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Minimum Wages and Youth Employment in France and the United States

John M. Abowd, Francis Kramarz, Thomas Lemieux, and David N. Margolis

11.1 Introduction

In this paper we examine the link between changes in the minimum wage and employment outcomes for the youth (under age 31) labor market, in France and the United States. We make use of longitudinal data on employment status and earnings to see how individuals are affected by real increases (in the case of France) or real decreases (in the case of the United States) in the minimum wage conditional on the individual's loca-

John M. Abowd is professor of labor economics at Cornell University, distinguished senior research fellow at the U.S. Bureau of the Census, research associate at the Centre de Recherche en Economie et Statistique (CREST, Paris), and a research associate of the National Bureau of Economic Research. Francis Kramarz is head of the research department at INSEE-CREST, the French statistical institute; an associate professor at Ecole Polytechnique; and a research fellow of the Centre for Economic Policy Research, London. Thomas Lemieux is associate professor of economics at the University of British Columbia, a research director of the Centre Interuniversitaire de Recherche en Analyse des Organisations (CIRANO, Montréal), and a research associate of the National Bureau of Economic Research. David N. Margolis is a researcher with the Centre National de la Recherche Scientifique, working at the Université de Paris 1 Panthéon–Sorbonne in the Laboratoire de Microéconomie Appliquée and the Centre de Recherche en Economie et Statistique. Part of this paper was written while Margolis was assistant professor at the Université de Montréal and a research associate at the Centre de Recherche et Developement en Economie and the Centre Interuniversitaire de Recherche et d'Analyse des Organisations.

The authors gratefully acknowledge financial support from CIRANO, the National Science Foundation (SBR-93-21053 to Abowd and Margolis), and the Fonds pour la Formation de Chercheurs et l'Aide à la Recherche (97-NC-1676 to Margolis). Much of this work was completed while Margolis was visiting the CREST Laboratoire de Microéconométrie. The authors thank David Blanchflower, Richard Freeman, Shulamit Kahn, Lawrence Katz, Alan Krueger, John Martin, and participants at the NBER Summer Institute, CIRANO Summer Workshop, CREST Département de la Recherche Internal Workshop, CREST Microéconométrie Workshop, and the Université de Paris 1 Panthéon–Sorbonne for comments on previous versions of this paper. The American data used in this study were taken from public-

tion in the earnings distribution. We take particular care to distinguish subpopulations that might be affected differently by the minimum wage, focusing in particular on low-wage workers and (in the case of France, where the data are available) on the use of employment promotion contracts that allow the payment of subminimum wages.

Although little attention has been paid to the situation in Europe,[1] some European countries provide interesting alternatives to the much studied U.S. case. France, in particular, seems a perfect contrast to the United States. Whereas in the United States the nominal federal minimum wage remained constant for most states during most of the 1980s (thus implying a declining real federal minimum wage), nominal minimum wages in France rose steadily over the 1980s, as did real minimum wages. In this paper we exploit the different growth patterns in real minimum wages in a symmetric manner to more clearly understand their effect on employment.

Most existing studies of the French minimum wage system use aggregate time-series data and find no effect of the minimum wage system on youth employment (see, e.g., Bazen and Martin 1991). This is surprising because, since the inception of the minimum wage, a significant percentage of the French labor force has been employed at wages close to that level. One reason for the orientation in the empirical analyses done in France is, certainly, the tendency of American applied researchers to rely on aggregate time-series analyses[2] prior to the widespread dissemination of public-use microeconomic data such as the Current Population Survey (CPS). Another reason is that research access to French microdata was extremely limited until the 1990s. In the present study we use microdata from France and the United States collected in household surveys that are quite comparable. In particular, we use longitudinal information on the workers. Consequently, we are able to analyze both French and American minimum wage systems using individual-level panel data.

Because of the dramatic differences between the evolution of both nominal and real French minimum wages and that of the national U.S. minimum,[3] we have designed statistical comparisons that address the same be-

use Current Population Survey (CPS) files provided by the Bureau of Labor Statistics and the Bureau of the Census. David Card graciously provided the computer code for implementing the Census Bureau CPS matching algorithms used in this paper. The French data were taken from the Enquête Emploi research files constructed by the Institut National de la Statistique et des Etudes Economiques (INSEE, the French national statistical agency). The French data are also public-use samples. For further information contact INSEE, Département de la Diffusion, 18 bd Adolphe Pinard, 75675 Paris Cedex 14, France. The opinions expressed in this paper are those of the authors, not the U.S. Census Bureau. The paper was completed before Abowd assumed his appointment.

1. See Dolado et al. (1996) for a summary of minimum wage studies for France, the Netherlands, Spain, and the United Kingdom.

2. See Brown, Gilroy, and Kohen (1982) for a review.

3. We do not consider state-specific minimum wages or youth subminimum wages in the United States, which became increasingly important at the end of the 1980s. See Neumark

havior using the different variations in the national minimum wage systems to identify the relevant effects. We use two different statistical approaches based on the same idea: analysis of employment transition probabilities conditional on the position of an individual in the wage distribution. In each approach, we decompose the wage distribution into four components (under, around, marginally over, and over the minimum wage). We then, in our first approach, use a multinomial logit model to analyze the factors that affect the probability of making a transition between a particular position in the wage distribution and employment or nonemployment (in the case of France) or between employment or nonemployment and the position in the wage distribution (in the case of the United States). We find that young workers paid around the minimum wage in France were more likely to transition to a nonemployment state (unemployment or inactivity) than those paid over the minimum wage and that, for French men, such differences were greater in years where major increases in the minimum wage occurred. In the United States, we find that among workers currently employed around the minimum wage, a larger share were in a nonemployment state the previous period than among workers above the minimum wage. In both cases, the effects are strongest for the youngest workers. We find some minor "spillover" effects in both cases and provide evidence to suggest that these effects capture some of the heterogeneity between low-wage and high-wage labor markets.

In the second approach, we exploit the size of the movements in the real minimum wage more directly.[4] For France, we use the automatic and legislated increases in the nominal minimum wage that occur (at least) each July to identify groups of workers whose current wage rate will fall below the new minimum wage rate after the increase. We also identify workers whose present employment is part of a special youth program that permits wage payments below the statutory minimum. We use the limited duration of employment spells in such programs to identify a second group of minimum wage employment effects. Our statistical analysis identifies the change in future employment probabilities given an individual's minimum wage status in the present period. We show that individuals whose reference year wage was between the two real minimum wages, as defined above, have substantially lower subsequent employment probabilities than those who were not. The conditional elasticity of subsequent nonemployment as a function of the real minimum wage for young male workers in France in this situation, evaluated at sample means, is -2.5. This effect is present even when unobserved labor market heterogeneity

and Wascher (1992) for an explicit treatment of this variation in the U.S. data. Similarly, we do not explicitly control for minimum wages specified by collective agreement in France that exceed the national minimum. See Margolis (1993) for a detailed treatment of the effects of the collective bargaining agreement salary grids on employment.

4. Our analysis bears some resemblance to that of Linneman (1982).

and supply behavior are partially controlled for by the inclusion of a separate category for workers marginally over the minimum. However, the impact of the minimum wage decreases with experience. We also show that youths who participated in employment programs had lower subsequent employment probabilities. For the United States we use the constancy of the nominal minimum wage between 1981 and 1987 to identify groups of employed workers whose real wage in the present period would have been below the real minimum wage in the previous period. We show that young men whose wages were between the two real minimum wages, as described above, had lower employment probabilities in the previous period than individuals who were not (the conditional elasticity, evaluated at sample means, is 2.2). These effects get worse with age in the United States and are mitigated by eligibility for special employment promotion contracts in France.

The structure of this paper is as follows. Section 11.2 provides some institutional background on the systems of minimum wages in both France and the United States and provides some preliminary indications of the potential impact in each case based on empirical wage distributions. Section 11.3 describes the data that we use to analyze the impact of minimum wages, and section 11.4 lays out the statistical models used to evaluate the employment effects of minimum wage changes. Section 11.5 details the results of our multinomial logit analysis, and section 11.6 discusses the conditional logit analyses. Section 11.7 concludes.

11.2 Institutional Background

11.2.1 France

The first minimum wage law in France was enacted in 1950, creating a guaranteed hourly wage rate that was partially indexed to the rate of increase in consumer prices. Beginning in 1970, the original minimum wage law was replaced by the current system (called the *salaire minimum interprofessionnel de croissance*—SMIC) linking the changes in the minimum wage to both consumer price inflation and growth in the hourly blue-collar wage rate. In addition to formula-based increases in the SMIC, the government legislated increases many times over the next two decades. The statutory minimum wage in France regulates the hourly regular cash compensation received by an employee, including the employee's part of any payroll taxes.[5]

5. In theory, no provisions in any of the minimum wage laws allow regional variation in the SMIC. In some sectors in the French economy, however, the effective minimum wage was determined by (often extended) collective bargaining agreements. These agreements typically covered entire regions and industries, especially when extended to nonbargaining employers. Although relatively important in the 1970s, these provisions became increasingly irrelevant

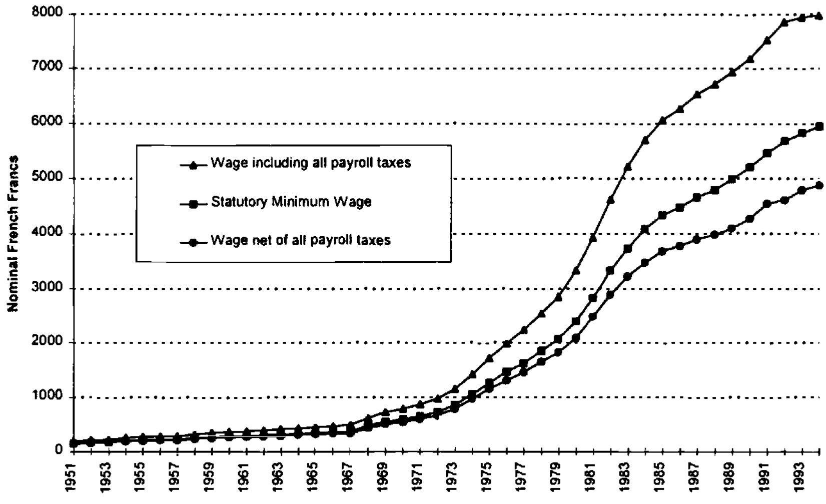

Fig. 11.1 Monthly minimum wage: France

Figure 11.1 shows the time series for the French minimum wage and the associated employee-paid and employer-paid payroll taxes. Because of the extensive use of payroll taxes to finance mandatory employee benefits, by the 1980s the French minimum wage imposed a substantially greater cost on the employer than its statutory value. Employees share in the legal allocation of the payroll taxes, as the figure shows; however, low-wage workers benefit substantially more than the average worker from the social security systems financed through these taxes in proportion to their revenue (unemployment insurance, health care, retirement income, and employment programs, in particular). Appendix table 11A.1 provides a complete statistical history of the real and nominal SMIC, including employer and employee payroll tax components.

Figure 11.2 shows the real hourly French minimum wage from 1951 to 1994. Although the original minimum wage program (called the *salaire minimum interprofessionnel garanti*—SMIG) was only partially indexed—in particular the inflation rate had to exceed 5 percent per year (2 percent from 1957 to 1970) to trigger the indexation—the real minimum wage did not decline measurably over the entire postwar period and increased substantially during most decades.

The French minimum wage lies near most of the mass of the wage rate distribution for the employed workforce. To show the location of the SMIC in this distribution, we plotted the empirical distribution of hourly

during the 1980s (our period of analysis) as the collective agreement nominal salary grids remained fixed in the face of an increasing nominal SMIC. See Margolis (1993) for a discussion of extended collective agreements and their relation to the SMIC.

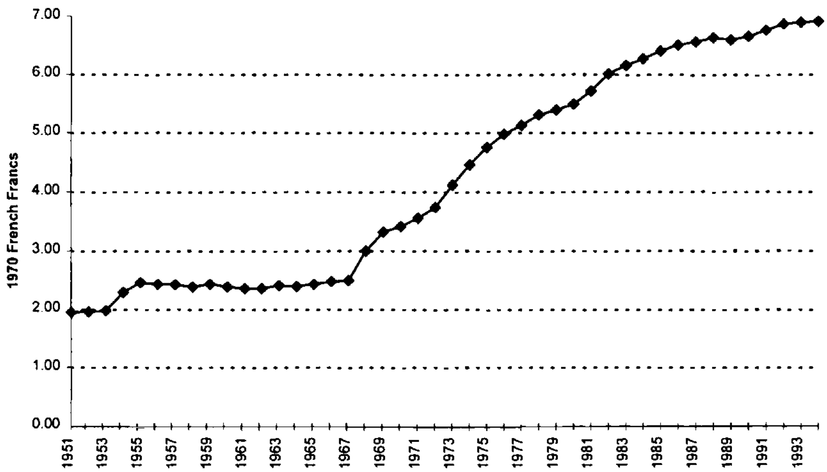

Fig. 11.2 Real hourly minimum wage: France

wage rates for 1990, the earliest year for which the Labor Force Survey reports continuous wage data. Figure 11.3 shows these data. We have indicated the SMIC directly on the figure. Notice that the first mode of the wage distribution is within F 5 of the minimum wage and the second mode is within F 10 of the minimum. In the overall distribution, 13.6 percent of the wage earners lie at or below the minimum wage and an additional 14.4 percent lie within an additional F 5 per hour of the SMIC.

Dolado et al. (1996) discuss the incidence of the SMIC with respect to household income. They find that although people employed at the SMIC do tend to be in the poorest households, the distribution of "*smicards*" (people paid the SMIC) is not monotonically decreasing in household income. For example, they find that the share of individuals paid the SMIC in each decile of household income increases from 10.1 percent in the lowest decile to 13.1 percent in the third lowest decile, then decreases to 6.6 percent for the fifth decile, increases to 7.4 percent for the sixth decile and declines monotonically to 0.6 percent in the highest decile of household income.

11.2.2 United States

The first national minimum wage in the United States was a part of the original Fair Labor Standards Act (FLSA) of 1938. The American national minimum wage has never been indexed and increases only when legislative changes are enacted. The national minimum applies only to workers covered by the FLSA, whose coverage has been extended over the years to include most jobs. The statutory minimum wage regulates the

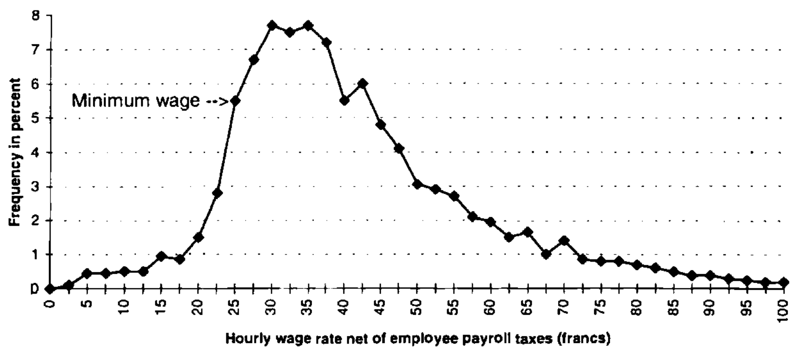

Fig. 11.3 Empirical distribution of hourly wages: France, 1990

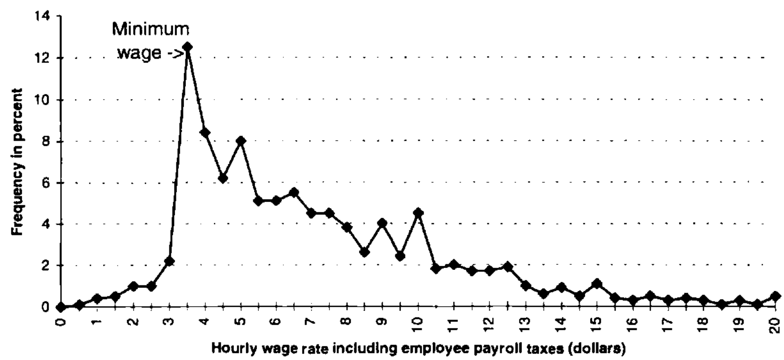

Fig. 11.4 Empirical distribution of hourly wages: United States, 1981

hourly regular cash compensation received by an employee including the employee's part of any payroll taxes.

Figures 11.4 and 11.5 show the distribution of the American hourly wage rate and the location of the minimum wage in that distribution for 1981 and 1987, the beginning and ending years of our analyses.[6] For 1981, 17.7 percent of the employed workforce had wage rates at or below the minimum wage and an additional 14.6 percent had wage rates within an additional $1.00 per hour of the minimum. For 1987, only 9.5 percent of employed persons had hourly wage rates at or below the minimum while an additional 9.9 percent lay within an additional $1.00 per hour of the minimum.

6. It should be noted that the federal minimum wage was increased to $3.35 per hour in 1980.

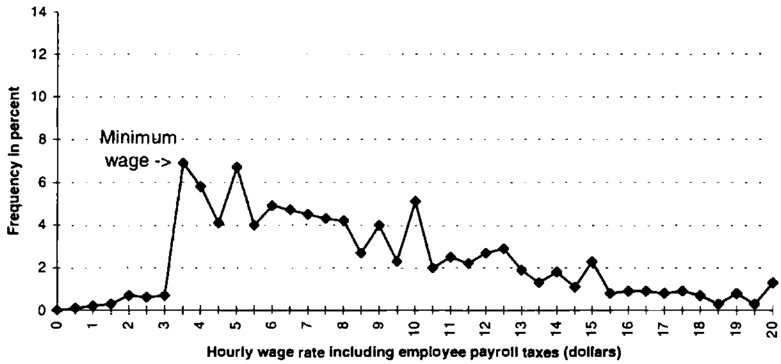

Fig. 11.5 Empirical distribution of hourly wages: United States, 1987

11.3 Data Description

11.3.1 France

The French data were extracted from the Enquête Emploi (Labor Force Survey) for the years 1982–89. The 60,000 households included in the Labor Force Survey sample are interviewed in March of three consecutive years with one-third of the households replaced each year. Every member of the household is surveyed and followed provided that he or she does not move during the three years. We used the INSEE research files for each of the indicated years. These files include identifiers that allowed us to follow individuals from year to year. Using these identifiers we created year-to-year matched files for the years 1982–83 to 1988–89.

The survey measures usual monthly earnings, net of employee payroll taxes but including employee income taxes, and usual weekly hours. Usual monthly earnings is measured in 20 intervals of widths varying from F 500 to F 5,000. It is important to note that the narrowest intervals were used for the lowest salaries. We take the categorical nature of our wage data explicitly into account in our analyses, in that we compare the declared wage category against the wage category in which an individual working the same number of hours per month at the SMIC would be found.

Certain young workers were employed in publicly funded programs that either combined classroom education with work ("*apprentis*," "*stage de qualification*," or "*stage d'insertion, contrat emploi-formation*") or provide subsidized low-wage employment ("*travaux d'utilité collective*" or "*stage d'initiation à la vie professionnelle*," both from 1985 to 1989). All of these programs provide a legal exemption from the SMIC and from certain payroll taxes. Most of these programs are limited to workers 25 years old and under.

The employment status in year t is equal to one for all individuals who are employed in March of the survey year and equal to zero otherwise. The French Labor Force Survey definition of employment is the same as the one used by the International Labour Office: a person is employed if he or she worked for pay for at least one hour during the reference week. The definition is thus consistent with the American Bureau of Labor Statistics (BLS) definition used below.

Our control variables consisted of education, labor force experience, seniority, region of France, date of labor force entry, and year. Education was constructed as eight categories: none, completed elementary school, completed junior high school, completed basic vocational/technical school, completed advanced vocational/technical school, completed high school (*baccalauréat*), completed technical college or undergraduate university, and completed graduate school or postcollege professional school. Labor force experience was computed as the difference between current age and age at school exit. Seniority was measured as the response to a direct question on the survey (years with the present employer). Region is an indicator variable for the Ile de France (Paris metropolitan area) as the region of residence.

The SMIC data were taken from Bayet (1994), which reports official INSEE statistics. We selected the hourly SMIC for March of the indicated year, net of employee payroll taxes.

11.3.2 United States

We used the official BLS public-use outgoing rotation group files from the CPS for the months January to May and September to December and the years 1981–87. We applied the Census Bureau matching algorithm to create year-to-year linked files for the years 1981–82 to 1986–87.

The outgoing rotation groups (households being interviewed for the fourth or eighth time in the CPS rotation schedule) are asked to report usual weekly wage and usual weekly hours. Individuals who normally are paid by the hour are asked to report that wage rate directly. We created an hourly wage rate using the directly reported hourly wage rate when available and the ratio of usual weekly earnings to usual weekly hours otherwise. Respondents are asked to report these wage measures gross of employee payroll taxes, so they are not directly comparable to the measures constructed from the French data, which are reported net of employee payroll taxes. We created real hourly wage rates by dividing by the 1982–84-based Consumer Price Index for All Urban Workers for the appropriate month.

We created a second set of hourly wage measures for the United States that included income from tips in the hourly wage. To do this we computed a second hourly wage rate as usual weekly earnings divided by usual weekly hours for workers who reported that they were paid by the hour.

When this second hourly wage rate exceeded the one directly reported, we used the computed measure. This measure of hourly wage rate is used below in the analysis labeled "including income from tips."

An individual is employed in year t if he or she worked at least one hour for pay during the second week of the survey month. We used the CPS employment status recode variable to determine employment. The BLS definition is thus consistent with the one used in the French Labor Force Survey.

Our control variables consist of education, potential labor force experience, race, marital status, and region. Education was constructed as the number of years required to reach the highest grade completed. For the multinomial logit analysis, this was decomposed into six categories: less than junior high school (no diploma), junior high school, high school, less than four years of college, four years of college, and more than four years of college. Potential labor force experience is age minus years of education minus five. Race is one for nonwhite individuals. Marital status is one for married persons. Region is a set of three indicator variables for the northeastern, north-central, and southern parts of the United States.

The U.S. national nominal minimum wage was \$3.35 throughout our analysis period.[7]

11.3.3 Empirical Transition Probabilities

A preliminary analysis of the empirical transition probabilities of young workers into or out of employment based on their positions in the wage distribution relative to the minimum wage suggests that one might expect to see significant impacts of the minimum wage on employment probabilities in both France and the United States. In the case of France, we are concerned with that probability that an individual is employed at the date $t + 1$ given the person's employment status and wage rate relative to the SMIC (if employed) at date t. In the case of the United States, the question is whether or not an individual was employed at date t given his or her employment status and wage rate relative to the minimum wage (if employed) at date $t + 1$.

Let miw_t be the nominal hourly minimum net wage in year t, $rmiw_t$ be the real hourly minimum net wage in year t, and h_t represent the number of monthly hours worked in the sample month in year t. For France let $wcat_t$ be the category in which the individual's nominal net monthly earnings falls in year t, and for the United States let w_t be the individual's hourly net wage rate in year t and rw_t be the real net wage for year t.

7. Throughout the period, and particularly toward the end, some states independently increased their nominal wages above the national level. We do not explicitly account for state-by-state variation in the nominal minimum wage. See Neumark and Wascher (1992) for an analysis, using a different methodology, of the effects of interstate variation of minimum wages in the United States.

For France define $micat_t$ as the earnings category into which expected nominal monthly earnings at the SMIC ($h_t \times miw_t$) would fall, and order the categories from 1 (less than F 500 per month) to 15 (over F 45,000 per month). Then we define the following six departure (occupied at date t) states:

out of the labor force at t,
unemployed at t,
employed at t and paid under the SMIC: $I(wcat_t < micat_t) = 1$,
employed at t and paid the SMIC: $I(wcat_t = micat_t) = 1$,
employed at t and paid marginally over the SMIC: $I(wcat_t = micat_t + 1) = 1$, and
employed at t and paid over the SMIC: $I(wcat_t > micat_t + 1) = 1$,

where $I(\cdot)$ is the indicator function taking the value one when the condition is true and zero otherwise. We also define two arrival (occupied at date $t + 1$) states:

employed at $t + 1$ and
not employed at $t + 1$.

For the United States recall that the nominal minimum wage was constant over the entire sample period at \$3.35 per hour. Thus we construct six arrival states:

out of the labor force at $t + 1$,
unemployed at $t + 1$,
employed at $t + 1$ and paid under the minimum wage: $I(w_{t+1} < \$3.25) = 1$,
employed at $t + 1$ and paid the minimum wage: $I(\$3.25 \leq w_{t+1} < \$3.50) = 1$,
employed at $t + 1$ and paid marginally over the minimum wage: $I(\$3.50 \leq w_{t+1} < \$4.00) = 1$, and
employed at $t + 1$ and paid over the minimum wage: $I(w_{t+1} \geq \$4.00) = 1$.

We have the same two departure states:

employed at t and
not employed at t.

Using these definitions, figures 11.6 and 11.7 describe the breakdown of the population and the change in the real hourly minimum wage for French young men and women, respectively, and figures 11.8 and 11.9 show the corresponding breakdowns and changes for U.S. young men and women, respectively. Table 11.1 describes the distribution of transitions over the sample periods for the French data, and table 11.2 describes the distribution of transitions for the American data.

In the case of the United States, it is clear from looking at the raw transition probabilities that minimum wage workers are different from

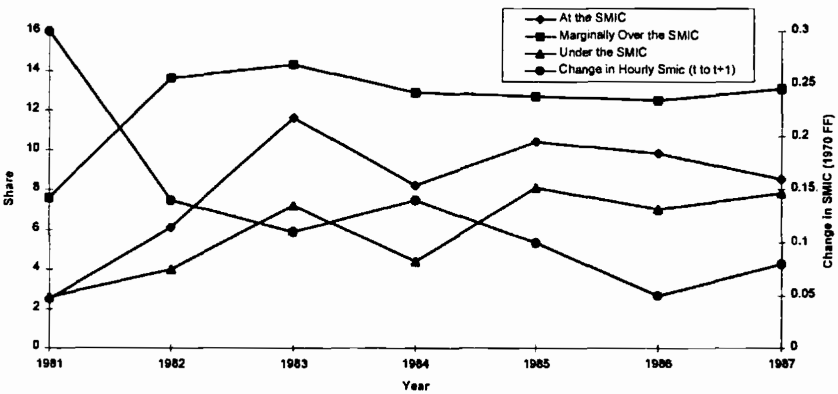

Fig. 11.6 Population breakdown by earnings and evolution of real SMIC: French young men

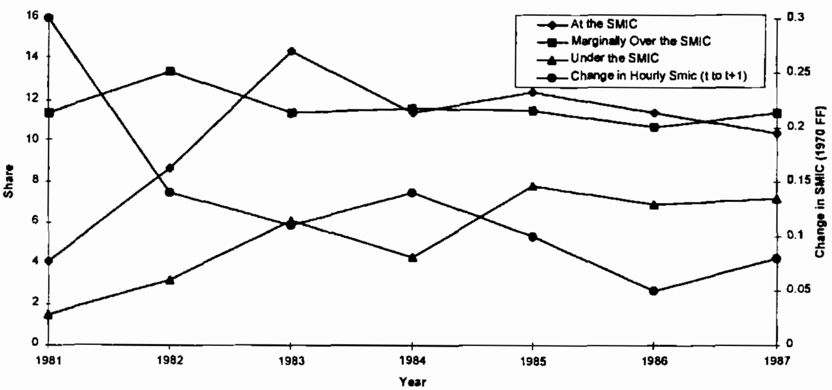

Fig. 11.7 Population breakdown by earnings and evolution of real SMIC: French young women

their higher paid counterparts. A much larger share of the population employed at the minimum wage at date $t + 1$ comes from the nonworking pool (42.92 percent) than does the share of the population employed far over the minimum wage (only 12.28 percent). The case in France is less clear, since the difference between the share of workers paid at the SMIC who are not employed the following period (6.63 percent) and the share paid over the SMIC who are not employed the following period (12.16 percent) is much less dramatic, and even goes in the opposite direction from the U.S. result. These effects may, however, be due to the presence of various sorts of employment promotion contracts, which might shield workers paid at or under the SMIC from layoffs. Such effects would not

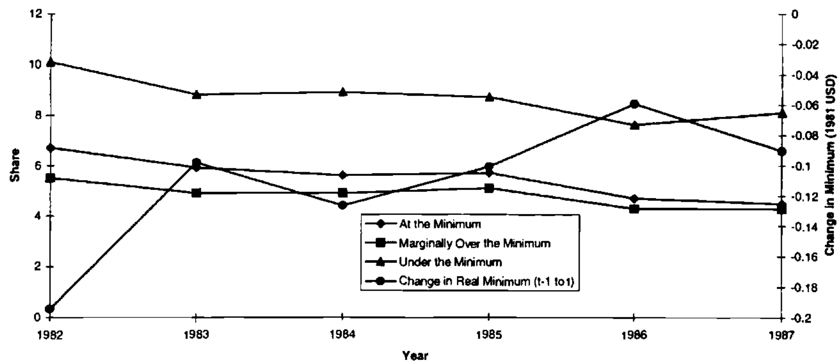

Fig. 11.8 Population breakdown by earnings and evolution of real minimum wage: U.S. young men

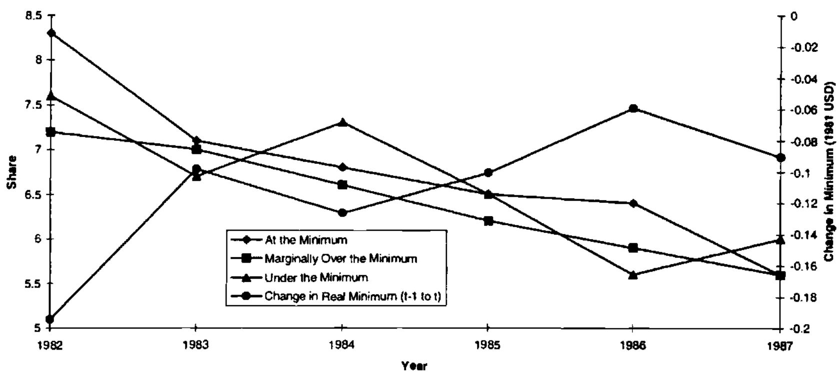

Fig. 11.9 Population breakdown by earnings and evolution of real minimum wage: U.S. young women

be visible in these cross-tabulations, and our conditional logit results go to great lengths to try to discriminate between the effects of the contracts and the effects of the minimum wage.

It should be noted that the transition behavior of workers paid marginally over the minimum is, in both countries, intermediate between the transitions made by those paid at the minimum and those paid over the minimum. This "spillover" effect could be capturing a degree of heterogeneity between low-wage and high-wage workers, and we will exploit this control group in what follows.

Clearly, this descriptive analysis is not sufficient to discredit the hypothesis that low-wage workers are, in some way, qualitatively different from high-wage workers; in fact, the spillover effect noted above suggests that

Table 11.1 **Transition Probabilities for France**

	To Nonemployment			To Employment					
From	Out of LF	Unemployed	Total	Under SMIC	At SMIC	Marginal SMIC	Over SMIC	Total	Total
Out of labor force	10,081	2,150	12,231	556	574	452	852	2,434	14,665
Overall %	11.67	2.49	14.15	0.64	0.66	0.52	0.99	2.82	16.97
Row %	68.74	14.66	83.40	3.79	3.91	3.08	5.81	16.60	100
Column %	67.54	18.10	45.64	11.85	7.40	4.64	2.28	4.08	111.8
Unemployed	2,328	5,733	8,061	856	723	595	1,041	3,215	11,276
Overall %	2.69	6.63	9.33	0.99	0.84	0.69	1.20	3.72	13.05
Row %	20.65	50.84	71.49	7.59	6.41	5.28	9.23	28.51	100
Column %	15.60	48.27	30.08	18.25	9.32	6.11	2.78	5.39	100.33
Under SMIC	6	14	20	1410	474	210	220	2,314	2,334
Overall %	0.01	0.02	0.02	1.63	0.55	0.24	0.25	2.68	2.70
Row %	0.26	0.60	0.86	60.41	20.31	9.00	9.43	99.14	100
Column %	0.04	0.12	0.07	30.06	6.11	2.16	0.59	3.88	39.1
At SMIC	133	150	283	880	2,144	661	300	3,985	4,268
Overall %	0.15	0.17	0.33	1.02	2.48	0.76	0.35	4.61	4.94
Row %	3.12	3.51	6.63	20.62	50.23	15.49	7.03	93.37	100
Column %	0.89	1.26	1.06	18.76	27.63	6.79	0.80	6.69	56.1
Marginal SMIC	175	451	626	540	2,465	3,194	1,166	7,365	7,991
Overall %	0.20	0.52	0.72	0.62	2.85	3.70	1.35	8.52	9.25
Row %	2.19	5.64	7.83	6.76	30.85	39.97	14.59	92.17	100
Column %	1.17	3.80	2.34	11.51	31.77	32.79	3.12	12.36	84.1
Over SMIC	2,202	3,378	5,580	449	1,380	4,630	33,837	40,296	45,876
Overall %	2.55	3.91	6.46	0.52	1.60	5.36	39.16	46.63	53.09
Row %	4.80	7.36	12.16	0.98	3.01	10.09	73.76	87.84	100
Column %	14.75	28.44	20.82	9.57	17.78	47.53	90.43	67.60	208.5
Total	14,925	11,876	26,801	4,691	7,760	9,742	37,416	59,609	86,410
	17.27	13.74	31.02	5.43	8.98	11.27	43.30	68.98	100
	99.8	82.62	182.4	100.1	114.72	82.91	119.85	417.62	600
	100	100	100	100	100	100	100	100	600

Source: French Labor Force Survey, 1982–89, matched year to year.
Note: Table reports on people aged 30 or under.

Table 11.2 Transition Probabilities for the United States

From	To: Out of LF	Unemployed	Under Minimum	At Minimum	Marginal Minimum	Over Minimum	Total
Nonemployment							
Out of labor force	25,245	3,124	1,586	2,278	1,617	4,547	38,397
Overall %	19.27	2.38	1.21	1.74	1.23	3.47	29.31
Row %	65.75	8.14	4.13	5.93	4.21	11.84	100
Column %	72.53	31.21	16.40	30.79	22.69	7.33	181.0
Unemployed	2,466	2,819	574	897	773	3,065	10,594
Overall %	1.88	2.15	0.44	0.68	0.59	2.34	8.09
Row %	23.28	26.61	5.42	8.47	7.30	28.93	100
Column %	7.09	28.16	5.94	12.12	10.85	4.94	69.10
Total	27,711	5,943	2,160	3,175	2,390	7,612	48,991
Overall %	21.15	4.54	1.65	2.42	1.82	5.81	37.40
Row %	56.56	12.13	4.41	6.48	4.88	15.54	100
Column %	79.62	59.37	22.34	42.92	33.53	12.28	250.06
Employment							
Under minimum	1,511	471	5,038	674	490	2,018	10,202
Overall %	1.15	0.36	3.85	0.51	0.37	1.54	7.79
Row %	14.81	4.62	49.38	6.61	4.80	19.78	100
Column %	4.34	4.71	52.11	9.11	6.88	3.25	80.40
At minimum	1,445	668	424	2,002	1,502	2,231	8,272
Overall %	1.10	0.51	0.32	1.53	1.15	1.70	6.31
Row %	17.47	8.08	5.13	24.20	18.16	26.97	100
Column %	4.15	6.67	4.39	27.06	21.07	3.60	66.95

(continued)

Table 11.2 (continued)

From	To: Out of LF	Unemployed	Under Minimum	At Minimum	Marginal Minimum	Over Minimum	Total
Marginal minimum	1,091	485	323	673	1,534	3,467	7,573
Overall %	0.83	0.37	0.25	0.51	1.17	2.65	5.78
Row %	14.41	6.40	4.27	8.89	20.26	45.78	100
Column %	3.13	4.85	3.34	9.10	21.52	5.59	47.53
Over minimum	3,046	2,443	1,723	874	1,211	46,674	55,971
Overall %	2.33	1.86	1.32	0.67	0.92	35.63	42.72
Row %	5.44	4.36	3.08	1.56	2.16	83.39	100
Column %	8.75	24.41	17.82	11.81	16.99	75.28	155.06
Total	7,093	4,067	7,508	4,223	4,737	54,390	82,018
Overall %	5.41	3.10	5.73	3.22	3.62	41.52	62.60
Row %	8.65	4.96	9.15	5.15	5.78	66.31	100
Column %	20.38	40.63	77.66	57.08	66.47	87.72	349.94
Total	34,804	10,010	9,668	7,398	7,127	62,002	131,009
	26.57	7.64	7.38	5.65	5.44	47.33	100
	141.2	58.21	71.4	55.66	56.89	216.70	600
	100	100	100	100	100	100	600

Source: U.S. Current Population Survey, 1981–87, January–May and September–December, matched year to year.
Note: Table reports on people aged 30 or under.

such heterogeneity may exist. To separate out this effect, we need to control for worker characteristics and analyze more carefully the transitions between employment and nonemployment.[8]

11.4 Statistical Models for the Minimum Wage Effects on Employment

In order to control for the impact that variables, including the minimum wage and its movements, might have on labor market transitions, we applied two different statistical techniques. In the first approach, we use a multinomial logit analysis to try to control for factors that might render low-wage workers different from other workers and could thereby affect their transition probabilities. We analyze the raw transitions and describe the factors that increase or reduce the probability of transitions involving nonemployment and how these factors differentially affect minimum wage and above minimum wage workers. In the second approach, we exploit the size of the increases to categorize workers as "between" old and new values of the real minimum wage (i.e., with an hourly real wage rate lying between the old and the new real minimum wage), and we use a logit analysis of subsequent (or prior) employment probabilities to see if workers who might be directly affected by minimum wage increases have significantly different subsequent (or prior) employment probabilities.

11.4.1 Multinomial Logit Analysis

Using the same definitions of states as in subsection 11.3.3, we regroup the unemployed and inactive states into a single state, nonemployment. Using the notation N = nonemployment, E = employment, U = under the minimum, A = at the minimum, M = marginally over the minimum, and O = over the minimum, we can define the set of possible transitions for each country. Thus for France there are 10 possible transitions: O to E or O to N, M to E or M to N, A to E or A to N, U to E or U to N, and N to E or N to N. For the United States there are 10 symmetric transitions: E to O or N to O, E to M or N to M, E to A or N to A, E to U or N to U, and E to N or N to N. We use a multinomial logit approach to control for observable factors while allowing for a common shock. For interpretation, however, we are particularly concerned with the conditional transition probabilities.

In the French case, we are interested in the probability of transition out of employment conditional on the position in the earnings distribution.

8. There remains a possibility that unobserved worker heterogeneity might bias our results in sections 11.5 and 11.6. Because of selection considerations and sample sizes, we were not able to use standard (Hsiao 1986) or nonstandard (Abowd, Kramarz, and Margolis 1999) techniques to control for these effects. Thus we are forced to suppose that the inclusion of the "marginally above" the minimum wage group is sufficient to capture any heterogeneity in transition rates that is correlated with wages.

For the United States, we are interested in the initial state of a worker conditional on his or her ex post position in the earnings distribution. In each of these cases, we have in mind the hypothesis of a competitive labor market, and thus a model in which a worker with a given marginal productivity (equal to the wage) closer to the minimum wage might be more at risk to transit out of employment in France or to have come from nonemployment in the United States than an observationally equivalent worker paid above the minimum wage. We suppose that those workers employed at wages marginally above the minimum share unobservable characteristics that affect transition probabilities in the absence of a minimum wage, and that all differences in their transition behavior can be attributed to the more direct impact of the minimum wage on those paid at it relative to those paid marginally over it. We can use our parameter estimates from the multinomial logit to see how the differences in these conditional transition probabilities evolve over time, thus seeing if the difference is correlated with movements in the real minimum wage. This approach is particularly useful not only for seeing how minimum wage movements affect the probability of job loss conditional on employment (or on having come from nonemployment conditional on being employed) but also for determining whether minimum wage movements play a role in excluding workers completely from the labor market. We can also see which workers are the most likely to transition out of employment in France or come from nonemployment in the United States based on observable characteristics, such as age, conditional on the individual's position in the earnings distribution. Furthermore, since our estimates are based on the entire population, interpretation of these results can be more easily generalized than the results based on the employed subsample of our data, as in the conditional logit analysis described below.

11.4.2 Conditional Logit Analysis

Once again, let $rmiw_t$ be the real hourly minimum net wage in year t and let rw_t be the real hourly net wage for year t. Let age_t represent an individual's age at the date t and $stage_t$ indicate that the person was employed under some employment promotion contract that allows for subminimum wages in year t. Finally, let e_t indicate the individual's employment status in year t ($e_t = 1$ if employed).

We define a person as "between" in France if the mean of the cell in which the person is located at the date t is at or above the minimum wage at date t but below the minimum wage (in date t francs) at date $t + 1$. Algebraically, after defining rw_t to be the mean of the cell in which the individual is located, this is equivalent to

$$I(rmiw_t \leq rw_t \leq rmiw_{t+1}) = 1.$$

We also break up the subminimum population (those for whom rw_t $rmiw_t$) into two groups in France: those on employment promotion contracts ($stage_t$) and those not on employment promotion contracts. Thus for France we estimate variants of the following equation for individuals:

$$
\begin{aligned}
&\Pr[e_{t+1} = 1 | e_t = 1] \\
&= F(x_t\beta + \alpha_1 I(rw_t < rmiw_t) \times stage_1 \times (rmiw_{t+1} - rmiw_t) \\
&+ \alpha_2 I(rw_t < rmiw_t) \times (1 - stage_1) \times (rmiw_{t+1} - rmiw_t) \\
(1) \quad &+ \alpha_3 I(rmiw_t \leq rw_t \leq rmiw_{t+1}) \times (rmiw_{t+1} - rmiw_t) \times age_t \\
&+ \alpha_4 I(rmiw_{t+1} < rw_t \leq (rmiw_{t+1} \times 1.1)) \times (rmiw_{t+1} - rmiw_t) \\
&\times age_t),
\end{aligned}
$$

where F(·) is the standard logistic function. The logit described in equation (1) allows us to test the hypothesis, implied by the theory of competitive labor markets, that if marginal productivity stays constant, increases in the real minimum wage render previously employed individuals, whose wages fall between the old and new minima, currently unemployable. In particular, this specification also us to see if the effects of the minimum wage vary with age, and we experiment with different degrees of age aggregation to evaluate particular labor market phenomena such as the end of eligibility for employment promotion contracts or mandatory military service.

We define a person as "between" in the United States if the person's wage at date $t + 1$ is at or above the minimum wage at date $t + 1$ but below the minimum wage (in date $t + 1$ dollars) at date t. Algebraically, this is equivalent to

$$I(rmiw_{t+1} \leq rw_{t+1} \leq rmiw_t) = 1.$$

We also define the variable $rmarg_t$ as the deflated value of \$4.00 at date t. Thus for the United States we estimate variants of the following equation:

$$
\begin{aligned}
&\Pr[e_t = 1 | e_{t+1} = 1] \\
&= F(x_t\beta + \alpha_1 I(rw_{t+1} < rmiw_{t+1}) \times (rmiw_t - rmiw_{t+1}) \times age_t \\
(2) \quad &+ \alpha_2 I(rmiw_{t+1} \leq rw_{t+1} \leq rmiw_t) \times (rmiw_t - rmiw_{t+1}) \\
&\times age_t + \alpha_3 I(rmiw_t < rw_{t+1} \leq rmarg_t) \times (rmiw_t - rmiw_{t+1}) \\
&\times age_t).
\end{aligned}
$$

The interpretation of equation (2) is symmetric to that of equation (1). Does a relatively large decrease in the real minimum wage allow previously unemployable individuals to be employed? Furthermore, in the United States, we explicitly examine the impact that tips might have on our measure of the position of a person in the wage distribution.

Notice that the equations for the United States have empirical content because the nominal minimum wage rate does not change during our sample period whereas the real minimum wage rate declines because of general price inflation. In contrast, the equations for France have empirical content because the indexation formula is tied to general price inflation and to the growth in average hourly earnings among blue-collar workers, and as noted in subsection 11.2.1, real minimum wages increased steadily throughout the sample period.[9]

11.5 Multinomial Logit Results

11.5.1 France

Appendix table 11A.2 shows some of the results of estimating the multinomial logit for France. We have reported only the coefficients on certain key variables; the reference state is the transition U to E. The multinomial logit models for both France and the United States were estimated on the entire population, and not just on the youth subpopulation (as is the case for the conditional logit models), in order to highlight differences between younger and older workers. A large number of the coefficients are significantly different from zero, and the differences in the intercepts are consistent with the raw transition probabilities (O-E is more probable than O-N, N-N is more probable than N-E, etc.). Having completed one's *baccalauréat* (roughly the equivalent of high school in the United States) is an advantage for those employed over the minimum wage (0.62 vs. 0.29 for men, 1.34 vs. 1.06 for women); however, men with *baccalauréats* who are employed at the minimum wage seem relatively worse off (−0.31 vs. −0.49). This might be coherent with a signaling explanation in which only the low-productivity *baccalauréat* holders are willing to accept jobs at the minimum wage.

In general, the coefficients corresponding to transitions from marginally over the SMIC are intermediate between transitions from at the SMIC and transitions from over the SMIC. This is consistent with the idea of using workers paid marginally over the SMIC as a comparison group for the purposes of analyzing the effects of the minimum wage on the popula-

9. Our conditional logit estimates are performed on the set of individuals who are employed at some point in the sample. Thus the coefficients should not necessarily be interpreted as representative of the entire potential labor force, but rather as appropriate for the sample of workers who satisfy the selection criterion.

tion of workers being paid at the minimum. For French women in particular, the time-series transition behavior of women paid marginally over the minimum strongly resembles that of women paid at the minimum. We exploit these results in the conditional logit models that follow in section 11.6.

Since the interpretation of the raw regression coefficients is not immediately informative, figure 11.10 explores the variation in conditional transition probabilities out of employment with age for a French man in 1984 who entered the labor market between 1962 and 1972, living in the Paris region with a *baccalauréat,* and figure 11.11 shows the same conditional transition probabilities for a French woman with the same characteristics. All conditional transition probabilities are conditional on the date *t* posi-

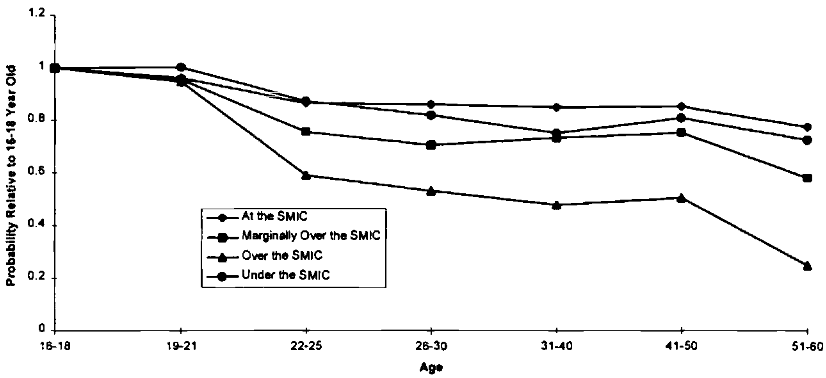

Fig. 11.10 Probability of leaving employment (relative to 16–18-year-olds): French men, 1984

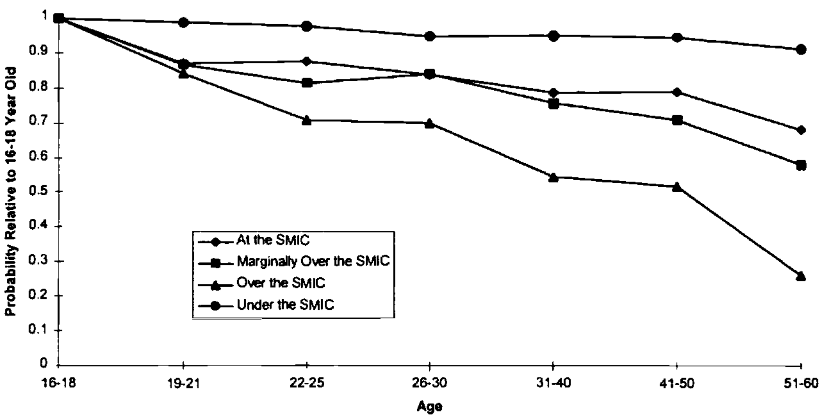

Fig. 11.11 Probability of leaving employment (relative to 16–18-year-olds): French women, 1984

tion in the earnings distribution. The general downward trends in both figures are due simply to the fact that young people are more likely to transition out of employment independent of position in the wage distribution. Still, it is worth noting that while 51–60-year-olds paid over the minimum are about a third as likely to transition out of employment than 16–18-year-olds, workers paid at the minimum seem to benefit much less from the reduction in the probability of transitioning out of employment as they age. Furthermore, it seems that aging does not reduce at all the probability of transitioning out of the labor force for women being paid under the minimum. This suggests that the subminimum population of older women is characterized by much weaker labor force attachment than comparable women paid elsewhere in the wage distribution.

11.5.2 United States

Appendix table 11A.3 shows some of the results of estimating the multinomial logit for the United States. Once again, we have reported only the coefficients on certain key variables; the reference state is the transition E to U. A certain number of the coefficients are significantly different from zero, and the differences in the intercepts are consistent with the raw transition probabilities (E-O is more probable than N-O, E-O is more probable than E-A, etc.). Having completed high school is associated with a relative higher share coming from employment for those employed over the minimum wage (0.75 vs. 0.49 for men, 0.65 vs. 0.37 for women); however, men with high school diplomas who are employed at the minimum wage come disproportionately from nonemployment (0.13 vs. 0.08) whereas the effect is opposite for women (−0.02 vs. 0.05), although the differences in the estimated coefficients are small. The subminimum transitions do not seem dramatically different from the at minimum transitions (the coefficients in the E-A column are rarely significantly different from zero), although a significantly smaller share of young women paid under the minimum were employed in the previous period, relative to those paid at the minimum. This suggests that low-wage employers hire relatively more from the pool of nonemployed, and it thus could be interpreted as running counter to the idea that subminimum sectors in the United States (particularly jobs that receive income from tips) provide more stable employment than jobs that pay the minimum wage.

As in the French case, the time-series behavior of the transitions of workers paid marginally over the minimum closely mimics that of workers paid at the minimum, further reinforcing the idea that the group of workers paid marginally over the minimum might be a reasonable control group for minimum wage workers. Also, as in the French case, the interpretation of the raw coefficients can be difficult. Figure 11.12 explores the variation in conditional (on arrival state) transition probabilities into employment with age for an American man in 1984 who entered the labor

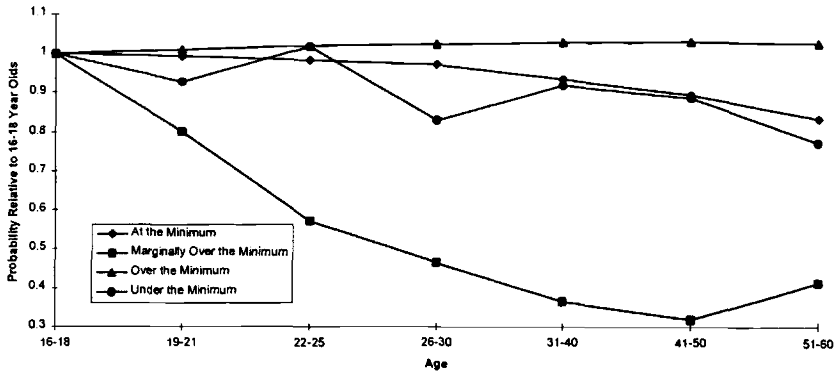

Fig. 11.12 Probability of moving into employment (relative to 16–18-year-olds): U.S. men, 1984

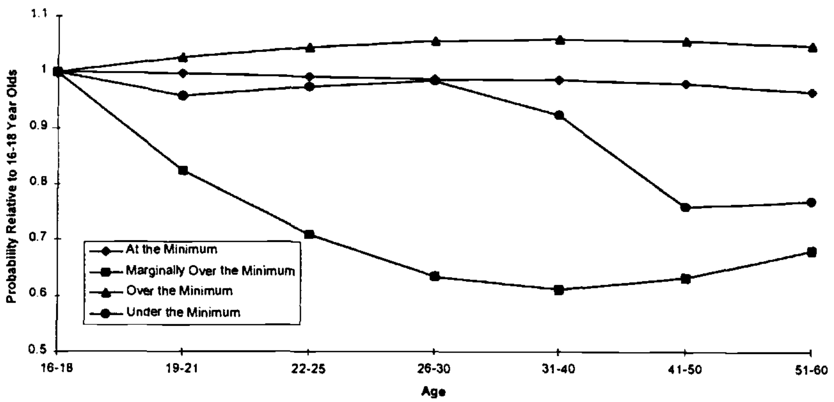

Fig. 11.13 Probability of moving into employment (relative to 16–18-year-olds): U.S. women, 1984

market between 1962 and 1972 with a high school diploma, and figure 11.13 shows the variation of the conditional transition probabilities for an American woman with the same characteristics.

Clearly, in the United States, the effect of age on the transition probabilities differs dramatically from the French case. The two figures are similar in form, although the relative reduction in the conditional probability of transitioning from nonemployed to marginally over the minimum is stronger for men and turns back up sooner for women. The most remarkable difference between the French and U.S. cases is that while in France the probability of making a O-N transition decreases with age, there is either no effect or a slight increase in the relative probability of N-O transitions (the U.S. equivalent) for older workers relative to younger workers in our results for the United States. This could be due to the high stability in general of jobs that pay substantially over the minimum wage; the intercepts for E-O transitions are significantly larger than all other estimated intercepts in the model. On the other hand, in the United States it seems that the probability of transitioning from nonemployment to marginally over the minimum wage is the transition the most affected by aging, while in France the order of magnitude of the change is about half for 31–40-year-olds relative to 16–18-year-olds (a 63 percent drop vs. a 27 percent drop for men, a 39 percent drop vs. a 24 percent drop for women). If workers paid marginally over the minimum are indeed a reasonable control group for minimum wage workers, the relatively feeble decline in the probability of having come from nonemployment experienced by workers paid at the minimum suggests that in the United States at least, the minimum wage is playing a role in determining the sorts of transitions that low-wage workers make in the labor market.

11.6 Conditional Logit Results

11.6.1 France

Table 11.3 shows the results of estimating equation (1) for France on young people, using broad age categories.[10] We have reported the coefficients for the key real minimum wage variables, as well as variables for several types of employment contracts in France.[11]

The coefficients show that French men aged 25–30 with real wage rates in period t that are above the real minimum in t but below the real minimum wage in period $t + 1$ have much lower subsequent employment probabilities than similar men paid substantially over the period $t + 1$ real minimum wage. The elasticity is very large: an increase of 1 percent in the minimum wage entails an decrease in the probability of keeping one's job of 4.6 percent, relative to men aged 25–30 who are paid marginally over the minimum. One interpretation of these results is that although low-wage workers do differ from high-wage workers (as the fairly consistent negative coefficients suggest), the minimum wage hits workers whose real wages are between the two minima much harder than other low-wage workers.

Similar results hold for women and people 20–24 years old, but these coefficients are less significant. In general, the employment loss effects worsen with age among the young employed population, but the level of

10. Appendix table 11A.4 provides descriptive statistics for the French data used in these regressions.

11. We explicitly consider fixed-term contracts (CDD), youth employment schemes (young *stagiaire*), and apprenticeships, with the reference being long-term contracts (CDI). See Abowd, Corbel, and Kramarz (1999) for more detail on the differences between CDD and CDI.

Table 11.3 **Estimated Effect of Real French Minimum Wage Increases on Subsequent Employment Probabilities: Broad Age Categories**

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
	Young Men, Hourly Wage			
Fixed-term contract	−.5129	.0819	.0001	−.0478
Young *stagiaire*	−.8777	.1263	.0001	−.0818
Apprentice	−.1490	.1364	.2747	−.0139
Real wage_t < Real SMIC_t and Not young *stagiaire*	2.9500	2.2341	.1867	.7765
Real wage_t < Real SMIC_t and Young *stagiaire*	9.0935	5.5130	.0990	5.4727
(Real $\text{SMIC}_t \leq$ Real $\text{wage}_t \leq$ Real SMIC_{t+1})*($16 \leq \text{Age}_t \leq 19$)	5.4614	8.5478	.5229	2.0094
(Real $\text{SMIC}_t \leq$ Real $\text{wage}_t \leq$ Real SMIC_{t+1})*($20 \leq \text{Age}_t \leq 24$)	−7.7651	8.2247	.3451	−1.2017
(Real $\text{SMIC}_t \leq$ Real $\text{wage}_t \leq$ Real SMIC_{t+1})*($25 \leq \text{Age}_t \leq 30$)	−33.2708	9.9755	.0009	−4.8928
(Real $\text{SMIC}_{t+1} \leq$ Real $\text{wage}_t \leq$ (1.1*Real SMIC_{t+1}))*($16 \leq \text{Age}_t \leq 19$)	2.9869	5.2162	.5669	1.1201
(Real $\text{SMIC}_{t+1} \leq$ Real $\text{wage}_t \leq$ (1.1*Real SMIC_{t+1}))*($20 \leq \text{Age}_t \leq 24$)	−3.4111	4.2892	.4264	−.4256
(Real $\text{SMIC}_{t+1} \leq$ Real $\text{wage}_t \leq$ (1.1*Real SMIC_{t+1}))*($25 \leq \text{Age}_t \leq 30$)	−3.7791	5.8713	.5198	−.2914
	Young Women, Hourly Wage			
Fixed-term contract	−.9351	.0826	.0001	−.0879
Young *stagiaire*	−1.4152	.1150	.0001	−.1331
Apprentice	−1.0683	.1954	.0001	−.1005
Real wage_t < Real SMIC_t and Not young *stagiaire*	−.8857	2.3804	.7098	−.1604
Real wage_t < Real SMIC_t and Young *stagiaire*	8.3441	5.0400	.0978	4.4279
(Real $\text{SMIC}_t \leq$ Real $\text{wage}_t \leq$ Real SMIC_{t+1})*($16 \leq \text{Age}_t \leq 19$)	−1.6553	9.8606	.8667	−.2759
(Real $\text{SMIC}_t \leq$ Real $\text{wage}_t \leq$ Real SMIC_{t+1})*($20 \leq \text{Age}_t \leq 24$)	−8.7397	6.8185	.1999	−1.2485
(Real $\text{SMIC}_t \leq$ Real $\text{wage}_t \leq$ Real SMIC_{t+1})*($25 \leq \text{Age}_t \leq 30$)	−11.6779	7.8799	.1383	−1.5537
(Real $\text{SMIC}_{t+1} \leq$ Real $\text{wage}_t \leq$ (1.1*Real SMIC_{t+1}))*($16 \leq \text{Age}_t \leq 19$)	−5.1875	7.6857	.4997	−.7447
(Real $\text{SMIC}_{t+1} \leq$ Real $\text{wage}_t \leq$ (1.1*Real SMIC_{t+1}))*($20 \leq \text{Age}_t \leq 24$)	.3164	4.4018	.9427	.0354
(Real $\text{SMIC}_{t+1} \leq$ Real $\text{wage}_t \leq$ (1.1*Real SMIC_{t+1}))*($25 \leq \text{Age}_t \leq 30$)	−1.6632	4.7962	.7288	−.1734

Source: French Labor Force Survey, 1982–89, matched year to year.

Note: Equations estimated by maximum likelihood logit. All equations include indicators for year, education (eight groups), region (Ile de France), and age (three groups), as well as the continuous variables labor force experience (through quartic), seniority, seniority squared, and hourly wage in year t (through cubic). All displayed coefficients except fixed-term contract, young *stagiaire*, and apprentice are equal to the indicated group multiplied by the real percentage increase in the SMIC between years t and $t + 1$ ($1981 = 100$). The coefficients and elasticities show the partial effects on the probability of employment in year $t + 1$, given t. A separate equation was estimated for each demographic panel. Sample sizes are young men, 30,804; young women, 26,434.

detail is not sufficient to speculate on why certain age groups are more affected than others. It is clear from the estimates of the coefficients on the different contract types that all of the types of contract studied here lead to more precarious labor force attachment than an indefinite term contract on average, but the employment promotion contracts (young *stagiaire*) seem to provide relative security for the subminimum population.[12] Looking at these populations in more detail, in particular at what happens to 25-year-olds (who will no longer be eligible for employment promotion contracts the following year), will give us more information on whether the dramatic differences seen between 25–30-year-old and 20–24-year-old men with wages between the two minima are due to the expiration of the protection provided by the employment promotion contracts. Table 11.4 gives these detailed results.

Looking first at the men, the most remarkable feature is in fact the huge negative coefficient affecting 25-year-old men whose wages are between the two minima. This elasticity of -15.9 (expressed as a difference from the marginally above category) and the subsequent negative coefficients for "between" men are consistent with the idea that the minimum wage has a strong negative impact on subsequent employment probabilities. However, the presence of employment promotion contracts, and the reduction in employer social insurance contributions that they imply, helps workers who are under age 25 to retain their jobs in the face of a steadily increasing real SMIC. When workers are no longer eligible for such contracts, their probability of losing their jobs increases dramatically. Relative to the control group of workers marginally above the SMIC, the coefficients for 25- and 26-year-olds are significantly larger. In fact, there is no significant bump in the coefficients at 25 years old for the marginally above workers, suggesting that this phenomenon is only pertinent to minimum wage workers. This further reinforces the interpretation that "between" workers who are eligible for employment promotion contracts are shielded from the negative effects of movements in the SMIC, but "older" young workers are not.

On average, the coefficients for workers between the two SMICs are more negative than for workers marginally over the date *t* SMIC. The average difference (excluding the 25-year-olds) is 7.8, suggesting that the "between" population might be different from the "marginal" population. Unfortunately, none of these differences (except for 25-year-olds) is significant, and in fact, none of the other coefficients for men are significantly different from zero. Although there are also a few significant coefficients in the results for women, interpretation of these results is much more difficult. Although 23-year-old women with wages between the two minima

12. See Bonnal, Fougère, and Sérandon (1997) for an analysis centered on the impact of the youth employment schemes.

Table 11.4 **Estimated Effect of Real French Minimum Wage Increases on Subsequent Employment Probabilities: Detailed Age Categories**

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
	Young Men, Hourly Wage			
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(16 ≤ Age$_t$ ≤ 19)	4.9184	8.5415	.5647	1.8096
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 20)	9.4237	17.3312	.5866	1.8847
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 21)	−14.4978	13.9315	.2980	−2.9995
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 22)	−16.5940	18.9398	.3810	−2.0742
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 23)	−21.2335	19.3804	.2732	−3.6252
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 24)	24.3191	32.6535	.4564	1.1581
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 25)	−63.8672	19.4477	.0010	−15.0276
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 26)	−48.3802	22.1020	.0286	−7.7408
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 27)	−10.1344	41.6355	.8077	−.8108
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(28 ≤ Age$_t$ ≤ 30)	−18.1628	15.4336	.2393	−2.0957
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(16 ≤ Age$_t$ ≤ 19)	2.9091	5.2114	.5767	1.0909
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 20)	−1.2895	7.5889	.8651	−.3281
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 21)	−5.3057	7.6142	.4859	−.8079
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 22)	−14.2510	9.4418	.1312	−1.3538
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 23)	9.8803	11.9823	.4096	.8084
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 24)	5.1411	12.0952	.6708	.3054
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 25)	7.3424	13.7843	.5943	.8811
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 26)	−2.0793	13.6645	.8791	−.1368
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 27)	−6.7963	13.8000	.6224	−.2281
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(28 ≤ Age$_t$ ≤ 30)	−8.2901	8.4234	.3250	−.6564

(continued)

Table 11.4 (continued)

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
		Young Women, Hourly Wage		
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(16 ≤ Age$_t$ ≤ 19)	−1.7276	9.8645	.8610	−.2879
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 20)	38.9118	23.1330	.0926	3.0882
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 21)	−2.5471	12.7138	.8412	−.3069
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 22)	−14.8695	14.2127	.2955	−2.2876
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 23)	−35.7959	14.0221	.0107	−7.8100
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 24)	−26.8167	17.8484	.1330	−4.3098
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 25)	4.9443	23.7480	.8351	.5494
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 26)	−17.3310	15.5787	.2659	−2.3788
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(Age$_t$ = 27)	.3354	18.9002	.9858	.0419
(Real SMIC$_t$ ≤ Real wage$_t$ ≤ Real SMIC$_{t+1}$)*(28 ≤ Age$_t$ ≤ 30)	−18.7008	11.4752	.1032	−2.6715
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(16 ≤ Age$_t$ ≤ 19)	−5.2027	7.6973	.4991	−.7469
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 20)	26.3323	11.6838	.0242	2.7296
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 21)	7.0573	8.8323	.4243	.7876
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 22)	−14.9729	8.3171	.0718	−1.7468
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 23)	−4.4278	9.8576	.6533	−.5009
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 24)	−6.0435	9.7212	.5341	−.6784
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 25)	−.0432	10.5009	.9967	−.0054
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 26)	1.5230	9.9692	.8786	.1488
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(Age$_t$ ≤ 27)	7.7465	12.2241	.5263	.7173
(Real SMIC$_{t+1}$ ≤ Real wage$_t$ ≤ (1.1*Real SMIC$_{t+1}$))*(28 ≤ Age$_t$ ≤ 30)	−7.2571	7.0661	.3044	−.7392

Source: French Labor Force Survey, 1982–89, matched year to year.

Note: Equations estimated by maximum likelihood logit. All equations include indicators for year, education (eight groups), region (Ile de France), and age (ten groups), fixed-term contract, young *stagiaire*, apprentice, paid under the SMIC and young *stagiaire*, and paid under the SMIC and not young *stagiaire*, as well as the continuous variables labor force experience (through quartic), seniority, seniority squared, and hourly wage in year t (through cubic). All displayed coefficients are equal to the indicated group multiplied by the real percentage increase in the SMIC between years t and $t+1$ (1981 = 100). The coefficients and elasticities show the partial effects on the probability of employment in year $t+1$, given employment in year t. Sample sizes are young men, 30,804; young women, 26,434.

are significantly more likely to be nonemployed the following year than women who are paid over the SMIC, the difference from 23-year-old women paid marginally over the SMIC is not significant. And the large, positive coefficients on 20-year-old women, again present in both the "between" and "marginal" populations, is hard to explain. These results may reflect the added opportunities available for women as men go off to perform their military service (and thus withdraw from the labor market), but such an interpretation can neither be accepted not rejected exclusively on the basis of the evidence presented here.

In addition to estimating the conditional logits with "marginally over" the SMIC defined as 1.10 times the SMIC, we also estimated these models with two alternative definitions (1.15 and 1.20 times the SMIC). Table 11.5 analyzes the robustness of the coefficients for the between and marginal categories to these changes in the definition of "marginally over." It seems clear that our results are quite robust to changes in the definition of "marginal."

11.6.2 United States

Table 11.6 shows the results of estimating equation (2) using both the hourly wage measure that excludes income from tips and the measure that

Table 11.5 Robustness of Conditional Logit Results to Variations in Definition of "Marginally over" the Minimum

	Narrow		Medium		Wide	
	Between	Marginally Over	Between	Marginally Over	Between	Marginally Over
French youth						
Men	4.0888	.7317	5.3906	4.0222	6.5107	5.0473
	(6.6196)	(3.8171)	(6.6543)	(2.6087)	(6.7083)	(2.4817)
Women	−6.0281	−.04525	−6.0108	−.4013	−5.8400	−.1178
	(8.2804)	(4.2333)	(8.3134)	(3.1828)	(8.3809)	(3.0601)
U.S. youth						
Men	1.9965	−1.6196	2.0827	−1.9342	1.5043	−2.6988
	(1.6373)	(1.8837)	(1.7436)	(1.7077)	(1.7871)	(1.6751)
Women	3.9599	−.8667	4.6514	−.5443	3.8852	−1.6244
	(1.5578)	(1.8022)	(1.6694)	(1.6615)	(1.7297)	(1.6484)

Sources: French Labor Force Survey, 1982–89, matched year to year, and U.S. Current Population Survey, 1981–87, January–May and September–December, matched year to year.

Note: Coefficients come from logistic regressions conditional on employment at date t for France and date $t+1$ for the United States. For France, the categories are defined as "narrow" = SMIC to 1.10*SMIC, "Medium" = SMIC to 1.15*SMIC, and "wide" = SMIC to 1.20*SMIC. For the United States, the categories are defined as "narrow" = $3.35 to $3.75, "medium" = $3.35 to $4.00, and "wide" = $3.35 to $4.25. For this table, "youth" is defined as ages 25 and under. See notes to tables 11.3, 11.4, 11.6, 11.7, and 11.8 for details on other variables in the regressions. Numbers in parentheses are standard errors.

Table 11.6 **Estimated Effect of Real U.S. Minimum Wage Decreases on Prior Employment Probabilities: Total Labor Market Experience**

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
Young Men, Hourly Wage—No Tips				
Real $\text{wage}_{t+1}<$ Real min_{t+1}	−.4567	2.5368	.8571	−.1498
Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t	−3.0723	1.6532	.0631	−1.3287
Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$	.3153	1.6178	.8455	.0977
(Real $\text{wage}_{t+1} \leq$ Real min_{t+1})*Experience	.2406	.4178	.5648	.4046
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*Experience	−1.4714	.2841	.0001	−2.5115
(Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*Experience	−.8961	.2497	.0003	−1.3746
Young Women, Hourly Wage—No Tips				
Real $\text{wage}_{t+1}<$ Real min_{t+1}	−.0535	2.1856	.9805	−.0340
Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t	−8.3538	1.5107	.0001	−4.8544
Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$	−2.6704	1.5055	.0761	−1.8436
(Real $\text{wage}_{t+1} \leq$ Real min_{t+1})*Experience	−.6488	.2570	.0116	−1.3900
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*Experience	−.9277	.2007	.0001	−1.8917
(Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*Experience	−.8574	.1894	.0001	−1.5564

Young Men, Hourly Wage—With Tips				
Real $\text{wage}_{t+1} <$ Real min_{t+1}	−2.6088	2.4905	.2949	−1.7404
Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t	−4.3814	1.6346	.0074	−2.4823
Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$	−.7521	1.6034	.6390	−.5111
(Real $\text{wage}_{t+1} \leq$ Real min_{t+1})*Experience	.1059	.4154	.7988	.1805
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*Experience	−1.5673	.2849	.0001	−2.6350
(Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*Experience	−.9464	.2491	.0001	−1.4794
Young Women, Hourly Wage—With Tips				
Real $\text{wage}_{t+1} <$ Real min_{t+1}	−3.0938	2.0570	.1326	−1.8775
Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t	−9.1702	1.4879	.0001	−5.2774
Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$	−3.3196	1.4939	.0263	−2.2658
(Real $\text{wage}_{t+1} \leq$ Real min_{t+1})*Experience	−.7841	.2570	.0023	−1.7565
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*Experience	−.9762	.2009	.0001	−1.9923
(Real $\text{min}_t \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*Experience	−.8851	.1894	.0001	−1.6186

Source: Current Population Survey, 1981–87, January–May and September–December, matched year to year.

Note: Equations estimated by maximum likelihood logit. All equations include indicators for year, region (three groups), nonwhite, and married; and years of schooling, labor force experience (through quartic), and log hourly real wage (1982 prices, through cubic). All displayed coefficients are equal to the indicated group times the real decrease (absolute value of the change in logarithms) in the minimum wage between years t and $t+1$. The coefficients and elasticities show the partial effects on the probability of employment in year t, given employment in year $t+1$. A separate equation was estimated for each panel. Sample sizes are young men, 41,001; young women, 38,992.

includes income from tips, and interacting with total labor market experience instead of age.[13] In every case, individuals who are employed in year $t + 1$ were more likely to have been unemployed or not in the labor force in t if their real wage in $t + 1$ was between the real minimum wage in years t and $t + 1$. The magnitudes of these effects are large, with elasticities for men with zero experience of -1.42 to -1.97 and for women with no experience of -3.01. Once again, we refer to comparisons with the "marginal" group—that is, workers who are paid marginally above the old (date t) minimum wage—to get at the direct effect of movements in the real minimum wage on transitions into employment. By weighting the different experience groups, a decrease in the real minimum wage of 1 percent between $t - 1$ and t is related to an increased probability of having been nonemployed at $t - 1$ of 2.2 percent (in difference from the marginal workers) for those men who are paid between the t and $t + 1$ minimum wages. These results are consistent with the neoclassical idea that decreases in the real minimum wage make nonemployed workers easier to employ and these workers enter disproportionately between the two minimum wages. This decreases the share of those employed at date $t + 1$ who were employed at date t for the "between" group more than for other groups.

It is interesting to note the differences, or rather lack of differences, between the results that measure wages with and without tips. None of the qualitative results seem sensitive to the manner in which we define wages; however, some intuition can be gleaned from how the coefficients seem to shift when passing from measures without tips to measures with tips. All of the coefficients shown in table 11.6 become more negative when tips are included in the wage measure. This is also consistent with the standard neoclassical model, which would imply that the measure with tips more accurately describes a worker's marginal productivity and would conclude that the less significant coefficients in the estimation without tips are affected by measurement error. Nevertheless, due to the lack of any qualitative difference between the results with and without tips, and because our measure without tips uses reported rather than constructed data,[14] the rest of our results for the United States will be based on the wage measure that excludes tips.

Table 11.7 reestimates equation (2) using the broad age categories, as in table 11.3. As was suggested by the negative coefficients on the experience interaction terms in table 11.6, the effects of the minimum wage worsen as young workers get older. The differences between workers paid between the two minima and workers paid marginally over the t minimum are still

13. Appendix table 11A.5 provides descriptive statistics for the U.S. data used in these regressions.

14. Welch (1997) provides evidence on various sorts of measurement error in the CPS and hints that hours are likely to be a greater source of measurement error than wages.

Table 11.7 **Estimated Effect of Real U.S. Minimum Wage Increases on Prior Employment Probabilities (excluding tips): Broad Age Categories**

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
Young Men, Hourly Wage				
Real wage$_{t+1}$< Real min$_{t+1}$	.6119	1.9147	.7493	.2007
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(16 ≤ Age$_t$ ≤ 19)	−6.1455	1.3807	.0001	−2.9233
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(20 ≤ Age$_t$ ≤ 24)	−11.8902	1.9536	.0001	−4.2095
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(25 ≤ Age$_t$ ≤ 30)	−19.4188	3.1495	.0001	−5.9588
(Real min$_t$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(16 ≤ Age$_t$ ≤ 19)	−.9696	1.3901	.4855	−.3767
(Real min$_t$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(20 ≤ Age$_t$ ≤ 24)	−5.9107	1.7693	.0008	−1.4697
(Real min$_t$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(25 ≤ Age$_t$ ≤ 30)	−9.8243	2.4330	.0001	−1.8055
Young Women, Hourly Wage				
Real wage$_{t+1}$< Real min$_{t+1}$	−3.2195	1.6924	.0571	−1.1762
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(16 ≤ Age$_t$ ≤ 19)	−9.1433	1.3730	.0001	−4.3346
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(20 ≤ Age$_t$ ≤ 24)	−14.0812	1.6675	.0001	−4.8644
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(25 ≤ Age$_t$ ≤ 30)	−19.8125	1.8812	.0001	−7.1220
(Real min$_t$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(16 ≤ Age$_t$ ≤ 19)	−3.0577	1.4261	.0320	−1.1732
(Real min$_t$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(20 ≤ Age$_t$ ≤ 24)	−8.4481	1.4757	.0001	−2.2399
(Real min$_t$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(25 ≤ Age$_t$ ≤ 30)	−12.5349	1.5423	.0001	−3.2334

Source: Current Population Survey, 1981–87, January–May and September–December, matched year to year.

Note: Equations estimated by maximum likelihood logit. All equations include indicators for year, region (three groups), nonwhite, married, and age (three groups); and years of schooling, labor force experience (through quartic), and log hourly real wage (1982 prices, through cubic). All displayed coefficients are equal to the indicated group times the real decrease (absolute value of the change in logarithms) in the minimum wage between years t and $t+1$. The coefficients and elasticities show the partial effects on the probability of employment in year t, given employment in year $t+1$. A separate equation was estimated for each demographic panel. Sample sizes are young men, 41,001; young women, 38,992.

significant for all age groups, and the elasticities are still large. For the oldest age group, a decrease of 1 percent in the real minimum wage at t is associated with a 5.96 percent higher chance that a given "between" worker came from nonemployment, whereas such a change is associated with only an 1.81 percent higher chance for "marginal" workers. Unlike the French case, although 25–30-year-olds with date $t + 1$ wages between the two minima have a higher chance of having come from nonemployment than do 20–24-year-olds, the difference is not nearly as dramatic. This is not surprising, as there existed no nationwide employment promotion schemes in the United States in the 1980s that would have induced effects similar to the French case.

One might think that our approach of considering *previous* employment in the United States could be subject to the possibility, especially among young people, that many of the transitions from nonemployment to employment are first jobs after the end of schooling.[15] Since we control for schooling as a set of regressors reflecting different levels of educational attainment, looking at the pattern of age coefficients for "between" workers and "marginal" workers should allow us to ignore such considerations to the extent that entry into the labor force does not occur disproportionately in a particular wage category. Table 11.8, which provides our conditional logit analysis at the same level of aggregation as table 11.4, therefore allows us to concentrate more precisely on how minimum wage movements affect the stability of early career employment at different points in the wage distribution.

As was the case in our earlier results, the probability that a worker came from nonemployment is higher among the set of workers with date $t + 1$ real wages between the two minima than among the set of workers with date $t + 1$ real wages marginally above the date t real minimum. The same holds true for a comparison of "between" workers with workers earning substantially more than the date t real minimum, and these differences are often significant. Despite a lot of variation across the different ages, there appears to be a secular trend toward a higher and higher share of workers coming from nonemployment as age increases, and this trend is steeper among "between" workers than among "marginal" workers, particularly for young men. This is not the case in France, and it may suggest that information is revealed faster in the United States and that as workers age, the sorts of low-wage jobs they can find become increasingly precarious.

Since there do not exist systematic, targeted programs that should affect transitions among young people throughout the United States in the same manner (with the exception of education), interpretation of these coefficients is not as straightforward as in the French case. However, if (as men-

15. See Topel and Ward (1992), among others, for an analysis of early career mobility in the United States.

Table 11.8 Estimated Effect of Real U.S. Minimum Wage Increases on Prior Employment Probabilities (excluding tips): Detailed Age Categories

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
Young Men, Hourly Wage				
Real wage$_{t+1}$ < Real min$_{t+1}$	.2962	1.9152	.8771	.0971
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(16 ≤ Age$_t$ ≤ 19)	−6.5106	1.3857	.0001	−3.0970
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 20)	−11.6092	3.1697	.0002	−4.4924
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 21)	−9.0680	3.4352	.0083	−3.2645
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 22)	−7.3453	4.7357	.1209	−2.0986
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 23)	−22.0209	5.2597	.0001	−8.4499
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 24)	−15.1148	5.2426	.0039	−4.6784
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 25)	−16.6557	6.2664	.0079	−4.7588
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 26)	−17.9004	6.9347	.0098	−5.3701
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(Age$_t$ ≤ 27)	−15.9424	8.5432	.0620	−5.1813
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real min$_t$)*(28 ≤ Age$_t$ ≤ 30)	−22.0514	4.5378	.0001	−6.9252
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(16 ≤ Age$_t$ ≤ 19)	−1.2309	1.3918	.3765	−.4783
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 20)	−4.7686	3.0687	.1202	−1.2724
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 21)	−4.4151	3.3797	.1914	−1.2184
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 22)	−5.2612	3.9467	.1825	−1.2314
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 23)	−9.3349	4.0392	.0208	−2.0277
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 24)	−8.6274	4.7811	.0712	−1.9071
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 25)	−6.4574	4.8991	.1875	−1.1170
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 26)	−8.4370	5.7535	.1425	−1.5576
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(Age$_t$ ≤ 27)	−12.1263	5.3991	.0247	−2.4804
(Real min$_{t+1}$ ≤ Real wage$_{t+1}$ ≤ Real ($4.00)$_t$)*(28 ≤ Age$_t$ ≤ 30)	−10.7899	3.5679	.0025	−1.9561

(*continued*)

Table 11.8 (continued)

Name of Effect	Coefficient	Standard Error	p-Value	Elasticity
Young Women, Hourly Wage				
Real wage_{t+1} < Real min_{t+1}	−3.7559	1.6913	.0264	−1.3722
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($16 \leq \text{Age}_t \leq 19$)	−9.8220	1.3730	.0001	−4.6564
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 20$)	−12.2205	2.8456	.0001	−4.6320
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 21$)	−12.8276	3.1141	.0001	−4.6853
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 22$)	−13.4058	3.6339	.0002	−4.4009
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 23$)	−14.1311	4.2524	.0009	−4.1771
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 24$)	−14.0301	4.2585	.0010	−4.1895
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 25$)	−23.5188	4.2544	.0001	−9.5817
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 26$)	−18.8257	4.0242	.0001	−6.4372
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($\text{Age}_t \leq 27$)	−20.1282	4.6770	.0001	−6.8814
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real min_t)*($28 \leq \text{Age}_t \leq 30$)	−19.6787	2.4999	.0001	−6.9948
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($16 \leq \text{Age}_t \leq 19$)	−3.4490	1.4233	.0154	−1.3234
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 20$)	−2.3108	2.8808	.4225	−.5968
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 21$)	−4.9630	2.9318	.0905	−1.3019
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 22$)	−9.1566	3.0945	.0031	−2.5897
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 23$)	−13.4398	3.2502	.0001	−3.4858
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 24$)	−14.1707	3.4026	.0001	−3.7468
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 25$)	−16.7514	3.1826	.0001	−4.7081
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 26$)	−7.2195	3.7185	.0522	−1.5576
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($\text{Age}_t \leq 27$)	−6.5597	3.5805	.0669	−1.4072
(Real $\text{min}_{t+1} \leq$ Real $\text{wage}_{t+1} \leq$ Real $(\$4.00)_t$)*($28 \leq \text{Age}_t \leq 30$)	−15.0802	2.0915	.0001	−4.2146

Source: Current Population Survey, 1981–87, January–May and September–December, matched year to year.

Note: Equations estimated by maximum likelihood logit. All equations include indicators for year, region (three groups), nonwhite, married, and age (ten groups); and years of schooling, labor force experience (through quartic), and log hourly real wage (1982 prices, through cubic). All displayed coefficients are equal to the indicated group times the real decrease (absolute value of the change in logarithms) in the minimum wage between years t and $t+1$. The coefficients and elasticities show the partial effects on the probability of employment in year t, given employment in year $t+1$. A separate equation was estimated for each demographic panel. Sample sizes are young men, 41,001; young women, 38,992.

tioned above) the coefficients corresponding to a given age are particularly strong, and if this age corresponds to the age at which many students typically finish a certain diploma, one might conclude that the coefficients are capturing disproportionate entry into the labor force at particular places in the wage distribution. Unfortunately, the most remarkable coefficients (23 years old for men and 25 years old for women) are not concurrent with ages at which a significant portion of the future workforce is in their last year of schooling. There does not seem to be any clear interpretation for the particular age pattern of the coefficients in the United States.

Finally, to promote comparability between our analysis, which is done conditional on the employment state in either year t (France) or year $t + 1$ (United States), and other analyses, which consider the effects of the minimum wage unconditional on the previous or future employment state, we compute the implied unconditional elasticities implied by our estimates. To calculate an unconditional elasticity we apply Bayes law to obtain the relation between the forms of the analysis equations we used for France and the United States. Hence, we have

$$\begin{aligned} &\Pr[e_{t+1} = 1 | e_t = 1, rmiw_t, rmiw_{t+1}] \\ &= \Pr[e_t = 1 | e_{t+1} = 1, rmiw_t, rmiw_{t+1}] \frac{\Pr[e_{t+1} = 1 | rmiw_t, rmiw_{t+1}]}{\Pr[e_t = 1 | rmiw_t]}. \end{aligned} \tag{3}$$

To calculate the elasticity we use the following derivative formula:

$$\begin{aligned} \frac{\partial \ln \Pr[e_{t+1} = 1]}{\partial \ln rmiw_{t+1}} = &\frac{\partial \ln \Pr[e_{t+1} = 1 | e_t = 1, rmiw_t, rmiw_{t+1}]}{\partial \ln rmiw_{t+1}} \\ &- \frac{\partial \ln \Pr[e_t = 1 | e_{t+1} = 1, rmiw_t, rmiw_{t+1}]}{\partial \ln rmiw_{t+1}}. \end{aligned} \tag{4}$$

Notice that the derivative in equation (4) simplifies because the denominator in the ratio of unconditional probabilities in equation (3) does not depend on the future minimum wage. The right-hand side of equation (4) has two terms. For France, we can estimate only the first of these two terms because the real minimum wage is always increasing. The conditions necessary for estimating the second term occur in the United States, where the real minimum wage is always decreasing. To estimate the unconditional elasticity in equation (4) we must make an assumption regarding the term that cannot be estimated in the particular country. We assume that this term is zero, which means that increases in the real minimum wage do not change the rate at which nonemployed workers become employed and, conversely, decreases in the real minimum wage do not change

Table 11.9 **Elasticity Estimates for Young Men and Women: Rate of Change of Employment Probability for 1 Percent Increase in Real Minimum Wage**

	France	United States
Conditional (aggregated over age groups)		
Young men	−2.489	−2.234
Young women	−1.044	−1.873
Unconditional (aggregated over age groups)		
Young men	−.203	−.123
Young women	−.108	−.127

Source: France, table 11.3, figs. 11.6 and 11.7, and Labor Force Survey. United States, table 11.7, figs. 11.8 and 11.9, and Current Population Survey.

Note: The conditional elasticity is the weighted average of the elasticities for each age group in tables 11.3 and 11.7 reported as the difference between the elasticity for the "at minimum" group as compared to the "marginally above" group. The unconditional elasticity is an estimate of the rate of change of the employment probability in period $t+1$ given a 1 percent increase in the real minimum wage between periods t and $t+1$.

the rate at which employed workers at t remain employed at $t + 1$. Our results are summarized in table 11.9. To take advantage of the structure of our estimates in tables 11.3 and 11.7, we computed the required conditional elasticities in equation (4) according to the following formula for France, which assumes that the appropriate control group is individuals who are marginally over the minimum wage:

$$\frac{\partial \ln \Pr[e_{t+1} = 1 | e_t = 1]}{\partial \ln rmiw_{t+1}}$$

$$= \Pr[\text{at minimum}] \sum_{\ell} \left[\frac{\partial \ln \Pr[e_{t+1} = 1 | e_t = 1, \ell, \text{at minimum}]}{\ell \ln rmiw_{t+1}} - \frac{\partial \ln \Pr[e_{t+1} = 1 | e_t = 1, \ell, \text{marginal}]}{\ell \ln rmiw_{t+1}} \right] \Pr[\ell],$$

where the summation is taken over the three age groups. We use the comparable formula for the United States.

11.7 Conclusion

This paper has shown that for young people in both France and the United States, movements in the real minimum wage are associated with significant employment effects, typically in the direction predicted by competitive labor market theory. In France, as the real SMIC increased over the period 1982–89, a certain share of young French workers had real wages that fell between the increasing consecutive real minimum wages.

For workers in this situation, subsequent employment probabilities fell significantly. However, participation in employment promotion programs seemed to shield these workers from some of the effects of the increasing real SMIC, and when this eligibility ended, the probability of subsequent nonemployment shot up dramatically. In the United States, a comparable effect of a real minimum wage moving in the opposite direction occurred, as many workers had market wage rates that were passed by the declining real minimum wage over the period 1981–87. American workers whose current real wage rate would have been below the real minimum wage in earlier periods were much less likely to have been employed in those earlier periods.

By comparing effects of minimum wage movements on workers employed at the minimum with the effects on those employed marginally above it, we identify the direct effects of the minimum wage, as distinct from heterogeneity across the wage distribution in labor force attachment and response to macroeconomic shocks. We suppose that these workers have identical labor supply behavior, but they also have much higher subsequent reemployment probabilities in France as well as much higher prior employment probabilities in the United States. Within the youth population, these strong effects increase with age in the United States, and the pattern in France is dominated by eligibility for employment promotion contracts. Across the population as whole, however, our multinomial logit results suggest that in both countries, it is youths who are most affected by movements in the real minimum wage.

Even if the conditional elasticities in question are large, the at-risk groups (workers between two minimum wages) are relatively small—8 percent of young men and 10 percent of young women in France, 6 percent of young men and 7 percent of young women in the United States. Thus overall unconditional elasticities tend to be much lower than the elasticities conditional on being between the two minima. If the relevant policy question concerns the impact of the minimum wage on those individuals most likely to be affected by it (i.e., those currently paid at the minimum wage), our results suggest that there are much larger negative employment effects on this group, especially as compared to the group in the wage distribution marginally above the minimum, than other research has found.

Appendix

Table 11A.1 **Statistical History of the *Salaire Minimum Interprofessionnel de Croissance* (SMIC)**

Year	Statutory Hours per Month	Gross: Hourly SMIC (francs)	Gross: Real Hourly SMIC (1970) (francs)	Gross: Monthly SMIC (francs)	Net Monthly SMIC (francs)	Monthly Total Compensation Cost (francs)	Employee Payroll Tax Rate (% at SMIC)	Employer Payroll Tax Rate (% at SMIC)	Consumer Price Index (1970 = 100)
1951	173.3	0.89	1.95	154.41	145.15	195.78	6.00	26.79	45.60
1952	173.3	1.00	1.96	173.33	162.93	220.74	6.00	27.35	50.98
1953	173.3	1.00	1.98	173.33	182.33	222.47	6.00	28.35	50.39
1954	173.3	1.15	2.29	199.98	187.98	256.67	6.00	28.35	50.21
1955	173.3	1.25	2.46	216.45	203.46	277.81	6.00	28.35	50.80
1956	173.3	1.26	2.43	218.40	205.30	280.32	6.00	28.35	51.80
1957	173.3	1.29	2.42	223.78	210.35	287.22	6.00	28.35	53.21
1958	173.3	1.46	2.39	253.87	238.64	319.50	6.00	28.85	61.19
1959	173.3	1.58	2.43	270.62	253.84	349.51	6.20	29.15	64.98
1960	173.3	1.61	2.39	279.19	261.88	360.57	6.20	29.15	67.40
1961	173.3	1.64	2.36	284.69	267.04	370.52	6.20	30.15	69.59
1962	173.3	1.72	2.36	298.77	278.45	393.33	7.05	31.65	72.91
1963	173.3	1.84	2.41	319.62	297.09	418.88	7.05	31.05	76.38
1964	173.3	1.89	2.39	328.27	305.13	430.20	7.05	31.05	78.98
1965	173.3	1.97	2.43	342.28	318.15	448.56	7.05	31.05	80.98
1966	173.3	2.06	2.48	358.27	331.15	468.00	7.05	31.36	83.22
1967	173.3	2.13	2.49	368.32	339.66	498.45	8.15	35.33	85.41
1968	173.3	2.68	3.00	484.81	426.84	617.17	8.17	32.78	89.28
1969	173.3	3.16	3.32	548.16	503.32	728.07	8.18	32.82	95.12

1970	173.3	3.42	3.42	591.92	543.50	786.13	8.18	32.81	100.00
1971	173.3	3.76	3.56	651.72	598.15	867.31	8.22	33.08	105.52
1972	173.3	4.19	3.74	725.96	668.00	971.62	8.26	33.84	111.99
1973	173.3	4.95	4.12	858.27	786.52	1,151.28	8.36	34.14	120.20
1974	173.3	6.10	4.46	1,053.74	967.78	1,421.63	8.42	34.53	136.71
1975	173.3	7.26	4.75	1,260.25	1,150.86	1,711.87	8.68	35.82	152.80
1976	173.3	8.34	4.98	1,466.01	1,306.18	1,981.47	9.67	37.03	167.49
1977	173.3	9.40	5.13	1,629.59	1,464.19	2,239.06	10.15	37.40	183.22
1978	173.3	10.61	5.31	1,839.61	1,650.68	2,536.45	10.27	37.88	199.82
1979	173.3	11.94	5.40	2,068.69	1,817.14	2,843.62	12.14	38.91	221.30
1980	173.3	13.80	5.49	2,391.67	2,085.54	3,324.42	12.80	39.00	251.30
1981	173.3	16.30	5.72	2,824.41	2,478.98	3,925.93	12.23	39.00	285.00
1982	169.0	19.17	6.02	3,323.46	2,892.07	4,623.60	12.98	39.12	318.70
1983	169.0	21.50	6.16	3,725.87	3,216.92	5,221.43	13.66	40.14	349.29
1984	169.0	23.53	6.27	4,077.88	3,465.79	5,693.33	15.01	39.62	375.19
1985	169.0	25.44	6.41	4,335.00	3,676.51	6,056.88	15.19	39.72	397.04
1986	169.0	26.53	6.51	4,482.87	3,777.27	6,270.64	15.74	39.88	407.62
1987	169.0	27.60	6.56	4,663.84	3,894.77	6,528.91	16.49	39.99	420.43
1988	169.0	28.65	6.64	4,791.71	3,977.60	6,715.10	16.99	40.14	431.74
1989	169.0	29.54	6.60	4,991.42	4,093.46	6,943.58	17.99	39.11	447.33
1990	169.0	30.80	6.66	5,205.20	4,269.83	7,182.13	17.97	37.89	462.38
1991	169.0	32.30	6.77	5,458.70	4,547.95	7,527.66	17.39	37.90	477.20
1992	169.0	33.58	6.87	5,674.46	4,606.38	7,860.94	17.98	38.53	488.60
1993	169.0	34.45	6.91	5,821.21	4,794.70	7,945.37	18.38	36.49	498.86
1994	169.0	35.20	6.92	5,947.96	4,881.38	7,981.57	18.64	34.19	508.84

Source: Friez and Julhès (1998).

Note: Data for 1950–69 are for the earlier minimum wage system (*salaire minimum interprofessionnel garantie*).

Table 11A.2 **Multinomial Logit Results for France**

	Transition																	
	Men									Women								
Effect	U-N	A-N	A-E	M-N	M-E	O-N	O-E	N-N	N-E	U-N	A-N	A-E	M-N	M-E	O-N	O-E	N-N	N-E
Intercept	1.39	−2.98	−3.81	−5.95	−5.77	−3.96	−3.36	−.42	−4.52	2.44	−1.55	−.93	−1.71	−1.02	1.03	2.77	2.08	−3.10
	(.91)	(.87)	(.53)	(.75)	(.45)	(.47)	(.40)	(.41)	(1.02)	(.73)	(.59)	(.37)	(.66)	(.35)	(.48)	(.32)	(.31)	(.96)
1982	−.11	−.13	−.18	.13	.01	.38	.40	.26	.32	.02	.04	−.02	.19	.13	.35	.35	.23	.30
	(.12)	(.12)	(.08)	(.09)	(.07)	(.07)	(.06)	(.06)	(.13)	(.10)	(.08)	(.05)	(.08)	(.05)	(.06)	(.04)	(.04)	(.12)
1983	−.28	−.11	−.23	.41	.18	.79	.79	.53	.64	−.05	.13	.00	.53	.40	.78	.69	.45	.64
	(.12)	(.12)	(.08)	(.09)	(.06)	(.06)	(.06)	(.06)	(.13)	(.10)	(.07)	(.05)	(.08)	(.05)	(.06)	(.04)	(.04)	(.13)
1984	−.07	−.14	−.21	.23	.13	.62	.57	.30	.45	.03	.08	.00	.40	.28	.65	.51	.34	.52
	(.12)	(.12)	(.08)	(.09)	(.07)	(.07)	(.06)	(.06)	(.13)	(.10)	(.07)	(.05)	(.08)	(.05)	(.06)	(.04)	(.04)	(.13)
1985	−.16	.06	−.07	.64	.37	1.00	.94	.59	.57	−.07	.30	.11	.62	.44	.87	.72	.51	.46
	(.12)	(.12)	(.08)	(.09)	(.06)	(.06)	(.06)	(.06)	(.13)	(.10)	(.08)	(.05)	(.08)	(.05)	(.06)	(.04)	(.04)	(.12)
1986	−.30	.00	−.08	.50	.33	.82	.81	.44	.64	−.23	.13	.08	.42	.36	.61	.57	.40	.30
	(.12)	(.12)	(.08)	(.09)	(.06)	(.07)	(.06)	(.06)	(.13)	(.10)	(.07)	(.05)	(.08)	(.05)	(.06)	(.04)	(.04)	(.12)
1987	−.39	.00	.05	.45	.34	.80	.79	.43	.33	−.26	.23	.12	.36	.37	.63	.57	.41	.38
	(.12)	(.12)	(.08)	(.09)	(.06)	(.07)	(.06)	(.06)	(.13)	(.10)	(.08)	(.05)	(.08)	(.05)	(.06)	(.04)	(.04)	(.12)
1988	−.34	.24	.14	.66	.47	.99	.92	.56	.49	−.24	.37	.26	.64	.54	.77	.74	.56	.46
	(.12)	(.13)	(.08)	(.11)	(.07)	(.07)	(.06)	(.06)	(.14)	(.10)	(.09)	(.05)	(.09)	(.05)	(.06)	(.05)	(.04)	(.14)
Baccalauréat	−.26	−.31	−.49	−.35	−.50	.29	.62	−.02	.42	.08	.01	.04	.30	.38	1.06	1.34	.38	1.04
	(.16)	(.13)	(.08)	(.13)	(.07)	(.07)	(.06)	(.06)	(.13)	(.11)	(.09)	(.06)	(.10)	(.05)	(.06)	(.05)	(.05)	(.12)
Age = 22–25	.60	−.68	−.81	−.90	−1.01	−1.50	−1.78	−1.29	−1.02	.77	.04	−.22	.00	−.19	−.53	−.69	−.39	.12
	(.11)	(.13)	(.07)	(.12)	(.06)	(.08)	(.06)	(.06)	(.16)	(.10)	(.09)	(.06)	(.11)	(.05)	(.08)	(.05)	(.05)	(.17)

Source: French Labor Force Survey, 1982–89, matched year to year.

Note: Equations estimated by multinomial logit. Transitions identified by U = under the minimum, A = at the minimum, M = marginally over the minimum, O = over the minimum, N = nonemployment, and E = employment. In addition to the coefficients shown, the regression included indicator variables for region (Ile de France), eight education categories, eight age categories, and three entry cohorts. The reference transition was U-E. The reference categories for the indicator variables were year = 1981, education = no degree, age = 41–50 years old, and year of entry into labor market = before 1961. Separate equations were estimated for men and women. Sample sizes were men, 145,646; women, 166,716. Numbers in parentheses are standard errors.

Table 11A.3 Multinomial Logit Results for the United States

	Transition																	
	Men									Women								
Effect	N-U	N-A	E-A	N-M	E-M	N-O	E-O	N-N	E-N	N-U	N-A	E-A	N-M	E-M	N-O	E-O	N-N	E-N
Intercept	−1.61	.11	.30	.80	1.20	4.28	6.60	−.80	−.39	−.89	2.74	.81	3.45	1.94	5.56	6.91	.34	1.40
	(1.06)	(.86)	(.58)	(.91)	(.54)	(.51)	(.44)	(.44)	(.46)	(.67)	(.57)	(.40)	(.61)	(.38)	(.42)	(.32)	(.32)	(.34)
1982	.04	.17	.02	.13	.10	.26	.07	.16	.06	.06	.09	.04	.09	−.02	.14	.09	.06	.02
	(.07)	(.05)	(.04)	(.06)	(.04)	(.04)	(.04)	(.04)	(.04)	(.05)	(.04)	(.03)	(.04)	(.03)	(.03)	(.03)	(.03)	(.03)
1983	−.07	.22	.00	.18	.05	.34	.09	.15	−.05	.00	.13	.00	.07	−.06	.17	.11	.05	−.01
	(.08)	(.05)	(.04)	(.06)	(.04)	(.04)	(.04)	(.04)	(.04)	(.05)	(.04)	(.03)	(.04)	(.03)	(.03)	(.03)	(.03)	(.03)
1984	.01	.22	.05	.20	.08	.30	.16	.20	.03	.01	.16	−.06	.12	−.01	.24	.19	.10	.05
	(.09)	(.06)	(.06)	(.07)	(.05)	(.05)	(.05)	(.05)	(.05)	(.06)	(.04)	(.04)	(.05)	(.04)	(.04)	(.03)	(.03)	(.03)
1985	.01	.16	.02	.22	.17	.34	.27	.27	.06	−.08	.11	−.12	.21	−.04	.26	.22	.06	.02
	(.12)	(.08)	(.07)	(.09)	(.07)	(.07)	(.06)	(.06)	(.06)	(.08)	(.06)	(.05)	(.06)	(.05)	(.05)	(.04)	(.04)	(.04)
1986	−.06	.09	−.08	.20	.06	.35	.20	.21	.05	−.05	.09	−.07	.16	−.06	.36	.29	.10	.08
	(.08)	(.05)	(.05)	(.06)	(.05)	(.05)	(.04)	(.04)	(.04)	(.05)	(.04)	(.03)	(.04)	(.03)	(.03)	(.03)	(.03)	(.03)
High school	−.12	.13	.08	.16	.12	.49	.75	.12	.38	−.14	.05	−.02	.19	.20	.37	.65	−.05	.22
	(.09)	(.07)	(.05)	(.07)	(.05)	(.05)	(.04)	(.04)	(.04)	(.06)	(.05)	(.04)	(.06)	(.04)	(.04)	(.03)	(.03)	(.03)
Age = 22–25	.09	.13	.10	.25	.14	.20	−.11	−.86	−.71	.13	.63	.21	.76	.25	.52	.21	−.64	−.21
	(.27)	(.21)	(.14)	(.22)	(.13)	(.12)	(.11)	(.11)	(.11)	(.16)	(.13)	(.10)	(.14)	(.09)	(.10)	(.08)	(.08)	(.08)

Source: Current Population Survey, 1981–87, January–May and September–December, matched year to year.

Note: Equations estimated by multinomial logit. Transitions identified by U = under the minimum, A = at the minimum, M = marginally over the minimum, O = over the minimum, N = nonemployment, and E = employment. In addition to the coefficients shown, the regression included indicator variables for six education categories, eight age categories, and three entry cohorts. The reference transition was E-U. The reference categories for the indicator variables were year = 1981, education = no diploma, age = 61 years old or older, and year of entry into labor market = before 1961. Separate equations were estimated for men and women. Sample sizes were men, 162,073; women, 199,682. Numbers in parentheses are standard errors.

Table 11A.4 **Descriptive Statistics Conditional on Employment: France**

	Entire Population				Youth			
	Men		Women		Men		Women	
Indicator	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
Age	37.5769	10.7732	36.9156	10.9116	25.0472	3.6006	24.9998	3.4430
Seniority	10.9995	9.0239	9.5347	8.1790	4.2628	5.7013	4.4259	6.1117
Experience	20.5879	14.7257	19.6161	13.1948	7.5144	12.0523	6.8722	10.8535
Fixed-term contract	.0144	.1191	.0178	.1322	.0342	.1697	.0357	.1756
Apprentice	.0064	.0795	.0021	.0462	.0216	.0671	.0065	.0389
Youth *stagiaire*	.0037	.0611	.0057	.0750	.0124	.0818	.0173	.0986
Paris region	.2031	.4023	.2350	.4240	.1907	.3924	.2217	.4151
Year = 1988	.0692	.2538	.0724	.2592	.0641	.2449	.0629	.2428
Year = 1987	.1386	.3455	.1444	.3515	.1343	.3410	.1349	.3415
Year = 1986	.1374	.3443	.1420	.3490	.1342	.3409	.1352	.3419
Year = 1985	.1413	.3484	.1430	.3501	.1385	.3453	.1401	.3471
Year = 1984	.1437	.3508	.1399	.3469	.1427	.3498	.1402	.3471
Year = 1983	.1455	.3526	.1411	.3481	.1481	.3552	.1490	.3561
Year = 1982	.1479	.3550	.1443	.3514	.1554	.3623	.1567	.3635
No education	.2407	.4275	.1821	.3859	.2361	1.4579	.1635	1.5493
Elementary school	.1845	.3879	.2114	.4083	.0978	.2941	.0963	.2908
Junior high school	.0610	.2394	.0920	.2890	.0794	.2701	.1055	.3072
Basic vocational/technical	.2997	.4581	.2344	.4236	.3989	.4896	.3129	.4632
Advanced vocational/technical school	.0509	.2199	.0689	.2533	.0477	.2103	.0887	.2831
Baccalauréat (high school)	.0434	.2038	.0709	.2566	.0460	.2050	.0921	.2878
Technical college or university	.0495	.2169	.0841	.2776	.0554	.2157	.1008	.2891
Grad school or postcollege professional school	.0639	.2445	.0541	.2263	.0387	.1727	.0401	.1815
Employed next period?	.9285	.2577	.9209	.2699	.9068	.2666	.9060	.2871
Observations under SMIC and *Stagiaire*	329		422		329		424	
Observations under SMIC and Not *stagiaire*	5,548		9,826		3,256		3,617	
Observations between two real SMICs	849		1,292		494		645	
Observations marginally over SMIC	4,146		5,441		2,155		2,206	
Total observations	104,081		80,993		30,804		26,434	

Source: French Labor Force Survey, 1982–89, matched year to year.

Note: S.D. = standard deviation. "Youth" is defined as ages 30 and under.

Table 11A.5 **Descriptive Statistics Conditional on Employment: United States**

	Entire Population				Youth			
	Men		Women		Men		Women	
Indicator	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
Years of education	12.8629	2.8842	12.8531	2.4770	12.7341	2.2514	12.9628	2.0845
Experience	20.5188	12.9620	20.1023	12.9669	7.3809	4.0105	6.9846	3.9707
Nonwhite	.1156	.3198	.1399	.3469	.1177	.3223	.1312	.3376
Married	.7055	.4558	.5973	.4905	.4366	.4960	4342	.4957
Year = 1981	.2005	.4004	.1937	.3952	.2048	.4035	.2019	.4014
Year = 1982	.2004	.4003	.2000	.4000	.2007	.4005	.2061	.4045
Year = 1983	.2049	.4036	.2023	.4017	.2076	.4056	.2043	.4032
Year = 1984	.1133	.3169	.1151	.3191	.1137	.3174	.1138	.3176
Year = 1985	.0706	.2561	.0721	.2587	.0694	.2542	.0695	.2542
Northeastern region	.2326	.4225	.2316	.4219	.2249	.4175	.2304	.4211
North-central region	.2618	.4396	.2613	.4393	.2679	.4429	.2684	.4431
Southern region	.3215	.4670	.3249	.4683	.3251	.4684	.3176	.4655
Employed previous period?	.9170	.2760	.8705	.3358	.8397	.3668	.7977	.4017
Observations under minimum wage	2,571		5,367		1,475		2,481	
Observations between two real minimum wages	4,085		7,645		3,177		4,434	
Observations marginally over minimum wage	6,799		13,218		4,664		6,097	
Total observations	121,356		110,287		41,001		38,993	

Source: Current Population Survey, 1981–87, January–May and September–December, matched year to year.
Note: S.D. = standard deviation. "Youth" is defined as ages 30 and under.

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