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Labor - Minimum Wage

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EXECUTIVE OFFICE OF THE PRESIDENT

COUNCIL OF ECONOMIC ADVISERS

WASHINGTON, D.C. 20500

MEMBER

July 18, 1996

MEMORANDUM FOR JOSEPH E. STIGLITZ
MARTIN N. BAILY

FROM: ALICIA H. MUNNELL

SUBJECT: Minimum Wage

Attached is a copy of the minimum wage exercise that the Fed staff prepared for the May Green Book. As you can see, this analysis suggests that the Federal funds rate would have to be raised by .75 percent immediately and held throughout 1997 to offset potential effect on inflation of a higher minimum wage.

Attachment

in the projection may encourage some short-term borrowing, but the bond market is expected to remain an important source of funds.

Net borrowing by households is projected to trend down gradually through 1997. Consumer credit growth is expected to slow, with tightening loan policies only a minor factor. The strong borrowing in recent years has left a legacy of rising repayments; and growth in nominal durable expenditures, which are often financed by credit, is expected to be moderate. Home mortgage debt growth is expected to edge down from its pace of recent years, restrained by the effects of higher interest rates and the associated falloff in home sales.

The Minimum Wage

Although we have not incorporated a hike in the minimum wage in the current projection, we have examined the effects of a 90 cent per hour increase from $4.25 per hour to $5.15 per hour, with half coming in July 1996 and the remainder in July 1997. Using both macroeconometric models and detailed data on the distribution of wages, we estimate that such a hike would have a direct effect on the level of hourly compensation of about 0.4 percent.[5]

The pass-through to prices, of course, would depend on the conduct of monetary policy. If the federal funds rate were held at its current level, the higher minimum wage would probably be passed gradually into prices and tend to get built into expectations; under this assumption, we estimate CPI inflation would be higher by about 0.3 percentage point in 1997 and a bit more than that in 1998. Simulations of the staff's quarterly econometric model suggest that the federal funds rate would have to be raised roughly 3/4 percentage point immediately and held at the higher level through 1997 to offset the potential effect on inflation of the higher minimum wage.

Alternative Simulations

Setting aside the minimum wage question, we have run two model simulations in which the funds rate is symmetrically raised or lowered relative to the assumption in the Greenbook. Deviations

5. Data from the household survey indicate that there are about 7 million hourly workers now earning between $4.25 and $5.15 per hour, and another 7-1/2 million hourly workers who currently earn between $5.15 and $6.00 per hour and whose wages would probably be raised to preserve at least some of the differential between them and their lesser-paid colleagues. In addition, there are 1-1/2 million hourly workers in the household survey who currently earn less than the minimum wage and whose wages also might be lifted by competitive pressures.

from baseline start at 10 basis points in the current quarter and increase to 50 basis points in the third quarter and to 100 basis points in the fourth quarter and beyond. In the lower-rate scenario, real GDP growth is little different this year and 0.6 percentage point higher in 1997. The unemployment rate is reduced 0.2 percentage point by the end of 1997, and core CPI inflation is 0.2 percentage point higher than in the baseline forecast. The effects of the tighter policy scenario are symmetric.

ALTERNATIVE FEDERAL FUNDS RATE ASSUMPTIONS
(Percentage change, Q4 to Q4, except as noted)

	1996	1997
Real GDP		
Baseline	2.5	2.2
Lower funds rate	2.6	2.8
Higher funds rate	2.4	1.6
Civilian unemployment rate (percent)[1]		
Baseline	5.5	5.5
Lower funds rate	5.5	5.3
Higher funds rate	5.5	5.7
CPI excluding food and energy		
Baseline	2.9	3.2
Lower funds rate	2.9	3.4
Higher funds rate	2.9	3.0

1. Average for the fourth quarter.

EXECUTIVE OFFICE OF THE PRESIDENT
COUNCIL OF ECONOMIC ADVISERS
WASHINGTON, D.C. 20500

SENIOR ECONOMIST

May 9, 1996

Labor (Min wage)

MEMORANDUM FOR JOSEPH STIGLITZ
ALICIA MUNNELL
MARTIN BAILY
MICHELE JOLIN
PETER ORSZAG

FROM TOM KANE TJK

RE TEAM ACT

Some in Congress have proposed linking the minimum wage increase with the TEAM Act proposal. As background for those discussions, I have attached a copy of a memo I wrote back in September regarding the TEAM Act.

Memorandum

To: Joseph E. Stiglitz
Alicia Munnell
Martin Baily
Michele Jolin
From: Tom Kane
Re: TEAM Act
Date: September 26, 1995

The Problem:

Under current law, the National Labor Relations Act appears to prohibit a very broad range of employer-sponsored employee-participation activities:

> Section 8(a)(2) of the National Labor Relations Act makes it an unfair labor practice to *"dominate or interfere with the formation or administration of any labor organization or to contribute financial or other support to it."*
>
> Section 2(s) defines a labor organization as *"any organization of any kind...in which employees participate and which exists for the purpose, in whole or in part, of dealing with employers concerning grievances, labor disputes, wages, rates of pay, hours of employment or conditions of work."*

In the NLRB's oft-cited Electromation case in 1992, plant workers had been discontent after a management-imposed cut in attendance bonuses. After 68 employees in the non-union shop signed a petition protesting the change, management chose from among employee volunteers to set up committees to discuss pay progression rules, a no-smoking ban and the company's communication network. Soon afterward, the Teamster's union began a membership drive and filed a petition charging a violation of section 8(a)(2). If any organizing efforts had been ongoing before the establishment of the committees, there was no evidence that management knew about them. In a decision that is currently under appeal, the NLRB ruled that the company had engaged in an unfair labor practice because the company had established and supported a "labor organization" to discuss wages and the conditions of work.

When employee-participation programs-- even those originally established to discuss productivity or efficiency issues-- discuss wages or working conditions in the course of their deliberations, such programs may be illegal under a strict interpretation of the NLRA. But they are also ubiquitous. In a survey of establishments with more than 50 employees in 1991, two-thirds were found to have some form of employee-participation program. It is unclear how many of these committees discuss productivity issues solely or also drift into discussions of wages or working conditions. In the wake of the Electromation decision and other controversial decisions, the business community has sought some loosening of what they believe to be a overly broad prohibition of employee-participation programs. A piece of legislation aimed at meeting those concerns, the TEAM Act, is scheduled for a vote on Wednesday, September 27.

The TEAM Act

The TEAM Act (H.R. 743) takes an extreme approach to resolving the ambiguity. Indeed, the language is quite direct:

> "..it shall not constitute...an unfair labor practice...for an employer to establish, assist, maintain or participate in any organization or entity of any kind in which employees participate to address matters of mutual interest, including, but not limited to, issues of productivity, efficiency and safety and health, and which does not have, claim or seek authority to be the exclusive bargaining representative of the employees or to negotiate or enter into collective bargaining agreements with the employer or to amend existing collective bargaining agreements..."

The bill has been amended to conclude with the following phrase:

> "...except that in a case in which a labor organization is the representative of such employees...this proviso shall not apply."

The above amendment would prevent employers from setting up their own committees to which they could grant concessions in an attempt to break an existing union. However, the latest amendment clearly does not go far enough in protecting existing unions for at least two reasons:

First, as in the usual 8(a)(2) case in the past, an employer may establish such a committee to forestall a union organization drive. The TEAM Act would go far beyond providing legitimacy to quality circles or productivity committees. It would legalize a potent union avoidance technique.

Second, one purpose of the NLRA was to define the circumstances under which employees could expect to reach a legitimate bilateral agreement with their employers. That is why the original act explicitly prohibits any arrangements under which employers essentially have veto power over the result. The TEAM Act not so subtly upsets that balance.

The *Status Quo*

If the House and Senate both pass some version of the TEAM Act and if the President vetoes the bill, there is not likely to be a crisis any time soon. Despite alarmist news coverage in the *Wall Street Journal*, the legal peril faced by benign employee-participation plans has been overstated for at least two reasons:

> First, in its opinions, the NLRB has made it clear that programs which are primarily intended to discuss product quality and efficiency-- which only incidentally touch upon issues of wages or working conditions-- would not represent 8(a)(2) violations.
>
> Second, while it is not necessary to prove an employer's union-avoidance motive to claim a 8(a)(2) violation, as a practical matter, an employer is very unlikely to be challenged except in cases where a union organizing campaign is ongoing or where a employer establishes such committees without union approval.

According to one recent analysis, an average of 3 cases involving 8(a)(2) violations have been brought to the NLRB over the past 20 years. If the union movement were to launch a campaign to unmask violations of Section 8(a)(2), they could probably force the need to resolve the ambiguity. However, there is little evidence of such activity to date. Indeed, it is not at all clear that it is in the union movement's interest to pursue that strategy, because it would only hasten the passage of legislation similar to the TEAM Act.

Alternatives to the TEAM Act

One approach would be to codify the case law emerging from the NLRB on 8(a)(2) and find someone to sponsor it on the Hill. However, I am under the impression that attempts to draft such a bill became bogged down. One potential reason is that the NLRB is currently considering several cases involving 8(a)(2) violations which it might use to provide further clarification.

However, simply codifying case law may not be sufficient over the long-term. Although we do not have much direct evidence, there are probably many employee-participation programs out there which go beyond quality and efficiency issues to discuss wages and the conditions of work. Therefore, it may be reasonable to seek to define quasi-representative structures for the 87% of the workforce that is not under a collective bargaining agreement. One proposal would be to define certain minimum principles that such programs could satisfy-- short of becoming a traditional labor union with legal power to negotiate a collective bargaining agreement-- which would legitimize their discussion of wages, grievances and working conditions with employers. For instance, there is a proposal by Charles Morris-- supported by some labor experts such as Tom Kochan, Harry Katz and Clyde Summers-- which would define the following set of conditions to legitimize such discussions:

(1) A majority of affected employees vote by secret ballot to select representatives and approve the body.
(2) These representatives are not supervisors or other members of management.
(3) There was no union organizational drive in process.
(4) The employer is not in violation of other unfair labor practices.

One would probably also want requirements that employers can not retaliate against employees that participate in these discussions.

In a recent survey conducted by Richard Freeman and Joel Rogers, many workers expressed an interest in such alternative, less adversarial forms of representation. In addition, it is interesting to note that 40% of the same workers reported that they would have voted for a union in a union election, roughly 3 times as many as were actually union members.

You should know that any such proposal is HIGHLY CONTROVERSIAL. Any change strikes near the heart of the union movement. In effect, such language would legitimize forms of representation other than unions to discuss wages and working conditions with employers. Perhaps not surprisingly, such proposals currently have little political backing. Employer representatives are holding out for language such as in the TEAM Act. And the labor movement-- engaged in an election of a new president of the AFL-CIO-- is not in a position to discuss so fundamental a reform.

If the TEAM Act passes both the House and the Senate and the President vetoes the bill, we will be left with the status quo. And, on this issue, the status quo is certainly livable for the medium term.

TO: JBS, AHM, MNB, MJ, SB
from: PO

BY GENE KORETZ

WANT TO CUT MEDICAL COSTS?

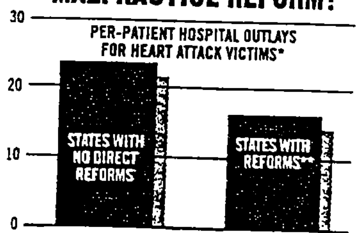
Put a lid on malpractice suits

One of the hottest issues in the U.S. debate over rising medical costs is the impact of malpractice suits. Critics claim the malpractice system vastly boosts health-care outlays by inducing doctors to practice defensive medicine—that is, prescribing tests and treatments with little or no benefit in order to ward off lawsuits. Defenders say it reduces negligence while providing injured patients with redress for doctors' mistakes.

Moreover, since the total cost of the malpractice system, including claims paid and lawyers' fees, comes to less than 1% of the nation's medical bill, its supporters argue that its effect on expenditures is highly exaggerated. Indeed, up to now, there has been little hard evidence to show how much—or even whether—doctors resort to defensive medicine.

In a new study, two Stanford University economists—Daniel P. Kessler, who is also a lawyer, and Mark McClellan, who is also a doctor—seek to fill this gap. Their findings suggest that defensive medicine has been a widespread and costly practice, one that could be reduced through liability reforms.

A PAYOFF FROM MALPRACTICE REFORM?



▲ PERCENT RISE 1984 TO 1990
DATA: DANIEL KESSLER AND MARK McCLELLAN
*MEDICARE RECIPIENTS
**REFORMS EFFECTIVE 1985-87

Kessler and McClellan analyzed hospital expenditures for all elderly Medicare patients hospitalized in 1984, 1987, and 1990 with recently diagnosed heart attacks and other serious heart ailments. Controlling for patients' age, race, and sex, they compared states that had implemented malpractice reforms with those that had not, in terms of how much their spending per patient grew and how the patients fared.

The results: In states that adopted reforms limiting liability awards, the growth in hospital outlays slowed appreciably for several years—lowering such expenditures by 5% to 9% relative to the levels reached in states with no reforms. In both groups of states, patients did equally well—measured by mortality rates and rehospitalizations.

Although the spending slowdowns that followed reforms may not have affected the long-term upward trend in medical costs, they did seem to provide a lasting one-time saving. Extrapolating from their results, Kessler and McClellan theorize that nationwide comprehensive limits on malpractice awards could reduce annual health-care outlays by as much as $50 billion.

In fact, the recent slowdown in U.S. medical expenditures may partly reflect the malpractice reforms adopted by many states over the past decade—as well as the huge growth in managed-care plans, which closely monitor physicians for signs of excessive use of tests and costly treatments.

CHINA'S OTHER CONSUMERS

Suburbanites are spending, too

The common view, says economist Gilbert Choy of ING Baring Securities Ltd., is that China's growing consumer market is tied to its urban population, which represents about 20% to 25% of the nation's 1.2 billion people. With annual per capita disposable income of about $465 (and much more in purchasing-power), urban dwellers reportedly earn about 2¼ times as much as their poverty-stricken rural cousins.

The trouble with this picture, says Choy, is that it not only undercounts city dwellers but also misses a huge adjunct to the main consumer market: an increasingly affluent suburban economy that has sprouted on the fringes of the cities but is often classified as part of the rural economy. According to ING Baring, suburbanites and better-off peasants number some 384 million and enjoy average per capita incomes of $450, almost as much as urbanites.

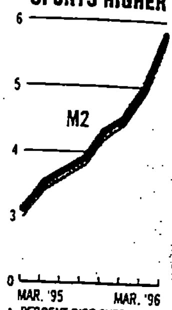
Counting the urban and suburban groups together, Choy estimates they exceed two-thirds of China's population and account for some 80% of its gross domestic product. Their purchasing power helps explain why real retail sales on the mainland continue to grow at a hefty 10% annual pace despite the nation's recent austerity policies.

MONEY, MONEY EVERYWHERE

The Fed won't ease again soon

If the Federal Reserve is pursuing a mildly restrictive monetary policy, as many financial market observers say, that stance isn't showing up in the various measures of the money supply. M1, the narrowest monetary measure, jumped at a 9.8% annual clip in March, while the annual growth rates of M2 and M3 both hit 10% to 11% and have hardly slackened in recent weeks.

THE MONEY SUPPLY SPURTS HIGHER



▲ PERCENT RISE OVER 12 MONTHS
DATA: FEDERAL RESERVE

Higher and speedier tax refunds may be behind some of this surge in liquidity, but not all of it. William V. Sullivan Jr. of Dean Witter Reynolds Inc. observes that since March, 1995, the 12-month growth rate of M2 has soared from 0.4% to 5.8%—"the most rapid acceleration on record." Moreover, both M2 and M3 are growing in inflation-adjusted terms for the first time in years. "The recent sharp pickup in monetary growth suggests that the economy is rebounding smartly," says Sullivan, "and that further easing moves by the Fed are unlikely any time soon."

A MINIMAL WAGE EFFECT

Will a hike raise labor costs?

Though they are not happy about a rise in the minimum wage, two Wall Street economists see only a slight impact on wage inflation. Bruce Steinberg of Merrill Lynch & Co. notes that only 5.3% of hourly workers earned the minimum wage or less last year. And nearly half of that group got less than the minimum—presumably because they are exempt from the law's coverage.

Maury N. Harris of PaineWebber Inc. observes that in seven of the nine times the minimum wage has been raised since 1968, hourly earnings slowed over the following year. "If labor markets tighten, wage pressures may emerge," he says. "But a higher minimum wage by itself won't have much of an effect."

ERIC HOFFMANN/BW

EXECUTIVE OFFICE OF THE PRESIDENT
COUNCIL OF ECONOMIC ADVISERS
WASHINGTON, D.C. 20500

April 22, 1996

SENIOR ECONOMIST

MEMORANDUM FOR JOSEPH STIGLITZ
MARTIN BAILY
ALICIA MUNNELL
PETER ORSZAG
LOUISE SHEINER

FROM TOM KANE

SUBJECT MINIMUM WAGE AND RELATED ISSUES

Senator Dole has mentioned two possible ways to make the minimum wage hike more palatable for the business community: a youth subminimum wage and some flexibility on the overtime provisions of the Fair Labor Standards Act. Below, I provide some background on each of these options.

1. Between 1990 and 1993, employers were allowed to pay a "subminimum wage" to youth aged 19 or under for up to 180 days. Little paperwork was required of employers seeking to take advantage of the subminimum for the first three months (90 days). To use the second 90 days of "subminimum wage" eligibility, an employer would be required to file a "training plan" with the Department of Labor. The "subminimum wage" provision expired in 1993 with little fanfare.

Very few employers have chosen to take advantage of the youth subminimum wage when it has been available in the past. For instance, Katz and Krueger found that only 4.8 percent of fast-food restaurants in Texas used the subminimum wage in 1991. Moreover, Katz and Krueger have also found that the introduction of the youth subminimum in 1990 had no discernible effect on the wages of youth.

Therefore, particularly if passed as a time-limited measure, a youth subminimum wage would not seriously weaken the effect of a minimum wage increase.

2. The Fair Labor Standards Act requires that employers pay overtime wages (1.5 times regular hourly pay) to covered workers working more than 40 hours in a given week. The largest exemptions are listed below:

- Executive, administrative and professional employees.
- Employees of seasonal amusement or recreational establishments.
- Agricultural workers on small farms.
- Persons employed as companions to the elderly or infirm.
- Most employees of air carriers, trucking companies, taxi services and shipping companies.

The May 1985 CPS contained a supplemental question to identify workers receiving overtime pay. Roughly 10 percent of all employees received overtime pay during the survey week in May 1985. Nearly half of those receiving overtime pay were in construction, manufacturing and mining (even though only about a quarter of all wage and salary workers worked in those industries).

The purpose of the original overtime-pay provision in the Fair Labor Standards Act of 1938 was to encourage employers to hire additional workers rather than make use of overtime. Indeed, the *Economic Report of the President* in 1964 advocated raising the overtime pay requirement from "time and one-half" to "double-time" to raise employment.

However, the expected impact of overtime pay restrictions on wages and hours is not clear. If employers and employees bargain over a *package* of compensation for an expected number of hours, it is unclear that such restrictions *would* have any effect on "effective" wages or employment. Employers and employees could simply agree on a straight-time wage and overtime pay that yielded a desired average wage-- and the right incentives for employers to smooth out the work. A recent paper by Trejo (attached) reported that straight-time wages were lower for covered workers working overtime than for workers with similar characteristics and similar hours who were not covered. Only those workers with a straight-time wage equal to the minimum wage would be constrained in making such a bargain with their employers. Trejo also found that covered workers with a straight-time wage equal to the minimum wage were less likely to be paid overtime when they worked more than 40 hours.

Therefore, in industries and occupations that are known to have fluctuating hours, we might expect the overtime pay provisions to have little effect since straight-time wages could adjust downward. However, such restrictions may have a larger impact when an employer has an atypical "spike" in demand, requiring employers to bear the cost of unanticipated inconveniences for workers. This may explain the fact that Trejo finds that only about half of the additional cost of overtime pay is offset by lowered straight-time wages.

A proposal has been floated to loosen the straight-time limit from 40 hours in a single week to 160 hours in 4 weeks. My initial reaction is that I do not believe that such a change would have a dramatic effect on overtime pay or employment for the following reasons.

First, given the Trejo results, the 40 hour limit may not be binding for many employer-employee bargains.

Second, a vast majority of those who ever work *more than* 40 hours in a week, *usually* work 40 hours per week, not less than 40. (In the 1993 CPS, 92 percent of those who reported working more than 40 hours in the previous week *usually* worked 40 hours or more.) In other words, an employer who wanted to avoid paying overtime would have to reduce an employee's work hours in subsequent weeks.

However, the politics of the proposal are likely to be complicated. Representatives of organized labor-- whose members probably benefit little from the minimum wage hike itself-- would probably resist any loosening of the overtime pay requirements. Such a proposal could drive a wedge into the coalition supporting the minimum wage.

cc: mJ

The Effects of Overtime Pay Regulation on Worker Compensation

By STEPHEN J. TREJO*

Proponents claim that a statutory overtime premium, by raising the relative cost of overtime, may encourage firms to substitute employment for overtime hours. I argue that there will be no real effects if firms reduce straight-time wages so as to offer the same package of weekly compensation and hours of work that was acceptable initially. Empirical analysis suggests that wage differentials do arise to mitigate the purely demand-driven effects predicted by previous models, but these differentials are not large enough to neutralize overtime pay regulation completely. (*JEL* J33, J38).

The Fair Labor Standards Act of 1938 (FLSA) introduced two forms of federal wage regulation into the U.S. labor market: a minimum wage and a requirement that an overtime wage of at least one and one-half times the straight-time hourly wage be paid for weekly hours worked in excess of 40. In contrast to the flood of research examining minimum wages, very little academic attention has been paid to the overtime pay provisions of the FLSA.[1] Such neglect is surprising because the overtime pay provisions potentially affect many more workers than the minimum-wage provisions: about a quarter of all employees work overtime during a typical week (Daniel E. Taylor and Edward S. Sekscenski, 1982), whereas only 11 percent of hourly paid workers earn the minimum wage or below (Earl F. Mellor and Steven E. Haugen, 1986). This paper investigates the economic consequences of overtime pay regulation.

This project acquires added urgency and policy significance from periodic proposals before Congress to increase the required overtime wage to twice the straight-time hourly wage.[2] Proponents argue that such a measure would induce firms to cut back on overtime hours assigned and instead hire more workers to take up the slack; thus, a kind of economy-wide work-sharing would ensue to reduce unemployment. Neoclassical labor-demand theory, appropriately adapted to treat this problem, lends some support to such a claim. If straight-time wages rates are taken as fixed, increasing the overtime premium reduces the cost of expanding output by hiring another worker relative to the cost of expanding output by employing the current work force for an additional hour, so firms are encouraged to substitute employment for overtime hours. Of course, increases in the overtime premium also raise the marginal cost of production, thereby generating an opposing scale effect, which leaves the net effect on employment theoretically indeterminate. Within this framework, the social cost of overtime pay regulation is the usual distortion introduced by any tax on a factor of production.

*Department of Economics, University of California, Santa Barbara, CA 93106. For advice and comments, I am grateful to Gary Becker, George Borjas, Steve Bronars, Donald Deere, Ron Ehrenberg, Ted Frech, Dan Hamermesh, Robert Hart, Ed Lazear, Tom Mroz, Kevin Murphy, Sherwin Rosen, Perry Shapiro, Bob Topel, two referees, and participants in various seminars.

[1] For an introduction to the vast literature on the effects of minimum wages, see the survey by Charles Brown et al. (1982). Previous inquiries into the effects of overtime pay regulation have been conducted by Ronald G. Ehrenberg (1971), Joyce E. Nussbaum and Donald E. Wise (1977), and Ehrenberg and Paul L. Schumann (1982b).

[2] An example of such a proposal is HR 2933 introduced in 1985 by Representative John Conyers of Michigan.

An alternative view of the labor market predicts that the overtime pay provisions of the FLSA will have little effect on real economic behavior. Following the ideas of H. Gregg Lewis (1969), a job is simply a tied package of compensation and working conditions. A job that usually requires a fixed number of hours of work per week must provide workers with enough compensation to make them at least as well off as any other job package they could obtain. According to this view, increasing the statutory overtime premium will have no real effects since, by reducing the straight-time hourly wage rate, the firm can offer the same package of total weekly compensation and hours of work that was acceptable to the worker initially. For example, suppose that in the absence of overtime pay regulation a job pays $550 per week and requires 50 hours of work, so the average wage is $11 per hour. An employer may satisfy the FLSA by reducing the straight-time wage to $10 per hour and paying time and a half for the 10 overtime hours. The employee works 50 hours and receives $550 per week in either case; consequently, the law has no real effect. Moreover, overtime pay regulation leaves total employment unaffected and imposes no social costs. In this alternative view of the labor market, overtime pay regulation only affects those workers with straight-time wage rates that cannot be fully adjusted, such as those of workers earning the legal minimum wage.

The next two sections present these alternative approaches to modeling the consequences of overtime pay regulation. The approach taken by previous studies, which I will call the "fixed-wage" approach to overtime pay regulation, attributes real economic effects to the overtime law and in particular predicts that raising the statutory overtime premium should reduce the use of overtime hours and increase the prevalence of a 40-hour work week. An alternative approach is developed, a "fixed-job" model, which suggests that hourly wage flexibility may be sufficient to rob overtime pay regulation of any substantive impact. The fixed-wage and fixed-job models are theories of overtime pay *regulation*, rather than of overtime pay *provision*, in that these models attempt to describe the labor-market effects which would result from the exogenous imposition of an overtime premium. Without further assumptions about worker preferences, they have nothing to say about why firms would voluntarily pay an overtime premium in the absence of regulation, or why in the presence of regulation firms would adopt an overtime pay schedule more generous than that required by law.[3]

The remainder of the paper attempts to determine whether the fixed-job model is of any empirical importance. My strategy is to compare workers covered by the overtime law with noncovered workers using microdata from the 1974, 1976, and 1978 May Current Population Surveys. Section III describes the data and discusses the classification of workers according to coverage status under the minimum-wage and overtime pay provisions of the FLSA. Section IV examines compliance with the overtime law, under the presumption that incentives to disregard the law should be greatest for those employers most adversely affected by the law. Section V searches for differences in distributions of weekly hours of work that are attributable to overtime pay coverage. Section VI compares the hourly wages and weekly earnings of covered and noncovered workers to investigate the extent to which adjustments occur to offset the statutory overtime premium. I test not only for the existence of straight-time wage differentials between the covered and noncovered sectors, but also whether these differentials are sufficiently large to neutralize overtime pay regulation completely. A final section summarizes and concludes.

I. The Fixed-Wage Model of Overtime Pay Regulation

Ehrenberg (1971) adapts the framework found in Rosen (1968) to illustrate the sub-

[3]Worker preferences may require that jobs with different hours schedules offer different wage schedules, as set out in the compensating-differential models of Lewis (1969), Sherwin Rosen (1974), and John M. Abowd and Orley Ashenfelter (1981).

stitution of employment for overtime hours induced by an increase in the statutory overtime premium. Capital is assumed to be fixed, and the firm faces an exogenous output requirement to be met by choosing the number of employees and weekly hours worked by each employee. The production technology allows the marginal product of an additional man-hour of work to differ depending on whether it is obtained by expanding employment or through increased utilization of the existing force. With the straight-time hourly wage rate taken as fixed, an increase in the overtime premium raises the marginal cost of overtime hours relative to the marginal cost of hiring an additional worker. Cost-minimizing firms respond by expanding employment and cutting back on overtime hours.[4]

One other potential consequence of overtime pay regulation is readily apparent. If the straight-time wage is exogenous, then the overtime pay structure imposed by the FLSA creates a kink in the firm's cost function at the statutory work week standard of 40 hours. It is easy to show that this kink induces some firms that would assign overtime in the absence of overtime pay regulation to choose instead the corner solution in which they avoid paying the overtime premium by limiting weekly hours of work to 40. Thus, the fixed-wage model of overtime pay regulation provides one possible explanation for the well-known fact that the hours-of-work distribution has a spike at 40 hours. This issue is investigated in the empirical work reported below.

Existing tests of the fixed-wage model have focused on the prediction that an overtime premium should discourage the use of overtime hours. Ehrenberg and Schumann (1982b) provide the most recent estimates and also review previous work. All studies use establishment-level data from various years of the Employer Expenditure for Employee Compensation surveys. The basic empirical methodology is to regress annual overtime hours per employee on control variables and the ratio of quasi-fixed labor costs to the overtime wage. Typical findings indicate a statistically significant positive association across establishments between this ratio and the use of overtime. However, Ehrenberg and Schumann (1982b) show that this result is driven by the strong correlation between quasi-fixed labor costs and overtime hours: when quasi-fixed labor costs and the overtime wage enter the regressions separately rather than in ratio form, the coefficients on the overtime wage often have the wrong sign and tend to be statistically insignificant. Moreover, virtually all firms are covered by the FLSA within most of the industries on which these regressions are run, so there is very little variation in the overtime premium with which to estimate directly the impact of overtime pay regulation. For these reasons, existing empirical analyses of the overtime-hours response generated by a statutory overtime premium are inconclusive.

II. A Fixed-Job Model

The fixed-wage model described above lends theoretical support to the idea that raising the overtime premium required by the FLSA would reduce overtime hours and possibly stimulate employment. However, this model assumes that the supply of hours is infinitely elastic at the given market wage rate and that this market wage is unaffected by overtime pay regulation. As suggested by Lewis (1969), the length of the work week can be viewed as a job aspect over which both firms and workers have preferences, with compensating wage differentials arising in equilibrium for jobs with differing hours of work. Under this alternative characterization of the labor market, differentials in straight-time hourly wages could arise to render overtime pay regulation almost completely inconsequential.

[4]This analysis ignores scale effects. With output endogenous, raising the overtime premium will reduce overtime hours so long as the scale effect is nonperverse, but the net impact on employment is indeterminate because of conflicting substitution and scale effects (Nussbaum and Wise, 1977). Similar remarks apply when capital is variable, because increasing the overtime premium encourages substitution out of labor inputs and into capital.

The basic idea is simple. A job consists of a given package of weekly hours of work and weekly earnings. The FLSA requires that weekly hours worked in excess of 40 be compensated at a rate of at least one and one-half times the straight-time hourly wage, but the act does not regulate the straight-time hourly wage except to require that it not fall below the legal minimum wage. So long as *hourly* wage rates are flexible, the overtime law does not in any way restrict the ability of workers and firms to contract over packages of *weekly* hours and earnings. Changes in the overtime premium or work-week standard set by law would produce perfectly offsetting changes in straight-time hourly wage rates so as to leave weekly hours and earnings unchanged.[5]

To see this point, consider the situation in which overtime pay regulation is in place and overtime hours are being worked. In this case, a worker's weekly earnings Y can be expressed as

$$(1) \qquad Y = w\left[\bar{H} + b(H - \bar{H})\right]$$

where H represents weekly hours of work, w is the straight-time hourly wage rate, b denotes the overtime premium required by law, and $\bar{H}$ is the weekly-hours standard at which the overtime premium must go into effect (the FLSA currently requires $b \geq 1.5$ and $\bar{H} \leq 40$). Holding hours of work constant, the firm is indifferent between any combinations of straight-time and overtime wage rates that produce the same level of weekly earnings. Exactly the same kind of indifference will hold for standard formulations of worker preferences, in which an individual's utility is a function of income for consumption and hours of work. That is, workers care only about how much weekly compensation they receive for a given amount of effort, and it does not matter to them how this compensation is split between straight-time pay and overtime pay. Suppose that in the absence of the FLSA a given employee works $H_0 > 40$ hours and is paid w_0 for each hour. When the FLSA is introduced, a straight-time wage of $w' = w_0 H_0/(1.5H_0 - 20)$ leaves the worker at precisely the same level of welfare as he was initially.

Under these circumstances, the firm and its workers may be expected to come to an agreement regarding the quantities that they jointly care about: weekly earnings and weekly hours. Regardless of exactly how this earnings–hours package is determined, equation (1) reveals that flexibility of the straight-time hourly wage is sufficient to neutralize overtime pay regulation. Whatever levels initially chosen for Y and H, w can adjust to legislative changes in b and $\bar{H}$ so as to maintain these initial levels. This extra degree of freedom allows weekly hours, weekly earnings, and employment to be unaffected by overtime pay regulation. Because in this model it is the job package of weekly hours and earnings (rather than the straight-time wage) that remains fixed, I refer to this as the fixed-job model of overtime pay regulation.

The preceding analysis has been entirely of a static nature, focusing on the determination of regularly scheduled overtime. It might well be objected that much of overtime is due to stochastic events such as unexpected demand, machine breakdowns, and absenteeism and that a dynamic model is necessary to capture these factors adequately. The simple fixed-job model developed above requires perfect wage flexibility to justify its strong conclusion that overtime pay regulation fails to have any real economic impact. This result will obviously be robust to dynamic specifications, provided that the straight-time hourly wage rate is flexible *in each period*. However, to the extent that variability in hours of work is not always accompanied by changes in the straight-time wage, the adjustment mecha-

[5]The seeds of this idea can be traced back as far as Paul H. Douglas and Joseph Hackman (1939 pp. 34–5). In addition, Ehrenberg and Schumann (1982b pp. 36–8) recognize that straight-time wage differentials could arise to offset the statutory overtime premium, but they do not argue that these wage differentials would necessarily be perfectly offsetting. They also suggest that compensating adjustments in fringe benefits could occur, which would accentuate the effects of the overtime premium.

nism specified in the simple fixed-job model must be at best incomplete.

These issues led me to consider an extended version of the fixed-job model in which firms are allowed to vary the work week after observing the state of demand but in which employment and the straight-time wage must be chosen *ex ante*. Within the context of this extended model, increases in the statutory overtime premium can produce nonzero effects on hours and employment, but the direction of these effects is theoretically indeterminate. Moreover, if workers can borrow and save to transfer earnings across states of nature, the neutrality result of the simple static fixed-job model will tend to hold even in this more complicated setting. Intuitively, the ability to borrow and save means that workers do not care about the variance in weekly earnings so long as the discounted value of their total pay equals the correct level over a longer time period, such as the year.

From a policy perspective, it may be desirable to focus on regularly scheduled overtime, since these are the hours of overtime that are easiest to convert into new jobs. In addition, static versions of the fixed-wage and fixed-job models yield sharper empirical implications. Therefore, the empirical tests conducted below concentrate on the "usual" work week and assume that a large fraction of overtime persists from week to week. The U.S. Department of Labor (1985) publishes hours data that can shed some light on the appropriateness of this assumption. Average weekly overtime hours of production workers are available on a monthly basis for manufacturing industries. For the period 1974–1978, which includes the years to be analyzed, overtime per worker averaged 3.2 hours per week for all of manufacturing. Over these 60 monthly observations, average weekly overtime ranged from a low of 2.2 hours to a high of 3.9 hours, with a standard deviation of 0.4 hours. Overtime variability is similar within more disaggregated manufacturing industries with two-, three-, and four-digit SIC codes. Thus, these numbers are consistent with the notion that a large part of overtime is regularly scheduled, but it must be cautioned that they are averages which mask variation across workers within a firm and across firms within an industry.

III. Data

The data analyzed below come from the 1974, 1976, and 1978 May Current Population Surveys (CPS).[6] These May versions of the CPS are especially useful for the present study because they contain information on whether or not individuals receive an overtime premium for weekly hours of work in excess of 40, in addition to data on straight-time hourly wages, weekly earnings and hours, and demographic characteristics. This section describes general features of these data, especially with regard to the classification of workers according to coverage status under the minimum-wage and overtime pay provisions of the FLSA.

The empirical analysis requires knowledge of whether or not a given worker falls under the jurisdiction of the overtime pay provisions of the FLSA. I first exclude unemployed or self-employed individuals and multiple jobholders and then classify remaining individuals into three groups: those believed to be subject to the overtime pay provisions of the FLSA (covered workers), those believed not to be subject (noncovered workers), and those workers whose coverage status could not be ascertained from the information available in the May CPS. This final group is excluded from the analysis. The detailed industry and occupation classifications provided in the CPS were coded to determine the coverage status of each individual.[7] This method of assigning

[6]These years were chosen for study because hourly-wage questions were not asked in the CPS until 1973, and beginning in 1979 these questions were only asked of a quarter of the sample. Also, the CPS sampling scheme is such that May surveys in adjacent years have about half of their households in common.

[7]Details of the coding algorithm used to assign coverage status to each individual are available from the author. The algorithm does not consider independent state regulations that supplement the FLSA, but little is lost by this omission. Very few states regulate overtime, and state minimum wages seldom exceed the

coverage under the FLSA has been used previously by Ashenfelter and Robert S. Smith (1979 p. 341) and Brigitte H. Sellekaerts and Stephen W. Welch (1981 tables 3 and 4) for the minimum-wage provisions, and by Ehrenberg and Schumann (1982b appendix C) for the overtime pay provisions. Other sources found helpful in interpreting the FLSA and its many exemptions were U.S. Congress (1974) and Conrad. F. Fritsch (1981).

The largest group of exempt employees comprises executive, ~~administrative,~~ and professional ~~workers. To qualify~~ for this exemption the employee must be paid on a salary basis. Since the covered sector does not include a comparable group of workers, individuals employed in these types of occupations are excluded from the analysis. As a further safeguard against misclassifying these exempt supervisory employees. I restrict the sample to those individuals who report being paid by the hour. Outside sales-workers and employees paid on a commission basis are also exempt under the law, necessitating the exclusion of virtually all sales-workers because the coverage status of these individuals could not be determined. Most employees in retail trade and many in service industries have their coverage status determined by the size of their employer. Because this information is not available for the years of the CPS I analyze, these individuals are also excluded unless they are exempt for other reasons.

Workers classified as definitely *not* subject to the overtime pay provisions of the FLSA include essentially all nonsupervisory agricultural workers, as well as employees in the transportation sector who fall under the jurisdiction of other federal statutes such as the Motor Carriers Act of 1935, the Interstate Commerce Act, and the Railway Labor Act.[8] Remaining noncovered workers in my data come from miscellaneous exemptions that apply to certain workers in retail trade, services, and state and local government. Virtually all nonsupervisory employees in mining, construction, and manufacturing are subject to the overtime pay provisions of the FLSA, as are federal government workers not exempt for other reasons.

The 1974 and 1977 amendments to the FLSA require that the rules used for assigning coverage status be slightly different in each of the three years studied. The most important differences concern the coverage status of state and local government employees. The 1974 amendments attempted to bring all public employees under the jurisdiction of the act, but the constitutionality of this extension was quickly challenged, and in June 1976 the Supreme Court ruled unenforceable all minimum-wage and overtime pay provisions of the FLSA as they apply to state and local public employees engaged in traditional governmental functions.[9] Therefore, most state and local government employees are classified as covered in 1974 and 1976 but noncovered in 1978.[10] Employees in police and fire protection constitute an exception to this statement, since these workers never were subject to the overtime pay provisions of the FLSA.

The empirical analysis also requires identifications of those individuals covered by

federal minimum and generally only apply to those workers already excluded from the current sample because their coverage status could not be determined. See Aline O. Quester (1981) for a discussion of state minimum-wage laws.

[8]None of these other federal statutes requires payment of an overtime premium, although they do sometimes restrict hours of work. For example, certain truck drivers are prohibited from working more than 70 hours in any period of eight consecutive days.

[9]This ruling was made in *National League of Cities* v. *Usery*, 426 U.S. 833. With its decision in *Garcia* v. *San Antonio Metropolitan Transit Authority et al.*, 105 S.Ct. 1005 (February 1985), the Supreme Court recently overruled its earlier decision and upheld the constitutionality of applying the FLSA to state and local government employees, but this ruling occurred well after the time period being studied.

[10]Except for the analysis of hours-of-work distributions to be discussed in Section V, none of the empirical results reported here is sensitive to the coverage status of state and local government employees. Similar results are obtained when these workers are excluded from the sample or if it is assumed that, because of the confusion surrounding the ensuing litigation, state and local governments never acted as if they were subject to the FLSA during the time period under study.

TABLE 1—VARIABLE DEFINITIONS

Variable	Definition
OTPAY	1 if received premium pay for weekly hours worked over 40; 0 otherwise
HOURS40	1 if worked exactly 40 hours during survey week, 0 otherwise
WAGE	straight-time hourly wage rate in dollars
EARNINGS	usual weekly earnings in dollars
HOURS	usual weekly hours of work
OTCOVER	1 if subject to the overtime pay provisions of the FLSA; 0 otherwise
MINWAGE	1 if subject to and earning the legal minimum wage; 0 otherwise
OTCOVG40	1 if OTCOVER = 1 and HOURS > 40; 0 otherwise
UNION	1 if member of labor union or employee association; 0 otherwise
AGE	years of age
EDUC	completed years of schooling
MARRIED	1 if married, spouse present; 0 otherwise
FEMALE	1 if female; 0 otherwise
BLACK	1 if black; 0 otherwise
NONWHITE	1 if nonwhite other than black; 0 otherwise
CITY	1 if reside in central city of SMSA; 0 otherwise
SUBURB	1 if reside in SMSA, but not in central city; 0 otherwise
GOVTEMP	1 if government employee; 0 otherwise

and earning the legal minimum wage. Since exemptions to the minimum-wage law constitute a subset of the overtime exemptions, it turns out that coverage under the minimum-wage law includes all those covered by the overtime law and some additional workers as well. In 1978, a single minimum wage of $2.65 per hour applied to all covered workers, so in this year determining which are minimum-wage workers is straightforward. In 1974 and 1976, however, separate minimum wages applied to agricultural workers, workers covered prior to the 1966 amendments, and workers newly covered by the 1966 and 1974 amendments. Workers in these years are first assigned to the proper minimum-wage category,[11] and then minimum-wage status is determined by comparing the individual's actual wage with the appropriate minimum wage. Minimum-wage rates in 1974 were $1.60 per hour for agricultural workers, $1.90 per hour for newly covered workers, and $2.00 per hour for workers covered prior to 1966. These same minimum-wage rates in 1976 were $2.00, $2.20, and $2.30 per hour, respectively.

Table 1 defines the main variables used in the empirical analysis. To give an idea how the covered and noncovered sectors differ, Table 2 presents means and standard deviations by status of overtime pay coverage for the May 1978 CPS sample employed in the earnings regressions reported below.[12] The data for 1974 and 1976 are similar except that, because of the change in coverage status for state and local government employees, overall coverage rates rise to just over 90 percent of the sample, and the

[11] Once again, by using the industry and occupation classifications provided in the CPS.

[12] Not all workers in the 1970 Census major occupation classifications "professional, technical, and kindred workers" and "managers and administrators, except farm" qualify for the FLSA exemption, which applies to executive, administrative, and professional employees (Sellekaerts and Welch, 1981 table 3). This is why the major occupation categories "professional" and "administrative" have nonzero means in Table 2. Examples of professional and administrative occupations that do not qualify for this exemption include computer programmers, technicians, surveyors, and railroad conductors.

TABLE 2—DESCRIPTIVE STATISTICS, 1978 EARNINGS SAMPLE

Variable	All Mean	All SD	Covered Mean	Covered SD	Noncovered Mean	Noncovered SD
EARNINGS	263.372	109.941	267.439	108.354	239.869	115.985
WAGE	6.209	2.348	6.294	2.321	5.710	2.443
OTCOVER	0.853	0.355	1.000	0.000	0.000	0.000
OTCOVG40	0.198	0.398	0.232	0.422	0.000	0.000
HOURS	41.888	6.291	41.947	5.984	41.552	7.828
AGE	35.276	12.693	35.274	12.594	35.290	13.258
EDUC	11.300	2.407	11.298	2.372	11.311	2.602
UNION	0.494	0.500	0.499	0.500	0.462	0.499
MARRIED	0.732	0.443	0.741	0.438	0.677	0.468
CITY	0.185	0.388	0.183	0.386	0.197	0.398
SURBURB	0.356	0.479	0.359	0.480	0.334	0.472
GOVTEMP	0.094	0.292	0.043	0.202	0.388	0.488
Major industry:						
Agriculture	0.022	0.146	0.006	0.075	0.114	0.318
Mining	0.034	0.182	0.040	0.196	0.000	0.000
Construction	0.178	0.382	0.208	0.406	0.001	0.030
Durable manufacturing	0.310	0.463	0.363	0.481	0.002	0.042
Nondurable manufacturing	0.146	0.353	0.171	0.376	0.000	0.000
Transportation and utilities	0.125	0.331	0.071	0.256	0.440	0.497
Wholesale trade	0.039	0.194	0.046	0.209	0.000	0.000
Retail trade	0.013	0.115	0.000	0.000	0.090	0.287
Finance, insurance, and real estate	0.009	0.097	0.11	0.105	0.000	0.000
Services	0.124	0.329	0.084	0.278	0.353	0.478
Major occupation:						
Professional	0.035	0.185	0.036	0.187	0.029	0.169
Administrative	0.001	0.034	0.000	0.000	0.008	0.089
Sales	0.001	0.030	0.000	0.000	0.006	0.079
Clerical	0.072	0.259	0.070	0.254	0.088	0.284
Crafts	0.386	0.487	0.410	0.492	0.251	0.434
Operatives	0.321	0.467	0.342	0.475	0.197	0.398
Laborers	0.110	0.313	0.106	0.308	0.131	0.338
Service workers	0.059	0.235	0.035	0.185	0.194	0.395
Farm workers	0.014	0.117	0.000	0.000	0.095	0.293
Sample size:	7,594		6,474		1,120	

Notes: The sample consists of white males aged 18–64 who usually work at least 20 hours per week and for whom coverage status could be determined. Because of missing values, the sample sizes for the WAGE variables are as follows: 7,222 for the full sample, with 6,166 in the covered sector and 1,056 in the noncovered sector.

fraction of noncovered workers who are government employees falls substantially.

IV. Overtime Pay Compliance

If the FLSA imposes real costs upon firms, some employers may choose to disregard the law. Incentives for noncompliance should vary directly with the cost of conforming to the law, so that investigating compliance with the FLSA can provide some information on whether the act imposes binding constraints upon firms. Actual monetary sanctions against first-time offenders of the FLSA are almost invariably substantially less than the amount of underpayment, and due to a limited enforcement budget the probability of detection is probably quite low, so it is not surprising that previous studies have uncovered significant amounts of noncompliance with both the minimum-wage and overtime pay provisions of the FLSA.[13] This section reexamines compliance with the overtime law in order to test a prediction of the fixed-job model.

The fixed-job model implies that overtime pay compliance should be lower for minimum-wage workers than for other workers. Incentives for noncompliance should be strongest in those sectors where the law has real economic effects, and the fixed-job model suggests that overtime pay regulation only affects those workers with straight-time hourly wage rates that cannot be fully adjusted to offset the law. This type of wage rigidity exists among workers covered by and earning the legal minimum wage. If this is the sole source of wage rigidity, then the fixed-job model implies that noncompliance should occur only with minimum-wage workers. For all other workers, adjustment of straight-time hourly wages provides a virtually costless means of complying with the overtime law, so employers of these workers would never risk having to pay a penalty for noncompliance. As a practical matter, however, ignorance and misinterpretation of coverage status under the law provide reasons for less than perfect compliance among workers earning more than the minimum wage. Furthermore, in the empirical work, it is inevitable that some workers have been mistakenly classified as being covered by overtime pay regulation, and there may also exist reporting error by CPS respondents as to their receipt of premium pay for overtime. For these reasons, the hypothesis to be tested is whether the overtime pay compliance rate among minimum-wage workers is significantly lower than that among other workers. Note that the fixed-wage model provides no justification for such a discontinuity in the relationship between overtime pay compliance and hourly wage rates.

The May CPS data described above provide the necessary information for examining the correlation between minimum-wage status and compliance with the overtime pay provisions of the FLSA. These surveys ask those individuals who worked more than 40 hours during the survey week for a single employer whether or not they received a higher rate of pay for weekly hours of work above 40. Table 3 reports, for covered workers who put in overtime during the survey week, the percentage of these workers in various wage intervals who received premium pay for overtime in compliance with the FLSA.[14]

[13]See Ashenfelter and Smith (1979) and Sellekaerts and Welch (1983, 1984) for evidence on noncompliance with the minimum wage and for discussions of the penalty and enforcement mechanisms that have been established to discourage noncompliance. Ehrenberg and Schumann (1982a) and Sellekaerts and Welch (1983) provide similar information concerning the overtime pay provisions.

[14]The minimum wage is assumed to be binding for those workers covered by the minimum-wage law who earn exactly the minimum wage. If the legal minimum wage coincidentally happens to be the equilibrium wage for some workers, so that the minimum wage is not binding for these minimum-wage workers, then the results reported in Table 3 are biased against the prediction of the fixed-job model. Note also that the numbers in Table 3 imply that some employers choose to comply with the minimum-wage provisions of the FLSA while at the same time violating the overtime pay provisions. Because the Wage and Hour Division of the Employment Standards Administration in the Department of Labor enforces both sets of provisions and therefore any inspection by enforcement agents is likely to uncover *all* violations of the FLSA, it might seem that rational employer behavior would dictate a package decision either to comply fully with the FLSA or else violate both sets of provisions simultaneously. However, most inspections are precipitated by employee complaints to the Wage and Hour Division (Ashenfelter and Smith, 1979), so employers might find it optimal to comply with one set of provisions and not the other if the provisions differ in the likelihood that noncompliance would induce an employee complaint. The overtime pay provisions are much less publicized and contain more numerous and complicated exemptions than the minimum-wage provisions, and these factors could result in employees being less likely to recognize and report employer noncompliance with overtime pay regulation.

Table 3 excludes the fewer than 15 workers in each year who earn less than the minimum wage. Although minimum-wage noncompliance is much larger than this (Ashenfelter and Smith, 1979), the subsample considered here misses many subminimum-wage workers because overtime pay data is only available for those who worked overtime during the survey week and also because subminimum-wage workers are often employed in industries that could not be classified as definitely subject to the overtime pay provisions of the FLSA.

TABLE 3—RATES OF OVERTIME PAY COMPLIANCE FOR COVERED WORKERS

	1974		1976		1978	
Hourly wage	Percentage compliance	Sample size	Percentage compliance	Sample size	Percentage compliance	Sample size
Minimum wage	67.2	58	77.4	53	66.0	53
Minimum wage–$3.00	82.3	541	81.4	338	85.0	180
$3.01–$4.00	93.4	683	88.7	558	87.5	528
$4.01–$5.00	94.5	598	90.4	510	92.1	520
$5.01–$6.00	95.2	395	93.7	396	92.0	427
$6.01–$7.00	93.1	131	94.5	290	93.4	422
$7.01 and above	91.7	84	89.1	221	92.8	755
All covered workers	90.8	2,490	89.3	2,366	91.0	2,885

Notes: The samples consist of covered employees who worked more than 40 hours during the survey week. The basic minimum hourly wages in the three years were $2.00 in 1974, $2.30 in 1976, and $2.65 in 1978.

The overall compliance rates of around 90 percent are very similar to what Ehrenberg and Schumann (1982a) report for the May 1978 CPS data. In all three years, rates of overtime pay compliance plunge for minimum-wage workers, even relative to the compliance rates of those workers earning just above the minimum wage, although the decline in 1976 is not as drastic as in the other years. In all three years, chi-square homogeneity tests performed on the cell frequencies overwhelmingly reject the hypothesis of identical compliance rates for minimum-wage and nonminimum-wage workers. Therefore, these compliance statistics provide strong support for the fixed-job model. However, the simple proportions presented in Table 3 do not control for other factors that might affect compliance with overtime pay regulation, and Ehrenberg and Schumann (1982a) have shown some of these other factors to be significantly correlated with compliance. Furthermore, it appears from the table that wage rates could be correlated with compliance, and it remains to be demonstrated that minimum-wage status has an effect on compliance separate from that of having a low wage.

To meet these objections, Table 4 presents compliance logits similar to the *non*-compliance probits reported by Ehrenberg and Schumann (1982a) for the May 1978 CPS data, except that I add a dummy variable identifying minimum-wage workers and replicate the analysis for the May 1974 and 1976 CPS data as well. The sample again consists of covered employees who worked more than 40 hours during the survey week, and the dependent variable is a dummy indicating whether or not the individual received premium pay for overtime.[15] The independent variables include the minimum-wage dummy of primary interest, a vector of socioeconomic characteristics of the individual, a set of dummies indicating major industry and major occupation, and a set of dummies indicating the census geographic division in which the worker resides. The industry, occupation, and geographic dummies are included to control for the possibility that government efforts to enforce the FLSA vary in intensity across sectors. The coefficients of these dummy variables are not reported. Standard errors are given in parentheses, and derivatives of the compliance probability with respect to the independent variables are also reported.

The results are consistent with the fixed-job model. The minimum-wage dummy has a negative coefficient in all three years, al-

[15]Table 4 includes the subminimum-wage workers excluded from Table 3, and this accounts for the slight differences in sample sizes across the tables. These subminimum-wage workers are classified as nonminimum-wage workers in the logits reported in Table 4, but the results do not change when these workers are excluded.

TABLE 4—LOGITS FOR OVERTIME PAY COMPLIANCE (DEPENDENT VARIABLE = OTPAY)

Variable	1974 Coefficient	1974 Derivative	1976 Coefficient	1976 Derivative	1978 Coefficient	1978 Derivative
MINWAGE	−0.6944 (0.3380)	−0.0597	−0.2979 (0.4075)	−0.0286	−1.0549 (0.3545)	−0.0886
WAGE	0.3567 (0.0970)	0.0307	0.4010 (0.1434)	0.0385	0.3590 (0.1438)	0.0302
WAGE²/100	−0.7379 (0.2467)	−0.0635	−2.5807 (1.0085)	−0.2477	−1.8054 (0.9733)	−0.1517
AGE	0.0148 (0.0066)	0.0013	−0.0050 (0.0061)	−0.0005	−0.0134 (0.0059)	−0.0011
EDUC	0.1060 (0.0332)	0.0091	0.0053 (0.0328)	0.0005	−0.0074 (0.0320)	−0.0006
UNION	0.5972 (0.2080)	0.0514	1.0742 (0.2125)	0.1031	0.8265 (0.1977)	0.0694
MARRIED	−0.1450 (0.1768)	−0.0125	0.0138 (0.1709)	0.0013	−0.0907 (0.1587)	−0.0076
FEMALE	−0.1042 (0.2224)	−0.0090	0.2487 (0.2172)	0.0239	0.0618 (0.2018)	0.0052
BLACK	−0.5305 (0.2689)	−0.0456	−0.5858 (0.2605)	−0.0562	−0.2307 (0.2542)	−0.0194
NONWHITE	−0.1487 (0.8622)	−0.0128	−0.1169 (0.6914)	−0.0112	0.3911 (0.7787)	0.0329
CITY	0.2874 (0.2053)	0.0247	−0.2074 (0.1997)	−0.0199	0.2299 (0.1972)	0.0193
SUBURB	0.0810 (0.1839)	0.0070	−0.1588 (0.1828)	−0.0152	0.1562 (0.1729)	0.0131
GOVTEMP	−0.3039 (0.2725)	−0.0261	0.1535 (0.2755)	0.0147	−0.5965 (0.3783)	−0.0501
Log likelihood:	−650.6		−683.7		−789.7	
Sample size:	2,504		2,379		2,895	

Notes: Standard errors are given in parentheses. The samples consist of covered employees who worked more than 40 hours during the survey week. Also included as indepenuent variables were dummy variables indicating major industry, major occupation, and census geographic division. Derivatives of the compliance probability with respect to the independent variables were calculated using the overall compliance frequency of each sample: 0.905 in 1974, 0.892 in 1976, and 0.908 in 1978.

though this effect is not statistically significant in 1976.[16] Furthermore, the estimated effects of this variable are relatively large, especially in 1974 and 1978. Within the covered sector, minimum-wage workers are associated with 3–9 percentage points lower probability of being paid an overtime premium, even after controlling for the wage and other characteristics of the worker.

However, Table 3 suggests that the relationship between overtime pay compliance and hourly wage rates may not be monotonic, so it could be objected that the minimum-wage dummy is simply picking up some of this nonlinearity rather than a true minimum-wage effect. It is for this reason that the wage is entered as a quadratic in the logits reported in Table 4, and the estimates do imply an inverted U-shaped relationship between compliance and wages, but even

[16]Reducing the straight-time wage to the minimum wage may not represent complete adjustment for some workers earning slightly above the minimum wage, so the fixed-job model implies that incentives for noncompliance also exist for these nonminimum-wage workers. Consistent with this, Table 3 reveals that compliance is relatively low for workers earning close to but above the minimum wage. However, this effect disappears in compliance logits that are identical to those reported in Table 4 except for the addition of a dummy variable identifying workers earning up to 25 cents above the basic minimum wage. In two of the three years, the coefficient on this dummy variable was not close to being statistically significant, and its inclusion did not affect the estimated coefficients of the minimum-wage dummy.

with this specification the minimum-wage dummy has important effects. As an additional specification check of this type, logits were estimated including another dummy variable identifying workers earning the arbitrarily chosen wage of \$3.00 per hour. The \$3.00-per-hour wage dummy was found to be statistically insignificant, and its inclusion did not change the coefficients of the minimum-wage dummy, indicating that the estimated minimum-wage effect is unlikely to be an artifact arising entirely from misspecification or from the similarity between minimum-wage workers and other low-wage workers.[17]

Another possibility is that the minimum-wage dummy captures nothing more than a firm-size effect, since small establishments employ disproportionate numbers of minimum-wage workers. To investigate this issue, County Business Patterns data on the average number of employees per establishment for two-digit industry classifications (U.S. Department of Commerce, 1978) were merged with the CPS data. The logits for overtime pay compliance were then reestimated with the average-establishment-size variable as an additional control. The establishment-size variable was not statistically significant, and its presence did not alter the coefficients of the minimum-wage dummy. Although this method of controlling for firm size is less than perfect, the results suggest that the compliance tests reported here are not seriously biased by the absence of micro-level data on employer size.

The estimated effects of the other variables on overtime pay compliance are qualitatively similar to those reported in the noncompliance probits of Ehrenberg and Schumann (1982a). Most notably, overtime pay compliance is considerably higher among union than nonunion workers. Under the fixed-wage model, in which overtime pay regulation imposes real costs upon firms and creates incentives for noncompliance, unions can be interpreted as policing organizations which increase the likelihood that noncompliance is recognized and reported to enforcement agencies. The fixed-job model, on the other hand, cannot easily explain why any variables other than minimum-wage status would be systematically related to overtime pay compliance, unless these variables are correlated with wage flexibility or the feasibility of making binding agreements over weekly earnings and hours packages.

V. Hours-of-Work Distributions

As discussed above, the fixed-wage model implies that the overtime pay provisions of the FLSA should increase the prevalence of a 40-hour work week. However, the well-known fact that the hours-of-work distribution has a spike at 40 hours does not by itself constitute supporting evidence for the fixed-wage model, since it is shown below that this spike occurs even among workers not subject to overtime pay regulation. Other factors obviously have much to do with the popularity of a 40-hour work week, so corroboration of the fixed-wage model requires finding a relatively larger spike in the covered sector.

In the simple fixed-wage model examined here, overtime pay regulation has no effect on those with work weeks below the statutory hours standard; therefore, the hypothesis to be tested is that, of those who work at least 40 hours during a given week, the

[17]The \$3.00-per-hour wage dummy remained statistically insignificant in logits that excluded the minimum-wage dummy. It still might be argued that minimum-wage workers in general are less likely to receive overtime pay, and so the pattern of overtime pay detected in the covered sector would also be found among noncovered workers. To test this conjecture, I reestimated the overtime pay logits including noncovered as well as covered workers and adding dummy variables to identify covered workers, minimum-wage workers, and covered minimum-wage workers separately. The results suggest that, as implied by the fixed-job model, being simultaneously covered by overtime pay regulation and at the minimum wage has a negative effect on the probability of receiving overtime pay which operates independently from either coverage status or minimum-wage status considered alone. This conclusion requires caution, however, because the data contain few noncovered minimum-wage workers, and therefore it is difficult to identify separately the effect of minimum-wage status for covered and noncovered workers.

probability of working exactly 40 hours should be higher for covered than for noncovered employees. Note that the fixed-job model, on the other hand, predicts that such an increased likelihood of working a 40-hour week should be found only for those covered employees earning the legal minimum wage, since, for all other employees, compensating adjustments in the straight-time hourly wage can take place to circumvent the overtime law.

Simple frequencies reveal that, although both the covered and noncovered sectors display a marked tendency for work weeks to mass at 40 hours, the tendency may be more pronounced in the covered sector. This lends some support to the fixed-wage model. For example, of those in the 1974 data who worked at least 40 hours during the survey week,[18] 68 percent of covered employees worked exactly 40 hours, whereas only 55 percent of noncovered employees worked a 40-hour week. The corresponding numbers for the 1976 data are 70 percent of covered employees and 55 percent of noncovered employees. For 1978, the situation is reversed, with a greater propensity to work a 40-hour week in the noncovered sector (71 percent) than in the covered sector (65 percent). However, this reversal arises primarily because of the switch in the coverage status of state and local government employees. If these workers are excluded, then the pattern in 1978 is similar to the earlier years.

Lending support to the fixed-job model, however, is the fact that in each year the proportion working 40-hour work weeks (rather than working overtime) is around 80 percent for covered minimum-wage workers, which is considerably higher than for nonminimum-wage workers. Although suggestive, these raw frequencies are not conclusive, because they do not control for other factors influencing the adoption of a 40-hour work week, and some of these other factors are likely to be correlated with coverage status.

Therefore, to provide a more convincing test of the implications just discussed, the logits reported in Table 5 were estimated. The sample now consists of all individuals (both covered and noncovered) who worked at least 40 hours during the survey week, and the dependent variable is a dummy indicating those individuals who worked exactly 40 hours. The independent variables include those used in the compliance regressions reported in Table 4, with the addition of dummy variables identifying individuals covered by the overtime pay provisions of the FLSA (OTCOVER) and covered minimum-wage workers (OTCOVER × MINWAGE). As before, the coefficients of the industry, occupation, and geographic dummies are not reported.

The estimates for 1974 and 1976 indicate that covered employees are much more likely to work a 40-hour week, even when other determinants of work-week length are held constant. These results corroborate the fixed-wage model and contradict the fixed-job model. On the other hand, the coefficient of OTCOVER is negative and statistically insignificant in 1978, which is consistent with the fixed-job model.[19]

Additional support for the fixed-job model is provided by the positive and relatively large coefficients estimated for the interaction term between OTCOVER and MINWAGE. This effect, which is somewhat smaller and less statistically significant in 1978, suggests that the positive influence of coverage status on the probability of working a 40-hour week is larger for minimum-wage than for nonminimum-wage

[18] May versions of the CPS provide information both on usual weekly hours of work and on hours actually worked during the survey week. Because respondents may be predisposed toward giving a standard answer such as 40 hours to the question about usual weekly hours of work, the results reported here use the data on hours worked during the survey week. However, the analysis was replicated using the alternative measure of weekly hours, and similar results were obtained.

[19] When state and local government employees are excluded, the coefficient of OTCOVER becomes positive but remains statistically insignificant in 1978, while the coefficients for 1974 and 1976 do not change. This sample exclusion does not affect the estimated impact of the interaction term between OTCOVER and MINWAGE.

TABLE 5—HOURS-DISTRIBUTION LOGITS (DEPENDENT VARIABLE = HOURS40)

	1974		1976		1978	
Variable	Coefficient	Derivative	Coefficient	Derivative	Coefficient	Derivative
OTCOVER	0.5728 (0.1350)	0.1257	0.6992 (0.1289)	0.1496	−0.1436 (0.0977)	−0.0322
MINWAGE	−1.1577 (0.8420)	−0.2540	−1.3064 (0.6643)	−0.2794	−0.2028 (0.4107)	−0.0454
OTCOVER × MINWAGE	1.5538 (0.8540)	0.3409	1.5260 (0.6843)	0.3264	0.7104 (0.4371)	0.1592
WAGE	0.0120 (0.0174)	0.0026	−0.0132 (0.0116)	−0.0028	−0.0449 (0.0128)	−0.0101
AGE	0.0029 (0.0020)	0.0006	0.0007 (0.0020)	0.0001	0.0038 (0.0019)	0.0009
EDUC	−0.0257 (0.0110)	−0.0056	−0.0212 (0.0108)	−0.0045	−0.0270 (0.0104)	−0.0061
UNION	0.3562 (0.0570)	0.0781	0.3865 (0.0568)	0.0827	0.4101 (0.0536)	0.0919
MARRIED	−0.2156 (0.0580)	−0.0473	−0.2161 (0.0571)	−0.0462	−0.1446 (0.0506)	−0.0324
FEMALE	0.9715 (0.0689)	0.2013	0.7779 (0.0667)	0.1664	0.7723 (0.0610)	0.1731
BLACK	0.5707 (0.0924)	0.1252	0.4586 (0.0869)	0.0981	0.2681 (0.0798)	0.0601
NONWHITE	0.3300 (0.2718)	0.0724	0.1070 (0.2184)	0.0229	0.7804 (0.1833)	0.1749
CITY	0.1128 (0.0647)	0.0247	0.1746 (0.0647)	0.0373	0.0807 (0.0609)	0.0181
SUBURB	0.0428 (0.0585)	0.0094	0.1031 (0.0583)	0.0221	0.0120 (0.0540)	0.0027
GOVTEMP	0.6721 (0.1104)	0.1474	0.5286 (0.1058)	0.1131	0.5180 (0.1097)	0.1161
Log likelihood:	−5,267.1		−5,291.3		−6,289.0	
Sample size:	9,006		9,175		10,470	

Notes: Standard errors are given in parentheses. The samples consist of employees who worked at least 40 hours during the survey week and for whom coverage status could be determined. Also included as independent variables were dummy variables indicating major industry, major occupation, and census geographic division. Derivatives of the probability of working exactly 40 hours with respect to the independent variables were calculated using the mean sample frequencies: 0.675 in 1974, 0.690 in 1976, and 0.661 in 1978.

employees. This additional effect for minimum-wage workers was confirmed in logits estimated on the sample of covered workers only. Of the remaining variables, it is apparent that workers who are less educated, unionized, unmarried, female, minority, employed by the government, or residing in a standard metropolitan statistical area (SMSA) are more likely than others to work the standard 40-hour week instead of a longer work week.

VI. Compensation

The fixed-job model presented in Section II asserts that straight-time hourly wages adjust to leave weekly earnings and hours unaffected by the overtime pay provisions of the FLSA. This section analyzes the wages and earnings of individual workers in order to determine whether such a straight-time wage differential exists and whether this differential is large enough to completely neutralize overtime pay regulation.

A. *Straight-Time Hourly Wage Rates*

The fixed-job model with perfect wage flexibility predicts weekly earnings to be equal for identical individuals working the *same weekly hours* at jobs that differ only in terms of whether or not they are subject to the overtime pay provisions of the FLSA. One can write this earnings equality as

$$(2) \quad wH = w^c H \qquad \text{if } H \le \bar{H}$$

$$(3) \quad wH = w^c\left[\bar{H} + b(H - \bar{H})\right] \qquad \text{if } H > \bar{H}$$

where w^c and w represent straight-time hourly wages in the covered and noncovered sectors, respectively, and the remaining variables are as defined in Section II. Note that, under the type of hedonic equilibrium envisioned by the fixed-job model, wages vary with hours of work, but for notational convenience the dependence of w on H has been suppressed above. Overtime pay regulation does not affect the wages of those who do not work overtime; therefore, equation (2) implies $w^c = w$ for identical workers in identical jobs with $H \le \bar{H}$. The above equations implicitly assume that overtime pay exists only to comply with the FLSA, and so no workers in the noncovered sector receive overtime pay, whereas all workers in the covered sector receive the minimum overtime pay required by law (i.e., an overtime premium of exactly b after exactly $\bar{H}$ weekly hours of work). The effects of relaxing this assumption will be reported below.

Using the fact that $\bar{H} + b(H - \bar{H}) = H + (b-1)(H - \bar{H})$, equation (3) can be rewritten as

$$(4) \quad w/w^c = 1 + (b-1)(H - \bar{H})/H \equiv V > 1 \qquad \text{if } H > \bar{H}.$$

V represents the straight-time wage premium necessary in the noncovered sector to produce equality in weekly earnings between this sector and the covered sector that receives overtime pay. The size of V depends on the length of the work week (H) as well as on the parameters of overtime pay regulation (b and $\bar{H}$), and it is intuitively clear why V must be increasing in H and b and decreasing in $\bar{H}$. Taking logs of equation (4) yields

$$(5) \quad \log(w^c) \approx \log(w) - (b-1)(H - \bar{H})/H \qquad \text{if } H > \bar{H}$$

where I have used the approximation that

$$\log(V) \approx (b-1)(H - \bar{H})/H$$

since $(b-1)(H - \bar{H})/H$ is small. Therefore, in the presence of overtime pay regulation, the structure of straight-time hourly wages is given by equation (5) for those who work overtime, and by $w^c = w$ for those who do not work overtime.

To derive an estimating equation from (5), specify straight-time hourly wages in the absence of overtime pay as $\log(w_i) = \mathbf{X}_i\boldsymbol{\beta} + \varepsilon_i$, where $\mathbf{X}$ represents a vector of worker characteristics, $\boldsymbol{\beta}$ is a vector of parameters, ε is a random error term, and i indexes the individual. Then the foregoing discussion suggests estimating a log(wage) equation of the form

$$(6) \quad \log(\text{WAGE}_i) = \mathbf{X}_i\boldsymbol{\beta} + \alpha\left[(H_i - \bar{H})/H_i\right] \times \text{OTCOVG40}_i + \varepsilon_i$$

where WAGE is the straight-time hourly wage and OTCOVG40 is a dummy variable identifying those individuals who are both covered by overtime law and actually working overtime (i.e., OTCOVER = 1 *and* $H > \bar{H}$). The term $(H - \bar{H})/H$ represents overtime as a percentage of total weekly hours and is calculated for each individual using the CPS data on usual weekly hours of work and setting $\bar{H} = 40$ as required by the FLSA.[20] From equation (5), it is clear that the fixed-job model predicts $\alpha = -(b-1) = -\frac{1}{2}$ in the above regression, since the FLSA currently requires $b = 1.5$. The fixed-wage model, on the other hand, provides no rationale for why straight-time wages should be inversely related to coverage status and overtime hours in this way. In other words,

[20] Because the earnings measure used in later analysis refers to usual weekly earnings, the hours measure employed here is usual weekly hours of work, rather than hours worked during the survey week. In addition, this measure should best reflect an individual's typical work week, since reported hours during the survey week can be affected by sick leave, holidays, vacation days, etc.

estimates of $\alpha = -\frac{1}{2}$ would provide evidence of the completely offsetting adjustment of straight-time wages to overtime pay regulation predicted by the fixed-job model, whereas $\alpha = 0$ implies the absence of any such adjustment, as is assumed by the fixed-wage model.

Before proceeding with the empirical analysis, however, it is necessary to discuss issues related to the specification and estimation of equation (6). First, note that the wage equation to be estimated is neither a labor-demand nor labor-supply function, but rather the type of equilibrium wage–hours locus described in Lewis (1969) and Rosen (1974). Because hedonic equilibrium generally requires that wage rates vary with hours of work, weekly hours enter this equation through the **X** vector as well as through the variable capturing the effects of overtime pay regulation. Therefore, the regressions reported below include usual weekly hours of work entering as a quadratic in the **X** vector.[21]

Other variables in the **X** vector are the human-capital measures and demographic characteristics typically included in wage equations. The human-capital variables are education, age, and age-squared. Dummy variables identifying marital status, union membership, and government employment are added because these variables are known to be correlated with wages. To control for regional cost-of-living differences, I include dummies indicating whether the individual resides in the central city of an SMSA or elsewhere in an SMSA, as well as a vector of dummies identifying the census geographic division of residence. Major industry and occupation dummies are included in an attempt to account for differences in the composition of the covered and noncovered sectors. Finally, I also introduce a dummy variable indicating coverage status under the overtime law (OTCOVER), thus allowing the intercept in the wage equation to differ across sectors. In effect, OTCOVER picks up straight-time wage differentials between covered and noncovered individuals who do *not* work overtime, and these wage differentials will be attributed to compositional differences between the covered and noncovered sectors. By contrast, the coefficient α in equation (6) measures straight-time wage differentials between overtime workers in the covered and noncovered sectors, and it is these wage differentials that provide a test of the fixed-job model.

The empirical work in Sections IV and V focused on the behavior of minimum-wage workers, and therefore it was necessary to keep the sample as large as possible in order to observe an adequate number of these workers. The empirical work in the present section, however, affords the luxury of restricting the sample to a more homogeneous group so that any wage or earnings differentials detected between covered and noncovered workers will not simply reflect crude demographic differences across the two sectors. Accordingly, all of the wage and earnings regressions reported here restrict the sample to white males aged 18–64 who usually work at least 20 hours per week.[22]

Table 6 presents ordinary least-squares estimates of equation (6).[23] Columns 1, 3,

[21] Experiments in which the **X** vector was generalized to include interactions between hours and other variables in the **X** vector produced similar results to the specification in which hours was not interacted, so only estimates of the simpler specification are presented here.

[22] I also exclude individuals covered by and earning the legal minimum wage, since these workers have their hourly wage rate fixed by law, and so the fixed-job model does not apply to them. Because they comprise a small fraction of the total sample, including these workers does not change the results. Regressions that included those who usually work fewer than 20 hours per week also produced similar estimates, as did regressions that included women and nonwhites.

[23] Because only estimates of the market wage-hours locus are sought, and not estimates of the underlying structure of worker preferences and firm technology, the empirical analysis presented here need not tackle the problems associated with structural estimation in implicit markets that have been described by James N. Brown and Harvey S. Rosen (1982) and Dennis Epple (1987). See Brown and Rosen (1987) for a recent study that also uses least squares on micro-data to estimate a hedonic wage locus. The hedonic locus estimated in

TABLE 6—HOURLY-WAGE REGRESSIONS [DEPENDENT VARIABLE = LOG(WAGE)]

Variable	1974 (1)	1974 (2)	1976 (3)	1976 (4)	1978 (5)	1978 (6)
$[(H-\bar{H})/H]$OTCOVG40	−0.2278 (0.0697)	−0.1903 (0.0704)	−0.1777 (0.0678)	−0.1281 (0.0688)	−0.2127 (0.0694)	−0.2448 (0.0701)
OTCOVER	—	−0.0638 (0.0173)	--	−0.0719 (0.0175)	—	0.0425 (0.0139)
HOURS	0.0161 (0.0031)	0.0166 (0.0031)	0.0180 (0.0027)	0.0181 (0.0027)	0.0138 (0.0027)	0.0134 (0.0027)
HOURS^2/100	−0.0155 (0.0034)	−0.0167 (0.0034)	−0.0148 (0.0028)	−0.0156 (0.0028)	−0.0092 (0.0029)	−0.0084 (0.0029)
AGE	0.0333 (0.0018)	0.0334 (0.0018)	0.0376 (0.0020)	0.0375 (0.0020)	0.0359 (0.0019)	0.0357 (0.0019)
AGE^2/100	−0.0370 (0.0023)	−0.0371 (0.0023)	−0.0413 (0.0024)	−0.0411 (0.0024)	−0.0386 (0.0023)	−0.0384 (0.0023)
EDUC	0.0231 (0.0014)	0.0230 (0.0014)	0.0266 (0.0015)	0.0266 (0.0015)	0.0287 (0.0014)	0.0286 (0.0014)
UNION	0.2273 (0.0072)	0.2265 (0.0072)	0.2738 (0.0076)	0.2731 (0.0076)	0.2812 (0.0071)	0.2807 (0.0071)
MARRIED	0.0856 (0.0087)	0.0856 (0.0087)	0.0817 (0.0089)	0.0815 (0.0088)	0.0741 (0.0081)	0.0739 (0.0081)
CITY	0.0724 (0.0090)	0.0710 (0.0090)	0.0541 (0.0093)	0.0536 (0.0093)	0.0074 (0.0092)	0.0074 (0.0092)
SUBURB	0.1071 (0.0077)	0.1056 (0.0077)	0.0882 (0.0081)	0.0870 (0.0081)	0.0541 (0.0078)	0.0542 (0.0078)
GOVTEMP	−0.0073 (0.0145)	−0.0021 (0.0146)	0.0116 (0.0141)	0.0166 (0.0141)	0.0317 (0.0157)	0.0482 (0.0166)
R^2:	0.4953	0.4964	0.5042	0.5055	0.5061	0.5067
Sample size:	6,407		6,521		7,357	

Notes: Standard errors are given in parentheses. The samples consists of white males aged 18–64 who usually work at least 20 hours per week and for whom coverage status could be determined. Also included as independent variables were dummy variables indicating major industry, major occupation, and census geographic division.

and 5 exclude the coverage dummy OTCOVER, whereas columns 2, 4, and 6 report results when this variable is included to help control for wage differences between the covered and noncovered sectors that are unrelated to whether overtime is worked. Estimates of α occupy the first row of this table. In all three years, the estimate of α is negative and statistically significant, providing evidence that straight-time wage differentials arise in response to overtime pay regulation. However, for all three years, one can also reject the hypothesis that $\alpha = -\frac{1}{2}$, which implies that the wage adjustment is less than complete. Including OTCOVER in the regressions causes the estimates of α to become less negative in 1974 and 1976, because in these years covered workers are found to earn lower wages independently of the effect of overtime pay regulation, whereas the opposite situation occurs in 1978.[24] However, in no year does α change by enough to overturn the previous conclusion that straight-time wages adjust significantly but incompletely to the presence of overtime pay regulation. These regressions support the existence of a partial wage adjustment that falls somewhere between the

that paper is between the mean level of the hourly wage and its variance.

[24] The sign reversal that occurs on the coefficient of OTCOVER in 1978 is related to the U.S. Supreme Court decision that caused many state and local government employees who were classified as covered in the 1974 and 1976 data to become noncovered in the 1978 data. See the discussion in Section III.

predictions of the fixed-wage and fixed-job models of overtime pay regulation.[25] The coefficients of the other variables generally appear reasonable in both sign and magnitude, although the estimated rate of return to education is quite low.

The empirical analysis conducted above required knowledge of the overtime pay coverage status of individual workers, as well as strong assumptions about the structure of overtime pay in the covered and noncovered sectors. It is therefore useful to discuss how sensitive the results are to these assumptions. As described in Section III, every effort was made to eliminate from the sample individuals whose coverage status under the overtime pay provisions of the FLSA could not be accurately assigned, but it is inevitable that some of the remaining individuals have been misclassified. If the fixed-job model is correct, classification errors of this type would bias estimates of α in the positive direction, since some covered workers with lower straight-time wages are being misclassified as otherwise identical noncovered workers, who would be expected to have higher straight-time wages, and vice versa. On the other hand, if the fixed-wage model best represents the truth, then no bias results from misclassification, since straight-time wages should be unrelated to coverage status. Therefore, any potential bias arising from errors in assigning coverage status works against the fixed-job model.

It was also assumed that no workers in the noncovered sector receive overtime pay, whereas all workers in the covered sector receive the minimum overtime pay required by law: an overtime premium of time and a half for weekly hours of work past 40. Although the assumptions that $b = 1.5$ and $\overline{H} = 40$ are probably not very objectionable, some union contracts specify an overtime premium of double time or above and a standard work week under 40 hours. Potentially more important assumptions are that overtime pay exists only to comply with the FLSA and that there exists perfect compliance within the covered sector. The compliance frequencies presented above indicated that about 10 percent of covered employees who worked overtime during the survey week were not paid an overtime premium as required by the FLSA, and many noncovered employees received overtime pay despite the lack of legal compulsion. To gauge the effect of these exceptions, the wage regressions were reestimated excluding union members from the sample and also excluding overtime workers in the covered sector who were not paid an overtime premium and those in the noncovered sector who were paid an overtime premium. These exclusions produced only minor changes in the estimates of α.

Finally, as noted above, weekly hours of work enter equation (6) through two avenues. The separate effect of overtime pay regulation is in large part identified by the fact that the reciprocal of weekly hours enters the term $[(H - \overline{H})/H]$OTCOVG40 but is excluded from the hedonic wage–hours locus of noncovered workers. The specific functional form of this term is a direct implication of the fixed-job model, but hedo-

[25]The extent to which straight-time wage differentials arise in response to the statutory overtime premium may vary across demographic groups, perhaps because wage flexibility or the feasibility of making binding agreements over weekly earnings and hours packages varies across sectors. In particular, the high-turnover jobs typically held by young, uneducated workers might not involve the implicit earnings–hours contracts envisioned by the fixed-job model. To investigate this possibility, wage regressions were estimated in which the variable capturing the effects of overtime pay regulation was interacted with a dummy variable identifying those workers 25 years of age or less without a high school diploma. Since only about 7 percent of the sample in each year meet this definition of "unskilled," it is not surprising that the estimates of α for the rest of the sample are very similar to the estimates in Table 6 that constrain α to be the same for skilled and unskilled workers. The point estimates of the interaction term for 1976 and 1978 imply that α is close to zero for unskilled workers in these years, which is consistent with the conjecture just discussed. However, in the 1974 regressions, unskilled workers receive a straight-time wage differential that is more than twice the size of that received by skilled workers, and in no year is this interaction term statistically significant. I also explored similar types of interactions with age, education, union status, and, for the extended sample, race and sex. These interactions generally failed to achieve statistical significance and did not exhibit consistent sign patterns across the three years.

nic theory provides little guidance as to the functional form of the wage–hours locus. Therefore, it could be argued that the reciprocal of weekly hours also belongs in the wage–hours locus of noncovered workers and that the estimates of α in Table 6 merely reflect this misspecification. When the reciprocal of weekly hours is added to the noncovered wage–hours locus, the estimates of α become statistically insignificant, ranging from -0.10 to 0.10 with standard errors of around 0.09. Thus, this alternative interpretation is also consistent with the data. On the other hand, the estimates of α reported in Table 6 make sense within the context of the fixed-job model, and there is no obvious reason for this to occur if the results are primarily due to a misspecified wage–hours locus.

B. *Weekly Earnings*

The usual-weekly-earnings data in the CPS provide an alternative way of testing for the effects of overtime pay regulation. The fixed-job model implies that the earnings–hours locus is unaffected by overtime-pay coverage status, whereas the fixed-wage model predicts that weekly earnings (given hours of work) should be higher for those covered individuals who work overtime and thereby qualify for the statutory overtime premium. In addition, the fixed-wage model implies that the earnings advantage enjoyed by covered workers should increase with the amount of overtime worked.

The implications of overtime pay regulation for the structure of weekly earnings can be derived in a fashion very similar to that used above for straight-time wages. For individuals who are both covered by overtime pay regulation and working overtime, the natural logarithm of weekly earnings can be written as

$$(7)\quad \log(Y^c) = \log(w^c) + \log(H) + (b-1)(H-\overline{H})/H \qquad \text{if } H > \overline{H}.$$

If the fixed-job model is correct, then $\log(w^c)$ adjusts to overtime pay regulation as specified in equation (5). Substituting (5) into (7), the term $(b-1)(H-\overline{H})/H$ vanishes, and therefore the fixed-job model predicts the statutory overtime premium to have no effect on weekly earnings. According to the fixed-wage model, on the other hand, straight-time wage rates are independent of overtime pay regulation, and so the earnings–hours locus is shifted outward for covered overtime workers.

Assuming as before that straight-time wages in the absence of overtime pay are given by $\log(w_i) = \mathbf{X}_i\boldsymbol{\beta} + \varepsilon_i$, one can test these implications of the fixed-job and fixed-wage models with a log(earnings) regression of the form

$$(8)\quad \log(\text{EARNINGS}_i) = \mathbf{Z}_i\boldsymbol{\gamma} + \delta\left[(H_i - \overline{H})/H_i\right] \times \text{OTCOVG40}_i + \varepsilon_i$$

where the vector $\mathbf{Z}$ includes $\log(H)$ in addition to all variables in the vector $\mathbf{X}$ described earlier and $\boldsymbol{\gamma}$ is a vector of coefficients. The fixed-job model implies $\delta = 0$, whereas the fixed-wage model predicts δ to be positive and on the order of $\frac{1}{2}$.

Table 7 reports estimates of equation (7).[26] Consistent with the fixed-job model, overtime pay regulation does not appear to have much effect on the earnings–hours locus. The estimates of δ never achieve statistical significance, are negative in 1974, and in the other years are always less than a third of the value predicted by the fixed-wage model.[27] The coverage dummy

[26]The weekly-earnings regressions in Table 7 are estimated on slightly larger samples than the hourly-wage regressions reported in Table 6 because of the fact that hourly-wage data are missing more often than weekly-earnings data. For the years being studied, reported weekly earnings of $1,000 or more are coded as $999 in the May CPS. Individuals for whom the usual-weekly-earnings variable has been truncated in this way are excluded from the sample.

[27]Unlike the straight-time wage regressions, estimates of the effects of overtime pay regulation from weekly-earnings regressions change very little when the reciprocal of weekly hours is added as a regressor.

TABLE 7—WEEKLY EARNINGS REGRESSIONS [DEPENDENT VARIABLE = LOG(EARNINGS)]

	1974		1976		1978	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
$[(H-\bar{H})/H]$OTCOVG40	−0.1202	−0.0325	0.0263	0.1254	0.1403	0.1132
	(0.0898)	(0.0920)	(0.0915)	(0.0946)	(0.0855)	(0.0879)
OTCOVER	—	−0.0756	—	−0.0733	—	0.0193
		(0.0175)		(0.0182)		(0.0146)
HOURS	−0.0217	−0.0393	−0.0440	−0.0619	−0.1453	−0.1404
	(0.0241)	(0.0245)	(0.0218)	(0.0222)	(0.0208)	(0.0211)
HOURS^2/100	0.0058	0.0142	0.0130	0.0214	0.0731	0.0709
	(0.0132)	(0.0133)	(0.0117)	(0.0118)	(0.0113)	(0.0115)
log(HOURS)	1.7675	2.1429	2.4167	2.7973	4.3824	4.2769
	(0.5022)	(0.5090)	(0.4550)	(0.4642)	(0.4328)	(0.4401)
AGE	0.0352	0.0352	0.0366	0.0364	0.0365	0.0364
	(0.0019)	(0.0019)	(0.0020)	(0.0020)	(0.0019)	(0.0019)
AGE^2/100	−0.0396	−0.0397	−0.0400	−0.0398	−0.0397	−0.0396
	(0.0023)	(0.0023)	(0.0025)	(0.0025)	(0.0024)	(0.0024)
EDUC	0.0221	0.0220	0.0257	0.0256	0.0289	0.0288
	(0.0015)	(0.0015)	(0.0015)	(0.0015)	(0.0015)	(0.0015)
UNION	0.2233	0.2222	0.2655	0.2644	0.2678	0.2677
	(0.0073)	(0.0073)	(0.0077)	(0.0077)	(0.0074)	(0.0074)
MARRIED	0.0809	0.0807	0.0862	0.0860	0.0818	0.0817
	(0.0088)	(0.0087)	(0.0091)	(0.0091)	(0.0085)	(0.0085)
CITY	0.0715	0.0697	0.0377	0.0374	0.0028	0.0028
	(0.0090)	(0.0090)	(0.0095)	(0.0095)	(0.0096)	(0.0096)
SUBURB	0.1079	0.1061	0.0875	0.0865	0.0574	0.0574
	(0.0078)	(0.0078)	(0.0083)	(0.0083)	(0.0081)	(0.0081)
GOVTEMP	−0.0098	−0.0046	0.0195	0.0245	0.0306	0.0380
	(0.0147)	(0.0147)	(0.0143)	(0.0143)	(0.0162)	(0.0171)
R^2:	0.5225	0.5238	0.5579	0.5590	0.5529	0.5530
Sample size:	6,854		6,774		7,594	

Notes: Standard errors are given in parentheses. The samples consist of white males aged 18–64 who usually work at least 20 hours per week and for whom converage status could be determined. Also included as independent variables were dummy variables indicating major industry, major occupation, and census geographic division.

OTCOVER is once again introduced to control for earnings differences between the covered and noncovered sectors that are unrelated to overtime, and OTCOVER picks up significant differences, but the same basic conclusion remains: overtime pay regulation does not significantly increase weekly earnings.

It is somewhat puzzling that the analyses of hourly wages and weekly earnings yield different estimates of the extent to which overtime pay regulation is circumvented. Straight-time wages appear to adjust in the direction implied by the fixed-job model, but by less than half as much as would be necessary to offset overtime pay regulation completely. The weekly-earnings regressions, on the other hand, suggest that adjustment is close to complete. There is evidence, however, that some respondents do not include overtime pay when answering the usual-weekly-earnings question in the May CPS, and this may explain why the regressions reported above find that coverage under overtime pay regulation has little effect on the weekly earnings of those who work overtime.[28] If the hourly-wage ques-

[28]Using the information provided on usual weekly hours, hourly-wage rates, and receipt of premium pay for overtime hours, I was able to calculate an alternative weekly-earnings measure for most individuals in the earnings sample. For overtime workers who receive premium pay, this alternative measure is on average 15–20 dollars higher than the response to the usual-weekly-earnings question, whereas the two measures are similar on average for those who do not work overtime or do not receive premium pay.

tion in the CPS provides accurate information on straight-time wage rates, then the regressions using this variable are likely to be more reliable than those using weekly earnings, and therefore it is probably prudent to favor the conclusion of a sizable but incomplete wage response to overtime pay regulation over the competing conclusion of full adjustment.

VII. Conclusion

This paper has analyzed the economic effects of the overtime pay provisions of the Fair Labor Standards Act. Two alternative approaches to modeling the impact of overtime pay regulation were shown to produce dramatically different implications. In particular, the labor-demand approach adopted by previous studies assumes that the straight-time wage is independent of the statutory overtime premium. This fixed-wage assumption implies that overtime pay regulation should reduce overtime hours, increase the prevalence of a 40-hour work week, and raise the weekly earnings of overtime workers. The fixed-job approach, on the other hand, assumes that firms and workers agree to packages of weekly hours and weekly compensation, and therefore changes in the statutory overtime premium can be neutralized by offsetting adjustments in the straight-time hourly wage. As a result, weekly hours and earnings are unaffected by overtime pay regulation. Because workers earning the minimum wage are legally precluded from experiencing reductions in their straight-time wage rates, the fixed-job model predicts overtime pay regulation to have a differential impact on minimum- and nonminimum-wage workers, with no substantive economic effects expected for nonminimum-wage workers.

The empirical analysis suggests that neither of the simple models of overtime pay regulation considered here can completely account for observed labor-market outcomes. In accordance with the fixed-job model, the overtime law appears to have a greater impact on minimum-wage workers, as reflected by the findings that overtime pay compliance is significantly lower for these workers and the prevalence of a 40-hour work week is significantly higher. On the other hand, overtime pay compliance is higher for union than for nonunion workers, a result more easily reconciled with the fixed-wage model of overtime pay regulation. In addition, the finding that overtime-pay coverage status systematically influences the hours-of-work distribution for nonminimum-wage workers is supportive of the fixed-wage model. No significant differences in weekly earnings were discovered between the covered and noncovered sectors, which is consistent with the fixed-job model. However, the analysis of hourly wages indicates a substantial but incomplete adjustment of straight-time wages to overtime pay regulation.

Overall, the data analyzed here suggest that wage adjustments occur to mitigate the purely demand-driven effects predicted by the fixed-wage model, but these adjustments are not large enough to neutralize overtime pay regulation completely. The offsetting adjustments that do occur serve to limit the potential usefulness of overtime pay regulation as a job-creation device.

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MEMORANDUM

COUNCIL OF ECONOMIC ADVISERS

Labor (Minimum wage)

April 1, 1996

MEMORANDUM FOR: JOSEPH STIGLITZ
MARTIN BAILY
ALICIA MUNNELL
TOM KANE
MICHAEL ASH

FROM: PETER ORSZAG PO

SUBJECT: Draft talking points on the minimum wage

Alicia asked me to put together some talking points on the minimum wage. Comments and corrections are welcome (as are suggestions for providing more detail on specific topics).

The value of the minimum wage

- The President has proposed to increase the minimum wage from its current level of $4.25 an hour to $5.15 an hour by July 1997. The increase will occur in two 45-cent steps: to $4.70 in July 1996, and to $5.15 in July 1997.

- The minimum wage has not been raised since April 1, 1991. On November 10, 1989, President Bush signed a 90-cent minimum wage raise. This increase, like the proposed one, took place in two 45-cent steps: from $3.35 to $3.80 on April 1, 1990, and from $3.80 to $4.25 on April 1, 1991.

- If the minimum wage is not increased within a year, it will fall to its lowest real level in 40 years. The attached table shows the value of the minimum wage in nominal dollars, real dollars, and as a percentage of the average private nonsupervisory wage. The Labor Department expects the real value of the minimum wage to fall below $3.92 (its 1955 level in 1995 dollars) in January 1997.

- The increase would raise a full-time, year-round minimum wage worker's annual income by $1,800 (90 cents over 2000 hours).

Workers directly affected by the increase

- In 1995, 9.7 million hourly-paid workers earned between $4.25 and $5.15 an hour.

    -- 59 percent of these workers are women

-- 69 percent are adults

-- According to the Economic Policy Institute, almost two-fifths are the sole wage-earners in their households

- These figures apply to those directly affected only. To the extent that a minimum wage increase shifts the wage distribution even for workers currently earning more than $5.15 an hour, the impact is broader. Research indicates that the "ripple" effect would affect at least the 3 million additional workers who earn between $5.15 and $5.75.

Current minimum wage workers

- The average share of household income earned by a minimum wage worker is one-half.

- In 1995, 2.0 million hourly-paid workers earned precisely $4.25 an hour; 3.7 million earned $4.25 or less (workers earning less than $4.25 an hour are either not covered by the $4.25 minimum wage or have incorrectly reported their wages or hours).

- The attached charts provide additional information on workers who were earning the minimum wage in 1993.

The employment debate

- There is an ongoing (and vitriolic) debate over the employment effects of an increase in the minimum wage. Roughly a dozen empirical studies published in peer-reviewed journals over the past 5 years find an insignificant effect (see attached list). While these results are not yet unanimously accepted in the academic community, there does seem to be a growing consensus that the employment effects are unlikely to be large. Robert Solow is often quoted as saying that the evidence "on job loss is weak. And the fact that the evidence is weak suggests that the impact on jobs is small."

Inflationary impact

- A WEFA study concluded that the President's proposal would raise the inflation rate during the transition period by less than 0.1 percentage point per year.

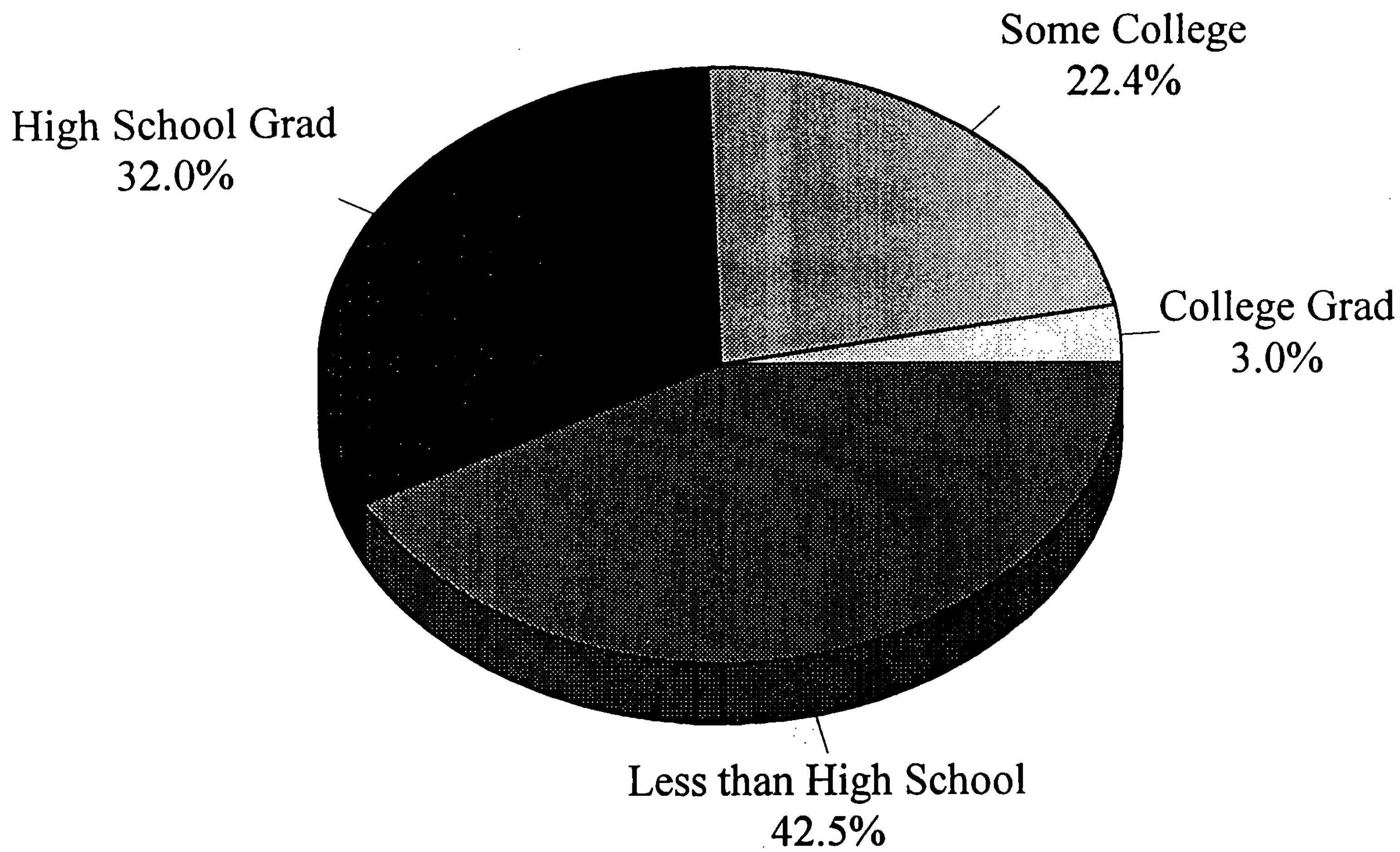
Value of the Minimum Wage, 1955-1995

Year	Value of the Minimum Wage, Nominal Dollars	Value of the Minimum Wage, 1995 Dollars	Minimum Wage as a Percent of the Average Private Nonsupervisory Wage
1955	$0.75	$3.92	43.9
1956	1.00	5.14	55.6
1957	1.00	4.99	52.9
1958	1.00	4.84	51.3
1959	1.00	4.81	49.5
1960	1.00	4.72	47.8
1961	1.15	5.38	53.7
1962	1.15	5.33	51.8
1963	1.25	5.71	54.8
1964	1.25	5.64	52.9
1965	1.25	5.56	50.9
1966	1.25	5.40	48.7
1967	1.40	5.86	52.3
1968	1.60	6.45	56.2
1969	1.60	6.17	52.6
1970	1.60	5.89	49.6
1971	1.60	5.64	46.4
1972	1.60	5.48	43.3
1973	1.60	5.15	40.6
1974	2.00	5.86	47.2
1975	2.10	5.68	46.3
1976	2.30	5.89	47.3
1977	2.30	5.53	43.9
1978	2.65	5.97	46.6
1979	2.90	5.96	47.1
1980	3.10	5.73	46.6
1981	3.35	5.65	46.2
1982	3.35	5.33	43.6
1983	3.35	5.11	41.8
1984	3.35	4.90	40.3
1985	3.35	4.73	39.1
1986	3.35	4.65	38.2
1987	3.35	4.48	37.3
1988	3.35	4.31	36.1
1989	3.35	4.11	34.7
1990	3.80	4.42	38.0
1991	4.25	4.74	41.2
1992	4.25	4.61	40.2
1993	4.25	4.47	39.2
1994	4.25	4.36	38.2
1995	4.25	4.25	37.1

Adjusted for inflation using the CPI-U-X1

Source: Center on Budget and Policy Priorities

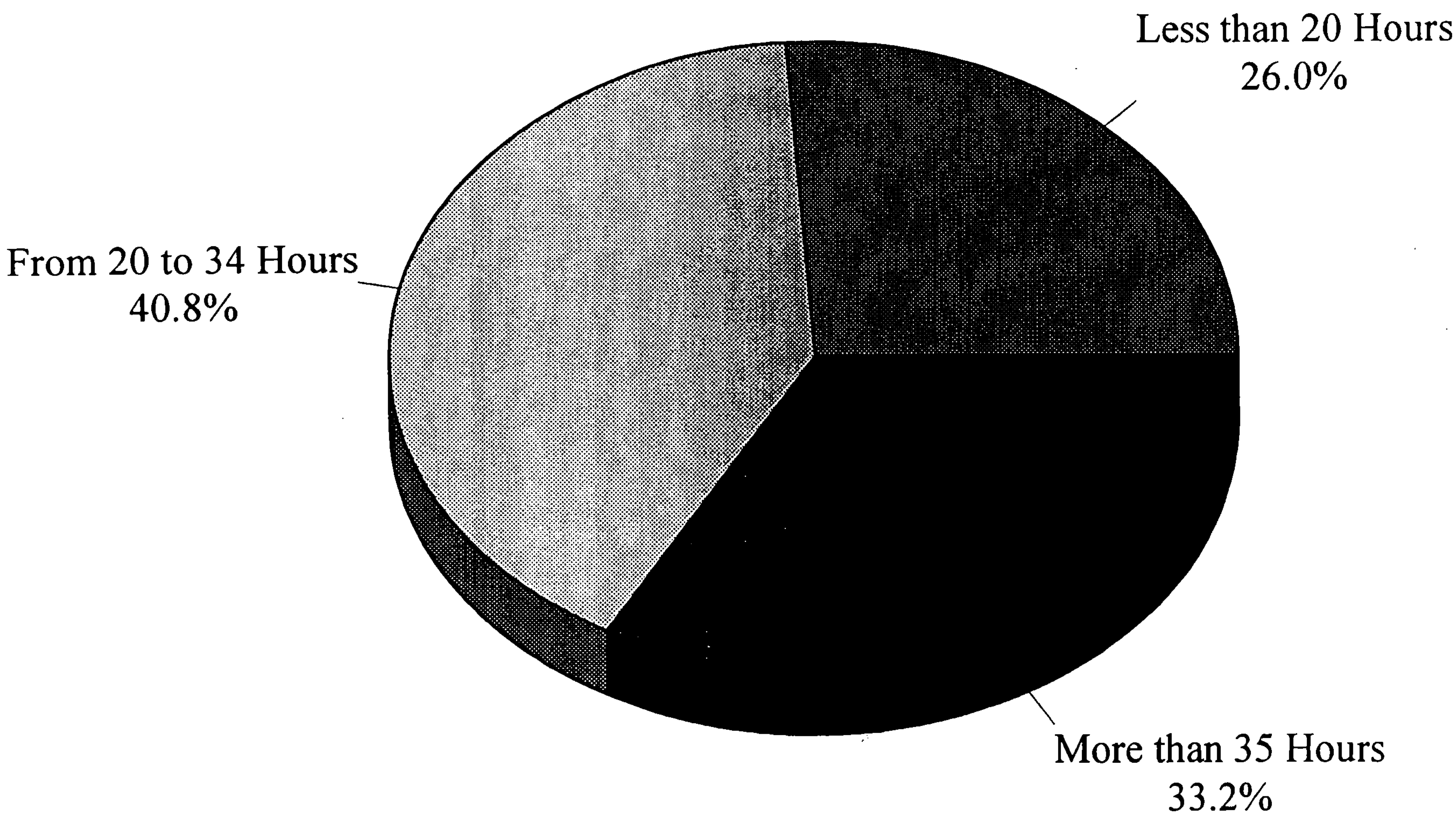
Distribution of Hourly-Paid Minimum Wage Earners by Education: 1993



Source: Bureau of Labor Statistics.

What Hours Do Minimum Wage Workers Work?

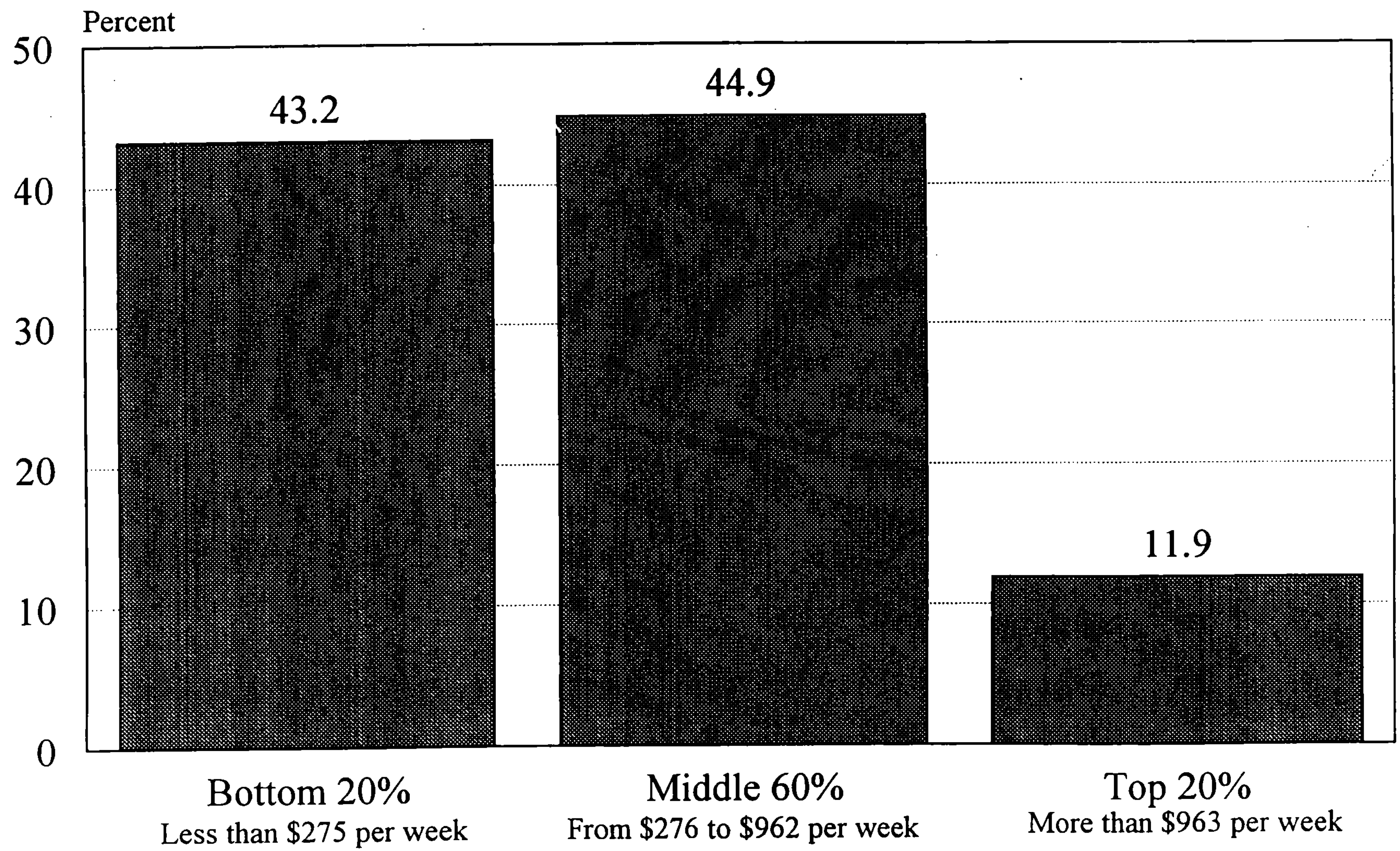
Distribution of Hourly-Paid Minimum Wage Earners by Hours Usually Worked per Week: 1993



Source: Bureau of Labor Statistics.

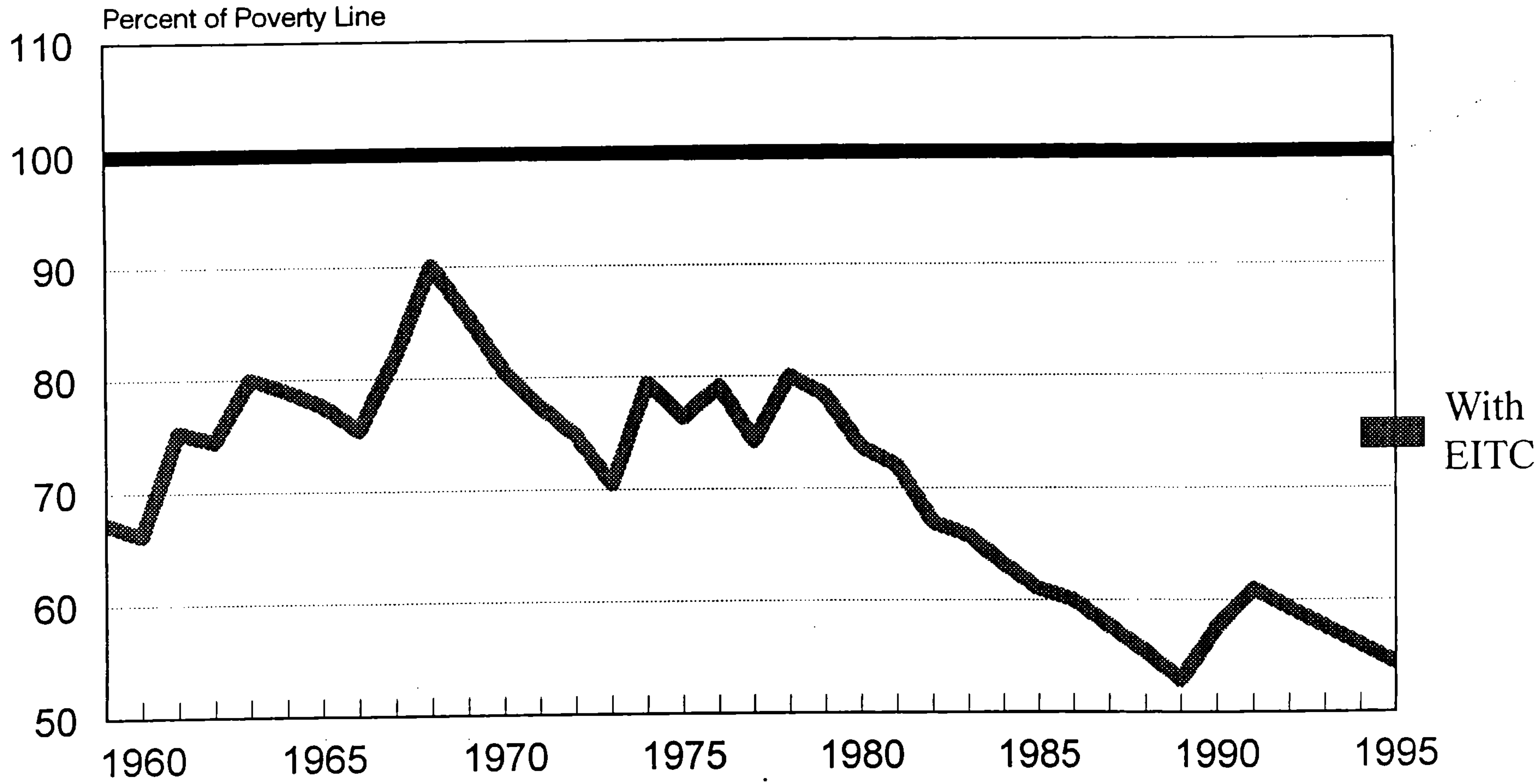
Who was Affected by the Last Minimum Wage Increase

Distribution by Family Earnings



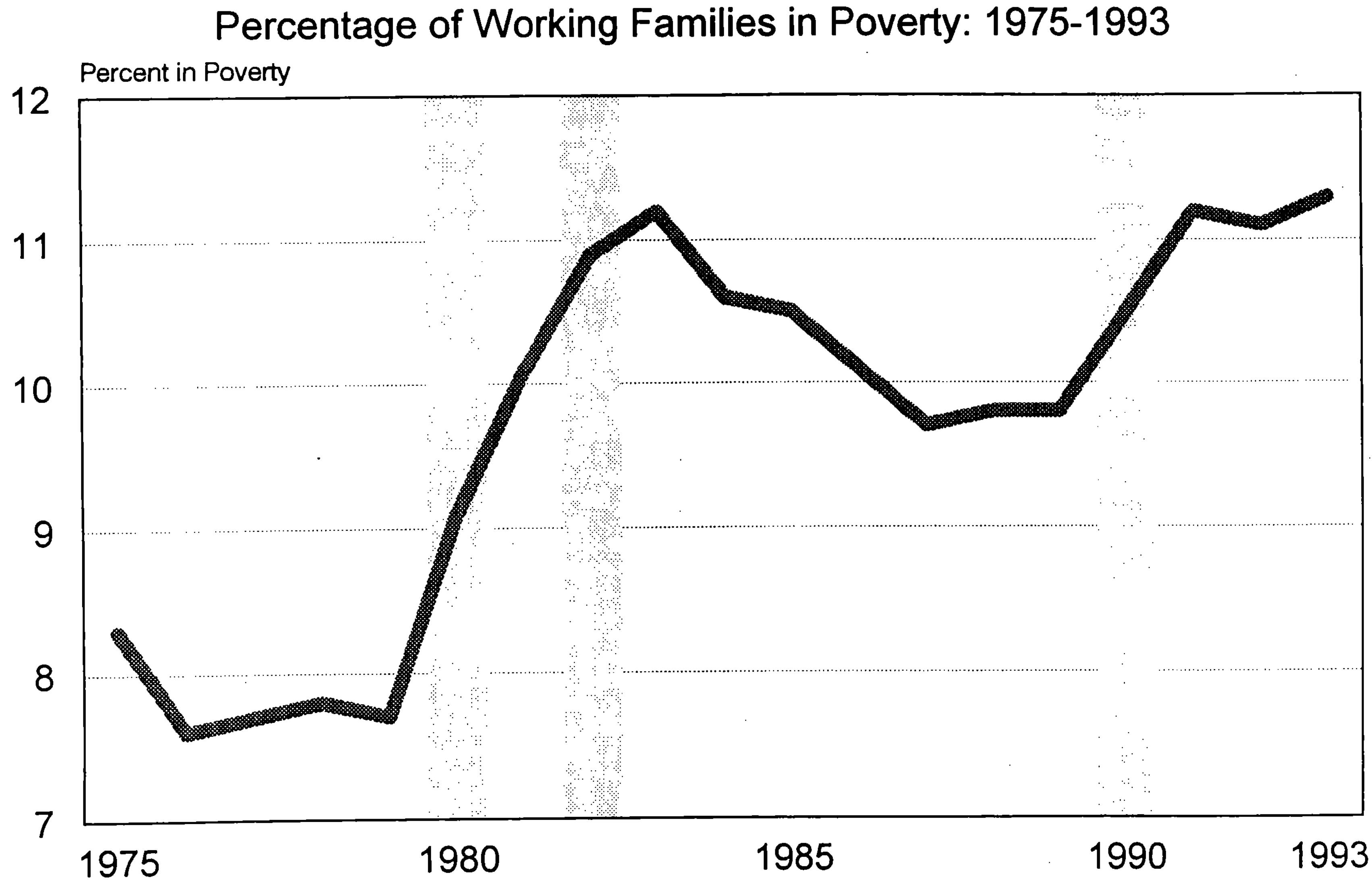
Minimum Wage Work Does Not Lift Families Out of Poverty

Annual Earnings at the Minimum Wage as a Percentage of the Poverty Line: 1959-1995



Note: Annual earnings for a family of four with one full-time, year-round minimum wage worker as a percentage of the four-person poverty line. The four-person poverty line for 1994 and 1995 are from Congressional Budget Office projections.

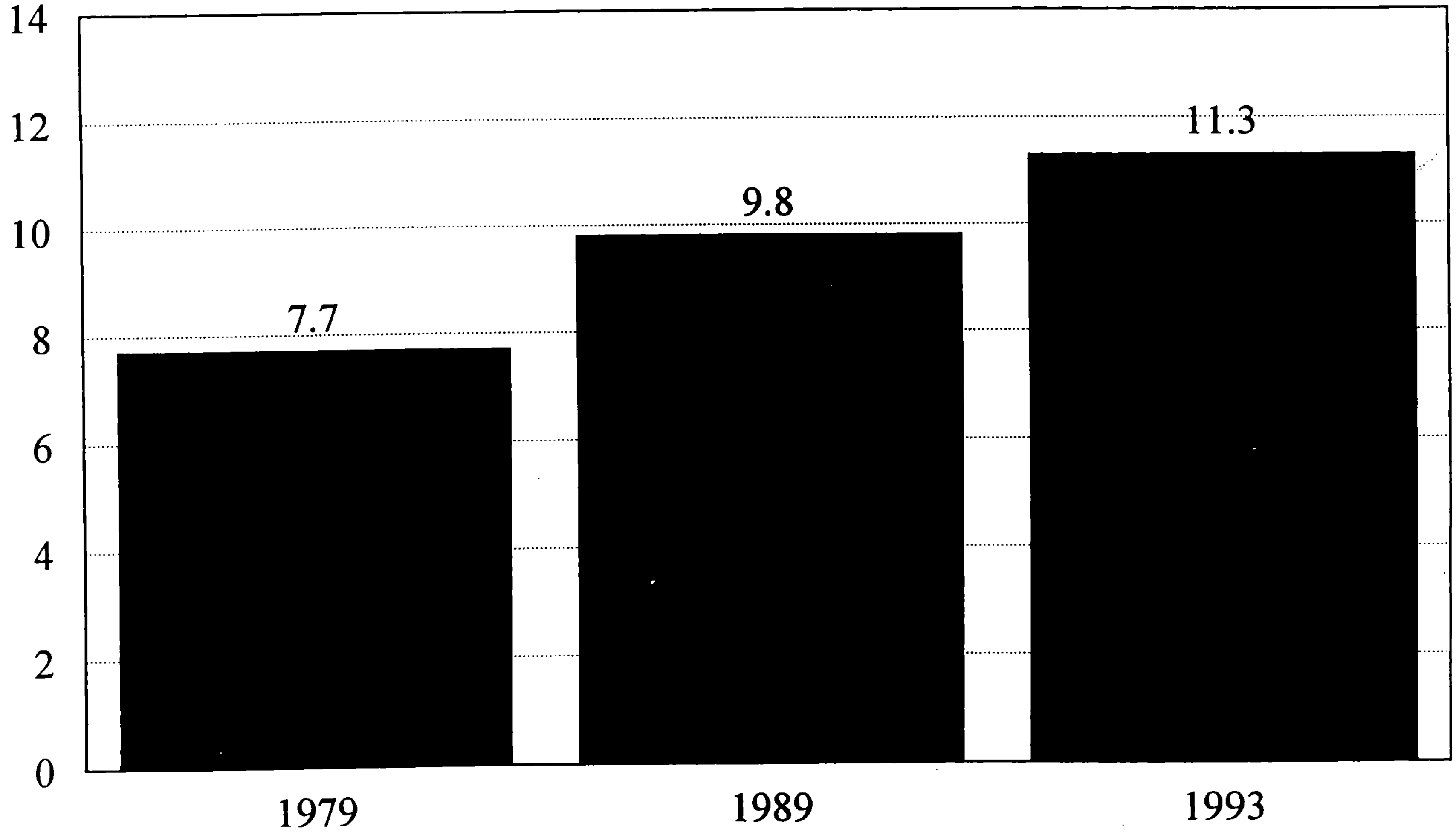
More Working Families are Poor



Source: Bureau of the Census, Current Population Survey. A working family is defined as one with children where someone in the household worked.

More Working Families are Poor

Percentage of Working Families in Poverty



Source: Bureau of the Census, Current Population Survey. A working family is defined as one with children where someone in the household worked.

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EXECUTIVE OFFICE OF THE PRESIDENT
COUNCIL OF ECONOMIC ADVISERS
WASHINGTON, D.C. 20500

SENIOR ECONOMIST

MEMORANDUM FOR	JOSEPH STIGLITZ MARTIN BAILY ALICIA MUNNELL
FROM	TOM KANE
RE	UPDATE ON IMPACTS OF THE MINIMUM WAGE
DATE	JANUARY 19, 1996

The purpose of this memorandum is to update you on several recent developments in the ongoing debate over the impact of the minimum wage on employment and school enrollment.

1. *New Neumark-Wascher Evidence*

The Neumark-Wascher research reporting negative effects of the New Jersey minimum wage hike was based upon actual payroll data for a subsample of the restaurants surveyed by Card and Krueger. The payroll data were provided to Neumark and Wascher by the Employment Policies Institute (EmpPI, not to be confused with the Economic Policy Institute). Although no direct evidence suggests that the sample supplied to Neumark and Wascher was specifically selected to yield negative employment impacts, suspicions have lingered, given EmpPI's connections to the fast-food industry.

Because of those suspicions, Neumark and Wascher collected payroll data from a second sample of fast-food restaurants in New Jersey and Pennsylvania. The new paper also reports significant negative impacts of the New Jersey minimum wage increase on employment. Unfortunately, throughout most of the paper, they simply pool their new data with the data supplied to them by EmpPI. However, in the text of the paper they note that they would have found a negative, but *insignificant* effect, if they had used the new data alone.

Neumark and Wascher defend the decision to pool their new data with the EmpPI data by arguing that one could not reject the hypothesis that the point estimates yielded by the two data-sets were the same. However, the samples do not seem to have been drawn from the same distributions, because one could reject the hypothesis that the residual error was the same in the two data-sets. (The suspect EmpPI data showed much more variation than the new data collected by Neumark and Wascher.) Unfortunately, the new sample did not include those firms in the original EmpPI data and, therefore, may not be representative by itself.

Therefore, we have three estimates of the effect of the New Jersey minimum wage increase on employment in fast-food restaurants:

1. Card and Krueger's positive and significant estimate based upon a phone survey of firms. The Card and Krueger survey has been attacked for asking imprecise questions about employment which proved more variable than actual payroll records. However, such criticisms may be misplaced. Although it may have made inference more difficult, measurement error in a dependent variable would not necessarily lead to bias unless, for some reason, employment growth was overstated in New Jersey.

2. An insignificant negative impact based upon new payroll data collected by Neumark and Wascher.

3. A significant negative estimate based upon the EmpPI sample which has been a subject of skepticism.

The most recent Neumark and Wascher paper suggests an elasticity of -.24. With a 12% percent increase in the minimum wage from $4.25 to $4.75 would imply only a 2 percent decline in employment in the fast-food industry, and probably smaller declines in industries less dependent upon minimum-wage workers. Therefore, even using the Neumark and Wascher results, it seems fair to say that the employment effects would be small. Furthermore, with any demand elasticity greater than -1, an increase in the minimum wage will raise total earnings (wage times hours) of low-wage workers, since a 10 percent increase in wage results in less than 10 percent decline in employment.

2. *School Enrollment Effects*

In a recent editorial in the *Wall Street Journal* (attached), Robert Barro highlights several other papers by Neumark and Wascher arguing that increases in the minimum wage lead some youth to drop out of school. The school enrollment impacts are vital since Neumark and Wascher have shown that it is primarily when one includes their measure of school enrollment as a regressor in a state panel analysis that one finds that the effect of the minimum wage is negative and significant.

Their work on school enrollment has been criticized primarily because their measure of school enrollment, excludes those who are in school but who report their major activity as "working". If an increase in the minimum wage encourages some youth to work a few more hours, students may switch their response on the major activity question from "school" to "work", although school enrollment may not have changed.

The October CPS, in contrast, identifies school enrollment independent of labor force activity. A recent analysis by Bill Evans from the University of Maryland and Mark Turner from the Urban Institute using October CPS data has suggested no statistically significant effect of the minimum wage on school enrollment. Further, using the alternative measure of school enrollment, the panel analysis reveals no statistically signficant effect of the minimum wage on employment. Neumark and Wascher are currently formulating a response to this critique.

It seems ironic that Barro would hold up Neumark and Wascher's result as saving the conventional wisdom, since a drop in school enrollment after an increase in the minimum wage would also seem to be inconsistent with neoclassical theory. If youth were simply paid their productivity, an increase in the minimum wage would not raise anyone's earnings potential and, therefore, would not lead anyone to drop out of school to take advantage of a "higher wage." It would simply eliminate any employment prospects for those with productivity below the minimum wage. In fact, I would think that the simplest neoclassical model would lead one to expect an increase in school enrollment with a rise in the minimum wage: Those who could no longer find work because they were laid off would have their opportunity costs lowered and may find schooling a worthwhile investment. Employment would decline and school enrollment and idleness would increase.

Nevertheless, I think the jury is still out on the school enrollment effects. Although I do not have a good story, in my own work with the CPS, I have noticed declines in school enrollment in California relative to other states after the April 1988 minimum wage increase in California. The finding is the same using administrative data on community college enrollments.

As new evidence is brought to light, I will keep you posted.