MINIMUM WAGE EFFECTS ON EMPLOYMENT AND SCHOOL ENROLLMENT: REPLY TO EVANS AND TURNER

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ABSTRACT

In earlier work, we presented results suggesting that minimum wage increases have important consequences for both the employment opportunities of youths and their decision to enroll in school. In this paper, we show that the recent claim made by William Evans and Mark Turner that our results are sensitive to changes in the definition of the enrollment rate is based upon an analysis that uses a mismeasured minimum wage index. When the data are constructed properly, our original conclusions are not affected by changes in the enrollment definition.

I. Introduction

In an earlier paper (Neumark and Wascher, 1995a), we presented results suggesting that minimum wage increases may have important consequences for both the employment opportunities of youths and their decision to enroll (or remain enrolled) in school. In particular, using a conditional logit model of alternative employment and enrollment outcomes applied to state-level data from 1977 to 1989, we found that although the net disemployment effects from minimum wages are small for teenagers as a group, there is a noticeable decline in the teenage school enrollment rate and a significant increase in the proportion of teenagers who are neither in school nor employed. We further argued that our results were consistent with the hypothesis that there is a shift in labor demand toward higher-productivity teenagers after a minimum wage increase, an assertion given stronger support from similar research we conducted using micro-level data from matched CPS panels (Neumark and Wascher, 1995b).

Recently, William Evans and Mark Turner (1995) argued that our results using the state-level aggregate data are sensitive to changes in the definition of school enrollment. In particular, they claim that the negative effects of the minimum wage on school enrollment become statistically insignificant when they switch from an enrollment measure based on major activity (such as we used) to a measure that is measured independently of employment and thus may also include part-time school activity. Evans and Turner further claim that estimated minimum wage effects from the employment equations estimated in Neumark and Wascher (1992,

1994) also become insignificant when they reestimated the equations using their alternative definition of enrollment.

Our first inclination with respect to the enrollment results is to question the assertion that the expanded definition of enrollment used by Evans and Turner is preferable to the measure based on reported major activity that we had used. That is, while their definition of enrollment obviously is more inclusive than ours, we are skeptical that an enrolled teenager who reports his major activity as work should be considered in the same group as one who reports his major activity as school. This turns out, however, to be only a minor issue. What drives Evans and Turner's results is not the specific definition of school enrollment that they use, but rather their (mis)use of our minimum wage variable.

II. Reassessing the Evidence

To generate estimates of their enrollment rate back to 1978, Evans and Turner relied on the enrollment supplements included in the October files of the Current Population Survey (CPS) rather than the basic May files that we had used in our earlier papers. As they note, only the October survey included an independent question on school enrollment that far back. Consistent with this change in months, they also constructed from the basic October CPS files new values for the prime-aged male unemployment rate and the fraction of the working-age population between 16 and 19 years of age. However, they continued to use our minimum wage variable (from the dataset that we supplied to them), which was explicitly constructed to measure the coverage-adjusted level of the minimum wage in May of each year relative to the average wage. This minimum wage variable is inconsistent with their use of October data for the other variables for two reasons. First, some states raised their minimum wage between May and October of particular years. Second, given that nominal wages have generally risen over time, the denominator generally understates the average wage by an amount that likely varies by state and year.

Evidence on the Mismeasurement of the Minimum Wage Variable

The top panel of Table 1 documents the statutory increases in state minimum wage laws misclassified by Evans and Turner because of their use of the May minimum wage index; given the definition of the minimum wage index we used (the maximum of the state or federal minimum wage multiplied by federal coverage in the state and divided by the average wage in the state), the relevant misclassifications are for states with a minimum wage higher than the federal minimum.[1] In all, there are 17 such minimum wage changes, all of which occurred in the mid- to late-1980s. A number of these reflect a series of small changes in several New England states. However, there also were some large changes that occurred between May and October, including a 90 cent per hour increase in California's minimum wage in July 1988, two increases in Connecticut totaling 88 cents per hour in October 1987 and 1988, and a 50

1. The dataset also includes federal minimum wage increases in 1979, 1980, and 1981, which–as they occurred on January 1 of each year--are correctly classified by Evans and Turner. However, because all of the specifications include year dummy variables, it is likely that much of the identifying information with respect to variation in the nominal minimum wage comes from the state minimum wage increases.

cent increase in Oregon's minimum wage in September 1989. More importantly, over this sample period, there were 34 statutory state minimum wage increases, so that Evans and Turner misclassify the timing of exactly half of these. (We excluded the District of Columbia from this list--despite statutory changes in the minimum wage for some occupations--because the minimum wage for D.C. in our previous papers was constructed as a weighted average of occupation-specific minimum wages and thus some of the variation in the minimum wage reflects changes in the distribution of employment across occupations.) In general, we would expect this misclassification to bias the estimated effects toward zero, since the misclassified minimum wage increases are assigned not to the year in which they occurred, but to the following year.

We also examined Evans and Turner's mismeasurement of the average wage, by comparing estimates calculated from the October CPS files to those from the May data (Panel B). As Evans and Turner indicate, wage data for October are only available beginning in 1979, and so these comparisons are based on a slightly shorter sample period than we used in our original paper (and Evans and Turner used in their comment). As can be seen in the first column, using the raw data, the correlation between the May and October wage is relatively high (0.92), and only about 13 percent of the state-year observations display a differently-signed 12-month change in the wage. However, once state and year means are removed from the data (which occurs in the regression analysis), the similarity in the two wage measures drops markedly. As shown in the right column, the correlation coefficient for the wage levels falls to 0.29, and the proportion of observations showing differently-signed changes increases to 0.40.

The combined effects of the misclassification of state minimum wage changes and the mismeasurement of the average wage are presented in Panel C. In the raw data, the levels of the May and October coverage-adjusted relative minimum wage display a correlation coefficient of 0.81, while 28 percent of the observations show a differently-signed change over the year. When state and year effects are removed, the correlation coefficient for the levels falls to 0.28, and the proportion of mis-signed changes rises to nearly one-half. Thus, it is clear from the data that the measurement errors associated with Evans and Turner's use of the May minimum wage variable with October employment data could lead to spurious differences in results.

Implications for Employment-Enrollment Outcomes

To assess the impact of mismeasuring the changes in the minimum wage variable on the multinomial logit model of employment and enrollment outcomes, we reestimated the model using a newly constructed minimum wage variable that is specific to October of each year, so that--in contrast to Evans and Turner--all of the data refer to the same month. As in our earlier paper (and as repeated by Evans and Turner), we use a four-category model that classifies teenagers as in one of four mutually exclusive states: in school and not employed, in school and employed, not in school and employed, and not in school and not employed. In addition, we report results separately using the GLS estimator that we employed in Neumark and Wascher

(1995a) and using the OLS estimator reported by Evans and Turner. Finally, for ease of interpretation, we have translated the multinomial logit coefficients into elasticities, evaluated at the sample means of the data.

The basic results regarding enrollment and employment are presented in Table 2. In Panel A, we report the results from Neumark and Wascher (1995a). These results indicate a negative effect of minimum wages on the proportion enrolled in school and employed, although this effect is significant only using the GLS estimator and even then, only at the ten-percent level. The results also indicate a statistically significant positive effect of the minimum wage on the proportion of teens neither in school nor employed (i.e. the proportion "idle").

In Panel B, we report results using the mixed October and May data with the two enrollment measures, repeating the OLS results reported by Evans and Turner (rows 4 and 6), as well as our reestimates using the less restrictive covariance matrix that permits heteroscedastic errors across states and first-order autocorrelation in the residuals within states. Using the enrollment supplement schooling measure (rows 5 and 6), the estimated effects on enrollment and idleness are much smaller than when the May data are used, especially as compared with the estimates in row 1; this result is evident with either estimator. In contrast, as shown in row 4, the positive and significant minimum wage effect on the percentage of teens in the not enrolled, not employed category in the models using major activity to classify teens by enrollment status is statistically significant when the

OLS estimator is used (Evans and Turner did not focus on minimum wage effects on idleness, which we view as perhaps the most important finding in our original paper). When the errors are permitted to exhibit serial correlation and heteroscedasticity, however, the estimates for the models using major activity are similar to those using the independent enrollment question, and more importantly, neither set of results is similar to those in panel A. This pattern, by itself, raises questions about Evans and Turner's claim that using their alternative enrollment measure, rather than using October data, is the source of the difference between their estimates and those in Neumark and Wascher (1995a).

In Panel C, we report the minimum wage elasticities from the model reestimated with all data--including the minimum wage index--measured as of October of each year. Because the average wage used in the minimum wage index is available beginning only in 1979, the sample period for these estimates is two years shorter than in panels A and B. Nonetheless, the differences between panels B and C--and the similarities between panels A and C--are striking. Consistent with Evans and Turner's claim (based on the results in row 4 of Panel B), the estimates using the major activity variable to classify enrollment status (rows 7 and 8) yield results very similar to those we reported in our earlier paper (although the negative elasticity for the proportion enrolled and employed is smaller). In contrast to their results, however, the estimates using the enrollment supplement to measure schooling also are very similar to our previous estimates (rows 9 and 10). In particular, the minimum wage is estimated to have a positive and significant effect on the proportion of teenagers neither in school nor employed, regardless of the month, enrollment measure, or covariance matrix used in the model, as long as the data are measured in a time-consistent manner. As in our original paper on this topic, the statistical significance of the negative effects of minimum wages on enrollment is not robust; but this has little, if anything, to do with the alternative enrollment measures. More generally, these results strongly suggest that the different findings that Evans and Turner report are not due to their use of a broader measure of enrollment, as they claim, but rather to their failure to accurately measure the timing and magnitude of changes in the minimum wage index.

We further illustrate this finding by highlighting in bold, in Table 2, the (preferred) GLS estimates from our original paper (row 1), the estimates emphasized by Evans and Turner using the mixed May/October data and the enrollment supplement (row 6), and the GLS estimates using the enrollment supplement and consistent data for October (row 9). We think it is fair to say that as long as consistent data are used, the results are relatively insensitive to the enrollment measure, and are as we originally reported.

In addition to reestimating the multinomial logit model, Evans and Turner report results from a set of simple enrollment regressions relating the fraction of teenagers enrolled to the same set of independent variables used in Table 2. This specification, which we reproduce in table 3, provides a simpler test of the effects of minimum wages on enrollment. Based on the evidence presented in the right-hand column of panels A and B, they again argue that the results are not robust to the choice of enrollment measure. In particular, when they use the enrollment supplement measure of schooling with their mixed October/May data, the coefficient on the minimum wage variable is small and insignificant (row 3), in contrast to their results using the major activity measure of enrollment (row 2).

However, their result obtains only under the assumption of serially independent and homoscedastic errors. As shown in the left-hand column of panel B, when we relax these restrictions on the covariance matrix, the effect of minimum wages on school enrollment are significantly negative and of a magnitude similar to those in panel A, regardless of the enrollment measure used. It does appear, however, that the results for the enrollment equations are somewhat sensitive to changes in the sample. When all data are based on October values, the minimum wage coefficient drops substantially in the equations using the major activity measure of enrollment (row 4), but in contrast to the claims made by Evans and Turner (based on the mixed May/October data), the estimated effect of minimum wages on enrollment is stronger (and significant at the five-percent level) using the enrollment supplement measure of schooling (row 5). (As in Table 2, we have highlighted the most pertinent estimates.) Again, it is the mixing of May and October data, rather than the different enrollment measure, that underlies Evans and Turner's results.

Implications for the Employment Effects of Minimum Wages

The final set of results presented by Evans and Turner reexamine the employment equations for teenagers and young adults that we presented in our original 1992 paper and in our 1994 reply to the comment by Card, Katz, and Krueger (1994). As in the case of the enrollment equations, Evans and Turner compare our results using the May data to those using a combination of May and October data and argue that the negative coefficients on the minimum wage variable that we reported in these earlier papers are not robust to the inclusion or definition of the enrollment rate.

As we had reported in Neumark and Wascher (1995a)--and as shown in row 1 of Panel A of Table 4--there is little evidence of a significant negative employment effect from minimum wages in the teenage specifications that exclude the enrollment rate, a result that motivated our research on joint enrollment and employment decisions. As we have suggested elsewhere, however, negative employment effects of minimum wages when we control for enrollment are implied by the hypothesis that minimum wage increases induce employers to substitute towards enrolled teenagers (Neumark and Wascher, 1995c). Thus, taken at face value, Evans and Turner's results appear to undermine this hypothesis. However, as is made clear by the results in Panel B (which only report results from Evans and Turner's paper), the small and insignificant effects of minimum wages found by Evans and Turner using various alternative definitions of the enrollment rate are once again due to their mixing of May and October data rather than to the choice of enrollment measure. In particular, as shown in row 3, when the narrowest definition of the enrollment rate (based on the employment status recode) is included in the regression, the estimated minimum wage effects are negative and statistically significant using consistent data for May. However, when the October data are used with the May minimum wage variable, these estimated effects become small and insignificant (row 4), and are similar to the results obtained using the enrollment supplement measure of schooling (row 6).

Moreover, as can be seen in Panel C, using an October-based minimum wage index largely reverses Evans and Turner's results. In the teenage specifications, the minimum wage effect is of the same magnitude and is statistically significant with a p-value of less than 0.1 regardless of the enrollment definition; the coefficients and standard errors are also quite similar in the equations for young adults (with p-values close to 0.1), although the estimates in panel C are smaller than in panel A. Indeed, in this sense, the results using the October data probably are more intuitively appealing than those using the May data, as the minimum wage effect is uniformly larger for teenagers than for young adults. As in the previous tables, we have highlighted the preferred estimates from our earlier work (row 2), those emphasized by Evans and Turner (row 6), and those using the enrollment supplement with consistent October data (row 9). Once again, the results using either enrollment rate are quite similar as long as consistent data are used and run counter to the conclusions reached by Evans and Turner using the mixed May/October data.

Thus, the evidence suggests that the sensitivity of minimum wage effects reported by Evans and Turner stems from using a month in which to measure the minimum wage variable that differs from the month for which the dependent variable (and the other independent variables) are measured. This may seem surprising given the nature of the models, as the specifications allow for both contemporaneous minimum wage effects and effects lagged one year, and the use of May--rather than October--wage data only shifts the timing of the effects by six months. However, this shift in timing could be quite important if the contemporaneous minimum wage effects are most prominent, or if lagged effects are strongest at lengths of about one year.

Sources of the Instability in the Evans-Turner Results

Finally, we can examine whether the instability in the estimates using May/October data is more closely associated with Evans and Turner's misclassification of minimum wage increases or with their mismeasurement of the average wage. Table 5 presents estimates of equations based on the October enrollment supplement using different constructs of the minimum wage variable, alternatively allowing the numerator and denominator of the index to be mismeasured. Turning first to the multinomial logit results shown in Panel A, it is evident that both the misclassification of minimum wage changes and the mismeasurement of the average wage matter to some extent. In particular, correcting individually for the mismeasurements, in both cases, leads to an increase in the estimated effect of minimum wages on idleness, with the effect becoming statistically significant (compare the second or third row to the first row). However, it is also clear that the mismeasurement of the average wage is the more important source of error. For all four enrollment-employment categories, the estimates using the May minimum wage with the October average wage are quite close to the estimates using October data for both the numerator and denominator. This same pattern is apparent in the employment equations (Panel C). In contrast, in the school enrollment equations (Panel B), the estimates are sensitive to both the misclassification of the minimum wage changes and the mismeasurement of the average wage.

III. Conclusion

In sum, a reexamination of the evidence presented by Evans and Turner indicates that there is no reason to believe that our earlier estimates of the effects of minimum wages on the enrollment and employment outcomes of teenagers were unduly influenced by our use of the major activity definition of school enrollment rather than the more inclusive definition available in the October CPS enrollment supplement. Instead, the small and statistically insignificant minimum wage effects produced by Evans and Turner result from their mismeasurement of the minimum wage index used in these studies. We thus stand by our original conclusions that increases in the minimum wage result in substitution by employers toward enrolled teenagers and in a significant increase in the proportion of nonenrolled teenagers without a job.

References

Card, David, Alan Krueger, and Lawrence Katz. 1994. "Employment Effects of Minimum and Subminimum Wages: Comment." Industrial and Labor Relations Review. Vol. 47, No. 3, pp. 487-497.

Evans, William and Mark Turner. 1995. "Minimum Wage Effects on Employment and School Enrollment: Comment." Mimeo. University of Maryland.

Neumark, David and William Wascher. 1992. "Employment Effects of Minimum and Subminimum Wages: Panel Data on State Minimum Wage Laws." Industrial and Labor Relations Review, Vol. 46, No. 1, pp. 55-80.

________________. 1994. "Employment Effects of Minimum and Subminimum Wages: Reply to Card, Katz, and Krueger." Industrial and Labor Relations Review. Vol. 47, No. 3, pp. 497-512.

________________. 1995a. "Minimum Wage Effects on Employment and School Enrollment." Journal of Business and Economics Statistics. Vol. 13, No. 2, pp. 199-208.

________________. 1995b. "The Effects of Minimum Wages on Teenage Employment and Enrollment: Evidence from Matched CPS Surveys." Research in Labor Economics. forthcoming.

________________. 1995c. "Reconciling the Evidence on Employment Effects of Minimum Wages: A Review of our Research Findings," Finance and Economics Discussion Series paper no. 95-53, Federal Reserve Board, December, 1995.

Table 1: Consequences of Using May Minimum Wage Variable for October

A. Misclassified Minimum Wage Increases

State	Minimum Wage Increase	Date of Increase
Massachusetts	$.20	July 1, 1986
Rhode Island	$.20	July 1, 1986
Vermont	$.10	July 1, 1986
Massachusetts	$.10	July 1, 1987
Rhode Island	$.10	July 1, 1987
Vermont	$.10	July 1, 1987
Connecticut	$.38	October 1, 1987
California	$.90	July 1, 1988
Massachusetts	$.10	July 1, 1988
Rhode Island	$.35	July 1, 1988
Vermont	$.10	July 2, 1988
Connecticut	$.50	October 1, 1988
Wisconsin	$.30	July 1, 1989
Vermont	$.10	July 2, 1989
Rhode Island	$.25	August 1, 1989
North Dakota	$.05	August 14, 1989
Oregon	$.50	September 1, 1989

Total number of state minimum wage increases, May 1978-October 1989:	34
Number incorrectly assigned to year after increase:	12
Number incorrectly omitted:	5

B. Mismeasured Average Wage

	Raw Data	Net of Year and State Effects
Correlation between May and October average wage	.92	.29
Proportion of observations with different-signed May-May and October-October changes in average wage	.13	.40

C. Overall Effects on Coverage-Adjusted Relative Minimum Wage

	Raw Data	Net of Year and State Effects
Correlation between May and October	.81	.28
Proportion of observations with different-signed May-May and October-October changes	.28	.45

Panel A excludes Washington, D.C. Changes in state minimum wages below the federal minimum wage are also ignored. Note that the minimum wage increases occuring in 1989 were not included in the data set used in Neumark and Wascher (1995a, 1992, 1994), because those data extended through May of 1989. However, these increases occurred by October 1989, and hence should have been included in the data used by Evans and Turner.

Table 2: Estimated Minimum Wage Effects (Current and Lagged) on Employment and Enrollment Status, Elasticities from Conditional Logit Estimates, 16-19 Year-Olds

Enrolled,	Enrollment Measure	Enrolled, Not Employed	Enrolled, Employed	Not Enrolled Employed	Not Not Employed
A. Earlier Results 1978-1989, all data from May:					
1) AR(1), heteroscedastic errors	***Major activity***	***-.02*** ***(.18)***	***-.47**** ***(.29)***	***.14*** ***(.26)***	***.64***** ***(.15)***
2) Independent, homoscedastic errors	Major activity	-.15 (.24)	-.17 (.38)	.12 (.12)	.60** (.20)
B. Results with Mixed May/October Data 1978-1989, enrollment and employment data from October, Minimum wage index from May:					
3) AR(1), heteroscedastic errors	Major activity	.00 (.21)	-.40 (.33)	.28 (.29)	.16 (.17)
4) Independent, homoscedastic errors	Major activity	-.05 (.27)	-.54 (.40)	.35 (.39)	.46** (.22)
5) AR(1), heteroscedastic errors	Enrollment supplement	.03 (.21)	-.13 (.29)	.04 (.34)	.15 (.17)
6) Independent, homoscedastic errors	***Enrollment supplement***	***-.02*** ***(.27)***	***-.04*** ***(.37)***	***-.07*** ***(.43)***	***.33*** ***(.22)***
C. New Results 1980-1989, all data from October:					
7) AR(1), heteroscedastic errors	Major activity	-.04 (.14)	-.19 (.22)	-.09 (.23)	.74** (.12)
8) Independent, homoscedastic errors	Major activity	.07 (.22)	-.28 (.33)	-.06 (.33)	.41** (.18)
9) AR(1), heteroscedastic errors	***Enrollment supplement***	***-.12*** ***(.16)***	***-.33*** ***(.23)***	***.43**** ***(.26)***	***.64***** ***(.13)***
10) Independent, homoscedastic errors	Enrollment supplement	.05 (.22)	-.24 (.31)	-.08 (.37)	.54** (.19)

'*' and '**' indicate statistically significant at the 10 and 5 percent levels, respectively. Earlier results are from Neumark and Wascher (1995a). The estimates in row 1 correct a slight programming error in that paper that Bill Evans and Mark Turner pointed out; the results are qualitatively similar. The estimates in Panel B are based on our specifications using the data in Evans and Turner (1995).

Table 3: Estimated Minimum Wage Effects (Current and Lagged) on Enrollment Rates, 16-19 Year-Olds

	Enrollment Measure	% In School: AR(1), heteroscedastic errors	% In School: Independent, homoscedastic errors
A. Earlier Results 1978-1989, all data from May:			
1)	***Major activity***	***-.26***** ***(.12)***	-.27* (.15)
B. Results with Mixed May/October Data 1978-1989, enrollment and employment data from October, Minimum wage index from May:			
2)	Major activity	-.25** (.13)	-.33** (.15)
3)	***Enrollment supplement***	-.20* (.11)	***-.08*** ***(.13)***
C. New Results 1980-1989, all data from October:			
4)	Major activity	-.06 (.10)	-.07 (.13)
5)	***Enrollment supplement***	***-.18***** ***(.09)***	-.09 (.11)

'*' and '**' indicate statistically significant at the 10 and 5 percent levels, respectively. Earlier results are based on the data in Neumark and Wascher (1995a).

Table 4: Estimated Minimum Wage Effects (Current and Lagged) on Employment Rates

	Enrollment Measure	16-19 Year-Olds: Enrollment Excluded	16-19 Year-Olds: Enrollment Included	16-24 Year-Olds: Enrollment Excluded	16-24 Year-Olds: Enrollment Included
A. Earlier Results					
1973-1989, all data from May:					
1)	Employment status recode	-.04 (.13)	-.24** (.08)	-.29** (.09)	-.28** (.07)
2)	*Major activity*	...	*-.13* *(.11)*	...	*-.27*** *(.08)*
B. Evans/Turner Results					
1978-1989, all data from May:					
3)	Employment status recode	-.06 (.15)	-.23** (.10)	-.16 (.10)	-.24** (.08)
1978-1989, enrollment and employment data from October, Minimum wage index from May:					
4)	Employment status recode	-.08 (.14)	-.13 (.10)	-.14 (.10)	-.06 (.08)
5)	Major activity	...	-.18* (.13)	...	-.07 (.09)
6)	*Enrollment supplement*	...	*-.09* *(.14)*	...	*-.04* *(.10)*
C. New Results					
1980-1989, all data from October:					
7)	Employment status recode	-.19 (.12)	-.17** (.08)	-.15* (.09)	-.13* (.07)
8)	Major activity	...	-.21* (.11)	...	-.13 (.08)
9)	*Enrollment supplement*	...	*-.22** *(.12)*	...	*-.14** *(.08)*

'*' and '**' indicate statistically significant at the 10 and 5 percent levels, respectively. Earlier results are from Neumark and Wascher (1992, 1994). Evans and Turner results are from Evans and Turner (1995). In Panel C, the data begin in 1980 because wage data in October are first available in 1979.

Table 5: Effect of Misclassified Minimum Wages and Mismeasured Average Wages, Using Enrollment Supplement

A. Table 2 Results

Minimum Wage Variable	Enrolled, Not Employed	Enrolled, Employed	Not Enrolled Employed	Not Enrolled, Not Employed
May minimum, May average (Row 5)	.03 (.21)	-.13 (.29)	.04 (.34)	.15 (.17)
October minimum, May average	-.02 (.20)	-.03 (.29)	-.10 (.31)	.27* (.17)
May minimum, October average	-.10 (.18)	-.34 (.25)	.40 (.29)	.64** (.15)
October minimum, October average (Row 9)	-.12 (.16)	-.33 (.23)	.43* (.26)	.64** (.13)

B. Table 3 Results

Minimum Wage Variable	% In School
May minimum, May average (Row 3)	-.20* (.11)
October minimum, May average	-.33** (.10)
May minimum, October average	-.12 (.10)
October minimum, October average (Row 5)	-.18** (.09)

C. Table 4 Results

Minimum Wage Variable	16-19 Year-Olds, Employment	16-24 Year-Olds, Employment
May minimum, May average (Row 6)	-.09 (.14)	-.04 (.10)
October minimum, May average	-.04 (.15)	-.02 (.10)
May minimum, October average	-.21* (.12)	-.13 (.08)
October minimum, October average (Row 9)	-.22* (.12)	-.14* (.08)

All data other than relative minimum wage variable are based on October CPS's. All specifications use enrollment supplement to measure enrollment, and allow AR(1), heteroscedastic errors. '*' and '**' indicate statistically significant at the 10 and 5 percent levels, respectively.