeed, after controlling for the relative wage gap in the overall employment regression, the coefficient on the average wage gap is larger and marginally significant ( $p=.07$ ). But second, to the extent that the legislation increased teenagers' relative wages, it significantly increased teenage employment and thus the overall level of employment (columns 3 and 9). Put another way, the negative employment effects were larger in markets where adult wages were closer to the minimum and where teenage wages were already close to the wages of adults. But where the legislation's primary impact of was on teenagers' relative wages, the employment effect was small or even positive.

## 6. Why did increases in teenage relative wages lead to increases in teenage employment?

This section considers why higher relative teenage wages led to higher levels of teenage employment. I focus on two types of model—one employs search costs and the other asymmetric information. Both types of model suggest that a minimum wage could lead to higher employment by inducing an increase in labor market participation. Hence they may be especially relevant to teenagers in markets where teenage wages are relatively low.<sup>30</sup>

Further analysis provides evidence that supports the idea that the minimum wage induced teenagers to enter the labor market. It also suggests that while all teenagers at a given store are paid similar wages, nevertheless the new teenage entrants were relatively productive. This finding provides support for the model with informational asymmetries.

### *6.1. Theories Linking Labor Market Participation to Labor Demand*

One possible link connecting minimum wages, labor market participation, and labor demand is the presence of search costs on the part of both jobs seekers and employers. For example, in models by Flinn (2006) and Ahn, Arcidiacono and Wessels (forthcoming), wages are set by Nash bargaining and a binding minimum effectively raises worker's bargaining power and the resulting wage. This induces more individuals to search for jobs (despite a lower probability of finding a “match” conditional on searching). From the firm's perspective, the increase in the number of searchers increases the probability of filling vacancies, and thus reduces the cost of creating new jobs. If this cost reduction is large enough to offset the wage increase, firms are induced to create more vacancies and thus to increase employment.

A second model that links increased labor market participation to increased labor demand assumes adverse selection in a labor market where individuals' reservation wages are increasing

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<sup>30</sup> Other models predict positive employment effects of minimum wages are based on monopsony-like firm behavior; but it is less clear how the mechanisms in these models could generate differential employment effects for teenagers vs. adults and for high-income vs. low-income teenagers (see below).

in the quality of their labor. Drazen (1986) demonstrates that asymmetric information on the part of both firms and potential job-seekers may lead to an inefficient equilibrium where firms offer low wages and only low-quality workers apply for jobs.<sup>31</sup> First, imperfect information about the quality of individual applicants prevents firms from conditioning wage offers on productivity. Second, potential job seekers have imperfect information about individual firms' wage offers, and so the decision of whether to enter the labor market depends on the average market wage that is observed. This prevents individual firms from attracting high-quality applicants by unilaterally offering higher wages—even if doing so would reduce the effective cost of labor. In this situation, a minimum wage can increase labor demand by inducing an improvement in the average quality of job applicants that more than offsets the wage increase.

The adverse selection model may be especially relevant to the findings of this paper because the informational asymmetries that it rests on are likely to be more severe in a market for teenage labor than in the adult labor market. First, because teenage applicants are more likely to lack an employment history, their productivity may be more difficult to observe. Second, because many non-working teenagers have never searched for a job, they may have little awareness of specific employers' wage offers.

## *6.2. The Effects of Relative Wage Increases on Teenage Labor Market Participation: Evidence on Teenagers from “High-Income” ZIP Codes*

The key feature common to both the search cost and adverse selection models of minimum wage effects is a positive effect on labor market participation. Both models thus suggest that the positive response of teenage employment to increased relative wages operated through an increase in labor market participation by teenagers. These theories, therefore, imply

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<sup>31</sup> Drazen's paper builds on earlier models of asymmetric information in labor markets by Stiglitz (1974), Weiss (1980), and Guasch and Weiss (1980).

other empirical predictions that can be examined with the personnel data. First, increases in teenage employment should be driven by individuals whose opportunity cost is most likely to exceed the benefit of employment at the initial wage. Second, changes in the composition of employment should be caused mainly by changes in the composition of new hires, and not simply by differential changes in turnover rates.

To test the first prediction, I analyze compositional changes in a measure of socio-economic status among teenagers. Socioeconomic status is likely to be negatively correlated with a teenager's surplus from being employed at a given wage. First, teenagers who live in more affluent households are likely to have a lower marginal utility of income.<sup>32</sup> Second, they may also have higher opportunity costs—for example, because access to higher quality schools increases their return to schooling. My measure of socio-economic status is constructed by merging employees' residential ZIP codes with data on median household income from the 1990 Census, and then ranking the ZIP codes of all of a store's employees by the median household income.<sup>33</sup> I then define as "high-income" those whose ZIPs are in the highest-income quartile and "low-income" those in the lowest quartile, and construct dummy variables based on these definitions.

Analysis that distinguishes between teenagers from high, middle, and low-income ZIP codes leads to a set of three results that, together, suggest an increase in the labor-market participation of teenagers. First, among teenagers, wage offers are not correlated with socio-economic status. Second, "high-income" teenagers are more likely to terminate employment to return to school. And third, the positive effect of higher relative wages on teenage employment

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<sup>32</sup> For evidence that parental income negatively affects teenagers' labor supply through an income effect, see Dustmann et al. (forthcoming).

<sup>33</sup> The ranking is constructed using all individuals employed at a store at any time during the sample period. The average number of residential ZIP codes per store is 28; the standard deviation is 18.

is driven mainly by “high-income” teenagers. The first two results suggest that high-income teenagers do indeed have lower benefits and/or higher opportunity costs of being employed; while the third result suggests that these teenagers in particular are drawn into the labor market by the minimum wage.

The first two results are shown in Table 9. First, column (1) shows the coefficients from regressions of starting wages on dummy variables indicating residence in high and low-income ZIP codes. These regressions control for store fixed effects and individual characteristics that include the employee’s month of hire, age, race, gender, fulltime vs. part-time employment status, and an indicator for previous employment with the company. Separate models are estimated for teenagers, young adults (ages 20 to 22), and adults over 22 years old. The results for the teenage sample show that the starting wages are very similar across all groups of teenagers. This suggests either that productivity is not correlated with the measure of socio-economic status or that the employer does not pay wages proportional to output (e.g. because of imperfect information). Interestingly, the lack of wage differentials is not seen in the adult samples; adults residing in high-income ZIP codes earn significantly more than those from lower-income locations.<sup>34</sup>

Column (2) of Table 9 shows the hazard ratios from a Cox proportional hazard model of the rate at which employees terminate employment to return to school. The hazard function is stratified by store and includes as regressors the same employee variables that are in the wage model (columns 1).<sup>35</sup> Among teenagers and young adults, employees from high-income ZIP

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<sup>34</sup> In the case of adults, the positive relationship between the wage and the high-income ZIP code indicator could be due to the influence of wages on socio-economic status. Such reverse causality is much less likely for teenagers, whose residence depends on their parents’ income and not their own.

<sup>35</sup> Stratifying by store controls for all fixed characteristics of the store and its location while allowing each store to have own, flexible baseline hazard. The hazards associated with various ways of terminating employment are treated as independent conditional on the covariates in the model. Thus, in modeling the hazard rate of dismissal,

codes are roughly 35 percent more likely than those in middle-income ZIP codes, and 50 percent more likely than those in low-income ZIP codes, to terminate employment to return to school.<sup>36</sup> Hence, the utility of employment relative to non-market time appears to be lower for teenagers from more affluent ZIP codes.

Evidence that the legislation-induced increase in teenage employment came disproportionately from “high-income” teenagers is shown in Table 10 and in Figure 3. Rows (1)-(4) of Table 10 show the coefficients on the store relative wage gap from regressions in which the dependent variables are the two-year changes in (1) the fraction of employees who are “high-income” teenagers; (2) the fraction of employees who are “middle” income teenagers; (3) the fraction of employees who are “low-income” teenagers; and (4) the fraction of *teenage* employees who live in high-income ZIP codes. The coefficients in rows (1)-(3) are all positive, suggesting that where stores were forced to increase the relative wages of teenagers, they subsequently increased their employment of *all* teenagers. But the coefficients in row (4), where the dependent variable is the fraction of teenagers from high-income ZIP codes, are also positive and statistically significant, and therefore indicate that the increase in teenage employment came disproportionately from high-income teenagers.

Figure 3 plots, by month, the coefficient on the store relative wage gap from a regression in which the dependent variable is the fraction of teenage employees who live in high-income ZIP codes, and which includes control variables as in Table 10, column (2). The time-series pattern here looks very similar to the pattern seen in Figure 2f, where the dependent variable is the teenage share of employment. There is no trend during the pre-legislation period; there is a

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exits for other reasons are treated as censored. For comparison, the table also shows estimates from a model of the hazard for quits. The hazard of being laid off is not correlated with employee ZIP code (results not shown).

<sup>36</sup> The table reports hazard ratios (exponentiated coefficients). E.g., a hazard ratio of 0.80 for a dummy variable implies that the daily rate of dismissal is 20 percent lower for the indicated group than for the omitted group.



jump during the post-legislation holiday season; and it then begins a steady increase in June 1997. This further supports the interpretation that the legislation-induced increase in teenage employment was driven mainly by the increased employment of more affluent teenagers.

### *6.3. Compositional Changes in Employment Flows: New Hires and Exit Rates*

Mechanically, an increase in the employment of teenagers could be driven either by an increase in the number of teenagers who are hired or by a reduction in the exit rate of teenagers who are already employed. But if teenage employment went up due to increased labor market participation, then the change must have been driven mainly by the hiring of more teenagers.

Table 11 shows some evidence that this is true. First, rows (1) and (2) show the estimated effects of the relative wage gap on changes in the shares of new hires who are teenagers (row 1) and 20-22 year-olds (row 2). Though the estimates are less precise than those from the regressions of changes in the employment stock, the pattern of change is similar—with positive effects for teenagers and negative effects for 20-22 year-olds. Next, the estimates in column (3) show the estimated effect of the relative wage gap on the share of new teenager hires who lived in high-income ZIP codes. Again, while not statistically significant, the results are consistent with the estimates based on employment shares and support the hypothesis that new teenage hires came disproportionately from high-income ZIP codes.

Estimates from regressions of changes in exit rates are shown in rows (4)-(6). The point estimates for both teenagers and young adults are positive and somewhat larger for teenagers, but all are statistically insignificant. Hence there is no evidence either that the legislation reduced teenagers' turnover rates or that it increased the turnover rates of young adults. It therefore is unlikely that the increase in teenage employment and decline in young adult employment were driven by differential effects of the legislation on overall turnover rates.

However, when attention is restricted to exits for the purpose of returning to school in “back-to-school months,” the legislation’s estimated effect is negative for teenagers and positive for young adults.<sup>37</sup> For teenagers, the negative effect suggests the relative wage increases not only induced labor market entry by students, but also induced more of them to *stay* in the labor force during the months that school was in session. The positive effect for young adults is more difficult to interpret because the data contains no information on college enrollment. On the one hand, the minimum wage may have induced more college students to end employment with the firm during the school year—perhaps because minimum-wage student jobs on college campuses were now relatively attractive. On the other hand, the increased rate at which young adults left employment for school could reflect a compositional change among young adults—perhaps because the firm substituted toward college students.

#### *6.4 Did the Minimum Wage Raise the Average Productivity of Teenage Employees?*

While both the search and adverse selection models feature increases in labor market participation, a key assumption that distinguishes the latter model is that new labor market entrants are relatively productive. In the current setting, the relevant assumption is that teenagers from high-income ZIP codes are more productive than those from low-income ZIP codes. Testing for such productivity differences can therefore shed more light on the plausibility of the adverse selection model.<sup>38</sup>

The data set offers two ways to test for productivity differences among these groups of teenagers. First, at the individual level, productivity can be measured by the probability of being

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<sup>37</sup> Back-to-school months are defined as January, February, August and September. For this analysis only, I examine the change from these back to school months of the 1996-97 school year to the same months one year later. This choice is motivated by the sample’s time frame (it does not contain January 1996) and by Figure 2F, which shows the permanent increase in teenage employment beginning in the summer of 1997.

<sup>38</sup> Unfortunately, the data set does not contain information on vacancies, so it is not possible to test the “search cost” assumption that the minimum wage caused vacancies to be filled more quickly.

fired.<sup>39</sup> Table 9, column (3) shows the results from a model similar to that estimated in column (1) except that the dependent variable is the hazard rate of being dismissed. Teenage employees from high-income ZIP codes are 17 percent less likely than teenagers from middle-income ZIPs to be fired and 28 percent less likely than those from low-income ZIPs. But moreover, results from a pooled regression (not shown) indicate that while the probability of being fired generally declines with employee age, teenagers from high-income ZIP codes are significantly *less* likely to be fired than a typical 20-22 year-old, and are only slightly more likely to be fired than adults on average. Thus by this measure, at least, teens from high-income ZIP codes are roughly as productive as adults.

The relationship between employee ZIP codes and productivity at the store level can be measured using data on monthly sales. Table 12 shows the results from a regression of sales on the employment shares of employees from each of four groups: teenagers from high-income ZIP codes, other teenagers, adults from high-income ZIP codes, and other adults. The estimation equation includes store fixed effects, month and year dummies, and controls for full-time equivalent employment and the fraction of employees that is part-time. The results indicate a significant positive relationship between sales and the fraction of employees who are from high-income ZIP codes, and especially high-income teenagers. While causality in this relationship is difficult to prove, it is plausible that hiring more high-income employees could cause sales to increase. For example, more affluent employees might attract more affluent customers.<sup>40</sup>

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<sup>39</sup> Dismissals are defined as involuntary terminations that result from dishonesty, substandard performance, tardiness, absenteeism, or violation of company policies; and they comprise roughly seven percent of the observed employment terminations in the data.

<sup>40</sup> In another study that analyzes data from a large U.S. service sector firm employing relatively young low-wage workers, Autor and Scarborough (2008) also find evidence that employees from high-income ZIP codes are more productive. Specifically, they find that the median household income of an employee's ZIP code is a significant predictor of the employee's score on a screening test that in turn predicts various measures of productivity.

In sum, the evidence presented here suggests that among teenagers in my data set, those from relatively high-income ZIP codes were more productive than those from lower-income ZIPs. The combination of this finding with the lack of wage differences among teenagers lends plausibility to the adverse selection model. Moreover, in the face of quality differences among teenagers, the adverse selection model better explains the finding that while the increased employment of “high-income” teenagers displaced some young adults, it did *not* displace teenagers from low-income ZIPs—who were apparently less productive, but were paid similar wages. Specifically, when the model is applied to the market for teenagers, it predicts that an increase in the labor market participation of more productive teenagers raises the demand for *all* teenagers by increasing their average productivity. In contrast, models relying solely on search costs do not offer an explanation for the lack of displacement effects among less productive teens.<sup>41</sup>

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<sup>41</sup> Flinn (2006) and Ahn et al. (2008) assume uniform worker quality. Lang and Kahn (1998) describe a model with search costs in which there are two types of employer and two types of workers, and in which the wage in low-wage jobs is set at a level that achieves a separating equilibrium. Like Drazen’s model, this model also predicts that a minimum wage can induce high-quality workers to apply for low-wage jobs earning the same wage as low-quality workers. In this model, however, the individual quality is observed and so the model predicts that low-quality workers are displaced by high-quality types.

## References

- Addison, John, McKinley Blackburn, and Chad Cotti. 2008. "The Effect of Minimum Wages on Wages and Employment: Country-Level Estimates for the United States." *IZA Discussion Paper No. 3300*. Bonn: Institute for the Study of Labor (IZA).
- Ahn, Tom, Peter Arcidiacono, and Walter Wessels. 2008. "The Distributional Impacts of Minimum Wage Increases when Both Labor Supply and Labor Demand are Endogenous." Forthcoming in *Journal of Business and Economic Statistics*.
- Allegretto, Sylvia, Arindrajit Dube, and Michael Reich. 2008. "Do Minimum Wages Really Reduce Teen Employment? Accounting for Heterogeneity and Selectivity in State Panel Data." *Institute for Research on Labor and Employment Working Paper Series*, No. 166-08. University of California, Berkeley.
- Autor, David and David Scarborough. 2008. "Does Job Testing Harm Minority workers? Evidence from Retail Establishments." *Quarterly Journal of Economics*, 123(1), pp. 219-256.
- Black, Rebecca M. 2002. "Evaluating Welfare Reform in the United States." *Journal of Economic Literature*. Vol. XL, pp. 1105-1166.
- Brown, Charles. 1999. "Minimum Wages, Employment, and the Distribution of Income.: In Orley Ashenfelter and David Card, eds. *Handbook of Labor Economics*, Vol. 3. pp. 2101-2163. New York: Elsevier.
- Card, David, Lawrence F. Katz, and Alan B. Krueger. 1994. "Comment on David Neumark and William Wascher, "Employment Effects of Minimum and Subminimum Wages: Panel Data on State Minimum Wage Laws." *Industrial and Labor Relations Review* 47: 487-496.
- Card, David. 1992a. "Using Regional Variation in Wages to Measure the Effects of the Federal Minimum Wage." *Industrial and Labor Relations Review* 46: 22-37.
- Card, David. 1992b. "Do Minimum Wages Reduce Employment? A Case Study of California, 1987-89." *Industrial and Labor Relations Review* 46: 38-54.
- Card, David and Alan B. Krueger. 1994. "Minimum Wages and Employment: A Case Study of the Fast-Food Industry in New Jersey and Pennsylvania." *American Economic Review* Vol. 84, No.4, pp.772-793.
- Card, David and Alan B. Krueger. 2000. "Minimum Wages and Employment: A Case Study of the Fast-Food Industry in New Jersey and Pennsylvania: Reply." *American Economic Review* Vol. 90, No.5, pp.1397-1420.
- Card, David and Alan B. Krueger. 1995. *Myth and Measurement*. Princeton University Press.

- Drazen, Allan. 1986. "Optimal Minimum Wage Legislation." *The Economic Journal*, Vol. 96, No. 383, pp.774-784.
- Dube, Arindrajit, T. William Lester, and Michael Reich. 2008. "Minimum Wage Effects Across State Borders: Estimate Using Contiguous Counties." *Institute for Research on Labor and Employment Working Paper Series*, No. 166-08. University of California, Berkeley.
- Dustmann Christian, John Micklewright · Arthur van Soest. Forthcoming. "In-school labour supply, parental transfers, and wages." *Empirical Economics*.
- Ehrenberg, Ronald G. 1995. "Review Symposium: *Myth and Measurement: The New Economics of the Minimum Wage*, by David Card and Alan B. Krueger.: *Industrial and Labor Relations Review* 48: 827-28.
- Fairris, David and Leon Fernandez Bujanda. 2008. "The Dissipation of Minimum Wage Gains for Workers through Labor-Labor Substitution." *Southern Economic Journal*, vol. 75, No. 2 (October), pp. 473-496
- Freeman, Richard B. 1996. "The Minimum Wage as a Redistributive Tool." *The Economic Journal* 106: 639-649.
- Guasch, J.L. and Weiss, A. 1980. "Wages as sorting mechanisms in competitive markets with asymmetric information: a theory of testing." *Review of Economic Studies*, vol. 47 (July) pp. 653-64.
- Guasch, J.L. and Weiss, A. 1981. "Self-selection in the labor market." *American Economic Review*, vol. 71 (June), pp. 275-84.
- Katz, Lawrence F., and Krueger, Alan B. 1992. "The Effect of the Minimum Wage on the Fast-Food Industry." *Industrial and Labor Relations Review*, vol. 46 (October): 6-21.
- Lang, Kevin and Shulamit Kahn. 1998. "The Effect of Minimum-Wage Laws on the Distribution of Employment: Theory and Evidence." *Journal of Public Economics* 69, pp.67-82.
- Nelson, Richard R. 1999. "State Labor Legislation Enacted in 1998." *Monthly Labor Review*, January.
- Nelson, Richard R. 1998. "State Labor Legislation Enacted in 1997." *Monthly Labor Review*, January.
- Nelson, Richard R. 1997. "State Labor Legislation Enacted in 1996." *Monthly Labor Review*, January.
- Neumark, David and William Wascher. 2008. *Minimum Wages*. MIT Press.

- Neumark, David and William Wascher. 2007. "Minimum Wages, the Earned Income Tax Credit, and Employment: Evidence from the Post-Welfare Reform Era." *NBER Working Paper 12915*. Cambridge: National Bureau of Economic Research.
- Neumark, David and William Wascher. 2006. "Minimum Wages and Employment: A Review of Evidence from the New Minimum Wage Research." *NBER Working Paper 12663*. Cambridge: National Bureau of Economic Research.
- Neumark, David and William Wascher. 1996. "The Effects of Minimum Wages on Teenage Employment and Enrollment: Evidence from Matched CPS Surveys." *Research in Labor Economics*, 15: 25-63.
- Neumark, David and William Wascher. 2000. "Minimum Wages and Employment: A Case Study of the Fast-Food Industry in New Jersey and Pennsylvania: Comment." *American Economic Review*, 90: 1362-96.
- Neumark, David and William Wascher. 1994. "Employment Effects of Minimum and Subminimum Wages: Reply to Card, Katz, and Krueger." *Industrial and Labor Relations Review* 47: 497-512.
- Neumark, David and William Wascher. 1992. "Employment Effects of Minimum Wages and Subminimum Wages: Panel Data on State Minimum Wage Laws." *Industrial and Labor Relations Review*, 90: 1362-96.
- Rebitzer, J. and L. Taylor. 1995. "The Consequences of Minimum Wage Laws: Some New Theoretical Ideas." *Journal of Public Economics*, 56, pp. 245-256.
- Stigler, George J. 1946. "The Economics of Minimum Wage Legislation." *American Economic Review* 36: 358-65.
- Stiglitz, J. 1974. "Alternative theories of the determination of wages and unemployment in LDC's." *Quarterly Journal of Economics* (May), pp. 194-227.
- Weiss, A. 1980. "Job queues in labor markets with flexible wages." *Journal of Political Economy*, vol. 88 (June), pp. 526-39.

**TABLE 1. STORES' WORKFORCE CHARACTERISTICS, FEB. '96-JULY '96**

<b>Variable</b>	<b>Mean</b>	<b>Std. Dev.</b>
Entry-level employees	27.3	14.0
F/T Equivalent employment <sup>a</sup>	14.7	7.9
% Part-time	93.7%	6.5%
% Teenagers	41.5%	12.9%
% ages 16-17	16.5%	9.0%
% ages 18-19	23.9%	9.7%
% Adults	58.5%	12.9%
% ages 20-22	22.1%	10.2%
% ages 23-29	17.1%	9.0%
% ages 30 & up	19.2%	13.1%
% Female	76.9%	13.5%
% White	71.9%	22.3%
% Black	11.4%	12.6%
% Hispanic	9.3%	12.8%
Square feet	6,978	3,845
Population with 2-mi. radius <sup>b</sup>	83,275	88,780
% population that is white <sup>b</sup>	79.4%	16.7%
% population that is black <sup>b</sup>	7.6%	9.5%
% population that is Hispanic <sup>b</sup>	5.4%	8.9%
Local Area Unemployment Rate <sup>c</sup>	5.1	1.6

*Notes:* Based on employment weighted averages from Feb. 1<sup>st</sup>, 1996-July 31<sup>st</sup>, 1996.

<sup>a</sup> Defined as the number of full-time employees plus ½ × the number of part-time employees.

<sup>b</sup> From 1990 Census; based on 2-mile radius from center of each store's ZIP code.

<sup>c</sup> Based on monthly unemployment rates for metropolitan areas from the U.S. Bureau of Labor Statistics Local Area Unemployment Statistics.



**TABLE 2. FEDERAL AND STATE MINIMUM WAGES DURING THE SAMPLE PERIOD (FEB. 1996-JULY 1998)**

	Synchronized state laws (3.2%)			High initial state minimums (8.6%)					Unsynchronized state laws (20.1%)					
	Fed	CT	AK	DC	RI	IA	WA	NJ	HI	CA	DE	MA	OR	VT
Feb '96	<b>\$4.25</b>	<b>\$4.27</b>	<b>\$4.75</b>	<b>\$5.25</b>	<b>\$4.45</b>	<b>\$4.65</b>	<b>\$4.90</b>	<b>\$5.05</b>	<b>\$5.25</b>	<b>\$4.25</b>	<b>\$4.25</b>	<b>\$4.75</b>	<b>\$4.75</b>	<b>\$4.75</b>
Mar '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	\$4.25	\$4.75	\$4.75	\$4.75
Apr '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	<b>\$4.65</b>	\$4.75	\$4.75	\$4.75
May '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	\$4.65	\$4.75	\$4.75	\$4.75
Jun '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	\$4.65	\$4.75	\$4.75	\$4.75
Jul '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	\$4.65	\$4.75	\$4.75	\$4.75
Aug '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	\$4.65	\$4.75	\$4.75	\$4.75
Sep '96	\$4.25	\$4.27	\$4.75	\$5.25	\$4.45	\$4.65	\$4.90	\$5.05	\$5.25	\$4.25	\$4.65	\$4.75	\$4.75	\$4.75
<b>Oct '96</b>	<b>\$4.75</b>	<b>\$4.77</b>	<b>\$5.25</b>	<b>\$5.75</b>	<b>\$4.75</b>	<b>\$4.75</b>	\$4.90	\$5.05	\$5.25	<b>\$4.75</b>	\$4.75	\$4.75	\$4.75	\$4.75
Nov '96	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$4.75	\$4.75	\$4.75	\$4.75	\$4.75
Dec '96	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$4.75	\$4.75	\$4.75	\$4.75	\$4.75
Jan '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$4.75	<b>\$5.00</b>	<b>\$5.25</b>	<b>\$5.50</b>	<b>\$5.00</b>
Feb '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$4.75	\$5.00	\$5.25	\$5.50	\$5.00
Mar '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	<b>\$5.00</b>	\$5.00	\$5.25	\$5.50	\$5.00
Apr '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$5.00	\$5.00	\$5.25	\$5.50	\$5.00
May '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$5.00	\$5.00	\$5.25	\$5.50	\$5.00
Jun '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$5.00	\$5.00	\$5.25	\$5.50	\$5.00
Jul '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$5.00	\$5.00	\$5.25	\$5.50	<b>\$5.15</b>
Aug '97	\$4.75	\$4.77	\$5.25	\$5.75	\$4.75	\$4.75	\$4.90	\$5.05	\$5.25	\$5.00	\$5.00	\$5.25	\$5.50	\$5.15
<b>Sep '97</b>	<b>\$5.15</b>	<b>\$5.18</b>	<b>\$5.65</b>	<b>\$6.15</b>	<b>\$5.15</b>	<b>\$5.15</b>	<b>\$5.15</b>	<b>\$5.15</b>	\$5.25	<b>\$5.15</b>	<b>\$5.15</b>	\$5.25	\$5.50	\$5.15
Oct '97	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.15	\$5.15	\$5.25	\$5.50	<b>\$5.25</b>
Nov '97	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.15	\$5.15	\$5.25	\$5.50	\$5.25
Dec '97	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.15	\$5.15	\$5.25	\$5.50	\$5.25
Jan '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.15	\$5.15	\$5.25	<b>\$6.00</b>	\$5.25
Feb '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.15	\$5.15	\$5.25	\$6.00	\$5.25
Mar '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	<b>\$5.75</b>	\$5.15	\$5.25	\$6.00	\$5.25
Apr '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.75	\$5.15	\$5.25	\$6.00	\$5.25
May '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.75	\$5.15	\$5.25	\$6.00	\$5.25
Jun '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.75	\$5.15	\$5.25	\$6.00	\$5.25
Jul '98	\$5.15	\$5.18	\$5.65	\$6.15	\$5.15	\$5.15	\$5.15	\$5.15	\$5.25	\$5.75	\$5.15	\$5.25	\$6.00	\$5.25

Source: Nelson, 1997; 1998; 1999.

**TABLE 3. STORE AVERAGE PRE-LEGISLATION WAGES AND MEASURES OF THE LEGISLATION'S IMPACT ON WAGES**

<b>Variable</b>	<b>Federal Min.</b>	<b>HI, IA, NJ, RI &amp; WA</b>	<b>CA</b>	<b>DE, MA, OR &amp; VT</b>	<b>Full Sample</b>
store average wage (all entry-level jobs)	\$5.67 (\$0.51)	\$6.00 (\$0.40)	\$5.93 (\$0.48)	\$5.78 (\$0.37)	\$5.75 (\$0.51)
store average P/T wage	\$5.56 (\$0.45)	\$5.90 (\$0.38)	\$5.79 (\$0.39)	\$5.67 (\$0.29)	\$5.64 (\$0.45)
store average teenage wage	\$5.31 (\$0.45)	\$5.67 (\$0.32)	\$5.58 (\$0.45)	\$5.50 (\$0.25)	\$5.40 (\$0.44)
store average adult wage	\$5.97 (\$0.67)	\$6.33 (\$0.55)	\$6.31 (\$0.69)	\$6.09 (\$0.63)	\$6.08 (\$0.68)
store average relative wage of teenagers	\$0.90 (\$0.08)	\$0.90 (\$0.07)	\$0.89 (\$0.09)	\$0.91 (\$0.08)	\$0.90 (\$0.08)
store wage gap (all entry-level jobs)	2.5% (3.1)	0.2% (0.5)	5.2% (3.4)	1.4% (1.5)	2.6% (3.1)
store P/T wage gap	2.6% (3.1)	0.2% (0.5)	5.5% (3.4)	1.4% (1.6)	2.2% (3.2)
store teenage wage gap	3.7% (4.1)	0.3% (0.6)	7.4% (4.9)	1.9% (2.0)	3.8% (4.3)
store adult wage gap	1.6% (2.7)	0.2% (0.5)	3.2% (2.9)	0.8% (1.2)	1.6% (2.5)
store relative wage gap	2.1% (2.7)	0.2% (0.4)	4.1% (3.6)	1.2% (1.3)	2.1% (2.7)

*Note:* Based on employment-weighted averages of wages during the pre-legislation period from Feb.1, 1996-July 31, 1996. Standard deviations in parentheses.

**TABLE 4. REGRESSION ANALYSIS OF CHANGES IN LOG OF AVERAGE ENTRY-LEVEL WAGE,  
(FEB.-JULY 1996) TO (FEB.-JULY 1998)**

	(1)	(2)	(3)	(4)	(5)
constant	0.044** (0.002)	0.078** (0.015)	0.070** (0.015)	0.105* (0.045)	0.028 (0.028)
<b>store wage gap (for all entry-level employees)</b>	<b>0.773**</b> (0.051)	<b>0.777**</b> (0.052)	<b>0.749**</b> (0.073)	<b>0.778**</b> (0.079)	<b>0.772**</b> (0.065)
% ages 18-19 in Feb-July 1996		-0.023 (0.025)	-0.012 (0.024)	-0.012 (0.024)	-0.019 (0.021)
% ages 20-22 in Feb-July 1996		-0.056* (0.022)	-0.053* (0.022)	-0.050* (0.023)	-0.053** (0.020)
% ages 23-29 in Feb-July 1996		-0.075** (0.022)	-0.064** (0.022)	-0.062** (0.022)	-0.060** (0.020)
% ages ≥30 in Feb-July 1996		-0.018 (0.021)	0.002 (0.021)	0.002 (0.021)	0.001 (0.019)
location type = open mall				0.005 (0.013)	-0.000 (0.010)
location type = street				-0.008 (0.006)	-0.008 <sup>‡</sup> (0.005)
location type = strip				0.012 (0.007)	0.010 (0.007)
store square footage (1,000s)				0.001 (0.001)	0.001 (0.001)
population in 2-mile radius of store (100,000s)				0.005 (0.005)	0.006 (0.004)
% white in 2-mile radius of store				-0.041 (0.043)	0.044 <sup>‡</sup> (0.025)
% black in 2-mile radius of store				-0.059 (0.048)	0.024 (0.032)
% Hispanic in 2-mile radius of store				-0.057 (0.061)	0.038 (0.039)
average local unemployment rate, Feb.-July 1996				-0.002 (0.002)	-0.002 (0.001)
<b>Region Fixed Effects</b>	No	No	Yes	Yes	Yes
<b>Sample Includes CA, DE, MA, OR &amp; VT</b>	No	No	No	No	Yes
<b>R-squared</b>	0.19	0.22	0.36	0.37	0.37

*Notes:* Omitted age category is % ages 16-17. Omitted location type category is "mall". Robust standard errors in parentheses. <sup>‡</sup> significant at 10%; \* significant at 5%; \*\* significant at 1%

**TABLE 5. REGRESSION ANALYSIS OF CHANGES IN AVERAGE FULL-TIME EQUIVALENT EMPLOYMENT,  
(FEB.-JULY 1996) TO (FEB.-JULY 1998)**

	(1)	(2)	(3)	(4)	(5)	(6)
constant	3.193** (0.216)	4.274** (1.528)	1.064 (1.456)	0.516 (4.703)	0.558 (4.710)	0.524 (2.639)
<b>store wage gap (for all entry-level employees)</b>	<b>-1.021</b> (4.697)	<b>-1.665</b> (4.927)	<b>-4.032</b> (7.123)	<b>-9.029</b> (7.505)	<b>-8.745</b> (7.541)	<b>-7.202</b> (6.736)
% ages 18-19 in Feb-July 1996		-0.880 (2.396)	1.581 (2.273)	1.346 (2.294)	1.298 (2.298)	1.234 (2.035)
% ages 20-22 in Feb-July 1996		-1.323 (2.203)	3.821 <sup>‡</sup> (2.235)	3.328 (2.200)	3.289 (2.210)	2.530 (1.976)
% ages 23-29 in Feb-July 1996		-1.149 (2.251)	2.228 (2.171)	1.885 (2.168)	1.796 (2.172)	0.836 (1.979)
% ages ≥30 in Feb-July 1996		-1.853 (2.059)	3.028 (1.986)	3.831 <sup>‡</sup> (1.982)	3.814 <sup>‡</sup> (1.987)	4.113* (1.770)
location type = open mall				-0.980 (1.520)	-0.989 (1.516)	-0.896 (1.147)
location type = street				-1.771** (0.518)	-1.786** (0.519)	-1.867** (0.443)
location type = strip				-0.938 (0.633)	-0.916 (0.637)	-0.952 (0.586)
store square footage (1,000s)				0.238** (0.085)	0.235** (0.086)	0.263** (0.078)
population in 2-mile radius of store (100,000s)				-0.422 (0.399)	-0.402 (0.401)	-0.006 (0.320)
% white in 2-mile radius of store				0.257 (4.499)	0.263 (4.496)	0.563 (2.196)
% black in 2-mile radius of store				-4.085 (4.715)	-4.130 (4.719)	-3.898 (2.775)
% Hispanic in 2-mile radius of store				3.885 (5.917)	3.852 (5.930)	3.308 (3.215)
average local unemployment rate, Feb.-July 1996				-0.023 (0.151)	-0.026 (0.151)	-0.084 (0.132)
Δ % Part-time					2.231 (3.507)	2.391 (3.025)
Region Fixed Effects	No	No	Yes	Yes	Yes	Yes
Sample Includes CA, DE, MA, OR & VT	No	No	No	No	No	Yes
R-squared	0.00	0.00	0.17	0.21	0.21	0.21

*Notes:* Omitted age category is % ages 16-17. Omitted location type category is "mall". Robust standard errors in parentheses. <sup>‡</sup> significant at 10%; \* significant at 5%; \*\* significant at 1%

**TABLE 6. REGRESSION ANALYSIS OF CHANGES IN RELATIVE TEENAGE WAGE,  
(FEB.-JULY 1996) TO (FEB.-JULY 1998)**

	(1)	(2)	(3)	(4)	(5)
constant	-0.020** (0.003)	-0.026 (0.025)	-0.042 (0.028)	0.025 (0.075)	-0.054 (0.050)
<b>store relative wage gap</b>	<b>1.021**</b> (0.104)	<b>1.042**</b> (0.103)	<b>1.280**</b> (0.120)	<b>1.321**</b> (0.124)	<b>1.267**</b> (0.110)
% ages 18-19 in Feb-July 1996		0.002 (0.039)	0.006 (0.042)	0.009 (0.042)	0.026 (0.037)
% ages 20-22 in Feb-July 1996		0.031 (0.041)	0.051 (0.043)	0.050 (0.042)	0.041 (0.037)
% ages 23-29 in Feb-July 1996		0.012 (0.035)	0.045 (0.039)	0.046 (0.040)	0.022 (0.036)
% ages ≥30 in Feb-July 1996		-0.020 (0.034)	-0.016 (0.037)	-0.028 (0.039)	-0.038 (0.034)
location type = open mall				0.006 (0.016)	-0.001 (0.015)
location type = street				0.005 (0.009)	0.009 (0.008)
location type = strip				-0.021* (0.010)	-0.015 (0.010)
store square footage (1,000s)				0.000 (0.001)	-0.000 (0.001)
population in 2-mile radius of store (100,000s)				-0.002 (0.010)	0.003 (0.008)
% white in 2-mile radius of store				-0.068 (0.074)	0.011 (0.042)
% black in 2-mile radius of store				-0.059 (0.078)	0.008 (0.051)
% Hispanic in 2-mile radius of store				-0.038 (0.094)	0.063 (0.070)
average local unemployment rate, Feb.-July 1996				-0.001 (0.003)	0.000 (0.002)
<b>Region Fixed Effects</b>	No	No	Yes	Yes	Yes
<b>Sample Includes CA, DE, MA, OR &amp; VT</b>	No	No	No	No	Yes
<b>R-squared</b>	0.12	0.13	0.21	0.23	0.22

*Notes:* Omitted age category is % ages 16-17. Omitted location type category is "mall". Robust standard errors in parentheses. ‡ significant at 10%; \* significant at 5%; \*\* significant at 1%

**TABLE 7. REGRESSION ANALYSIS OF CHANGES IN TEENAGE SHARE OF EMPLOYMENT, (FEB.-JULY 1996) TO (FEB.-JULY 1998)**

	<i>Dependent Variable: Change in Fraction of Employees Who are Ages...</i>											
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	<b>16-19</b>	<b>16-19</b>	<b>16-19</b>	<b>16-19</b>	<b>16-19</b>	<b>16-19</b>	<b>16-19</b>	<b>16-17</b>	<b>18-19</b>	<b>20-22</b>	<b>23-29</b>	<b>30 &amp; up</b>
constant	-0.036** (0.006)	-0.161** (0.039)	-0.214** (0.041)	-0.287* (0.118)	-0.281* (0.118)	-0.239* (0.115)	-0.195** (0.065)	-0.566** (0.056)	0.371** (0.059)	0.074 (0.050)	0.146** (0.054)	-0.024 (0.049)
<b>store relative wage gap</b>	<b>0.621**</b> (0.204)	<b>0.735**</b> (0.191)	<b>0.925**</b> (0.212)	<b>0.790**</b> (0.214)	<b>0.798**</b> (0.215)	<b>0.729**</b> (0.199)	<b>0.747**</b> (0.173)	<b>0.497**</b> (0.133)	<b>0.250<sup>‡</sup></b> (0.133)	<b>-0.702**</b> (0.136)	<b>-0.189</b> (0.126)	<b>0.144</b> (0.106)
% ages 18-19 in Feb-July 1996		-0.401** (0.065)	-0.391** (0.063)	-0.404** (0.063)	-0.409** (0.063)	-0.466** (0.061)	-0.417** (0.055)	0.653** (0.059)	-1.070** (0.048)	0.422** (0.044)	0.001 (0.040)	-0.005 (0.037)
% ages 20-22 in Feb-July 1996		0.372** (0.059)	0.460** (0.063)	0.458** (0.062)	0.453** (0.061)	0.406** (0.057)	0.412** (0.051)	0.515** (0.047)	-0.103* (0.044)	-0.636** (0.041)	0.198** (0.035)	0.025 (0.035)
% ages 23-29 in Feb-July 1996		0.309** (0.061)	0.409** (0.064)	0.422** (0.064)	0.409** (0.063)	0.303** (0.059)	0.290** (0.054)	0.477** (0.052)	-0.187** (0.046)	0.127** (0.041)	-0.585** (0.037)	0.168** (0.034)
% ages ≥30 in Feb-July 1996		0.403** (0.053)	0.457** (0.055)	0.472** (0.058)	0.470** (0.057)	0.330** (0.057)	0.354** (0.051)	0.588** (0.051)	-0.234** (0.045)	-0.060 (0.038)	0.050 (0.034)	-0.344** (0.036)
location type = open mall				-0.007 (0.025)	-0.008 (0.025)	-0.019 (0.021)	-0.020 (0.017)	0.007 (0.012)	-0.027* (0.013)	-0.007 (0.013)	0.032* (0.014)	-0.005 (0.014)
location type = street				-0.023 <sup>‡</sup> (0.013)	-0.025* (0.012)	-0.024* (0.012)	-0.017 (0.011)	-0.011 (0.009)	-0.006 (0.009)	0.004 (0.009)	0.011 (0.008)	0.002 (0.008)
location type = strip				-0.022 (0.017)	-0.020 (0.017)	-0.019 (0.017)	-0.016 (0.016)	0.007 (0.013)	-0.023* (0.010)	-0.016 (0.011)	0.001 (0.011)	0.031* (0.013)
store square footage (1,000s)				0.000 (0.002)	-0.000 (0.002)	-0.000 (0.002)	0.000 (0.001)	0.001 (0.001)	-0.001 (0.001)	0.001 (0.001)	-0.001 (0.001)	-0.001 (0.001)
population within 2-mi. of store (100,000s)				-0.017 (0.011)	-0.014 (0.011)	-0.015 (0.011)	-0.009 (0.009)	-0.003 (0.007)	-0.005 (0.007)	-0.008 (0.007)	0.015* (0.006)	0.002 (0.006)
% white in 2-mile radius of store				0.143 (0.114)	0.144 (0.114)	0.150 (0.114)	0.100 <sup>‡</sup> (0.057)	0.113** (0.041)	-0.013 (0.051)	-0.017 (0.040)	-0.118* (0.050)	0.036 (0.041)
% black in 2-mile radius of store				0.018 (0.125)	0.012 (0.125)	0.012 (0.121)	-0.043 (0.071)	-0.000 (0.051)	-0.043 (0.058)	0.090 <sup>‡</sup> (0.053)	-0.059 (0.062)	0.012 (0.049)
% Hispanic in 2-mile radius of store				0.341* (0.155)	0.337* (0.153)	0.392** (0.148)	0.334** (0.088)	0.314** (0.068)	0.021 (0.076)	-0.180** (0.068)	-0.211** (0.075)	0.057 (0.067)
avg. local unemployment rate (Feb.-Jul.'96)				-0.007 (0.005)	-0.007 (0.005)	-0.004 (0.005)	-0.008* (0.004)	-0.005+ (0.003)	-0.003 (0.002)	0.002 (0.003)	0.004 (0.003)	0.001 (0.002)
change in percent part-time					0.307** (0.085)	0.252** (0.082)	0.256** (0.073)	0.217** (0.065)	0.040 (0.057)	-0.069 (0.054)	-0.177** (0.051)	-0.010 (0.053)
change in percent single female						0.378** (0.041)	0.320** (0.037)	0.206** (0.032)	0.114** (0.026)	0.007 (0.028)	-0.086** (0.025)	-0.242** (0.026)
Region Fixed Effects	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Sample Incl. CA, DE, MA, OR, VT	No	No	No	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.01	0.31	0.41	0.44	0.45	0.51	0.48	0.38	0.54	0.50	0.43	0.38

Notes: Omitted age category is % ages 16-17. Omitted location type category is "mall". Robust standard errors in parentheses. <sup>‡</sup> significant at 10%; \* at 5%; \*\* at 1%.

**TABLE 8. REGRESSION ANALYSIS OF TWO-YEAR CHANGES IN FULL-TIME EQUIVALENT EMPLOYMENT,  
(FEB.-JULY '96) TO (FEB.-JULY '98)**

	<u>Δ Teenage</u> FTE employment			<u>Δ Adult</u> FTE employment			<u>Δ FTE</u> employment		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
<b>Store average wage gap</b>	4.474 (4.138)		-2.439 (4.528)	-13.203* (5.341)		-12.575* (5.848)	-8.737 (7.639)		-15.019 <sup>‡</sup> (8.315)
<b>Store relative wage gap</b>		13.892** (4.016)	15.000** (4.444)		-7.075 (5.178)	-1.361 (5.672)		6.807 (6.582)	13.631 <sup>‡</sup> (7.164)
R-squared	0.34	0.35	0.35	0.20	0.19	0.20	0.20	0.20	0.21

*Note:* Estimation equations include control variables as in Table 5, columns (5) & (6). <sup>‡</sup> significant at 10%; \* significant at 5%; \*\* significant at 1%.

**TABLE 9. COMPARISONS OF EMPLOYEES LIVING IN HIGH-INCOME VS. LOW-INCOME ZIP CODES**

<b>Sample: Teenagers</b>				
	(1) Log of	Hazard Rate for Employment Termination Due to:		
	Starting Wage	(2) School	(3) Dismissal	(4) Quit
employee lives in high-income ZIP code	<b>-0.000</b> (0.001)	<b>1.247**</b> (0.026)	<b>0.828**</b> (0.031)	<b>0.970*</b> (0.014)
employee lives in low-income ZIP code	<b>0.000</b> (0.001)	<b>0.847**</b> (0.026)	<b>1.146**</b> (0.046)	<b>1.031<sup>‡</sup></b> (0.018)
age 17	0.011** (0.001)	1.325** (0.053)	1.023 (0.048)	1.172** (0.022)
age 18	0.031** (0.001)	2.756** (0.106)	1.117* (0.053)	1.413** (0.028)
age 19	0.051** (0.001)	2.986** (0.117)	1.174** (0.057)	1.560** (0.031)
full time	0.060** (0.007)	0.527** (0.077)	1.437* (0.224)	1.413** (0.082)
prior company experience	0.021** (0.001)	1.843** (0.041)	0.402** (0.024)	0.528** (0.011)
Observations	> 50,000	> 50,000	> 50,000	> 50,000
<b>Sample: 20-22 Year Olds</b>				
	(1) Log of	Hazard Rate for Employment Termination Due to:		
	Starting Wage	(2) School	(3) Dismissal	(4) Quit
employee lives in high-income ZIP code	<b>0.002</b> (0.001)	<b>1.279**</b> (0.037)	<b>0.841**</b> (0.047)	<b>0.991</b> (0.018)
employee lives in low-income ZIP code	<b>-0.005**</b> (0.001)	<b>0.868**</b> (0.029)	<b>1.162**</b> (0.056)	<b>1.037*</b> (0.019)
age 21	0.015** (0.001)	0.894** (0.024)	0.887** (0.041)	1.079** (0.018)
age 22	0.024** (0.001)	0.594** (0.021)	0.788** (0.042)	1.149** (0.021)
full time	0.170** (0.006)	0.241** (0.037)	1.124 (0.146)	0.815** (0.042)
prior company experience	0.038** (0.001)	1.694** (0.044)	0.401** (0.025)	0.499** (0.010)
Observations	> 50,000	> 50,000	> 50,000	> 50,000
<b>Sample: Over 22 Years Old</b>				
	(1) Log of	Hazard Rate for Employment Termination Due to:		
	Starting Wage	(2) School	(3) Dismissal	(4) Quit
employee lives in high-income ZIP code	<b>0.010**</b> (0.002)	<b>1.094</b> (0.068)	<b>0.730**</b> (0.044)	<b>0.959*</b> (0.016)
employee lives in low-income ZIP code	<b>-0.011**</b> (0.001)	<b>0.941</b> (0.060)	<b>1.343**</b> (0.064)	<b>1.022</b> (0.016)
age	0.007** (0.000)	0.750** (0.010)	0.940** (0.009)	0.956** (0.003)
age <sup>2</sup>	-0.000** (0.000)	1.003** (0.000)	1.001** (0.000)	1.000** (0.000)
full time	0.257** (0.005)	0.292** (0.047)	0.868 (0.090)	0.674** (0.023)
prior company experience	0.055** (0.002)	1.389** (0.080)	0.477** (0.031)	0.609** (0.011)
Observations	> 50,000	> 50,000	> 50,000	> 50,000

*Notes:* Column (1) shows coefficients from linear regressions predicting the log of starting wage. Columns (2)-(4) show estimated hazard ratios from Cox proportional hazard models predicting the likelihood of (1) terminating employment to return to school; (2) being dismissed; (3) quitting (for job-related reason). All regression models include store fixed effects (hazard models are stratified by store) and all control for dummy variables indicating month of hire. Omitted ZIP code category is ZIPs in middle quartiles of median household income. Omitted age categories: age 16 for teenage sample; age 20 for 20-22 year-olds. Robust standard errors in parentheses; <sup>‡</sup> significant at 10%; \* significant at 5%; \*\* significant at 1%.



**TABLE 10. ESTIMATED EFFECTS OF STORE RELATIVE WAGE GAP ON EMPLOYMENT SHARE OF TEENAGERS, BY ZIP CODE TYPE**

<u>Dependent Variable:</u>	<u>Coefficient on store relative wage gap</u> <u>(Std. Error)</u>		
	<u>(1)</u>	<u>(2)</u>	<u>(3)</u>
(1) $\Delta$ % Employees who are teenagers from high-income ZIP codes	0.390* (0.157)	0.467** (0.159)	0.347* (0.130)
(2) $\Delta$ % Employees who are teenagers from middle-income ZIP codes	0.064 (0.184)	0.040 (0.197)	0.119 (0.166)
(3) $\Delta$ % Employees who are teenagers from low-income ZIP codes	0.151 (0.095)	0.221* (0.142)	0.281* (0.125)
(4) $\Delta$ % Teenagers who are from high-income ZIP codes	0.755** (0.270)	0.864* (0.338)	0.567* (0.271)
Estimation equation includes control variables as in Table 7 , col. (6)	NO	YES	YES
Sample Includes CA, DE, MA, OR & VT	NO	NO	YES

Notes: ‡ significant at 10%; \* significant at 5%; \*\* significant at 1%.

**TABLE 11. ESTIMATED EFFECTS OF STORE RELATIVE WAGE GAP ON COMPOSITION OF NEW HIRES AND EXIT RATES**

<u>Dependent Variable:</u>	<u>Coefficient on store relative wage gap</u> <u>(Std. Error)</u>		
	<u>(1)</u>	<u>(2)</u>	<u>(3)</u>
(1) $\Delta$ % New hires who are teenagers	0.499 (0.352)	0.985* (0.465)	0.512 (0.391)
(2) $\Delta$ % New hires who are ages 20-22	-0.292 (0.271)	-0.598 (0.379)	-0.415 (0.321)
(3) $\Delta$ % Teenage hires who are from high-income ZIP codes	0.548 (0.467)	0.482 (0.605)	0.347 (0.687)
(4) $\Delta$ Avg. daily exit rate for teenagers	0.017 (0.028)	0.008 (0.016)	0.031 (0.021)
(5) $\Delta$ Avg. daily exit rate for ages 20-22	0.003 (0.008)	0.005 (0.010)	0.007 (0.008)
(6) $\Delta$ Avg. daily rate at which teens exit employment for school (back-to-school months in 1996-97 vs. 1997-98 school years)	-0.016‡ (0.009)	-0.017‡ (0.010)	-0.003 (0.009)
Estimation equation includes control variables as in Table 7 , col. (6)	NO	YES	YES
Sample Includes CA, DE, MA, OR & VT	NO	NO	YES

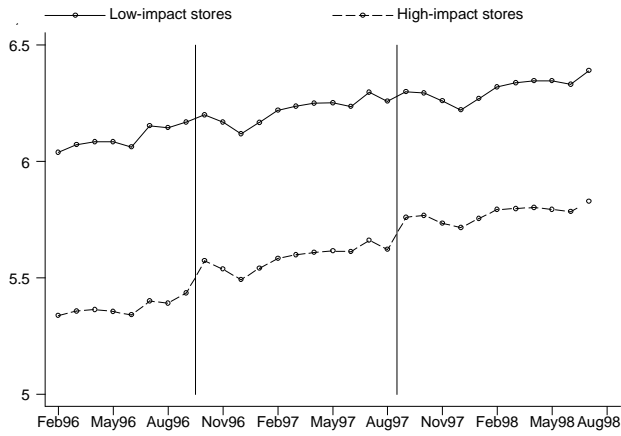
Notes: ‡ significant at 10%; \* significant at 5%; \*\* significant at 1%.

**TABLE 12. ESTIMATED EFFECTS OF EMPLOYEE AGE AND ZIP CODE COMPOSITION ON SALES**

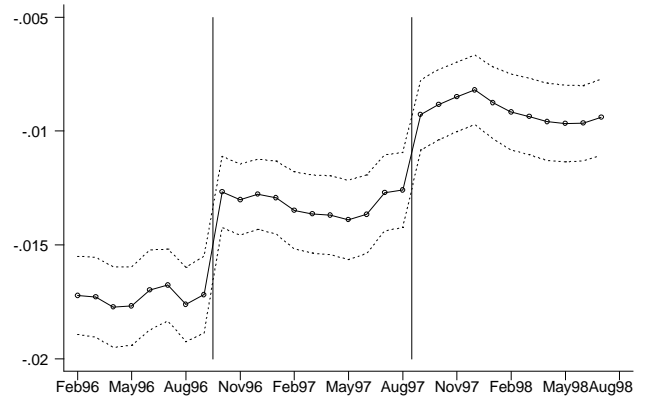
	<u>Log Real Monthly Sales</u>
intercept	2.636** (0.055)
% employees who are teenagers from high-income ZIP codes	0.109** (0.036)
% employees who are adults <i>not</i> from high-income ZIP codes	-0.011 (0.026)
% employees who are adults from high-income ZIP codes	0.042 (0.035)
<hr/>	
R-squared	0.90
<hr/>	
F test of H <sub>0</sub> : %not-rich adults=%rich adults ( <i>Prob. &gt;F</i> )	2.21 (0.137)

Notes: Table shows coefficients and robust standard errors from a linear regression with store fixed effects and controls for full-time equivalent employment, the fraction of employees who are part-time, and month indicators. The omitted category is % employees who are teenagers *not* from high-income ZIP codes. ‡ significant at 10%; \* significant at 5%; \*\* significant at 1%

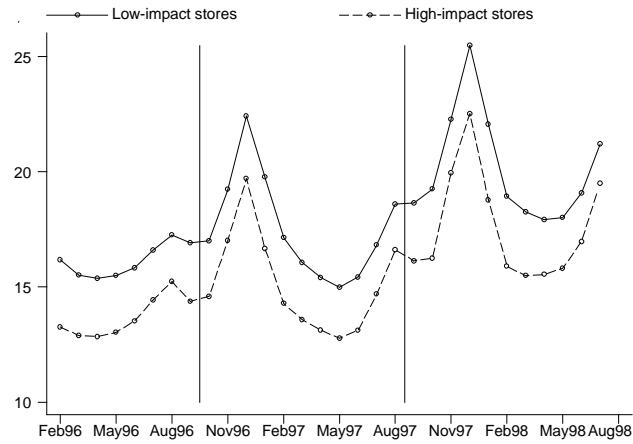
**FIGURE 1A.  
AVERAGE ENTRY-LEVEL WAGE**



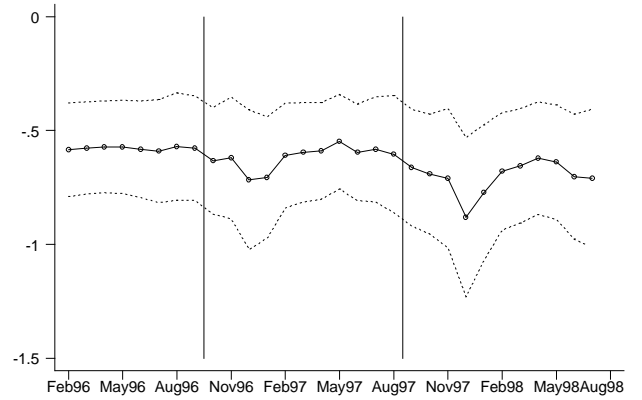
**FIGURE 1B. EFFECT OF A .01 INCREASE IN STORE WAGE GAP ON LN(AVG. ENTRY-LEVEL WAGE)**



**FIGURE 1C.  
FULL-TIME EQUIVALENT ENTRY-LEVEL EMPLOYMENT**

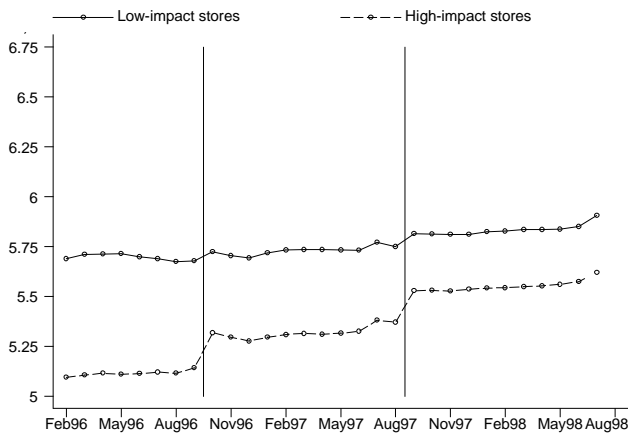


**FIGURE 1D. EFFECT OF A .01 INCREASE IN STORE WAGE GAP ON FULL-TIME EQUIVALENT EMPLOYMENT**

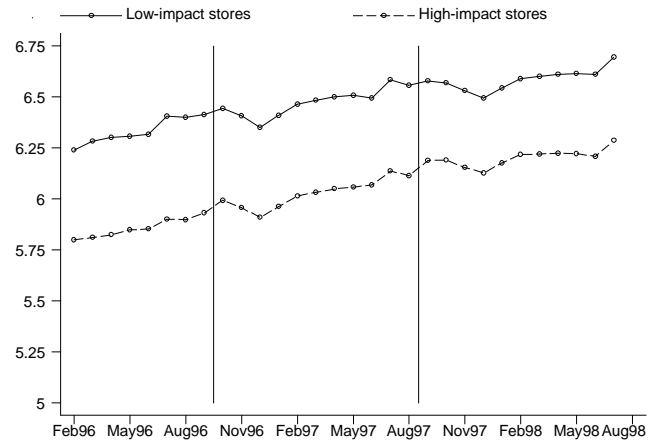


Notes: Vertical lines indicate dates of October 1996 and September 1997 federal minimum wage increases. Dotted lines in Fig. 1b, 1d show 95 percent confidence interval.

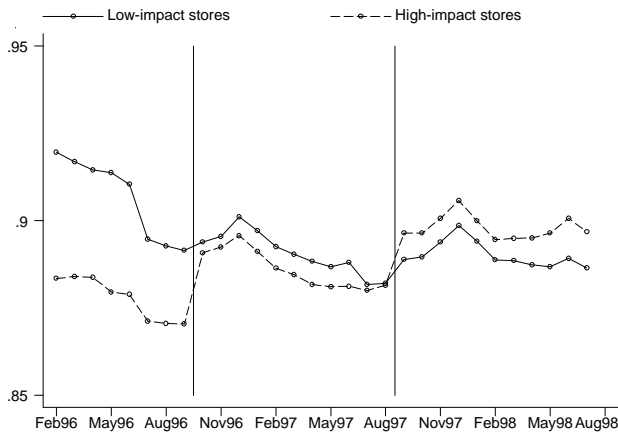
**FIGURE 2A. AVERAGE TEENAGE WAGE**



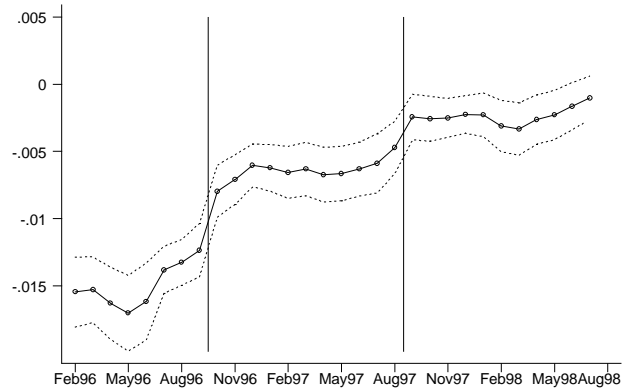
**FIGURE 2B. AVERAGE ADULT WAGE**



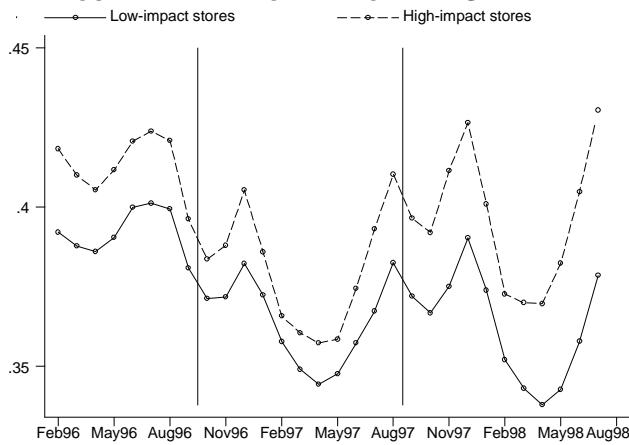
**FIGURE 2C. TEENAGE WAGE/ ADULT WAGE**



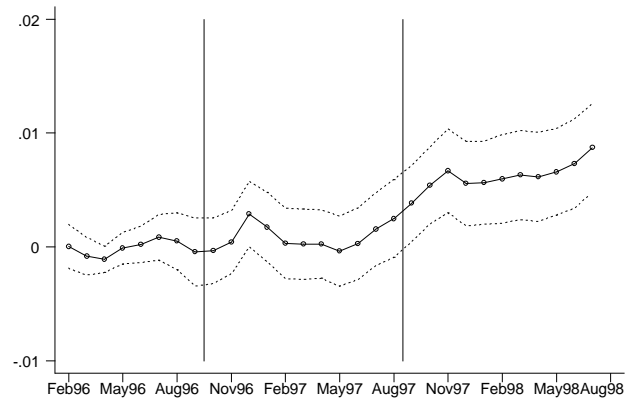
**FIGURE 2D. EFFECT OF A .01 INCREASE IN RELATIVE WAGE GAP ON TEENAGE RELATIVE WAGE**



**FIGURE 2E. TEENAGE EMPLOYMENT SHARE**

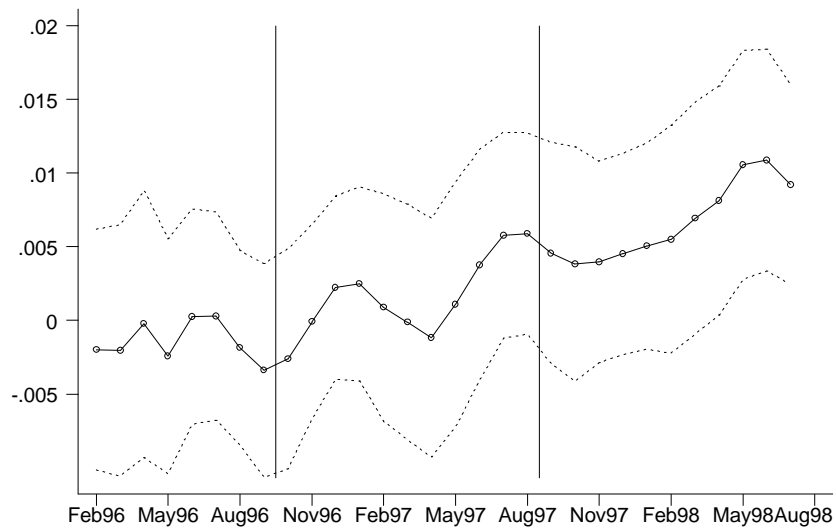


**FIGURE 2F. EFFECT OF A .01 INCREASE IN RELATIVE WAGE GAP ON TEENAGE EMPLOYMENT SHARE**



*Notes:* Vertical lines indicate dates of October 1996 and September 1997 federal minimum wage increases. Dotted lines in Fig. 2d, 2f show 95 percent confidence interval.

**FIGURE 3. ESTIMATED EFFECT OF .01 INCREASE IN RELATIVE WAGE GAP ON FRACTION OF TEENAGERS WHO ARE FROM HIGH-INCOME ZIP CODES**




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*Notes for figures:* Vertical lines indicate dates of October 1996 and September 1997 federal minimum wage increases. Dotted lines show 95 percent confidence interval.

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