The Minimum Wage:
Maximum Controversy over a Minimal Effect?

by

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Abstract

This thesis examines several aspects of the minimum wage in the United States. This work is motivated by recent empirical work that fails to find negative employment effects of minimum wage increases. The first two chapters of this thesis develop models in which an increase in the minimum wage might not reduce employment and investigate the empirical validity of these models. In Chapter 1, I propose a dynamic monopsony model in which binding increases in the minimum wage increase hires and reduce quits. A minimum wage increase can raise both firm-level and total employment in the model. The validity of the model is then tested using data on employment, prices, and the number of establishments in the restaurant and retail clothing industries. I find evidence of positive employment effects in the restaurant industry, but the other results are not consistent with the dynamic monopsony model.

Chapter 2 examines another reason why minimum wage increases might not reduce employment: political endogeneity. If politicians time increases in minimum wages to occur when any adverse effects are minimized, then traditional analyses underestimate the true disemployment effect of the minimum wage by failing to control for the endogeneity of the minimum wage. I examine the determinants of state and federal minimum wages and find that economic conditions are correlated with wage floors. Two methods are used to control for endogeneity: controlling for lagged economic conditions, and using potentially exogenous political variables that affect the minimum wage as instrumental variables to identify the employment effect of the minimum wage. The estimated results are not significantly different from estimates without controlling for endogeneity bias.

Chapter 3 examines the effects of two increases in Washington state's minimum wage in low-wage retail trade and service industries. The effective minimum wage in Washington increased by 90 cents over 1988-1990. Changes in average wages and employment in Washington are compared to changes in the rest of the United States. The results indicate that the minimum wage increases raised average wages but may not have had substantial disemployment effects. A political economy analysis of voting behavior on the ballot initiative that raised the minimum wage is consistent with the possibility that the minimum wage increase was not exogenous, which may explain the failure to find conclusive evidence of adverse employment effects.

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Table of Contents

Introduction 5

Chapter 1: Empirical Evidence on Dynamic Monopsony and Minimum Wages 11

Chapter 2: Controlling for Endogeneity in Estimates of Minimum Wage Effects 55

Chapter 3: Effects of the Minimum Wage in Low-Wage Industries in Washington 79

References 106

Introduction

Whether the minimum wage should be raised is periodically the subject of intense political debate. Politicians have argued about whether raising the wage floor reduces employment opportunities--and by how much--and ameliorates poverty since Franklin D. Roosevelt proposed establishing a federal minimum wage during the 1930s. Until a few years ago, however, the effect of the minimum wage on employment was not a controversial subject among economists. The vast majority of economists accepted the prediction of the neoclassical model that increases in the minimum wage reduced the demand for labor (Kearl et. al. 1979 and Frey et. al. 1984). Disemployment effects were believed to be particularly pronounced among teens; the conventional estimate was that a 10 percent increase in the minimum wage reduced teen employment by 1 to 3 percent (Brown, Gilroy, and Kohen 1982). A series of recent papers that find that increases in federal and state minimum wages appear to not reduce employment have challenged the economic orthodoxy and stimulated research on the effects of the minimum wage. Studies of the effects of the minimum wage are particularly relevant at the moment as President Clinton calls for making the minimum wage a "living wage" and Congress debates a 90-cent increase in the federal minimum wage.

A study by David Card and Alan Krueger (1994) of the effects of a minimum wage increase in New Jersey in 1992 is receiving substantial attention in the current debate over whether to raise the minimum wage. Based on a phone survey of fast-food restaurants, Card and Krueger find that the 80-cent minimum wage hike did not appear to reduce employment in New Jersey relative to Pennsylvania. Both the data and the differences-in-differences in

methodology used by Card and Krueger have been criticized. While some of the criticisms appear valid, other studies find the same result using a variety of data and methodologies. Wellington (1991) uses a traditional time-series methodology to estimate the relationship between changes in the federal minimum wage and the employment rates of teens and young adults over 1954-1986. Card (1992a) compares the effect of the 45-cent federal minimum wage increase in 1990 on teen employment in low-wage and high-wage states. Similarly, Katz and Krueger (1992) examine the effects of the 1990 federal minimum wage hike on fast-food restaurants in Texas. Card, Katz, and Krueger (1994) examine the correlation the effective minimum wage (the higher of state and federal minimum wages) and teen employment rates in the 50 states over 1973-1989. Card (1992b) studies the effects of a 90-cent increase in California's minimum wage in 1988 relative to five other areas. None of these studies finds evidence of a significant negative relationship between minimum wages and employment. In addition, Machin and Manning (1992) find no evidence that minimum wages and employment are inversely related in Britain.

Several recent studies that find a negative correlation between the minimum wage and employment are cited by opponents of a minimum wage increase to counter claims that raising the minimum wage would not reduce employment. Some of these papers use similar methodologies and examine the same minimum wage increases as those that find that minimum wages increases do not reduce employment. Kim and Taylor (1995) report that California's minimum wage increase in 1988 reduced employment in retail trade relative to the rest of the U.S. Their study relies on establishment data, whereas Card (1992b) uses individual data on teens and a smaller comparison group. Neumark and Wascher (1992 and

1994) report that increases in effective minimum wages are negatively correlated with teen employment rates over 1973-1989; they use the same data as Card, Katz, and Krueger (1994), but Neumark and Wascher control for changes in enrollment. Neumark and Wascher (1995a) also find that New Jersey's minimum wage increase in 1992 reduced employment when payroll data is used for a subset of the firms surveyed by phone by Card and Krueger. Other studies finding negative employment effects include Deere, Murphy, and Welch (1995), Currie and Fallick (1995), Williams (1993), and Abowd et. al. (1995). In addition, Baker, Benjamin, and Stanger (1995) report a negative correlation between minimum wages and employment in Canada.

Studies that fail to find negative employment effects are not necessarily incompatible with the neoclassical model and may not be generalizable. First, the studies that failed to find negative employment effects may have examined minimum wage increases that were not binding. However, most of these studies report that average wages rose among the affected population when the minimum wage was increased. Second, the econometric techniques used on data with measurement error may not be able to capture the negative employment effect--which is likely to be small--predicted by the neoclassical model. As John Kennan (1996) explains, the large cyclical swings and high degree of serial correlation in teen employment rates over time makes looking for employment rate changes of about one percent like looking for a needle in a haystack. This criticism clearly begs the question of why some studies are able to find negative effects while others are not when using the same methodology and similar data. Third, firms may be slow to adjust to minimum wage changes, and adverse effects may show up with a lag and not be captured by the

contemporaneous data used in most studies. Neumark and Wascher (1992 and 1995b) report that lagged employment effects are more negative than contemporaneous effects. Card and Krueger (1995) re-estimate some of their previous results while including lagged terms and still do not find adverse employment effects.

Other explanations for why increases in the minimum wage might not reduce employment fall outside the neoclassical model. Models in which an increase in the wage floor may raise employment include static monopsony and dynamic monopsony.[1] In the traditional static monopsony model, a firm faces an upward-sloping labor supply curve. A binding wage floor makes the marginal cost of labor constant over some range and can increase the amount of labor hired by the firm.[2] The static monopsony model is typically applied to concentrated markets, such as "one company" coal towns, and is of limited relevance for unconcentrated, minimum-wage industries such as retail trade. In the dynamic monopsony model, a minimum wage increase can raise employment because workers engage in imperfect search. In the model formulated by Burdett and Mortensen (1989), workers have imperfect information about the wages offered by firms, making the model particularly plausible for industries with many small firms. An increase in the minimum wage can increase employment by reducing quits and raising hires, particularly at smaller, low-wage firms.

The political process in which minimum wages are determined may also cause

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[1]The efficiency wages model has also been offered as an explanation for why minimum wage increases might not reduce employment. This model is discussed in Chapter 1 of this thesis and in Rebitzer and Taylor (1995) but may be of limited relevance for minimum wage industries.

[2]Increases in the minimum wage raise employment until the wage floor exceeds the marginal revenue product of labor.

minimum wages to appear to have no employment effect. If politicians are concerned about adverse effects, they may not pass large, or even binding, minimum wage increases. Minimum wage increases may be a symbolic gesture aimed at political effect, not a mechanism for raising wages. In this case, it is not surprising that researchers fail to find evidence of substantial adverse effects.

The first two chapters of this thesis develop and attempt to empirically validate models in which increases in the minimum wage do not have adverse employment effects. In Chapter 1, I develop a dynamic monopsony model in which minimum wage increases may increase employment. This paper makes two main contributions to the minimum wage literature: it distinguishes between aggregate and firm-level effects of minimum wage increases, and it uses several sources of panel data to examine the effect of the minimum wage on employment, prices, and the number of establishments in the restaurant and retail clothing industries. The results do not support the dynamic monopsony model, but there does not appear to be a negative correlation between employment and wages in the restaurant industry.

In Chapter 2, I examine whether political endogeneity explains why minimum wage increases do not appear to be negatively correlated with employment changes. A political economy analysis of federal and effective minimum wage levels shows that minimum wages are negatively correlated with the teen unemployment rate and positively correlated with the average manufacturing wage. Controlling for endogeneity should lower the coefficient on a minimum wage variable in teen employment rate regressions. However, neither of the two methods used here appreciably changes the estimated relationship between employment and

the minimum wage.

The employment and wage consequences of two minimum wage increases in Washington State are discussed in Chapter 3. I focus on low-wage retail trade and service industries. Although the minimum wage increases appear to have raised average wages, there does not appear to be a significant disemployment effect. A political economy analysis of voting behavior developed in the chapter potentially explains the failure to find negative employment effects. The minimum wage increases were approved by the majority of voters in every county in Washington through a ballot initiative, so perhaps it is not surprising that there were no substantial adverse effects.

This thesis perhaps raises more questions than it answers. Neither dynamic monopsony nor political endogeneity explains the failure to find empirical evidence that increases in the minimum wage reduce employment among affected groups. This leaves open the question of the validity of the neoclassical model in low-wage labor markets. Fortunately, it reveals many avenues for future research.

Chapter 1: Empirical Evidence on Dynamic Monopsony and Minimum Wages

"I've studied the arguments and the evidence for and against a minimum wage increase. I believe the weight of the evidence is that a modest increase does not cost jobs, and may even lure people back into the job market."
-- President Bill Clinton, 1995 State of the Union Address

1. Introduction

Recent research on the minimum wage has renewed interest in the monopsony model. Several studies which find that increases in the minimum wage do not decrease employment among teens and restaurant workers suggest that firms may face an upwards-sloping labor supply curve.[1] There are two primary models in which firms face a labor supply curve that is not perfectly elastic with respect to the wage: static monopsony and dynamic monopsony. The traditional example of a static monopsonist is a "company town" in which a firm has monopoly power in the labor input market. The dynamic monopsony model developed by Burdett and Mortensen (1989) assumes that there are frictions in the job search process due to imperfect information. This paper develops a similar model, adds a product market, and explores the effect of the form of product market competition on the predictions of the model. The validity of the model is then tested in the restaurant and retail clothing industries using establishment data on employment and prices.

In the traditional static monopsony model, the firm faces an upward-sloping labor supply curve and a downward-sloping marginal revenue product of labor curve. The firm

[1]See, for example, Card and Krueger (1994), Card, Katz, and Krueger (1994), Card (1992b), and Katz and Krueger (1992). In addition, Machin and Manning (1994) do not find an increase in employment in Britain as minimum wages fell there during the 1980s and suggest imperfect labor markets as an explanation. Card (1992a) also finds that binding increases in the federal minimum wage do not appear to reduce teen employment rates.

sets employment to equalize the marginal cost and marginal revenue product of labor. The equilibrium wage is then determined from the labor supply function, and the marginal revenue product exceeds the wage. Imposition of a minimum wage above the equilibrium wage increases employment until the wage floor exceeds the intersection of the labor supply and marginal revenue product of labor curves. This model has been empirically validated in several concentrated markets, including nurses (Sullivan, 1989) and school teachers (Landon and Baird, 1971). The static monopsony model does not apply very well to minimum-wage labor markets such as retail stores and restaurants, which are usually characterized by many small establishments and low skill requirements.

The dynamic monopsony model, in contrast, assumes that workers have imperfect information about the wages offered by firms, making it particularly plausible in an industry with many small firms. A worker receives offers at random intervals and accept an offer if it exceeds the current wage or the reservation wage if the individual is not working. Firms with lower wages have higher quit rates and lower hire rates. As Burdett and Mortensen (1989) show, imposition of a binding minimum wage acts as a lower bound on the wage distribution and raises the highest wage in the wage distribution. Under certain conditions, a minimum wage may increase employment, particularly at smaller, low-wage firms.

The dynamic monopsony model is particularly applicable in the restaurant and retail clothing sectors, which pay low wages and experience high turnover rates. Annual turnover at restaurants is about 250 percent, while employee turnover at retail stores is as high as 60 percent per year.[2] Vacancies are also high in the restaurant industry, with over 40 percent

[2]See Nation's Restaurant News (1988) and Friedman (1988).

of surveyed restaurant operators reporting they had at least one position vacant for more than two weeks in December 1989.[3] If turnover and vacancies are costly, a model of perfect competition predicts that the equilibrium wage will increase.[4] However, firms may not raise the offered wage if labor market frictions are present, as in the dynamic monopsony model. Offering a higher wage increases the cost of inframarginal workers but may not help a firm recruit additional workers if searchers are not aware of the higher wage. A binding minimum wage increase forces firms to raise the wage and overcomes the information problem if workers are aware of the wage floor. Employment may then increase.

Some recent empirical evidence accords with this view of the labor market. Card and Krueger (1994) and Katz and Krueger (1992), for example, find no disemployment effects of minimum wage increases in fast-food restaurants. Moreover, the National Restaurant Association (1990) reports that despite restaurant operators' concern about turnover and vacancies, managers are more likely to increase training than to increase wages or benefits to deal with the labor shortage.

This paper presents a dynamic monopsony model in which a binding increase in the wage floor reduces the quit rate and increases the number of potential hires. I assume that a minimum wage increase causes firms to increase employment. If firms face perfect competition in the product market, the minimum wage increase causes aggregate employment to fall and prices to rise if the industry faces a downward-sloping demand curve. The number of establishments decreases. In this model, a minimum wage can cause aggregate

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[3]See National Restaurant Association (1990).

[4]Hiring an entry-level manager or hourly employee who does not work out can cost up to $10,000 in wasted training, insurance, and productivity (Nation's Restaurant News, 1988).

employment to increase only if firms have market power in the labor input market *and* in the output market. Data on employment and prices at eating and drinking places and clothing stores over 1987-1992 are then used to test the model.

The rest of the paper proceeds with a brief review of the recent literature on the effects of the minimum wage on firms and employment. Section 3 presents a basic dynamic monopsony model and compares it to models of perfect competition and efficiency wages. Section 4 describes the data and Section 5 presents the empirical specifications and results. Section 6 concludes.

2. Summary of the Literature

Recent research on the relationship between the minimum wage and employment reaches disparate conclusions. Using establishment-level survey data, Card and Krueger (1994) and Katz and Krueger (1992) find either positive or no employment effects from state and federal minimum wage increases.[5] Similarly, Card (1992a and 1992b) and Card, Katz, and Krueger (1994) find no disemployment effects of federal and state minimum wage increases in data from the Current Population Survey (CPS). Neumark and Wascher (1992, 1994, and 1995b), in contrast, report that the minimum wage has a negative effect on teen employment based on CPS panel data. Using CPS data, Wellington (1991) reports small negative employment effects for teens based on time-series data, and Deere, Murphy, and Welch (1995) find large negative employment effects among low-wage workers. Kim and Taylor (1995) also report negative employment effects in California using establishment data

[5]Several studies, including Neumark and Wascher (1995a), have criticized the data and the methodology used by Card and Krueger (1994).

from the County Business Patterns.

Previous research has generally found no effect or small positive effects of minimum wage increases on food prices. Card (1992b) finds some positive but insignificant effects of the increase in California's minimum wage in 1988 on hamburger prices and the consumer price index for food away from home in several California cities. Katz and Krueger (1992) find no significant effect of the 1991 federal minimum wage increase on fast-food prices in Texas. Card and Krueger (1994) report that prices in New Jersey increased relative to Pennsylvania after the 1992 minimum wage hike in New Jersey; however, price increases within New Jersey do not bear any consistent relationship to the impact of the minimum wage increase. Card and Krueger (1995) find that the 1990 and 1991 federal minimum wage increases had a positive, although generally not significant, effect on hamburger prices.

The effect of the minimum wage on the number of firms and on hire and quit rates has received little attention. Card and Krueger (1994) find positive correlations between measures of the change in the minimum wage in a state over 1986-1991 and the growth rate of the number of McDonald's restaurants. Holzer, Katz, and Krueger (1991) investigate the effect of the minimum wage on the number of job applicants and find that jobs that pay exactly the minimum wage have more applicants than jobs that pay either slightly more or slightly less.

3. The Model

A simple model of dynamic monopsony is proposed.[6] Following the models of Burdett and Mortensen (1989), Machin, Manning, and Woodland (1993), and Chapter 11 of Card and Krueger (1995), an individual firm has monopsony power in that it faces a labor supply curve that is not perfectly elastic. A firm can recruit and retain more workers if it increases its wage. The model is used to derive testable predictions for the effect of a minimum wage on establishment-level and aggregate employment and output, the number of establishments, and price. The product market is initially assumed to be perfectly competitive, and the predictions of the model under other forms of product market competition are discussed later.

Firms are assumed to be identical and to each have a Cobb-Douglas production function with two inputs, capital and labor. Production at each firm at time t is

$$q_t = \beta K^{\alpha} L_t^{1-\alpha}. \tag{1}$$

As in the dynamic monopsony model in Layard, Nickell, and Jackman (1991), the capital stock of each firm, K, and the cost of capital, r, are fixed. The firm chooses a production level by setting a wage to obtain the desired number of workers, L_t. Firms are assumed to face perfect competition in the output market and to accept the product price, p_t, as given.

A firm may experience turnover in its labor force through quits and/or hires and can influence this turnover through its wage. The number of quits at a firm at time t is $s(w_t)L_t$ where the quit rate $s(w_t)$ is decreasing in the wage, with $s'(w_t) < 0$ and

[6]The model presented here does not incorporate the imperfect search process detailed by Burdett and Mortensen (1989). The use of a search model is not necessary to derive the predictions of a dynamic monopsony model for employment, prices, and establishments. Only an upwards-sloping labor supply curve is needed, as in Machin, Manning, and Woodland (1993) and Card and Krueger (1995).

$s''(w_t) > 0$. The number of hires is $H(w_t)$, with $H'(w_t) > 0$ and $H''(w_t) < 0$.[7] It is assumed that a firm cannot fire workers, so all attrition occurs through quits.

Although the model presented below is in continuous time, it may be easier to think of the dynamics in discrete time. The timing is as follows: at the beginning of period t, the firm has L_t workers and announces a wage w_t. Production then takes place and output is $q_t = \beta K^{\alpha} L_t^{(1-\alpha)}$. At the end of the period, workers decide whether to quit based on that period's wage, and $s(w_t)L_t$ workers quit. Simultaneously, the firm hires $H(w_t)$ workers.[8] At the beginning of the next period, the number of workers at the firm is $L_{t+1} = (1 - s(w_t))L_t + H(w_t)$.

The firm wants to maximize the present discounted value of its profit stream with respect to the wage, or

$$\max_{w_t} \int_0^{\infty} \{p_t q(L_t) - w_t L_t - rK)\} \, e^{-\theta t} \, dt \tag{2}$$

$$s.t. \quad \dot{L} = H(w_t) - s(w_t)L_t$$

where θ is the discount rate and L is the law of motion for labor: the change in a firm's employment level is the number of hires less the number of quits.

The equilibrium can be found by first solving the intertemporal profit-maximization problem for a law of motion for the wage, $\dot{w}$. The derivation of $\dot{w}$ is shown in the Technical Appendix. The two laws of motion L and $\dot{w}$ characterize the dynamics of the system. In

[7]Making either hires or quits invariant to the wage does not change any of the implications of the model. Making both of them invariant to the wage, of course, results in a perfectly inelastic labor supply curve.

[8]The hire function actually gives the number of potential hires available to the firm. The firm may not opt to hire all potential employees. In the unconstrained equilibrium, the firm will hire all potential employees; otherwise, the firm could profitably lower its wage.

equilibrium, the quantity of labor employed and the wage are both constant, or $\dot{L} = 0$ and $\dot{w} = 0$. The $\dot{L} = 0$ curve is positively-sloped under the assumptions of the model and corresponds to an upward-sloping labor supply curve in the static monopsony model. The slope of the $\dot{L} = 0$ curve depends on the elasticities of the hire function and the quit rate with respect to the wage. The smaller the two elasticities, the more steeply-sloped is the $\dot{L} = 0$ curve. Jointly solving $\dot{L} = 0$ and $\dot{w} = 0$ gives the equilibrium level of employment, L^*, which is expressed here as a function of the equilibrium wage w^*:

$$L^* = \frac{H(w^*)}{s(w^*)}. \tag{3}$$

In equilibrium, each firm has an excess demand for labor in that the firm would prefer to hire more workers at the equilibrium wage. It can easily be shown that the marginal revenue product of labor exceeds the wage in the unconstrained equilibrium. The equilibrium is similar to the static monopsony model: the firm does not hire more workers since it faces an upwards-sloping labor supply curve. The cost of paying inframarginal workers the higher wage outweighs the benefit of hiring additional labor, thereby increasing output and revenue.

Equilibrium firm-level output and price can be derived as a function of the wage. Equilibrium output is

$$q^* = \beta K^{\alpha} \left\{\frac{H(w^*)}{s(w^*)}\right\}^{1-\alpha}. \tag{4}$$

Under perfect competition in the output market, price equals minimum average cost and each firm earns zero profits. The unconstrained equilibrium price as a function of the wage is

$$p^* = \frac{1}{\beta K^\alpha} \left\{\frac{H(w^*)}{s(w^*)}\right\}^{\alpha-1} \left\{\frac{w^* H(w^*)}{s(w^*)} + rK\right\} \tag{5}$$

In addition, perfect competition in the output market implies that firms must be producing at the minimum average cost, or that average cost is equal to marginal cost.[9] A functional form for industry demand in the product market, which is necessary to derive the equilibrium number of firms and aggregate employment, is assumed below. But first I examine the effect of a minimum wage on firm-level employment.

Minimum Wage Imposed

Suppose a minimum wage w_{min} that is greater than the equilibrium wage w^* is imposed.[10] There are three possibilities: the firm can increase employment, decrease employment, or keep the same number of workers at a higher wage. The firm is able to increase its employment level since the higher wage causes the quit rate to fall and the number of potential hires to increase. However, the increase in the wage floor may be large enough that the firm does not wish to hire more workers.[11] Because I am interested in developing the predictions of the model for the case in which employment increases, I assume that the minimum wage increase is large enough to be binding but small enough to

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[9]This yields a constraint on the equilibrium wage: $w^* = \frac{1-\alpha}{\alpha} \frac{s(w^*)}{H(w^*)} rK$.

[10]The analysis is the same if a wage floor at or below the equilibrium wage already exists and is increased. The analysis also does not change if the minimum wage increase is anticipated by firms.

[11]This is equivalent to a wage floor above the intersection of the marginal revenue product and labor supply curves in the static monopsony case model--any further increases in the minimum wage will reduce employment as the firm moves along the marginal revenue product curve instead of along the labor supply curve.

not lower the quantity of labor employed. It should of course be remembered that an increase in the minimum wage can reduce employment even in a monopsony model.

How much employment changes if the minimum wage raises firm-level employment depends on the slope of the $\dot{L} = 0$ curve; the lower the elasticities of the hire and quit functions with respect to the wage, the steeper the curve and the smaller the increase in employment in response to a given increase in the wage. If the minimum wage increase raises employment, it must also raise price since average cost has increased. Quantity produced should also increase as employment increases. If the firm remains on the $\dot{L} = 0$ curve, the constrained equilibrium values of employment, output, and price can be found by simply substituting w_{min} for w^* in Equations 3, 4, and 5.

Employment is only assumed to increase at the firm level. We need to assume a demand structure for the product market to determine how the minimum wage affects the number of firms and aggregate employment. Although each firm acts a price taker, the industry as a whole is assumed to face a downwards-sloping demand curve in the product market. I assume the industry demand curve has constant elasticity, or

$$p_t = (N_t q_t)^{\rho} \tag{7}$$

where N_t is the number of identical firms at time t and $\rho < 0$. Under free entry and exit, the number of firms adjusts to keep profits equal to zero; if price is greater than average cost and firms are earning profits, entry occurs until the industry returns to zero profits. Inspection of Equation 7 shows that the number of firms must fall if the minimum wage

increases employment since both price and quantity produced at each firm increase.[12] If the firm remains on the $\dot{L} = 0$ curve after the minimum wage is increased, the equilibrium number of firms can be found by substituting w_{min} for w^* in

$$N^* = \{\beta K^{\alpha}\}^{-\frac{1+\rho}{\rho}} \{\frac{H(w^*)}{s(w^*)}\}^{\frac{(\alpha-1)(\rho+1)}{\rho}} \{\frac{w^* H(w^*)}{s(w^*)} + rK\}^{\frac{1}{\rho}}. \tag{8}$$

Although employment at each firm is assumed to increase, aggregate employment decreases if the industry faces a downward-sloping demand curve and firms earn zero profits. Each firm experiences an increase in average cost when the wage floor increases, so price must increase for firms to continue to break even. For price to increase, either demand must increase or supply must decrease. If the demand curve does not shift, price can increase only if industry supply falls. The increase in marginal cost causes the industry supply curve to shift up. Total output and therefore total employment must fall. The model thus generates the surprising result that total employment can fall even though firm-level employment is assumed to increase.

In summary, the dynamic monopsony model predicts that a binding minimum wage increase causes the quit rate to fall and the number of potential hires to increase. If the minimum wage is assumed to raise firm-level employment and firms are perfectly competitive in the output market, price increases since average cost increases. Aggregate employment and output fall. The number of firms also falls.

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[12]If the number of firms remains constant when a minimum wage increase causes firm-level employment to rise, firms earn negative profits. Output at each firm increases, so total output increases if the number of firms is fixed. Price must fall if the industry faces a downward-sloping demand curve. Price falls while average cost rises, so profits are less than zero. The zero-profit equilibrium can be restored if the number of firms falls sufficiently.

There must be imperfect competition in the product market for an increase in the minimum wage to raise aggregate as well as firm-level employment.[13] An increase in the minimum wage that increases firm-level employment raises average cost regardless of the form of product market competition. For aggregate employment to increase, price must decrease, as is evident from Equation 7. Price can fall when average cost increases only if firms have market power and earn positive profits before the minimum wage increase; if firms are initially earning zero profits, an equilibrium with lower prices and higher costs is unsustainable. Firms must have enough product market power to absorb the loss in profits due to an increase in costs and a decrease in prices as a result of the minimum wage. A minimum wage increase thus can raise total employment and lower prices only if there is market power not only in the labor input market but also in the output market. The effect on the number of firms is indeterminate, but any decrease in the number of establishments must not offset the increase in firm-level employment if aggregate employment is to increase.

Other Models of the Labor Market

To assess the empirical validity of the dynamic monopsony model, we need to know how other models predict the minimum wage will affect employment, prices, and the number of firms. There are two primary alternatives to the monopsony model: perfect competition and efficiency wages. Both models are nested in the dynamic monopsony model developed here.

Under perfect competition, firms are price takers in the labor input market and hire

[13]Product market power can be incorporated in the above model by making price a function of quantity, and thus implicitly a function of employment, in Equation 2.

workers until the marginal revenue product of labor is equal to the wage. In terms of the model developed above, both the number of hires and the quit rate are infinitely elastic with respect to the wage. Each firm will decrease employment when a binding wage floor is imposed. The employment effect is larger the more elastic is labor demand. Output at each firm will fall, and average cost will increase. If firms are also perfectly competitive in the output market, price will increase as cost increases. Aggregate output and employment will also fall. We cannot a priori determine whether the number of firms increases or decreases when the minimum wage increases since the price and quantity effects are in different directions. The main predictions in which the perfect competition and dynamic monopsony models differ are thus firm-level employment and output.

The efficiency wage model has been offered by Rebitzer and Taylor (1995) as an explanation for positive employment effects of the minimum wage. Firms may offer a wage above the market equilibrium wage to reduce shirking by workers. The efficiency wage model assumes that firms can fire workers caught shirking and that workers reduce the level of shirking as the opportunity cost of being fired increases. The minimum wage can act as such an efficiency wage. The efficiency wage model can be viewed as a variant of the dynamic monopsony model if workers caught shirking and fired are viewed as quits. A shirking worker is caught and fired with some positive probability, and as the wage rises, the worker is less likely to shirk and therefore the probability of "quitting" by getting caught shirking falls. The predictions are the same as in the dynamic monopsony model since efficiency wages can be viewed as a special case of the quit function.[14]

[14]One way to determine whether the minimum wage acts as an efficiency wage is to examine whether the minimum wage has different effects on firms with different monitoring costs.

Table 1 summarizes the predictions of the perfect competition and monopsony models. For the monopsony predictions, it is assumed that an increase in the minimum wage increases firm-level employment.

Table 1
Summary of Predictions

Labor Market:	Dynamic Monopsony		Perfect Competition
Output Market:	Perfect Competition	Product Market Power	
Firm-level Employment	+	+	-
Aggregate Employment	-	+	-
Price	+	-	+
Number of Establishments	-	?	?

4. Data Description

This analysis focuses on eating/drinking places and retail clothing stores during the years 1987-1992. Restaurants and retail clothing stores are two of the largest employers of minimum wage workers; in 1988, about 32 percent of workers paid the federal minimum wage were employed in these two sectors.[15] This period is chosen because it includes two increases in the federal minimum wage and considerable variation in minimum wages across states. As the real federal minimum wage fell by one-third during the 1980s, an

[15]Based on the CPS Outgoing Rotation Group files for 1988.

unprecedented number of states raised their minimum wages above the federal level. The variation in effective minimum wages began with Maine raising its minimum wage above $3.35 in 1985, and 16 states exceeded the federal minimum wage by January 1990.[16] The federal minimum wage was increased from $3.35 to $3.80 in April 1990 and to $4.25 in April 1991. The year 1987 is chosen as a starting point for this analysis because minimum wages began to vary across more than a handful of states that year.

The restaurant industry is particularly likely to be affected by minimum wage changes since average wages are low and labor makes up a large proportion of costs. The restaurant industry traditionally has the highest payroll-to-sales ratio within the retail trade industry, with labor costs averaging about 30 percent of sales (Alpert, 1986). About 20 to 25 percent of all hourly workers earning the minimum wage are employed at eating and drinking places.[17] Wages are closely tied to the minimum wage, with between 20 and 30 percent of restaurant workers earning between the federal minimum wage and 25 cents above the federal wage floor over 1987-1992.[18] Fast-food places account for over 40 percent of sales at all eating/drinking places, and restaurants account for about 45 percent of sales.[19]

The minimum wage is also likely to have an impact on retail clothing stores. Over 5

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[16]See Chapter 1 of this thesis and Neumark and Wascher (1992) for a chronology of state minimum wage increases.

[17]The figures are based on the Current Population Survey Outgoing Rotation Group files over 1987-1992 for all workers paid an hourly wage.

[18]The federal minimum wage for tipped employees is lower than the wage floor used in the summary statistics presented here, so even more workers are affected by the federal minimum wage than these statistics imply. About 15 to 20 percent of employees at eating/drinking places earn less than the federal minimum wage and are likely to be covered by the tipped minimum wage. Figures are based on CPS data.

[19]See U.S. Bureau of the Census (1992), *Census of Retail Trade*.

percent of all hourly workers paid the federal minimum wage are employed at retail clothing stores, and 10 to 25 percent of all clothing store workers earned between the federal minimum wage and 25 cents above the federal minimum wage during the period 1987-1992.[20] The ratio of labor costs to sales ranges from about 10 percent at discount stores to almost 25 percent at specialty clothing stores (Bluestone, 1981).

The validity of the dynamic monopsony model is tested using several types of data. The effect of the minimum wage on employment is examined using establishment data on the number of employees at eating/drinking places and clothing stores. I also examine the effect of the minimum wage on the prices of several entrees at franchise restaurants, the prices of several items of clothing, and two price indexes. The relationship between the minimum wage and the number of eating/drinking places and clothing stores is also examined. The data are discussed in more detail below.

The U.S. Bureau of Labor Statistics publication *County Business Patterns* (CBP) is used to examine the effect of the minimum wage on the number of persons employed at eating/drinking places and clothing stores.[21] The CBP includes annual state-level data on the number of workers in retail trade in mid-March, the number of establishments, and total first quarter payroll. The data are collected from surveys of medium- and large-sized firms and from surveys and federal administrative records for small firms.

Card and Krueger (1995) criticize the CBP data for having large, random fluctuations

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[20]Retail clothing stores include department stores, apparel stores, and shoe stores. These figures are based on CPS data.

[21]As used here, clothing stores encompasses the CBP categories of general merchandise stores and apparel and accessories stores.

in industry employment that make inferences about the effect of the minimum wage sensitive to the choice of sample period and estimation technique. However, establishment data such as the CBP has several advantages over individual-level data such as the CPS. First, the administrative nature of some of the CBP data may make it more reliable than data form household surveys; some of the CBP data comes from Internal Revenue Service records, for example. In addition, the CBP provides more insight into firm-level behavior. The CBP data also has several advantages over phone-based surveys of firms such as those used in Card and Krueger (1994). The universe sample nature of the CBP avoids problems of having an unrepresentative sample, and the answers are more likely to be accurate than those given in phone surveys. However, the CBP data can only be used to examine a few outcomes, such as employment and the number of establishments. CBP data at the 2-digit SIC-code level from 1987-1992 are used in this study, which should reduce any concerns about the choice of an unrepresentative year or the use of data so disaggregated that measurement error becomes a serious problem.

The payroll figures given in the CBP are comparable to estimates based on other data sources, including the CPS Merged Outgoing Rotation Group files. Table 2 compares wage data for the CBP, the CPS, and the Current Employment Statistics (CES) program over the period 1986-1992.[22] Shown are national average pay per employee during the first quarter at eating/drinking places and clothing stores in the CBP, annual average hourly earnings and average hours in the CPS, and annual average weekly earnings in the CES. Both the levels

[22]The CES program collects data from monthly payroll records for a sample of non-farm business establishments. In 1992, the sample included over 370,000 establishments. The data are published in the U.S. Bureau of Labor Statistics publication *Employment, Hours, and Earnings, United States, 1981-1993* (1993). The CPS data used here are from the Merged Outgoing Rotation Group files and are not weighted.

and trends are fairly similar if the CPS and CES data are converted into quarterly payroll by multiplying average hours by wages by 13 in the CPS, or weekly wages by 13 in the CES. For example, nominal average payroll per employee at eating/drinking places is $1798 in the CBP for first quarter of 1990. The comparable figure in the CPS is $1776, and $1628 in the CES. The levels in both the CPS and the CES are consistently slightly lower than in the CBP, and the trends are smaller in the CES. One possible reason for the difference between payroll and hourly or weekly wages is that the CBP payroll data is supposed to include bonuses, benefits, tips, and the value of payments in kind (e.g. free meals and employee discounts).

Figure 1 shows the trends in aggregate employment and real average payroll per employee at eating/drinking places and retail clothing stores from the CBP data. Employment at eating/drinking places increased by 18 percent while employment at clothing stores increased by 8 percent over 1986-1992. Real average payroll in both sectors was relatively constant during this period. These figures provide the first means of examining the effect of the minimum wage on employment. The federal minimum wage increases in April of 1990 and 1991 should be reflected in the 1991 and 1992 data since the CBP data are from the first quarter. Total employment at eating/drinking places falls slightly over 1990-1991 and increases over 1991-1992. Real average payroll per employee at eating/drinking places falls over 1990-1991 and increases over 1991-1992. From a first glance at the data, employment and wages appear to move in the same direction in the restaurant industry. Total employment at clothing stores fell over 1990-1992, and real average payroll fell

slightly over 1990-1991 and increased slightly over 1991-1992.[23] If the minimum wage had an effect on clothing stores, it appears to have lowered employment.

Figure 2 shows the number of eating/drinking places, clothing stores, and gross national product over 1986-1992. The number of eating/drinking places increased by 18 percent, and the number of clothing stores was fairly constant. The number of eating/drinking places does not fall after the 1990 and 1991 minimum wage increases, and the growth rate of the number of eating/drinking places also does not slow down. The number of clothing stores remains virtually unchanged over 1990-1992. The minimum wage increases thus do not appear to be negatively correlated with the total number of establishments.

I use two types of price data: actual prices of six items and price indexes for two series. The American Chamber of Commerce Research Association (ACCRA) publishes a list of product prices in 200-300 cities on a quarterly basis. I use first-quarter data on the prices of three fast-food items and three clothing items in over 300 cities over 1987-1992. The items are a McDonald's cheeseburger, a two-piece chicken entree, a cheese pizza, a man's shirt, a package of boy's underwear, and a pair of Levi's jeans. These items are homogeneous since they are the same across cities and change little over time. The nominal average fast-food prices trend upwards during 1987-1992, with burger prices increasing the most of the three items.[24] The nominal price of underwear remains constant while the

[23]The declines in real average payroll over 1990-1991 are surprising since the minimum wage increased in most states during this period. CPS data also fail to indicate an increase in the real average hourly wage at eating/drinking places and clothing stores over 1990-1991.

[24]McDonald's dominates the restaurant industry, accounting for about 7 percent of total sales of food consumed away from home (Emerson 1990)0.

average prices of a shirt and jeans increase by 22 and 39 percent, respectively.

The price indexes are from the Bureau of Labor Statistics publication *CPI Detailed Report*, which includes price indexes for major cities on at least a bi-monthly basis. I use the annual CPI for food away from home and apparel for 26 cities over 1987-1992. The apparel index increases more slowly than the other two indexes during this period; the average increase in the apparel price index is 17 percent, compared with 23 percent in the overall CPI and 21 percent in the food away from home price index.[25]

Data on business-cycle conditions are also used in the empirical analysis. The data on the financial, insurance, and real estate sector are also from the CBP, and the unemployment rates for executives and population data are from the *Geographic Profile of Employment and Unemployment*.

5. Empirical Specification and Results

Data on prices and the number of firms are used to assess the validity of the dynamic monopsony model. The first issue is whether firm-level employment increases when a minimum wage increase raises wages. The model gives two sets of predictions if firm-level employment increases, and which one applies depends on the form of product market competition. If firms are perfectly competitive in the product market, aggregate employment should fall when the minimum wage is increased. Prices should increase, and the number of firms should fall. If firms have monopsony and monopoly power, then firm-level and total

[25]The clothing series is reported to have an unintentional downward bias relative to other industries (Pryor, 1995).

employment should increase and prices should fall as the minimum wage is increased. The effect on the number of firms is indeterminate, but any negative effect should not be large enough to offset the increase in firm-level employment. Firms may have monopsony power even if an increase in the minimum wage does not raise firm-level employment, but most of the minimum wage changes during this period are probably not large enough to reduce the number of jobs in the monopsony model.

A general empirical framework is used here to examine the effect of the minimum wage on wages, employment, prices, and the number of establishments in annual panel data over 1987-1992. The predictions of the dynamic monopsony model can by tested by regressing an outcome variable, such as employment, on the minimum wage and control variables:

$$y_{st} = \alpha + \beta_1 X_{st} + \beta_2 X_{st-1} + \gamma_1 MW_{st} + \gamma_2 MW_{st-1} + \delta_s + \sigma_t + \epsilon_{st} \tag{9}$$

where y_{st} is the log of an outcome variable, the vector X_{st} is a set of variables to control for changes in the business cycle, and MW_{st} is the log of the real effective minimum wage (the higher of the state and federal minimum wages) in state s in year t. Both the contemporaneous and the lagged value of the minimum wage and business-cycle controls are included to allow for lagged effects.[26] The variables δ_s and σ_t are included to capture any state-specific effects that are invariant over time and any year effects that are common to all states. Several sets of variables are used to control for business cycle effects, including regional time trends, state-specific time trends, and outcomes in other industries. The

[26]Neumark and Wascher (1992 and 1995b) report significant lagged effects of minimum wages on employment.

standard errors are corrected for heteroscedasticity in all estimates. For some outcomes, estimation results without including lagged terms in Equation 9 are also reported:

$$y_{st} = \alpha + \beta_1 X_{st} + \gamma_1 MW_{st} + \delta_s + \sigma_t + \epsilon_{st} \qquad (9')$$

Equations 9 and 9′ are first estimated with payroll data to see if the minimum wage had an effect on wages at eating/drinking places and clothing stores.

5.1 Wage Effects

Data from the County Business Patterns are used to examine the effect of the minimum wage on wages, but the CBP data do not include an average hourly wage. Instead, average payroll per employee is created by dividing first-quarter payroll by the number of employees in mid-March. As in Kim and Taylor (1995), this serves as the wage variable. The real effective minimum wage (the higher of the state and federal minimum wages) is used as the minimum wage variable. One of three sets of business-cycle controls is used: linear time trends for three of the four Census regions; state-specific linear time trends; and average payroll per employee in the finance, real estate, and insurance sector, and the unemployment rate among executives and managers. The two state-level variables used in the last method are intended to capture the effect of business-cycle conditions without being correlated with shocks to the retail trade and restaurant sectors.[27] A full set of state and

[27]Employment in the financial, insurance, and real estate (FIRE) sector and the unemployment rate of executives and managers are unlikely to be well-correlated with the minimum wage but should reflect aggregate economic changes. These variables are used instead of the unemployment rate of prime-aged males, the usual business-cycle control, because employment and wages of that group are more likely to be affected by the minimum wage than executives/managers. For example, union wage scales are typically based on the minimum wage.

year dummy variables is included in each regression. Equations 9 and 9′ are estimated separately for each industry for the period 1987-1992.

There is a clear positive relationship between the minimum wage and average payroll per employee at eating and drinking places. As shown in Table 3, the contemporaneous minimum wage is positively correlated with wages at restaurants in all specifications. The estimates indicate that a 10 percent increase in the minimum wage raised average wages by about 2 percent. The lagged minimum wage is also positively correlated with the average wage, also with an estimated elasticity of about 0.2. The cumulative effect is also positive and statistically significant. The estimates imply that the 18 percent increase in the real federal minimum wage over 1990-1991 increased real average payroll per employee by about 7 percent at eating/drinking places. The estimated coefficients on the contemporaneous minimum wage are larger if the effect of the lagged minimum wage is not controlled for, with an estimated elasticity of about 0.27.

The minimum wage is not as highly correlated with average payroll at clothing stores. The contemporaneous minimum wage is positively but weakly correlated with wages in all specifications. The cumulative effect of the current and the lagged minimum wage on payroll is small and not significantly different from zero. The contemporaneous minimum wage is also only weakly positively correlated with average payroll when the lagged minimum wage is not included in the regressions. These results are surprising given that over 15 percent of workers at clothing stores nationally were earning between $3.35 and $3.80 in 1989, according to the CPS. Given the weak estimated relationship between the minimum wage and average payroll per employee, it would be surprising if the minimum

wage had much of an effect on employment, prices, or the number of stores in the retail clothing industry over 1988-1992.

5.2 Employment Effects

Data on aggregate and establishment-level employment at eating and drinking places and clothing stores are used to examine the validity of the monopsony model. An increase in firm-level employment is consistent with the dynamic monopsony model. but aggregate employment does not necessarily increase even if an increase in the minimum wage raises firm-level employment. The CBP gives total employment and the number of establishments, so average employment per establishment in a state is used as a proxy for establishment-level employment here. Equations 9 and 9′ are estimated with the log of either total or average employment as the dependent variable. The equations are estimated for the 50 states over the years 1987-1992. As in the wage equations, one of three sets of variables is used to control for changes in the business cycle. The last set of business cycle controls includes the log of total or average employment in the financial industry, the unemployment rate of professionals, and a full set of state and year dummy variables.

Table 4 displays the estimation results. The contemporaneous minimum wage is positively correlated with total employment at eating/drinking places in all specifications. The estimated elasticities of employment with respect to the current minimum wage range from below 0.1 to above 0.15. The lagged minimum wage is negatively correlated with employment when regional time trends or economic conditions are controlled for. The cumulative effect of the current and the lagged minimum wage is statistically significant in

only one specification, and the estimated elasticity of total employment is about 0.15 when state time trends are controlled for.

The minimum wage is not highly correlated with average employment at eating/drinking places. The results in Table 4 indicate that the contemporaneous minimum wage is positively correlated with average employment when controlling for regional time trends or economic conditions. However, the negative effect of the lagged minimum wage on average employment appears to be larger than the positive contemporaneous effect. The estimated cumulative effects are negative but not statistically different from zero.

The minimum wage generally is not correlated with either total or average employment at clothing stores, which is not surprising since wages also are not correlated with the minimum wage during this period. The estimation results imply a positive relationship between the contemporaneous minimum wage and total employment and a negative relationship between the lagged minimum wage and total employment at clothing stores. The cumulative effects are not statistically different from zero. When the lagged minimum wage is not included in the regression, the estimated relationship between the current minimum wage and total employment is positive if economic conditions are controlled for; however, the result is not robust to the inclusion of regional or state time trends. None of the estimated coefficients on the minimum wage variables are statistically significant at conventional levels in the average employment regressions, but all of the estimated coefficients are negative.

5.3 Price Effects

Both the dynamic monopsony and perfect competition models predict that prices will be positively correlated with the minimum wage if firms do not have monopoly power. Prices should be negatively correlated with the minimum wage if firms have monopsony and monopoly power and an increase in the minimum wage raises total employment. I use two types of data to investigate the relationship between prices and the minimum wage: product prices and consumer price indexes.

There are several reasons why prices for the same item or inflation rates might differ across locations. Input or wholesale prices for clothing items and food products should not vary greatly across locations, so price differences across cities are likely to reflect differences in rents, transportation costs, and labor costs. Differences in prices may also reflect varying degrees of market power across geographic markets. If differences across cities in rents, transportation costs, and market power do not change much over time, differences in prices are likely to reflect differences in labor costs. As discussed earlier, labor-to-sales ratio range from 10 to 30 percent in the two industries examined here.

Equation 9 is used to estimate the effect of the minimum wage on prices. For the product price estimates, the log of the first-quarter product price in a city is the dependent variable.[28] All of the independent variables are state-level in the product price regressions. For the CPI estimates, the dependent variable is the log of the city consumer price index for food away from home or apparel in 26 cities. All of the independent variables in the CPI

[28]Data from the first quarter are used since there are likely to be seasonal effects that differ across quarters, and using data only from the first quarter avoids problems with back-to-school and holiday sales at clothing stores.

regressions are city-level except the minimum wage, which is state-level. Again on of three sets of business cycle controls is used: regional time trends, state time trends, or the overall inflation rate. A full set of year and state or city dummy variables is included in each regression.[29] For the product price estimates, the regional consumer price indexes for all items and for either food away from home or apparel are included in the last set of controls.[30] For the CPI estimates, the city overall CPI is included.

Table 5 gives the regression results for the product prices. As shown in the first column, the price of a McDonald's hamburger is positively correlated with the contemporaneous minimum wage but not highly-correlated with the lagged minimum wage. A 10 percent increase in the minimum wage is correlated with about a 1 percent increase in the price of a hamburger.[31] This estimate is similar to results reported for a fast-food main course in Katz and Krueger (1992) and for a subsample of the ACCRA data in Card and Krueger (1995). The contemporaneous minimum wage is also positively correlated with the price of a chicken entree. The estimated cumulative effects imply that a 10 percent increase in the minimum wage is correlated with about a 1.5 percent increase in entree prices at fast-food chicken places. The minimum wage is generally not highly correlated with the price of a cheese pizza.[32] More extensive use of tipped employees at pizza places than at other fast-

[29]State instead of city dummy variables are used in the product price estimates because of the large number of cities (338).

[30]Inflation rates are not available at the state level.

[31]A previous version of this paper reported a negative relationship between the annual change in the real price of a hamburger and the change in the real effective minimum wage. This negative relationship does not hold unless a first-difference estimation technique is used.

[32]All of the results are the same if the lagged minimum wage is not included in the regressions.

food franchises may account for the difference in the relationship between the minimum wage and entree prices at pizza places and other fast-food places.

The prices of clothing items generally are not correlated with either the contemporaneous or the lagged minimum wage. As shown in last three columns of Table 5, few of the estimated coefficients for the minimum wage variables are statistically significant. The results imply a positive relationship between the minimum wage and the price of a man's shirt and the price of jeans and a negative relationship between the minimum wage and the price of a package of boys' underwear, but few of the estimated coefficients are statistically different from zero. Interestingly, the weak positive price effects for a shirt and jeans are due to the lagged, not the contemporaneous, minimum wage.[33] These results are consistent with the weak relationship between the minimum wage and average payroll per employee at clothing stores and general merchandise stores.

Table 6 gives the regression results for the city-specific consumer price indexes. The food away from home price index is positively correlated with the contemporaneous minimum wage in most specifications. The estimated coefficients indicate that a 10 percent increase in the minimum wage increases the inflation rate for food away from home by about 0.7 percentage points. The estimated coefficients on the lagged minimum wage are not statistically significant, and the estimated coefficients on the current minimum wage are also not statistically significant if the overall CPI is included as a regressor. These results are consistent with the findings from the regressions for product prices; since the food away from home price index includes prices at sit-down restaurants with tipped workers as well as

[33]If the lagged minimum wage is not included in the regressions, the estimated coefficients on the current minimum wage are similar to those reported in Table 5 and are not statistically significant.

at fast-food places, the effect of the minimum wage is likely to be smaller when restaurants and fast-food places are combined than when only fast-food places are examined. The apparel price index is not highly correlated with the minimum wage, as shown in the second column of Table 6.[34] Again, this result is consistent with the lack of wage effects of the minimum wage in this sector.

5.4 Number of Establishments Effects

The dynamic monopsony model predicts that the minimum wage should be negatively correlated with the number of establishments if an increase in the minimum wage raises firm-level employment but lowers aggregate employment. The model does not give a clear prediction if an increase in the minimum wage reduces both average and total employment or increases both. Data from the CBP on the number of eating/drinking places and general merchandise and apparel stores in each state over 1987-1992 are used to examine the relationship between the minimum wage and the number of establishments.

Equation 9 is again used to estimate the effect of the minimum wage, and the dependent variable is the log of the number of establishments. One of three sets of variables is used to control for business-cycle conditions: regional time trends, state time trends, or three state-specific economic variables. The three economic variables included in the last method are contemporaneous and lagged values of the log of the number of financial, insurance, and real estate establishments, the unemployment rate of executives and managers, and the log of population. A full set of state and year fixed effects is included in each

[34]All of the results are similar if the lagged minimum wage is not included as a regressor.

regression.

The estimation results are shown in Table 7. The number of eating/drinking places is positively correlated with the contemporaneous minimum wage, and the estimated cumulative effect is positive and statistically significant when controlling for regional or state time trends. The results indicate that a 10 percent increase in the minimum wage is correlated with about a 1 to 1.7 percent increase in the number of eating/drinking places. The results are similar when the lagged minimum wage is not included as a regressor. The estimation results for clothing stores indicate a positive relationship between minimum wages and the number of apparel and general merchandise stores, but none of the estimated coefficients are statistically significant at conventional levels.

5.5 2SLS Variables Estimates of Effects

The above estimates are reduced-form relationships between the minimum wage and employment, prices, and the number of establishments. The strong relationship between the minimum wage and average payroll per employee at eating/drinking places can be used to estimate the relationships between average wages and employment or the number of establishments in the restaurant industry.[35] The minimum wage serves as exogenous source of variation in wages across states over time since wages should be the only mechanism through which the minimum wage affects employment and the number of restaurants. The relationship between wages and employment or the number of establishments can be

[35]The 2SLS regressions are not estimated for the clothing industry because of the weak relationship between the minimum wage and payroll. They are also not estimated for the price regressions since wage data is not available at the establishment or city level.

estimated with two-stage least squares (2SLS) estimation. The first-stage regression is Equation 9 with the log of contemporaneous or lagged real average payroll per employee as the dependent variable. Either the log of total employment, average employment per establishment, or the number of establishments is the dependent variable in the second-stage regression

$$y_{st} = \alpha + \beta_1 X_{st} + \beta_2 X_{st-1} + \gamma_1 WAGE_{st} + \gamma_2 WAGE_{st-1} + \delta_s + \sigma_t + \epsilon_{st} \qquad (10)$$

where the log of the contemporaneous wage and the lagged wage are both instrumented.[36] One of several sets of variables is again used to control for business cycle effects in X_{st}: regional time trends, state-specific time trends, or outcomes in other industries. A full set of state and year dummy variables is also included in all regressions.

The results do not indicate positive cumulative effects of wages on total or average employment at eating/drinking places. The contemporaneous wage is positively correlated with total employment and average employment in almost all specifications, as shown in Table 8, and many of the estimated coefficients are statistically significant. However, the negative correlation between the lagged wage and average or total employment is larger than the positive correlations between the current wage and the employment variables in most specifications. The estimated cumulative elasticities of total employment with respect to wages range from about -0.4 to 0.3. The estimated cumulative elasticities of average employment range from about -0.5 to -0.2. Thus, the results do not provide strong support for a positive relationship between employment and wages in the restaurant industry.

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[36]Equation 10 is also estimated without including $WAGE_{st-1}$ and X_{st-1}. The current wage, $WAGE_{st}$, is then instrumented with only the current minimum wage. The first-stage regression is equivalent to Equation 9′ with the log of real average payroll per employee as the dependent variable.

The results do indicate a positive relationship between wages and the number of eating and drinking places. As shown in the last two columns of Table 8, the current wage is positively correlated with the number of establishments, and most of the estimated coefficients are statistically significant. The cumulative effect is also positive and significant when regional or state time trends are included in the regressions.

6. Conclusion

The model presented in this paper differs from previous dynamic monopsony models in that it develops predictions at the industry level as well as at the firm level and examines how the type of competition in the output market affects the predictions of the model. A binding minimum wage increase may result in an increase in firm-level employment and a decrease in total employment if there is perfect competition in the output market. Positive employment effects can occur at both the firm- and the aggregate-level only if firms have market power in both the input and output markets.

Panel data on payroll at eating/drinking places indicate that the minimum wage is strongly correlated with wages during the years 1987-1992. The contemporaneous minimum wage is positively correlated with total employment at eating/drinking places, a result that might indicate the presence of monopsony. However, the lagged minimum wage is negatively correlated with total employment in several specifications, and the cumulative effect is positive and statistically significant in only one specification. In addition, the minimum wage is generally not significantly positively correlated with firm-level employment--total employment divided by the number of establishments--over 1987-1992.

The prices of two entrees at eating/drinking places and the food away from home price index are positively correlated with the minimum wage in most specifications investigated here. This result is not consistent with the monopsony model in the case where total employment increases as a result of a minimum wage increase. The minimum wage is positively correlated with the number of eating/drinking places in some specifications. The 2SLS results also do not consistently indicate the presence of monopsony power and positive employment effects, particularly when the lagged wage is included in the regressions.

The minimum wage is not strongly correlated with average payroll per employee at clothing stores over 1987-1992, which is surprising given that the retail clothing industry is one of the largest retail employers of low-wage workers. The regression results for clothing stores generally fail to show any relationship between the minimum wage and employment, prices, or the number of stores, which is not surprising if wages in the retail clothing sector are not affected much by the minimum wage.

The combined results thus do not consistently indicate the presence of monopsony in the restaurant industry. However, there do appear to be robust positive relationships between the contemporaneous minimum wage and wages, total employment, prices, and the number of establishments. Potential explanations for this empirical phenomenon include substantial income effects as a result of minimum wage increases that have particularly strong effects on the demand for food away from home; and the possibility that the minimum wage is endogenous, perhaps with respect to conditions in the eating/drinking places industry, and minimum wage increases occur when employment, prices, and the number of establishments are rising. The apparent negative relationship between the lagged minimum wage and many

of the outcomes investigated in this paper may indicate the presence of adjustment costs in the eating/drinking places industry, and firms may replace workers with equipment over the long-run as wages rise. The validity of these hypotheses should be further investigated. In addition, a true test of the dynamic monopsony model requires data on hire and quit rates. Collection and analysis of such data is a natural avenue for future research.

Technical Appendix

The intertemporal profit-maximization problem can be expressed as the Hamiltonian

$$\mathcal{H}_t = \{p_t q(L_t) - w_t L_t - rK + \lambda_t(H(w_t) - s(w_t)L_t)\}\, e^{-\theta t} \tag{T1}$$

The first-order conditions for the intertemporal profit-maximization problem under perfect competition are

$$\frac{\partial \mathcal{H}_t}{\partial w_t} = 0 \Leftrightarrow \lambda_t = \frac{L_t}{H'(w_t) - s'(w_t)L_t} \tag{T2}$$

and

$$\dot{\lambda} = -\frac{\partial \mathcal{H}_t}{\partial L_t} e^{\theta t} + \theta\lambda_t \Leftrightarrow \dot{\lambda} = -p_t q'(L_t) + w_t + \lambda_t(s(w_t) + \theta) \tag{T3}$$

where $q'(L_t)$ is the marginal product of labor.

The transversality condition is

$$\lim_{t\to\infty} L_t\, \lambda_t\, e^{-\theta t} = 0 \Leftrightarrow \lim_{t\to\infty} \frac{L_t^2}{e^{\theta t}\,(H'(w_t) - s'(w_t)L_t)} = 0 \tag{T4}$$

Differentiating (T2) with respect to t and combining the results with (T3) yields the law of motion for the wage, $\dot{w}$:

$$\dot{w} = \frac{(H'(w)-s'(w)L)^2(pq'(L)-w) + H'(w)(H(w)-s(w)L) + L(H'(w)-s'(w)L)(s(w)+\theta)}{L(H''(w)-s''(w)L)} \tag{T5}$$

where the subscripts are omitted for notational convenience.

In equilibrium, $\dot{L} = 0$ and $\dot{w} = 0$. We cannot solve explicitly for L^* and w^* as a function of the parameters of the model without assuming functional forms for the hire and

quit functions. Employment, output, and price are therefore expressed as a function of w^*, the equilibrium wage, and can be derived using either of the $\dot{L} = 0$ and $\dot{w} = 0$ equations. The equations given in the text are derived from the $\dot{L} = 0$ equation since the $\dot{L} = 0$ locus will hold in the constrained as well as the unconstrained equilibrium.

Figure 1
Aggregate Employment and Real Average Payroll per Employee

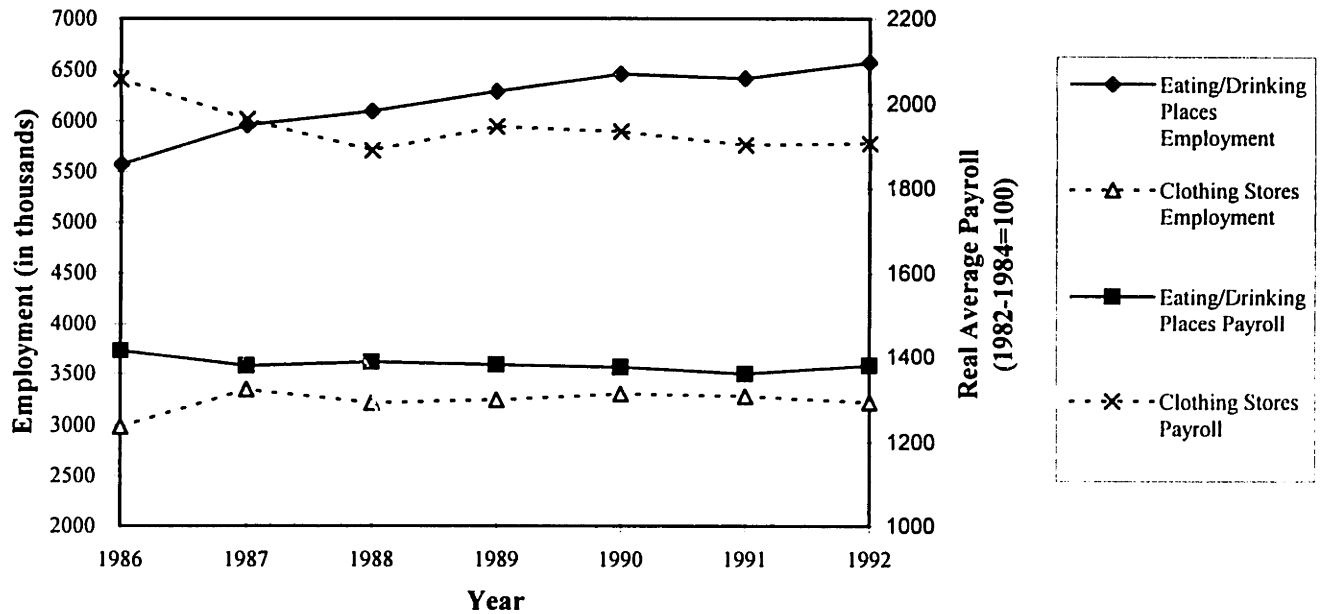

Figure 2
Number of Establishments and Real GNP

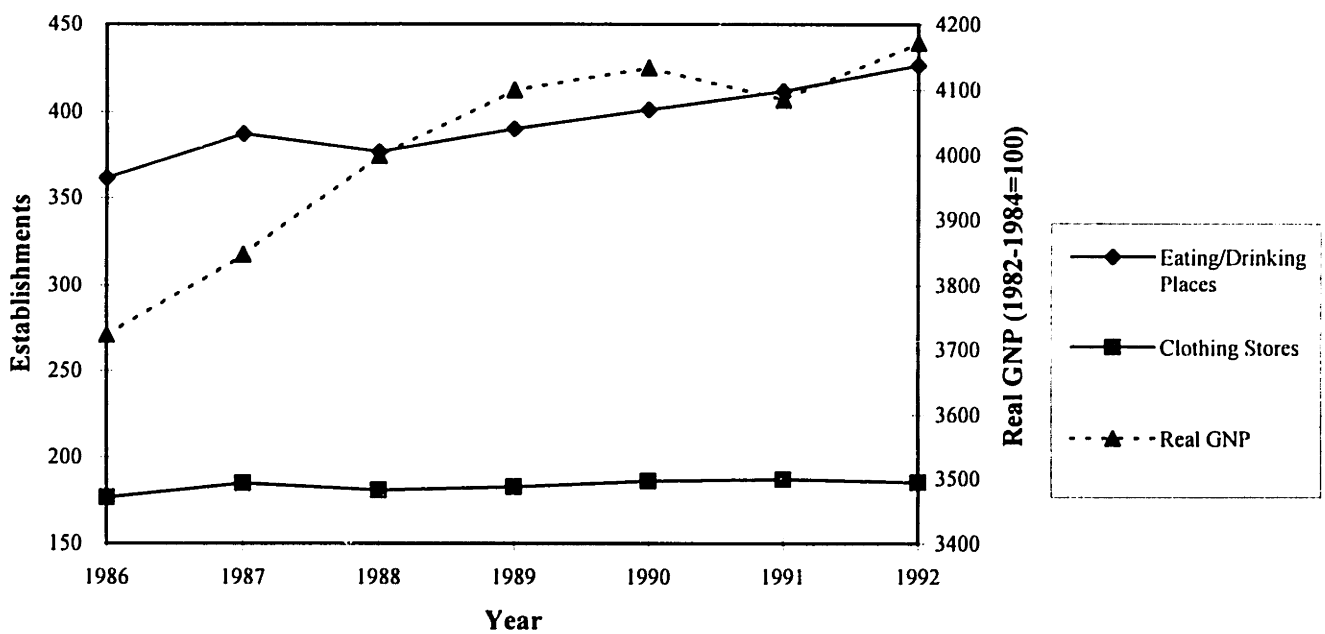

Table 2
Comparison of County Business Patterns Data to Alternate Sources

Year	CBP Quarterly Payroll	CPS Hourly Wages	CPS Hours	CPS Hours*Wages*13	CES Weekly Earnings*13
			Eating/Drinking Places		
1986	1551.07	3.88	28.7	1447.63	1447.68
1988	1641.38	4.13	29.1	1562.38	1526.85
1990	1797.56	4.63	29.5	1775.61	1628.12
1992	1935.18	4.97	29.1	1880.15	1719.25
			Clothing Stores		
1986	2222.20	4.95	30.1	1936.94	1870.96
1988	2280.50	5.31	30.1	2077.80	2024.75
1990	2557.90	6.59	30.5	2612.94	2193.75
1992	2725.97	6.64	31.8	2744.98	2370.16

Notes: All wage data are nominal. The CBP data is total national first-quarter payroll divided by mid-March employment for eating/drinking places and general merchandise and apparel and accessories stores. The CPS data are unweighted annual averages from the Outgoing Rotation Group files for eating/drinking places and department stores, apparel stores, and shoe stores. The CES data are annual averages for eating/drinking places and apparel and accessories stores.

Table 3
Effect of the Minimum Wage on Average Payroll

Specification	Eating/Drinking Places		Clothing Stores	
Regional Trends				
Minimum Wage	.193	.267	.058	.065
	(.037)	(.037)	(.041)	(.038)
Lagged Minimum Wage	.193	--	.020	--
	(.036)		(.037)	
Cumulative Effect	.386	--	.078	--
	(.041)		(.045)	
State Trends				
Minimum Wage	.221	.270	.072	.047
	(.031)	(.031)	(.038)	(.037)
Lagged Minimum Wage	.191	--	-.104	--
	(.038)		(.038)	
Cumulative Effect	.412	--	-.031	--
	(.044)		(.044)	
Economic Conditions				
Minimum Wage	.198	.268	.072	.052
	(.034)	(.035)	(.041)	(.033)
Lagged Minimum Wage	.172	--	-.020	--
	(.034)		(.036)	
Cumulative Effect	.370	--	.052	--
	(.039)		(.040)	

Notes: The dependent variable is the log of real average payroll per employees at eating and drinking places or at general merchandise and apparel stores. All specifications include a constant and a full set of state and year dummy variables. The data are state-level observations for a maximum of 300 observations over 1987-1992. Standard errors are in parentheses.

Table 4
Effect of the Minimum Wage on Total and Average Employment

Specification	Eating/Drinking Places Total		Eating/Drinking Places Average		Clothing Stores Total		Clothing Stores Average	
Regional Trends								
Minimum Wage	.142	.096	.082	.019	.064	.004	-.217	-.344
	(.047)	(.041)	(.042)	(.042)	(.056)	(.053)	(.237)	(.274)
Lagged Minimum Wage	-.120	--	-.163	--	-.165	--	-.354	--
	(.050)		(.048)		(.065)		(.204)	
Cumulative Effect	.022	--	-.081	--	-.102	--	-.571	--
	(.051)		(.055)		(.072)		(.364)	
State Trends								
Minimum Wage	.087	.102	-.004	-.011	.039	.033	-.316	-.337
	(.032)	(.031)	(.032)	(.031)	(.038)	(.036)	(.276)	(.262)
Lagged Minimum Wage	.058	--	-.027	--	-.025	--	-.097	--
	(.044)		(.046)		(.039)		(.180)	
Cumulative Effect	.145	--	-.031	--	.014	--	-.412	--
	(.046)		(.046)		(.045)		(.257)	
Economic Conditions								
Minimum Wage	.161	.131	.111	.087	.094	.150	-.165	-.238
	(.054)	(.055)	(.050)	(.052)	(.062)	(.072)	(.230)	(.284)
Lagged Minimum Wage	-.190	--	-.166	--	-.092	--	-.318	--
	(.057)		(.055)		(.073)		(.231)	
Cumulative Effect	-.029	--	-.055	--	.002	--	-.483	--
	(.054)		(.061)		(.083)		(.384)	

Notes: The dependent variable is the log of the total or average number of employees at eating and drinking places or at general merchandise and apparel stores. All specifications include a constant and a full set of state and year dummy variables. The data are state-level observations for a maximum of 300 observations over 1987-1992. Standard errors are in parentheses.

Table 5
Effect of the Minimum Wage on Product Prices

Specification	Hamburger	Chicken	Pizza	Shirt	Underwear	Jeans
Regional Trends						
Minimum Wage	.083	.180	.015	-.001	-.074	.062
	(.036)	(.076)	(.053)	(.095)	(.118)	(.096)
Lagged Minimum Wage	-.001	.013	.089	.142	-.003	.160
	(.026)	(.061)	(.045)	(.083)	(.116)	(.078)
Cumulative Effect	.082	.193	.103	.141	-.077	.222
	(.042)	(.087)	(.069)	(.118)	(.148)	(.111)
State Trends						
Minimum Wage	.089	.172	.025	-.006	-.070	.088
	(.039)	(.078)	(.055)	(.100)	(.121)	(.101)
Lagged Minimum Wage	.037	-.011	.026	.160	-.010	.118
	(.031)	(.071)	(.053)	(.090)	(.135)	(.097)
Cumulative Effect	.126	.161	.051	.154	-.080	.206
	(.051)	(.098)	(.080)	(.135)	(.176)	(.140)
CPI Controls						
Minimum Wage	.116	.156	.032	-.009	-.014	.056
	(.035)	(.075)	(.054)	(.096)	(.122)	(.098)
Lagged Minimum Wage	.005	-.017	.070	.135	.002	.148
	(.027)	(.061)	(.045)	(.083)	(.115)	(.077)
Cumulative Effect	.121	.139	.102	.127	-.012	.204
	(.041)	(.082)	(.066)	(.118)	(.149)	(.113)

Notes: The dependent variable is the log of the real first-quarter product price. All specifications include a constant and a full set of state and year dummy variables. The data are an unbalanced panel of quarterly observations for 338 cities over 1987-1992, and the number of observations is 1447. Standard errors are in parentheses.

Table 6
Effect of the Minimum Wage on Price Indexes

Specification	Food Away from Home	Apparel
Regional Trends		
Minimum Wage	.068	-.045
	(.033)	(.066)
Lagged Minimum Wage	.035	.127
	(.032)	(.093)
Cumulative Effect	.103	.082
	(.033)	(.116)
City Trends		
Minimum Wage	.071	-.056
	(.029)	(.051)
Lagged Minimum Wage	-.001	-.053
	(.026)	(.077)
Cumulative Effect	.070	-.109
	(.026)	(.100)
CPI Controls		
Minimum Wage	.109	-.031
	(.176)	(.063)
Lagged Minimum Wage	.040	-.005
	(.181)	(.073)
Cumulative Effect	.148	-.036
	(.173)	(.096)

Notes: The dependent variable is the log of the consumer price index for food away from home or apparel. All specifications include a constant and a full set of state and year dummy variables. The data are a panel for 26 cities, and the number of observations is a maximum of 156 over 1987-1992. Standard errors are in parentheses.

Table 7
Effect of the Minimum Wage on the Number of Establishments

Specification	Eating/Drinking Places	Clothing Stores
Regional Trends		
Minimum Wage	.060	.239
	(.039)	(.224)
Lagged Minimum Wage	.044	.249
	(.033)	(.194)
Cumulative Effect	.104	.488
	(.040)	(.341)
State Trends		
Minimum Wage	.091	.330
	(.031)	(.243)
Lagged Minimum Wage	.085	.083
	(.033)	(.170)
Cumulative Effect	.175	.413
	(.038)	(.231)
Economic Conditions		
Minimum Wage	.047	.218
	(.037)	(.210)
Lagged Minimum Wage	-.036	.216
	(.031)	(.198)
Cumulative Effect	.011	.434
	(.036)	(.335)

Notes: The dependent variable is the log of the number of eating and drinking places or general merchandise and apparel stores. All specifications include a constant and a full set of state and year dummy variables. The data are state-level observations for a maximum of 300 observations over 1987-1992. Standard errors are in parentheses.

Table 8
2SLS Estimates of the Effect of Wages

Specification	Total Employment		Average Employment		Establishments	
Regional Trends						
Wage	.867	.359	.557	.071	.311	.288
	(.381)	(.228)	(.339)	(.225)	(.203)	(.143)
Lagged Wage	-1.088	--	-1.034	--	-.054	--
	(.385)		(.342)		(.205)	
Cumulative Effect	-.220	--	-.477	--	.257	--
	(.241)		(.214)		(.128)	
State Trends						
Wage	.415	.376	.044	-.040	.371	.417
	(.326)	(.212)	(.300)	(.194)	(.163)	(.113)
Lagged Wage	-.123	--	-.257	--	.135	--
	(.535)		(.492)		(.267)	
Cumulative Effect	.292	--	-.214	--	.506	--
	(.330)		(.303)		(.165)	
Economic Conditions						
Wage	.941	.519	.798	.318	.509	.397
	(.374)	(.278)	(.350)	(.251)	(.189)	(.158)
Lagged Wage	-1.376	--	-1.225	--	-.332	--
	(.332)		(.335)		(.173)	
Cumulative Effect	-.434	--	-.427	--	.177	--
	(.254)		(.224)		(.139)	

Notes: The dependent variable is the log of the total or average number of employees at eating and drinking places or log of the number of eating and drinking places. The wage and the lagged wage are instrumented (see text for details). All specifications include a constant and a full set of state and year dummy variables. The data are state-level observations for a maximum of 300 observations over 1987-1992. Standard errors are in parentheses.

Chapter 2: Controlling for Endogeneity in Estimates of Minimum Wage Effects

"State minimum wages more closely reflect the distinct dynamics of the Nation's regional economics markets and work force than does the Federal minimum wage."
--Rep. David Drier (R, CA)

1. Introduction

The impact of the minimum wage on employment is one of the most hotly-disputed topics in economics. Whereas as recently as a decade ago 90 percent of economists agreed that an increase in the minimum wage reduces employment, several recent studies have challenged the economic orthodoxy.[1] A series of papers by Card, Katz, and Krueger find at most minimal disemployment effects from several federal and state minimum wage increases during the 1980s and 1990s.[2] However, a number of studies continue to support the conventional view that increases in the minimum wage adversely affect employment, including studies by Neumark and Wascher (1992 and 1994), Kim and Taylor (1995), and Deere, Murphy, and Welch (1995). All of these studies contribute to our understanding of the employment effects of the minimum wage, but virtually none of them have stopped to consider the political process that determines minimum wage rates.[3] By accepting minimum wage rates as given and focusing on employment outcomes, the existing literature may be underestimating the disemployment effects of the minimum wage. This paper examines the

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[1]See Kearl et. al. (1979) and Frey et. al. (1984) for statistics on economists' views on the minimum wage. Colander and Klamer (1987) show a similar degree of consensus among doctoral students in economics in 1985.

[2]See, for example, Card (1992a and 1992b), Card and Krueger (1994 and 1995), and Card, Katz, and Krueger (1994). Machin and Manning (1994) also find no evidence of positive employment effects in Britain when the minimum wage declined there during the 1980s.

[3]Card and Krueger (1994) is a notable exception. They discuss the political process underlying the 1992 New Jersey minimum wage increase and explain why it can plausibly be viewed as unrelated to economic conditions.

determinants of state and federal minimum wages and controls for possible endogeneity bias in estimates of the effect of the minimum wage on teen employment rates.

Estimates of the employment effects of minimum wage increases are likely to be biased if politicians time changes in the minimum wage to minimize the potential disemployment effect. If politicians accept the traditional view that minimum wage hikes reduce employment, they may be reluctant to increase the wage floor when unemployment, particularly among teens, is high. Legislatures may instead increase the minimum wage when unemployment is low and wages are rising. If minimum wages depend on economic conditions, the estimated effect of minimum wage increases on employment is likely to be biased upwards. Studies that examine exogenous increases in minimum wages, such as those that examine the impact of federal minimum wage hikes at the state level, avoid this bias.[4] However, time-series studies that use aggregate federal data may be underestimating the true disemployment effect of minimum wage increases if the level of the federal minimum wage depends on economic conditions. Similarly, panel data studies that exploit the variation in state minimum wages may be biased if the variation in minimum wages across states depends on economic conditions within states.

Studies that examine the effects of state minimum wage increases during the 1980s are particularly prone to this bias since many of the states that raised their minimum wages above the federal level--most notably California and the New England states--were experiencing relatively strong economic growth at the time. For example, Card (1992a)

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[4]If members of Congress focus on the employment effects in large states, such as California and New York, when deciding whether to raise the federal minimum wage, studies of the effects of federal minimum wage increases in large states would also be prone to endogeneity bias.

finds that the 90-cent increase in the minimum wage in California in 1988 had no negative effect on the teen employment rate in the state relative to five other regions.[5] Similarly, Card and Krueger (1994) report that the 80-cent hike in the minimum wage in New Jersey in 1992 did not reduce employment in fast food restaurants in New Jersey relative to neighboring Pennsylvania.[6] Neumark and Wascher (1994) find that increases in state and federal minimum wages reduced teen employment rates in state-level panel data over 1973-1989; however, a reanalysis of the data by Card, Katz, and Krueger (1994) concludes that the minimum wage had at most a minimal disemployment effect. If politicians imposed these minimum wage increases because the employment effects would be small, these studies may be underestimating the disemployment effects because they do not control for the endogeneity of minimum wage rates.

This paper examines the economic and political determinants of minimum wages at both the federal and state levels and then controls for possible endogeneity bias in estimates of the employment effect of the minimum wage. At the federal level, I focus on the influence of economic and political conditions on the timing of increases in the minimum wage and on the level of the real federal minimum wage; at the state level, I examine the determinants of states' decision whether to impose a minimum wage above the federal minimum wage and of the level of the real effective minimum wage. This approach follows

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[5]Kim and Taylor (1995) report opposite results, however. Using data from the County Business Patterns, they find that the increase in California's minimum wage did adversely affect employment. Kim and Taylor find that employment in retail trade grew relatively more slowly in industries in which the wage increase was the most rapid and in counties that were the most affected by the minimum wage increase. Card, in contrast, uses data from the Current Population Survey and focuses on teen workers.

[6]Card and Krueger's study has sparked a controversy over their methodology and results. See, for example, Hammermesh (1995) and Neumark and Wascher (1995).

the suggestion of Besley and Case (1994) to first investigate the determinants of policy before using a policy outcome as an independent variable. Two methods are then used to control for possible endogeneity bias in time-series and panel data models: inclusion of lagged economic conditions as right-hand-side control variables and instrumental variables estimation. This analysis is similar to Besley and Case (1994), who model the political and economic determinants of workers' compensation policies at the state level and then discuss several means of correcting for bias.

The paper is organized as follows. Section 2 replicates previous time-series and panel data estimates of the relationship between the minimum wage and employment rates and then examines the determinants of federal and state minimum wage policy. The findings serve to motivate the two methods then used to control for potential endogeneity. Section 3 discusses the results of controlling for lagged economic conditions to correct for endogeneity bias, while Section 4 shows the results of using potentially exogenous components of the minimum wage determination process to identify the true effect of the minimum wage on employment. I find that both economic and political variables appear to influence minimum wages. The results of the IV estimates and the estimates that control for lagged economic conditions do not indicate the presence of substantial bias in traditional estimates of the employment effect of the minimum wage.

2. Previous Research on Employment Effects of Minimum Wages

Studies of the effect of the minimum wage on employment focus on teens since a relatively large proportion of young workers earn low wages. Until recently, the empirical

literature relied on a time-series methodology to estimate employment effects because the federal minimum wage exceeded the state minimum wage in the vast majority of states.[7] As the real federal minimum wage declined by one-third during the 1980s, an unprecedented number of states passed minimum wages exceeding the federal rate. By 1990, sixteen states and the District of Columbia had minimum wages above the federal rate.[8] Several studies have used a panel data methodology to exploit the cross-state variation in minimum wages over time, most notably Neumark and Wascher (1992 and 1994). Typical time-series and panel data studies are replicated below to provide a basis for comparison of estimates after correcting for possible endogeneity bias.

2.1 Time-Series Estimates

The majority of time-series studies that use data from the 1960s and early 1970s find that a 10 percent increase in the minimum wage results in a 1 to 3 percent decrease in teen employment (Brown, Gilroy, and Kohen, 1982). In a time-series study incorporating data from 1954-1986, Wellington (1991) finds that a 10 percent increase in the minimum wage results in a less than 1 percent decrease in the employment of teens aged 16-19. She concludes that previous studies may overestimate the disemployment effect of the minimum wage. However, these studies may actually be *underestimating* employment losses due to

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[7]Cullen (1961) suggests that most state legislatures viewed the federal minimum wage as a ceiling for state minimum wages.

[8]The states that had minimum wages above $3.35 before April 1990 are AK, CA, CT, HI, IA, ME, MA, MN, NH, ND, OR, PA, RI, VT, WA, and WI. In 1989, about 25 percent of all U.S. workers were covered by state minimum wages that exceeded the federal minimum (Card and Krueger, 1995). The percentage of nonsupervisory workers in private industry covered under the federal minimum wage increased from 43 percent in 1938 to over 86 percent in 1991 (*Statistical Abstract*, 1994). See Schiller (1994) for a discussion of some of the complexities of coverage under state minimum wage laws.

increases in the minimum wage because they do not address the possibility of endogeneity bias.

The typical time-series model estimates the employment effect of the minimum wage by regressing the employment-to-population ratio for teens on a measure of the minimum wage and a set of business-cycle controls. The results are interpreted as a labor demand equation since the minimum wage is hypothesized to affect employers' labor costs but to not affect labor supply. The Kaitz index, a coverage-weighted ratio of the minimum wage to the average wage, is often used to gauge how binding the minimum wage is. As Card, Katz, and Krueger (1994) discuss, use of the Kaitz index may lead to overestimates of the disemployment effect of the minimum wage. The relative minimum wage--the federal minimum wage divided by the average manufacturing wage--or the real minimum wage is used in this analysis instead.

The time-series model estimated here is

$$y_t = \alpha + \beta X_t + \gamma MW_t + \epsilon_t \tag{1}$$

where y_t is the log of the employment-to-population ratio for teens aged 16-19 at time t. The minimum wage variable MW_t is the log of either the real minimum wage or the relative minimum wage. The relative minimum wage is instrumented using the real minimum wage since the average manufacturing wage and teen employment may be simultaneously determined. The teen-to-total population ratio and the unemployment rate of adult males are included in X_t to control for a "baby boom" effect and for business-cycle conditions that

affect labor demand. The equation is estimated using quarterly data from 1954-1993.[9] As in Wellington (1991), the vector X_t also includes quarter dummy variables, a linear time trend, the time trend squared, and interactions of the quarter dummy variables with the time trend and the squared time trend. The error term is corrected for serial correlation using the Cochrane-Orcutt iterative procedure.

The teen employment rate is negatively correlated with the relative and the real minimum wage, as shown in the first two columns of Table 1. The estimates imply that a 10 percent increase in the real or the relative minimum wage reduces teen employment rates by about 1.5 percent, and the estimated coefficients are statistically significant at the 95 percent confidence level.[10]

2.2 Panel Data Estimates

Panel data models are particularly appropriate for examining the employment effects of the minimum wage during recent years because of the cross-state variation in minimum wages during the 1980s. The panel data model estimated here is similar to the time-series model except that the data are state-level observations:

$$y_{it} = \alpha + \beta X_{it} + \gamma MW_{it} + \epsilon_{it} \qquad (1')$$

where i indexes states and t indexes years. Annual data for the 50 states over 1979-1989 are

[9]The employment, unemployment, and average wage data are from the Bureau of Labor Statistics publication *Employment and Earnings*, and the population data are from the *Statistical Abstract*. Data from the Bureau of Labor Statistics publications *Employment, Hours, and Earnings: United States, 1909-94* and *Labor Force Statistics Derived from the Current Population Survey, 1948-1987*, also were used.

[10]These estimates differ from those in Wellington (1991) because the years 1987-1993 are included here. These years include two increases in the federal minimum wage as well as many increases in state minimum wages. Including data from the late 1980s and the early 1990s slightly lowers the estimated coefficients and reduces the standard errors.

used to estimate the equation. The employment rates in the variable y_{it} are from the Bureau of Labor Statistics publication *Geographic Profile of Employment and Unemployment*.[11] The log of either the real or the relative effective minimum wage (the higher of the state and federal minimum wages in effect on May 1 divided by the average adult wage in the state) is used as the minimum wage variable, and the relative minimum wage is again instrumented using the real minimum wage. The vector X_{it} includes the annual average unemployment rate, the log of the real annual average manufacturing wage in a state, and a full set of state and year dummy variables. The regressions are estimated using the Prais-Winston method to correct for serial correlation and heteroscedasticity.[12]

The real and the relative minimum wage are negatively correlated with state-level teen employment rates, as shown in Columns 3 and 4 of Table 1. The estimated employment elasticities are again near -0.15, but both coefficients are only marginally statistically significant (p-value=0.101).

Instead of concluding from the results that the minimum wage has at most a small effect on teen employment, the possibility that the minimum wage variable is endogenous and the coefficients are biased upwards should be investigated. I next examine whether minimum wages appear to depend on economic conditions and whether there is an exogenous source of

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[11]The typical panel data study uses employment rates and other data from the May CPS. Card, Katz, and Krueger (1994) argue that estimates based on data from the CPS may be problematic since the small sample size leads to imprecise estimates in smaller states. The *Geographic Profiles* data is based on the full CPS sample instead of the one-quarter sample in the outgoing rotation group in the May CPS. *Geographic Profile* data on annual averages of total population, teen population, and teen employment are used here. Data from *Employment and Earnings* on the annual average unemployment rate and annual average manufacturing wage for each state are used instead of estimates of the adult male unemployment rate and average adult wage from the CPS.

[12]The Prais-Winston method used here corrects for autocorrelation within cross-sectional units. An AR(1) process is assumed, and the serial correlation parameter is constrained to be the same for all states. The panel data used here is not long enough to allow for accurate estimation of a separate serial correlation parameter for each state.

variation that could serve as an instrumental variable for the minimum wage.

2.3 Political Economy of Minimum Wage Determination

Although there is a large literature on the effects of minimum wage increases, little work has been done on the determinants of minimum wages at either the federal or the state level. In research on the determinants of state minimum wages, Cox and Oaxaca (1982) find that state minimum wages in 1970 and 1975 are positively correlated with the potential political power of organized labor and negatively correlated with the potential political base of organized capital.[13] There are also several studies that find that Congressional legislators' votes on minimum wage increases are correlated with economic conditions.[14] Krehbiel and Rivers (1988) and Poole and Rosenthal (1991), however, suggest that ideological and partisan interests, not economic conditions, are the primary determinants of legislative support for minimum wage increases. Political variables are potential instruments for the minimum wage if they are good predictors of minimum wages since they are unlikely to affect teen employment levels except through the minimum wage.

Several aspects of the minimum wage and its relationship with economic and political variables are of interest. First, do economic conditions and partisan control affect whether the federal minimum wage is increased in a given year? Second, do economic and political variables affect whether a state opts to exceed the federal minimum wage? And third, what

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[13]Blais, Cousineau, and McRoberts (1989) apply a similar model to minimum wages in the Canadian provinces. They find that the relative minimum wage (the minimum wage divided by the average wage in a province) tends to be lower the higher the average wage, the higher the unemployment rate, and the more conservative the provincial government.

[14]See, for example, Bloch (1993, 1980), Colberg (1981), Kau and Rubin (1978), and Silberman and Durden (1976).

role do economic and political variables play in the determination of minimum wage levels? These questions are addressed through a general empirical framework by regressing a minimum wage outcome on a set of economic and political variables:

$$MW_t = \pi + \delta X_{t-1} + \rho W_{t-1} + \sigma Z_{t-1} + \eta_t \qquad (2)$$

where MW_t is a minimum wage outcome at time t. The vector X_{t-1} includes annual lags of the two variables in X_t in the employment equations: the teen-to-total population ratio and the overall unemployment rate. The vector W_{t-1} includes two other lagged economic variables: the log of the real average hourly manufacturing wage and the unemployment rate of teens. The vector Z_{t-i} is political variables that are potential instruments for the minimum wage. It includes variables representing Democratic control of each of the legislative and the executive branches.[15] The independent variables are lagged in most of the specifications for two reasons. First, this reflects the information on which politicians are likely to be basing their votes. Also, lagging avoids potential feedback effects of the minimum wage on the economic and political variables. The error term is assumed to be serially uncorrelated and orthogonal to the regressors.

The question of what affects the timing of increases in the federal minimum wage is examined through maximum-likelihood Probit estimation of Equation 2, where the dependent variable is 1 if the minimum wage was increased during a given year and zero otherwise. The equation is estimated for the years 1954-1993, the same years used in the time-series

[15]The three variables are the percentage of seats held by Democrats in both houses of the legislature and a dummy variable equal to one if the executive is a Democrat and zero otherwise.

employment model.[16] Six pieces of legislation to raise the minimum wage were passed during this time period. As shown in the first column of Table 2, none of the variables appear to be correlated with timing of minimum wage increases. The coefficient on the lagged teen unemployment rate is negative and marginally statistically significant, but the marginal effect of an increase in the teen unemployment rate is small. The lack of a strong correlation between the timing of increases and the political variables seems surprising given that all six laws were passed when the Democratic party controlled both houses of Congress, but there was little variation in partisan control of Congress during this period.

The second model examines how political and economic variables affect whether states choose to exceed the federal minimum wage. A dummy variable indicating whether the state minimum wage is higher than the federal minimum wage on January 1 of a given year is regressed on state-level data for the variables described above for the 50 states during the period 1979-1993.[17] As shown in the second column of Table 2, economic variables appear to play a sizable role in states' decision to exceed the federal minimum wage. The lagged manufacturing wage is positively correlated with the dependent variable, while the lagged teen unemployment rate is negatively correlated with the probability a state exceeds the federal minimum wage. A 10 percent increase in the average wage raises the probability that a state will exceed the federal minimum wage by 2.5 percent, while a 10 percentage

[16]The data are from the U.S. Bureau of the Census publications *Statistical Abstract* and *Historical Statistics of the U.S.: Colonial Times to 1970*, and the Bureau of Labor Statistics publication *Labor Force Statistics Derived from the Current Population Survey 1948-1987*. The political variables are not lagged in this specification because whether the minimum wage is increased in a given year should depend on contemporaneous partisan control. The specification also includes a linear time trend.

[17]A full set of year dummy variables is also included. A full set of state dummy variables could not be included because many of the states do not exhibit variation in the dependent variable.

point increase in the teen unemployment rate lowers the probability by considerably less than 1 percent. The coefficients on the teen-to-total population ratio and total unemployment rate variables are also statistically significant; the positive sign of the coefficient for the total unemployment rate is likely to be due to the collinearity between the teen and total unemployment rates.

All three Democratic control variables are positively correlated with the probability a state exceeds the federal minimum wage, and the coefficient on the Democratic governor variable is statistically significant. These results are not surprising given that only one state with a Republican majority in both houses of the legislature and a Republican governor--New Hampshire--ever exceeded the federal minimum wage over 1979-1993. The estimated marginal effects are small, however; changing from a Republican to a Democratic governor increases a state's probability of exceeding the federal minimum wage by 4 percent.

The last two models examine the relationship between minimum wage levels and the economic and political variables. Equation 2 is estimated for the federal minimum wage over 1954-1993 and the effective minimum wage in the 50 states over 1979-1993; the dependent variable in Equation 2 is the log of the real federal or the real effective minimum wage.[18] Of the economic and demographic variables, only the teen-to-total population ratio is significantly-correlated with the real federal minimum wage. The real effective minimum wage is positively correlated with the average manufacturing wage, with an estimated elasticity of about 0.4. The effective minimum wage is negatively correlated with both the teen and the overall unemployment rate. The effective minimum wage is also positively

[18]A linear time trend is included in the regression for the federal minimum wage, and a full set of state and year dummy variables is included in the regression for the effective state-level minimum wage.

correlated with the teen-to-total population ratio.

Most of the political variables are highly correlated with the minimum wage levels. The percentage of seats in the senate held by Democrats and the indicator variable of a Democratic President or governor are positively correlated with both the federal and the effective minimum wage. Switching from a Republican to a Democratic President is correlated with almost a 6 percent increase in the federal minimum wage. Surprisingly, the percentage of Democratic seats in the U.S. House of Representatives is negatively correlated with the real federal minimum wage.

The results of the simple political economy models estimated here indicate that the minimum wage determination process is influenced by economic variables such as the average wage and the teen unemployment rate, particularly at the state level. This finding, combined with the small estimates of the disemployment effect of the minimum wage, gives reason to be concerned that legislators may increase minimum wages when any negative effects will be minimal. Minimum wages may not be an exogenous variable in employment equations. The political economy models also indicate the importance of controlling for lagged economic conditions, particularly the teen unemployment rate, when estimating the employment effect of the minimum wage and suggest that political variables may allow for identification of the true employment effect of the minimum wage.

3. Using Lagged Variables to Control for Endogeneity

Estimates of the employment effect of the minimum wage from the standard time-series and panel data methodologies may be biased for several reasons. First, the minimum

wage may be endogenous with respect to economic conditions, creating endogeneity bias. Second, failure to include variables that affect minimum wage policy and have an independent effect on employment leads to omitted variable bias. One way to control for these biases is to include any variables that affect minimum wages and employment on the right-hand-side of the employment regression, or

$$y_t = \alpha + \beta_1 X_t + \beta_2 X_{t-1} + \theta W_{t-1} + \gamma MW_t + \epsilon_t \tag{3}$$

where y_t is the log of the employment-to-population ratio for teens, X_t and X_{t-1} include the contemporaneous and annual lags of the teen-to-total population ratio and the overall unemployment rate, and W_{t-1} includes annual lags of the teen unemployment rate and the lagged real manufacturing wage. The minimum wage variable MW_t is again the log of either the real minimum wage or the relative minimum wage, and the relative minimum wage is instrumented using the real minimum wage. Equation 3 is estimated with quarterly time-series data for 1954-1993 and with annual panel data for 1979-1993.[19]

Including the lagged variables has little effect on the time-series estimates, as shown in the first two columns of Table 3. The estimated coefficients on the relative and the real minimum wage variables are not statistically different from the estimates without including the lagged variables. The estimated employment elasticities are again -0.15. The lagged variables are also not jointly significantly different from zero at conventional levels in either specification.

[19]The time-series regressions also include three quarter dummy variables, a linear time trend, the time trend squared, and quarterly dummy variables interacted with the time trend variables. The time-series regressions are estimated using the Cochrane-Orcutt iterative procedure to correct for AR(1) serial correlation. The panel data regressions include a full set of state and year dummy variables and are estimated using the Prais-Winston transformation method.

The estimated employment effects of the minimum wage in the panel data are slightly more negative when the lagged variables are included in the regressions. The estimated elasticity of employment with respect to the relative minimum wage changes from -0.189 to -0.202 and the standard error falls when the lagged variables are included. The estimated coefficient on the real minimum wage falls from -0.149 to -0.190 when the lagged variables are included. The four lagged variables are jointly significantly different from zero above the 95 percent confidence level in both specifications.

Including lagged variables to control for endogeneity or omitted variable bias thus slightly lowers the estimated coefficients, particularly in the panel data, but does not result in significantly different estimates of the disemployment effects of the minimum wage.

4. Using Instrumental Variables to Control for Endogeneity

The results of the political economy models suggest that minimum wages depend on the teen unemployment rate and other economic conditions and on the partisan composition of the government. Since the teen employment rate is also likely to be influenced by economic conditions, it is difficult to interpret the coefficient on a minimum wage variable without also controlling for these conditions. In addition to controlling for contemporaneous and lagged economic conditions, it is desirable to instrument for the minimum wage with variables that influence teen employment only through the minimum wage. Instrumental variables estimation controls for bias from two sources: endogeneity of the minimum wage, and unobservable variables that affect both the minimum wage and teen employment.

The political economy models indicate that minimum wages are strongly positively

correlated with the Democratic composition of the senate and the executive branch. These variables are unlikely to affect teen employment except through the minimum wage and are therefore potential instrumental variables for the real and the relative minimum wage. The system estimated is

$$y_t = \alpha + \beta_1 X_t + \beta_2 X_{t-1} + \theta W_{t-1} + \gamma MW_t + \epsilon_t$$

$$MW_t = \pi + \delta X_{t-1} + \rho W_{t-1} + \sigma Z_{t-1} + \eta_t$$

where y_t, X_t, X_{t-1}, W_{t-1}, and MW_t are as defined above. The vector of instruments, Z_{t-1}, includes the proportion of Democrats in the Senate and an indicator variable of whether the President is a Democrat for the time-series specifications and an indicator variable of whether the governor is a Democrat for the panel data specifications.[20] The system is again estimated with time-series data over the period 1954-1993 and with panel data over the period 1979-1993. The second-stage estimates are corrected for serial correlation.

The estimated coefficients on the minimum wage variables are not lower than the OLS estimates. As shown in the first two columns of Table 4, the estimated coefficients for the federal minimum wage are slightly more positive but not significantly different from the OLS estimates. Although the two instrumental variables are strongly correlated with the minimum wage in the first-stage regressions, the standard errors on the minimum wage variables increase considerably when 2SLS is used.[21] Thus, there does not appear to be any

[20]The percentage of Democrats in the House is not included in the time-series regressions because it is negatively correlated with the minimum wage. Neither of the legislative variables is included in the panel data regressions because including them as instruments gives positive, usually statistically significant estimates of the effect of the minimum wage on the teen employment rate. Since there is no theoretical reason why correcting for endogeneity should make the coefficients positive, those results are not included here.

[21]The p-value of the F-test that the coefficients on the two political variables are jointly equal to zero is 0.0000 in both first-stage regressions. The variables also "pass" the over-identification test of the orthogonality of residuals and instruments in F-tests of whether the coefficients on the political variables are jointly equal to zero in a regression of the fitted residuals on the variables included in the second-stage regression. The p-values of the F-

significant amount of endogeneity bias that 2SLS with these political instruments is able to correct.

The estimated effect of the effective real or relative minimum wage is imprecisely estimated in the panel data specifications because the first-stage regressions are weak. As shown in the third and fourth columns of Table 4, the estimated elasticities are not significantly different from zero. The hypothesis that the IV estimates are the same as the OLS estimates certainly cannot be rejected, nor the hypothesis that the minimum wage has no effect on teen employment rates. A strong instrument for the minimum wage at the state level is clearly needed to better ascertain the effect of the minimum wage on teen employment.

5. Conclusion

Several recent studies of the effect of the minimum wage on employment suggest that the traditional view that a 10 percent increase in the minimum wage reduces employment by 1 to 3 percent may be incorrect. However, these studies ignore the political process in which minimum wages are determined and are prone to endogeneity bias if politicians increase the minimum wage when the disemployment effect will be minimal. Studies that rely on the "natural experiments" created when a state opts to raise its minimum wage and compare outcomes in that state to an unaffected control state are particularly prone to this endogeneity bias since states that choose to raise the minimum wage are inherently different from states that chose to remain covered by the federal minimum wage. The political

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tests are 0.8510 and 0.9198, respectively, for the F-tests in the relative and the real minimum wage specifications.

economy models developed here indicate that state and federal minimum wage levels depend on economic conditions and are endogenous.

However, I find that controlling for the endogeneity bias does not have a significant effect on estimates of the disemployment effects of the minimum wage. This paper examines the effects of including possible omitted variables--lagged economic conditions--and using political variables as instruments to identify the effect of the minimum wage on employment. The results of both techniques fail to demonstrate the presence of bias in traditional estimates of the effect of the minimum wage on teen employment rates.

Table 1

Estimates of the Employment Effect of the Minimum Wage

	Time-Series Data, 1954-1993		Panel Data, 1979-1993	
Covariate	(1)	(2)	(3)	(4)
Log of Relative	-.153	--	-.189	--
Minimum Wage	(.075)		(.115)	
Log of Real	--	-.145	--	-.149
Minimum Wage		(.071)		(.091)
Unemployment Rate	-.025	-.026	-.038	-.037
	(.005)	(.005)	(.002)	(.002)
Teen-to-Total	-.021	-.021	.008	.006
Population Ratio	(.026)	(.026)	(.004)	(.003)

Notes: The regressions for Columns 1 and 2 also contain 3 quarter dummy variables, a time trend and squared time trend, and quarterly dummy variables interacted with the time trend variables. The data are quarterly observations over 1954-1993 for a total of 160 observations. The regressions are estimated using the Cochrane-Orcutt iterative process to correct for serial correlation. The regressions for Columns 3 and 4 contain a full set

of state and year dummy variables. The data are annual observations for the 50 states over 1979-1993 for a total of 750 observations. The regressions are estimated using the Prais-Winston method to correct for serial correlation and heteroscedasticity. The relative minimum wage is instrumented using the real minimum wage in all regressions that contain the relative minimum wage, and the first-stage regressions include all other independent variables from the second stage. Standard errors are shown in parentheses.

Table 2
Determinants of Minimum Wages

Covariate	Probability Federal MW Increased (1)	Probability State Exceeds Federal MW (2)	Real Federal MW (3)	Real Effective MW (4)
Teen-to-Total Population Ratio	.539 (.483) {.109}	-.356 (.080) {.038}	.056 (.024)	4.102 (.241)
Total Unemployment Rate	.585 (.544) {.119}	.182 (.076) {.020}	-.007 (.021)	-.005 (.002)
Log of Real Manufacturing Wage	-8.798 (10.621) {1.782)	2.303 (.672) {.247}	.009 (.517)	.384 (.040)
Teen Unemployment Rate	-.424 (.292) {.086}	-.188 (.030) {.020}	.010 (.013)	-.0011 (.0006)
Percentage Democratic Seats in the Senate	.013 (.058) {.003}	.008 (.007) {.001}	.005 (.002)	.0009 (.0002)
Percentage Democratic Seats in the House	.009 (.020) {.002}	.007 (.008) {.001}	-.005 (.002)	.0003 (.0003)
Democratic Executive	.537 (.529) {.121}	.405 (.148) {.041}	.056 (.021)	.015 (.004)
Log Likelihood	-15.206	-194.815	--	--
Adj. R^2	--	--	.777	.881

Notes: The dependent variable in Column 1 is one if legislation to increase the federal minimum wage was approved that year and zero otherwise. The dependent variable in Column 2 is one if a state exceeds the federal minimum wage on January 1 of a given year and zero otherwise. The dependent variable in Columns 3 and 4 are the log of the real federal or the real effective minimum wage. Most of the independent variables are lagged (see text for details). Standard errors are shown in parentheses and marginal effects in brackets. The number of observations is 40 in Columns 1 and 3 and 750 in Columns 2 and 4.

Table 3
Estimates of the Employment Effect of the Minimum Wage Controlling for Lagged Economic Conditions

Covariate	Time-Series Data, 1954-1993 (1)	(2)	Panel Data, 1979-1993 (3)	(4)
Log of Relative Minimum Wage	-.154 (.074)	--	-.202 (.095)	--
Log of Real Minimum Wage	--	-.147 (.071)	--	-.190 (.089)
Unemployment Rate	-.023 (.005)	-.024 (.005)	-.033 (.002)	-.033 (.002)
Lagged Unemployment Rate	-.001 (.005)	.0001 (.005)	-.010 (.003)	-.010 (.003)
Teen-to-Total Population Ratio	-.042 (.028)	-.042 (.028)	.008 (.004)	.008 (.004)
Lagged Teen-to-Total Population Ratio	.034 (.030)	.036 (.030)	-.004 (.004)	-.004 (.004)
Lagged Teen Unemployment Rate	-.002 (.002)	-.002 (.002)	.001 (.001)	.001 (.001)
Log of Lagged Real Manufacturing Wage	.436 (.297)	.449 (.296)	-.105 (.104)	.085 (.053)

Notes: The regressions for Columns 1 and 2 also contain 3 quarter dummy variables, a time trend and squared time trend, and quarterly dummy variables interacted with the time trend variables. The data are quarterly observations over 1954-1993 for a total of 160 observations. The regressions are estimated using the Cochrane-Orcutt iterative process to correct for serial correlation. The regressions for Columns 3 and 4 contain a full set of state and year dummy variables. The data are annual observations for the 50 states over 1979-1993 for a total of 750 observations. The regressions are estimated using the Prais-Winston method to correct for serial correlation and heteroscedasticity. The relative minimum wage is instrumented using the real minimum wage in all regressions that contain the relative minimum wage, and the first-stage regressions include all other independent variables from the second stage. Standard errors are shown in parentheses.

Table 4
Instrumental Variables Estimates of the Employment Effect of the Minimum Wage

Covariate	Time-Series Data, 1954-1993 (1)	(2)	Panel Data, 1979-1993 (3)	(4)
Log of Relative Minimum Wage	-.117 (.148)	--	-1.328 (.977)	--
Log of Real Minimum Wage	--	-.132 (.167)	--	-1.799 (1.325)
Unemployment Rate	-.023 (.005)	-.023 (.005)	-.030 (.002)	-.031 (.002)
Lagged Unemployment Rate	-.004 (.007)	-.004 (.007)	-.011 (.003)	-.014 (.004)
Teen-to-Total Population Ratio	-.042 (.029)	-.041 (.029)	.006 (.004)	.002 (.006)
Lagged Teen-to-Total Population Ratio	.033 (.031)	.035 (.031)	-.003 (.004)	-.002 (.004)
Lagged Teen Unemployment Rate	-.001 (.002)	-.0002 (.003)	.001 (.001)	-.0000 (.001)
Log of Lagged Real Manufacturing Wage	.421 (.316)	.466 (.302)	-1.126 (.891)	.151 (.073)

Notes: The regressions for Columns 1 and 2 also contain 3 quarter dummy variables, a time trend and squared time trend, and quarterly dummy variables interacted with the time trend variables. The data are quarterly observations over 1954-1993 for a total of 160 observations. The relative and the real minimum wage are instrumented with the percentage of Democrats in the Senate and an indicator variable of a Democratic President. The regressions are estimated using the Cochrane-Orcutt iterative process to correct for serial correlation. The regressions for Columns 3 and 4 contain a full set of state and year dummy variables. The data are annual observations for the 50 states over 1979-1993 for a total of 750 observations. The relative and the real minimum wage are instrumented with an indicator variable of a Democratic governor. The regressions are estimated using the Prais-Winston method to correct for serial correlation and heteroscedasticity. Standard errors are shown in parentheses.

Appendix Table 1
Federal Minimum Wage Chronology

Date of Legislation	Date of Increase	Nominal Minimum Wage[21]	Percent of Employees in Private Industry Covered
1938	October 1938	0.25	43.4
	October 1939	0.30	47.1
	October 1945	0.40	55.4
1949	January 1950	0.75	53.4
1955	March 1956	1.00	53.1
1961	September 1961	1.15	62.1
	September 1963	1.25	62.1
1966	February 1967	1.40	75.3
	February 1968	1.60	72.6
1974	May 1974	2.00	83.7
	January 1975	2.10	83.3
	January 1976	2.30	83.8
1977	January 1978	2.65	85.1
	January 1979	2.90	85.1
	January 1980	3.10	86.1
	January 1981	3.35	86.0
1989	April 1990	3.80	86.8
	April 1991	4.25	86.5

Source: U.S. Bureau of the Census, *Statistical Abstract*, various years.

Appendix Table 2
States Exceeding the Federal Minimum Wage, January 1979-1993

Year	States
1979-1984	AK, CT
1985	AK, CT, ME
1986	AK, CT, ME
1987	AK, CT, ME, MA, NH, RI, VT
1988	AK, CT, HI, ME, MN, MA, NH, RI, VT
1989	AK, CA, CT, HI, ME, MA, MN, NH, RI, VT, WA
1990	AK, CA, CT, HI, IA, ME, MA, MN, NH, ND, OR, PA, RI, VT, WA, WI
1991	AK, CA, CT, HI, IA, ME, MA, MN, NH, OR, RI, VT, WA
1992	AK, CT, IA, OR, RI
1993	AK, CT, HI, IA, NJ, OR, RI

Source: R. Nelson, *Monthly Labor Review*, various years, and Quester (1981).

[21]Reflects the highest federal minimum wage in effect. Lower rates often applied to workers newly covered by the Federal Labor Standards Act or to teen workers.

Chapter 3: Effects of the Minimum Wage in Low-Wage Industries in Washington

1. Introduction

The creation of a federal minimum wage in 1938 and subsequent expansion of coverage made state minimum wage rates irrelevant for most workers until the 1980s. As the federal minimum wage remained unchanged from 1981 until 1990, an unprecedented number of states enacted state minimum wages above the federal wage floor. By 1990, sixteen states had state minimum wages that exceeded the federal minimum wage. These state-level minimum wage increases offer a series of "natural experiments" for examining the effect of the minimum wage on employment. Several studies have examined the effect of state minimum wage increases in California and New Jersey, and this paper adds to the literature a discussion of the effects of a 90-cent increase in Washington's state minimum wage in 1989 and 1990 on wages and employment. Whereas most previous studies have concentrated on the effects of minimum wage increases on teens and fast-food restaurants, I examine the impact of Washington's minimum wage increase on low-wage retail trade and service sectors.

Previous studies of state minimum wage increases have focused on the employment effects of a 90-cent increase in California's minimum wage in 1988 because of the state's large population and the size of the minimum wage change. Using individual-level data from the Current Population Survey (CPS), Card (1992a) reports that the increase in California's minimum wage raised teen wages but did not lower teen employment rates. Kim and Taylor (1995), however, find that the California minimum wage increase had a negative effect on

employment in retail trade when establishment data from the County Business Patterns are used to examine the effect of the increase. Kim and Taylor find that wages grew more quickly and employment more slowly in California compared to the rest of the United States. In addition, counties that experienced relatively larger wage increases had larger disemployment effects. They conclude that there is a strong negative correlation between relative wage changes and employment growth. Card and Krueger (1995) re-examine the data used by Kim and Taylor and report several flaws with the methodology; after correcting these errors, they find small and statistically insignificant employment effects of California's minimum wage increase. Card and Krueger (1994) also find from establishment survey data that employment in fast-food restaurants did not fall in New Jersey relative to Pennsylvania after an 80-cent increase in New Jersey's minimum wage in 1992.[1]

Empirical studies of the employment effects of federal minimum wage increases also yield contradictory results. The older studies surveyed in Brown, Gilroy, and Kohen (1982) commonly found small negative employment effects for teens. A 10 percent increase in the minimum wage appeared to reduce teen employment by 1 to 3 percent. Using more recent data, Neumark and Wascher (1992) find that increases in the federal minimum wage reduced teen employment rates. Deere, Murphy, and Welch (1995) report adverse relative employment effects among low-wage groups such as teens, blacks, and high-school dropouts. However, a growing body of literature fails to find disemployment effects. Using a variety of data and methodologies, Wellington (1991), Card (1992b), Katz and Krueger (1992), and

[1]Several papers, including Hammermesh (1995) and Welch (1995), have criticized the data and methodology used by Card and Krueger. Using payroll data instead of the phone survey results used by Card and Krueger, Neumark and Wascher (1995) claim to find significant negative employment effects as a result of the increase in New Jersey's minimum wage.

Card, Katz, and Krueger (1994) find that increases in the federal wage floor did not have noticeable adverse employment effects.

The effects of Washington's minimum wage increase are of interest for several reasons. First, the increase was substantial, particularly for tipped employees. The effective minimum wage for most workers was increased from $3.35 to $3.85 in January 1989 and to $4.25 in January 1990. Unlike most minimum wage legislation, this minimum wage increase applied to most agricultural workers and did not include a tip credit.[2] The 39 counties in Washington exhibit considerable variation in average wages, allowing for a comparison of the effects of the increase in low-wage and high-wage counties. The unique nature of the ballot initiative that raised the minimum wage also allows for an examination of the political economy of voting behavior. In response to the state legislature's repeated rejection of bills that would have raised the state's wage floor, a coalition of labor, church, women's and community groups backed a ballot initiative in 1988 to raise the state wage floor.[3] The initiative was approved in November 1988 with about 76 percent of the vote.

Data from the County Business Patterns and the unemployment insurance program are used to examine the effects of the minimum wage increases in low-wage retail trade and service sectors. I first discuss the data and examine whether the two increases in Washington's minimum wage appear to have raised wages. I then examine the industry-level effect on employment in Washington relative to the rest of the United States. The results are imprecise but generally fail to indicate that the minimum wage increases had a negative

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[2]The maximum allowable credit tip credit for the federal minimum wage was increased from 40 percent of the minimum wage to 45 percent in April 1990 and to 50 percent in April 1991.

[3]Bureau of National Affairs, *Daily Labor Report*, 27 July 1988.

employment effect. I then discuss the county-level effects of the minimum wage increases. The last section of the paper develops a political economy model of the effects of political and economic variables on votes on the ballot initiative. The results are consistent with the view that economic interests affected voting behavior on the minimum wage initiative and offer a potential explanation for why the minimum wage increases appear to have not reduced employment opportunities.

2. Discussion of the Data

This study focuses on low-wage retail trade and service sectors, which are the primary employers of minimum-wage workers in the United States. According to the CPS, over 68 percent of workers paid the minimum wage in Washington in 1988 were employed in retail trade and services, and 24 percent of workers in those industries earned between $3.35 and $3.85. Twenty-two percent of workers in retail trade and services earned between $3.85 and $4.25 in 1989.[4] These workers--if they remained employed--should have experienced wage increases when the minimum wage rose.[5]

This study uses establishment data from the U.S. Bureau of the Census publication *County Business Patterns* (CBP). The CBP includes data on total first-quarter payroll, the number of employees in mid-March and the number of establishments at the county level.

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[4]These figures are based on the CPS Outgoing Rotation Group files. Only workers paid an hourly wage are included. Services does not include workers in personal, business, auto, miscellaneous repair, health, legal, or educational services.

[5]Unlike the federal minimum wage, Washington's minimum wage does not include an exemption for small businesses with sales below some cutoff level. The ballot initiative also gives the state Department of Labor and Industries the authority to set minimum wages for workers under age 18. The youth minimum wage for 16 and 17 year-olds was set equal to the adult minimum wage, while the wage floor for workers under age 16 was set at 85 percent of the state minimum wage (McIntire, 1991).

The data are generally available at the 3- or 4-digit SIC code level and are from a universe sample.[6] Eleven sectors at the 2-digit SIC code level are examined here, including eight retail trade industries (building materials and garden supplies, general merchandise stores, food stores, automotive dealers and service stations, apparel and accessory stores, furniture stores, eating/drinking places, and miscellaneous retail) and three service industries (hotels, motion pictures, and entertainment/recreation).

Data from the U.S. Bureau of Labor Statistics publication *Employment and Wages* are also used to evaluate the wage and employment changes in Washington relative to the rest of the US. This data source reports average annual employment, total annual pay, and the number of reporting units in each state at the 3- or 4-digit SIC code level. The data is based on quarterly unemployment insurance reports filed by employers with state employment agencies and is from a universe sample. The data only include workers who are covered by the unemployment insurance (UI) program and are likely to give a higher average annual wage per worker than the CBP.

The effective minimum wage for most workers in Washington increased from $3.35 to $3.85 on January 1, 1989, and to $4.25 on January 1, 1990. The CBP employment and payroll data are from the first quarter, so the CBP data for 1989 and 1990 should show the effects of Washington's minimum wage increases. The federal minimum wage increased from $3.35 to $3.80 on April 1, 1990, and any effects of the federal minimum wage increase should not be reflected in the CBP data until 1991. The UI data are annual, so the 1990 UI data reflect the effect of Washington's second minimum wage increase and the effect in other

[6]All data used here at the 2-digit SIC code level and were created by aggregating up from the 3- or 4-digit level.

states of the federal minimum wage increase. Since Washington's minimum wage increase occurred four months earlier, there should be a larger wage effect in Washington than in the rest of the US, but the difference may be small. The 1989 UI data may better indicate the effect of the minimum wage increase in Washington relative to the rest of the US than the 1990 data.

The Washington minimum wage increases should have raised wages in industries in which a significant fraction of workers were initially earning less than the new wage floor. As shown in Table 1, this is the case in most industries examined here. The first column gives the fraction of workers in Washington earning $3.35 to $3.85 an hour in the 1986-1988 CPS Outgoing Rotation Group files. Over one-quarter of workers at eating and drinking places and in the motion picture industry should have been affected by the increases, while apparently no workers were directly affected at building materials stores. The more workers earning below the new wage floor, the greater should be the change in payroll. Multiplying the percentage of workers affected by the minimum wage increase times the percentage change in the minimum wage gives an upper bound on the direct effect of the minimum wage changes on payroll.[7] Washington's minimum wage increased by 15 percent in 1989 and 12 percent in 1990.

The next four columns of Table 1 show the percentage changes in nominal payroll per employee over 1988-1989 and 1989-1990 in the UI and the CBP data. All of the industries experience nominal wage growth in average pay per worker in the UI data, but several show

[7]The estimates are an upper bound since many workers in the affected region are earning more than the previous wage floor but less than the new wage floor. There may also be indirect effects on the wages of workers earning more than the new wage floor (the ripple effect). In addition, changes in employment or hours may offset wage changes and cause payroll to increase by less than the percentage change in the minimum wage.

decreases in average pay per workers in the CBP data. The correlation between the wage changes in the two datasets is 0.181. Industries with a higher fraction of workers affected generally experience larger pay increases, particularly in the UI data. However, there are several exceptions and the relationships are not monotonic. If the one-year change in the payroll data is regressed on the fraction of workers affected in an industry, the estimated coefficients range from 0.086 to 0.263.

Although the wage changes in Washington appear to be related to the fraction of workers affected by the minimum wage increases, the correlation might be spurious and the wage changes might have occurred in the absence of the minimum wage increases. Comparing the changes in Washington to changes elsewhere allows for an assessment of whether the observed wage changes are due to the minimum wage increases. Table 2 shows the differences between the percentage changes in wages in Washington and the rest of the United States (the US less Washington) over 1988-1989 and 1989-1990. The "differences in differences" show that average pay per worker increased in Washington relative to the rest of the US in several industries. In the UI data, average pay per worker either increased or fell only slightly in Washington relative to the rest of the US. The differences in differences are more likely to be negative in the CBP data, indicating that average pay fell in Washington relative to the rest of the US. The differences in differences in the UI data over both periods are positively correlated with the percentage of workers affected by the minimum wage increases, while the differences in differences in the CBP are not highly correlated with the percentage of workers affected. The correlation between the changes in the UI and the CBP is 0.359.

The differences in differences estimates assume that the low-wage retail trade and service sectors would have grown at similar rates in Washington and the rest of the US in the absence of the increases in Washington's minimum wage. There are several reasons to be concerned this might not have been the case. Retail sales, median family income, and population grew more quickly over 1988-1990 in Washington than in the US as a whole.[8] However, employment in the low-wage retail trade and service industries was growing at similar rates before Washington's minimum wage increases. The correlation between the percentage changes in employment in Washington and the rest of the US is 0.623 over 1986-1987 and 0.896 over 1987-1988 in the CBP data. The percentage changes in average payroll per employee before Washington's minimum wage increases are not positively correlated in the CBP data. The correlation between the percentage changes in average pay in Washington and the rest of the US is -0.157 over 1986-1987 and -0.533 over 1987-1988.

The UI data appear to be a better measure of wage changes than the CBP data. The annual changes in average pay and the differences in differences in the UI data are better correlated with the fraction of workers affected by the minimum wage increases than in the CBP data. The UI data also uniformly indicate nominal wage increases in Washington while the CBP data show considerable variation at the 2-digit SIC code level. I use the CBP data in addition to the UI data for several reasons. The CBP data are available for more years than the UI data, which are only available for 1988-1990. This allows for the comparison of trends in Washington and the rest of the US before Washington's minimum wage increases.

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[8]Nominal retail sales increased by 21.9 percent in Washington and 10.3 percent in the U.S. over 1988-1990. Median nominal income for a 4-person family increased by 12.7 percent in Washington and 6.1 percent for the entire US over 1988-1990. Population increased by 4.9 percent in Washington and 1.7 in the US (U.S. Bureau of the Census, *Statistical Abstract of the United States*, various years).

The CBP data are readily available at the county level and can be used to investigate differences in the impact of the minimum wage across counties. The federal minimum wage increase in 1990 confounds the estimates for 1989-1990 in the annual UI data but not in the first-quarter CBP data. Kim and Taylor (1995) use CBP data for their estimates of the employment effect of California's minimum wage increase in 1988, and the results obtained using the CBP here are easier to compare to theirs than those using a different dataset.[9] In addition, the UI data and the CBP data give fairly similar results in the comparisons reported below.

3. Employment Effects of the Minimum Wage Increases

The neoclassical model of labor markets predicts that imposition of a binding wage floor should reduce the quantity of labor employed in a competitive market. The model assumes that each firm's demand for labor is decreasing in the wage and workers are willing to supply more labor as the wage increases. This creates a downward-sloping market labor demand curve and an upward-sloping labor supply curve. The intersection of the market demand and supply curves determines the equilibrium wage. Imposition of a wage floor above this wage reduces the quantity of labor demanded. The magnitude of the effect of a minimum wage increase can be determined from the slope of the labor demand curve. The slope of the labor demand curve can be estimated by regressing the change in employment on the change in the wage for a given industry:

[9]Card and Krueger (1995) use CPS data to evaluate Kim and Taylor's study and the quality of CBP data. The number of observations in the CPS for Washington is too small--about 1000 a year--to obtain reasonable estimates. The measurement error in annual wage changes in the CPS data for Washington appears to be substantial. For example, the estimated annual average wages changes are over 100 percent in some retail trade industries.

$$\Delta e_i = \alpha + \beta \Delta w_i + \epsilon_i \tag{1}$$

where Δe_i and Δw_i are the percentage changes in employment and wages in industry i. However, it is difficult to interpret equation (1) as a demand curve because shifts of either the supply or the demand curve can change employment. As Kim and Taylor (1995) explain, an exogenous variation in wage changes across industries allows us to interpret Equation 1 as a demand curve. A binding increase in the minimum wage in competitive markets is one such source of variation.

This paper uses a differences-in-differences estimation technique to control for industry-specific effects that are likely to affect employment growth in a given industry. As in Kim and Taylor (1995), the change in employment in Washington in industry i can be represented by

$$\Delta e_{wi} = \alpha_w + \alpha_i + \beta \Delta w_{wi} + \epsilon_{wi} \tag{2}$$

where α_i is an industry-specific effect and α_w is a state-specific effect. If the industry-level effect also occurs in a control group that has a similar demand function, then differencing will eliminate the industry-specific effect:

$$(\Delta e_w - \Delta e_c)_i = \alpha_w - \alpha_c + \beta(\Delta w_w - \Delta w_c)_i + \epsilon_{wi} - \epsilon_{ci} \tag{3}$$

where w denotes Washington and c the control group. If the employment and wage variables are in logs, then Δe and Δw are percentage changes and β can be interpreted as the elasticity of demand. The sign on estimates of β is expected to be negative since increases in wages should lower the demand for labor.

Equation 3 is estimated separately with the CBP and the UI data on employment and average pay per worker at the 2-digit SIC code level using weighted ordinary least squares

(OLS). The weights are lagged employment in Washington. Washington's minimum wage increases occurred in 1989 and 1990, so the equation is estimated in first differences for 1988-1989 and 1989-1990. The analysis cannot be extended to include an examination of any long-term effects of the minimum wage increases using 1991 data because the federal minimum wage was increased during that time period.

Table 3 shows the differences in differences in employment and the OLS estimation results when the differences in employment changes are regressed on the differences in wage changes. It is readily apparent that the differences in differences in employment are not monotonically declining as the percentage of workers affected by Washington's minimum wage increase rises. The regression results also do not indicate a clear negative relationship between wage changes and employment changes. The regression results are imprecisely estimated since the sample size is only eleven observations. The UI data suggest a negative correlation between wages and employment over 1988-1989; the estimated coefficient is negative and statistically different from zero at the 90 percent confidence level. The UI data do not indicate a negative relationship between wages and employment over 1989-1990, but the control group also experienced an increase in the federal minimum wage during this period. The CBP data are not confounded by changes in the federal minimum wage, but they do not indicate a statistically-significant negative correlation between wages and employment.

The increases in Washington's minimum wage should generate the exogenous variation in wages necessary to identify the slope of the demand curve, but there are two reasons to be concerned about endogeneity bias. First, there may be sources of exogenous

variation in wages due to supply-side factors that would make it difficult to interpret estimates of β as a demand elasticity. Second, the wage measures used here are constructed by dividing total payroll by total employment, automatically generating negative correlation between the independent and the dependent variables. Equation 3 therefore also should be estimated using instrumental variables (IV) estimation. At least one variable correlated with the difference in the change in wages but uncorrelated with the error term in Equation 3 is needed. The change in Washington's minimum wage seems like the natural instrument for the difference in the wage changes, but we lack the necessary variation across industries in Washington's minimum wage.

The instrument used for the differences in differences in wages between years t and t-1 is the fraction of workers affected by the minimum wage increase.[10] This variable is created by estimating the percentage of workers in Washington paid an hourly wage between \$3.35 and \$3.85 over 1986-1988 in the CPS. The first-stage regression is

$$\Delta w_{wit,\,t-1} - \Delta w_{cit,\,t-1} = \delta + \gamma p_{wi} + \eta_{it} \tag{4}$$

where the wage variables are in logs. This instrument should be positively correlated with the difference in the change in wages. Industries with more workers affected by the minimum wage increases should experience a greater increase in wages relative to the rest of the US. The first- and second-stage regressions are weighted by lagged employment in Washington. The first-stage regression results are similar to those shown in the last row of

[10]Kim and Taylor (1995) use lagged real average payroll per worker and average establishment size in California as instruments. Lagged average pay should be negatively correlated with the difference in the change in wages, while the sign on firm size is theoretically ambiguous. Neither instrument is used here because both are very weak predictors of the differences in differences in wages.

Table 2.[11]

None of the estimated coefficients is statistically different from zero in the second-stage regressions. As shown in the last row of Table 3, the estimated coefficients are negative for the UI data over 1988-1989 and the CBP data over 1989-1990 but are imprecisely estimated.[12] The small sample size--eleven observations--again makes it difficult to precisely determine the effect of the minimum wage increases on employment in Washington.

The sample size can be increased by using more disaggregated data. The CBP and the UI data are both available at the 3- or 4-digit SIC code level. The CPS can also be used to estimate the percentage of workers earning between $3.35 and $3.85 at the 3- or 4-digit SIC code level. The CPS offers data on 37 sectors at the 3- or 4-digit SIC code level for the eleven 2-digit SIC code low-wage industries examined here. However, the 3- and 4-digit SIC code cells in the CPS are extremely small if only data on Washington is used. I therefore use the percentage of workers earning $3.35 to $3.85 in all states that did not exceed the federal minimum wage before 1989 as an instrument for the differences in differences in average pay per employee in the IV regressions.[13] The wage distributions in the entire US and Washington are similar; the correlation between the percentage of workers affected at the 2-digit level in Washington and the entire US is 0.829. Equation 3 is therefore estimated with data at the 3- or 4-digit SIC code levels, which increases the number

[11]The first-stage regression results differ slightly because a constant is included and the observations are weighted by lagged employment. The correlation between the differences in differences in wages and the fraction of workers affected is positive and statistically significant except in the CBP data for 1988-1989.

[12]If differences in differences in employment are regressed directly on the fraction of workers affected by the minimum wage increase, none of the estimated coefficients are statistically different from zero.

[13]The excluded states are AK, CA, CT, HI, ME, MA, MN, NH, RI, and VT.

of observations to 37.

The results are slightly more suggestive of a negative relationship between employment and wages than the results with the aggregate data but are not conclusive. As shown in the first row of Table 4, all of the estimated coefficients in the OLS regressions are negative, although none is statistically different from zero. The estimated coefficients in the first-stage regression results, shown in the second row, are positive and generally statistically significant. The second-stage results in the IV regressions are negative in the three of the four specifications, and the estimated coefficient for the UI data over 1988-1989 is statistically significant.[14]

The UI and CBP data indicate that the two increases in Washington's minimum wage raised wages in several low-wage industries relative to the rest of the US but may not have adversely affected employment opportunities in Washington. The differences in differences in employment between Washington and the rest of the US are not highly negatively correlated with the differences in differences in wages or with the fraction of workers affected by the minimum wage increase in most specifications. These results are dissimilar to those reported by Kim and Taylor (1995), who find employment elasticities of about -0.8 using a similar methodology and data for California.

4. Intercounty Effects of the Minimum Wage Increases

In addition to raising wages in Washington relative to the rest of the US, the

[14]If the differences in differences in employment are regressed directly on the fraction of workers affected in the entire US, only the coefficient for the UI data over 1988-1989 is statistically significant. It is -0.079 (standard error=0.034).

minimum wage increases should have had a relatively larger effect on wages in counties that initially paid lower wages--the rural counties in eastern Washington--than in higher-wage counties. This does not appear to be the case in the CBP data even though average wages vary widely across Washington. For example, average payroll per employee in retail trade ranged from $1414 in Garfield County, the smallest county in Washington, to $3244 in King County, which includes Seattle, in the first quarter of 1988. If the minimum wage increases raised wages more in counties that initially had lower wages, average wages at time t-2 should be negatively correlated with the change in wages between years t and t-1 if the minimum wage increased during that period. If average payroll per employee in 1987 or 1988 is regressed on the county-level change in wages over 1988-1989 or 1989-1990, respectively, the estimated coefficient on the lagged wage is not negative and statistically significant.[15] The minimum wage increases thus do not appear to have raised wages in low-wage counties relative to high-wage counties.

The minimum wage increases also should have reduced the dispersion in wages in low-wage industries across counties. However, the standard deviation of wages in low-wage retail trade and services in the CBP data across counties rose from 1988 to 1990. The standard deviation in average payroll per worker increased from 0.338 in 1988 to 0.440 in 1990.

Washington's minimum wage increases should have had a particularly large impact on the restaurant industry since the ballot initiative did not include a tip credit. The minimum

[15]The regressions use CBP data pooled across the eleven low-wage retail trade and service industries examined here for the 39 counties in Washington. The likelihood of measurement error precludes the use of county-level data at the 2-digit SIC code level; the annual changes in average wages are over 100 percent in many industries in the county-level CBP data.

wage for tipped employees rose from $2.30 to $3.85 on January 1, 1989. The effect on wages should have been particularly pronounced in low-wage counties. However, average payroll per employee at eating/drinking places in 1987 is positively, not negatively, 0correlated with the change in average wages over 1988-1989 in the CBP county-level data.[16] The result implies that counties with initially high wages experienced larger average wage increases when the minimum wage rose than counties with initially low wages.

Since the minimum wage increases did not raise average wages in low-wage counties relative to high-wage counties, I do not examine the potential disemployment effects of the minimum wage increases at the county level. There are several possible explanations for the failure to find relative wage effects in the CBP data. The CBP wage data is total payroll divided by total employment, not an average hourly wage. Employers in low-wage counties may have reduced employment or hours considerably more than employers in high-wage counties. Payroll therefore may not have increased relatively more in low-wage counties even though average hourly wages may have increased. Data on hours and hourly wages at the county-level are necessary to test this hypothesis.

5. Political Economy Analysis of Voting Behavior

The above analysis of the effects of the minimum wage assumes that the increase was exogenous to economic conditions, particularly the growth rate of low-wage employment. If the increase in the minimum wage was endogenous, then my estimates of the elasticity of labor demand may be biased. If Washington increased its minimum wage because the

[16]McIntire (1991) reports that establishments in western counties are more likely to employed tipped workers than establishments in eastern Washington.

economy was growing and an increase would not have a substantial disemployment effect, my estimates would understate the effect of an exogenous change in the minimum wage. There are several reasons to be concerned this might be the case. First, nominal retail sales grew by 7.9 percent in Washington over 1987-1988, compared with 6.3 percent for the entire U.S.[17] Washington's unemployment rate fell by 0.4 percentage points over 1987-1988, while the U.S. average increased by 0.3 percentage points. Similarly, Washington's teen unemployment rate fell by 4.6 percentage points and the U.S. average fell by 1.6 percentage points over 1987-1988.[18] In addition, fewer workers were earning low wages in Washington than in other states; only 9.9 percent of workers in Washington were earning between $3.35 and $3.85 an hour in Washington in 1987 and 1988, compared to 12.6 percent in the rest of the US.[19] The unique nature of the ballot initiative that raised Washington's minimum wage allows for a deeper investigation of whether support for the minimum wage increase depended on economic conditions.

Previous studies of the political economy of minimum wage legislation have usually found that votes on minimum wage increases are correlated with economic conditions. Studies of the effect of state-level economic conditions and political party on Congressional votes on minimum wage legislation typically find that support for higher minimum wages is positively associated with average wages.[20] Krehbiel and Rivers (1988) and Poole and Rosenthal (1991), however, argue that political variables better explain legislators' votes than

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[17]U.S. Bureau of the Census, *Statistical Abstract of the United States*, various years.

[18]U.S. Bureau of the Census, *Geographic Profile of Employment and Unemployment*.

[19]Figures are based on the CPS Outgoing Rotation Groups and only include workers paid an hourly wage. States that exceeded the federal minimum wage by 1989 are excluded from the comparison group.

[20]See, for example, Bloch (1993, 1980), Colberg (1981), Kau and Rubin (1978), and Silberman and Durden (1976).

do measures of constituent economic interests. In research on the determinants of state minimum wages, Cox and Oaxaca (1982) find that state minimum wages in 1970 and 1975 are positively correlated with the potential political power of organized labor and negatively correlated with the potential political base of organized capital. Blais, Cousineau, and McRoberts (1989) apply a similar model to minimum wages in the Canadian provinces over 1975-1982. They find that the relative minimum wage (the minimum wage divided by the average wage in a province) tends to be lower the higher the average wage, the higher the unemployment rate, and the more conservative the provincial government.

Support for Washington's minimum wage increase is likely to depend on voters' political ideology and on economic conditions. Counties that vote more Democratic are more likely to support the initiative. Both Presidential and gubernatorial elections were also held in 1988, and the Democratic vote in those elections is likely to be positively correlated with the percentage of voters supporting the initiative. Economic conditions that might affect support for the initiative include population size, income, unemployment rates, the rate of growth in the low-wage retail trade and service sectors, and the percentage of workers earning low-wages who are likely to be affected by a minimum wage increase. A variety of variables measuring economic conditions over 1980-1988 are used to capture these effects. If economic growth leads to support for minimum wage increases, the growth rates of population and average real income per capita over 1980-1988 and the level of average income in 1988 are likely to be positively correlated with votes in favor of the initiative. Similarly, the unemployment rate in 1988 is likely to be negatively correlated with support for the minimum wage increase. The growth rates of employment and the number of

establishments in retail trade and services over 1987-1988 are likely to be positively correlated with support for the initiative.

Three variables are used to capture the percentage of low-wage workers: the percentage of workers in retail trade and services, the percentage of the population living on farms, and the percentage of the population that are college graduates.[21] The expected correlation between the number of low-wage workers and support for the minimum wage is ambiguous. Low-wage workers may support the initiative because they want higher wages or may oppose it because of potential disemployment effects. If a minimum wage increase would have a negative overall effect on the economy, areas with a large number of low-wage workers might not support the initiative.

The effect of political and economic variables on support for the initiative is examined by regressing the county-level percentage of votes in favor of the initiative on the variables described above. Ordinary least squares estimation is used. The minimum wage ballot initiative won a majority of votes in each county, and the percentage voting in favor of the initiative ranged from 63.9 percent to 81.1 percent. The regression results shown in Table 5 indicate that support for Democratic candidates, particularly the Presidential candidate, is positively correlated with the percentage of votes in favor of the initiative. The correlation between votes for the initiative and for the Democratic Presidential candidate remains statistically significant after controlling for the economic variables.

If the political variables are not included in the regression, the results indicate that

[21]The population, income, and unemployment rate data are from the U.S. Bureau of the Census. The growth rate of employment and establishments in the retail trade and services sectors and the percentage of workers in those sectors are based on the 1987 and 1988 CBP. Jim Snyder generously supplied the voting data.

support for the minimum wage is correlated with economic growth. The growth rate of employment in retail trade and services is positively correlated with support for the initiative, and the growth rate of income is also marginally positively correlated with the vote in favor of the initiative. As shown in Column 2 of Table 5, average income is also positively correlated with support for the minimum wage. The percentage of the population on farms is negatively correlated with the vote. Washington's minimum wage increase extended coverage to most agricultural workers for the first time, and the estimation result implies that agricultural areas may have been particularly concerned about negative effects of the minimum wage increase. Surprisingly, the unemployment rate is marginally positively correlated with support for the initiative.

The regression results for the economic variables change slightly if partisan support is controlled for.[22] As shown in Column 3, population growth over 1980-1988 and average income in 1988 are positively correlated with support for the ballot initiative. The percentage of the population on farms is again negatively correlated with votes in favor of the initiative, as is the percentage of college graduates. Again the unemployment rate is positively correlated with support for the initiative. Although employment growth in retail trade and services is no longer positively correlated with the vote, the growth rate of the number of establishments is positively correlated with support for the initiative.

These results provide are consistent with the proposition that areas in which the minimum wage increase is less likely to have adverse effects are more likely to vote to increase the minimum wage. Either the growth rate of employment or of the number of

[22]All of the results reported in Table 5 are similar if the observations are weighted by the number of votes in the county.

establishments in retail trade and services is positively correlated with support for the initiative. The extension of minimum wage coverage to most agricultural workers would have had a particularly large effect on agricultural areas, and those areas were less likely to support the initiative. Counties with higher incomes, in which the minimum wage increase is less likely to be binding, voted more strongly in favor of the minimum wage increase.

Changes in the retail trade and service sectors after the minimum wage increase provide additional support for the view that counties that would be less adversely affected by the minimum wage increase voted more heavily in favor of the initiative. The growth rates of employment and of the number of establishments in retail trade and services over 1988-1990 are both positively correlated with the county-level vote for the initiative. The correlation coefficient between the percentage of votes in favor the increase and the percentage change in establishments is 0.498, and 0.193 for the percent change in the number of employees. These results are consistent with the possibility that support for the minimum wage increase was higher in areas where it would have less of an adverse effect.

6. Conclusion

This paper examines the effect of two increases in Washington's minimum wage on wages and employment in low-wage retail trade and service industries. Comparisons of wage changes in Washington to the rest of the US generally indicate that the minimum wage increases raised wages in Washington. The results do not conclusively show that the minimum wage increases reduced employment opportunities in Washington. The minimum wage increases also appear to have not raised average payroll per employee in low-wage

counties relative to high-wage counties. These results are surprising given the size of Washington's minimum wage increase, particularly for tipped employees.

The possibility that Washington's increase in the minimum wage was not exogenous with respect to economic conditions may explain the finding that employment did not fall in Washington relative to the rest of the US. If Washington increased its minimum wage because the state's economy was growing and any negative impacts would be minimal, it is not surprising that a comparison of differences in growth rates between Washington and other areas fails to reveal any adverse effects. A comparison of county-level effects might reveal more negative effects of the increase if the ballot initiative passed because of support in areas where it would have a small impact but was imposed on the entire state. The ballot initiative was supported by a majority of voters in all counties, however, and appears to have had no effect on average pay in low-wage counties relative to high-wage counties.

Table 1
Comparison of Wage Data Sources
Washington State Earnings Data

SIC	Industry	1986-88 CPS % Earning 3.35-3.85	1988-89 UI % Δ Pay	1988-89 CBP % Δ Pay	1989-90 UI % Δ Pay	1989-90 CBP % Δ Pay
78	Motion Pictures	40.2	5.4	-1.7	11.7	5.4
58	Eating/Drinking Places	38.1	7.2	3.6	7.5	9.0
59	Misc. Retail	25.6	3.0	4.3	3.0	6.4
54	Food Stores	20.2	1.6	0.3	6.0	2.5
56	Apparel	19.2	5.5	12.2	3.3	-2.6
55	Auto Retail	19.1	5.4	-2.9	7.3	10.2
79	Entertainment	17.4	4.5	-7.1	5.9	-3.1
70	Hotels	14.5	2.7	8.4	6.6	2.4
53	General Merchandise	12.6	3.7	6.8	5.9	2.6
57	Furniture	12.1	5.2	3.2	5.3	5.5
52	Building Materials	0.0	1.8	3.1	6.3	5.3
Coefficient in Regression of Change in Nominal Payroll on Percentage Affected			0.182 (0.021)	0.086 (0.064)	0.263 (0.029)	0.175 (0.042)

Notes: The CPS data are from the Outgoing Rotation Group files for Washington and are weighted. The UI data are from the Bureau of Labor Statistics publication *Employment and Wages*, various years. The regressions of changes in payroll on the percentage of workers earning $3.35 to $3.85 do not include a constant.

Table 2
Differences in Percentage Changes in Average Payroll, 1988-1990
Washington and Rest of the U.S.

SIC	Industry	1988-89 UI	1988-89 CBP	1989-90 UI	1989-90 CBP
78	Motion Pictures	3.6	0.3	2.9	3.6
58	Eating/Drinking Places	3.2	-0.7	3.7	4.3
59	Misc. Retail	-1.3	-0.8	-0.9	1.9
54	Food Stores	-0.9	-1.6	2.3	-0.8
56	Apparel	2.8	7.8	-0.7	-6.1
55	Auto Retail	4.9	-4.6	4.6	7.1
79	Entertainment	4.2	-2.8	-3.7	-7.2
70	Hotels	-1.2	4.3	1.7	-2.6
53	General Merchandise	1.4	-2.7	1.1	-1.3
57	Furniture	2.1	-0.7	1.8	2.1
52	Building Materials	-0.5	-1.1	3.2	1.1
Coefficient in Regression of Difference in Payroll Changes on Percentage Affected		0.082 (0.024)	-0.006 (0.035)	0.062 (0.025)	0.037 (0.048)

Notes: The percentage of workers affected is based on CPS data for workers paid an hourly wage in Washington over 1986-1988 and is the (weighted) fraction of workers earning $3.35 to $3.85. The regressions do not include a constant.

Table 3
Differences in Percentage Changes in Employment
Washington and Rest of the U.S.
And OLS and IV Regression Results, 1988-1990

SIC	Industry	1988-89 UI	1988-89 CBP	1989-90 UI	1989-90 CBP
78	Motion Pictures	3.3	23.3	0.4	-10.2
58	Eating/Drinking Places	1.3	1.8	3.0	3.9
59	Misc. Retail	-3.8	2.2	1.7	1.5
54	Food Stores	5.6	4.5	0.1	1.8
56	Apparel	-1.7	-5.1	2.7	9.1
55	Auto Retail	3.4	6.1	4.9	3.7
79	Entertainment	-1.6	2.7	6.1	-0.7
70	Hotels	3.5	8.5	1.1	-3.0
53	General Merchandise	-1.5	-2.1	0.0	5.6
57	Furniture	10.3	1.0	8.3	9.0
52	Building Materials	4.8	-0.7	11.0	12.8
Coefficient in Regression of Difference in Employment Changes on Difference in Wage Changes (OLS)		-0.562 (0.301)	-0.298 (0.455)	0.446 (0.327)	0.120 (0.209)
Coefficient in Regression of Difference in Employment Changes on Difference in Wage Changes (IV)		-0.449 (0.545)	3.412 (7.380)	0.465 (0.642)	-0.357 (0.506)

Notes: All regressions include a constant and are weighted by industry employment. Standard errors and shown in parentheses, and the number of observations is 11.

Table 4
Differences in Percentage Changes in Employment Washington and Rest of the U.S. And OLS and IV Regression Results, 1988-1990

Specification	1988-89 UI	1988-89 CBP	1989-90 UI	1989-90 CBP
Coefficient in Regression of Difference in Employment Changes on Difference in Wage Changes (OLS)	-0.149 (0.184)	-0.205 (0.303)	-0.227 (0.501)	-0.252 (0.150)
Results of First-Stage Regression of Difference in Wage Changes on Percentage Affected	0.099 (0.035)	0.031 (0.028)	0.080 (0.020)	0.141 (0.041)
Coefficient in Regression of Difference in Employment Changes on Difference in Wage Changes (IV)	-0.803 (0.341)	-0.423 (1.609)	0.235 (0.662)	-0.056 (0.423)

Notes: All regressions include a constant and are weighted by industry employment. Standard errors are shown in parentheses, and the number of observations is 37.

Table 5
Determinants of Votes for the Ballot Initiative

Covariates	(1)	(2)	(3)
Percentage Vote for Democratic Presidential Candidate	.386 (.105)	--	.420 (.125)
Percentage Vote for Democratic Gubernatorial Candidate	.191 (.086)	--	.049 (.075)
Percentage Change in Population, 1980-1988	--	.025 (.064)	.129 (.051)
Percentage Change in Real Income Per Capita, 1980-1988	--	.180 (.100)	-.026 (.070)
Income Per Capita	--	.497 (.203)	.386 (.116)
Unemployment Rate	--	.463 (.266)	.324 (.155)
Percentage Change in Employment in Retail Trade and Services, 1987-1988	--	.122 (.034)	-.007 (.030)
Percentage Change in Establishments in Retail Trade and Services, 1987-1988	--	.205 (.127)	.187 (.070)
Percentage of Workers in Retail Trade and Services	--	-.033 (.054)	.050 (.035)
Percentage of Population on Farms in 1980	--	-.842 (.091)	-.339 (.118)
Percentage of Population College Graduates	--	-.092 (.069)	-.095 (.041)
Adjusted R-squared	.589	.579	.812

Notes: The dependent variable is the percentage of votes in a county in favor of the ballot initiative to increase Washington's minimum wage. Standard errors are shown in parentheses, and the number of observations is 39.

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