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The Impact of the Minimum Wage Law on the Employment, Occupational Choice, and School Enrollment Decisions of Graduating High School Seniors

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Abstract

This study analyzes the impact of minimum wages on the educational and occupational choices of Maryland high school graduates from 1951 though 1969. Individuals are assumed to choose an occupational or educational option so as to maximize expected lifetime utility. Minimum wages are assumed to affect this choice in three ways: minimum wages alter the distribution of earnings in each occupation, they alter the length of time required to find a job in each occupation and they alter the amount of on-the-job training provided in each occupation. We develop a simple model which embeds these minimum wage effects in a system of equations representing the range of activities available to high school graduates. The activities studied include enrolling in one-to-three year vocational or technical colleges; enrolling in a four-year college; selecting one of six occupations: professional/managerial, clerical/sales, crafts, operative/laborer, service, or farming; or remaining unemployed. We estimate this system of equations separately for male and female high school graduates.

The results illustrate that the effects of minimum wages vary with the level of school quality and home inputs. Graduates of better schools or from wealthier school districts benefit from increasing minimum wages through increased job opportunities whereas those from poorer quality schools and poorer school districts are harmed. There is some evidence that an increase in the minimum wage reduces on-the-job training opportunities in occupations and hence induces individuals to enter junior colleges and vocational programs. However, we fail to find strong evidence linking minimum wages to changes in enrollments at four-year colleges or changes in unemployment status in this sample.

Introduction

Few topics in labor economics have been as widely studied as the minimum wage. A recent survey of the literature (Brown et al., 1982) has 111 citations on the topic. Most of these studies have attempted to measure the impact of raising the minimum wage on employment, unemployment, or labor force participation. Estimates of minimum wage effects have been obtained for the aggregate economy and have been disaggregated by age, race, sex, and industry. A common finding is that the adverse impacts of minimum wages on employment are most severe for younger cohorts, and particularly for young blacks. On the other hand, more skilled workers are not hurt and may even gain from minimum wages.[1]

A common explanation for the differential impacts of minimum wages on various demographic groups is that these groups differ in their average level of embodied human capital. Presumably the groups most adversely affected by minimum wages will be those whose skill levels are lowest, on average. Thus the young, who have relatively low levels of both education and on-the-job training, will be the most likely to lose job opportunities, while more highly skilled workers may gain employment opportunities through substitution effects.

A second connection between human capital and the minimum wage is through the opportunity for on-the-job training. Since minimum wages increase the costs of training for firms, it is commonly hypothesized that firms will cut back on training when minimum wages are increased. Therefore some of the gains of minimum wages are reversed through slower wage growth.[2] The lower opportunity for employer financial training coupled with reduced job opportunities may induce members of the low-skilled demographic groups to seek self-financed training through vocational, junior college, or college opportunities.[3]

This paper examines the entry-level occupational and educational decisions of successive cohorts of high school graduates. It seeks to answer several questions that have not been adequately addressed in the literature. First, we study whether changes in minimum wages alter the mix of occupational entry decisions. Previous studies have not considered the distributional impacts of minimum wages on occupational choice. Second, we investigate the extent to which minimum wages increase the incentive to seek postsecondary education. Our inclusion of vocational and junior college training is, to our knowledge, unique to our study. Third, we determine whether differences in endowments of school quality or family inputs into human capital production and the resulting heterogeneity in skills result in differential impacts of minimum wages.

The study makes use of an explicit model of occupational choice under uncertainty. An agent decides which occupation or training option to select, given knowledge of his current level of human capital, the moments of the distribution of earnings per unit of human capital in all occupations and education categories, and the level and coverage of the minimum wages across all occupations. By specifying the form of the utility function, we are able to derive structural occupational supply equations that allow us to extract the impact of minimum wage policy on occupational choice. Occupational decisions should reflect the agent's perceptions of the effects of minimum wages on on-the-job training opportunities, on the distribution of earnings per unit of human capital, and on the length of job search. Therefore, we may implicitly obtain estimates of these effects by observing how the distribution of occupational choice varies in response to changes in the minimum wage in the context of our structural model.

Model

We assume that agents live for two periods. In the first period, the agent specializes in production of human capital. In the second period, the individual initially accumulates on-the-job training or other occupation-specific training, such as education or vocational training, and then works full time in the occupation for the remainder of the period. In this study, we treat the college or vocational training choice as if it was an occupational choice. For this reason, we use the more general term "activities" to represent the range of occupational or educational choices available to individuals who have just completed high school.

Life earnings from each activity are not known with certainty, but the moments of the distribution of earnings are known. At the beginning of period two, the agent decides which activity to choose so as to maximize lifetime expected utility, based on human capital accumulated in period one and the training opportunities and the earnings distributions across all activities. The minimum wage alters the probability of choosing an activity through its impact on the wage distribution, on the waiting time to obtain a job in the occupation, and on the on-the-job training opportunities in the occupation.

At the end of period one, an individual has an endowment of human capital obtained in school. This stock of human capital may not be equally applicable to all activities, so we designate the human capital specific to activity i as h_i. The individual also has an endowment of wealth from period one equal to Y_1.

In period two, the individual selects an activity. The activity will further endow the individual with activity-specific human capital, h_{2i}, which represents on-the-job training or education associated with activity i. Earnings in the activity are determined by the individual's stock of human capital, $h_i = h_{1i} + h_{2i}$, and the earnings per unit of human capital, W_i. The individual uses his period two earnings along with his wealth carried over from period one to purchase consumer goods in period two. In the absence of any bequest motive, the individual sets consumption in period two, C_{2i}, equal to total income available in activity i in period two, $W_i h_i + Y_1$.

If earnings per unit of human capital are known with certainty, the individual chooses activity i so as to maximize lifetime earnings and thus lifetime utility. In general, earnings in activity i will not be known with certainty. Future labor demand and supply conditions are unknown, as are future hours of work available in the activity. Finally, the individual will not know his degree of luck or success in the activity. This implies that, while the individual can observe the distribution of earnings in the activity, the individual cannot know with certainty where he will end up in that distribution.

The individual's choice is therefore to choose an activity so as to maximize expected utility in period two, given knowledge of his stock of activity-specific human capital produced in period one, the availability of further human capital accumulation in period two, and the distribution of earnings per unit of human capital. The choice may be written as

(1) maximize $E\{U_i(C_{2i})\} = E\{U_i(Y_1 + W_i h_i)\}$.

Minimum wages affect the expected utility of selecting an activity in three ways. First, minimum wages may alter the distribution of earnings per unit of human capital in the activity. We cannot predict *a priori* the direction of change in the moments of the distribution of earnings in the activity. Some workers will work longer hours and others will work shorter hours as firms adjust their skill-specific labor demand to the presence of minimum wages.[4] To the extent that spillover from the covered to the uncovered sectors of the occupation occurs, wages in the uncovered sector of the occupation will be depressed as minimum wages in the covered sector are increased.

A second effect is that minimum wages may alter the amount of on-the-job training available in an occupation. To the extent that minimum wages increase the firm's costs of providing training, firms will reduce their investment in the human capital of their employees.[5] This lowers the amount of h_{2i} available in the occupation, lowering the lifetime human capital, h_i, and lifetime earnings, $W_i h_i$, in the occupation.

The third effect of minimum wages on expected utility is to alter the waiting time or search costs associated with finding a job in activity i. If new entrants must wait for premium jobs in a given occupation, minimum wages may alter the length of waiting time. Minimum wages may also establish queues for jobs in the covered sector relative to the uncovered sector. New entrants into the occupation may "wait" in the uncovered sector for openings to develop in the covered sector.

The standard argument is that increasing the minimum wage increases the length of the queue. In fact, the effect of minimum wages is more complex and depends upon the level of human capital embodied in the job entrant. If the job entrant has a high level of skill relative to the minimum wage (due perhaps to high-quality schooling or high levels of household investment in human capital in period one), minimum wages will not lengthen and may shorten the individual's queuing time as employers substitute toward higher-skilled applicants and away from lower-skilled applicants. On the other hand, if the job entrant has a low skill level relative to the minimum wage, the person's queuing time will lengthen. In other words, we posit that for an individual with human capital h_{1i}, length of queue, q_i, will fall as the minimum wage, MW_i, increases until the minimum wage equals the marginal product associated with h_{1i}. Thereafter, we expect that the individual's waiting time will increase as the gap between the minimum wage and the individual's marginal product rises. This relationship is illustrated in Figure 1.

Adding the effects of minimum wages to (1), we have

(2) maximize $E\{U_i[Y_1+W_i(MW_i)h_i(MW_i),\ q_i(MW_i)]\}$.

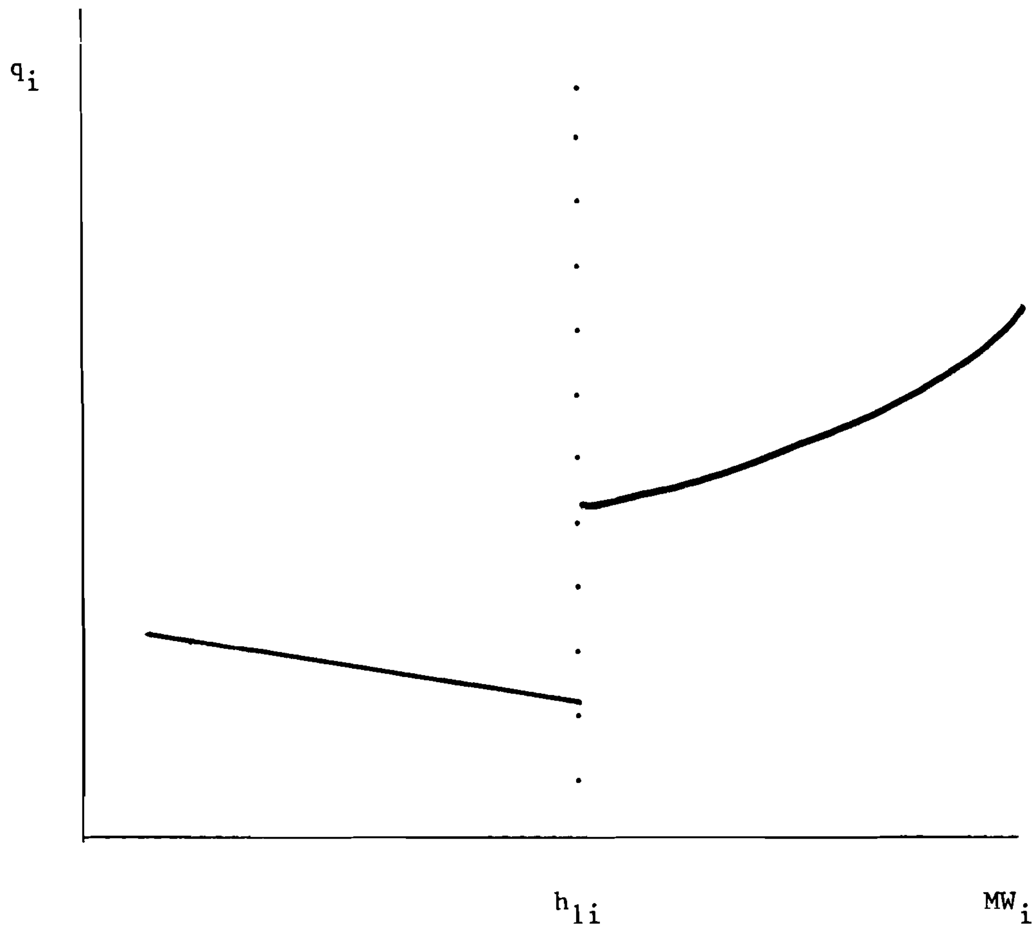
Equation (2) represents the expected utility from consumption in activity i, taking into account the disutility from waiting in queues associated with occupation i.

In order to make the theory empirically tractable and thus derive explicit empirical measures of the effects of minimum wages on activity choice, it is necessary to impose additional structure on equation (2). We assume an exponential utility function of the form:

(3) $U_i = \alpha' X_i - \phi q_i - \beta \exp(-\gamma W_i h_i - Y_1)$,

where X_i is a vector of variables measuring the hedonic attributes of activity i, α is a vector of taste parameters associated with these hedonic attributes, ϕ is a taste parameter for waiting time, β is a taste parameter for consumption, and γ is a measure of absolute risk aversion. ϕ, β and γ are assumed to be positive. The individual is assumed to know X_i, q_i, h_i, and Y_1 with certainty, but he knows only the distribution of W_i. Equation (3) has several intuitively appealing implications. It is a monotonically increasing concave function in consumption. Utility asymptotically approaches $\alpha' X_i - \phi q_i$, the individual's tastes for activity i, as W_i approaches infinity. Individuals facing the same wage distribution

FIGURE 1



q_i is the length of an applicant's queue for covered sector jobs in occupation i. MW_i is the minimum wage in occupation i. h_{1i} is the applicant's human capital, measured in terms of his marginal productivity.

may choose different activities due to heterogeneity in activity-specific skills, wealth, degree of risk aversion, or tastes for the occupation. Heterogeneity in taste becomes more important in explaining activity choice as the degree of risk aversion rises, as wealth rises, as the taste for consumption falls and as human capital falls. In the limit, activity choice will depend only on taste as either Y_1 or γ approaches infinity and as either h_i or β equals zero. In other words, agents who base their activity choice solely on taste will be those who are extremely wealthy or extremely risk averse, or those who have no marketable skills or those who derive no utility from consumption.

Because the individual knows Y_1, X_i and q_i with certainty, we can take the expectation of U_i as

$$(4) \quad E(U_i) = \alpha' X_i - \phi q_i + E\left[-\beta_1 \exp(-\gamma W_i h_i)\right],$$

where $\beta_1 = \beta \exp(-Y_1)$ and E is the expectation operator. We can approximate the expected utility from activity i by taking a MacLaurin series expansion about γh_i. Assuming that the distribution of W_i has at least m moments, we have

$$(5) \quad E(U_i) = \alpha' X_i - \phi q_i - \beta_1 - \beta_1 \sum_{k=1}^{m} \frac{(-\gamma h_i)^k}{k!} E(W_i^k) + \beta_1 \xi_i,$$

where $\beta_1 \xi_i$ is the error of our approximation. Equation (5) shows that the expected utility from selecting activity i depends on the hedonic characteristics of activity i and on the m moments of the earnings distribution of occupation i, $E(W_i)$, $E(W_i^2)$, ..., $E(W_i^m)$. The first moment of the distribution will increase expected utility. The second moment of the distribution will decrease expected utility. A positive (negative) third moment will increase (decrease) expected utility.

The probability of selecting activity i over another activity n depends upon the relative expected utility of selecting the activities. This probability may be written

$$(6)\quad P_i = P_r[E(U_i) > E(U_n)]$$
$$= P_r[(-\xi_i+\xi_n) < (\frac{\alpha'}{\beta'}(X_i-X_n) - \frac{\phi}{\beta'}(q_i-q_u)$$
$$- \sum_{k=1}^{m} [\frac{(-\gamma h_i)^k}{k!} E(W_i^k) - \frac{(-\gamma h_n)^k}{k!} E(W_n^k))].$$

To derive an estimable form of (6), it is necessary to specify the distribution of the error terms ξ_i and ξ_n. It is convenient to assume that these random variables have independent Weibull distributions. This allows us to derive equations for n-1 equations which show the logarithm of the odds of selecting an activity over an arbitrary n^{th} activity. The equation is

$$(7)\quad \log\left(\frac{P_i}{P_n}\right) = \sum_{k=1}^{m} [\frac{(-\gamma h_i)}{k!} E(W_i^k) - \frac{(-\gamma h_n)^k}{k!} E(W_n^k)] + \frac{\alpha'}{\beta'}(X_i-X_n) - \frac{\phi}{\beta'}(q_i-q_n)$$

for i = 1, 2, 3, ..., n-1.

Although the individual knows h_i, X_i and q_i, these are not directly observable by the econometrician. We must therefore posit some underlying functional relationship between these unobservables and some measurable instruments including the minimum wage. We approximate the activity-specific human capital as[7]

$$(8)\quad h_i = \delta_{0i} + \delta_{1i}S_1 + \delta_{2i}H_1 + \delta_{3i}MW_i,$$

where S_1 is a measure of school quality in period one and H_1 is a measure of household inputs into human capital in period one. S_1 and H_1 control for human capital production in period one. The minimum wage MW is a control for human capital production in period two. If minimum wages reduce a firm's incentive to provide on-the-job training, then increasing the minimum wage in activity i should reduce activity i specific human capital.

Waiting time for premium jobs or covered sector jobs in activity i is also a function of the minimum wage. In our formulation,

(9) $q_i = \pi_{0i} + \pi_{1i} MW_i$.

As argued above, we cannot sign π_{1i} without knowledge of the level of skill relative to the level of the minimum wage.

The hedonic factors X_i are not observed directly. However, such factors are unlikely to change rapidly over time. To the extent that these hedonic characteristics are time invariant, their effects are captured by the constant term in the regression. We may also have variation in tastes for activities across geographic areas. These differences in tastes are captured by a vector of county-specific dummy variables, D_i. Other possible proxies for variation in taste such as sex, years of education, or age are not relevant to this study since our sample holds all these factors fixed.

In general, we may estimate equation (7) with any number of moments of the distribution of earnings per unit of human capital. In applications, it becomes impossible to extend the approximation very far. The problem is that each added moment of the earnings distribution requires that we add the corresponding higher-order moment of the human capital production process as well. Using even the simple form in equation (8) involves four interaction terms with the first moment of the earnings distribution, ten interaction terms with the second moment, twenty interaction terms with the third moment and so on. As higher-order moments of the distribution are admitted to the equation, problems of multicollinearity become increasingly severe, eventually leading to an uninvertable moment matrix of the regressors. For this reason, we constrained our estimation to include only the first two moments of the earnings distribution.

Data

We estimate equation (7) by utilizing a unique and rich set of data. Each year between 1951 and 1969 the Maryland State Board of Education has published data (in its Annual Report) on the number of high school graduates, broken down by sex and by their major type of activity in the year following graduation.[8] This survey categorized graduates into those who continued their education in schools requiring one to three years for a degree, (junior colleges, vocational schools; art, music, and drama schools) or in schools requiring four years for a degree (universities; state teachers colleges), or who took jobs broken down into six occupational groups (farming, fishing, and lumbering; operatives and laborers; service workers; craftsmen; clerical and sales; professionals and managers), and an "other" group of those who were not enrolled or employed (at home; miscellaneous).[9]

Our dependent variable is measured as the logarithm of the number of persons in each activity group i divided by the number in the farmer/farm laborer group. We have eight activities and dependent variables for men (two enrollment categories, five occupational employment categories, and the "other" group) but only seven for women (one less occupation, as discussed later in this paper). Farming was used as the reference group because there were very few instances in which no one entered that occupation. This was computationally convenient because we did not then have to worry about dividing by zero. The problem of taking the log of zero still remained in some instances when there were no graduates taking jobs in an occupation.

In such cases, we set the numerator's zero value equal to 0.1.[10] Descriptive statistics for the variables used in our regression analysis for men are summarized in Table I.

The measure of school quality (S) used in this study is the real value of instructional salaries per student enrolled, weighted by a school utilization ratio. To be more specific, we divided "salaries and wages for instruction at the secondary level" by the number of students enrolled in the school district and then expressed it in constant dollars. This value was multiplied by the attendance rate. In order to take account of the fact that a student's schooling investment occurred over several years, a six-year moving average was computed.[11]

A measure of average "family" wealth (H), which is used to control for home inputs into educational production, is also available from the Maryland State Board of Education. The board reports the total value of property other than corporate or railroad property which is assessable for tax purposes by county school districts. We compute a six-year moving average of real property values per capita as our proxy for "A."[12]

We used national income data from the U.S. Bureau of the Census (Series P-60) to develop our measures of the distribution of earnings by occupation and schooling. Income distributions by discrete intervals have been published annually for men and women by major occupation and by level of educational attainment. From these distributions, we estimated the annual mean and the second moment of income (each expressed in real terms) for each of our six occupational groupings and for persons completing 1-3 years of college and 4 or more years of college.[13]

Unfortunately, these Census Bureau data do not hold constant the level

TABLE I
DESCRIPTIVE STATISTICS FOR MEN

	Mean	Standard Deviation
Dependent Variable: $\text{Log}(P_i/P_F)$		
Professionals/Managers	-2.309	2.011
Clerical/Sales	.090	1.385
Craftsmen	-.678	1.587
Operatives/Laborers	.772	1.294
Service	-1.556	1.795
1-3 Year Schools	-.723	1.568
4 Year Schools	4.433	1.419
Other (Unemployed)	-.138	1.509
Differences in Income Means $(\mu_i - \mu_F)$		
Professionals/Managers	.459	.081
Clerical/Sales	.213	.064
Craftsmen	.215	.074
Operatives/Laborers	.094	.069
Service	-.023	.077
1-3 Year Schools	.175	.075
4 Year Schools	.366	.110
Other (Unemployed)	-.554	.123
Differences in Income Second Moments $(\sigma_i - \sigma_F)$		
Professionals/Managers	.601	.147
Clerical/Sales	.127	.098
Craftsmen	.074	.125
Operatives/Laborers	-.055	.131
Service	-.134	.142
1-3 Year Schools	.080	.085
4 Year Schools	.384	.161
Other (Unemployed)	-.380	.185
School Quality (S)	253.473	67.368
Assets (H)	1.869	.661
Miniumum Wage Index $(MW_i - MW_F)$		
Professional/Managers	.187	.059
Clerical/Sales	.163	.067
Craftsmen	.213	.068
Operative/Laorers	.248	.071
Service	.106	.050
1-3 Year Schools	.187	.059
4 Year Schools	.187	.059
Other (Unemployment)	.187	.059

of human capital within an occupation. As a consequence, an increase in the proportion of those who are college educated will tend to alter the published distribution of income, even though high school graduates taking jobs may have experienced no change in their actual income distribution. The method of purging the occupational income distribution data of educational attainment effects was to regress the logarithm of income on the logarithm of education, for each moment of their respective distributions. The anti-logarithm of the residuals from such a regression should yield the corresponding income moment purged of the education effect.[14] A similar procedure was used in purging of age effects the means and second moments of earnings for those completing 1-3 years and 4+ years of college. That is, the moments were computed for persons age 25-64 and these moments were regressed on corresponding moments of the age distribution for college-educated workers. Again, the anti-log of the residuals was utilized.

This procedure yielded eight time series, each around a mean of 1.0. Information on the mean and second moment of income for each occupation for high school graduates and for persons having completed some or 4 years of college was obtained from the 1950, 1960, and 1970 Census of Population. These data were used to establish the relative rankings of the eight categories in those years. Trends were built back into the series using the 1950, 1960, 1970 relative values as benchmarks.[15] Finally, we constructed a simple two-year moving average of the occupational means and second moments. This adjusts for the fact that high school graduates in 1951, for instance, are observed in their occupational and schooling pursuits in late 1951 and early 1952.[16]

Our minimum wage index is closely patterned after the index which is widely used in the minimum wage literature. Unfortunately, coverage data are not available by occupation. As a result, we approximated an occupation-specific index by using the distribution of occupational employment across industries as our weights.[17] To be precise, our index was computed as

$$(10)\quad MW_{jt} = \sum_i \{ [\frac{M_{mt}}{W_t} C_{mt} + \frac{M_{pt}}{W_t} C_{pt}]_i \frac{E_{ij}}{E_j} \},$$

where j indexes occupation, i indexes industry, t indexes the year, m indexes newly covered workers, p indexes previously covered workers, M is the minimum wage, W is the average industry wage, C is the proportion covered, and E is employment of those aged 18-19 in 1960. During 1961-1964 and 1967-1969 different minimum wage rates were in effect for <u>newly</u> covered workers and for those who were covered <u>prior</u> to the amendments of 1961 and 1967. National data on coverage and on average wage rates were used rather than data for Maryland alone. This is consistent with our use of nationwide data in estimating the means and second moments of the earnings distribution. National data were used in part owing to data limitations for Maryland and in part owing to the view that a national job market is not implausible for young, mobile high school graduates. National data may also be interpreted as an instrumental variable for local labor market measures.

Since the minimum wage variable is a key variable in this study, some comparisons between Maryland and the U.S. are appropriate. We have sufficient data to estimate the proportion of all nonsupervisory workers (excluding government and domestic workers) who were covered by the law for <u>both</u> Maryland and the U.S. in three years 1964, 1967, and 1969.[18] These proportions are reported at the top of Table II and are virtually identical.

TABLE II

COMPARISON OF MARYLAND AND U.S. MINIMUM WAGE COVERAGE DURING SELECTED YEARS OF THE 1960s

Proportion of Non-supervisory Employees
(Exc. Government* and Domestic Workers)
Covered by Federal Minimum Wage

	U.S	Maryland
1964	.660	.664
1967	.790	.779
1969	.818	.813

*Non-supervisory employees are not available for government and hence could not be included. None were covered until 1967. Domestics were never covered until 1974.

INDUSTRIAL DISTRIBUTION OF EMPLOYEES COVERED BY THE FEDERAL MINIMUM WAGE LAW

	1964		1967	
	U.S.	Maryland	U.S.	Maryland
Agriculture, Forest, Fishing	.000	.000	.012	.007
Mining	.019	.004	.013	.003
Construction	.082	.116	.079	.111
Manufacturing	.514	.464	.419	.368
Transportation, Communications, and Utilities	.117	.124	.093	.095
Wholesale Trade	.071	.064	.058	.050
Retail Trade	.088	.116	.105	.129
Finance, Insurance and Real Estate	.063	.066	.049	.055
Service (excluding domestics)	.047	.046	.112	.103
Government	.000	.000	.060	.079
TOTAL	1.000	1.000	1.000	1.000

Source: Published and unpublished data supplied by U.S. Wage and Hour and Public Contracts Division, U.S. Department of Labor.

More detailed data are available by industry coverage for 1964 and 1967. Since total non-supervisory employment is not available by individual industry for Maryland, we can only report the distributions of covered workers across industries. These comparisons are available at the bottom of Table II. The Fair Labor Standards Act covered slightly more workers in Maryland in construction, retail trade, and government and slightly fewer workers in manufacturing than was true nationally. But overall, the distributions are very similar. We conclude that the use of national coverage data is reasonable for Maryland.

It is more difficult to judge the suitability of using national average industry wage rates in computing our minimum wage index. Unfortunately the U.S. Bureau of Labor Statistics only publishes manufacturing wage rates for Maryland. One other industry series is available, average wages for hired farm labor in Maryland, from the U.S. Department of Agriculture. These wage rates are compared in Table III. Aside from some relative increase in Maryland manufacturing wage rates in the early 1950s, this series has been essentially identical in the U.S. and Maryland. Relative farm wage rates have varied more: Maryland rates rose somewhat in the early 1950s and fell in the 1960s. Since agriculture was not covered until 1967, these differences should have little impact on our index. Although the evidence is limited, manufacturing wage rates suggest that national data should be a good proxy for Maryland.

Most professional and managerial employees have not been covered by the minimum wage law. Interpreted literally, the minimum wage index should be zero for this occupation as well as for the one-to-three year and 4-year school groups and for the "other" category (those not enrolled and not

TABLE III

AVERAGE WAGE RATES IN MARYLAND AND THE U.S. BY MANUFACTURING AND AGRICULTURE FOR SELECTED YEARS, 1951-1969

	MANUFACTURING			AGRICULTURE		
	U.S.	Maryland	Maryland / U.S.	U.S.	Maryland	Maryland / U.S.
1951	$1.56	$1.47	.942	$ .77	$ .80	1.039
1953	1.74	1.66	.954	.82	.87	1.061
1955	1.85	1.82	.984	.82	.90	1.098
1957	2.04	2.06	1.010	.88	.94	1.068
1959	2.19	2.20	1.005	.95	1.01	1.063
1961	2.32	2.34	1.009	.99	1.04	1.051
1963	2.45	2.48	1.012	1.05	1.07	1.019
1965	2.61	2.62	1.004	1.14	1.14	1.009
1967	2.82	2.82	1.000	1.33	1.27	.955
1969	3.19	3.22	1.009	1.55	1.50	.968

Sources: U.S. Bureau of Labor Statistics, _Employment and Earnings_; U.S. Department of Agriculture, _Farm Labor_.

employed). In addition, farm workers were not covered until 1967, so their index is zero for 1951-1966. Even after 1967, only a minority of farm laborers (on large farms) were covered. Since farming is our base or reference group, the $(MW_i - MW_{farm})$ index would be essentially zero for these four categories and we could not estimate any impact of minimum wages on professional/managers, school enrollment, or on the "other" category. This is unfortunate in that one of our major objectives was to determine whether youths who are displaced from employment by minimum wages spill over into these alternative activities.

Our theoretical analysis suggests another approach. The notion that the length of the queue (q) is important implies that some young people who discover that the length of the queue for operative jobs in factories has lengthened may go to vocational school to invest in more human capital, or may take managerial jobs (in small businesses) to avoid the law, or may remain at home. All of these strategies may be interpreted more broadly as "waiting time" strategies. That is, youths may enter these alternative states while waiting for jobs to open up in the covered occupations. (Of course, they may also "wait" in lower-paying, non-covered operative--or other--jobs until higher-paying, covered operative--or other--jobs become available.) This suggests that it is appropriate to include a minimum wage variable in our professional/managerial, schooling, and "other" equations. An index averaged over all industries and occupations seems most relevant and was employed in our initial estimates. The sensitivity of this approach to alternative specifications is discussed below.

Zellner "Seemingly Unrelated" Regressions for Men

A. Queuing Effects

In this section, we report our estimates of a simplified version of equations (7-9). If we modify equation (8) by imposing that the human capital production parameters are equal across activities, by excluding the on-the- job training effects of minimum wages (so $h = \delta_0 + \delta_1 S + \delta_2 H$) and by limiting the system to the first two moments, then we obtain equation 11:

$$\text{(11)}\quad \log \frac{P_i}{P_n} = \gamma\delta_o[\mu_i-\mu_n] + \gamma\delta_1 S[\mu_i-\mu_n] + \gamma\delta_2 H[\mu_i-\mu_n] - \frac{\gamma^2}{2}\delta_o^2[\sigma_i-\sigma_n]$$

$$- \gamma^2\delta_o\delta_1 S[\sigma_i-\sigma_n] - \gamma^2\delta_o\delta_2 H[\sigma_i-\sigma_n] - \gamma^2\delta_1\delta_2 SH[\sigma_i-\sigma_n]$$

$$- \frac{\gamma^2}{2}\delta_1^2 S^2[\sigma_i-\sigma_n] - \frac{\gamma^2}{2}\delta_2^2 H^2[\sigma_i-\sigma_n] + \frac{\alpha'(X_i-X_n)}{\beta} + \frac{(\phi\pi_i)'(MW_i-MW_n)}{\beta} + \varepsilon_i,$$

where μ_i and σ_i represent respectively the first and second moments of the wage distribution of activity i, and ε_i is the approximation error due to the missing higher-order moments of the human capital production function and the wage distribution. In theory, more general activity equations may be derived by admitting higher-order moments of the earnings distribution and the production functions. In practice, however, as higher-order moments of these functions are added, problems of multicollinearity become severe.

Because of the possibility of contemporaneous correlation between the errors of the eight activity equations, equation (11) was estimated using Zellner's seemingly unrelated regression technique. A set of county dummy variables was added to each equation in order to control for county-specific effects. These dummy variables help to control for time-invariant taste differences and costs of job entry across counties and across activities. Other typical controls for taste such as sex, age, education, or region of the country are unnecessary in this application, since these characteristics are uniform across individuals in the sample.

In addition, the minimum wage variable is interacted with school inputs (S) and home inputs (H) respectively. In our theory, we noted that a youth may face a shorter queue when the minimum wage rises if the value of his productivity exceeded the minimum wage. This is more likely for youths who attend higher-quality high schools and who accumulate more human capital at home. Hence, minimum wages should have a positive impact on choice of most activities when interacted with S and H.

The results as reported in Table IV are quite good in general, with a weighted R^2 of .548. Magnitudes of the regression coefficients and associated t-statistics are given at the top of the table. The pattern of coefficients is reasonably consistent across equations. Technically, we could use equation (11) to reclaim the structural parameters. However, because of the complexity of the interaction terms, the estimated coefficients are easier to interpret if we compute the regression slopes (i.e., the first derivatives of the dependent variables with respect to each of the independent variables). For instance, the derivative with respect to mean income $(\mu_i - \mu_F)$ is reported in the first row below the dashed line of Table IV. If we look back to equation (11), we see that this derivative equals $\gamma(\delta_0 + \delta_1 S + \delta_2 A)$, which is the level of human capital embodied in the worker multiplied by the measure of constant absolute risk aversion. This term should be positive if agents are risk averse. Similarly, the derivative with respect to the second moment of the wage distribution equals $-\frac{\gamma^2 h^{*2}}{2}$, which must be negative if agents are risk averse. The derivatives

TABLE IV

MEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS AND ASSETS

	Professionals and Managers	Clerical and Sales	Craftsmen	Operatives and Laborers	Service	1-3 Year Schools	4 Year Schools	Other
$(\mu_i-\mu_F)$	22.7 (1.88)*	11.21 (1.65)*	16.70 (1.84)*	2.36 (.30)	-21.47 (1.57)	6.96 (2.24)**	.99 (.43)	31.19 (4.08)***
$(S)*(\mu_i-\mu_F)$	-.10 (1.69)*	-.02 (.79)	-.02 (.69)	-.01 (.23)	.15 (2.34)**	-.01 (1.45)	-.10 (.72)	-.10 (3.18)***
$H*(\mu_i-\mu_F)$	6.68 (1.50)	-.03 (.02)	1.13 (.52)	-.03 (.01)	-3.91 (.69)	.93 (1.12)	1.69 (1.88)*	2.12 (.81)
$(\sigma_i-\sigma_F)$	-11.54 (1.31)	-35.41 (2.46)**	-18.93 (1.05)	2.07 (.16)	35.20 (1.74)*	-31.81 (3.61)***	-2.67 (1.18)	-36.81 (3.10)***
$S*(\sigma_i-\sigma_F)$	.003 (.09)	.27 (2.37)**	.16 (1.25)	-.04 (.50)	-.28 (2.22)**	.11 (1.39)	-.002 (.18)	.16 (2.44)**
$(S)^2*(\sigma_i-\sigma_F)$	$1.5x10^{-4}$ (1.57)	$-6.2x10^{-4}$ (2.16)**	$-4.7x10^{-4}$ (1.71)*	$.6x10^{-4}$ (.38)	$4.3x10^{-4}$ (2.16)**	$-.8x10^{-4}$ (.37)	$.4x10^{-4}$ (1.04)	$-1.2x10^{-4}$ (1.49)
$H*(\mu_i-\sigma_F)$	-.48 (.18)	-7.46 (1.32)	-10.39 (1.53)	5.19 (1.02)	9.34 (1.08)	10.78 (2.22)**	1.36 (1.65)*	-2.90 (.64)
$H^2*(\sigma_i-\sigma_F)$	-.57 (.90)	-.14 (.13)	.30 (.30)	-.56 (1.08)	.27 (.35)	-.88 (.62)	-.57 (2.49)**	-.17 (.59)
$(S)*H*(\sigma_i-\sigma_F)$	-.003 (.31)	.03 (1.19)	.04 (1.39)	-.004 (.27)	-.03 (1.59)	-.02 (.83)	.00 (.08)	.005 (.64)
(MW_i-MW_F)	-11.48 (1.56)	-8.81 (1.95)*	-1.67 (.31)	-11.58 (3.05)***	-.25 (.02)	-11.90 (2.74)***	-9.02 (3.49)***	1.24 (.18)
$(S)*(MW_i-MW_F)$	.04 (1.17)	.04 (1.88)*	.03 (1.45)	.04 (2.59)***	-.02 (.40)	.04 (3.02)***	.03 (2.81)***	-.02 (.79)
$(H)*(MW_i-MW_F)$	-.24 (.08)	-.13 (.09)	-.51 (.33)	.93 (.95)	3.66 (1.69)*	1.32 (1.16)	1.00 (1.38)	2.81 (1.42)
County Dummies	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.
Intercept	-5.03 (5.27)***	-1.36 (2.42)**	-4.56 (4.87)***	.21 (.31)	-3.23 (4.61)***	-1.93 (3.30)***	.02 (.05)	2.11 (1.65)*

TABLE IV (Continued)

MEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS AND ASSETS

	Professionals and Managers	Clerical and Sales	Craftsmen	Operatives and Laborers	Service	1-3 Year Schools	4 Year Schools	Other
Estimated Derivatives:								
$\frac{\partial Z}{\partial(\mu_i-\mu_F)}$	11.05	6.64	13.26	.42	8.00	4.89	1.84	10.57
	(2.58)***	(2.08)**	(3.20)***	(.13)	(1.48)	(3.12)***	(2.75)***	(2.06)**
$\frac{\partial Z}{\partial(\sigma_i-\sigma_F)}$	-5.43	-5.73	-9.87	.69	-2.39	-4.34	-.32	-7.43
	(1.99)**	(1.98)**	(2.67)***	(.25)	(.51)	(3.44)***	(.67)	(1.75)**
$\frac{\partial Z}{\partial H}$	.97	-.09	-.11	.18	.28	.65	.52	.24
	(2.01)**	(.37)	(.38)	(.84)	(.91)	(2.56)**	(3.46)***	(.82)
$\frac{\partial Z}{\partial(S)}$	.0053	.0096	.0053	.0085	.0049	.0081	.0082	.0098
	(1.61)	(4.03)***	(1.73)*	(3.46)***	(1.29)	(3.11)***	(7.48)***	(3.41)***
$\frac{\partial Z}{\partial(MW_i-MW_F)}$	-1.92	-.13	4.79	-.90	2.71	1.05	-.30	.87
	(.90)	(.10)	(2.79)***	(.70)	(1.06)	(.58)	(.37)	(.44)

Joint Summary Statistics:

Weighted R^2 = .548	
F-test for $(\mu_i-\mu_F)$	2.86***
F-test for $(\sigma_i-\sigma_F)$	2.76***
F-test for H	1.60**
F-test for S	2.32***
F-test for (MW_i-MW_F)	4.53***

NOTE: Dependent variable: Z = log odds of selecting occupation i over farming. μ_i represents the mean of the distribution of earnings in occupation i. σ_i represents the second moment of the distribution of earnings in occupation i. S is a measure of real expenditures on teacher salaries per pupil, taken to proxy school quality in the country. H is a measure of the real value of taxable personal assets per person, taken to be a measure of household wealth in the county. MW_i is the minimum wage in occupation i.

(t-statistics in parentheses)
* significant at 10% level
** significant at 5% level
*** significant at 1% level

with respect to school and home inputs are $\gamma\delta_1(\mu_i-\mu_N) - \gamma^2\delta_1 h^*(\sigma_i-\sigma_N)$ and $\gamma\delta_2(\mu_i-\mu_N) - \gamma^2\delta_2 h^*(\sigma_i-\sigma_N)$ respectively, neither of which may be signed. The derivative with respect to minimum wages is $-(\phi\pi_i)/\beta$ plus the coefficients of the interaction terms times S and H respectively. This derivative cannot be signed, in general, since the sign of π_i is indeterminate. All of these derivatives were evaluated at the sample means of the variables. Each of these slopes may be interpreted as the impact of a unit increase in the independent variable on the log odds of choosing occupation i relative to farming. Asymptotic t-statistics are shown in parentheses below each slope.

As expected, an increase in mean income increases the likelihood of being in that activity. The slope with respect to $(\mu_i-\mu_F)$ is positive in all eight cases and is statistically significant (using standard criteria) for six of eight activities. The impact on selection of operative/labor and service jobs is not significant.

F-tests of joint hypotheses across all five equations were computed and reported at the bottom of Table IV. The first tests whether the three terms per equation which contain mean income $(\mu_i-\mu_F)$ are jointly zero across all equations. This test is of interest in that a variable may be an important determinant of choice of activity and yet happen to have a near-zero slope when evaluated at the sample means, owing to nonlinearities. The test statistic (F=2.86) allows us to reject the hypothesis that mean income differentials do not matter.

The derivatives with respect to the second moment of income are also reported in Table IV. Seven of the eight slopes are negative, as expected, and five of eight are significant. The major anomaly occurs in the operative/laborer occupation, which is positive but not significant. The F-test for the joint significance of all $(\sigma_i-\sigma_F)$ terms across equations (F=2.76) allows us to reject the hypothesis that the second moment does not matter. We take these results as strong evidence that choice of activity is subject to uncertainty and that higher-order moments of the wage distribution enter an agent's decision process.

Partial effects with respect to assets and school quality were also computed. All of these slopes except two were positive. The anomalies occur in the clerical/sales and crafts occupations for home inputs. Three of the asset slopes and five out of eight school quality slopes were statistically significant. The F-tests show that the joint hypothesis that school quality and home inputs do not enter the activity choice decision (i.e., the hypothesis that $\delta_1 = 0$ and $\delta_2 = 0$ in all equations respectively) can be rejected.

These results suggest that heterogeneity in human capital production due to differences in school quality does affect choice of activity. In particular, the increased human capital production in higher-quality schools induces greater entry into other occupations and into post-high school training relative to farming. In addition, there is evidence that human capital produced as a result of higher family assets increases the probability of choosing the professional/managerial occupation and of entering post-high school training. Because S and H do not have uniform effects on choice of activity, these results lead to the speculation that school and home inputs into human capital production may have occupation-specific as well as general training components.[19]

The set of 22 county dummy variables were jointly significant for each of the activities. Although not shown in Table IV, the orderings of the individual county dummies were quite consistent across activity groups. As one would expect, most young men enter nonagricultural occupations and enter college in the suburban counties of Anne Arundel, Baltimore, Montgomery, and Prince George's. Young men are more likely to enter agriculture in the rural counties of Garret, Kent, and Queen Anne's.

We are most interested in the impact of the minimum wage variable. As discussed above, our theory does not yield clear predictions as to the signs of the coefficients involving π_i.[20] However, our results have reasonable interpretations, with some exceptions.

The individual regression coefficients for MW are negative in seven out of eight cases and significant in four cases. This is quite reasonable for the occupational choice equations in the sense that previous studies have found that higher minimum wage rates and coverage decrease teenage employment. It is particularly notable that this coefficient is significantly negative for the operative/laborers category. The positive coefficient (1.24) for the "other" category is also reasonable if we interpret this group of youths (who reported "at home" and miscellaneous" activities) as primarily being unemployed. Harder to understand are the negative and significant coefficients on MW for the one-to-three year and four-year school categories. We would have expected positive signs in these regressions, in the sense that higher minimum wages should induce youths to enter vocational schools to overcome the employment barriers created by the law. On the other hand, it is possible that those more able youths who otherwise would have gone to vocational school find that they are more in demand (relative to less able youths, such as high school dropouts) and decide to take a job rather than continue in school.

The interaction term with school quality (S*MW) has a positive sign six out of eight times. As predicted, higher-quality schools allow their graduates to move up in the queue and take jobs in four of the five occupations. The negative, insignificant coefficient for the "other" category is also reasonable. The one-to-three year and four-year school categories have positive and significant coefficients, although it is less clear what sign should be expected.

The interaction terms with home inputs (H*MW) yielded mixed results. Five coefficients are positive and three are negative. None are significant at the 5% level. The positive coefficients for one-to-three year schools, four-year schools, and the "other" category may be interpreted as indicating that youths from higher-income families can better afford to enter these activities when minimum wages rise.

The derivatives with respect to each minimum wage index are also reported in Table IV. Three of five occupations have negative slopes, although none is significant. The large positive and significant slope for craftsmen is surprising and difficult to explain. The positive slopes for one-to-three year schools and for "other" activities are reasonable, as explained earlier, although they are not significant. Youths who can't get jobs may take additional vocational training or remain unemployed while searching for jobs. They may possibly go to four-year colleges and universities, although this is less likely than vocational schools and junior colleges. Indeed, the slope for four-year schools is negative, although also insignificant.

Overall, the joint F-test shown at the bottom of Table IV is significant (F=4.53) for all of the minimum wage terms. This indicates that minimum wages do play an important role in the choice of postgraduation activities. However, the underlying mechanism is complex and not easily summarized by the reported slopes.

The reader will recall that the minimum wage variable ($MW_i - MW_{farm}$) was treated specially for the professional/managers, one-to-three year and four-year schools, and "other" categories. Rather than setting $MW_i = 0$ for these four cases, MW_i was set equal to the nationwide minimum wage index. This was the index used in the Table IV regressions. We tested the sensitivity of our results by replicating Table IV but replacing MW_i with 0 for these four categories only. Not surprisingly, the results did not change much for the four occupations for which there was <u>no</u> change in the minimum wage index. For professional/managers, the slope became more negative and significant at the 10% level. For one-to-three year schools, the (H*MW) interaction became negative but insignificant while the slope became slightly negative. Likewise the slope for the "other" category became negative but insignificant. The slope for four-year colleges became slightly more negative. These results are consistent with those of Table IV in the sense that the expected slope is negative when using ($0-MW_{farm}$) whenever a positive sign is expected for ($MW-MW_{farm}$). We place more confidence in the Table IV results in that ($0-MW_{farm}$) is zero for 1951-1966 and nearly zero for 1967-1969.

Although the results in Table IV are quite strong, we were concerned that they may have been clouded by intercorrelation between home inputs and school quality, owing to the large number of interaction terms. In order to test this, we reestimated (11) without the home input variables and without the school quality variable. The results from the former regressions were nearly identical to those in Table IV and are not reported. The only significant difference was that the t-ratios with respect to school quality (S) generally increased in magnitude and were statistically significant in all eight equations. This supports our hypothesis that assets and school quality are competing with each other for explanatory power.

The results obtained by deleting the school quality variable are reported in Table V. The results are extremely strong. The second moment slope has the expected negative sign and is statistically significant all eight times. The slope with respect to mean income is positive and statistically significant all eight times. The home input slope is positive and statistically significant in all cases except crafts. The F-tests over all equations reconfirm that means, second moments, and assets each have significant impacts on choice of activity. This formulation provides even stronger support for the hypothesis that uncertainty matters and that agents are risk averse.

A comparison of the minimum wage (MW) coefficients between Tables IV and V reveals that, in most cases, the magnitudes of the coefficients diminish and that two categories (craftsmen and other) switch sign, although remain insignificant. In all eight cases, the interaction term (H*MW) is positive, as expected, and is significant five times. This reinforces our previous conclusion that youths from higher-income counties find it easier to enter all categories when the minimum wage rises.

There were some changes in the minimum wage slopes, however. The most important change is that the slope for one-to-three year schools increases

TABLE V

MEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS, ASSETS, AND MINIMUM WAGES

	Professionals and Managers	Clerical and Sales	Craftsmen	Operatives and Laborers	Services	1-3 Year Schools	4 Year Schools	Others
$(\mu_i-\mu_F)$	1.53	3.00	14.46	5.97	-1.38	2.91	-4.85	19.14
	(.23)	(.89)	(3.56)***	(1.64)	(.18)	(1.54)	(4.01)***	(4.02)***
$H*(\mu_i-\mu_F)$	5.20	1.62	.72	.18	6.62	.90	4.26	-4.18
	(2.07)**	(1.55)	(.53)	(.12)	(1.89)*	(1.41)	(7.17)***	(2.60)***
$(\sigma_i-\sigma_F)$	-2.90	-5.07	-8.08	-8.60	3.74	-19.23	-.32	-15.26
	(.59	(1.28)	(1.64)	(2.22)**	(.52)	(5.18)***	(.30)	(3.34)***
$H*(\sigma_i-\sigma_F)$	-1.44	.41	-2.76	3.56	-6.68	14.00	.79	4.67
	(.72)	(.13)	(.74)	(1.35)	(1.50)	(3.91)***	(1.07)	(2.03)**
$H^2*(\sigma_i-\sigma_F)$	-.37	.07	.61	-.62	.62	-2.92	-.73	-.28
	(1.21)	(.09)	(.79)	(1.33)	(.91)	(3.53)***	(4.83)***	(1.07)
(MW_i-MW_F)	-5.08	-2.53	1.00	-5.76	-6.95	-5.01	-6.81	-.13
	(1.11)	(1.12)	(.35)	(2.97)***	(1.64)	(1.93)*	(4.75)***	(.04)
$H*(MW_i-MW_F)$	2.99	1.79	1.55	3.24	4.60	4.49	3.62	1.18
	(1.51)	(1.98)**	(1.41)	(4.57)***	(2.70)***	(5.09)**	(6.29)***	(.80)
County Dummies	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.
Intercept	-5.31	-2.23	-5.20	-1.53	-3.96	-2.95	-.91	.92
	(6.60)***	(6.05)***	(9.39)***	(3.82)***	(9.85)***	(8.32)***	(4.03)***	(.85)
Estimated Derivatives:								
$\frac{\partial Z}{\partial(\mu_i-\mu_F)}$	11.26	6.03	15.81	6.31	11.00	4.59	3.10	11.31
	(2.76)***	(2.84)***	(6.19)***	(3.34)***	(3.69)***	(4.35)***	(5.24)***	(2.65)***
$\frac{\partial Z}{\partial(\sigma_i-\sigma_F)}$	-7.07	-4.04	-10.86	-4.37	-6.32	-4.54	-1.73	-7.64
	(2.79)***	(2.41)***	(5.53)***	(3.27)***	(3.19)***	(5.36)***	(4.44)***	(2.45)***
$\frac{\partial Z}{\partial H}$	1.26	.72	.41	.80	1.19	1.33	1.47	1.20
	(6.17)***	(4.38)***	(1.92)*	(4.75)**	(3.61)***	(7.12)***	(18.78)***	(5.59)***
$\frac{\partial Z}{\partial(MW_i-MW_F)}$	.50	.82	3.90	.29	1.64	3.38	-.04	2.08
	(.27)	(.86)	(3.26)***	(.34)	(1.03)	(2.67)***	(.07)	(1.39)

Joint Summary Statistics:

Weighted R^2 = .520

F-test for $(\mu_i-\mu_F)$	8.22***	F-test for H	13.41***
F-test for $(\sigma_i-\sigma_F)$	5.72***	F-test for (MW_i-MW_F)	6.94***

NOTE: See Table IV for definitions.

(t-statistics in parentheses)

*Significant at 10% level **Significant at 5% level ***Significant at 1% level in positive magnitude and becomes significant. This supports our hypothesis that youths enter vocational and other schools when minimum wages rise as a strategy to (1) enhance their human capital and (2) to wait for jobs to become available in the covered sector. On the other hand, there is no evidence that minimum wages increase four-year college enrollments.[21]

The minimum wage slope and associated t-ratio increase in magnitude for the "other" group, although remaining insignificant. Although still weak evidence, it is consistent with our expectations for this group of young men who remain at home and who, in our view, are primarily unemployed.

Craftsmen continue to have a positive and significant coefficient. There is the possibility that some youths may find it easier to take positions as apprentices or learners in crafts, relative to other jobs, following legislative changes. The Fair Labor Standards Act has continously provided a certification program under which "learners in occupations involving enough skill to necessitate an appreciable training period may be paid subminimum rates." Generally they have been paid 85% of the normal rate. Indeed, the U.S. Bureau of Labor Statistics (1970) reported: "Temporary increases in the number of certificates and learners authorized have occurred at the end of the fiscal year coincident with the effective dates of increases in the minimum wage to $1 in 1956, $1.15 in 1961, and $1.25 in 1963. No similar spurts occurred in 1967 and 1968 following the $1.40 and $1.60 rates." On the other hand, this BLS report indicates that the numbers involved were small, varying from 21,000 to 73,000 nationally. Of course, it is also possible that noncompliance with the law may be significant among apprentices and learners in the crafts area. Apprentices may feel that the law is not applicable or may be more willing to overlook it in return for the training. We do not claim that this explains all of our results for craftsmen, but it may contribute to the large positive slope.

The other four occupational categories continue to have insignificant minimum wage slopes, although three switch from negative to positive. Once again, we interpret this as an indication that the minimum wage had little adverse impact on high school graduate employment in Maryland. High school dropouts may have been more likely to have experienced difficulty finding jobs than graduates.

As before, Table V was rerun substituting $(0\text{-}MW_{farm})$ for $(MW_i\text{-}MW_{farm})$ for the professional/manager, one-to-three year and four-year schools, and the "other" category. This caused the operative/laborer and "other" minimum wage slopes to switch from positive to negative, although remaining insignificant. The slope for one-to-three year schools remained positive but lost significance.

Several additional tests of the sensitivity of our results were run. None of these have caused us to alter our conclusions. We reran Tables IV and V substituting a minimum wage index which eliminated all coverage effects. This index focuses on the ratio of the minimum wage to average industry wage rates. Because it rises with congressionally amended rates and falls in other years owing to inflation, this series eliminated much of the upward trend which is associated with the expanded coverage of 1961, 1967, and 1969. The resulting series was less highly correlated with the other explanatory variables. There was little impact on the minimum wage slopes and coefficients, however. If anything, the slopes were more likely to be positive than those of Table IV, but insignificant.

Inclusion of <u>both</u> this "pure" minimum wage index and a separate coverage variable had slightly more interesting effects. The separate coverage variable had positive and significant coefficients for one-to-three year and four-year schools and for "other" persons, as expected. Four of the five occupations also had positive coefficients, although only one was significant (services). The "pure" minimum wage index was more likely to be negative or closer to zero in this specification.

B. On-the-Job Training Effects

In all of the previous estimates, we have ignored on-the-job training (OJT) aspects of the model. In this section, OJT is reintroduced into our model by including the minimum wage (M_i) as a variable in the human capital equation ($h_i = \delta_0 + \delta_1 S + \delta_2 M_i$) replacing home inputs (H). We substituted M_i for H, rather than adding a third variable, because of the large number of interaction terms which result from including three terms in the human capital equation. Previous experience has indicated that collinearity becomes a severe problem when we have so many interactions; indeed, in several experiments the moment matrix would not invert.

This minimum wage variable (M) is the group's own minimum wage index and is <u>not</u> net of the farm minimum wage. We interpret "M" as controlling for the on-the-job training implications of the law. In particular, an increase in M should reduce training opportunities in the four occupations (clerical/sales; crafts; operatives/laborers; service) which are covered. We expect <u>no</u> OJT effect for the other four categories (professional/managers; one-to-three year and four-year schools; "other"), so $h = \delta_0 + \delta_1 S$ in these cases. No interaction terms involving M are included. All eight equations include two additional terms (MW and S*MW) as were included in Table IV. We refer to these as the queuing effects, as opposed to the OJT effects of the law. Of course, there is a question as to whether our data will allow us to disentangle these two effects, in practice.

Our regression estimates are presented in Table VI. The data and estimation methods are identical to those of Table IV except for the substitution of M_i or O for H and except for dropping the (M_i*MW) term, which would have introduced even more collinearity. The signs and magnitudes of the mean, second moment, and school quality slopes are very similar in Tables IV and VI. There is loss of significance in Table VI, particularly for the second moments for crafts and clerical occupations. Joint F-statistics, however, increase in value.

Our primary interest is in the minimum wage effects. First comparing the queuing effects ($\partial Z/\partial MW$), we see a great deal of similarity between Tables IV and VI. All signs and significance tests are identical, although magnitudes change slightly. Nothing here causes us to modify our earlier conclusions. We expected the OJT effects ($\partial Z/\partial M_i$) to be negative, and three of four terms are negative. None of the slopes, however, are close to conventional levels of significance individually. The joint F-test is significant at the 1% level over all terms involving M_i.

We ran some additional sensitivity tests of our model. First, we reran Table VI dropping the (S*MW) variable from all eight equations. This tended to reduce the collinearity problem and raise t-ratios. In particular, the "queuing" coefficient on the remaining MW variable became positive and significant for one-to-three year and four-year schools, as well as craftsmen. In addition, all four OJT slopes ($\partial Z/\partial M_i$) were negative as expected, including clerical/sales.

TABLE VI

MEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS, SCHOOL QUALITY, AND MINIMUM WAGES (BOTH QUEUE AND ON-THE-JOB EFFECTS)

	Professionals and Managers	Clerical and Sales	Craftsmen	Operatives and Laborers	Service	1-3 Year Schools	4 Year Schools	Others
$(\mu_i-\mu_F)$	18.2 (1.57)	11.9 (1.25)	5.63 (.32)	-5.19 (.40	-33.8 (1.80)	8.07 (2.61)**	.17 (.07)	30.81 (4.17)**
$(S)*(\mu_i-\mu_F)$	-.03 (.79)	-.02 (.53)	.014 (.24)	.03 (.51)	.20 (1.72)*	-.01 (1.36)	.01 (.75)	-.08 (3.99)**
$M_i(\mu_i-\mu_F)$	-- --	-1.49 (.08)	3.03 (.11)	-1.56 (.08)	-10.9 (.10)	-- --	-- --	-- --
$(\sigma_i-\sigma_F)$	-9.00 (1.05)	-28.8 (1.37)	.78 (.02)	28.4 (.66)*	37.4 (1.65)*	-35.46 (4.11)***	-1.85 (.84)	-36.65 (3.28)***
$S*(\sigma_i-\sigma_F)$	-.003 (.12)	.17 (1.60)	-.007 (.04)	-.08 (.63)	-.18 (1.43)	.21 (3.50)	-.005 (.46)	.14 (3.08)**
$(S)^2*(\sigma_i-\sigma_F)$	$.7x10^{-4}$ (1.44)	$-2.1x10^{-4}$ (.91)	$-.9x10^{-4}$ (.37)	$.1x10^{-4}$ (.06)	$.9x10^{-4}$ (.57)	$-3.3x10^{-4}$ (3.08)	$.1x10^{-5}$ (.06)	$-1.0x10^{-4}$ (2.24)**
$M_i(\sigma_i-\sigma_F)$	-- --	-25.5 (.25)	7.87 (.04)	-90. (.53)	-77. (1.19)	-- --	-- --	-- --
$M_i^2*(\sigma_i-\sigma_F)$	-- --	226. (1.21)	-6.74 (.02)	60. (.25)	136. (1.02)	-- --	-- --	-- --
$(S)*M_i^*(\sigma_i-\sigma_F)$	-- --	-.20 (.65)	-.19 (.35)	.17 (.50)	.09 (.38)	-- --	-- --	-- --
(MW_i-MW_F)	-11.41 (1.55)	-13.37 (1.41)	5.75 (.47)	-7.54 (.91)	.76 (.05)	-11.76 (2.71)***	-8.72 (3.31)***	-.017 (.00)
$(S)*(MW_i-MW_F)$	.04 (1.54)	.04 (1.54)	.01 (.30)	.03 (1.17)	.004 (.09)	.05 (4.58)***	.03 (4.27)***	.002 (.09)
County Dummies	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.
Intercept	-4.37 (5.39)***	-1.03 (1.37)	-5.08 (4.45)***	.13 (.15)	-3.12 (4.33)***	-1.39 (2.59)***	.27 (.98)	2.53 (2.07)**

TABLE VI (Continued)

MEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS, SCHOOL QUALITY, AND MINIMUM WAGES (BOTH QUEUE AND ON-THE-JOB EFFECTS)

	Professionals and Managers	Clerical and Sales	Craftsmen	Operatives and Laborers	Service	1-3 Year Schools	4 Year Schools	Others
Estimated Derivatives:								
$\frac{\partial Z}{\partial(\mu_i-\mu_F)}$	10.41	7.28	9.90	1.47	14.52	5.17	1.98	10.33
	--	(1.78)**	(1.87)**	(.37)	(1.91)	(3.31)**	(2.94)***	(2.03)**
$\frac{\partial Z}{\partial(\sigma_i-\sigma_F)}$	-5.33	-4.71	-7.10	.80	-5.53	-4.46	-.48	-7.20
	(1.97)**	(1.30)	(1.46)***	(.91)	(.96)	(5.53)***	(1.00)	(1.72)**
$\frac{\partial Z}{\partial M_i}$	--	.62	-1.41	-1.56	-1.83	--	--	--
	--	(.23)	(.43)	(.68)	(.61)	--	--	--
$\frac{\partial Z}{\partial(S)}$	.0105	.0091	.0042	.0091	.0093	.0127	.0111	.0117
	(5.73)***	(4.61)***	(1.49)	(4.34)***	(2.55)***	(6.44)***	(17.14)***	(4.89)***
$\frac{\partial Z}{\partial(MW_i-MW_F)}$	-1.52	-2.74	8.47	-.50	1.80	.97	-.20	.48
	(.71)	(.94)	(2.37)**	(.20)	(.51)	(.54)	(.23)	(.24)

Joint Summary Statistics:

Weighted R^2 = .534	
F-test for $(\mu_i-\mu_F)$	2.57***
F-test for $\sigma_i-\sigma_F)$	2.85***
F-test for M_i	3.01***
F-test for S	14.35***
F-test for (MW_i-MW_F)	5.69***

NOTE: See Table IV for definitions.

$\frac{\partial Z}{\partial M_i}$ represents the indirect effect of minimum wages on the log odds of selecting occupation i. $\frac{\partial Z}{\partial(MW_i-MW_F)}$ represents the direct effect of minimum wages on the log odds of selecting occupations i. The former is interpretted to be the on-the-job training effect and the latter is interpretted to be the queuing effect.

(t-statistics in parentheses)
* significant at 10% level
** significant at 5% level
*** significant at 1% level

Although not conclusive, we take this as evidence that the impact of the minimum wage on choice of activity is complex. Entry of high school graduates into occupations may have been reduced by the OJT effect but increased by the queuing effect.

Zellner "Seemingly Unrelated" Regressions for Women

Similar regression estimates were made for women. Some modifications had to be made owing to data limitations. First, there are relatively few women in certain occupations (crafts; farming), so the number of occupational activities dropped from six to four.[22] Second, this necessitated changing the reference group from farming to clerical/sales. Third, we did construct income means and second moments for women, using methodology similar to that for men. Unfortunately, such data for women are of questionable usefulness. The major problem is that income is reported for all women, including part-time and part-year workers. As a result, year-to-year variation as well as long-term trends may reflect changes in hours worked by women, rather than changes in the unit value of their human capital. We did attempt to adjust for changes in the average weeks worked per year, but with limited success.[23] As a consequence, we substituted the male income moments, which yielded somewhat more consistent results than the female income moments. This difficulty suggests that less confidence should be placed on the results for women. Nevertheless, we do believe that we can draw a few interesting, if tentative, conclusions. Table VII provides means and standard deviations of the variables.

Table VIII reports our Zellner regression estimates for the full model, which includes both school quality (S) and home inputs (H). The estimated

TABLE VII
DESCRIPTIVE STATISTICS FOR WOMEN

	Mean	Standard Deviation
Dependent Variable: $\text{Log}(P_i/P_C)$		
Professionals/Managers	-4.829	1.430
Operatives/Laborers/Crafts	-2.515	1.678
Service	-1.739	.782
1-3 Year Schools	-.962	.722
4 Year Schools	-.292	.641
Other (Unemployed)	-.331	.570
Differences in Income Means $(\mu_i - \mu_C)$		
Professionals/Managers	.246	.034
Operatives/Laborers/Crafts	-.119	.023
Service	-.236	.035
1-3 Year Schools	-.036	.086
4 Year Schools	.155	.095
Other (Unemployed)	-.768	.089
Differences in Income Second Moments $(\sigma_i - \sigma_C)$		
Professionals/Managers	.473	.066
Operatives/Laborers/Crafts	-.183	.052
Service	-.262	.070
1-3 Year Schools	-.046	.046
4 Year Schools	.259	.150
Other (Unemployed)	-.510	.115
School Quality (S)	254.549	67.588
Assets (H)	1.872	.659
Miniumum Wage Index $(MW_i - MW_C)$		
Professional/Managers	.024	.021
Operative/Laborers/Crafts	.085	.028
Service	-.057	.020
1-3 Year Schools	.024	.021
4 Year Schools	.024	.021
Other (Unemployment)	.024	.021

TABLE VIII

WOMEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS, SCHOOL QUALITY AND HOME INPUTS

	Professional and Managerial	Operatives, Laborers, and Crafts	Services	1-3 Year Schools	4+ Year Schools	Other
$(\mu_i-\mu_C)$	-21.6	16.1	29.9	4.45	2.42	9.02
	(.94)	(.36)	(1.95)*	(1.36)	(.83)	(2.43)**
$S*(\mu_i-\mu_C)$	.13	-.002	-.12	-.03	-.01	-.06
	(.93)	(.01)	(2.10)**	(2.37)**	(.85)	(3.88)***
$H*(\mu_i-\mu_C)$	-5.94	-8.97	2.26	.98	1.06	4.21
	(.44)	(.64)	(.48)	(1.23)	(.78)	(2.82)***
$(\sigma_i-\sigma_C)$	16.53	-6.73	-23.3	11.72	.08	-16.23
	(1.37)	(.21)	(1.34)	(1.64)	(.03)	(3.66)***
$S*(\sigma_i-\sigma_C)$	-.10	-.05	.05	-.03	.003	.11
	(1.32)	(.28)	(.59)	(.50)	(.20)	(4.20)***
$S^2*(\sigma_i-\sigma_C)$	$.0x4x10^{-4}$	$2.0x10^{-4}$	$.9x10^{-4}$	$.9x10^{-4}$	$.1x10^{-4}$	$-.6x10^{-4}$
	(.04)	(1.10)	(.95)	(.59)	(.37)	(2.75)***
$H*(\sigma_i-\sigma_C)$	1.04	9.30	3.80	-6.38	-1.18	-5.57
	(.13)	(.79)	(.62)	(1.20)	(1.33)	(2.18)**
$H^2*(\sigma_i-\sigma_C)$	-.33	.84	-.03	.16	.15	-.29
	(.45)	(.90)	(.07)	(.19)	(.88)	(2.42)**
$S*H*(\sigma_i-\sigma_C)$	.014	-.028	-.012	.007	.0002	.005
	(1.25)	(1.48)	(1.41)	(.38)	(.06)	(2.03)**
(MW_i-MC_C)	-52.8	-9.59	-28.5	-30.9	-35.1	-4.21
	(1.55)	(.69)	(2.09)**	(4.14)***	(6.05)***	(.46)
$S*(MW_i-MW_C)$	.12	.009	.11	.10	.10	-.02
	(.76)	(.12)	(2.11)**	(2.62)***	(3.23)***	(.48)
$H*(MW_i-MW_C)$	8.99	-2.44	-1.97	.11	1.33	4.27
	(.91)	(.44)	(.57)	(.04)	(.64)	(1.31)
County Dummies	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.
Intercept	-3.75	-.90	-1.04	-.11	.08	.61
	(4.28)***	(1.21)	(1.46)	(.71)	(.68)	(.66)

TABLE VIII (Continued)

WOMEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS, SCHOOL QUALITY AND HOME INPUTS

	Professional and Managerial	Operatives, Laborers, and Crafts	Services	1-3 Year Schools	4+ Year Schools	Other
Estimated Derivatives						
$\frac{\partial Z}{\partial(\mu_i-\mu_C)}$	.64	-1.32	3.89	-.99	.72	1.22
	(.09)	(.11)	(.56)	(1.17)	(0.92)	(.44)
$\frac{\partial Z}{\partial(\sigma_i-\sigma_C)}$	-1.83	.28	-3.43	31.7	-.02	1.59
	(.44)	(.03)	(.74)	(1.65)*	(.05)	(.63)
$\frac{\partial Z}{\partial H}$	.55	-.49	-.78	1.48	.13	-.35
	(1.31)	(.72)	(2.37)**	(.88)	(1.57)	(2.36)**
$\frac{\partial Z}{\partial S}$	-.0006	.0009	-.0037	-.0021	.0023	.0019
	(.16)	(.22)	(1.17)	(2.56)**	(3.65)**	(1.59)
$\frac{\partial Z}{\partial(MW_i-MW_C)}$	-5.71	-11.87	-3.24	-5.09	-7.73	-1.73
	(.94)	(2.17)**	(1.00)	(2.85)***	(6.45)***	(.89)

Joint Summary Statistics:

Weighted R^2 = .599	
F-test for $(\mu_i-\mu_C)$	2.54***
F-test for $(\sigma_i-\sigma_C)$	1.99***
F-test for H	1.90***
F-test for S	2.33***
F-test for (MW_i-MW_C)	

NOTE: See Table IV for definitions.

(t-statistics in parentheses)
* Significant at 10% level
** Significant at 5% level
*** Significant at 1% level

derivatives are less consistent in terms of sign and significance in contrast with the results for men (see Table IV). Four of six means are positive, as expected, but none are significant, while one-half of the second moment slopes are negative (none significant). Although we found primarily positive schooling and home input slopes for men, three of these slopes are negative and significant for women. On a more positive note, women in higher-quality schools are more likely to go on to four-year colleges.

The most consistent and significant derivatives occur for the minimum wage variable. It is interesting that all six slopes are negative and three are significant. This suggests that female high school graduates may have been hurt more by minimum wages than their male counterparts. The negative slopes for the two school groups could be interpreted as implying that women face (financial) barriers to using additional education as a means of overcoming the minimum wage constraint. It is less clear to us why the data suggest that women have flowed into clerical/sales jobs rather than home activities (included in "other").

The individual minimum wage coefficients are similar to those estimated for males in that MW is negative in all cases (and significant three times) while the interaction terms with S and H are positive in general. The positive interaction with school quality (S) is particularly notable, being positive in all five cases where we expect a positive sign and being significant in three of those cases. Like the men, female graduates of higher quality schools do better than graduates of lower-quality schools in the face of rising minimum wages.

In Table IX, we reestimate the model deleting the school quality variable. Even more than for men, this improves the estimated derivatives

TABLE IX
WOMEN: ZELLNER REGRESSIONS CONTROLLING FOR INCOME MOMENTS, AND ASSETS

	Professional and Managerial	Operatives, Laborers, and Crafts	Services	1-3 Year Schools	4+ Year Schools	Other
$(\mu_i-\mu_C)$	-11.54 (.69)	13.99 (.74)	14.58 (2.17)**	-1.74 (1.13)	-.15 (.08)	3.42 (1.16)
$H*(\mu_i-\mu_C)$	6.22 (.64)	-4.92 (.46)	-3.70 (1.32)	.08 (.12)	.86 (.88)	-1.08 (1.16)
$(\sigma_i-\sigma_C)$	6.46 (.71)	-10.18 (.79)	-12.06 (1.82)*	10.31 (2.54)**	.86 (.76)	-7.08 (2.27)**
$H*(\sigma_i-\sigma_C)$	-6.07 (1.07)	3.21 (.40)	4.13 (1.14)	-6.49 (2.02)**	-1.04 (1.46)	3.66 (2.53)**
$H^2*(\sigma_i-\sigma_C)$	.75 (1.80)*	-.11 (.14)	-.24 (.68)	.88 (1.42)	.19 (1.70)	-.41 (4.36)***
(MW_i-MW_C)	-26.96 (1.84)*	-2.89 (.33)	-10.13 (1.72)*	-20.70 (4.99)***	-21.81 (7.89)***	-3.04 (.63)
$H*(MW_i-MW_C)$	12.21 (1.65)*	-3.19 (.80)	4.23 (1.66)*	6.17 (2.78)***	6.41 (4.15)***	.56 (.23)
County Dummies	Inc.	Inc.	Inc.	Inc.	Inc.	Inc.
Intercept	-4.13 (6.12)***	-1.42 (2.24)**	-1.00 (2.22)**	-.21 (1.57)	-.14 (1.46)	.34 (.38)
Estimated Derivatives:						
$\frac{\partial Z}{\partial(\mu_i-\mu_C)}$	.11 (.02)	4.79 (.65)	7.64 (1.94)**	-1.59 (2.86)***	1.46 (2.45)***	1.40 (.52)
$\frac{\partial Z}{\partial(\sigma_i-\sigma_C)}$	-1.95 (.61)	-4.61 (.61)	-5.30 (2.20)**	1.62 (1.94)**	-.31 (.86)	-1.86 (.78)
$\frac{\partial Z}{\partial H}$	.59 (2.48)**	-.24 (.48)	-.52 (2.01)**	.68 (2.29)**	.30 (5.19)***	-1.13 (3.13)***
$\frac{\partial Z}{\partial(MW_i-MW_C)}$	4.10 (.89)	-8.87 (2.03)**	-2.22 (1.21)	-9.16 (6.38)***	-9.82 (10.46)***	-2.00 (1.04)

Joint Summary Statistics:

Weighted R^2 = .584	
F-test for $(\mu_i-\mu_F)$	2.18***
F-test for $(\sigma_i-\sigma_F)$	2.27***
F-test for H	3.72***
F-test for MW	16.79***

NOTE: Dependent Variables: Z = Log odds of selecting occupation i over Clerical/Sales
See Table IV for definitions. (t-statistics in parentheses)
*Significant at 10% level **Significant at 5% level ***Significant at 1% level

for means and second moments. The mean is positive five of six times and significant in two cases. It is difficult to explain the negative and significant result for 1-3 year schools. The second moment is negative five of six times, with 1-3 year schools as the exception, once again. The derivative with respect to home inputs remains mixed, with three positive signs (all significant) and three negative signs (two significant).

The minimum wage variable has similar effects in Tables VIII and IX. Overall, women appear to be hurt by minimum wages, although this is less true for women from more wealthy communities.

Structural Estimates

In order to summarize the information contained in the previous tables, it is useful to obtain estimates of the structural parameters, δ_{0i}, δ_{1i}, δ_{2i}, δ_{3i}, π_{0i} and π_{1i}.[24] The full model outlined in equations (7), (8), and (9) would have 50 parameters with nine activities. Given the complication of estimating such a large model using maximum likelihood methods, we impose further restrictions on the system. We impose the restriction that the higher-skill occupations (professional, crafts, and clerical) have the same structural parameters. Similarly, the lower-skill occupations (service; farming, fishing and lumbering; and operatives and laborers) have the same structural parameters. Finally, we reduce the complexity of the human capital production function by restricting the household asset effect, δ_{2i}, to zero. These restrictions reduce the number of structural parameters to 20. We have 10 production parameters, δ_{0J}, δ_{0C}, δ_{0H}, δ_{0L}, δ_{1J}, δ_{1C}, δ_{1H}, δ_{1L}, δ_{3H}, and δ_{3L}, where the subscripts are J $\equiv$ junior college or vocational school, C ≡ four-year college, H ≡ high-skill jobs and L ≡ low-skill jobs. δ_{3i} represents on-the-job training effects, which applies only to the high-skill clerical and craft occupations and to the low-skill occupations. We cannot obtain the production coefficients for the unemployed (U) group. Since earnings for this group are zero, $E(W_U^k) = 0$ for all k, which means that the terms δ_{0U} and δ_{1U} fall out of the equation.

We have 10 queuing parameters, π_{0U}, π_{0J}, π_{0C}, π_{0H}, π_{0L}, π_{1U}, π_{1J}, π_{1C}, π_{1H}, and π_{1L}. The vector π_{0i} gives the level effect of the minimum wage on the length of queue and π_{1i} represents the effect of school quality on the minimum wage effect. By inserting (9) into (7), we see that a positive (negative) sign on these parameters will imply an increased (decreased) length of queue and a lower (higher) probability of entry into the activity.

The estimates are reported in Table X. The results are generally consistent with those reported in the unconstrained estimation--increasing school quality increases activities--specific human capital for junior college, college, and high-skill occupations, but it lowers human capital for low-skill occupations. Therefore, increasing school quality lowers entry into the low-skill occupations as it increases entry into the other activities.

These parameters imply that the level of human capital embodied in entrants into the high-skill occupations is higher than that of entrants into the low-skill occupations. Consistent with this finding is the result that increasing the minimum wage further increases the human capital of the entrants into the high-skill occupations (presumably by increasing the availability of on-the-job training) but it decreases the human capital of

TABLE X

MAXIMUM LIKELIHOOD ESTIMATES OF THE STRUCTURAL EQUATIONS (7-9)

δ_{0J}	-11.80
δ_{0C}	-2.58
δ_{0H}	1.54
δ_{0L}	6.56
δ_{1J}	5.66
δ_{1C}	1.61
δ_{1H}	2.53
δ_{1J}	-2.55
δ_{3H}	12.62
δ_{3L}	-19.65
π_{0U}	2.68
π_{0J}	-6.51
π_{0C}	2.51
π_{0H}	7.67
π_{0L}	7.89
π_{1U}	-1.80
π_{1J}	1.42
π_{1C}	-1.65
π_{1H}	-4.37
π_{1L}	-5.40

Likelihood Value -4165.51

NOTES: δ_{0i}, δ_{1i}, and δ_{3i} are respectively the constant term, school quality term, and minimum wage term in the human capital production function. δ_{2i} is restricted to be zero. π_{0i} and π_{1i} are respectively the direct effect of minimum wages on length of queue and the interaction effect of school quality and minimum wages on length of queue. Activities include J for one-to-three year schools, C for four-year colleges, H for high-skill occupations, L for low-skill occupations and U for unemployed or other.

entrants into low-skill occupations. The interpretation is that the negative impacts of minimum wages on on-the-job training are concentrated in the lower-skill occupations, whereas increasing minimum wages will actually increase training opportunities for higher-skilled individuals in high-skill occupations.

In looking at the queuing coefficients, we find that minimum wages increase the incentives to go to junior college or vocational school, but reduce the incentives to go to college or to high- or low-skill occupations. Contrary to expectations, increasing the minimum wages appears to reduce the number going into unemployment and "other" activities as well. Increasing school quality acts to reverse all these affects. As school-generated human capital increases, the queue for high- and low-skill jobs is reduced and the incentives to attend four-year rather than two-year colleges increases. Once again, the anomaly is that a higher level of school quality appears to increase entry into unemployment and "other" activities. However, taken as a whole, these results indicate that increasing the minimum wage does affect queues for jobs: doing so increases queues for people from poorer schools and reduces queues for those from higher-quality schools. These results, coupled with those above for the on-the-job training effects, point to a compelling explanation for the frequent finding of differential impacts of minimum wages across various demographic groups. Variation in average levels of embodied human capital across demographic groups may well explain the variation in the direction and magnitude of the effects of minimum wages on employment and on-the-job training of those demographic groups.

Summary

Our major objective in this study has been to analyze the impact of minimum wages on the choice of activity of youths following high school graduation. There has been little previous research in this area. As such, our results must be interpreted as preliminary but suggestive.

One interesting finding is that the impact of the minimum wage law varies with quality of schooling and home environment. Graduates of better high schools, having more human capital, find it easier to take jobs when the minimum wage rises. Graduates of lower-quality high schools have a more difficult time finding nonagricultural jobs. Although our data set excludes them, by inference high school dropouts also suffer as the minimum wage rises.

A second important result is that junior college and vocational school enrollments increase in response to higher minimum wages and coverage. This may be interpreted as being due to efforts of youths to increase their human capital so as to surpass the minimum wage barrier or as a place to wait for acceptable jobs to become available. We find less evidence of an effect of minimum wages on entry into four-year colleges and universities. There is somewhat more evidence that minimum wages induce more youths to remain at home and undertake other nonemployment, nonschooling activities (such as unemployment status).

Third, we find some support for the notion that the adverse effects of minimum wages on on-the-job training reduce entry into the occupations in our data set which were covered by the law. We found generally consistent negative impacts of reduced training opportunities on occupational entry. While the individual effects are only marginally significant, the joint hypothesis of no on-the-job effect across regressions is strongly rejected.

If we interpret our results literally, we find that minimum wage laws have offsetting effects on high school graduates. Evaluated at the average level of school quality in our sample, the negative effect of minimum wages on on-the-job training tends to reduce the incentives to enter occupations. On the other hand, increasing the minimum wage appears to reduce the length of queue for jobs in this sample, encouraging occupational entry. The latter result implies that Maryland high school graduates have sufficiently high levels of human capital on average to benefit from increases in the minimum wage through an increased demand for skilled labor relative to unskilled labor.

Our results suggest that future studies of the minimum wage should use disaggregated data. The impact on young people varies with their accumulated level of human capital. Those with more human capital may be in a more advantageous position, while those with less human capital may be more constrained. This fact may serve to explain common findings of differential minimum wage effects across demographic groups. Future studies may also profit by including further schooling as an alternative to employment. We find evidence of increased enrollments in vocational schools and junior colleges resulting from higher minimums.

The policy implications are closely related to these observations. Our results reinforce the findings of other studies. Youths from richer families gain at the expense of youths from poorer families. We extend this by noting that youths from better schools also gain. In addition, we find stronger evidence than do other studies of additional schooling as an alternative to employment. If this schooling is productive, it will help to minimize the losses of economic efficiency caused by a minimum wage law. On the other hand, if schools are simply places to wait for job openings, then this strategy is no better than waiting at home.

NOTES

1. See Brown, Gilroy, and Kohen (1982) for a review of the theory.

2. See Hashimoto (1981, 1982) or Leighton and Mincer (1981) for evidence supporting this proposition.

3. Cunningham (1981), Ehrenberg and Marcus (1980), and Mattila (1981) all explore this question with different data sets. The results are not consistent across the studies so no general conclusions may be drawn.

4. See Linneman (1982) for evidence of this hours adjustment.

5. See note 2.

6. This exponential utility form is commonly used in the empirical literature on asset choice. See Deaton and Muellbauer (1980) for a review of the methodology.

7. More complex forms of the production function are admissible, although the estimated activity choice equation becomes overburdened with higher-order terms as more terms are added to the production equation.

8. Some might wonder if the Maryland data are representative of the United States. The following comparison indicates that they are quite similar. The Maryland figures are averages for high school graduates over the period 1951-70. The U.S. data are averages for males 18-19 having exactly four years of high school in 1960 (U.S. Bureau of Census, 1960 Census of Population, Educational Attainment, Report PC(2)-5B, Subject Reports, Table 8).

	Maryland	U.S.
Professionals & Managers	.020	.058
Clerical & Sales	.215	.251
Craftsmen	.101	.120
Operatives & Laborers	.426	.414
Service	.041	.076
Farm & Farm Laborers	.197	.081
	1.000	1.000

The discrepancy for "farmers" may be due, in part, to the fact that the classification scheme for Maryland includes fishing and lumbering in addition to farming, whereas the U.S. classification is exclusively farmers and farm laborers. The largest discrepancy for the other occupational categories is four percentage points.

9. A seventh occupational category in the published data is titled "Protective Service Workers Including Armed Forces." A substantial number of young men are in this classification, especially during the Korean War years and the late 1950s. Some decline in the early 1960s and a rise in the late 1960s suggests that variation in this category is largely exogenous, changing with national defense requirements and draft calls. Hence, we exclude these men from our data set. Our employed population is best interpreted as the employed civilian population.

10. These zero observations were retained since they provide useful information about occupational choice and since deletion could cause sample selection bias. Setting N = 0.1 in the expression Log(N/F), where F is the number of farmers and N is the number in occupation i, allows us to maintain a relative ranking as compared to the case in which N = 1.0 or higher. Other specifications (e.g., setting N=.01 or N=.001) yielded similar results. Correlations varied between .936 and .995 for dependent variables utilizing these approximations. In Table III, the only notable impact of changing N to .001 was to eliminate the significance of the asset derivatives for the craftsmen and service occupations.

11. We considered other measures such as substituting "total school expenditures" for "salaries and wages" and substituting the number of students "attending" for the number "belonging." Each pair was so highly correlated that this made little difference in the results. The six-year average covers the junior high and high school years. The measure is not greatly affected by varying the number of years in the moving average.

12. Once again, this measure is not highly sensitive to varying assumptions regarding the number of years in the moving average. The correlations were .990 or higher among the 2, 4, and 6 year moving averages.

13. For specifics, see Orazem and Mattila (1983). Identical methodology was used in estimating means and second moments for school attainment.

14. Assume that the distributions of human capital (h) and earnings per unit of human capital (W) are independent. If W and h each have moment-generating functions, the joint moment-generating function is M(W,h) = M(W,0)M(0,h) (see Hogg and Craig, 1970, p. 80). This implies that $\mu_{W,h} = \mu_W \mu_h$ and $\sigma_{W,h} = \sigma_W \sigma_h$, for the first two moments. Taking logarithms, we get

    (i) $\ln \mu_{W,h} = \ln \mu_h + \ln \mu_W$

    and (ii) $\ln \sigma_{W,h} = \ln \sigma_h + \ln \sigma_W$.

    Equation (i) suggests regressing the observed values of $\ln \mu_{W,j}$ on the observed values of $\ln \mu_h$ and then estimating μ_W as the anti-log of the residuals. This was our approach with the addition of an intercept, which corresponds to the constant or slowly changing distribution of age and experience within an occupation (see Orazem and Mattila, 1983 for more details).

15. Data on earnings by occupations for high school graduates were available for 1959 and 1969, providing the 1960s trend. Due to the lack of directly comparable data for 1949, the 1949-59 trend was established by analyzing changes in the income moment and education attainment moments.

16. Additional detail on the adjustment process is available in Orazem and Mattila (1983).

17. Coverage by industry 1951-1960 was based on data published in the 1954 Annual Report of the U.S. Wage and Hour and Public Contract Divisions.

Industry coverage 1961-69 was based on data published in the Annual Reports to Congress on Minimum Wage and Maximum Hour Standards under the FLSA. A major increase in coverage in construction occurred in 1956 due to court decisions which extended coverage beyond the 1953 interpretation of the U.S. Department of Labor. More minor adjustments in coverage were made due to court rulings in wholesale trade, services, and finance during the 1950s. In addition, care was taken to eliminate workers phased into coverage during 1969 from those covered during 1967. The published statistics for 1967 include both groups. Some other minor inconsistencies in the published data for agriculture and for finance, insurance, and real estate were rectified.

18. Once again, those phased into coverage in 1969 are separated out from those phased in under the 1967 amendments. We are grateful to Joyce Cooper for supplying the unpublished FLSA coverage data by industry and state for 1964. Similar data for 1967 were published in Minimum Wage and Maximum Hours Standards under the FLSA, 1967.

19. It may be hypothesized that these results are obtained merely because rural areas have lower levels of assets and school quality, and thus the probability of selecting farming is inversely related to S and H. However, the county-specific dummy variables should control for the rural versus urban nature of a county school district. Thus, the parameters on S and H should capture effects that are independent of fixed county characteristics.

20. We should note at the outset that in Linneman's (1982) study of the probability of membership in the subminimum population, the mean educational attainment for the population above the minimum wage was 12.1. If this result carries over to Maryland, then we would expect that the average high school graduate would benefit from increases in the minimum, especially those from higher-quality schools or wealthy families.

21. This may account for the mixed findings in the literature. Mattila, using time-series data, found that minimum wages significantly increased school enrollments in his 1979 paper, but found no significant impact on school enrollments in his 1982 paper. Ehrenberg and Marcus (1980) and also Cunningham (1981) find mixed results using cross-sectional data. We believe that our tests are more powerful in that (a) we can separate two-year school enrollments from four-year colleges, and (b) we have combined time-series and cross-sectional data. Cross-sectional data alone miss the impacts of changes in the law. Aggregate time-series data may hide real effects which are more apparent using disaggregated data.

22. The four groups are (1) professional and managers, (2) clerical and sales, (3) service (including private household workers) and (4) operatives, laborers, crafts, and farming.

23. Distributions of workers are available by sex and occupation for weeks worked per year and for full-time and part-time workers from the annual U.S. Bureau of Labor Statistics "Work Experience" surveys. Using fixed hours worked per week weights and weeks worked mid-points within each

category, we computed the mean and second moment of work weeks per year by occupation. These variables were then added to the regression equation in which we "purged" income of changes due to weeks worked as well as education. See note 14 above.

24. These estimates use the overall average minimum wage index for the college, junior college and vocational school, professional and managerial, and "other" groups. The occupation-specific minimum wage indices are used for the remaining groups.

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