Minimum Wages and Employment in France and the United States

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Abstract

We use longitudinal individual wage and employment data in France and the United States to investigate the effect of changes in the real minimum wage rate on an individual's employment status. We focus on workers employed at wages close enough to the minimum in a reference year as to be illegal in an adjacent comparison year as a result of movements in the real minimum wage. We find that movements in the American real minimum wage are associated with no employment effects, whereas movements in the cost of French minimum wage workers are associated with very strong negative employment effects. Our analysis is based upon identifying the direct effect of the change in the real minimum wage rate on exits from (entry into) employment when the real minimum wage rate increases (respectively, decreases) and identifying the heterogeneity in the behavior of our treatment and control groups using a pseudo-experimental contrast. We relate the difference-in-difference estimator directly to demand and supply elasticities for the two groups.

Keywords: Minimum Wage, International Comparisons, Labor Demand, Labor Costs

JEL Classifications: J31, J23

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1. Introduction

In this paper we examine the link between changes in the real minimum wage rate and employment outcomes for men and women in France and the United States. We make use of longitudinal data on employment status and earnings to study how individuals are affected by real increases or real decreases in the minimum wage, conditional on the individual's location in the earnings distribution. We focus on low-wage workers and take particular care to distinguish sub-populations that might be affected differently by the minimum wage. We are also careful to distinguish workers in states where the real minimum wage rate increased from those in states where it decreased in the United States, and to explicitly model changes in the subsidies for minimum and low wage employment in France.

Although little attention has been paid to the situation in continental Europe,[1] some European countries provide interesting alternatives to the much-studied U.S. case. France, in particular, provides a stark contrast to the United States. 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) but the nominal minimum wage rate in France rose steadily over the 1980s and 1990s, as did real minimum wages and the cost of employing minimum wage workers (for most, but not all, years). In this paper we exploit the different growth patterns in real minimum wage rates in a symmetric manner to better understand their effects on employment.

As in the U.S., the original studies of the French minimum wage system used aggregate time-series data and found no effect of the minimum wage system on employment.[2] This could be considered surprising because, since its inception, a significant percentage of the French labor force has been employed at wages close to the minimum wage. One reason for use of time series models in the original empirical analyses for France was, certainly, the tendency of American applied researchers to rely upon aggregate time series analyses[3] prior to the widespread dissemination of public use micro-economic data such as the Current Population Survey (CPS). Another reason is that research access to French micro-data was extremely limited until the 1990s. In the present study we use micro-data from France and the United States that were collected using household surveys that are quite comparable. Both of our data sources have a longitudinal design that we exploit extensively to analyze both French and American employment changes in relation to changes in their minimum wage rates and payroll taxes. This paper is a substantial extension of the study we conducted for young workers in both of these countries using similar data (Abowd, Kramarz, Lemieux, and Margolis, 1999).

We use a statistical approach based on the analysis of employment transition probabilities conditional on the position of an individual in the wage distribution. We decompose each year's wage distribution into 4 regions: under, around, marginally over and over the minimum wage. By the definition of our categories and through interactions with changes in the real minimum wage, our analysis of exploits the size of the movements in the real minimum wage directly.[4] Real minimum wage variation in the United States comes from nationally legislated increases, state-specific legislated increases, and inflation. For France, we use the automatic and legislated increases in the nominal

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[1] See Dolado *et. al.* (1996) for a summary of minimum wage studies for France, the Netherlands, Spain and the United Kingdom. See also Brown (1999) for a comprehensive review of recent minimum wage research.

[2] See, for example, Bazen and Martin (1991).

[3] See Brown, Gilroy and Kohen (1982) for a review.

[4] Our analysis bears some resemblance to that of Linneman (1982) and Currie and Fallick (1996).

minimum wage rate that occur (at least) each July, as well as inflation and legislated subsidies for minimum wage labor, to provide variation in the equivalent minimum wage. This variation serves to identify groups of workers whose current wage will fall below the future real minimum wage (in year-pairs when the real minimum wage increases between years), or whose current wage fell below the previous minimum wage (in year-pairs when the real minimum wage declines between years).

Our statistical analysis identifies the change in future or previous employment probabilities given an individual's minimum wage status in the reference period. This change in employment probability is compared to the change in employment probability for a "control" group of workers whose wage in the reference year is marginally above the real minimum wage in the comparison year in an attempt to purge the estimates of the impact of unobserved worker heterogeneity that differs according to the position in the wage distribution. We further compare this difference to an alternative scenario, namely the probability of future employment when minimum wages are decreasing or the probability of previous employment when minimum wages were increasing. This allows us to derive difference-in-difference estimates that further control for unobserved heterogeneity. Our theoretical model provides a direct structural interpretation of the difference-in-difference estimator.

We show that, when one considers pairs of years when the real minimum wage increased, individuals whose current real wage was between the current real minimum wage and the future real minimum wage have significantly lower future employment probabilities than those whose real wages were not similarly situated in France but not in the United States. These effects are slightly larger for French men than for French women, with the difference-in-difference elasticity of future employment with respect to changes in the equivalent minimum wage being -2 for men and -1.5 for women. For the United States, the equivalent (insignificant) figures are 0.4 for men and 0.1 for women. On the other hand, we find that between pairs of years when the real minimum wage decreased, individuals whose current real wage was between the current real minimum wage and the previous real minimum wage do not have substantially lower prior employment probabilities than those whose real wages were not similarly situated.

The paper is structured as follows. Section 2 provides some institutional background on the systems of minimum wages in both France and the United States. Section 3 describes the data that we used to analyze the impact of minimum wages. Section 4 provides an economic interpretation of our natural experiment and its associated pseudo-experiment. Section 5 lays out the statistical models used to evaluate the employment effects of minimum wage changes. Section 6 discusses the results. Section 7 concludes.

2. Institutional Background

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 SMIC "Salaire Minimum Interprofessionnel de Croissance," 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 also 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, there are no provisions in any of the minimum wage laws that would allow regional variation in the SMIC. In some sectors in the French economy, however, the effective minimum wage was determined by collective bargaining

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 upon the employer than its statutory value. Employees share in the legal allocation of the payroll taxes; however, low wage workers benefit substantially more than the average worker from social security benefits financed through these taxes. In general, the payroll taxes are proportional to employee's gross salary; however, the social programs—particularly, unemployment insurance, health care, retirement income and employment programs—benefit low wage workers substantially more (Abowd and Bognanno, 1995).

During the 1990s, France experimented with subsidies for minimum wage and low wage employment, implemented as reductions in payroll taxes paid by employers on earnings of workers who received between the minimum wage and 1.33 times the minimum wage (depending upon the year). Figure 1 shows the relation between total labor costs and real minimum wage rates for the analysis period. The payroll tax reductions implied variations in the cost of employing minimum wage workers that did not move in the same direction as the real minimum wage rate. We constructed an equivalent real minimum wage rate that reflects the time series variation in the total labor costs as shown in Figure 1. See Kramarz and Philippon (2001) for additional details.

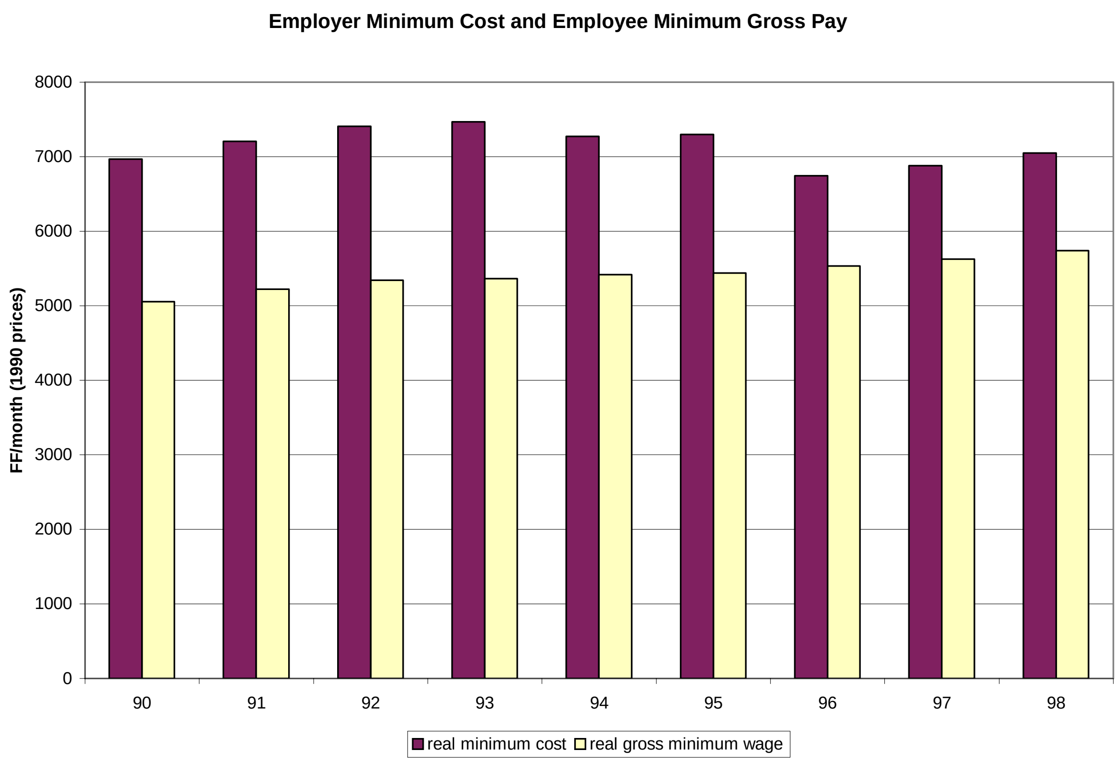

Figure 1

The French minimum wage lies near most of the mass of the wage rate distribution for the employed work force. The first mode of the French wage distribution is within five francs of the minimum wage and the second mode is within 10 francs of the minimum. For example, in 1990 for the

agreements. Because they were often extended by the Minister of Labor to include employers who were not party to the original negotiations, these agreements typically covered entire regions and industries. Although relatively important in the 1970s, these provisions became increasingly irrelevant during the 1980s and 1990s (our period of analysis) since the collectively bargained nominal salary grids remained fixed in the face of an increasing nominal SMIC (Margolis, 1993).

overall distribution, 9.9% of the wage earners in our data lie at or below the minimum wage and an additional 10.0% lie within an additional 5F per hour of the.[6]

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% in the lowest decile to 13.1% in the 3rd lowest decile, then decreases to 6.6% for the 5th decile, increasing to 7.4% for the 6th decile and then declining monotonically to 0.6% in the highest decile of household income.

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.[7] The statutory minimum wage regulates the hourly regular cash compensation received by an employee including the employee's part of any payroll taxes. Card and Krueger (1995) provide an extensive discussion of the effects of the American minimum wage rate on employment and other outcomes.

For 1981, 17.7% of the employed work force had wage rates at or below the minimum wage and an additional 14.6% had wage rates within an additional $1.00 per hour of the minimum. For 1987, only 9.5% of employed persons have hourly wage rates at or below the minimum while an additional 9.9% lie within the next $1.00 per hour (Abowd, Kramarz, Lemieux, and Margolis, 1999).

3. Data Description

3.a. France

The French data were extracted from the "Enquête Emploi" (Labor Force Survey) for the years 1990 to 1998. Approximately sixty thousand 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 over age 14 is interviewed and followed provided that he or she does not change domiciles during this three-year period. We used the INSEE research files for each of the indicated years. These files include the identifiers that allow us to follow individuals from year to year. Using these identifiers we created year-to-year matched files for the years 1990-91 to 1997-98.

The survey measures usual monthly earnings, net of employee payroll taxes but including employee income taxes, and usual weekly hours. The minimum wage is defined on an hourly basis,

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[6] These figures, derived from our household survey data, are roughly consistent with those found in data from enterprise-based surveys. The French ACEMO enterprise survey data suggest approximately 11% of workers were paid at or below the SMIC in both 1987and 1997, with a dip in the intervening period (CSERC, 1999).

[7] Over the period covered by our data, the only significant change in coverage rules occurred in 1989 when the dollar volume of sales threshold necessary for coverage was raised from $250,000 ($362,000 in retail trade and services) to $500,000 for both retail and non retail businesses. Employees of firms covered by the previous thresholds remained covered even if their sales volume did not exceed the new threshold. Small retail businesses became covered in any workweek in which they engaged in commerce or the production of goods for commerce. Thus the remaining uncovered workers were, for the most part, non-hourly executive, administrative, professional or outside sales personnel with salaries above a minimum level and agricultural workers under certain precisely-defined circumstances, most notably piece-work provisions and family labor.

unfortunately the usual weekly hours measure appears to be somewhat noisy. Many respondents report that they work more than 39 hours per week, the legal limit.[8] If one calculates an hourly wage based on these reports, an unreasonable fraction of the employed population is paid below the minimum. For instance, some high-paid young engineers declare more than 50 hours a week. Therefore, we used the monthly wage together with the full-time or part-time status to compute the total labor cost. For workers employed part-time, we used the reported weekly hours to compute their full-time equivalent monthly earnings. For full-time workers, we use the reported monthly earnings.

All young workers 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 (such as "SIVP, stage d'initiation à la vie professionnelle") were excluded from the database. All of these programs provide a legal exemption from the SMIC and from certain payroll taxes. These programs are limited to workers 25 years old and under. In addition, all workers who declared a wage below 95% of the minimum wage without reporting employment on a special scheme were eliminated from the analysis file (they represent less than 5% of the original file). Most correspond to reporting or coding errors as well as workers on special contracts who did not specify the type of contract. We also eliminated workers employed as civil servants or in the public sector since they cannot become non-employed, owing to their status.

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 0 otherwise. The French Labor Force Survey definition of employment is the same as the one used by the International Labor 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 BLS definition.

Our control variables consist of education, age, sex, seniority, type of contract, wage, and year. Education was constructed as six categories: none; completed elementary school, junior high school, or basic vocational/technical school; completed advanced vocational/technical school; completed high school (baccalauréat); completed technical college; completed undergraduate or graduate university. Seniority was measured as the response to a direct question in the survey (years with the present employer). The type of contract was constructed as 3 categories: short-term contracts (CDD), temporary work, and long-term contracts (CDI)[9] as in Abowd, Corbel, and Kramarz (1999).

The data on minimum wage rates, price indices and taxes were taken from "Les Retrospectives", BMS (Bulletin Mensuel de Statistiques, INSEE) in March of each year. The data on tax subsidies were taken from "Liaisons Sociales" (DARES) and "Séries longues sur les Salaires" (INSEE Résultats, édition 1998). We use the information on taxes and the legislated minimum wage to construct an "equivalent minimum wage", which we call the minimum wage below for simplicity. The equivalent minimum wage is calculated as the minimum wage rate that would provide the same cost to the employer using 1990 tax rates as the current legislated minimum wage does with the current tax rates. Algebraically, the equivalent minimum wage is defined as $\widetilde{miw}_t = {}^{\tau_t}\!/_{\tau_{1990}}\, miw_t$, where τ_t is the employer

[8] In France, the first 8 hours of work per week beyond the legal limit are paid with a 25% premium and all additional overtime hours beyond the first 8 are paid with a 50% premium.

[9] The different sorts of contracts mentioned (CDD, CDI and Temporary work) are each associated with different sorts of firing costs. As one objective of this paper is to study exit from employment in France as a function of minimum wage variations, it is important to control for potential sources of variation in exit rates that may be related to the probability that an individual is employed at or near the minimum wage.

payroll tax rate on a worker paid the minimum wage in year *t*, miw_t is the legislated minimum wage in year *t*, and $\widetilde{miw}_t$ is the equivalent minimum wage in year *t*.[10]

3.b. United States

We used the NBER extracts of the outgoing rotation group files from the Current Population Survey for the years 1981 to 2000. We applied the U.S. Census Bureau matching algorithm to create year-to-year linked files for the years 1981-82 to 1999-2000.[11]

The outgoing rotation groups (households being interviewed for the fourth or eighth time in the CPS rotation schedule) are asked to report the usual weekly wage and usual weekly hours. Individuals who normally are paid by the hour were 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 from the U.S. Bureau of Labor Statistics International Labor Statistics.[12]

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. 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 northeast, north-central and southern parts of the U.S. In all of our analyses we also control for the real hourly wage rate in the analysis period.

The U.S. federal minimum wage was increased to \$3.35/hour in 1980, to \$3.80 in 1990 and finally to \$5.15 in 2000. We accounted for state-specific increases in nominal minimum wages (but not youth sub-minimum rates), as well as the federal increases in 1990 and 1991, starting from the data set furnished by Neumark and Wascher (1992).[13]

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[10] The equivalent minimum wage is the minimum wage that would provide the same compensation cost for a minimum wage worker after normalizing the payroll and subsidy structure to that of a base year. In years when the tax structure changes (relative to the base year) so as to increase the cost of employing a minimum wage worker *ceteris paribus*, the equivalent minimum wage will be higher than the legislated minimum wage. Conversely, in years when additional subsidies render minimum wage employment less costly (for the same real minimum), this would imply an equivalent minimum wage below the legislated minimum. Given the enormous statewide diversity in payroll taxes, and in particular the presence of experience rating that makes each firm's compensation cost different for workers with the same gross wage (Margolis and Fougère 2000), we do not attempt to control for this source of variation in the United States.

[11] David Card graciously provided the computer code for implementing the U.S. Census Bureau CPS matching algorithms.

[12] The data were taken from the web site: ftp://ftp.bls.gov/pub/special.requests/ForeignLabor/flscpian.txt

[13] David Neumark graciously provided us with updated versions of the tables that appear in the paper and continue the accounting through 1992. We thank Melissa Bjelland for further extending this data through the rest of our sample period.

4. An Economic Model of Minimum Wage and Employment Changes

The precise economic question that we attempt to answer in this paper is the following: "Is a person whose reference year real wage is between the year t and the year t+1 real minimum wages employed with a significantly lower probability in the comparison year than someone whose reference year real wage is marginally above the comparison year real minimum wage?" If, for example, the wage paid in year t reflects the value of the individual's marginal product, and if this value of marginal product were to increase slower (in real terms) than the real minimum wage increases, such an individual should be unemployed in year t+1. On the other hand, if the individual's value of marginal product increases fast enough for the year t+1 minimum wage to not be binding, or if the individual is paid below the value of his or her marginal product in year t[14] and the firm can increase his or her wage to a level consistent with the t+1 minimum wage and still not make a loss on the worker, then the employment prospects of these individuals should be no different than those of other workers with identical characteristics who are employed at other (higher) points in the wage distribution. Similar reasoning can be used to accommodate pairs of years in which the real minimum wage declines by focusing on entry into employment in year t based on the real wage in year t+1.

Our econometric analysis therefore consists of two natural experiments and two alternative scenarios (pseudo-experiments) designed to test our specifications. The natural experiments occur when the real minimum wage rate increases or decreases between two successive years of data. When the real minimum wage increases, we define a treatment group as those individuals whose real wage rate in year t is between the year t real minimum wage and the new, year t+1 real minimum wage. We define a control group as those individuals whose year t real wage rate is between the year t+1 real minimum wage and a multiple of the year t+1 real minimum wage rate. Both treatments and controls are employed in year t. Then, we use our exit model to predict which of these individuals will remain employed in year t+1. The structure of this natural experiment is illustrated in Figure 2 in the panel labeled "Natural Experiment" and "Exit Model." The figure plots the distribution of log wages. Along the log wage axis, the log real minimum wage rate at date t and is denoted by the symbol $rmiw_t$. One could interpret the difference between the treatments and controls in this natural experiment as the effect of the increased real minimum wage rate.

To test this specification, we consider the following alternative scenario. For year pairs between which the real minimum wage rate is decreasing, define the treatment group in this setting as those individuals whose year t wage rates lie in the interval between the year t minimum wage rate and a multiple of that wage rate that equals the average increase in minimum wage rates over year pairs when they increase. Define the control group in this setting as those individuals whose year t wage rate lies between the upper bound of the alternative scenario treatment group and a multiple of this wage rate. Then, we use our exit model to predict which of these individuals will remain employed in year t+1. The structure of this scenario is comparable to that of the natural experiment and is illustrated in Figure 2 in the panel labeled "Pseudo-experiment" and "Exit Model." One could interpret the difference between the treatments and controls in this alternative scenario as being due to heterogeneity in the employment responsiveness of individuals in these two regions of the wage distribution. Thus, a difference-in-difference estimator of the effect of the increased real minimum wage rate on subsequent employment probabilities can be constructed by subtracting the estimated (treatment-control) effect in the alternative scenario from the estimated (treatment-control) effect in the natural experiment, since

[14] One often-cited explanation for such a phenomenon is that employers may have monopsony power; see Card and Krueger (1995), Manning (2003).

the pseudo-experiment removes the effect of unobserved heterogeneity from the global estimates of the employment effect derived from the natural experiment.

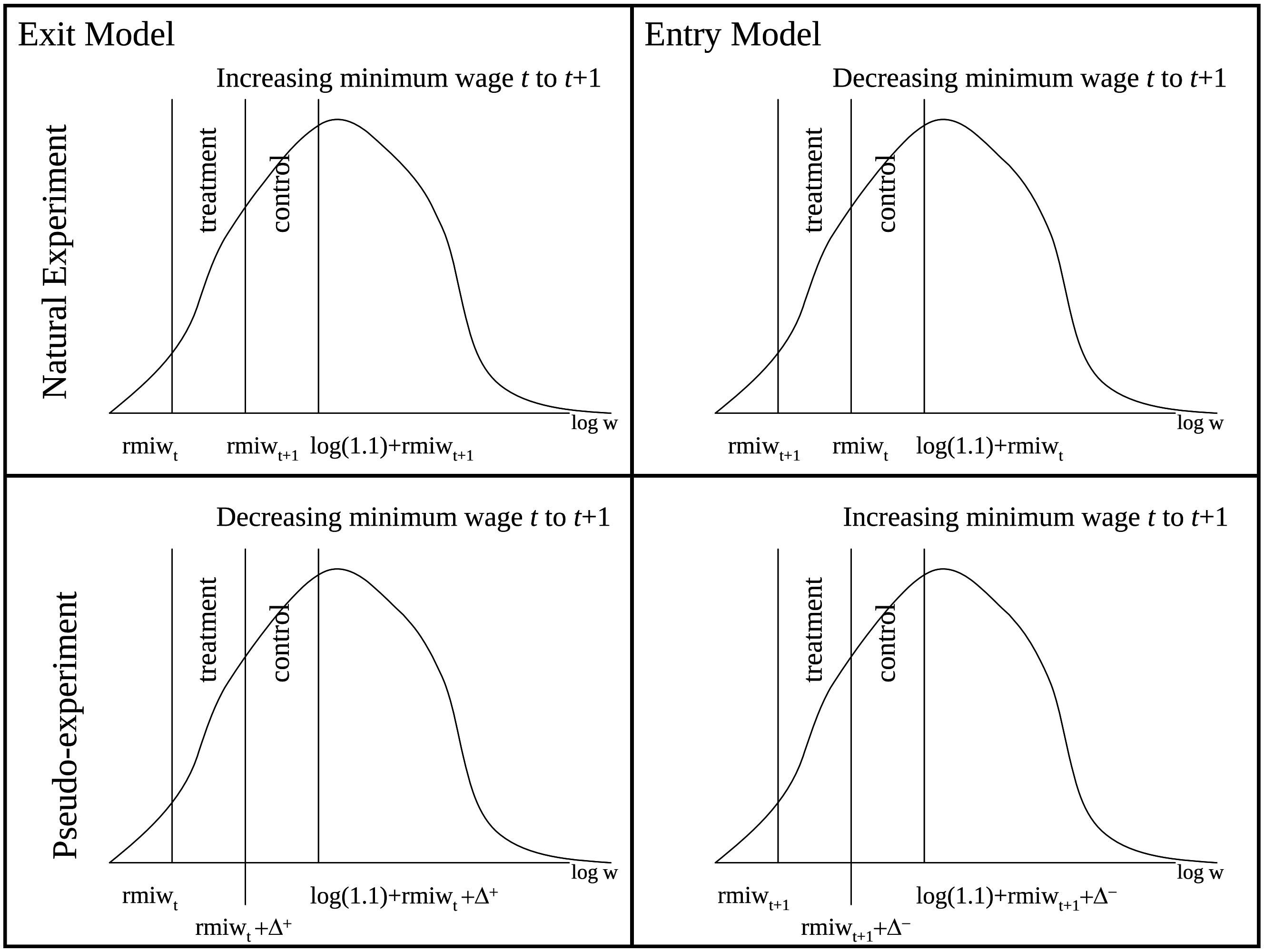

Figure 2

Our second natural experiment is constructed symmetrically. When the real minimum wage decreases, we define a treatment group as those individuals whose wage rate in year *t*+1 is between the year *t*+1 real minimum wage and the year *t* real minimum wage rate. We define a control group as those individuals whose year *t*+1 wage rate is between the year *t* real minimum wage and a multiple of the year *t* real minimum wage rate. Both treatments and controls are employed in year *t*+1. Then, we use our entry model to analyze which of these individuals were employed in year *t*. The structure of this natural experiment is illustrated in Figure 2 in the panel labeled "Natural Experiment" and "Entry Model." One could interpret the difference between the treatments and controls in this natural experiment as the effect of decreasing the real minimum wage rate.

To test this specification, we consider a second alternative scenario. For years in which the real minimum wage rate is increasing, define the treatment group in this setting as those individuals whose year *t*+1 wage rates lie in the interval between the year *t*+1 real minimum wage rate and a multiple of that wage rate that corresponds to the average decrease in minimum wage rates over year pairs when they decrease. Define the control group in this setting as those individuals whose year *t*+1 real wage rate lies between the upper bound of the alternative scenario treatment group and a multiple of this wage rate. Then, we use our entry model to analyze which of these individuals was employed in year *t*. The structure of this alternative scenario is illustrated in Figure 2 in the panel labeled "Pseudo-experiment" and "Entry Model." Once again, one could interpret the difference between the treatments and controls in this setting as heterogeneity in the non-employment responsiveness of individuals in these two regions of the wage distribution. Thus, a difference-in-difference estimator of the effect of the decreased real minimum wage rate on prior employment probabilities can be constructed by subtracting the estimated (treatment-control) effect in the second alternative scenario from the estimated (treatment-control) effect in the second natural experiment.

In order to interpret the outcomes of our quasi-experimental framework, we consider the effects of minimum wage changes on demand and supply conditions affecting the treatment and control groups. Suppose that the treatments and controls are two distinct types of labor. Then, the demand for each labor type depends upon its own wage rate, the other group's wage rate, and other factors, which we ignore below. The supply of each type of labor depends upon its own wage rate and other factors, which we also ignore.

There are two demand equations:

$$\log L_T = \eta_{TT} \log w_T + \eta_{TC} \log w_C$$

$$\log L_C = \eta_{CT} \log w_T + \eta_{CC} \log w_C$$

where the coefficients on the log wage rates represent Hicks-Allen demand elasticities. Similarly, there are two supply equations:

$$\log L_T = \varepsilon_T \log w_T$$

$$\log L_C = \varepsilon_C \log w_C$$

where the coefficient on the log wage rate in each supply equation is the Allen elasticity of supply. Consider a change in the real minimum wage rate. In the natural experiment, this change increases w_T but there is only movement along the demand curve for the treatment group (the minimum wage rate is binding). There is both a demand and supply response in the market for controls. Hence, the equilibrium quantity changes are:

$$\frac{d \log L_T}{d \log w_T} = \left(\eta_{TT} + \eta_{TC} \left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}} \right) \right) + u_T$$

and

$$\frac{d \log L_C}{d \log w_T} = \varepsilon_C \left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}} \right) + u_C$$

where u_T and u_C represent unmeasured heterogeneity in the response of the treatment and control groups to the change in the minimum wage rate. The treatment-control contrast is:

$$\left[\frac{d\log L_T}{d\log w_T}-\frac{d\log L_C}{d\log w_T}\right]_N=\left(\eta_{TT}+\left(\eta_{TC}-\varepsilon_C\right)\left(\frac{\eta_{CT}}{\varepsilon_C-\eta_{CC}}\right)\right)+u_T-u_C\,.$$

For the simple pseudo-experiment where the change in the minimum wage does not affect the relevant portion of the wage distribution, we have a treatment-control contrast of

$$\left[\frac{d\log L_T}{d\log w_T}-\frac{d\log L_C}{d\log w_T}\right]_P=u_T-u_C\,.$$

The difference-in-difference contrast, therefore, identifies

$$\left[\frac{d\log L_T}{d\log w_T}-\frac{d\log L_C}{d\log w_T}\right]_N-\left[\frac{d\log L_T}{d\log w_T}-\frac{d\log L_C}{d\log w_T}\right]_P=\left(\eta_{TT}+\left(\eta_{TC}-\varepsilon_C\right)\left(\frac{\eta_{CT}}{\varepsilon_C-\eta_{CC}}\right)\right)$$

or alternatively

$$\left[d\log L_T-d\log L_C\right]_N-\left[d\log L_T-d\log L_C\right]_P=\left(\eta_{TT}+\left(\eta_{TC}-\varepsilon_C\right)\left(\frac{\eta_{CT}}{\varepsilon_C-\eta_{CC}}\right)\right)d\log w_T\,. \qquad (1)$$

As long as the minimum wage is binding, equation (1) holds for a decrease in the minimum wage rate for both the natural experiment and the pseudo-experiment.

Our actual data analysis is slightly more complex. When the minimum wage rate increases between years *t* and *t*+1, we use those periods to generate the natural experiment and we use the other periods, say *s* and *s*+1, to generate the alternative scenario. Similarly, when the minimum wage rate decreases between *s* and *s*+1 we use those periods to generate the natural experiment and the periods *t* and *t*+1 to generate the pseudo-experiment. As a result, each pair enters the contrasting equations twice, once as a natural experiment (when the minimum wage rate moves in a direction that should affect employment for entry or exit, as appropriate) and once as a pseudo-experiment (when the minimum wage rate moves in a direction that should not directly affect employment for entry or exit, as appropriate). Since equation (1) holds in both directions (with the sign of $d\log w_T$ determining the direction of the effect), we can rewrite equation (1) to make the distinction between increasing and decreasing pairs more explicit:

$$\begin{aligned}&\left[d\log L_T-d\log L_C\right]_N-\left[d\log L_T-d\log L_C\right]_P=\\&\left(\eta_{TT}+\left(\eta_{TC}-\varepsilon_C\right)\left(\frac{\eta_{CT}}{\varepsilon_C-\eta_{CC}}\right)\right)\left(I\left(d\log w_T>0\right)d\log w_T+I\left(d\log w_T<0\right)\left|d\log w_T\right|\right)\end{aligned} \qquad (2)$$

5. Minimum Wage Effects on Employment: Conditional Logit Analysis

Let $rmiw_t$ be the log of the real minimum wage in year *t* and rw_{it} be the log of the real wage earned by individual *i* in year *t*.[15] Furthermore, let $\Delta_t^{rmiw} = I(rmiw_{t+1} - rmiw_t \geq 0)$ denote that year *t* is the first year of an increasing real minimum wage year pair (an increasing year pair implies $\Delta_t^{rmiw} = 1$), $e_{i,t}$ denote the individual's employment status (employed=1) in year *t*; and $X_{i,t}$ denote a vector of covariates that also affect the probability of employment, measured for individual *i* at date *t*.[16] After conditioning on employment in the reference year, we consider separate (but symmetric) models for entry and exit.

5.1. Exit

In the case of exit, we are interested in analyzing the probability of future employment, conditional on current employment, as a function of the position in the wage distribution and the size of the change in the minimum wage. The variables that determine the position in the wage distribution are a function of whether or not the real minimum wage increases or decreases between *t* and *t*+1. For year pairs between which the real minimum wage increases, we define the indicator variables $B_{i,t}$ and $M_{i,t}$ as follows:

- $B_{i,t} = I\left(rmiw_t \leq rw_{i,t} < rmiw_{t+1}\right)$
- $M_{i,t} = I\left(rmiw_{t+1} \leq rw_{i,t} < \log(1.1) + rmiw_{t+1}\right)$.

Hence, $B_t = \left\{i \middle| B_{i,t} = 1\right\}$ represents the treatment group and $M_t = \left\{i \middle| M_{i,t} = 1\right\}$ the control group. We also define $\Delta^+ = \left(\frac{\sum_{i,t} (rmiw_{t+1} - rmiw_t) \times \Delta_t^{rmiw}}{\sum_{i,t} \Delta_t^{rmiw}} \right)$, which corresponds to the average (over person-years) log increase in the real minimum wage for year pairs in which it increases.

Our first model estimates the probability that $e_{i,t+1} = 1$ given that $e_{i,t} = 1$, where *t* is the first year of an increasing minimum wage pair. The functional form is

$$\Pr\left[e_{i,t+1} = 1 \middle| e_{i,t} = 1, \Delta_t^{rmiw} = 1\right] = \Lambda \begin{pmatrix} X_{i,t}\beta \\ + b \times B_{i,t} \times \left(rmiw_{t+1} - rmiw_t\right) \\ + m \times M_{i,t} \times \left(rmiw_{t+1} - rmiw_t\right) \end{pmatrix}, \qquad (3)$$

[15] 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 divided usual weekly earnings 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. However, since this measure of wages with tips is constructed by dividing earnings by hours, caution is advised when interpreting results that use it. Welch (1997) provides evidence on various sorts of measurement error in the Current Population Survey, and hints that hours are likely to be a greater source of measurement error than wages. As a result, all of the results we present here use the declared hourly wage measure (when available), i.e. without tips. Results using the measure with tips are very similar to those presented here, and are available upon request.

[16] In the U.S. data the real minimum wage rate varies by state; however, we suppress the subscript *i* for clarity in our comparisons with France.

where Λ is the standard logistic function. Equation (3) is the natural experiment for the exit model associated with increasing minimum wage rate. The results are reported in Tables 1 and 2 (for men and women, respectively).

For year pairs between which the real minimum wage decreases, the definitions of $B_{i,t}$ and $M_{i,t}$ provided above are no longer valid. We redefine the indicator variables $B_{i,t}$ and $M_{i,t}$ as follows:

- $B_{i,t} = I\left(rmiw_t \leq rw_{i,t} < rmiw_t + \Delta^+\right)$
- $M_{i,t} = I\left(rmiw_t + \Delta^+ \leq rw_{i,t} < \log(1.1) + rmiw_t + \Delta^+\right)$.

Using these definitions, we estimate a similar model in years of decreasing real minimum wage year pairs:

$$\Pr\left[e_{i,t+1} = 1 \middle| e_{i,t} = 1, \Delta_t^{rmiw} = 0\right] = \Lambda\begin{pmatrix} X_{i,t}\beta \\ + b \times B_{i,t} \times \left|rmiw_{t+1} - rmiw_t\right| \\ + m \times M_{i,t} \times \left|rmiw_{t+1} - rmiw_t\right| \end{pmatrix}. \quad (4)$$

Equation (4) corresponds to the pseudo-experiment for the exit model associated with the decreasing minimum wage rate. The results are also reported in Tables 1 and 2 (for men and women, respectively). Note that equations (3) and (4) represent a generalization of equation (2), in that the estimation samples in each case are determined by the sign of the change in the minimum wage. As a result, we allow for the elasticities in the case of an increase or a decrease in the minimum wage to differ,[17] whereas equation (2) suggests a specification in which the elasticities are symmetric but of opposing sign.[18]

An alternative extension of equation (2) imposes that the elasticities be symmetric and of opposite sign but allows for indicator variables corresponding to the position in the wage distribution:

$$\Pr\left[e_{i,t+1} = 1 \middle| e_{i,t} = 1\right] = \Lambda\begin{pmatrix} X_{i,t}\beta + b_0 B_{i,t} + m_0 M_{i,t} \\ + b \times B_{i,t} \times \left(rmiw_{t+1} - rmiw_t\right) \\ + m \times M_{i,t} \times \left(rmiw_{t+1} - rmiw_t\right) \end{pmatrix}. \quad (5)$$

The results of equation (5) are reported in Tables 5 and 6 (for men and women, respectively). Note that the sign of the difference in the minimum wages in equation (5) corresponds to the interaction of the indicator variable in equation (2).

[17] One could also allow for asymmetric elasticities by interacting Δ_t^{rmiw} with $b \times B_{i,t} \times \left(rmiw_{t+1} - rmiw_t\right)$ and $m \times M_{i,t} \times \left(rmiw_{t+1} - rmiw_t\right)$ and introducing absolute values where appropriate. Such a specification (which we have also estimated, and for which the results are also available upon request) constrains the other coefficients of the model to be equal, a hypothesis which is rejected by a Hausman-type specification test on the basis of the results of estimating equations (3) and (4).

[18] Our results suggest that an asymmetric specification is more appropriate. Nevertheless, we have also estimated a specification that corresponds directly to equation (2). Results are available from the authors upon request.

A final specification that encompasses both of these extensions[19] of the model in equation (2) can be written as

$$\Pr\left[e_{i,t+1}=1\middle|e_{i,t}=1\right]=\Lambda\begin{pmatrix} X_{i,t}\beta+b_0B_{i,t}+m_0M_{i,t} \\ +b^{inc}\times B_{i,t}\times\left(rmiw_{t+1}-rmiw_t\right)\times\Delta_t^{rmiw} \\ +m^{inc}\times M_{i,t}\times\left(rmiw_{t+1}-rmiw_t\right)\times\Delta_t^{rmiw} \\ +b^{dec}\times B_{i,t}\times\left|rmiw_{t+1}-rmiw_t\right|\times\left(1-\Delta_t^{rmiw}\right) \\ +m^{dec}\times M_{i,t}\times\left|rmiw_{t+1}-rmiw_t\right|\times\left(1-\Delta_t^{rmiw}\right)\end{pmatrix}. \qquad (6)$$

The results of estimating this specification are also reported in Tables 5 and 6. It is worth noting that the identification of b_0 separately from b in equation (5) and the b^{inc}- b^{dec} pair in equation (6), as well as the identification of m_0 separately from m or the m^{inc}- m^{dec} pair, relies on sufficient variation in $\left(rmiw_{t+1}-rmiw_t\right)$. Given that this term is positive when $\Delta_t^{rmiw}=1$ and negative otherwise, all of the coefficients in equation (5) should be identified. Although the cross-state variation in the minimum wages should be sufficient to identify the coefficients in equation (6) for the United States, there is may be insufficient variation within increasing and decreasing year pairs in $\left(rmiw_{t+1}-rmiw_t\right)$ in France to identify $B_{i,t}$ and $M_{i,t}$ in equations (3) and (4).

5.2. Entry

Entry models consider the probability of previous employment, conditional on current employment as a function of the individual's position in the wage distribution and the size of the change in the minimum wage. As before, the variables that determine the position in the wage distribution are a function of whether or not the real minimum wage increases or decreases between t and $t+1$. For year pairs between which the real minimum wage decreases, we define the indicator variables $B_{i,t+1}$ and $M_{i,t+1}$ as follows:

- $B_{i,t+1}=I\left(rmiw_{t+1}\le rw_{i,t+1}<rmiw_t\right)$
- $M_{i,t+1}=I\left(rmiw_t\le rw_{i,t+1}<\log(1.1)+rmiw_t\right)$.

Hence, $B_{t+1}=\left\{i\middle|B_{i,t+1}=1\right\}$ represents the treatment group and $M_{t+1}=\left\{i\middle|M_{i,t+1}=1\right\}$ the control group. We define $\Delta^-=\left(\frac{\sum_{i,t}\left(rmiw_t-rmiw_{t+1}\right)\times\left(1-\Delta_t^{rmiw}\right)}{\sum_{i,t}\left(1-\Delta_t^{rmiw}\right)}\right)$, which corresponds to the average (over person-years) log decrease in the real minimum wage for year pairs in which it declines.

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[19] This specification, unlike the specifications of equations (3) and (4), constrains the coefficients on the other variables in the model (β) to be identical in increasing and decreasing year pairs.

Our first entry model estimates the probability that $e_{i,t} = 1$ given $e_{i,t+1} = 1$, where t+1 corresponds to the second year of a decreasing minimum wage year pair. The functional form is

$$\Pr\left[e_{i,t} = 1 \middle| e_{i,t+1} = 1, \Delta_t^{rmiw} = 0\right] = \Lambda\begin{pmatrix} X_{i,t+1}\beta \\ + b \times B_{i,t+1} \times \left(rmiw_t - rmiw_{t+1}\right) \\ + m \times M_{i,t+1} \times \left(rmiw_t - rmiw_{t+1}\right) \end{pmatrix}. \qquad (7)$$

Equation (7) corresponds to the entry model natural experiment and is the entry version of equation (3). The results are reported in Tables 3 and 4 (for men and women, respectively).

In the case of entry, for year pairs between which the real minimum wage increases, the definitions used above of $B_{i,t+1}$ and $M_{i,t+1}$ are again no longer valid. Thus we redefine the indicator variables $B_{i,t+1}$ and $M_{i,t+1}$ as follows:

- $B_{i,t+1} = I\left(rmiw_{t+1} \le rw_{i,t+1} < rmiw_{t+1} + \Delta^-\right)$
- $M_{i,t+1} = I\left(rmiw_{t+1} + \Delta^- \le rw_{i,t+1} < \log(1.1) + rmiw_{t+1} + \Delta^-\right)$.

Using these definitions, we estimate a similar logit model in years of increasing real minimum wage year pairs:

$$\Pr\left[e_{i,t} = 1 \middle| e_{i,t+1} = 1, \Delta_t^{rmiw} = 1\right] = \Lambda\begin{pmatrix} X_{i,t+1}\beta \\ + b \times B_{i,t+1} \times \left|rmiw_t - rmiw_{t+1}\right| \\ + m \times M_{i,t+1} \times \left|rmiw_t - rmiw_{t+1}\right| \end{pmatrix}. \qquad (8)$$

Equation (8) corresponds to the pseudo-experiment for the entry model associated with the increasing minimum wage rate and is the counterpart to equation (4). The results are also reported in Tables 3 and 4 (for men and women, respectively).

We can again apply the alternative generalization of the model in equation (2) that involves using symmetric elasticities and including indicator variables for the position in the wage distribution.

$$\Pr\left[e_{i,t} = 1 \middle| e_{i,t+1} = 1\right] = \Lambda\begin{pmatrix} X_{i,t+1}\beta + b_0 B_{i,t+1} + m_0 M_{i,t+1} \\ + b \times B_{i,t+1} \times \left(rmiw_t - rmiw_{t+1}\right) \\ + m \times M_{i,t+1} \times \left(rmiw_t - rmiw_{t+1}\right) \end{pmatrix}. \qquad (9)$$

The results of equation (9) are reported in Tables 7 and 8 (for men and women, respectively). Similarly, we can again attempt to estimate the specification that encompasses both extensions, but in the case of entry:

$$\Pr\left[e_{i,t} = 1 \middle| e_{i,t+1} = 1\right] = \Lambda\begin{pmatrix} X_{i,t+1}\beta + b_0 B_{i,t+1} + m_0 M_{i,t+1} \\ + b^{dec} \times B_{i,t+1} \times \left(rmiw_t - rmiw_{t+1}\right) \times \left(1 - \Delta_t^{rmiw}\right) \\ + m^{dec} \times M_{i,t+1} \times \left(rmiw_t - rmiw_{t+1}\right) \times \left(1 - \Delta_t^{rmiw}\right) \\ + b^{inc} \times B_{i,t+1} \times \left|rmiw_t - rmiw_{t+1}\right| \times \Delta_t^{rmiw} \\ + m^{inc} \times M_{i,t+1} \times \left|rmiw_t - rmiw_{t+1}\right| \times \Delta_t^{rmiw} \end{pmatrix}. \qquad (10)$$

These results are also shown in Tables 7 and 8, and are also likely to be subject to the same identification issues as the results for exit insofar as concerns the variability in the change in minimum wages within increasing and within decreasing year pairs.

6. Discussion of the Results

In all our tables we report only the coefficients and elasticities on the key real minimum wage rate variables. The differences in elasticities reported in the tables are contrasts of partial elasticities based on the formulas:

$$\text{Difference in exit elasticities} = \frac{\partial \ln \Pr[e_{t+1} = 1 \mid e_t = 1, rmiw_t, rmiw_{t+1}, B_t = 1]}{\partial rmiw_{t+1}} - \frac{\partial \ln \Pr[e_{t+1} = 1 \mid e_t = 1, rmiw_t, rmiw_{t+1}, M_t = 1]}{\partial rmiw_{t+1}}$$

$$\text{Difference in entry elasticities} = \frac{\partial \ln \Pr[e_t = 1 \mid e_{t+1} = 1, rmiw_t, rmiw_{t+1}, B_t = 1]}{\partial rmiw_{t+1}} - \frac{\partial \ln \Pr[e_t = 1 \mid e_{t+1} = 1, rmiw_t, rmiw_{t+1}, M_t = 1]}{\partial rmiw_{t+1}}$$

Where the reader is reminded that $rmiw_t$ is the log real minimum wage rate at date *t*. Both formulas are evaluated at the sample means of the exit and entry rates for the treatment and control groups.

6.a. Exit Results

Tables 1 and 2 present the basic results of our exit model for men and women, respectively, distinguishing between pairs of years when the real minimum wage increases and pairs of years when it decreases. For increasing real minimum wage year pairs, we are estimating equation (3), while for decreasing real minimum wage year pairs, we estimate equation (4). We report only the coefficients and elasticities on the key real minimum wage variables.

These two tables show a result that is common to all of our analyses—namely, that the French minimum wage laws result in substantial employment loss while the American laws do not. For American men (Table 1) the natural experiment estimate of the exit elasticity of subsequent employment with respect to the real minimum wage rate is essentially zero (-0.0189±0.4392) whereas the comparable estimate for French men is substantial (-1.9672±0.7598). The exit elasticity in the natural experiment may be interpreted as an estimate of $\left(\eta_{TT} + (\eta_{TC} - \varepsilon_C)\left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}}\right)\right) + u_T - u_C$. The pseudo-experiment estimate of the difference in exit elasticities for American men is -0.4813±0.3509 and the comparable estimate for French men is 0.3136±0.3116. The pseudo-experimental differences in exit elasticities can be interpreted as an estimate of $u_T - u_C$. Finally, we apply the difference-in-difference estimator. For American men the resulting contrast in exit elasticities is 0.4624±0.5622, which is essentially zero. For French men the contrast is -2.2809±0.8212, a very substantial estimate. The difference-in-difference estimator provides a direct estimate of $\left(\eta_{TT} + (\eta_{TC} - \varepsilon_C)\left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}}\right)\right)$, which is the sum of η_{TT}, the Hicks-Allen own elasticity of demand for the treatment group (those subject to the increase in the minimum wage rate) and a term whose sign depends upon the difference between the Hicks-Allen elasticity of demand for the treatment group with respect to the wage rate of the control group, η_{TC}, which is positive provided that the treatments and controls are demand substitutes, and the Allen elasticity of supply, ε_C, for the control group. The magnitude of the departure of the estimated exit elasticity from the own elasticity of demand for the treatment group also depends upon the ratio $\left(\left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}}\right)\right)$, which must be positive. Table 1 shows that η_{TT} dominates the result for France but for the U.S. the cross-elasticity of demand, η_{TC}, must dominate the elasticity of supply, ε_C in order for the difference between them to be positive. The results for women are shown in Table 2. Going directly to the difference-in-difference estimator, we find that American women are essentially unaffected by changes in the minimum wage rate (0.1151±0.4366). On the other hand, French women are strongly affected by changes in the minimum wage rate (-1.5350±0.5747). Comparable entry results are presented in Tables 3 and 4 (for men and women, respectively) and are discussed in the "Entry Results" subsection below.

Consider next the specification that imposes symmetric effects and allows for indicator variables corresponding to the position in the wage distribution, as in equation (5). The results are presented in Tables 5 and 6 (for men and women, respectively) in the panels labeled "Identical coefficients." When applying equation (5), the difference-in-difference estimator is produced directly by using the difference in coefficients or elasticities. For American men the difference in exit elasticities is 0.1050±0.1683, which is essentially zero. For French men this difference is -0.4041±0.2506, which is smaller than the estimate presented in Table 1 but still substantially negative. For American women (Table 6) the difference in exit elasticities is -0.2164±0.1465, again essentially zero but closer to statistical significance, whereas for French women we find -0.2983±0.1997, again smaller than in Table 2 but still substantially negative.

It should be noted that the difference in the coefficients on the indicator variables that reflect the position in the wage distribution is generally significant, suggesting that this generalization is likely to be important and potentially useful in other contexts. Furthermore, it also seems unlikely that the elasticity of employment with respect to increases in the minimum wage would be identical (but of opposite sign) to the elasticity of employment with respect to reductions in the minimum wage, as allowed for in equations (3) and (4). To accommodate both of these considerations, we estimate the encompassing model of equation (6), whose results also appear in Tables 5 and 6 in the panels labeled "Different Coefficients."

As expected, the more general specification is better identified for the United States than for France, although neither set of results is precisely estimated. In the United States, the difference in difference estimator of the exit elasticity is -0.8025±0.5436 for men and 0.1283±0.4361 for women. For France, the corresponding (statistically insignificant) figures are -1.1605±1.2273 for men and -1.3077±0.9290 for women. These point estimates are closer to the difference in difference estimates of Tables 1 and 2 than those of tables 5 and 6 when we impose symmetric employment elasticities, suggesting that the first generalization (asymmetric elasticities) of the model in equation (2) may be the more relevant. Furthermore, recall that the separate identification of a main effect for the treatment and control groups as well as an interaction effect with the change in the minimum wage rate requires sufficient variation in the change in the minimum wage rate across years (recall that there is also an unrestricted year effect in the model). In the United States, the variation in state minimum wage rates contributes to this identification; however, in France, there must be sufficient year-to-year variability in the changes of the national minimum wage rate. Our results indicate that this year-to-year variability may not be sufficient to well-identify the asymmetric estimator for France.[20]

6.b. Entry Results

Tables 3 and 4 present the basic results of our entry model for men and women, respectively, distinguishing between pairs of years when the real minimum wage decreases and pairs of years when it increases. For decreasing real minimum wage year pairs, we are estimating equation (7), while for increasing real minimum wage year pairs, we estimate equation (8). We report only the coefficients and elasticities on the key real minimum wage variables.

These two tables show a result that is common to all of our analyses—namely, that neither the French minimum wage tax subsidies nor the American real minimum wage decreases clearly facilitate employment entry. Since the interpretation of the entry elasticity components is identical to the exit model discussion, we discuss only the difference-in-difference estimator. For American men the contrast in entry elasticities is -0.2398±0.3711, which is essentially zero. For French men the contrast is 0.2499±0.9775, again, essentially zero. Since the difference-in-difference estimator provides a direct estimate of $\left(\eta_{TT} + \left(\eta_{TC} - \varepsilon_C\right)\left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}}\right)\right)$, the results show that neither part clearly dominates. The results for women are shown in Table 4. Going directly to the difference-in-difference estimator, we find that American women are essentially unaffected by changes in the minimum wage rate (-0.1180±0.3063) as are French women (-0.0710±0.5509).

Consider next the specification that imposes symmetric effects and allows for indicator variables corresponding to the position in the wage distribution, as in equation (9). The results are presented in Tables 7 and 8 (for men and women, respectively) in the panel labeled "Identical Coefficients." As in the case of equation (5), the difference-in-difference estimator is produced directly by using the difference in coefficients or elasticities when applying equation (9). The analysis periods for the American data are identical to the periods used in the comparable exit tables. However, we consider a different set of years for the French data in Tables 7 and 8. When we used the same year pairs as in Tables 3 and 4 for the analysis of the pooled French data, our estimates were essentially identical to those in Tables 3 and 4. We decided to focus the pooled French analysis on the years that followed the large tax subsidies for employers of minimum wage workers that were enacted in the mid-1990s. Recall that the real minimum wage rate cannot fall in France (by law), and thus our declining equivalent real minimum wage rate results from the effects of these tax subsidies on the cost of employing a minimum wage worker. These subsidies began in 1993 and, as Figure 2 shows, produced decreases in the cost of employing a minimum wage worker for the pairs 1993-1994 and 1995-1996, of which 1995-1996 was much larger.

Considering Table 7, part of which shows the pooled analysis with symmetric coefficients as described by equation (9), the estimated difference in entry elasticities for American men is -0.0247±0.1301 while the difference is -0.3224±0.3683 for French men. For American women (Table 8) the difference in entry elasticities is 0.2939±0.1143, whereas for French women we find -0.3701±0.2919. None of these estimated elasticities is significantly different from zero, even given our focus on years in which one might most expect to see effects of the minimum wage in France. That said, as with our exit models, we again find important differences in the main effects in these

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[20] The absence of a significant effect in the difference in difference estimator of Tables 1 and 2, where the main effect is absent from the specification, does not allow us to conclude that the insignificant result for the United States is due exclusively to insufficient variability in minimum wage changes in the data.

specifications in the United States, although even these differences disappear in France when one turns one's attention from exit to entry.

Finally, Tables 7 and 8 also estimate the general model of equation 10 in the panel labeled "Different Coefficients." Since there did not appear to be enough variability in minimum wage movements to identify the effects in the exit models for France, it is not surprising that the entry models for France suggest a similar identification problem. For the United States, however, these results are quite different from the other entry models, with a male employment elasticity of 0.8870±0.3791 and a female employment elasticity of 0.7638±0.3272. These elasticities for both sexes are significantly positive, suggesting that the effect of the Hicks-Allen own elasticity of demand for the treatment group η_{TT} is dominated by the term $(\eta_{TC} - \varepsilon_C)\left(\frac{\eta_{CT}}{\varepsilon_C - \eta_{CC}}\right)$. Since this latter term is positive only when $\eta_{TC} > \varepsilon_C$ (and since $\varepsilon_C > 0$), this further implies that the Hicks-Allen elasticity of demand for the treatment group with respect to the wage rate of the control group, η_{TC}, which is positive provided that the treatments and controls are demand substitutes, is larger than the Allen elasticity of supply, ε_C, for the control group. In other words, our results for the United States suggest important spillover effects for the minimum wage insofar as concerns entry into employment when the real minimum wage declines.

7. Conclusion

By comparing effects of minimum wage movements on workers employed at the minimum with those employed marginally above it, we identify the direct effects of changes in the real minimum wage rate on exits from employment and entry into employment. By constructing an appropriate pseudo-experimental contrast, we identify the heterogeneity in the responses of the individuals employed near the minimum wage rate as compared with those who are marginally above them in the wage distribution. Our difference-in-difference estimator, which removes the estimated heterogeneity from the direct effect of the minimum wage rate is directly interpretable as a function of the demand and supply elasticities associated with the treatment and control groups. We find that exits from employment are not very sensitive to changes in the minimum wage rate in the U.S. whereas in France there is a strong negative effect. Entry into employment is not very sensitive to changes in the minimum wage rate in either country. There is not much difference in the responses of men and women in either country even though more women are paid near the minimum wage rate in both countries.

Even when the conditional exit and entry elasticities are large, the treatment groups are small, 3-5% of men and 8% of women in both countries. Thus, unconditional elasticities of employment are much lower than our estimated conditional ones. 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 large negative employment effects on this group in France but not in the United States.

Our results, which are based on direct data evidence from households, are compatible with the results of Card and Krueger (1994, 2000), which are based on direct data evidence from American establishments. Kramarz and Philippon (2001) have analyzed the French data for 1990 to 1998, focusing carefully on the effects of targeted payroll tax subsidies on the total labor cost of minimum wage and low-wage workers. Their results, for a period of analysis that contains intervals in which the total labor cost of minimum wage workers rises and falls, are essentially the same as the ones we find here for France.

Our major contribution to the minimum wage debate consists of carefully analyzing both the direct effects of a change in the real minimum wage and the effects of heterogeneity in the behavior of individuals who are near each other in the wage distribution under circumstances of both increasing and decreasing real minimum wage rates. If a single set of behavioral parameters were able to explain gains and losses of employment surrounding changes in the minimum wage rate, this methodology would have detected it. Because we obtain very different results for the two countries when we study the exit model, it is clear that a single set of demand and supply parameters is not consistent with the data. There is no "employment effect of changing the minimum wage rate," properly defined. Rather, it appears to depend upon the level of the real minimum wage rate inclusive of both employer and employee payroll taxes, which is much higher in France and the direction of the change of the real minimum, which regularly moves up and down in the U.S. but almost always goes up in France.

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