Looking for a needle in a haystack? A structural time series model of the relationship between teenage employment and minimum wages in the United States*

Stephen Bazen*

Velayoudom Marimoutou**

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*GRAPE (CNRS UMR 5113) Université Montesquieu Bordeaux IV, Avenue Léon Duguit, 33608 Pessac, France - Corresponding author.
e-mail: bazen@montesquieu.u-bordeaux.fr

** GREQAM , Université de la Méditerranée, Marseille, France.
email: marimout@ehess.cnrs-mrs.fr

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Abstract

The work of Card and Krueger has cast doubt on the nature of the relationship between the minimum wage and teenage employment. The earlier "consensus" finding of a small but statistically significant negative effect was based on time series data whereas Card and Krueger's findings are based mainly on cross section data. In this article, we re-examine the time series relationship between minimum wage and teenage employment. We find that previous models break down due to their inability to capture changes in the trend, cyclical and seasonal components of teenage employment. We propose a structural time series model in which these components are treated as stochastic components and which contains the traditional approach as a special case. The model when estimated up to 1979 accurately predicts what happens to teenage employment subsequently, when the minimum wage was frozen after 1981 and then increased quite substantially in the early 1990s. Moreover, we find that there is a significant, negative effect of the minimum wage on teenage employment and its size and significance have hardly changed during the 1980s and early 1990s. Finally, the model remains robust in an out-of-sample test for 1993-99 containing two further minimum wage hikes.

Introduction

In their survey of studies covering the period upto 1981, Brown, Gilroy and Kohen (1982) (BGK) concluded that a 10% increase in the minimum wage reduces teenage employment by one to three per cent in the United States. The vast majority of the studies covered used time series data. However, recent work on the effect of increases in minimum wages on employment has shed considerable doubt on these earlier findings for the United States. There appears to be no significant negative employment effect associated with increases in minimum wages that were implemented in the early 1990s, and in certain sectors employment may actually have increased (see Card and Krueger,1995; Bernstein and Schmitt, 1998). The latest research has used mainly cross-section data and often refers to the experience of a particular sector or particular states. Not unexpectedly, these conclusions have been questioned on a number of fronts such as the quality of the data used and the interpretation of the results obtained (Deare, Murphy and Welch, 1995; Neumark and Wascher, 1996).

However, as the results are at odds with previous findings, it is important to see whether this is because there were problems with previous research, or whether the underlying relationship between teenage employment and minimum wages has changed. The distinguishing feature of the work summarised in the survey by Brown, Gilroy and Kohen (1982) is that it largely concerns time series data for teenagers up to the end of the 1970s. In their recent book, David Card and Alan Krueger re-examine the earlier research and claim that compared to earlier work, extending the observation period up to 1993, on the basis of time series evidence "the minimum wage has a numerically smaller and statistically insignificant effect on employment" (1995, p.205). But as Daniel Hamermesh (1995) points out, all of the estimates are negative and John Kennan (1996) likens isolating the impact of minimum wages on teenage employment using time series data to "looking for a needle in a haystack" (p.1955). Charles Brown (1995), for example, suggests that there may be a significant negative impact in the long run which is simply not picked up in a cross-section setting.

In this paper, we analyse whether or not the information that can be extracted from a time series analysis tells a different story from the cross section studies. First, we re-examine the basis of the apparent consensus observed by studies undertaken upto 1980. We analyse the stability of these models and expose their inability to account for observed changes in teenage employment in the 1980s and early 1990s (in Section I). Secondly, we propose an alternative empirical model that overcomes the shortcomings of previous studies which use time series data (Section II). We find that a structural time series model that treats the unobservable trend, cyclical and seasonal components of teenage employment as stochastic not only successfully accounts for changes in teenage employment in the period since 1980, but also attributes a significant, negative employment effect to rises in the minimum wage (in Section III). Finally, in an out-of-sample test for 1993-99 containing two further minimum wage hikes, the model remains robust.

I The Basis of the "Consensus"

As the studies surveyed in BGK (1982) used different time periods and equation specifications, in a separate paper, the same authors (BGK,1983) used a common quarterly data set (1954:1 to 1979:4) to examine the impact of different specifications on the size and significance of the estimated effect of minimum wages on teenage employment. In this work they found that (a) the elasticity was more often around -0.1, and (b) there was substantial residual autocorrelation in the different specifications. Estimation using AR(1) "correction" however was found to reduce the significance of the estimated effects below conventional levels in the majority of specifications.

In a subsequent study, Gary Solon (1985) found that residual autocorrelation was due in part to the inadequate manner in which seasonality had been modelled and managed to restore the significance of the estimated impact by adding seasonal dummies interacted with a time trend and time trend squared. A further study by Alison Wellington (1991) re-estimated the relationship with Solon’s specification up to 1986 and found that the minimum wage elasticity had decreased in size and statistical significance. Card and Krueger (1995) again using the same specification, find that at conventional levels the significance of the effect disappears when the data are extended to include the period up to 1993, which covers the upratings studied in their cross-section work.

In order to evaluate these early models, we examine the model's ability to account for the evolution of teenage employment over the period 1980 to 1993. This period immediately follows the period treated by BGK and Solon. In line with the above-mentioned studies, we use quarterly data for the period 1954-93 to re-examine this earlier work and Figures 1(a) to1(d) show how the teenage employment rate, nominal minimum wage, real minimum wage and real average wage, respectively, evolve over the period.

We use the standard baseline specification (see BGK, 1982) for the teenage employment-population ratio (EP):

EP = g(MWK, X, Cycle, Time Trend, Time Trend Squared, Seasonal dummies) (1)

where X contains supply side "control" variables (proportion in the armed forces, proportion of teenagers aged 16 and 17, proportion of teenagers in the total working population) and the cycle is represented by the unemployment rate of males aged 25-54. The relationship is assumed to be linear with all variables expressed in logarithms (although Card and Krueger (1995) and Wellington (1991) for some unexplained reason do not transform the latter two variables). The minimum wage variable MWK is the Kaitz index - the minimum wage relative to the average wage weighted by the coverage of the minimum wage. The data set is that used in Card and Krueger (1995) where the data sources are described in detail[1].

A simple static model estimated by ordinary least squares is recognised by several authors as inappropriate, as there is clear evidence of residual autocorrelation. When estimated by maximum likelihood allowing for first order autocorrelation in the error term (which is almost universal practice in studies since 1983 (Card and Krueger, 1995; Bernstein and Schmitt, 1998)), the effect of the minimum wages is not significant at a 5% critical value (Table 1, column 1), thus confirming the finding of BGK (1983) mentioned above. In fact as Gary Solon (1985) pointed out the correlogram for the AR(1) model shows that the residual correlation is mainly of seasonal nature and thus the issue of how seasonality is modelled arises. Solon suggested adding a time trend and time trend squared interacted with seasonal dummies while maintaining the first order "correction". This addition restores the statistical significance of the minimum wage effect (Table 1, second column) and the Box-Pierce test suggests that these two transformations of the basic equation are sufficient to eliminate residual autocorrelation. The transformation proposed by Solon demonstrates that the seasonal and trend components of teenage employment are inadequately modelled by a set of determininstic seasonal dummies and a deterministic time trend and its square. As Koopman et al (1995) point out, the misspecification of seasonal and trend components often shows up as residual autocorrelation, and this is exactly what happened in the models considered by BGK. An AR(1) correction itself will not eliminate this type of autocorrelation.

A key test is how well the different models explain what happened in the 1980s - when the minimum was frozen in nominal terms - and in the early 1990s - when it was increased in sizeable steps. As Figure 1(c) shows, the real value of the minimum wage fell by around 20% between 1982 and 1990, only to rise sharply by about 10% in the following two years. The decline in the value of the minimum relative to average earnings (and therefore in the Kaitz index) should lead to an increase in the teenage employment-population ratio as the negative effect of increases in the minimum wage goes into reverse. Basically, the models upon which the consensus was based predict an increase over and above the cyclical and trend paths in teenage employment during the 1980s and a decrease relative to trend thereafter.

Figures 2(a) and (b) present the in-sample and out-of-sample prediction errors for the period up to the fourth quarter 1993 of the baseline model estimated with an AR(1) correction and with Solon's seasonality and AR(1) correction, respectively. Both approaches fail signally to account for the changes in the observed teenage employment rate. The AR(1) model estimates of teenage employment simply diverge from the observed value and substantially over-estimate it. The prediction errors from model with Solon's seasonality correction oscillate explosively after 1980 and there is a trend of over-estimation.

The reasons for the inability of the models to account for movements in the dependent variable over the period can be easily diagnosed. In particular the treatment of dynamics, the trend component and seasonality appears to be important. The former in terms of the inapproriateness of the autocorrelation correction and the latter in terms of the evolving nature of the seasonality. An AR(1) correction - retained by Solon - is itself very questionable, being obtained by imposing nonlinear restrictions on a general autoregressive distributed lag model. In Solon's model, the five common factor restrictions implied by an AR(1) error model against a general autoregressive distributed lag model are rejected by a likelihood ratio test at a 1% significance level (the test statistic is 19.8 against a critical value of 15.1). This would suggest that the residual autocorrelation is symptomatic of dynamic misspecification and that an autocorrelation "correction" is not appropriate (see Mizon, 1995). As the AR(1) correction is rejected against a more general autoregressive, distributed lag representation, this suggests that a more flexible lag structure is appropriate. When the AR(1) model with Solon's seasonality correction is estimated over the whole period up to 1993:4, the residual correlogram suggests that there is 4th and 8th order residual autocorrelation. The seasonality correction therefore appears to be valid for the sample period up to the end of 1979 but is no longer adequate in the 1980s and 1990s.

Another problem concerns the use of the Kaitz index of the minimum wage. As pointed out by Card and Krueger (1995) the fact that the minimum wage is expressed as a ratio of average earnings implies that simultaneously doubling the minimum wage and doubling average earnings would leave the teenage employment-population ratio unchanged. When expressed in logarithms, only if their coefficients are of equal and opposite sign, would the ratio form be appropriate. Furthermore, the minimum wage itself is an administered, discrete variable and moves up in nominal increments (see Figure 1b). When, divided by prices, the minimum wage (MW) in real terms becomes a continuous variable. It takes on an economic significance once it is recognised that there is a high correlation between teenagers' real earnings and the real value of the minimum wage. Real average earnings (AW) then represents the cost of substitute, adult workers. A possible alternative is that there may have been a regime change during the 1980s which altered agents' behaviour. Two of the principal explanatory variables, the minimum wage and (real) average earnings evolve in a completely different way compared to the period prior to 1980. In the next section we explore possible alternative specifications of the basic equation.

II An alternative approach

One way of avoiding the pitfalls faced by the traditional approach is to adopt a more flexible approach to the specification of various components of the model. Driving a highly inflexible, deterministic model through a sample covering a relatively long period can be too restrictive. For example, Harvey et al (1986) introduce a stochastic trend into an employment function for British manufacturing industry, and found that it represents the effect of underlying productivity growth in a far more satisfactory manner than the standard deterministic time trend. More recently, Harvey and Scott (1994) show that the inclusion of a stochastic seasonal component in a consumption function for the United Kingdom gives rise to a stable overall relationship between consumers'expenditure and disposable income whereas models without this stochastic component break down in the 1980s. In the US context, Krane and Wascher (1999) use this type of approach to examine cyclical-seasonal interactions in US payroll employment data. We have also pointed to the inappropriateness of assuming an AR(1) error process to model what is essentially a dynamic relationship.

In the current context, the models hitherto used in the analysis of the impact of the minimum wage in the United States of the type presented in equation (1), include seasonal, trend and cyclical components. The former two enter in the equation in deterministic fashion as dummy variables and time trends - and in Solon's case interactions of the two. The cyclical component is usually represented by a proxy variable - the unemployment rate of prime age males. Each of these essentially unobservable components can be represented in a far more flexible manner in what have come to be called 'structural time series models' (see for example, Harvey (1989, 1997)). A series y_t can be represented as a normal regression model - that is with explanatory variables represented by the vector x_t - with stochastic unobservable components as follows:

$$y_t = \alpha' x_t + \mu_t + \gamma_t + \psi_t + \varepsilon_t \qquad t = 1,2, ..., T$$

where μ_t is the trend component, γ_t the seasonal, ψ_t the cyclical component, and ε_t the white noise, irregular component.

The unobserved components also have an economic interpretation in the current context. The trend could be interpreted as incorporating technical progress, underlying trend productivity growth, social changes such as an increasing tendency to remain in education, demographic changes and generally any influential, unobserved factor that evolves in a fairly continuous manner. Its stochastic nature means that shifts can be taken into account, without altering the parameters of interest - the coefficients associated with the explanatory variables (α). The cyclical component has an obvious interpretation, given the apparent cyclicality of teenage employment in Figure 1(a). Finally the seasonal component reflects the fact that over a year employment in general will vary for demand-side reasons, and teenage employment in particular will vary due for example to breaks in the school year. Its stochastic nature allows the impact of seasonal factors to change over the sample period which runs from 1954 to the 1990s. As Solon (1985) has already pointed out: "The seasonality of the OLS residuals suggests that the constant seasonal model is indeed overly rigid" (p.294).

II.1 The structure of the unobservable components

The stochastic specification of the trend component is represented by the following equations:

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t \qquad \eta_t \sim NID\ (0, \sigma^2_\eta)$$

$$\beta_t = \beta_{t-1} + \zeta_t \qquad \zeta_t \sim NID(0,\sigma^2_\zeta)$$

where β_t is the slope of the trend μ_t, and η_t and ζ_t are error terms each of which is normally, and identically distributed (NID) with mean zero and variance σ^2_η and σ^2_ζ, respectively. If σ_η is different from zero, then the level of trend term is stochastic. If σ_ζ is non-zero, then the slope of the trend term is stochastic. It is assumed that the three error terms ζ_t, η_t and ε_t are uncorrelated.

The cyclical component is given by:

$$\begin{pmatrix} \psi_t \\ \psi_t * \end{pmatrix} = \rho \begin{pmatrix} \cos\lambda & \sin\lambda \\ \sin\lambda & \cos\lambda \end{pmatrix} \begin{pmatrix} \psi_{t-1} \\ \psi_{t-1} * \end{pmatrix} + \begin{pmatrix} \kappa_t \\ \kappa_t * \end{pmatrix}$$

where ρ is the scale factor such that $0 < \rho < 1$and λ is the frequency in radians such that $0 < \lambda < \pi$. κ_t and κ^*_t are NID and uncorrrelated with zero means and constant variance σ^2_κ.

Finally, the seasonal component can be defined in a similar, trigonometric manner. With quarterly data the seasonal frequency is equal to 4. There are thus four values for the index s (s=4) and for the trigonometric formulation we define j=s/2 when s is even, and j=(s-1)/2, when s is odd, and include the seasonal components γ_{jt} which are defined as follows:

$$\begin{pmatrix} \gamma_{jt} \\ \gamma^*_{jt} \end{pmatrix} = \begin{pmatrix} \cos\lambda_j & \sin\lambda_j \\ -\sin\lambda_j & \cos\lambda_j \end{pmatrix} \begin{pmatrix} \gamma_{j,t-1} \\ \gamma^*_{j,t-1} \end{pmatrix} + \begin{pmatrix} \omega_{jt} \\ \omega^*_{jt} \end{pmatrix}$$

for j=1,2, ...s/2, where $\lambda_j = 2\pi j/s$ is the frequency in radians and ω_t and ω^*_t are seasonal error terms and are NID with zero means and constant variance σ^2_ω. The seasonal components are stochastic when the latter is non-zero.

The specification of the trend, cycle and seasonal components in this way avoids the arbitrary use of proxies, introduces greater flexibility in the measure of what are basically unobservable influences, and contains the deterministic specification (constant trend, deterministic cycle and fixed seasonal effects) as a special case when $\sigma_\omega = \sigma_\zeta = \sigma_\kappa = 0$.

II.2 The explanatory variables

The baseline specification used in previous work includes the Kaitz index, the unemployment rate of prime age males and various demographic variables as controls. In the specification adopted here, we follow the reasoning presented in Card and Krueger (1995) referred to above and enter the real minimum wage and real average earnings as separate variables. We test and reject the restriction that the two variables should be entered as a ratio. The specification of the cyclical component above obviates the need for a proxy variable such as the unemployment rate of prime age males although we also present estimates using the latter as representing the cyclical component. The stochastic trend will pick up any slowly changing demographic factors. The dependent variable is defined in terms of the overall teenage population, which already controls for the key demographic factor affecting teenage employment. Finally, in view of the universal presence of autocorrelation in previous models and the rejection of an AR(1) "correction" as a solution, we include a lagged dependent variable which will reflect any sluggishness in the adjustment of teenage employment to its desired level[2]. As quarterly data are used, it is unlikely that firms' desired adjustment to a change in average real wages or changes in the real value of the minimum wage will be instantaneous. If this specification is in any way inadequate as a dynamic representation of the determinants of teenage employment rates, we would expect the various tests of misspecification to indicate this (such as the presence of residual autocorrelation and parameter instability).

The model we retain is therefore of the form:

$$EP_t = \alpha_1 MW_t + \alpha_2 AW_t + \alpha_3 EP_{t-1} + \mu_t + \gamma_t + \psi_t + \varepsilon_t$$

All observable variables are in logarithms. Estimation of the parameters of a model with unobservable effects is more complicated than traditional least squares methods. For estimation purposes, the model is written in the state space form, and estimates of the various parameters are obtained from a smoothing algorithm using maximum likelihood methods and the Kalman filter (see Harvey, 1989). Estimates were obtained using the STAMP 5 software package (Koopman et al, 1995).

III Results

III.1 Comparisons with previous studies

As a first step, we estimate the model over the period 1954:1 to 1979:4, in order to compare the results of those obtained earlier (and which figure in the work of BGK) and to see if it provides more accurate predictions than previous models. The importance of allowing the various components of teenage employment to contain a stochastic element can be seen immediately from the results in the first column of Table 2. An estimated standard deviation (and therefore variance) of zero indicates that a component contains no stochastic element, and is therefore deterministic. All of the components have stochastic elements except the slope of the trend and the cyclical component. The cycle has an estimated periodicity of 23 quarters or 5 ¾ years and is deterministic. In Figure 3(a) the stochastic trend component can be seen to 'bottom out' after 1965. The stochastic nature of the seasonal component confirms the finding of evolving seasonality by Solon. Figure 3(b) reveals that seasonal variations in the 1960s are far more pronounced than in the 1970s. The various diagnostic tests suggest that there is no evidence of misspecification: the hypotheses of the absence of autocorrelation and heteroscedasticity, and the normality of the errors are not rejected.

The parameters of economic interest are those of the explanatory variables which are obtained as 'final state' parameters, and these are presented in the lower half of Table 2. All three explanatory variables have significant coefficients, and the signs are consistent with the basic neoclassical model. If the minimum wage increases by 1% in real terms, the teenage employment rate will fall by 0.1% in the same quarter other things being equal. In the long run, after firms fully adjust to the rise, a 1% increase in the real value of the minimum wage will reduce the teenage employment-population ratio by about 0.2% for a given level of real average earnings. If the latter are growing in the long term, the size of the effect will obviously be lower, since a 1% increase in real average earnings will increase teenage employment by 0.54% in the same quarter and by 1.1% in the long run. This asymmetry justifies the inclusion of the minimum wage and average earnings as separate variables and not as a ratio. The restriction that the parameters are of equal size but opposite sign is rejected at a 5% significance level (t = 2.02, p value =0.042)[3]. The adjustment parameter of 0.514 suggests that half of the adjustment occurs in the same quarter as the minimum wage increase, a further quarter in the second quarter and by the fourth quarter nearly all the adjustment will have taken place. All of these estimates are evaluated relative to what would have been the case without a minimum wage or average earnings shock[4].

III.2 Predictive performance and stability

In order to check that the alternative model satisfies the same criteria used to judge traditional models considered above, the forecasting ability and stability of the model is next investigated. Figures 4(a) and 4(b) provide forecast errors and a CUSUM plot, respectively, obtained by assuming that the values of the seasonal and trend components obtained in the final state remain constant over the period 1980:1 to 1993:4. The forecasts of the teenage employment rate are satisfactory: they are never more than two standard errors from the observed value, and rarely more than 3% out. The CUSUM plot in Figure 4(b) and the tests of parameter stability and forecast accuracy presented at the bottom of Table 2 confirm the satisfactory forecast performance of this model. The model using the unemployment rate as a proxy for the cyclical component performs less well (see Appendix, Figures A.1 and A.2).

Overall, this model with stochastic trend and seasonal components represents an improvement on previous models. It also identifies a significant negative effect of the minimum wage on teenage employment. However, the minimum wage was increased again in 1980 and 1981 and thereafter frozen, and its real value reduced by some 20% over the period 1982-90. Furthermore real average earnings stopped increasing in the mid 1980s and actually declined thereafter (see Figure 1(d)). Wellington (1991), using the baseline model with the AR(1) and seasonality corrections suggested by Solon with data up to 1986, found that the minimum wage elasticity was smaller in (absolute) magnitude and generally insignificant at conventional levels. As already pointed out above, Card and Krueger found an even smaller effect.

Before estimating the model over the whole sample (up to 1993), we first estimate it using data up to 1989, the last year for which the nominal minimum wage was frozen. The only notable difference from earlier results is that the slope of the stochastic trend term also becomes stochastic. This has no impact on the estimated values of the other parameters including the minimum wage elasticity although the average real wage elasticity is slightly higher.

The final estimated model is for the whole sample period considered by Card and Krueger, 1954:1 to 1993:4. The period after 1989 incorporates two substantial increases in the minimum wage (to $3.80 on April 1st 1990 and $4.25 on April 1st 1991) at a time when real average earnings were falling. The addition of these last four years of data alters quite noticeably the stochastic nature of the model. Along with the trend, slope and seasonal components, the cyclical component also becomes stochastic. The seasonal components remain fairly constant over the period 1980-93 (see Figure 5). The model is not misspecified according to the diagnostic test statistics, and the variances of the various disturbances associated with the irregular term, the seasonal component and the slope of the trend are not very different from the values obtained for the period up to 1989. The variance of the disturbance of the trend component somewhat smaller, but the most noticeable changes occur in the values of the minimum wage elasticities. This is firstly due to a larger impact effect - the short run elasticity increases slightly in absolute value to 0.12 - and secondly due to a larger coefficient on the lagged dependent variable suggesting more rapid adjustment to the desired level of youth employment. This combines with the short run elasticity to give a long run minimum wage elasticity of -0.299 which is 45% larger in absolute value than that obtained using data up to 1979. This is the opposite of numerically smaller effect found by Card and Krueger using the Solon's approach. Furthermore, the effect is highly significant.

III.3 The impact of the 1996 and 1997 increases in the federal minimum

After the two increases implemented in the early 1990s, the federal minimum wage was again left unchanged for several years. Two further increases were implemented on October 1st 1996 from $4.25 to $4.75 and September 1st 1997 to $5.15, and these enable our model to be tested out-of-sample (these data were only recently made available to us[5] and were not used in a previous version of this paper (Bazen and Marimoutou, 1999)). The forecast error is presented in Figure 6 along with the CUSUM plot for the 22 quarters that make up the period 1994:1 to 1999:2. The forecast error is always inside a two standard errors of zero and the CUSUM plot remains close to the horizontal axis. Estimation for the whole period confirms the overall findings of a negative impact of the minimum wage on teenage employment. The short and long run elasticities are –0.117 and –0.340 respectively (Table 2, column 4). The latter is slightly higher than the value found for the period up to 1993.

III.4 Interpretation of the findings

The clear message that emerges from these results is that the time series data do suggest that federal minimum wage increases have a statistically significant, negative effect on teenage employment. The results confirm the findings of pre-1980 studies and show why subsequent studies using time series data found no significant effect. The question therefore arises of how can our results be squared with the findings of Card and Krueger (1995). In various reviews of their cross section-based work - which forms the main basis for claims that there was no negative employment effect associated with increases in the minimum wage - there was concern expressed that their data did not enable the long run effect to be identified and that this effect would be negative and significant. For example, Charles Brown (1995) maintained: "I would still expect employment effects of a minimum wage increase to be more negative in the long run than in the short run, but this is a matter that demands further investigation." (p.829). Our results support this prediction, the long run effect being more than twice the size of the short run effect.

In this study, we have successfully modelled the purely seasonal variation in teenage employment simply by adopting a more flexible approach which has been found to work well in other areas such as modelling consumption patterns (Harvey and Scott, 1994). The interpretation of the stochastic nature of the trend is less clear. Brown (1995) suggests that "a future study will show that teenage employment responds to the declining real minimum wage in the 1980s if one takes account of technologically driven declines in demand for low-wage workers over the period. " (p.829). A stochastic trend term in an employment-output relation would indeed pick up this kind of change in the industrial sector, but here we are examining the teenage labour market. Figure 5 shows the estimated trend component for the whole sample period and indicates that through the 1980s there was a certain stagnation in the factors that make up the trend or that positive elements outweigh negative elements.

The reconciliation of the time series and cross section evidence is difficult in this context because the two approaches identify different aspects of the minimum wage-employment relationship. When time series data are used, a rise in the nominal minimum wage will cause a rise in the real value of the minimum wage, and this initially reduces the teenage employment-population ratio. However, unless the nominal minimum wage is continually uprated or unless prices remain constant, this real increase will be subsequently eroded. As the federal minimum wage is not indexed to prices, and as there is no automatic mechanism that gives rise to regular upratings as in France[6], the fact that price inflation is non-zero means that in a time series context, the two-way variation of the teenage employment rate is linked to two-way variations in the real value of the minimum wage. In contrast, the studies of Card and Krueger and others examine changes in employment just after minimum wage hikes (ie one-way variations).

IV Conclusions

These results contrast with those obtained by updating the standard model undertaken by Card and Krueger in that the effect of the minimum wage is found to be negative and significantly different from zero. The basic model is built on similar foundations to that used in previous studies with the key difference that the cyclical, trend and seasonal influences are treated as unobservable stochastic components rather than deterministic components. This increased flexibility enables the effect of changes in the real value of the minimum wage to be more successfully isolated. Furthermore, unlike previous models, the model remains stable during a period in which the nominal value of the minimum wage was frozen, and subsequently increased quite substantially.

The effect of a rise in the minimum wage of 10% in real terms with real average wages constant is to reduce teenage employment by 2 to 3%. As the real value of the minimum wage fell over the 1980s, teenage employment increased relative to what it otherwise would have been. The increases implemented at the beginning of the 1990s will have reduced teenage employment - a conclusion contrary to that obtained by Card and Krueger in a number of cross section contexts. The size of the estimated effect however is small and is not out of line with estimates found for the period prior to 1980. Unlike previous estimates though the effect remains statistically significant and relatively stable over the 1980s and appears to be slightly higher in the early 1990s.

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Table 1 – Baseline model estimates for the period 1954-1979		
	AR(1) model	AR(1) with Solon's seasonality correction
Kaitz index	-0.0897 (0.046)	-0.0957* (0.041)
Unemployment rate	-0.1156** (0.019)	-0.1079** (0.0179)
Share of 16-17 in teenage population	-0.6733** (0.212)	-0.6252** (0.210)
Share of teenagers in working age population	-0.4431 (0.275)	-0.5904 (0.323)
Proportion of teenagers in armed forces	0.0179 (0.0429)	0.00369 (0.0392)
'rho'	0.5528 (0.090)	0.7318 (0.0766)
* significant at 5% ** significant at 1% 'rho' is the estimated first order autocorrelation coefficient		

rho - first order autocorrelation coefficient
* significant at 5%. Estimated standard errors in parentheses.

Table 2 Estimated parameter values and stability tests for the coefficients of the explanatory variables				
	1954:1 to 1979:4	1954:1 to 1989:4	1954:1 to 1993:4	1954:1 to 1999:2
Lagged dependent variable	0.514 (0.081)	0.547 (0.068)	0.605 (0.059)	0.656 (0.049)
Real Minimum Wage	-0.101 (0.051)	-0.0983 (0.045)	-0.118 (0.039)	-0.117 (0.033)
Real Average Wage	0.540 (0.243)	0.615 (0.189)	0.611 (0.166)	0.779 (0.143)
Long run minimum wage elasticity	-0.208	-0.217	-0.299	-0.340
Hyperparameters				
Irregular	0.0108462	0.0102673	0.0110284	0.011698
Level	0.0117246	0.0103783	0.0082175	0.006998
Slope	0	0.000138216	0.0001328	0.000220
Cycle	0	0	0.002146	0
Seasonal	0.00297503	0.00271531	0.002719	0.002844
Misspecification tests				
Box-Ljung Q statistic (a)	5.063	7.079	8.101	9.480
Heteroscedasticity (d degrees of freedom) (b)	0.59425(32)	0.4512(46)	0.384 (51)	0.3866 (58)
Bowman-Shenton normality test (c)	0.90802	0.22468	0.194	0.6604
Prediction error variance	0.00052054	0.00044966	0.0004394	0.000415
Parameter stability (Chow test) (d)	0.465 [F(56,98)= 1.46]	0.495 [F(16,138)= 1.71]	0.807 [F(22,154)=	N/A
Forecast test (d)	28.09 [χ^2(56)=73]	11.06 [χ^2(16)=26.3]	18.05 [χ^2(22)=26.3]	N/A

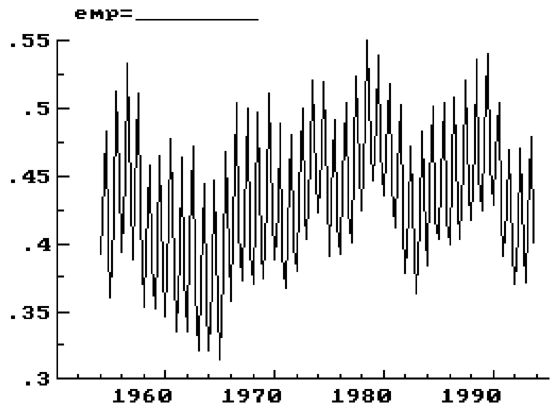
Estimated standard errors in parentheses.

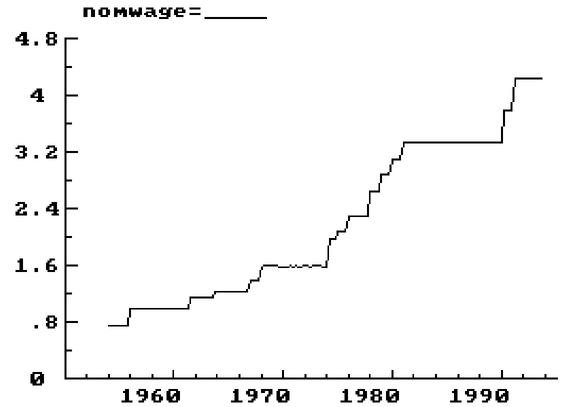
Notes to Table 2:

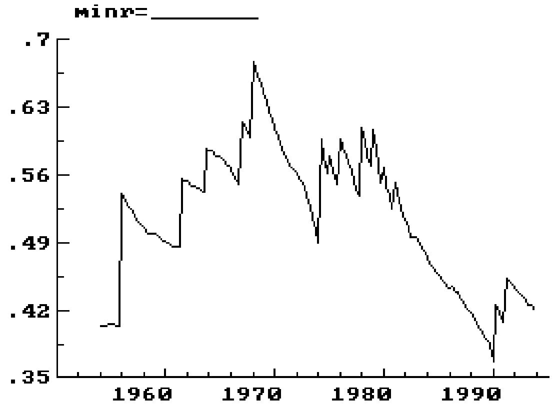
(a) Test for 6th order auto-correlation distributed as χ^2_6 under the null hypothesis of no autocorrelation.
(b) Distributed as χ^2_d under the null hypothesis of homoscedasticity.
(c) Distributed as χ^2_2 under the null hypothesis of normality.
(d) These tests are for the period up to 1993:4 except for the results using data up to 1993:4 where the period 1994:1-1999:2 is used.

Figure 1(a) Teenage employment population ratio

Figure 1(b) Nominal minimum wage







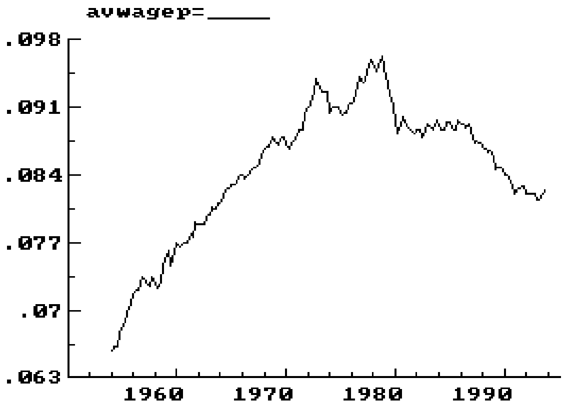
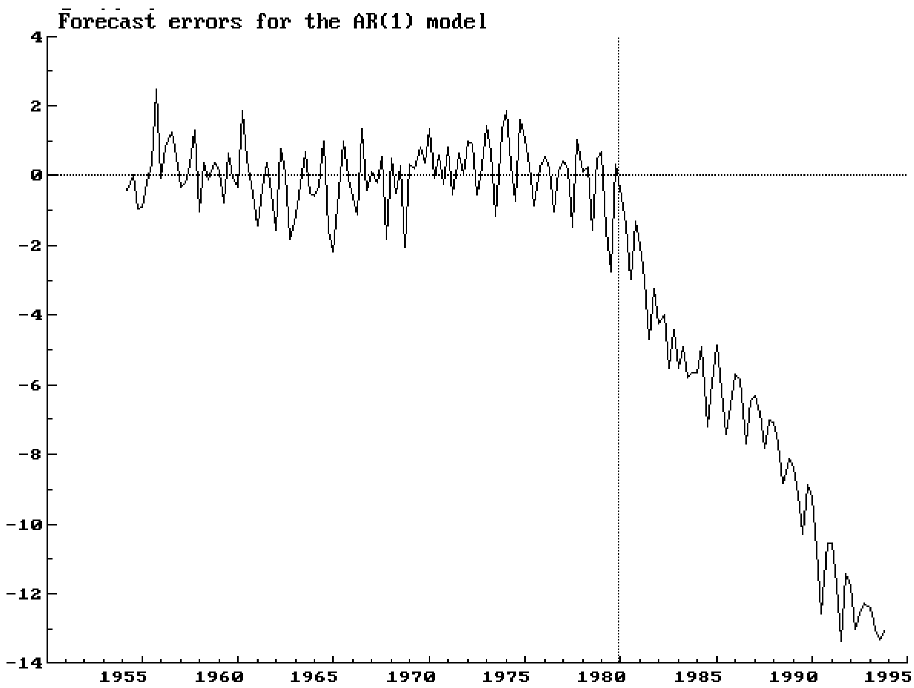


Figure 1(c) Real value of minimum wage

Figure 1(d) Real average earnings

Figure 2 Prediction errors from the baseline and Solon model



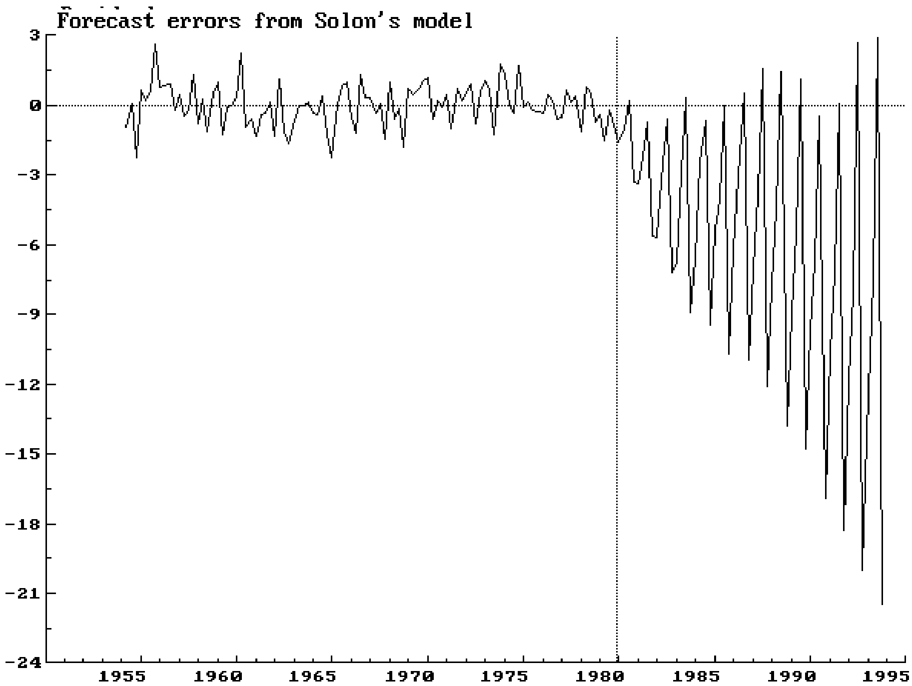


Figure 3 - Trend and seasonal components for model estimated over 1954-79

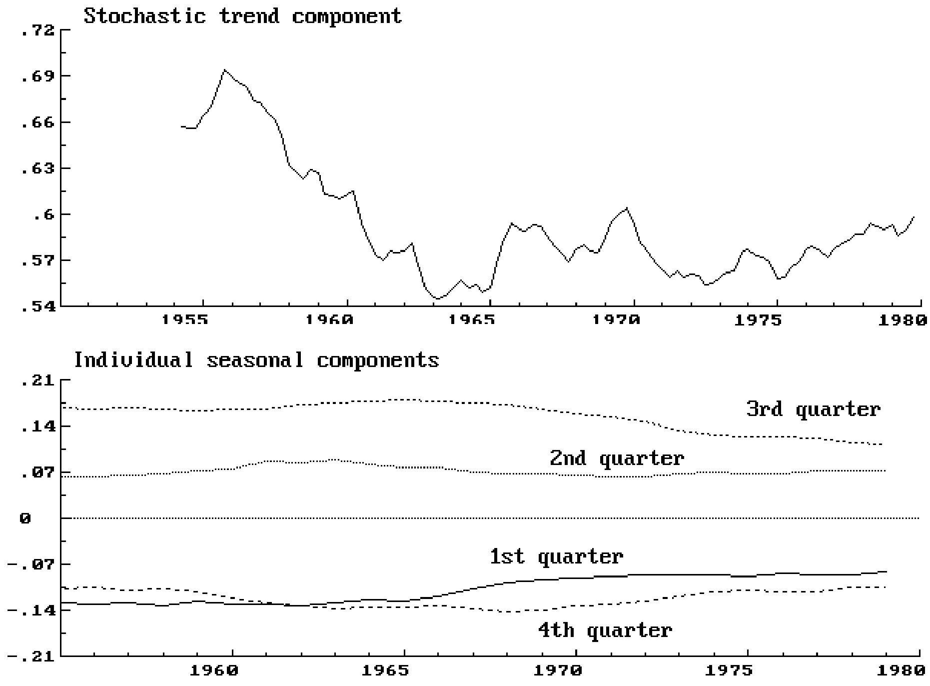


Figure 4 Prediction quality for the period 1980-93 from the model estimated over 1954-79

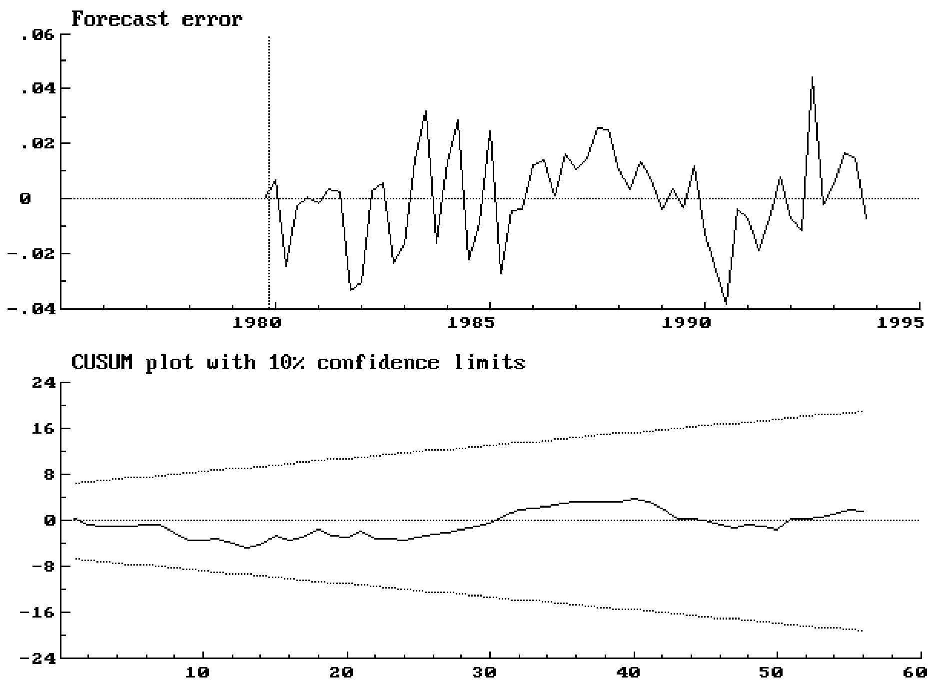


Figure 5 - Trend and seasonal components for model estimated over 1954-79

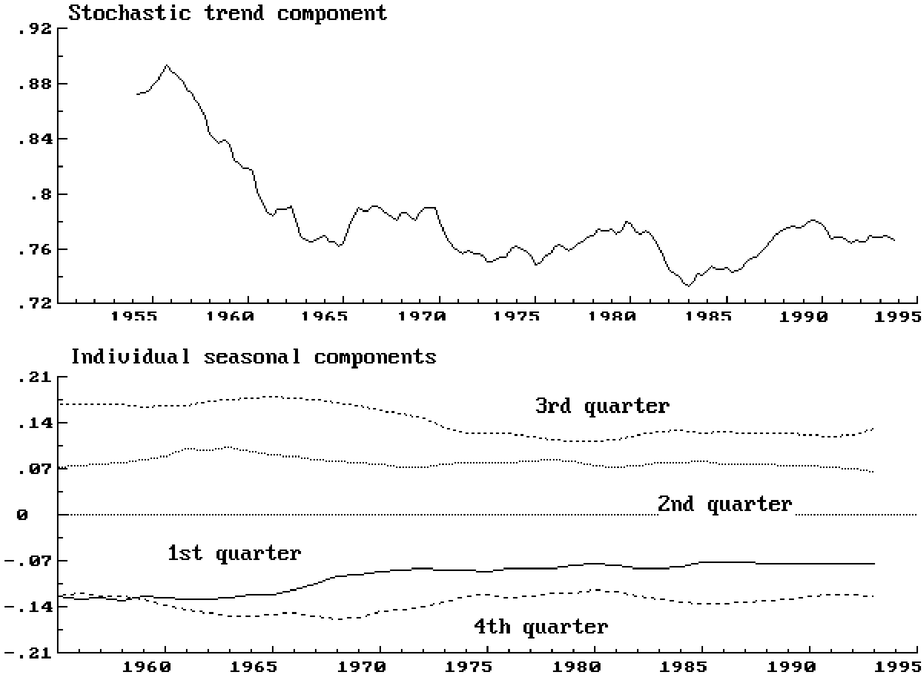
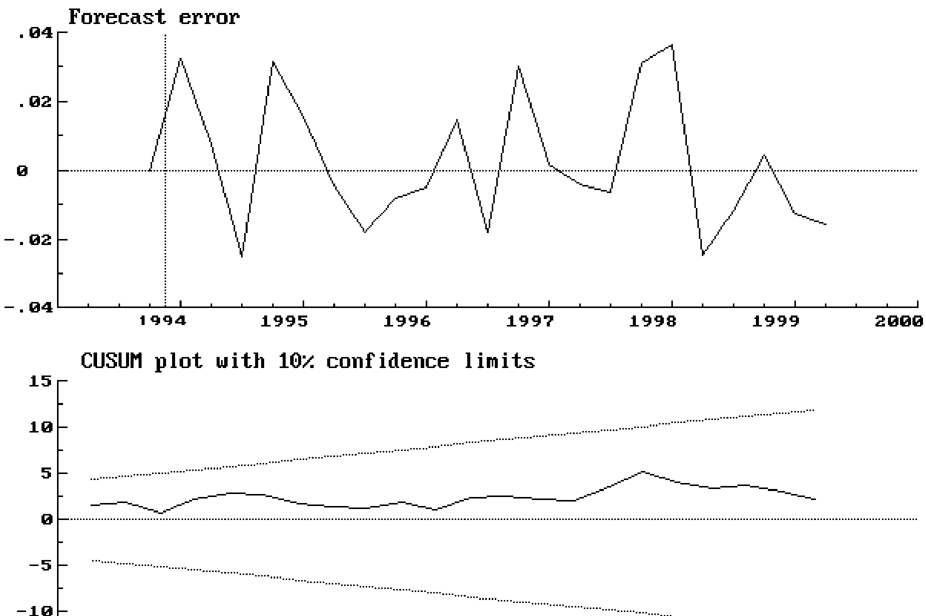


Figure 4 Prediction quality for the period 1994-99 from the model estimated over 1954-93



[1] We are very grateful to David Card for providing these data.

[2] We experimented with several lag specifications but found that only a single lagged dependent variable was appropriate.

[3] The restriction is even more decisively rejected as the sample size increases to incorporate the 1980s and the 1990s. For 1954-89, t = 2.31 (p value = 0.022) and for 1954-93, t = 3.13 (p value =0.002).

[4] When the unemployment rate is used to proxy the cyclical component, the results are slightly different (see Appendix, Table A.1.). The minimum wage coefficient is not significant at conventional levels (p value = 0.0599) and the real wage elasticity is smaller and statistically insignificant.

[5] We are very grateful to Jared Bernstein for providing the additional data.

[6] See Bazen and Martin (1991) for a description