Abstract

This paper investigates the effect of minimum wages on employment using a panel of US state-based data. We estimate a minimalist dynamic version of the specification implied by neo-classical theory. We find statistically and economically significant effects of minimum wages on youth employment. Unlike many other studies we find also significant effects on aggregate state employment. These results re-establish the conventional wisdom as existing before the work of Card-Krueger-Katz. The paper meets the methodological criticisms of this sort of panel study made by CKK. An important econometric innovation in this paper is to produce estimates allowing for cross-sectional correlation, which offers unbiased inference and potential efficiency gains.

Key words: Minimum wages, panel data, cross sectional correlation, factor analysis.
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MINIMUM WAGES AND EMPLOYMENT

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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 labour 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 labour markets. Card and Krueger have been seriously challenged on methodological grounds (see, in particular, the comments by Brown; Hammermesh; 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]. The struggle for the hearts and minds of policy-makers seems to have been won in the US - and in the UK as well. The Federal minimum wage was increased substantially in late 1996 and there are plans for further increases as we write[3]. The British Labour government instituted a national minimum wage in 1999 and has recently increased it further. How permanent is this shift in fashion in ideas remains to be seen.

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 say - with a rather stronger and easier-to-find effect on youth employment. Typically, this evidence was derived from time series or panel studies of labour markets. In contrast, CK typically study an *event*. For example Card and Krueger (1994, 2000) study the increase in minimum wages in New Jersey in 1992. They find that, compared to Pennsylvania, where there was no such increase, employment in fast-food restaurants increased somewhat. While in some

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[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.

[3] The Federal minimum wage stood at \$4.25 in September 1996, \$4.75 in October 1996 and \$5.15 in September 1997. In the House of Representative version of the bill, this approximately 20 percent increase is to be followed by another roughly 20 percent increase to \$6.15 in two increments over two years.

specifications this effect is not statistically significantly positive, it is certainly not negative[4]. 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. One might compare it more precisely to establishing the link between smoking and lung-cancer, wherein the hypothesised effect is a modest increase in the probability of an already fairly infrequent event. The increased disease-rate for smokers might be easily observable in whole-population data but difficult to detect by study say of the fates of siblings in a handful of families, no matter how carefully conducted.

In this paper we provide empirical estimates of the effect of minimum wages on aggregate adult and youth employment rates from a panel of US states. Neumark and Wascher's 1992 article is the closest to ours in data and specification. For a panel of US states, they fit the employment rate for teenagers and young workers to the minimum wage relative to the average wage, and some other controls over the period 1975-1989. Essentially they find that in a simple specification, the minimum wage turns out to have a significantly negative effect on employment, but this can be reversed by the inclusion of fixed-effects. Minimum wages are found to be significantly negative in the presence of time and state fixed-effects only when they include as a right hand side variable the fraction of teenagers in school and not working. They were subject to an onslaught from Card and Krueger for the inclusion of this variable (see Card, Katz and Krueger 1994,and Neumark and Wascher 1994). The estimated equation is not developed from a precisely specified economic theory and, even if this were given, the schooling variable would surely be endogenous and require instruments. We steer clear of such criticism by estimating a minimalist specification: essentially the labour-supply schedule, modified by the inclusion of a variable to reflect the distortion in relative wages caused by a mandatory minimum wage. Our model thus contains two endogenous right hand side variables, the real wage and the relative minimum wage. One instrument is given by the nominal minimum wage itself; to obtain a second instrument, we argue that, since a large part of capital accumulation within specific states is essentially irreversible, lags of the real wage should be valid instruments. Our major econometric innovation is to allow for cross-sectional correlation in the US states. Conventional panel estimators impose zero cross-sectional correlation in the error process which, for employment regressions using state-based data, is almost certain to be violated - the unobservables are unlikely to be independent. Clearly if independence is assumed in estimation then potential SUR-type efficiency gains will be lost and inference will be suspect. Our solution is to specify a general factor structure for the error process, as described in Robertson and Symons (2000).

Our results support the conventional wisdom: employment responds to minimum wages with a measurable negative elasticity. In our preferred specification, a change in minimum wages has an elasticity of -0.11 for total employment and -0.37 for youth employment in the short-run, with long-run elasticities of -0.19 and -0.69

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[4]Though Neumark and Wascher (1995a,b,c) have argued that, while total employment may not have fallen, there was significant replacement of black and Hispanic youths by white youths dropping out of school, a remarkable example of the law of unintended consequences.

respectively. This difference between impact and total effects is a key finding and demonstrates the importance of the dynamic approach[5].

Card and Krueger argue in their book that time-series or panel studies of the sort we conduct are intrinsically dubious because of the opportunities offered for tendentious data-mining. It is presumably impossible to eliminate completely such doubts, but we limit potential for manoeuvre by basing our analysis on a stock-standard specification and experiment with a large number of estimation techniques and representations of the forcing variables. All in all, results appear to be robust to plausible perturbations in specification and estimation procedure.

2. Theoretical Analysis

2.1. **Employment of Different Types of Workers.** Panel studies develop response parameters from observations of an economy through time and we begin with a review of the appropriate neo-classical theory which might be taken to generate these observed data. We assume real state GDP is derived from a constant-returns production function

$$(1) \qquad\qquad y = f(k, s^*, u^*)$$

where k is physical capital and s^* and u^* are quantities in efficiency units of two sorts of labour ("skilled" and "unskilled" say). We assume

$$(2) \qquad\qquad s^* = \lambda_s s, \; u^* = \lambda_u u$$

where s and u are the observed quantities of the two sorts of labour and λ_s and λ_u are indices of labour-augmenting technical progress. We define the bias γ in technological progress as

$$(3) \qquad\qquad \gamma = \lambda_s / \lambda_u.$$

The analysis is facilitated by use of Samuelson's factor-price frontier, the relationship between the marginal products implied by CRS (plus some regularity):

$$(4) \qquad\qquad f_k = \phi(f_{s*}, f_{u*})$$

where $\partial \phi / \partial f_x = -x/k$, for $x = s^*, u^*$. Figure 1 depicts the factor-price frontier for a constant marginal product of capital. If the factors s and u are paid their marginal products, then the marginal product of capital is the profit rate per unit of capital.

We shall consider as a benchmark the case when the profit rate is constant, determined ultimately by the rate of subjective time-preference. With this assumption, each factor proportion s^*/u^* determines a position on the factor-price frontier in Figure 1 and hence the marginal products $f_x = \lambda_x w_x$ where w_x is the real wage of factor x. This implies

$$\gamma \frac{w_s}{w_u} = \frac{f_{s*}(\gamma s/u)}{f_{u*}(\gamma s/u)}$$

[5]This is also emphasised by Baker *et al* (1999) who find strikingly similar estimates of the long-run youth elasticity in a study of Canadian data. Fortin *et al* (2001) estimate dynamic unemployment rate equations for Canadian regions and find significant minimum wage effects for both female teenagers and females 20 and older, with long-run effects roughly three times the impact effect.
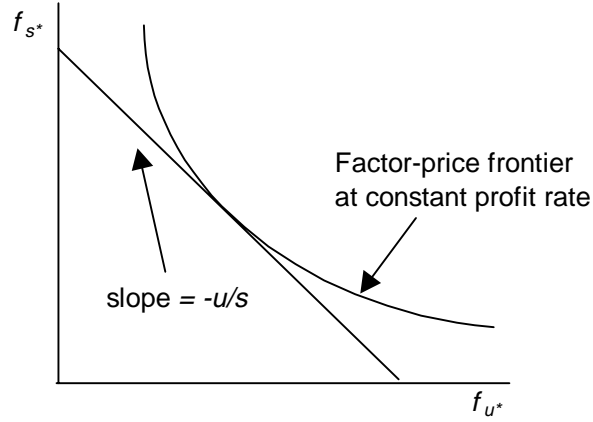
FIGURE 1. Factor Price Frontier at Constant Profit Rate

whence one obtains

$$(5) \qquad \frac{s}{u} = \mu(\frac{w_s}{w_u}, \gamma)$$

for some function μ. This relationship is what becomes of the familiar relative-wages-equals-MRS condition in a three-factor model when the assumption of separability in the labour inputs is not made. In this case one would usually find that the MRS depends on the stock of capital; conducting the analysis at a constant profit rate eliminates this dependence and leads to (5).

2.2. **Separability and no technological progress.** If we abstract from technological progress and assume the production function is separable in the labour inputs

$$(6) \qquad y = f(k, n(s, u))$$

where the index n is homogeneous of degree one in its inputs, then it is possible to give a geometrical representation of short- and long-run impacts of exogenous changes in w_u. In this case, the factor-price frontier takes the form

$$(7) \qquad f_k = \phi(f_n)$$

where f_n is the marginal product of the index n. For fixed k, the demand for n is

$$n^d = k\, g(w_n) \qquad g' < 0$$

Taking s as fixed (supplied inelastically) and regarding n as produced in a separate labour sector at price w_n, we obtain a demand for u

$$u^d = s\, h(w_u/w_n) \qquad h' < 0$$

and thus a supply of n

$$n^s = n(s, sh(w_u/w_n)) = s\, q(w_u/w_n) \qquad q' < 0$$

Figure 2 illustrates. The schedules n^s and n^d are indexed by k, and s and w_u, respectively. Initially the economy resides at A whereupon an increase in the
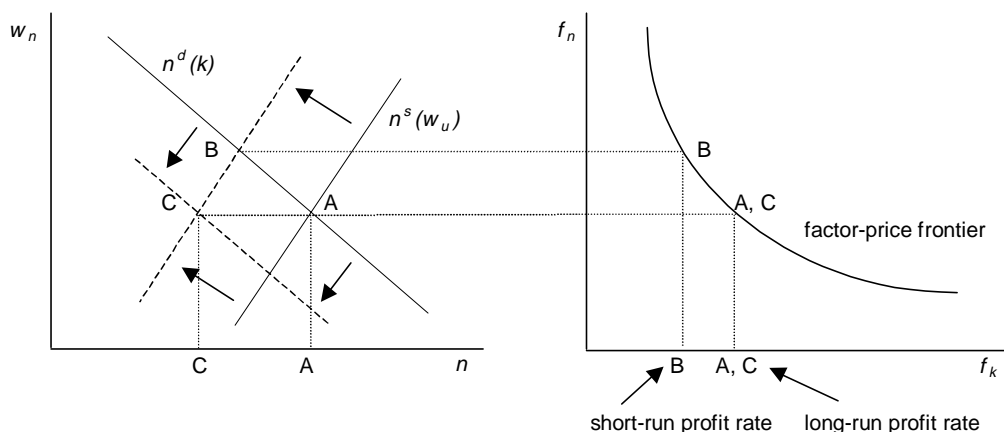
FIGURE 2. Effect of a shift in minimum wage

minimum wage causes w_u to rise and the economy moves to B. But at B the profit rate is less than long-run levels. As capital is eliminated, the n^d schedule shifts left until the long-run profit rate is restored at C. If s is inelastically supplied, reductions in n are associated with reductions in u via the sub-production function $n = n(s, u)$. In the long-run, w_s falls (from the factor price frontier - this implication does not require separability) but may rise or fall in the short-run depending on substitution possibilities, in particular the sign of f_{su}. In the separable case, w_s rises for fixed k if and only if

$$c_{su} - \sigma_k c_{nk} < 0$$

where $c_{..}$ refers to Hicks's elasticity of complementarity and σ_k is the share of capital[6].

One implication of this example is that one cannot infer the response to a once-and-for-all increase in the minimum wage from MRS relationships such as (5) because w_s cannot be taken as fixed. However one may use MRS conditions to infer the effect of permanent increases in the minimum wages relative to average wages say, which is perhaps the most natural experiment to consider.

3. METHODOLOGICAL ISSUES

3.1. **The estimating equation.** One could estimate (5) directly (or even the marginal productivity conditions) but it is not really clear what the appropriate measures of s and u should be. It is common to identify u with the young, but not all young workers receive the minimum wage and not all who receive the minimum

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[6]For more than two sorts of labour, one can obtain permanent increases in the real wages of some workers if they are sufficiently substitutable for u. For example, if a group of workers is perfectly substitutable for u, then their wages will always be proportional to w_u, provided some u are employed at all. (Note that the existence of perfect substitutes could result in all u-workers becoming unemployed, even in the short-run.)

wage are young. We argue instead as follows. Write $E = s + u$ for total employment. Then

$$(8) \qquad E/P \quad = \quad (s/P_s)(1 + u/s)(1 + P_u/P_s)^{-1}$$
$$(9) \qquad\qquad = \quad g(w_s, w_s/w_u, \gamma, P_u/P_s)$$

for some function g, where P denotes the relevant population. This is the equation we shall estimate. Essentially, (8) is a labour-supply relationship, modified to reflect any involuntary unemployment created by minimum wages.[7]

We propose that w_s be measured by state average hourly earnings in manufacturing relative to the US CPI, w_u by the mandatory minimum wage relative to the CPI, and P_u/P_s by the share of youth in the population. Technological progress γ is assumed to be non-state specific and appropriately modelled by (national) time trends (captured by time fixed effects). Another factor one must take into account in longitudinal data is the increase in female participation. Ideally one would like to have a measure of the deep factor, whatever it was, that caused female participation to increase so much over our period but, failing this, we shall include trend state relative female labour force.[8]

We shall treat both w_s and w_s/w_u as endogenous. Abstracting from minimum wage effects, one can think of w_s as the outcome of the interaction of labour demand and supply, so that the shift variables from the labour demand schedule are available as instruments, in particular, the stock of physical capital. Since physical capital is relatively immobile, it is natural to take *lags* of w_s as instruments, failing measurements of state capital stocks. In fact we construct an instrument for w_s as the fitted value from a *state-specific* regression of w_s on its lag. For w_s/w_u we take the nominal minimum wage as an instrument. Specifically, we take as the instrument in this case the fitted value from a state-specific regression of w_s/w_u on its lag and the change in the nominal minimum wage from the previous period.

3.2. **Dynamics.** So far we have been concerned essentially with equilibrium relationships but it seems fairly plausible that economies will at times be observed out of long-run equilibrium. It is sometimes suggested that lengthy adjustment to equilibrium is not plausible (Brown *et al.*, 1982). This is reasonable if the only implication of an increase in the real wage is to sack a few "burger flippers" but is not so clear if such an increase sets in train extensive substitution between low-skilled workers and capital and skilled workers, or, on the other hand, reallocation of capital and skilled workers to other activities. Consider for example a garment manufacturer who employs a number of workers at sweat-shop wages. If the minimum wage were to be increased, he may wish he had not bought the sewing-machines and taken a lease on the building, but if these are sunk costs, he may prefer to stay in business while the lease lasts, gradually disposing of the sewing-machines as opportunities arise, and letting employment run down. When the adjustment is complete, the physical capital (the building and the machines) and the skilled workers will be employed by other firms and some or all of the unskilled workers will be out of a

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[7]Note that the derivative of $g(.)$ with respect to its second argument does not reflect the full effect of a minimum wage change, because as argued above subsequent changes in capital stock will lead to level changes in w_s. However these are likely to be minor given that the minimum wage affects a relatively small proportion of the population.

[8]We use the fitted values from a state by state regression of female labour force relative to total on a quadratic in time.

job. Alternatively, the manufacturer may decide to replace the unskilled workers with computer-controlled equipment, which might well entail a lengthy period of planning, during which the firm might operate largely as before.

The most natural way to introduce dynamics is via adjustment costs which typically leads to adding a lag of the dependent variable and replacing the forcing variable by a weighted sum of all *future* values. In practice, one is forced to measure this weighted sum by the current value of the forcing variable which thus introduces measurement errors into the forcing variable and consequent downwards biases to estimated elasticities. An alternative is to specify an adjustment-cost technology and to estimate the Euler equation resulting from the inter-temporal maximisation. We prefer a more transparent methodology and shall merely append a lagged dependent variable to our estimating equation. The resulting bias towards zero of the estimated elasticities may not be severe since our relative minimum wage variable exhibits a great deal of variance.

3.3. Econometric Considerations. A stochastic and dynamic version of (8) is

$$(10) \qquad e_{it} = \lambda e_{it-1} + \beta' z_{it} + \varepsilon_{it} \qquad i = 1, ... n; \ t = 1, ... T$$

where e_{it} is the employment rate in state i at time t, z_{it} represents a vector of forcing variables, one of which will be the minimum real wage relative to average hourly earnings, another real hourly earnings, and ε_{it} is some error process that may contain state and time fixed effects, with perhaps contemporaneous correlation between different ε_{it}, ε_{jt}. We shall start with straightforward OLS estimation of (10) and these form our base results. But there are several further econometric considerations. Firstly one may be concerned about serial correlation in the error process which would lead to biases in estimation because of the lagged endogenous variable. There is some evidence of low-order serial correlation in the residuals from typical OLS experiments. This problem can be solved effectively[9] by using the second lag of the dependent variable as instrument for the lagged dependent variable and we report these results as well. A further problem with OLS estimation of dynamic fixed-effects panels is the bias described by Nickell (1981). This is particularly severe for small T panels, in which case it would be usual to take first-differences to remove the fixed-effects and then use an IV estimator to overcome the induced correlation between the differenced lagged dependent variable and the differenced errors (as suggested by Anderson and Hsiao 1981,1982). GMM techniques are also available that exploit the increased set of instruments as the panel advances (Arellano and Bond 1988, 1991). In this area there is always a compromise between the desire to reduce the bias of OLS, and the possible biases and lack of efficiency that may be introduced by the instrumental variables procedure, where the instruments may not be perfectly orthogonal to the errors and may be only poorly correlated with the endogenous variable in question. Even with quite short panels, it may well be preferable in root-mean-squared-error terms to use OLS estimation of the dynamic model. When the time dimension is more substantial, as in this paper (here T is about 20), or when there are strong forcing variables (as is also the case here), Nickell-bias tends to disappear and the case

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[9]This is true for the case of an MA(1) error structure, and will be approximately true for an AR process where the second-order correlation is small, as with an AR(1) of parameter about 0.2.

for straightforward OLS estimation is even stronger[10]. Despite this we shall also discuss Anderson-Hsiao estimation below.

The error process ε_{it} captures all unobserved effects on employment levels and any mismeasurement of the z_{it}. It is thus likely to have a complicated dependence structure across i as well as t.[11] Dependence over t can be treated by a combination of instrumental variables as above and perhaps quasi-differencing. Dependence over i is more problematical. One possible solution is to model spatial dependence by assuming some measure of distance between the cross-sectional units, thus imposing an order on the data and allowing dependence patterns to be modelled in the covariance matrix of the errors. This seems unlikely to be an accurate description of the dependence patterns for state data, in that certain states will be close by some economic criteria but distant by others *i.e.* economic distance is not a univariate concept. An alternative is the SUR technique wherein it is left to the data to determine the cross-sectional dependence. The problem with SUR here, where there are more cross-sections than time periods, is that the standard technique of estimating the covariance matrix will not produce an invertible matrix, and, in any case, tends to be too profligate with the number of estimated parameters. Robertson and Symons (2000) propose implementing SUR type regressions in a situation where the standard estimator of the covariance matrix is rank-deficient by imposing a factor structure on the residuals and using maximum likelihood techniques to recover an invertible estimate of the matrix.

To see how this is done, re-write (10) as

$$(11) \qquad e_t = \lambda e_{t-1} + z_t \beta + \varepsilon_t \qquad t = 1, ...T$$

where $e_t = (e_{1t}, ..., e_{nt})'$ and $\varepsilon_t = (\varepsilon_{1t}, ..., \varepsilon_{nt})'$ are $n \times 1$ vectors, z_t is an $n \times v$ matrix of explanatory variables observed at t, β is a vector of unknown parameters to be estimated, ε_t is a vector white-noise process with $E(\varepsilon_t \varepsilon_t') = \Sigma$, and $E(z_{ijt} \varepsilon_{kt}) = 0$ all i, j, k, t.

The factor assumption is that

$$(12) \qquad \Sigma = \Lambda \Lambda' + \Psi$$

where Λ is a $n \times p$ matrix of so-called factor loadings and Ψ is a diagonal $n \times n$ matrix with diagonal elements $\psi_1, \psi_2, ...\psi_n$, where $\psi_i > 0$ reflects idiosyncratic effects. This allows for some contemporaneous correlation between shocks, but expressed as a function of fewer parameters than the unconstrained Σ if $p < n$. Note that the factor model generalises fixed-effects models directly, as the fixed effects can be entered as elements of z.

The factor model amounts to specifying that the residuals take the form $\varepsilon_t = \Lambda_1 \phi_t^1 + \Lambda_2 \phi_t^2 + .. + \Lambda_p \phi_t^p + u_t$ where $E(\phi_t^i \phi_s^j) = \delta_{ij} \delta_{ts}$ and $E(u_t u_s') = \delta_{ts} \Psi$ and Λ_k is a column vector of weights. The ϕs can be interpreted as p common shocks and the elements of each Λ_k give the loading or impact of each of these shocks on each of the cross-sectional units. These common shocks provide cross-sectional correlation

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[10]See Grubb and Symons (1987) for a discussion of bias in dynamic models, in particular the bias-reducing properties of significant forcing variables..

[11]Note also that parameter heterogeneity in the underlying model (i.e. the slope coefficients β vary across cross sectional units) could be interpreted as a homogenous slope model with a more complex error covariance structure. That is, if the model is $e_{it} = \beta_i' z_{it} + \varepsilon_{it}$ we can rewrite as $e_{it} = \beta' z_{it} + ((\beta_i' - \beta') z_{it} + \varepsilon_{it})$ so that even if the ε_{it} are uncorrelated across time and units, the compound error in the homogenous model $(\beta_i' - \beta') z_{it} + \varepsilon_{it}$ will have a correlation structure reflecting correlations in the forcing variables z

in the error structure, with the u_t adding an idiosyncratic term. Estimation of the model (12) is by maximum likelihood. The details, including selection of the number of factors p, are discussed in Robertson and Symons (*op. cit.*).

4. EMPIRICAL RESULTS

We estimate our model on annual data for the 48 contiguous US states over the period 1977-1995. Full data sources and description are in the data appendix. We present two sets of results, one with dependent variable the ratio of employment to population, the other with youth employment relative to youth population, since it is often found that minimum wages impinge more strongly on youths[12]. Results for employment/population are in Table 1; results for youth employment/population are in Table 2.

In Column 1 of each Table we regress the employment rate on a lagged dependent variable, the relative minimum wage and two demographic variables, the trend share of women in the labour force and the share of youth in the population. State fixed effects are included, but not time fixed effects. The minimum wage is negative and significant in this regression for both the aggregate employment specification and for youths. When, however, we add time fixed effects in Column 2, the sign on the minimum wage is reversed in both cases. This essentially reproduces the result of Neumark and Wascher (for teenagers and young adults) discussed in the Introduction. In Column 3, in accordance with the interpretation of the equation as a quasi labour-supply relationship, we add the real consumption wage. The sign of the minimum wage becomes negative again. In columns 4, 5 and 6, we experiment with Instrumental Variables. Serial correlation in the residuals, as we tend to find in all earlier experiments[13], will in general invalidate lags of the endogenous variables as instruments, but when this serial correlation is low and of first order only, as we also find, *second* lags of endogenous variables will be approximately orthogonal to the equation error (assuming of course that the innovation in the error process is orthogonal to variables dated $t-1$). In Column 4 we use as instruments state-specific forecasts of the two endogenous variables based on information dated $t-1$. We instrument as well the lagged dependent variable by its lag. In these experiments the minimum wage parameters are both negative and both comfortably more than two standard errors from zero. In Column 5 we include as extra instruments forecasts based on $t-2$. The equation is now over-identified so it is possible to examine the orthogonality between the instruments and the residuals. A regression of the residuals on the instruments gives a P-value of 0.52 for the aggregate specification and 0.55 for youths so in both cases the instruments do seem appropriately orthogonal. Column 6 then gives results for instruments dated only at $t-2$. Throughout the IV experiments, for both the aggregate employment and youth specifications, the minimum wage parameter is significantly negative and of an important magnitude.

We proceed to the Factor-GLS estimates as follows. There is evidence of low-order serial correlation in the residuals in Columns 6 of both Tables 1 and 2. Since such correlation could be equally well modelled as an MA(1) or an AR(1), we choose

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[12]The specification is in levels, not the more usual logs. Neumark and Wascher (1992) find a level specification preferable in a Box-Cox analysis.

[13]The estimated first, second and third order serial correlations of the pooled residuals are reported in the Tables.

the latter and quasi-first-difference the data in (10) using the estimated first-order serial correlations. These quasi-differenced residuals represent the shocks. For the aggregate employment series, the average modulus of off-diagonal elements in the residual correlation matrix is 0.216. Under the null of zero cross-sectional correlation one can base a χ^2 test on the sum of squared elements below the main diagonal; using Fisher's $tanh$ transformation and approximating the average of χ^2 by a normal, one obtains a statistic distributed as standard normal, with large values being evidence of dependence. Here $z = 3.1$. This suggests that if one neglects cross-sectional dependence there could be problems for inference. To deal with this, a factor-decomposition along the lines of (12) is sought for the covariance matrix of these transformed residuals. The first step is to find the appropriate number of factors. Robertson and Symons (2000) report from Monte-Carlo study that information criteria such as Hannan-Quinn or AIC applied to the maximised log-likelihood (as the chosen number of factors varies) give good results in identifying the correct number of factors in (12). Care is needed in the case when $n > T$ (as here) because the likelihood grows without bound as the number of factors is increased to T, but it is natural to choose a local maximum of the information criterion. In our data the AIC shows a local maximum at three so that our estimate of Ψ is for a three-factor model. Applying this estimate of the variance-covariance matrix to the quasi-first-differenced data yields the model in Table 1, Column 7. The fit is broadly similar to Columns 4 - 6 but is less dynamic, has smaller standard errors, and, as indicated by the Jarque-Bera test, has normally distributed errors[14]. The long-run elasticity is -0.190 and is quite tightly estimated. All in all, we feel the estimates in Column 7 are the most reliable. Throughout these experiments the elasticity of participation with respect to real wages is estimated to be of the order of -0.4.

For the youth employment specification the same procedure was used. Here the IV experiments point to negative serial correlation. Whilst this may not have an obvious economic rationale, it could plausibly arise from the aggregation of sub-processes with different dynamics, as might well be the case here. Again there are substantial off-diagonal elements in the residual correlation matrix; which in this case are well-represented by a one factor model. The results are given in Table 2 Column 7. As before, the dynamics are somewhat reduced, the parameters are more precisely estimated, and the Jarque-Bera suggests normality of the residuals. For youths the effect of the minimum wage is found to be quite quick-acting - a mean lag of a little over a year typically - with rather larger elasticities than Table 1 implies for aggregate employment.

We conclude that the employment effects of the minimum wage act similarly for youths and the whole population, with a somewhat larger and faster-acting effect for youths. The results for youths have been found before by others; the results for aggregate employment are more novel.

5. Robustness of Results

The above section has set out empirical results and considered the various econometric issues that impinge on this type of dynamic panel study. Different estimation

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[14]The failure of the Jarque-Bera tests elsewhere could be taken as evidence, not of non-normality of the errors, but of their cross sectional dependence, which is allowed for only in Column 7.

techniques by and large deliver the same message. In this section we consider the sensitivity of our results to the variables included in the regression, the presence or absence of trending variables, the treatment of instruments etc. We concentrate on the regression for aggregate employment throughout. See Table 3.

As benchmark we employ the model from Table 1, Column 6, which has time and state fixed-effects and instruments dated $t-2$ as discussed above: This model is potentially less efficient than the factor model of Column 7 but the methodology is more transparent. In Model (2) we add union density and the unemployment benefit replacement rate as explanatory variables. The elasticity barely changes. In Model (3) the demographics are excluded so that the included variables are forced, in particular, to explain the trend increase in employment participation. This seems obviously false, but, nevertheless, the minimum wage elasticity stays much the same. In Model (4) we exclude the lagged dependent variable so that the static regression is estimated. The minimum wage parameter remains negative and nominally significant. This estimate is by IV: in Model(5) the same specification is estimated by OLS. Now the minimum wage switches sign and becomes insignificant. Thus endogeneity is an important issue in detecting minimum wage effects, as we found in Table 1. In these last two experiments the residuals exhibit strong serial correlation ($\rho(1) = 0.73$ in Model (4)). Model (6) addresses the issue raised by Card and Krueger in their book that the choice of deflator for the minimum wage might itself bogusly introduce the business cycle into the minimum wage variable. In the benchmark, minimum wage is measured relative to average earnings in manufacturing, with lags of this variable used to create the instrument. It could be argued that this procedure does not remove the business cycle so in Model (6) we replace average earnings by the fitted values of state-specific regressions on quadratic trends and similarly for the CPI in the denominator of the real wage, and estimate by OLS. The point estimate of the minimum wage effect is largely as before though the standard error is much increased. The final experiment in Table 3 excludes fixed-effects but introduces US-wide variables into the equation, specifically, the US nominal interest rate, the US inflation rate, the US aggregate unemployment rate and a US-wide quadratic trend. We find that the minimum wage elasticity is much the same as in the benchmark (as are the unreported point estimates).

We have experimented as well with varying the sample, 1977-1995 in the benchmark. Holding the first observation fixed at 1977 and reducing the endpoint by three year intervals, we find that the point estimate remains negative and more than two standard errors from zero until the sample ends at 1983 whereupon the point estimate, though negative, becomes less than its standard error. We repeated this exercise for the beginning-point of the sample. When the model is fitted 1980-1995, the elasticity falls to -0.124(0.089); when the sample begins at 1983 or 1986, the point estimate of the elasticity is *positive*, about 0.1 with a standard error of about 0.05 in both cases. Further increases in the beginning-point yield negative elasticities again, though the standard errors now become large. Thus minimum wage effects appear not to be readily detectable if one relies only on more modern data.

Finally it may be argued that our estimates suffer from dynamic fixed effect panel biases. We performed a variety of Anderson-Hsiao experiments using either lagged levels or lagged differences as instruments; where serial correlation of the residuals means that the lags must be deeper than usual. Confidence intervals for

the forcing variables tended to be very wide, but more importantly for our purpose, whilst the lagged dependent variable was estimated to be somewhat larger than its value in the benchmark (roughly 0.77 against 0.72), consistent with a downward bias in the OLS, the Anderson-Hsiao estimate is within two standard errors of the benchmark supporting the view advanced above that Nickell biases are likely to be minor.

6. Conclusion

There are major disagreements amongst economists about the impact of minimum wages, and this debate has serious policy implications. In this paper we obtain a specification consistent with an underlying theory of the labour market, and stress the importance of allowing for dynamic adjustment in the reponse of employment to minimum wage changes. Our data embody the full experience of aggregate US labour markets over the past two decades and our error specification allows a rich structure for those influences that the econometrician can never observe.

In our preferred specification, a change in minimum wages has an elasticity of -0.11 for total employment and -0.37 for youth employment in the short-run, with long-run elasticities of -0.19 and -0.69 respectively. These results are robust to reasonable perturbations in specification and statistical methodology. Positive elasticities appear only in models that are clearly misspecified. Our finding of significant negative elasticities for total employment, and the magnitude of the long run elasticites, demonstrate that the cost of minimum wage legislation, far from being negligible as claimed by its apologists, may be higher still than even the minimum wage hawks have argued.

Table 1 Estimation Results for Aggregate Employment
Dependent variable=employment/population

	Estimation Method						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
estimation method	OLS	OLS	OLS	IV	IV	IV	IV-GLS
constants fitted	α_i	α_i & f_t	α_i & f_t	α_i & f_t	α_i & f_t	α_i & f_t	α_i & f_t
instruments dated t	-	-	-	-1	-1 & -2	-2	-1
lagged dep var	0.766 (0.022)	0.785 (0.023)	0.794 (0.023)	0.709* (0.030)	0.711* (0.030)	0.720* (0.031)	0.652 (0.027)
min wage/ahe	−0.053 (0.013)	0.051 (0.017)	−0.050 (0.033)	−0.089* (0.036)	−0.078* (0.035)	−0.129* (0.044)	−0.105* (0.028)
ahe/cpi	-	-	−0.676 (0.187)	−0.951* (0.212)	−0.870* (0.207)	−1.270* (0.276)	−1.074* (0.179)
share women in lab force	0.278 (0.049)	0.112 (0.072)	0.074 (0.072)	0.075 (0.073)	0.080 (0.073)	0.048 (0.075)	0.177 (0.063)
share youth in population	0.086 (0.060)	0.056 (0.058)	0.014 (0.059)	−0.013 (0.060)	−0.007 (0.060)	−0.033 (0.061)	−0.006 (0.051)
$\rho(1)$	0.16	0.15	0.13	0.20	0.20	0.20	-0.04
$\rho(2)$	−0.09	−0.08	-0.10	−0.04	−0.04	−0.04	0.04
$\rho(3)$	−0.13	−0.04	−0.04	-0.02	-0.02	-0.02	0.01
Jarque-Bera	0.82	11.37	12.45	9.80	10.09	8.70	3.08
min wage elast	−0.140 (0.037)	0.147 (0.053)	−0.150 (0.101)	−0.190 (0.080)	−0.168 (0.079)	−0.286 (0.106)	−0.190 (0.051)
real wage elasticity	-	-	−0.448 (0.138)	−0.446 (0.114)	−0.411 (0.112)	−0.620 (0.161)	−0.424 (0.076)

Notes.

(i) standard errors in parentheses

(ii) variable descriptions:

minwag/ahe is state minimum wage divided by state average hourly earnings in manufacturing

ahe/cpi is state average hourly earnings in manufacturing divided by US national CPI.

share women in lab force is state female labour force as proportion of state total labour force, fitted by state specific quadratics in time

share youth in population is state youth population as proportion of total state population

(iii) variables denoted * are treated as current endogenous and instrumented by their fitted values from state-by-state regressions. For the real wage these regressions were on the lagged dependent variable (lagged once or twice as indicated in the Table) and a constant. For the relative minimum wage, the regression included as well the change in the nominal minimum wage (from $t-1$ or $t-2$, as appropriate).

(iv) Estimated over 48 States, 1977-1995 annual data.

(v) IV-GLS uses Robertson and Symons's (2000) SUR method by fitting a three factor model to the covariance matrix of quasi-differenced residuals from column (6). This (invertible) matrix is then used in a SUR procedure on the quasi-differenced data.

(vi) $\rho(1)$ etc. are estimates of the serial autocorrelations in the residuals. Jarque-Bera is test for normality, distributed as χ_2^2.

(vii) Elasticities are calculated at sample means.

Table 2 Estimation Results for Youths
Dependent variable=youth employment/population of youth

		Estimation Method					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
estimation method	OLS	OLS	OLS	IV	IV	IV	IV-GLS
constants fitted	α_i	$\alpha_i \& f_t$	$\alpha_i \& f_t$	$\alpha_i \& f_t$	$\alpha_i \& f_t$	$\alpha_i \& f_t$	$\alpha_i \& f_t$
instruments dated t	-	-	-	-1	-1 & -2	-2	-1
lagged dep var	0.560 (0.029)	0.482 (0.031)	0.486 (0.031)	0.608* (0.063)	0.618* (0.062)	0.614* (0.063)	0.565* (0.028)
min wage/ahe	-0.122 (0.036)	0.062 (0.053)	-0.133 (0.101)	-0.290* (0.110)	-0.242* (0.109)	-0.433* (0.136)	-0.367* (0.094)
ahe/cpi	-	-	-1.307 (0.573)	-2.452* (0.650)	-2.118* (0.636)	-3.265* (0.855)	-3.001* (0.619)
share women in lab force	0.377 (0.128)	0.620 (0.220)	0.547 (0.221)	0.368 (0.230)	0.379 (0.230)	0.338 (0.233)	0.605* (0.158)
share youth in population	0.695 (0.169)	0.213 (0.179)	0.131 (0.182)	0.083 (0.185)	0.107 (0.185)	0.036 (0.190)	0.187* (0.145)
$\rho(1)$	-0.01	-0.03	-0.04	-0.15	-0.17	-0.15	-0.01
$\rho(2)$	-0.03	-0.01	0.01	-0.04	-0.05	-0.04	-0.06
$\rho(3)$	-0.09	-0.05	-0.05	0.03	0.03	0.02	0.02
Jarque-Bera	13.44	18.60	18.61	23.32	23.76	23.24	2.34
min wage elast	-0.226 (0.070)	0.098 (0.083)	-0.212 (0.101)	-0.603 (0.252)	-0.517 (0.251)	-0.915 (0.324)	-0.692 (0.174)
real wage elasticity	-	-	-0.457 (0.138)	-1.123 (0.351)	-0.996 (0.347)	-1.520 (0.464)	-1.241 (0.253)

Notes as for Table 1. For the IV-GLS a one factor model was specified.

Table 3 Robustness Experiments
Dependent variable=employment/population

Model	Specification	Elasticity of Min Wage
(1)	Benchmark from Table 1, Col 6	$-0.286(0.106)$
(2)	Benchmark with benefits & union density	$-0.286(0.104)$
(3)	Benchmark without women & youth	$-0.285(0.106)$
(4)	Benchmark without lagged dependent variable	$-0.147(0.044)$
(5)	Benchmark without ldv, estimated by OLS	$+0.015(0.032)$
(6)	Benchmark with ahe and cpi replaced by trends	$-0.226(0.170)$
(7)	No fixed effects but US-wide macro variables	$-0.243(0.092)$

Appendix -Data Sources and Construction

The data used in this paper were obtained as follows

1. Employment and population: The source for labor force, population, and employment for youth, females, and adults is *Geographic Profiles of the Employed and Unemployed*. Some of the data can be retrieved from the U.S. Department of Labor Web site at http://stats.bls.gov:80/top20.html ("most requested series").

2. Average hourly earnings: State manufacturing average hourly earnings. The data are available at the U.S. Department of Labor Web site at http://stats.bls.gov:80/top20.html ("most requested series").

3. Benefits data are the product of the replacement rate (average weekly benefits in covered employment/average weekly wage) and UI coverage (average number of weekly insured unemployed people/annual average number of unemployed). All series from U.S. Department of Labor, Bureau of Labor Statistics. These data were kindly provided by Louis Pantuosco.

4. Union density: Percentage of the state labor force who are union members. Prior to 1982, the data is from *Statistical Abstract of the United States* (U.S. Department of Labor, Bureau of Labor Statistics). After that, the data were extracted from the CPS tapes by Hirsch and McPherson (unpublished data).

5. State minimum wages: The data were kindly provided by William Wascher, Federal Reserve of Philadelphia.

6. US consumer price index and unemployment: *Economic Report of the President*.

7. Nominal interest rate: We chose the six months commercial paper rate from the *Economic Report of the President*.

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