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MINIMUM WAGES AND THE DEMAND FOR LABOR

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ABSTRACT

I formulate measures of the effective minimum wage, based on broad definitions of the labor costs that face employers, and use these measures in reestimating some simple equations relating the relative employment of youths and adults to the U.S. minimum wage using aggregate data for 1954-78. I then ground the model more closely in the theory of factor demand, first by adding the relative wages of youths and adults to the equation describing their relative employment, and then by specifying a complete system of demand equations for these two types of labor. Teen employment responds quite robustly to changes in the effective minimum in these specifications, with an elasticity of -0.1. A translog cost function defined over young workers, adults, and capital shows that the effective minimum wage reduces employers' ability to substitute other factors for young workers. Using both sets of results, I find that a subminimum wage for youths would have increased their employment with at most a small loss of jobs among adults.

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I. INTRODUCTION

The vast amount of research on the effects of the minimum wage on employment requires that we justify the presentation of additional results. Economists have examined the issue: (1) Recognizing the importance of incomplete and changing coverage of the minimum (Kaitz, 1970; Goldfarb, 1974); (2) Considering how to define the real effective minimum wage in a manner consistent with economic theory (Welch, 1974); and (3) Acknowledging the role played by spillovers and turnover in completely and incompletely covered markets (Hashimoto-Mincer, 1971; Welch, 1974). Yet, as Siskind (1977) has shown, the findings about the magnitude of the effect of higher wage minima on employment seem sensitive to minor changes in the data and specification. A more careful specification of the underlying theoretical model and more attention to the data used could have a substantial payoff in terms of the confidence one can place in the estimates produced. Thus, though a great deal of research precedes this study, that research is not as thorough as current theory and techniques enable the analysis to be.

In the next section we describe some new measures of the effective minimum wage and examine how they affect estimates of *ad hoc* models of the minimum wage's impact on youth labor markets, one of which introduces factor prices directly into the estimating equations. In Section III we estimate a complete system of demand equations for teen and adult labor that incorporates the effective minimum wage within a framework based entirely on the theory of factor demand. In Section IV we examine how changes in the effective minimum wage change the structure of firms' costs, using a translog approximation to a three-factor cost function involving youths, adults, and capital. In Section V we show how to calculate the net effects of higher minima and use the method to simulate the impact of the minimum wage on youth and adult employment and on factor shares. As does all previous empirical work on this subject, our results present only the net employment effects; they cannot show the larger gross effects as workers are displaced from some firms and find employment elsewhere.

II. MEASURES OF EFFECTIVE MINIMUM WAGES AND THEIR EFFECTS ON EMPLOYMENT

The most thorough previous work on the employment effects of the minimum wage is based on weighted averages of legislated minimum wages relative to average hourly earnings in particular sectors, with the weights dependent on coverage of the minimum and on youth employment in each sector. Earnings per hour paid for, though, are not a good measure of the cost of employment. Required payments such as payroll taxes for social insurance, and negotiated and unilateral payments such as bonuses, are a cost of employment. Increases in paid holidays and vacations have imposed a growing wedge between hours paid for and hours worked. Changes in the user cost of training will also cause the correlation between broader measures of labor costs and average hourly earnings to differ.

These are not minor distinctions. Real average hourly earnings (AHE) in manufacturing rose 55 percent between 1953:I and 1978:IV. In the same period real compensation per hour actually worked (COSTWK) rose 93 percent, while the sum of this measure and real user costs of training (ECNT) rose 92 percent. Comparable figures for the private business sector are 67 percent, 94 percent and 98 percent. Clearly, there is room for substantial differences in the estimated effects of higher legislated minimum wages on employment depending upon the labor cost measure to which the minimum is compared.[1/]

For each of the eight private nonfarm one-digit industries we form:

$$MIN1_i = (B_i COVB_i + N_i COVN_i)/AHE_i = MIN_i/AHE_i \quad , \tag{1}$$

where B and N are the minimum wage rates applying to previously and newly covered workers, and COVB and COVN are the corresponding fractions of workers covered in the i'th industry.[2/] Alternative measures of the effective minimum, $MIN2_i$ and $MIN3_i$, are defined by dividing (1) by the ratio of COSTWK to COMP, hourly compensation (wages, social insurance and pension payments), and the ratio ECNT to COMP respectively. Since our focus is on teenage employment as affected by changes in the minimum wage, we calculate effective minimum wage measures for the private nonfarm sector as weighted averages over the eight industries of $MINi_i$, $MIN2_i$ and $MIN3_i$ using each industry's share of teen employment as its weight.[3/]

In the top half of Table 1 we list the means and standard deviations of the minimum wage measures for the years 1954 - 1978 for the private nonfarm sector and for the three industries — services (except private household workers), retail trade and manufacturing — for which there are large samples of teen workers. The more inclusive nature of these labor cost measures ensures that the effective minimum wage variables based upon them have lower means than do those based on AHE. Also, it is worth noting that the effective minimum is highest in manufacturing among the three industries considered: This occurs, even though labor costs are higher in manufacturing, because the coverage rate has historically been far higher there than in services or retail trade. Finally, note that in manufacturing, though not in the other industries, the coefficient of variation of MIN3 (.095), based on ECNT, is much larger than that of MIN1 (.079), based on AHE.

These measures of the effective minimum wage compare the minimum price employers must pay for an hour of labor to the average cost of an hour of labor. Since most of the interest in the employment effects of the minimum wage is in the labor market for youths, they have serious problems insofar as they are not specific to that market. To circumvent these problems we replace AHE (COSTWK or ECNT) in (1) by RY*AHE (or RY times COSTWK or ECNT), where RY is the ratio of the weekly earnings of full-time workers age 16 to 24 to those of all full-time

Table 1

Means and Standard Deviations of
Effective Minimum Wage Variables, 1954:I - 1978:IV

	Services	Retail	Manufacturing	Private Nonfarm
Based on average labor cost:				
AHE	.243 (.113)	.237 (.166)	.440 (.035)	.302 (.089)
COSTWK	.210 (.095)	.204 (.141)	.388 (.034)	.263 (.073)
ECNT	.176 (.078)	.171 (.117)	.336 (.032)	.220 (.060)
Based on teen labor cost:				
AHE	.320 (.162)	.317 (.230)	.568 (.047)	.396 (.134)
COSTWK	.278 (.136)	.273 (.196)	.501 (.039)	.344 (.111)
ECNT	.232 (.112)	.228 (.163)	.434 (.037)	.288 (.091)

workers.[4/] This redefinition could have large effects: The ratio RY was at its highest value, .86, in the years 1954-1978 in 1955, and fell to .70 in 1977. In addition to its use in forming a better effective minimum wage measure, we use RY to define hourly labor cost measures.[5/]

The means and standard deviations of the redefined effective minimum wage variables are shown in the bottom half of Table 1. Because the teen wage is below the average, the effective minimum measures based upon it exceed those based on average labor costs. Also, because the coverage rate outside manufacturing (which affects the numerator of the effective minimum wage variable) was rising at the same time RY (which affects the denominator) was falling, the effective minimum measures for nonmanufacturing industries based on teen labor costs have much more variance than do those based upon average labor costs.

Our modifications of the effective minimum wage variables in (1) have been concerned chiefly with broadening the terms included in the denominator. The only adjustment of the numerator has been the inclusion of social insurance and pension payments (through COMP). If persons at the minimum wage receive nonwage benefits (reduced hours, specific training) at the same rate as does the average worker, the ratios MIN3 and MIN2 should equal MIN1. Whether this extreme assumption or our partial adjustment of the numerator is correct is unknowable *a priori*. However, by comparing fits of employment equations using the different measures, we can infer which assumption is superior.

Our initial approach is to estimate equations describing the behavior of teen relative to adult employment over time. The equations have the form:

$$ER_t = \alpha_0 + \alpha_1 MINJ_t + \alpha_2 U_t + \alpha_3 t + \alpha_4 DUMS_t + \nu_t \quad , \tag{2}$$

where ER is the logarithm of relative teen/adult employment; MINJ is the logarithm of one of the effective minimum wage variables we constructed; U is the logarithm of the adult unemployment rate; DUMS is a vector of three quarterly dummy variables, and ν is a disturbance term. This equation was developed by Welch (1974) and used by Siskind (1977). The employment data are monthly CPS data, seasonally unadjusted, averaged into quarterly observations. Teenagers are persons 14-19; adults are those 20 and over. Because of the limits on the availability of data on coverage of the minimum wage by industry, and because previous studies that used (2) started their samples in 1954, our estimates too begin with that year. Since it is likely that the disturbances are autocorrelated, (2) is estimated in each case using the Cochrane-Orcutt iterative technique.[6/]

The results of estimating (2) for the four samples (three industries and the aggregate of private nonfarm employment) for 1954 - 1978 are presented in Table 2.[7/] Examining first the coefficients of MINJ in the specification based on AHE, we find that, with the exception of services, the coefficients on the MINJ variables are all significantly negative at the 99 percent level of confidence. Comparing these to Siskind's (1977) estimates for 1954-68, we observe quite similar effects. The main difference is that our elasticities have greater statistical significance in the equations for the private nonfarm sector and for services, and less in that for manufacturing. Despite the addition of the extra years of data, the adjustment for serial correlation, and the use of corrected series on teen employment, the estimated minimum wage elasticities differ only slightly.[8/]

The discussion above is based on results using an effective minimum wage variable whose denominator is average hourly earnings. When equations (2) are reestimated using MIN2 and MIN3, we find uniformly that the explanatory power of the model increases. Moreover, the best fits for all four data sets are in the equations based on an effective minimum variable that includes

Table 2

Basic Equations for Relative Teen-Adult Employment, 1954:I - 1978:IV*

Industry and Labor Cost Series	Minimum Wage Elasticity	$\hat{\rho}$	$\hat{\sigma}_e$
Private Nonfarm:			
AHE	-.1107 (-2.23)	.677 (9.16)	.033587
COSTWK	-.1062 (-2.12)	.702 (9.80)	.033478
ECNT	-.1214 (-2.56)	.666 (8.88)	.033152
Services (Except private Household):			
AHE	-.0219 (-.43)	.459 (5.14)	.071622
COSTWK	-.0227 (-.44)	.460 (5.15)	.071618
ECNT	-.0302 (-.58)	.464 (5.21)	.071566
Retail Trade:			
AHE	-.0410 (-2.95)	.661 (8.76)	.037242
COSTWK	-.0411 (-2.96)	.660 (8.75)	.037234
ECNT	-.0418 (-3.04)	.657 (8.67)	.037151
Manufacturing:			
AHE	-.3786 (-2.86)	.611 (7.67)	.058076
COSTWK	-.3798 (-2.86)	.610 (7.65)	.058083
ECNT	-.3988 (-3.28)	.591 (7.30)	.057568

*t-statistics are in parentheses here and in Tables 3-8.

ECNT, the most complete of the three labor-cost measures, in the denominator. Not only are the fits better, the estimated minimum wage elasticities are higher as well, by about 10 percent in the private nonfarm sector as a whole and in manufacturing, by 50 percent (on a low base) in services, and by a tiny fraction in retail trade.

Equation (2) is a strange hybrid whose basis in theory is quite difficult to discern.[9/] There are three problems:

(1) It appears to be a relative demand equation, yet the relative price measure cannot be claimed to reflect the prices of the two types of employee. Implicitly the equation states that the price of adults (the denominator of the effective minimum) is average hourly earnings (or labor costs), while the coverage-weighted minimum (MIN) is the price of teenagers.

(2) If equation (2) is in part based on the theory of factor demand, it puts substantial restrictions upon the adjustment of the employment of youths and adults. Implicitly it states that employers are concerned only about the ratio of employment in these two groups, and that there are no separate disturbance terms that reflect random effects in the adjustment of employment in the two groups.

(3) If the equations are intended to reflect the demand for labor, they should include a scale effect, measured by the demand for output. From this viewpoint the trend can be seen as reflecting changes in factor productivity, but the unemployment rate is difficult to rationalize as a good measure of shifts in demand.

As a first step toward grounding (2) in the theory of factor demand, we add the relative prices of teen labor and adult labor, based on RY.

We also replace the MINJ variables with MINT (the variables whose means and standard deviations are listed in the bottom half of Table 1), whose denominators are based on those labor costs specific to teenagers. These two modifications force us to reinterpret the meaning of the minimum wage variable. Increases in that term produced by legislated increases in MIN imply the truncation of the distribution of the marginal productivity of teen labor. Essentially, the labor cost measures based on RY show the average prices of teen and adult labor, while the minimum wage variable shows how the distribution of productivity of teens is truncated from below by changes in the legislated minimum.[10/] In terms of Figure 1, the relative price variable is based upon an average of the wages of teens in the shaded area beyond MIN_o, while MIN, the numerator of MINT, reflects the truncation point. This suggests that the net effect of any increase in the minimum wage must be calculated very carefully. An increase in MIN from MIN_o to MIN_1 will affect both MINT and relative prices (because the truncation point of the distribution of teen wages is changed). The cross-hatched area in Figure 1 will drop out of the observed distribution of wages.

The revised version of (2) is:

$$ER_t = \beta_0 + \beta_1 MINT_t + \beta_2\ WR_t + \beta_3\ U_t + \beta_4\ t + \beta_5\ DUMS + \nu'_t, \qquad (2')$$

where WR is the log of the relative teen-adult wage or labor cost, and MINT is in logs. The coefficient β_1 can be interpreted as showing the effect of a higher effective minimum on relative employment if WR is unchanged. Conceptually it shows the extra impact of a higher minimum once that effect has been compensated for by adjusting WR to account for the increased average wage of teenagers produced when the truncation point in Figure 1 moves rightward. The compensating change

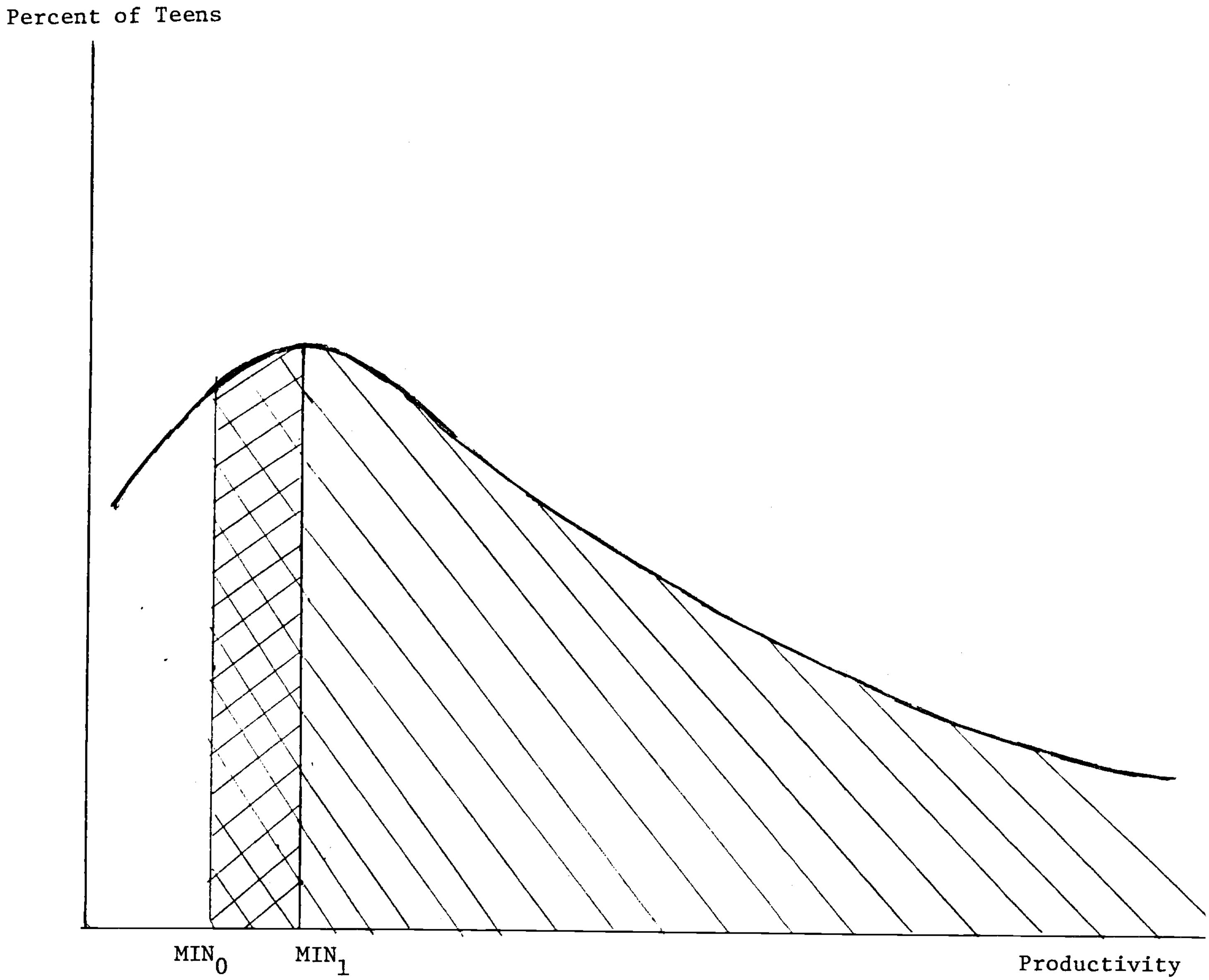

Figure 1: The Effective Minimum Wage and the Distribution of Teenagers by Productivity

to hold WR constant when MIN increases must occur through a drop in wages of high-wage teens sufficient to offset the effect of the truncation. That $\beta_1 < 0$ follows from the assumption, based on the studies surveyed in Hamermesh-Grant (1979), that the demand elasticity for low-wage workers exceeds that for high-wage workers; the net negative effect on teen employment of a higher MIN, holding WR constant, results from the partly offsetting positive and negative effects on high- and low-wage teens respectively. This suggests that the net effect of higher MIN must be calculated using β_1 and β_2 both (see Section V).

Table 3 presents estimates of (2') using labor cost and effective minimum wage variables based on AHE and ECNT.[11/] Despite the drastic decline in RY since the 1960s, and the different bases of MINT and MINJ, the addition of the relative labor cost variable has little impact on the estimated minimum wage elasticity, as a comparison of the estimates in Tables 2 and 3 clearly shows.[12/] Moreover, as in Table 2 the equations based on the most complete labor-cost measure, ECNT, produce slightly better fits and slightly higher minimum wage elasticities. Despite the stability of the minimum wage elasticity, the inclusion of a relative price measure is justified in terms of achieving a better fit to the data (except in manufacturing). The other t-statistics on relative labor costs exceed one, and all the estimated relative price elasticities are negative.

Since coverage and the legislative minimum are separate issues, we experimented with separate variables for each. The logs of the fraction of teen employment covered and the ratio of the minimum wage to AHE_i (or to $ECNT_i$) were both entered in (2'). (For the private nonfarm sector these measures were teen-employment weighted averages of the variables for each industry.) For retail trade and the private nonfarm sector only did $\hat{\sigma}_\varepsilon$ decrease. In all four cases the larger effects were through the relative minimum; their elasticities were -.21, -.14, -.21 and -.43 for the four equations using ECNT, and they were significantly negative except for services.[13/] Since coverage is now fairly

Table 3

Estimates of (2') for Relative Teen-Adult Employment, 1954: I - 1978: IV

Industry and Labor Cost Series	Minimum Wage Elasticity	Relative Labor Cost Elasticity	$\hat{\sigma}_e$
Private Nonfarm:			
AHE	-.1027 (-2.09)	-.4116 (-1.55)	.033505
ECNT	-.1131 (-2.40)	-.3995 (-1.52)	.033266
Services (Except private Household):			
AHE	-.0272 (-.46)	-1.94 (-3.05)	.069207
ECNT	-.0383 (-.64)	-1.96 (-3.07)	.069144
Retail Trade:			
AHE	-.0403 (-2.84)	-.4666 (1.43)	.037087
ECNT	-.0411 (-2.93	-.4601 (-1.42)	.037006
Manufacturing:			
AHE	-.4016 (-3.05)	-.5311 (-.91)	.058099
ECNT	-.4185 (-3.46)	-.5652 (-.99)	.057572

complete, any future increases in the effective minimum wage must come through higher legislated minima. These estimates suggest that the employment effects of such increases would be more severe than implied by estimates based on increases in the MINT variables (that combine coverage and the legislated minimum).

III. THE MINIMUM WAGE IN A COMPLETE SYSTEM OF DEMAND EQUATIONS FOR LABOR

In this section we generalize the model of Section II by transforming it into a complete system of demand equations for the two factors of production, teen and adult labor:

$$ET_t = a_1 + \alpha_1 WT_t + \beta_1 WA_t + \gamma Q_t + \delta MINT_t + \kappa_1 X_t + \varepsilon_1 \, , \tag{3a}$$

$$EA_t = a_2 + \alpha_2 WT_t + \beta_2 WA_t + \gamma Q_t + \kappa_2 X_t + \varepsilon_2 , \tag{3b}$$

where ET and EA are logarithms of employment of teenagers and adults respectively; WT and WA are logarithms of labor costs per hour; Q is the log of output; X is a vector including a time trend and quarterly dummy variables; and the ε are random disturbance terms.[14/] It implicitly assumes, as did Figure 1, that an increase in the effective minimum wage facing employers of teenagers directly affects only their employment. This equation system respecifies (2) further to account for the objections in Section II.

As it is written, system (3) imposes no restrictions on the effects of one wage rate on employment in the other group. This allows the testing of hypotheses stemming from factor demand theory. The theory implies the symmetry of cross-price effects, $\alpha_2 = R\beta_1$, where R is the ratio of factor shares; and it also requires that there be homogeneity in the responses of employment to changes in all prices, i.e., $\alpha_1 + \beta_1 = 0$ and $\alpha_2 + \beta_2 = 0$. (We assume here and in Section IV that the legislated minimum is included in the phrase "all prices.") The model in (3) is estimated using the data for the private nonfarm sector underlying the estimates in Tables 2 and 3. Separate first-order autoregressive processes are assumed for ε_1 and ε_2, and the parameters ρ describing these processes are estimated. Because the fits of (3) were always slightly better when ECNT was used, we present here estimates in which WT, WA and MINT are based upon that measure.[15]

The restrictions of homogeneity and symmetry cannot be rejected at the 99 percent level of significance ($\chi^2(3) = 10.72$), though they can at the 95 percent level. Since the restricted system is more consistent with economic theory, and the estimate of the coefficient on MINT from the unconstrained model differs only slightly from that from the model in which homogeneity and symmetry have been imposed, we present the restricted estimates in Table 4.[16] The equations were estimated by iterative least-squares, a procedure that is asymptotically equivalent to maximum likelihood.

As a result of the imposition of the constraints, there is only one independent coefficient on the labor cost terms, α_1. Though this coefficient is negative, its t-statistic is very low. Further, the elasticity is far below that found in Section II, and far below values that seem reasonable in light of recent research (see Hamermesh-Grant, 1979). The output elasticity is also quite low in light of those found in previous work (Hamermesh, 1976). It is impossible to believe in the degree of increasing returns implied by the estimate of γ. The trend coefficients are positive and always significant. This too is disturbing in view of the usual interpretation of them as reflecting increases in productivity.

We added a variable like MINT, but based on adult labor costs and employment weights, to (3b) to test whether a change in the effective minimum wage directly affects the employment of adults. The χ-test of this hypothesis is 1.50, not significantly different from zero.[17] We may conclude that our interpretation of the effective minimum wage variable here and in Section II

Table 4

System of Labor Demand Equations, Private Nonfarm, 1954: I - 1978: IV

$\hat{\alpha}_1$	-.118 (-.42)	$\hat{\gamma}$	.269 (4.35)
$\hat{\beta}_1$	.118 (.42)	$\hat{\delta}$	-.0834 (-1.62)
$\hat{\alpha}_2$	.0067 (.42)	$\hat{\kappa}_1$ (Trend)	.0094 (8.09)
$\hat{\beta}_2$	-.0067 (-.42)	$\hat{\kappa}_2$ (Trend)	.0033 (4.63)
$\hat{\rho}_T$	.804 (15.15)		
$\hat{\rho}_A$	.935 (23.07)		
R^2_T	.988		
R^2_A	.998		

as a reflection of the truncation of the distribution of labor costs for teenagers is not inconsistent with the data. This finding allows us to interpret an increase in the effective minimum wage in the context of the models in (2') and (3) as directly affecting only the employment of teenagers. There is, though, an indirect effect on the employment of adults: With a rightward movement in the truncation point of the distribution of teenagers' labor costs, their average labor cost increases, and there is some substitution toward adult workers.

The elasticity of the effective minimum wage variable is negative and almost significantly different from zero, though its size is somewhat below that in Section II. The basic message is that, even if we take the theory of factor demand seriously and modify it to include the effect of the minimum wage, we still find a negative employment effect on teenagers as the effective minimum rises. No matter what formulation we have used - from the hybrid nontheoretical model in (2) to system (3) - increased coverage and higher legislated minima are found to reduce the employment of teenagers.

IV. THE MINIMUM WAGE AND FACTOR SUBSTITUTION

The theoretically based estimating models we have constructed must stem from some underlying production or cost function. Here we examine how the minimum wage affects the structure of firms' costs or the nature of production using annual data on the employment of teenagers and adults, and on services of capital. The view implicit in the model is that a higher minimum wage constrains the factor choices of the firm and thus raises its costs at a given output.

The work in this section is based on the flexible translog form (see Berndt-Christensen, 1974, for an early application). Throughout the discussion we use a cost rather than a production function. Though these are dual to each other, and thus should theoretically give identical results, this does not in practice occur. Both the nature of the translog form as an approximation, and the problem of finding the appropriate terms - prices or quantities - that can be treated as exogenous for estimation purposes, have been cited as causing differences between estimates produced using the cost or production function appraoches. Although Grant-Hamermesh (1981) argue that the production function approach, in which quantities are taken as exogenous, is more appropriate for estimating substitution parameters among groups of workers disaggregated by age and sex, this argument rests on the assumed relative inelasticity of labor supply in most groups. Since that assumption is likely to be incorrect for teenagers, the main focus of interest of this study, and since our estimates must in any case involve a price term in the form of the effective minimum wage, we use an approximation to a generalized cost function.

The translog cost function for this study is:

$$\begin{aligned} C = {} & Q + \alpha_0 + \alpha_1\, WY + \alpha_1'\, [WY \cdot MINT] + \alpha_2\, WA + \alpha_3\, PK \\ & + \frac{\beta_{11}}{2}\, [WY]^2 + \frac{\beta_{11}'}{2}\, [WY]^2\, MINT + \frac{\beta_{22}}{2}\, [WA]^2 + \frac{\beta_{33}}{2}\, [PK]^2 \\ & + \beta_{12}\, WY \cdot WA + \beta'_{12}\, WY \cdot WA \cdot MINT + \beta_{13}\, WY \cdot PK \\ & + \beta'_{13}\, WY \cdot PK \cdot MINT + \beta_{23}\, WA \cdot PK \quad , \end{aligned} \tag{4}$$

where C are the typical firm's costs, Q is output, PK is the user cost of capital, WY and WA are the wages of young and older workers, and the α_i, α_i', β_{ij} and β_{ij}' are parameters describing the firms' costs. (All variables are in logarithms; MINT is based on the series shown in the bottom part of <u>Table 1.</u>). The interaction terms between MINT and WY and the three price variables reflect the assumptions that a higher effective minimum affects costs by constraining firms' choice of inputs, and that this effect works only through the price of teen labor.

We can use (4) to derive equations describing the shares of total output accruing to each of the three inputs. Implicit in this derivation are the assumptions of constant returns to scale and price-taking firms. The derivation yields:

$$S_Y = \alpha_1 + \alpha_1' \text{ MINT} + \beta_{11} \text{ WY} + \beta_{11}' \text{ WY}\cdot\text{MINT} \tag{5a}$$

$$+ \beta_{12} \text{ WA} + \beta_{12}' \text{ WA}\cdot\text{MINT} + \beta_{13} \text{ PK}\cdot\text{MINT};$$

$$S_A = \alpha_2 + \beta_{12} \text{ WY} + \beta_{12}' \text{ WY}\cdot\text{MINT} + \beta_{22} \text{ WA} + \beta_{23} \text{ PK}; \tag{5b}$$

and

$$S_K = \alpha_3 + \beta_{13} \text{ WY} + \beta_{13}' \text{ WY}\cdot\text{MINT} + \beta_{23} \text{ WA} + \beta_{33} \text{ PK}, \tag{5c}$$

where S denotes the share of the particular factor. We expect β_{1i}' to be such as to imply price elasticities that are closer to zero as MINT is higher. For example, as MINT increases, the workers who are disemployed are the lowest-skilled, for whom the demand is likely to be most elastic (see Hamermesh-Grant, 1979); thus η_{YY}, the own-price elasticity will rise toward zero. Similarly, the substitutability of skilled and unskilled workers suggests the measured η_{YA} will fall toward zero when the least-skilled youth are disemployed.

The symmetry of cross-substitution effects has already been imposed in (5) by assumption in (4). However, homogeneity restrictions must also be imposed if the share equations are to make economic sense. These are:

$$\beta_{11} + \beta_{11} \overline{\text{MINT}} + \beta_{12} + \beta_{12}' \overline{\text{MINT}} + \beta_{13} + \beta_{13}' \overline{\text{MINT}} = 0; \tag{6a}$$

$$\beta_{12} + \beta_{12}' \overline{\text{MINT}} + \beta_{22} + \beta_{33} = 0; \tag{6b}$$

$$\beta_{13} + \beta_{13}' \overline{\text{MINT}} + \beta_{22} + \beta_{33} = 0; \tag{6c}$$

and

$$\alpha_1 + \alpha_2 + \alpha_3 + \alpha_1' \overline{\text{MINT}} = 1. \tag{6d}$$

These restrictions are quite standard in the empirical literature, though one should note that they are modified here by our inclusion of the effective minimum wage in (4).

One more homogeneity constraint is needed to complete the model. If the effective minimum increases, factor shares must still sum to one; the restriction must hold that:

$$\alpha_1 + \beta'_{11}\,\overline{WY} + \beta'_{12}\,[\overline{WY} + \overline{WA}] + \beta'_{13}\,[\overline{WY} + \overline{PK}] = 0. \tag{6e}$$

Restrictions (6a through e) cannot be valid for all values of the factor price variables and the effective minimum wage, so that there is some problem in interpreting them. We make the assumption that each constraint holds at the sample means of the factor prices and the effective minimum wage, implicitly assuming that the stochastic process generating (5) conforms with the restrictions imposed by theory only at the mean of the process. To denote this, we write superior bars over the price terms in (6); because of the constraints (6), equations (5) contain eight independent parameters. This model too is estimated using iterative least squares.

The capital stock data cover both private and government capital, and are from Freeman (1979). The user cost of capital is computed accounting for changes in the tax treatment of capital, depreciation and capital gains. Data on the labor quantities and prices are based on the Money Incomes of Families and Persons (CPR Series P-60). The estimates cover twenty-one annual observations, 1955-1975. The input prices WY and WA are based on the annual incomes of full-time, year-round workers ages 14-24 and 25+ respectively. Both WY and WA were deflated to constant 1972 dollars using the deflator for the private business sector. Factor quantities were computed as full-time equivalent employment by prorating the total number of persons in each age group who reported some earnings by the ratio of their earnings to those of year-round, full-time workers. Thus we are implicitly assuming that each person in the two labor subaggregates works the same number of hours.[18/]

Table 5 shows the estimates of the parameters in (5). Those in the first column are based on a model in which all terms involving MINT have been deleted (in which α_1' and the β_{1j}' have been set equal to zero); those in the second column are based on the complete model in (5). It is worth noting that the fit of the complete model is statistically better than that of the model from which the minimum wage terms have been excluded: The χ^2-statistic describing this test is 29.01, significantly different from zero at the 99 percent level. Most of the parameter estimates in the full model are quite significant, though $\hat{\beta}_{12}$ and some from the terms in MINT are not. (This undoubtedly results from the instability induced by the small share of costs accounted for by young labor. As shown in Grant-Hamermesh, 1981, it is difficult to get sensible parameter estimates from systems like (5) when the average shares become small.) Given this problem, we should not expect high levels of significance for any of the estimated effects of the minimum wage on the substitution parameters that we calculate below.

We can use the estimates in the second column of Table 5 to calculate the implied partial elasticities of substitution, own substitution elasticities, and cross- and own-price elasticities. Partial elasticities of substitution are calculated from (5) as:

$$\sigma_{ij} = \frac{\beta_{ij} + \beta_{ij}' \text{ MINT}}{S_i S_j} + 1 \quad , \tag{7a}$$

while own-substitution elasticities are:

$$\sigma_{ii} = \frac{\beta_{ii} + \beta_{ii}' \text{ MINT}}{S_i^2} + 1 - \frac{1}{S_i} \quad . \tag{7b}$$

Cross- and own-price elasticities are calculated from (7a) and (7b) respectively by multiplying by the share of the factor whose price is assumed to change.

Table 5

Estimates of Parameters for the Three-Factor Translog Cost Functions with Symmetry and Homogeneity Imposed, 1955 - 1975

α_1	-.0778 (-.67)	-.356 (-.86)
β_{11}	.0368 (2.80)	.0625 (1.39)
β_{12}	-.0228 (-1.23)	-.0134 (-.73)
β_{13}	-.0141 (-1.38)	-.0103 (-1.53)
α_1'		.334 (1.22)
β_{11}'		.0306 (.95)
β_{12}'		-.0090 (-8.10)
β_{13}'		.0074 (7.46)
α_2	-2.00 (-3.53)	-1.91 (-7.28)
β_{22}	.258 (4.06)	.229 (7.07)
β_{23}	-.235 (-4.66)	-.228 (-9.74)
α_3	3.07 (6.61)	3.72 (15.71)
β_{33}	.249 (6.01)	.267 (8.71)
$\ln L$	142.64	157.14

Table 6 lists the values at the sample means of all the substitution and price elasticities involving youths.[19] The former are also presented as linear functions of the logarithm of the effective minimum wage. The estimated demand elasticity for young workers is quite low, -.59, though not nearly so low as that produced in the system in Section III (a system, though, that excluded capital). We find here that workers in the two groups are substitutes on average during the sample period. Young workers and capital are found to be complements, though the cross-price elasticity is essentially zero, and its accompanying t-statistic is tiny.

The most important finding of this section is implicit in the representation of the substitution elasticities as linear functions of MINT in Table 6. Increases in the effective minimum wage during the period 1955-1975 reduced the own-substitution elasticity of demand for young workers and decreased the extent to which employers were able to substitute older for young workers in response to an exogenous increase in the price of young workers. Based upon the value of MINT in 1955, $\eta_{YY} = -.718$, and $\eta_{YA} = .643$; for 1975 the comparable elasticities are -.233 and .500. We observe the same result for η_{YK}, though the very low t-statistic attached to the estimate prevents us from drawing any useful inferences from it. These estimates provide evidence for our rationale for including the minimum wage in the cost function (4). They imply that a higher effective minimum wage induces a rigidity into firms' responses to exogenous changes in factor prices by restricting the range of choices. This inference is strengthened by the calculation that the estimated $\partial C/\partial MIN > 0$; we estimate that increases in the legislated minimum wage raise the estimated total cost, as logic suggests they should.

V. THE NET EMPLOYMENT EFFECT OF THE MINIMUM WAGE AND SOME POLICY SIMULATIONS

Here we use the results of Sections II-IV to analyze the effects of changes in the FLSA. The effects cannot simply be computed on the basis of the estimated

Table 6

Substitution Parameters and Price Elasticities from the Translog Cost Model, 1955-75

	Partial Elasticities of Substitution		
	σ_{YY}	σ_{YA}	σ_{YK}
As function of MINT	1.180 +8.00 MINT	.654 -.233 MINT	.465 +.383 MINT
At mean of MINT	-9.53 (-2.30)	.966 (1.98)	-.049 (-.08)

	Demand Elasticities		
	η_{YY}	η_{YA}	η_{YK}
At mean of MINT	-.590 (-2.30)	.605 (1.98)	-.0156 (-.08)

minimum wage elasticities, for changes in the legislated minimum or its coverage will change average labor costs. This will have an additional effect on teen employment through the variable WT included in (3) or in (2').

Writing all prices in logs, assume that the distribution of WT (teen labor costs) is normal. Then, following Johnson-Kotz (1970, Volume 2, p. 81), the observed mean of this variable is :

$$E(WT) = \mu + \frac{f\left(\frac{MIN - \mu}{\sigma}\right)}{1 - F\left(\frac{MIN - \mu}{\sigma}\right)} , \tag{8}$$

where f is the normal density function; F the normal distribution function; μ is the mean of the untruncated distribution; and σ is its standard deviation. From (8) the derivative of the mean of WT with respect to an increase in the effective minimum wage produced by an increase in MIN is:

$$dWT/dMIN = \frac{1}{\sigma} \{f'(\cdot)/[1-F(\cdot)] + f^2(\cdot)/[1-F(\cdot)]^2\} , \tag{9}$$

where $(\cdot)$ denotes the argument has been suppressed.

Remembering that ER and WR are differences in logs of teen and adult employment and labor costs reprectively, and treating MINT as MIN - WT, we can write:

$$dET/dMIN = \partial ET/\partial MINT \; [1-dWT/dMIN] + [\partial ET/\partial WT] \; [dWT/dMIN]. \tag{10}$$

The partial derivatives in (10) are either the regression coefficients from (2') under the assumption that adult employment is not directly affected by changes in MIN, or from (3). Therefore, if we evaluate the effect of an increase in MIN on the truncated mean of the distribution of teen labor costs, we can evaluate the net effect of changes in MIN on teen employment.

We make three alternative assumptions about how changes in the legislated minimum wage truncate the distribution of teen labor costs:

(1) All unemployed teens owe that status to the effects of the minimum wage, but teens who are out of the labor force are unaffected;

(2) The fraction truncated is equal to the highest fraction of teens

(.115) inferred as disemployed in the Meyer-Wise (1981) estimates of wage distributions of teens;

(3) Same as (1), but using the teen labor force, L, as a base rather than the teen population, P.

Based on averages from 1954:I to 1978:IV the fractions truncated under assumptions (1) and (3) are .069 and .149 respectively. We assume that σ, the standard deviation of the untruncated wage distribution, equals 1 or .5. (The latter figure is roughly in line with the Chiswick-Mincer (1972) estimate for the truncated distribution for adults, and with Meyer-Wise (1981) for the untruncated distribution for teens.)

In Table 7 we examine the employment effect of a youth subminimum wage equal to 75 percent of the adult minimum. The calculations ignore scale effects; only substitution between teen labor and other factors is dealt with. The estimated impacts of the 75 percent subminimum are not small, especially if we assume that the average wage of teens would decline as low-wage teens become employed. We have, though, ignored any changes in compliance and in the use of student exemptions that might occur.[20]

The estimates in Table 7 can be used with additional assumptions to gauge the total impact on teen and adult employment of a 75 percent subminimum. The substitution effects on teens are as listed in that table; the disemployment effect on adults is:

$$dEA = (\partial EA/\partial WT)(dWT/dMIN)\ dMINT.$$

Assuming that truncation of the teen wage distribution is based on U/P (.069 of the distribution is truncated), and that $\sigma = .5$, dEA = -.089 percent. This compares to dET = 2.86 percent, shown in Table 7. Making a conservative assumption about the share of output accruing to teens at the minimum wage, the scale effect is .161 percent.[21] Based on 1979 employment of 9356 thousand teenagers and 88,961 thousand adults, we infer that a 75 percent subminimum would create 283 thousand jobs for teens and 62 thousand

Table 7

Percentage Effect of a 75 Percent Youth Subminimum on Teen Employment or Relative Teen/Adult Employment

	Based on:			
	Teen Employment (Table 4)		Relative Teen-Adult Employment (Table 3)	
Standard Deviation of the Wage Distribution	σ		σ	
	1	.5	1	.5
Assumption about Truncation of the Teenage Wage Distribution (and Fraction Truncated)				
No Truncation	2.40	2.40	2.95	2.95
Unemployed Ratio (.069)	2.63	2.86	5.01	7.07
16-24 Year-olds Disemployed (.115)	2.71	3.02	5.72	8.49
Unemployment Rate (.149)	2.76	3.12	6.14	9.33

jobs for adults. Using the same estimates of $\partial ET/\partial MINT$ and dWT/dMIN, but using σ_{YY} and σ_{YA} from Table 6 (implying greater own -and cross-price effects), the job creation estimates are 523 thousand and -230 thousand respectively.[22]

Different estimates will be produced depending upon assumptions about dWT/dMIN. However, though the scale effect may be small, the direct effect on teen employment (through dET/dMINT) is large enough that with reasonable estimates of substitution possibilities between youths and adults, far more teen jobs would be created than adult jobs lost.

Finally, though we cannot draw any direct inferences about how a higher effective minimum affects the size distribution of income, we can use the results in Table 5 to calculate the effect of a given increase on the shares of income of each of the three factors. These are given by:

$$\partial S_A/\partial MIN = \beta'_{12} \left\{\frac{dWY}{dMIN} [MINT - WY] + WY\right\} + \beta_{12} \frac{dWY}{dMIN} \quad ; \qquad (11a)$$

$$\partial S_K/\partial MIN = \beta'_{13} \left\{\frac{dWY}{dMIN} [MINT - WY] + WY\right\} + \beta_{13} \frac{dWY}{dMIN} \quad ; \qquad (11b)$$

$$\partial S_Y/\partial MIN = 1 - \partial S_A/\partial MIN - \partial S_K/\partial MIN \quad . \qquad (11c)$$

In Table 8 we list these partial derivatives and some estimates of what intrasample changes in MIN have done to the estimated factor shares, using $\sigma = 1$ and the same three assumptions about the truncation point. The estimates imply that the gradual increase in the effective minimum wage raised the shares of capital and youth and lowered adults' share of total factor returns. That this happened is implicit in our earlier findings that the demand for younger workers is less than unit elastic, and that younger and older workers are substitutes.

VI. CONCLUSIONS AND IMPLICATIONS

We have provided several advances over the previous literature on employ-

Table 8

Effects of Increases in the Legislated Minimum Wage on Factor Shares, Based on Table 5

	$\partial S/\partial MIN$			Increase of MINT from 1968: IV to 1978: IV		
			Share of:			
Assumption about Truncation of the Teenage Wage Distribution (and Fraction Truncated)	Y	A	K	Y	A	K
No Truncation	.0138	-.0769	.0631	.0021	-.0115	.0095
Unemployed Ratio (.069)	.0156	-.0594	.0438	.0023	-.0089	.0066
16-24 Year-olds Disemployed (.115)	.0162	-.0533	.0371	.0025	-.0081	.0056
Unemployment Rate (.149)	.0165	-.0498	.0333	.0025	-.0075	.0050

ment demand and the minimum wage. The more complete measures of labor costs we have developed uniformly improve the fit of equations describing relative teen-adult employment and increase the estimated (negative) response to increases in the effective minimum wage. The employment elasticities generally remain significant and of roughly the same magnitude when these equations are respecified to give them a basis in demand theory.

If one views increases in the effective minimum wage as constraining firms' choices on factor inputs by restricting the range of employees who may be hired, one can model a cost function that includes the minimum wage. We estimate equations implied by such a function for three inputs---workers 14-24, workers 25+, and capital. Higher effective minima have reduced firms' ability to substitute among groups of workers in response to exogenous changes in their relative wages. This is consistent with the notion that, by restricting employers' choice sets, higher minima add rigidity to the labor market.

The most striking finding of this study is the remarkable robustness of the negative teen employment elasticity in response to higher minimum wages and expansions of the coverage of the minimum wage, holding output constant. Regardless of the choice of wage measures or the choice of models, the elasticity for the private nonfarm sector is on the order of -.1. Though these minimum wage elasticities do not seem very large, one must remember that they are estimated over a period that saw a tremendous increase in the effective minimum wage. Thus the implied effect of expansions of the minimum wage law on teen employment has been substantial. A youth subminimum wage would have offset some of these effects, with relatively little displacement of adult workers.

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FOOTNOTES

1. The labor cost series are described and listed in Hamermesh (1981). They are based on combining information from the biennial Chamber of Commerce series on employee benefits with BLS quarterly series on compensation and monthly series on average hourly earnings by sector. Data on the user cost of training are constructed using assumptions about the burden of cost of specific training and estimates of cross-section regressions that include terms for experience and tenure, and thus allow the estimate of the effects of firm-specific experience.

2. The data on coverage by sector are unpublished and were provided to me by the staff of the Minimum Wage Study Commission. Data on the legislated minima applicable in each sector are from published information from the U.S. Department of Labor, Employment Standards Administration.

3. None of the previous studies of the employment effects of the minimum wage has gone beyond using AHE as the denominator of an effective minimum wage measure.

4. For 1967-1978, RY is the ratio of the usual weekly earnings of full-time workers ages 16-24 to those of all full-time workers. The ratio is based on unpublished data provided by the Bureau of Labor Statistics. Since those series only began in 1967, for 1955-1966 we prorated the ratio of incomes of full-time, year-round workers, ages 14-24, relative to incomes of all full-time, year-round workers, by the usual weekly earnings ratio for 1967. Since no data on the incomes of young full-time, year-round workers were available for 1954, we arbitrarily assumed that the ratio RY was identical in 1954 and 1955.

5. For youths the labor cost measure is calculated as $AHE_{it} \cdot RY_t$, where i is the industry and t the quarter. ($COSTWK_i$ or $ECNT_i$ could be used in place of AHE_i.) For adults, labor costs are (W-eWT)/(1-e), where W is the average labor cost in the industry at time t, WT is the labor cost for youths, and and e is the fraction of young workers in the industry at time t.

6. Welch (1974) used ordinary least squares estimation for the private nonfarm sector, but took into account contemporaneous correlation of the residuals in the equations for the three industries. (He also estimated an equation over a composite of all other industries.) Since autocorrelation is likely to be the most severe problem in time-series estimation, the use of the Cochrane-Orcutt procedure is probably the best choice if one wishes to go beyond least squares.

7. For the private nonfarm sector and each of the larger industries the time trends in (2) were positive and significant. All the coefficient estimates on the adult employment variable were negative and significant.

8. Mincer (1976) found the addition of lagged effective minimum wage terms significant, though Welch (1974) did not. Though the sum of the terms in an eight-quarter Almon lag structure estimates using a quadratic without end-point constraints differed little from the coefficients of MINJ in Table 2, the specification did lower $\hat{\sigma}_{\varepsilon}$ slightly except in retail trade.

9. Other studies, such as Mincer (1976) and Ragan (1977), mix elements of demand and supply models. Only Welch-Cunningham (1978) has a sound basis in the theory of factor demand, and that study has problems with its attempts to disaggregate teen labor into three subgroups to find substitution elasticities within the teen group.

10. Clearly, there may be some "bunching" of the distribution at the minimum. For our argument to hold, though, we only require that a higher minimum cause a greater truncation of the distribution.

11. To examine whether induced disemployment elsewhere affects employment in a specific sector, MINT for the private nonfarm sector was added to the equations for each of the three industries. In no case was the coefficient on this variable significantly different from zero, nor did its addition ever change the coefficient of $MINT_i$ in Table 3 by more than one standard error.

12. With the introduction of the relative price variable the importance of a simultaneous-equations bias may be increased. (Insofar as the effective minimum wage variable includes an average wage, it exists already in (2) and in equations estimated by others.) To account for this (2') for the private nonfarm sector was reestimated using an instrumental estimate for WT in the relative price and the MINT variables. (The instrumental equation included the numerator of (1), DUMS, and teenage and adult population. The coefficient of the minimum wage in this equation was .032; its t-statistic was 1.58.) The reestimation of (2') yielded a relative price elasticity of -2.41 (t= -2.34), but a much lower elasticity on MINT, -.027 (t = -.42).

13. The other coefficients in the equations changed only slightly, and, as before, the fits in the equations using ECNT were better than in those using AHE.

14. The output measure is gross domestic business product deflated by the gross domestic product deflator. These series were from the CITIBASE file.

15. Variables based on AHE were also used in estimating (3). As in Section II, we found that the fits were slightly inferior, and the estimated minimum wage elasticities were slightly lower in absolute value.

16. In the system in which only homotheticity has been imposed, the coefficient on MINT, along with it t-statistic, is -.056 (-1.13).

17. We know from Grossman (1980) that increases in the minimum wage have only slight effects on wages above the minimum. Insofar as young workers have less human capital, this evidence for the assertion that attention be directed toward the effect of higher minima on the employment of youths corroborates our result.

18. As a check on the validity of using capital stock and user cost series together with labor input and price data constructed from an entirely different source, it is worth reporting some statistics describing these data. The mean shares are .0619, .6263, and .3118 for youths, adults, and capital, respectively. Moreover, the mean annual full-time earnings seem quite reasonable in light of previous work.

19. Implicit in the calculations of η_{YY} and η_{YA} based on (7) is the assumption that the effective minimum wage stays unchanged as WY varies.

20. Ashenfelter-Smith (1979) build a model that suggests firms will decrease compliance as the effective minimum rises. While they present no direct evidence on this, they do show the widespread nature of noncompliance.

21. This is derived by assuming that the share of youths earning at or below the minimum is .56 percent, based on the assumption that one-third of all teens earn the minimum or less, that their average wage is half that of other teens, and that teens' share of output is 3.3 percent.

22. The implied η_{TT} is calculated as σ_{YY} (-9.53) times the share of teens (.033). This latter is calculated as teens' share of labor earnings from Section III times labor's share from Section IV. η_{AT} is just σ_{YA} (.966) times .033.