UNIVARIATE REGRESSIONS OF EMPLOYMENT ON MINIMUM WAGES IN THE PANEL OF U.S. STATES

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14 February 2009

Key words: Minimum Wages, Panel Data.

JEL Classification: J23, C33.

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MANFRED KEIL, DONALD ROBERTSON AND JAMES SYMONS

ABSTRACT. The paper finds a strong negative correlation between youth employment and minimum wages for the panel of U.S. states, 1976-2007. Such a correlation is not observed in earlier panels. The source of the new results is traced to the greater variance of minimum wages across states emerging after about 2000.

1. INTRODUCTION

The effect of minimum wage legislation on employment represents an area where the predictions of simple economic theory are hotly contested by both economists and policymakers. The competitive model of the labor market predicts unambiguously that an increase in minimum wages would reduce employment, and for many years this was a touchstone of economic orthodoxy. This orthodoxy explained popular support for minimum wages as springing from a benign but misguided desire to help the poor – misguided because some of the poor are made better-off only by making others of the poor worse-off, the losses outweighing the gains. Card, Katz and Krueger's influential set of papers[1] however seems to find empirical evidence of the opposite effect if anything, consistent perhaps with some sort of large-scale monopsony power in labor markets. Card and Krueger have been seriously challenged on methodological grounds (see, in particular, the comments by Brown; Hamermesh; and Welch in Ehrenberg (1995)). Perhaps more importantly, the result of a positive minimum wage-employment relationship has been subsequently questioned by other studies using data which cover identical time periods.[2] In any case, the struggle for the hearts and minds of policy-makers seems to have been won in the United States.

[1] For a convenient summary of Card and Krueger (1994); Katz and Krueger (1992); Card (1992a,b); see Card and Krueger (1995) chapters 2-4.

[2] For Card (1992a), see Taylor and Kim (1995); for Card (1992b), see Deere, Murphy and Welch (1995a,b); for Card and Krueger (1994), see Neumark and Wascher (2000), although Card and Krueger (2000) should also be viewed in response. A comprehensive recent summary of the evidence is provided by Neumark and Wascher (2007).

Federal minimum wages were increased by a combined 21% in 1996 and 1997. The 1999 *Economic Report of the President* stated that "the weight of the evidence suggests that modest increases in the minimum wage have had very little or no effect on employment." (Yellen, 1999; 112). While the Federal Minimum Wage remained constant over the following decade, State Minimum Wages increased so that, by 2007, over half of the U.S. States had minimum wages above the federal level. On October 11, 2006, just before the November election, five Economics Nobel laureates, six past Presidents of the American Economic Association and over 650 economists signed a statement in support of a "modest" increase in Federal Minimum Wages.[3] Following the 2006 election, Federal Minimum Wages were increased in 2007 and 2008, with another planned increase in 2009. The 2009 Federal Minimum Wage of $7.25 will be 41% above the 1997 level. Likewise in the U.K., the British Labour government instituted a national minimum wage in 1999 and increased it steadily over the last decade, by about 18% in total**.** Moreover it appears the citadel itself is about to fall: the Hong Kong Head of Government has recently announced that the territory is to impose a minimum wage for the first time (*Financial Times*, October 15, 2008)**.** There has clearly been a shift in fashion.

When Brown *et al.* published in 1982 their detailed review of the literature existing up to that date, they concluded that the balance of evidence was that minimum wages exert a detectable, though small, negative effect on employment – a short-run elasticity of -0.1 or less say - with a rather stronger and easier-to-find effect on youth employment.[4] Typically, this evidence was derived from time series studies of labor markets. In contrast, the "New Minimum Wage Research" era, starting with

[3] Quoted as saying "We believe that a modest increase in the minimum wage would improve the well-being of low-wage workers and would not have the adverse effects that critics have claimed."

[4] "[Time] series studies typically find that a 10 percent increase reduces teenage employment by one to three percent." Brown *et al.* (1982).

the publication of four influential papers by Card (1992a, b), Katz and Krueger (1992), and Neumark and Wascher (1992) in the *Industrial and Labor Relations Review*, has seen no such consensus. The key papers are Card and Krueger (1994, 2000). They study an *event*, the increase in minimum wages in New Jersey in 1992, and find that, compared to Pennsylvania where there was no such increase, employment in fast-food restaurants increased somewhat. While in some specifications this effect is not significantly positive, it is certainly not negative.[5] An intrinsic problem with event studies in this context is that the sought effect is acknowledged to be small compared to ambient fluctuation in employment rates. Annual changes in state employment rates have a standard deviation of around 2%, which is the expected fall in employment for a 20% increase in the real wage if the received wisdom is correct. Clearly this could easily be missed by chance. Kennan (1995) has likened the quest for minimum wage effects to trying to find a needle in a haystack.

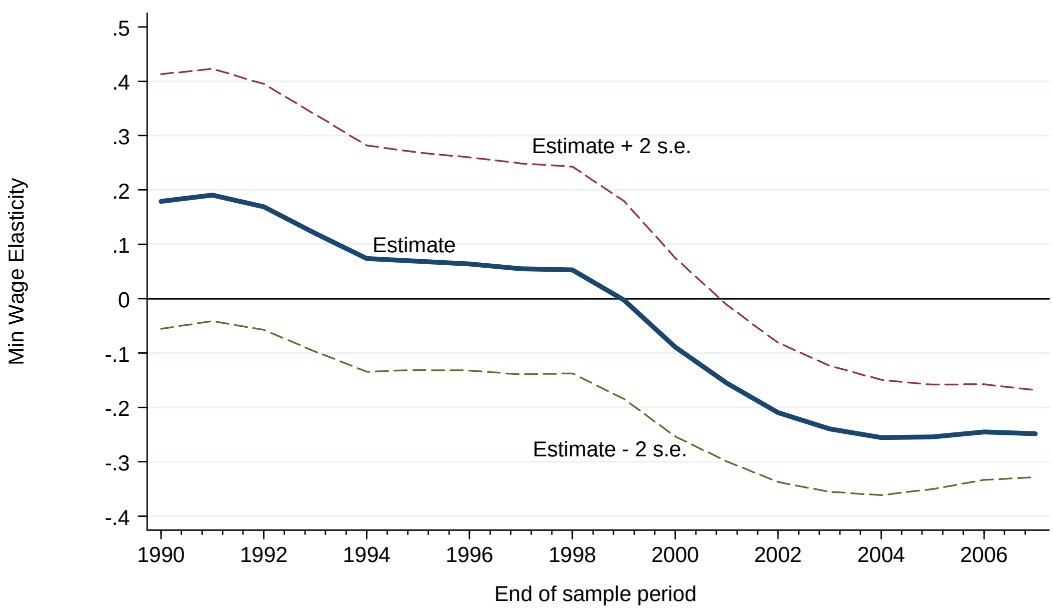
In this paper we shall consider the employment effects of minimum wages in the panel of U.S. states in annual data from 1976 to 2007. In many ways, our approach is much in the spirit of the original panel study by Neumark and Wascher (1992) and it is appropriate to offer some justification for a return to this venerable and - one might have thought - exhausted topic.[6] The point is simply that the observed behaviour of state minimum wages over the last few years contains a vast increase in the amount of information bearing on the effects of minimum wages on employment. The reason is that, following the turn of the millennium, an increasing number of states enacted

[5] Though Neumark and Wascher (1995a,b), (1996) have argued that, while total employment may not have fallen, there was significant replacement of black and Hispanic teenagers by white teenagers dropping out of school, which would be a remarkable example of the law of unintended consequences.
[6] Other influential studies that have used state panels similar to this paper are Card and Krueger (1995), Deere *et al*. (1995), and Burkhauser *et al*. (2000). For a review of the literature in general up to this date, see Neumark and Wascher (2008).

their own minimum wage laws, as distinct from merely acquiescing in the federal mandate. This led to such an increase in the variance of the independent variable that the confidence interval for a regression ending in 2007 is *half* the length of that of a regression ending in 2000.

We find that, fitted between 1976 and 2007, a regression of the youth employment-population ratio on state minimum wages (in logs), allowing only for two-way fixed effects (we call this the minimalist regression), returns a parameter estimate of -0.248 with a computed standard error of 0.040. This estimate of the elasticity is in the middle of the range of those from the pre-Card-Krueger literature; what is new is that an estimate of this magnitude and apparent statistical significance has not previously been observed in these data.[7] Figure 1 shows how the estimated coefficient and its confidence interval vary with the end-point of the sample.

Figure 1: Estimates and Standard Error Bands for the Minimalist Model with Different Ends of Sample.



[7] While some panel studies have found elasticities of this magnitude and higher (in absolute terms), they have had to resort to either omitting time fixed effects (Burkhauser *et al*., 2006), use additional control variables (Sabia, 2008), choose a later sample period and quarterly data plus a different employment category (*inter alia*, Dube *et al*. (2007), or focus on a smaller number of states (Sabia and Burkhauser, 2008).

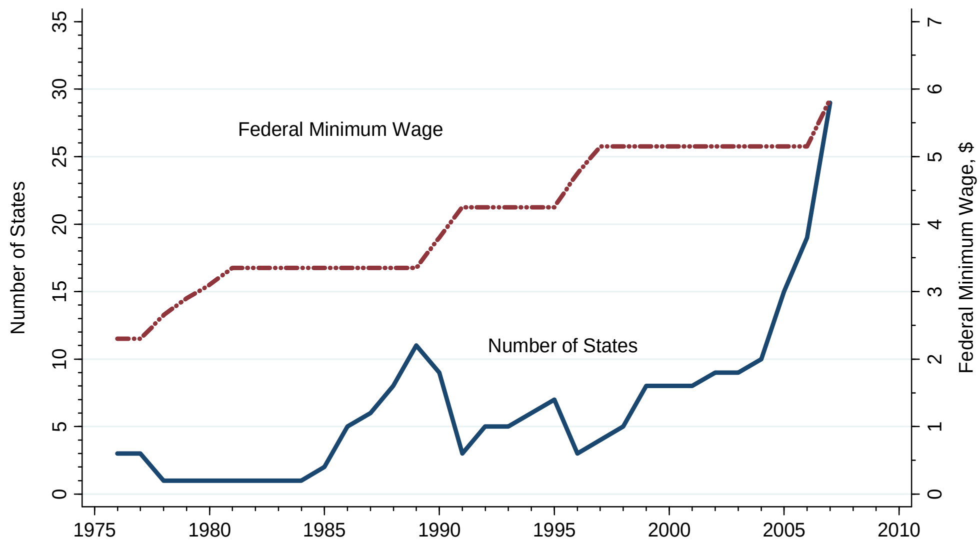
One observes the not-obviously-negative estimate from 1992, with the bulk of the confidence interval being positive. Over time, and particularly after 2000, the confidence interval has shrunk and migrated south, now living exclusively in the negative zone. Clearly the larger dataset has changed the game.

The paper is organised as follows. In the next section we discuss the evolution of state minimum wages and demonstrate the manner in which the post-2000 data make possible much stronger inferences regarding the parameter of interest. In section 3 we analyze the econometric problems involved with the interpretation of parameter estimates obtained from versions of (1). Section 4 presents the results. A final section concludes.

2. THE BEHAVIOUR OF STATE MINIMUM WAGES

In a regression of state employment on state minimum wages where one controls for time fixed effects, one obtains identifying variance only to the extent that some state minimum wages are greater than the federally mandated minimum wage. Figure 2 shows how the number of states with higher minimum wages has evolved 1976-2007.

Figure 2: States with Minimum Wages Greater than the Federal Level (December of Given Year)



At the beginning of 1976 just two states, Connecticut and New Jersey, had higher-than-Federal rates. In fact Connecticut state law specifies that its minimum wage will be at least one half of one percent greater than the Federal rate, often just one or two cents.[8] In October 1976, these two were joined by California, but the Federal rate was raised steadily between 1978 and 1981, and, since only Connecticut met this increase, she was left as the single exemplar until 1985. In January 1985 Maine raised her

[8] One naturally wonders why Connecticut has had such a strange arrangement. Communication with various scholars suggested plausible explanations, such as that Connecticut was relatively wealthy (it has had the highest per capita income for many years) with a liberal legislature and a relatively high cost of living. However, correspondence with the Assistant Director Ronald Marquis of the Connecticut Department of Labor revealed quite a different reason: the legislature changed the law as far back as 1969 in response to the so called Davenport Taxi case. The then Commissioner of Labor Renato Ricciuti had sued the company for failing to pay overtime wages to its dispatchers. “Davenport’s defence was that in crafting the FLSA, Congress denied individual states jurisdiction over the subject issues unless those states enacted laws providing *higher* coverage for the employees. Since Connecticut law provided only *equal* coverage, the employer prevailed.” (Ronald Marquis, communication by e-mail, July 22, 2008). Mr. Marquis then speculates that as a result of this ruling the legislature was concerned to establish its jurisdiction in future cases involving minimum wage levels and thus ensured that the state’s minimum wage would always be higher than the federal minimum wage. The point here is that Connecticut did not have a higher than federal minimum wage because of some conscious labor market policy, but rather as a result of an unrelated law suit, which, as it turned out, was subsequently found to be erroneous and of no legal consequence by a court in Alaska.

minimum wage above the Federal rate and, by the beginning of 1990, with the Federal minimum flat, the number of such states had grown to twelve. The Federal government raised its minimum wage twice in 1990 and 1991 which led to the number of states with higher rates falling to two by mid-1991. This cycle was repeated between 1991 and 1996, with the Federal minimum constant and an increasing number of states breaking ranks to join Connecticut with minimum wage rates higher than the Federal level. Once again the cycle was repeated: the Federal rate was increased in 1996 and 1997 and the number of states with rates higher than the Federal minimum fell from nine to three by September 1997. The Federal rate remained constant after this until August 2007, by which time 29 states had enacted higher-than-Federal rates. In the first part of 2007, fully ten states raised minimum wages above the Federal rate.

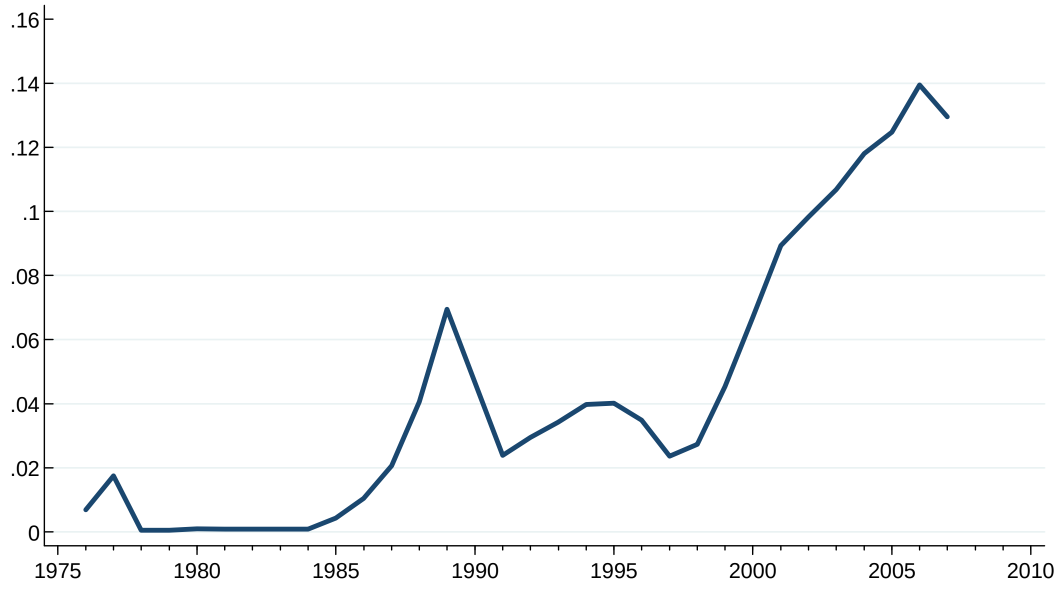
Active minimum wage policies tend to be pursued by states with a tradition of political liberalism, in particular the North-East, the West Coast, and an enclave to the west of Lake Michigan (Wisconsin, Illinois, Iowa, Minnesota). The South, those western states formed from U.S. Territories after the Civil War, and an enclave to the east of Lake Michigan (Michigan, Indiana, Ohio) typically do not pursue activist minimum wage policies. In the mass conversion of 2007 (plus the last months of 2006), 11 states enacted higher-than-Federal minimum wages for the first time:[9] Arkansas, Maryland, North Carolina, Missouri and West Virginia from the South, Arizona, Colorado, Montana and Nevada from the West, as well as Michigan and Ohio (Indiana held firm). These events followed the election victory by the Democrats in November 2006, having promised to raise the federal minimum wage during their campaign. However, independently six states (Arizona, Colorado,

[9] This would rise to 12 if Kentucky were included. Kentucky raised its minimum wage on June 26 in anticipation of the Federal increase in August. Since 1976, the month of July 2007 is the only time Kentucky's wage has exceeded the Federal level.

Montana, Missouri, Nevada, Ohio) had minimum-wage measures on the ballot, all of which passed, some of them by a relatively large margin (one poll quoted in *The Economist* saw 85% of respondents in favour of a minimum wage increase). This was a new development as economic policy of this nature had been typically set by state legislatures. Furthermore, the ballots included cost-of-living indexing. As a result 10 states in total had indexing by 2007.[10] Following the election, President Bush signalled that he would no longer oppose a federal minimum wage increase. A policy of high minimum wages seems to have become a hugely popular social movement.[11]

These cycles in minimum wages act to increase the cross-state variance in minimum wages. Figure 3 shows the coefficient of variation.

Figure 3: Coefficient of Variation of Cross-State Minimum Wages



Up until the mid '80s there was only limited variation in minimum wages across the states. The break-out of 1989 increased this variance. The most prominent feature in the graph is the large monotonic increase in the coefficient after 1998.

[10] The *Los Angeles Times* reports on November 8, 2006: "… [Critics] saw the measures as misguided attempts to put complex economic issues before voters when legislators really should have been deciding them." There was some subsequent speculation that Democrats added the measures to ballots to increase voter turnout, 'rallying the troops' so to speak.
[11] Dube *et al.* (2007) provide further useful discussion on the pattern of state minimum wage increases.

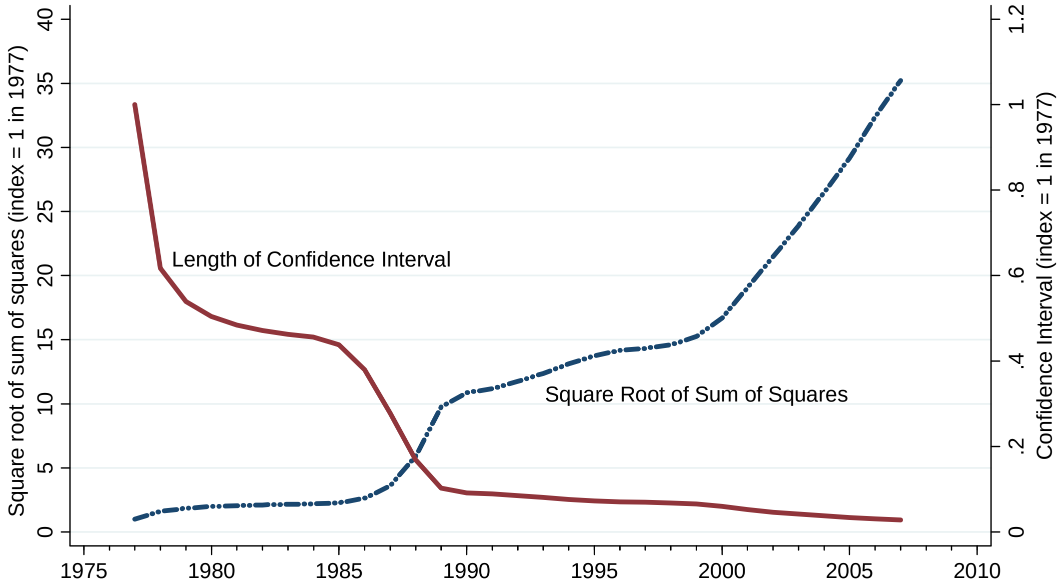
We intend to consider univariate panel regressions of employment on minimum wages (but controlling for two-way fixed effects and state specific trends). The OLS estimated standard error for the estimate of the elasticity is

$$\text{Standard error of estimate} = \sigma_\varepsilon \left(\sum_{i,t} mw_{it}^2 \right)^{-1/2} \qquad (1)$$

where mw_{it} is now the minimum wage having taken out two-way fixed effects, and σ_ε^2 is the equation error variance. The length of confidence intervals of estimates is proportional to (2). The first term σ_ε expresses the outcomes from experiments; the second term is independent of outcomes and depends only on the experimental design: it is thus an index of the effectiveness of the experiment for any given σ_ε. Figure 4 shows how this index varies with the sample period.

The graph gives the square root of the sum of squares as well as its reciprocal, a measure of the length of confidence intervals. These are normalised to be unity in 1977. Between 1985 and 1989, increased cross-state variation in minimum wages reduced confidence intervals by a factor of four. The next decade saw a reduction in confidence intervals by only about 30%. But, since 2000, confidence intervals have fully halved. Thus the years after the millennium offer a major increase in information concerning the relationship between minimum wages and employment.

Figure 4: Confidence Intervals for Estimates over different Time-Samples, 1976-end



3. ECONOMETRIC ISSUES

To obtain plausible predictions from such thought-experiments as a 1% increase in the state minimum wage, one requires a structural relationship that can in principle remain stable while such experiments are performed. Such a relationship will inevitably contain other conditioning variables. In a simple world where the wage was always set above the supply-demand equilibrium, so that employment was determined along a labor demand curve, these conditioning variables would be those that describe production technology: capital stocks and indices of technological progress are natural candidates. However as soon as one allows for the fact that most markets pay wages above the minimum wage (of the order of 90% of young workers according to BLS, 2006), all of the variables from the labor-supply side become relevant to determining aggregate state employment rates. Here the natural candidates are vectors of wealth, financial and non-financial (perhaps human capital). Note as well that, once one considers sub-markets, the conditioning technology variables are now specific to those sub-markets. One is ultimately led to an estimating equation of the form

$$ep_{it} = \beta mw_{it} + \alpha s_{it} + \varepsilon_{it} \qquad (2)$$

where the conditioning set *s* is a large vector of the stocks of capital and knowledge discussed above.

In practice, nothing like this detailed set of variables is observed and one is forced to resort to proxies. Allowing for two-way fixed effects takes account of U.S.-wide conditioning variables as well as once-and-for-all state effects. In the past (Keil *et al.*, 2000) we have added state trends in female labor participation, the share of youth in the population, and real wages; Neumark and Wascher (1992) include the proportion of youth in school but not working. One can interpret these and other approaches as seeking to account for non-minimum-wage trends in labor supply by readily-available data. The problem is that the wage elasticities so derived are not particularly convincing because the specification-search underlying them tends not to be an objective statistical procedure. Card and Krueger (1995), in particular, have made this a centre-piece of their criticism of the existing literature. We propose in this study to model the conditioning variables simply as error-components: two-way fixed effects plus state-specific trends. Specifically, we shall assume the effect of the conditioning variables is given by:

$$\text{Effect of conditioning variables} = p_i(t) + f_t + u_{it} \qquad (3)$$

where $p_i(t)$ is a state-specific polynomial in time, f_t is a time fixed-effect and u_{it} is a potentially autocorrelated error process. In this formulation we expect $p_i(t)$ to measure slow-changing stock variables and f_t to measure U.S.-wide variables, in particular the national business-cycle. For the most part, $p_i(t)$ will be a linear trend, so that this component consists of a state fixed-effect (the constant of the polynomial) and a state-specific trend. Thus we propose to represent the conditioning variables s by two-way fixed effects and state trends.

The inclusion of time and state fixed effects can be viewed as follows. First, from each state time-series, remove the average. Then, from the resulting variable, remove at each date the U.S. average. The regression parameter is recovered from a univariate regression between the transformed variables. In a logarithmic specification this amounts to regressing the proportional deviation of the employment-population ratio of each state at each date from the U.S. average at that date on the proportional deviation of the nominal state minimum wages from the U.S. average.

The main virtue of removing two-way fixed-effects (2WFEs) in our context is that the procedure will purge the underlying structural equation of U.S.-wide variables, which would otherwise be consigned to the residual. Such omitted variables will not create bias when they are uncorrelated with the variable of interest, but removing them may lead to greater econometric efficiency, provided the ensuing reduction in error variance more than off-sets the reduction in variance of the transformed variable of interest.

A perhaps more pertinent case arises when the omitted variables are correlated with the variable of interest, when their omission will lead to bias in the estimate of the parameter of interest. Assume that the observed youth employment-population ratio is comprised of three components:

1. macroeconomic factors associated with the business cycle (e.g. monetary shocks);
2. deep labour-supply variables determining the trend in youth participation;
3. minimum wage effects in sub-markets.

In large part, the observed behaviour of the employment-population ratio will be determined by the first two components, which are essentially unrelated to the minimum wage. Consider the first component. It turns out that the (real) minimum wage is somewhat pro-cyclical in the time-series of aggregate data and similarly pro-cyclical in panel data (see below). A natural interpretation of this is that legislatures are reluctant to increase minimum wages when unemployment is high. In the absence of any neo-classical minimum-wage mechanism, this form of endogeneity would cause the employment-population ratio to be *positively* correlated with the minimum wage. This will create a positive bias to the minimum wage parameter (i.e. towards zero if the true parameter is negative) which would be removed or ameliorated by time fixed-effects, to the extent that the Federal government is the major player in minimum-wage setting. This does not solve the problem of state sensitivity to the business cycle. Some experiments with instrumental variables suggest this source of bias is not particularly important. In any case, since the bias is towards zero, assuming the minimum wage is exogenous should make it harder to find negative minimum wage effects on unemployment.

The second component may lead to a similar bias. Assume for example the observed downwards trend in youth participation is driven by a trend increase in household wealth.[12] Given that compassion for the poor, even if misguided, is income-elastic, then trends in minimum wages will also be driven by wealth.[13] As with the business cycle, the inclusion of time fixed-effects will eliminate at least the U.S. component of the common wealth-effect, and thus reduce bias, even when state-

[12] This is presumably the most plausible explanation for reduced labor supply in youth and old age observable longitudinally in all capitalist economies.

[13] In 2006 the correlation between state GDP/capita and state minimum wages was 0.46 across the states.

specific trends are not included. Note that the likely direction of bias from wealth effects is negative, so that estimates are higher *in magnitude*.

A final virtue of time fixed-effects is that they turn the RHS of the regression equation into a *real* variable, namely the state minimum wage relative to the U.S. average, to match the LHS. An alternative is to deflate the minimum wage by average wages or prices, the choice here depending on the specific structural model underpinning the empirical analysis. The problem is that this introduces a strongly endogenous variable on the RHS. In some models the appropriate deflator for the minimum wage will be a U.S.-wide variable (a monetary variable say) in which case the inclusion of time fixed-effects with the nominal minimum wage gives the correct result. Even when this is not the case, it is quite likely that serious bias will not arise from assuming it is.

Finally dynamics. Equation (2) will produce an estimate of the instantaneous own-price elasticity of the minimum wage. An increase in minimum wages may in principle have a larger effect via second-round effects on the stock and structure of capital, but these are missed when these variables are treated as given. Even with given technology, adjustment costs will generate dynamic rather than static relationships, but it has been argued that such costs are likely to be small because of the ease with which employment levels may be adjusted for high-turnover workers (Brown *et al*., 1982; Card and Krueger, 1995). This would apply especially to young workers, the focus of our analysis. Adjustment dynamics may well be more important for older workers. We shall make some rudimentary allowance for dynamic relationships by allowing for serial persistence in the error structure.

4. YOUTH EMPLOYMENT AND THE MINIMUM WAGE

We begin with some data description. Figure 5 graphs the U.S. aggregate unemployment rate and the state averages of the (logs of the) youth employment-population rate and the minimum wage relative to the U.S. GDP deflator, 1976-2007. Longitudinal means have been removed.

Figure 5: Time Series of state Averages of the Real Minimum Wage, the Youth Employment-Population Ratio (logs), and the Aggregate Unemployment Rate.

There is perhaps a slight downwards trend observable in all three variables, most pronounced for the employment-population rate. The latter is well documented especially for the post 1998 period (on this see, *e.g.* Aaronson *et al.*, 2006). Taking the unemployment rate as a negative indicator of the business cycle, one sees that the youth employment-population ratio is strongly pro-cyclical, but that the real minimum wage is not obviously pro- or anti-cyclical. These impressions are confirmed by the correlation matrix in Table 1, where all variables have had a linear trend removed:

Table 1: Contemporaneous Correlations of Detrended U.S.-Wide Variables

	Unemployment rate	Real minimum wage	Employment/ population
Unemployment rate	1.00	-0.09	-0.79
Real minimum wage	-0.09	1.00	-0.08
Employment/ population	-0.79	-0.08	1.00

Rather than pass immediately to the model foreshadowed in (2) and (3), we find it instructive to begin with the simple models one might consider if one were to approach the data afresh. Table 2 reports a number of regressions of versions of the log of the youth employment-population ratio on versions of the log of the nominal minimum wage. Our sample consists of annual observations from the 48 contiguous states, 1976-2007, using averages constructed from monthly observations.

Model 0 is the benchmark, where the employment-population ratio is regressed on two-way fixed-effects only. One sees that these two explain about 86% of the variation in the (log of the) employment-population ratio. Model 1 is the "minimalist" regression where the log of the minimum wage is added. The minimum wage parameter is substantial and apparently well-determined. This is in contrast to both Card and Krueger (1995) and Burkhauser *et al*. (2000) who find the coefficient to be insignificant once time fixed effects (or year dummies in monthly data) are added. However, Addison *et al*. (2008) find a coefficient similar to ours in size using county level quarterly data for the restaurants-and-bars sector.

The tabulated (unadjusted) regression R^2 is calculated as the proportion of the variation in the dependent variable explained by the model, including any fixed-effects and any AR structure fitted to the residuals: thus the addition of the minimum wage increases the proportion explained by only 0.3%, despite a calculated *t*-statistic for the estimate of 6.2.

Model 2 considers the regression when only state fixed-effects are included with the minimum wage. The estimated parameter falls somewhat: this is consistent with the possibility raised in Section 3, namely that time fixed-effects eliminate a national business-cycle component which is positively associated with minimum wages. In contrast, the state component of minimum wages does not appear to create bias: if we instrument the minimum wage in the minimalist model by lagged minimum wages and the lagged state aggregate employment rate (both of which are good predictors of minimum wages), the minimum wage parameter is effectively unchanged at -0.230 (0.047). Model 3 fits no fixed effects. The minimum wage parameter remains negative and the regression R^2 indicates that the minimum wage on its own explains about 5% of the total variation in the youth employment-population ratio across time and states. Our elasticity is roughly half of that typically found by Burkhauser *et al*. (2000), who use monthly data, a shorter time period, and additional control variables. Sabia (2008), working with data for retail and small business employment, produces elasticities similar to ours over a longer sample period (1979-2004).

In the first three models the Durbin-Watson statistic[14] indicates substantial serial correlation in the residuals, which means that inference cannot be based on the computed standard errors. In Model 4 we allow the error process to be an AR(1): this can be seen as correcting the estimate in Model 1 for serial correlation. The estimated parameter is about a fifth smaller and the standard error about a quarter larger. Nevertheless, the estimate remains significant at conventional levels.

[14] Computed as the average of 48 state-specific Durbin-Watsons.

Table 2: Regressions of the log of the Youth Employment-Population Ratio on Variants of the Minimum Wage and Some Trend Controls

Model	***Specification***	***Current minimum wage***	***Time fixed effects***	***State fixed effects***	***R^2***	***Durbin-Watson Statistic***
0	no min wage 2WFE only	-	Yes	Yes	0.864	0.85
1	2WFE “minimalist”	-0.248 (0.040)	Yes	Yes	0.867	0.88
2	SFE only	-0.175 (0.010)	No	Yes	0.777	0.68
3	No fixed-effects	-0.166 (0.018)	No	No	0.053	0.16
4	2WFE and AR(1) errors	-0.203 (0.054)	Yes	Yes	0.909	2.14
5	2WFE and state trends	-0.134 (0.049)	Yes	Yes	0.900	1.14
6	2WFE, trends, AR(1) errors	-0.110 (0.057)	Yes	Yes	0.918	2.05
7	1st differences TFE	-0.111 (0.071)	Yes	No	0.245	2.64
8	2WFE and Kaitz’s index	-0.060 (0.027)	Yes	Yes	0.864	0.71
9	As in (6) for adults	-0.001 (0.012)	Yes	Yes	0.965	1.74

Note: standard errors in parenthesis, 1536 state-year observations.

The next model considers the possibility that state employment is conditioned by a state-specific trend which might be itself correlated with the state minimum wage. Model 5 allows for this with a state-specific linear trend. We observe that the estimate is about half that of the comparable (but no-state-trend) Model 1 lending some support to the above suggestion. However, the minimum wage coefficient remains negative and statistically significant within the consensus range. Some recent papers have argued that the negative effect of minimum wages is an artefact of the omission of state- (or county-) specific trends (Allegretto *et al*., 2008, Addison *et al*., 2008; Dube *et al*., 2007). In our data, including trends diminishes the effect of minimum wages but does not eliminate it[15].

There is evidence of serial correlation in the residuals and this is catered for in Model 6 where we allow for an AR(1) errors. The estimated parameter here falls to -0.110, with a one-sided *p*-level of 0.027.

Since there is no guarantee that omitted state-specific trends are well-modelled by a simple linear trend, one can consider polynomial trends, selecting the order by information criteria. In fact, the Schwarz Bayesian information criterion (BIC) favors the linear trend, as we have considered in Models 5 and 6.[16]

Model 7 considers another approach to the omitted variable problem by differencing the levels data rather than fitting trends. If the missing variables are I(1) processes, then this transformation will yield an unbiased estimator (under extra assumptions). This estimate is in strikingly similar to Model 6.

Model 8 replaces the minimum wage with Kaitz's index (minimum wage relative to state average earnings[17]). The estimate is about half those of Models 6 and 7 with a

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[15] The cited studies use a shorter sample period: from 1990 to 2007 for Allegretto *et al*. (2008), and from 1990 to 2005, quarterly data, for Addison *et al*. (2008). Dube *et al*. (2008), using quarterly data from 1990 to 2006, look only either at contiguous county pairs that are located on opposite sides of a state border with different minimum wages, or at multi-state metropolitan areas that have differences in minimum wages across their component counties. Addison *et al*. (2008) find some evidence of a negative minimum wage elasticity in the fast food ("limited service") sector when disaggregating their data further, even for the case when county specific trends are included.

[16] Using the Akaike criterion (AIC), however, would result in a ninth-order trend. If this is used in place of the linear trend one finds a parameter estimate (standard error) of -0.094 (0.090). There is no evidence of residual autocorrelation in this model and the estimate is of the same order as that in Model 6. There is however a marked increase in the standard error of the estimate, which is to be expected because the procedure will remove state-specific ninth-order time-polynomials from each state time series, and thus cause a large reduction in the variance of the explanatory variable. One perceives here the trade-off between bias-reduction and estimate-precision.

[17] Strictly speaking, the Kaitz index is the relative minimum wage multiplied by the percent of workers covered under the federal minimum wage in a given state. However, as Neumark and Wascher (2007; 14) point out: "given that the combined coverage of federal and state laws has been very high for some time, changes in coverage are not likely to offer much in the way of identification for samples limited to the 1980s and 1990s, suggesting that it may not be unreasonable for more recent studies to ignore coverage altogether. Indeed, much of the literature in the past decade or so has followed this approach."

p-level of 0.013. As we have pointed out above, the RHS variable is likely to be endogenous in this specification.

Finally, Model 9 considers adult employment. The estimate here is negative but very small. Note however that the standard error is small as well and the 95% confidence interval is (-0.025, 0.023). Thus the estimate and its standard error imply an upper bound on the magnitude of the adult elasticity of -0.025, so that, for adults, if there is a negative effect of the minimum wage, it must be small.

The standard panel estimator imposes three restrictions on a model:

1. the structural parameters are constant across cross-sections and over time;
2. the equation errors are independent over cross-sections;
3. the equation errors are independent over time.

Allowing for an AR(1) residual as in some specifications in Table 2 relaxes the third assumption; in Table 3 we consider the first two.

Table 3: Estimates of the Average Minimum Wage Elasticity when the Parameter is Allowed to Vary Across States

Model	***Specification***	***Average of 48 state estimates***	***Time- fixed effects***	***State- fixed effects***	***State-specific trends***	***Avg. Durbin-Watson***	***CSI z-test***
10	2WFE, no state trends	-0.231 (0.127)	Yes	Yes	No	1.08 (0.057)	1.12
11	2WFE and state trends	-0.200 (0.130)	Yes	Yes	Yes	1.31 (0.057)	0.61
12	SFE and state trends	-0.005 (0.252)	No	Yes	Yes	0.90 (0.057)	41.01
13	2WFE, state trends, GLS	-0.202 (0.114)	Yes	Yes	Yes	1.92 (0.146)	-

Model 10 allows for 2WFE and differing state minimum wage parameters. The estimation method is simply OLS applied to each state time-series. We report the unweighted average of parameter estimates together with the computed standard error of the average. This model differs from the minimalist Model 1 only in that the latter imposes parameter equality. One observes that the average estimate is similar to the minimalist estimate, though the associated standard error is much larger. There is rather less serial correlation in the residuals, suggesting that the imposition of parameter equality diverts some of the effect of the minimum wage in some of the cross-sections into the error process. Model 11 adds state-specific trends and corresponds to Model 5. Compared to Model 10, the minimum wage elasticity is somewhat lower and serial correlation in the residuals is lower.

Model 12 controls only for state fixed-effects and trends. As with the corresponding Model 2, the minimum wage parameter is much reduced. An interpretation of the estimates in Models 10-12 is that they describe the average over states of what a researcher in a single state would find using only the data from that state. Since the estimated elasticity in Model 12 is close to zero, we can infer that a researcher in a randomly-chosen state would have approximately only a 50% chance of finding a negative effect from a regression of de-trended employment on de-trended minimum wages. In Model 11 the researcher adds time fixed-effects which have the effect of transforming the variables to measures relative to the U.S. average (de-trended). In this case, though the mean of the population of estimates is negative and reasonably well-determined, this typical researcher would still have about a 40% chance of finding a *positive* elasticity (based on the family of estimates we have obtained).

The last column of Table 3 presents *z*-tests of the independence of error processes over cross-sections (cross-sectional independence, CSI), based on the sum of squared terms below the main diagonal of the 48×48 residual correlation matrix.[18] There is

[18] More specifically, the correlation matrix is formed from the residuals of state-specific regressions. Before computing the correlation matrix we first filter the residuals using the parameters from a state-specific AR(1). R. A. Fisher's tanh transformation is applied to the correlations to reduce them to approximate normality so that the sum-of-squares will be chi-square under the null of independence.

only limited evidence of cross-sectional dependence in Model 10 and even less in Model 11. In contrast in Model 12, one finds that not eliminating time fixed-effects leads to very strong cross-sectional dependence. It is natural to associate this with the macro U.S. business cycle, perhaps with federal monetary policy.

CSI effects a great simplification in the econometric environment because it facilitates GLS estimation of the panel. If CSI holds, the covariance matrix of the error process is block-diagonal with the 32×32 temporal covariance matrices of the states ranged down the diagonal. These can be estimated from the time series of each state and GLS straight-forwardly applied.[19] Estimating these temporal covariance matrices under the assumption that they are generated by state-specific AR(1)s, we obtain the GLS estimates in Model 13. The average elasticity is estimated as -0.202 with a standard error of 0.114, implying a one-tailed *p*-level of about 4%.

Starting with Model 13, one can test the assumption, implicit in the models in Table 2, that the elasticities are common across states. This entails 47 restrictions on the estimates in Model 13. Holding constant the covariance matrix, the Wald-test of these restrictions has a *p*-level less than 10^{-7}.

As feasible GLS, the estimator in Model 13 has asymptotically optimal properties and dominates the others considered so far. Its chief rival is Model 6, but the latter imposes parameter-equality, which we have seen is rejected by the data. An interesting facet of this is that the misspecification implied by parameter equality

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To produce a z-statistic, we use the approximation, also due to Fisher, that if X is distributed as chi-square with k degrees of freedom then $\sqrt{2X}$ is approximately normal with mean $\sqrt{2k-1}$ and unit variance.

[19] Things become trickier when there are correlations between states. Even if the errors are independent over time, so that simple SUR techniques are apparently available, the sample covariance matrix will be singular in our case, since we have more cross-sections than time periods. While there are recent techniques for dealing with this (Bai, 2006; Peseran, 2006), cross-sectional independence means we do not need to trawl these waters.

manifests itself as cross-sectional dependence: the z-test applied to the residuals from Model 6 yields a value of z = 3.78.

4. CONCLUDING DISCUSSION

The conclusion of this paper is that there now appears to be a strong negative correlation between minimum wages and youth employment detectable in the current panel of U.S. states. Our results are not dependent on a particular choice of RHS conditioning variables and are robust to the omission of any set of such that would be captured by state-specific time-polynomials and common U.S.-effects. Figure 1 describes how the increased variance in minimum wages has allowed more certain inference. A 1992 researcher would have found no significant minimum wage effects but, transported to the present, would find nothing inconsistent between his 95% confidence interval from the 1992 version of the minimalist regression (appropriately corrected for residual serial correlation) and a significant minimum wage parameter of around -0.2.[20]

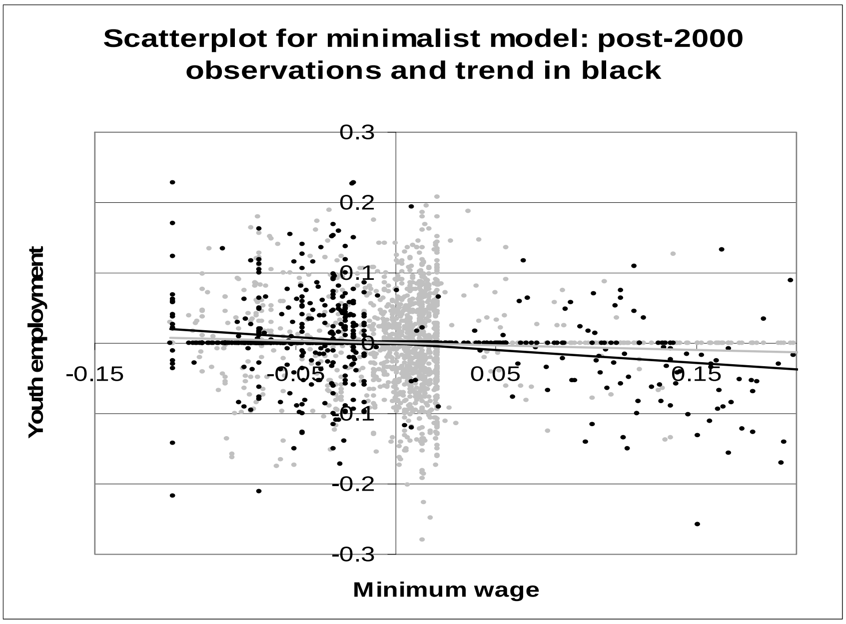
Figure 6 gives a scatter plot of employment and minimum wages (2WFEs removed) with the post-2000 observations in black and the earlier data in grey. One can discern in the South-East section of the diagram a collection of post-2000 points with a pronounced negative relationship between minimum wages and employment. It is precisely these new data that have proved crucial.

A natural question is whether these new data genuinely shed light on a pre-existing relationship or constitute a new phenomenon. We can offer an answer to this by a Chow test, splitting the data at 2000. We base the test on model 6 and hold constant the fixed effects, the trend parameters, and the autoregressive parameter at their full-sample values, thus allowing only the minimum wage elasticity to vary

[20] Note the standard errors in Figure are from the OLS and are not corrected for residual correlation.

across sub-samples. We find an *F*-statistic of 1.23, indicating no inconsistency across sub-samples. The point here is that the earlier data are essentially uninformative about the relationship and thus are consistent with stronger results in the later data. Similar results are found for splits at earlier years such as 1992.

Figure 6: Scatter Plot for the Minimalist Model.



Appendix – Data Sources and Construction

The data used in this paper were obtained as follows

MW – State Minimum Wages: Most of the monthly data were kindly provided by William Wascher, Federal Reserve of Philadelphia. Some of the earlier data in our sample period were collected from various state sources (e.g. Department of Industrial Relations, State of California, www.dir.ca.gov/Iwc/MinimumWageHistory.htm; or Department of Labor, www.state.me.us/labor/labor_laws/minwagehistory.htm). Monthly observations were used to calculate annual averages.

AHE – Average Hourly Earnings: Bureau of Labor Statistics, Quarterly Census of Employment and Wages (stats.bls.gov.cew.home.htm). The two series used are (i) SAEHE (state and area employment, hours, and earnings, SIC basis), 1976 – 2001; and (ii) EHCES (employment, hours, and earnings from the current employment statistics survey, 2001 – 2007). The two series were chain linked using the overlapping observations for 2001. There are some missing data: CO (1976), IN (1976-1988), KS (1976-1978), MD (2001-2002), MI (2001-2002), SC (2001-2002), and WV (2001-2002). We generated observations for the missing observations by back-casting using the following sample period regressions: CO, 1977-2006, explanatory variables: AHE(AZ) and AHE(NV); IN, 1989-2006, AHE(MI); KS, 1976-2006, AHE(MO); MD, 2003-2006, AHE(VA); MI, 2003-2006, AHE(IN); SC, 2003-2006, AHE(NC); WV, 2003-2006, AHE(KY).

EY, *E*, *POPY*, *POP*, –
Youth (16-19 years) Employment, Employment, Youth Population, Population: The source is *Geographic Profile of Employment and Unemployment*, Bureau of Labor Statistics, stats.bls.gov/gps/home.htm.

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