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Real Wages over the Business Cycle: Some Further Evidence*

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I. Introduction

Over the past 50 years economists have been puzzled by the movement of real wages over the business cycle. Neoclassical and early Keynesian theories advanced the view that real wages are countercyclical, varying inversely with employment and output over the business cycle [12; 13]. On the other hand, early empirical evidence by Dunlop [9] and Tarshis [25] suggested that in fact real wages move procyclically. The macroeconomic disequilibrium analysis of Barro and Grossman [3] provided a theoretical explanation for the possibility of procyclical real wages. Despite this theoretical work, numerous recent empirical tests of the movement of real wages over the business cycle have reached no definite conclusions.[1]

One reason these studies have reached no definite conclusion lies in the fact that they have been performed at the aggregate level.[2] They have ignored the possibility that the age-sex-race composition of the labor force may change considerably over the course of the business cycle. But the human capital theory of Becker [4] and Oi [21] and the more recent implicit contract theory of Azariadis [1] and Baily [2] predicts that employment shares of different demographic groups will vary over the business cycle. Consider the consequences of changing employment shares. Suppose, for instance, that firms tend to lay off lower skill and/or less senior workers during cyclical downturns and retain workers who, on the average, are higher skilled and/or more senior than the workers laid off. The result will be that the quality of the workers still employed will rise. As a consequence the aggregate real wage, which is typically constructed by dividing the aggregate straight-time payroll by aggregate straight-time manhours, will tend to rise even when no real wage growth has in fact occurred per unit of quality adjusted labor. To reverse the example, during cyclical expansions employment increases may come disproportionately from among lesser skilled, less senior workers. Because employment is becoming more skewed towards lower paid workers, the aggregate real wage may show a decline even when no workers have actually

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1. Studies by Neftci [20], Otani [22], Canzoneri [7], Chirinko [8], Mehra [18], and Leiderman [17] provide evidence of countercyclical real wages. However Kuh [16], Bodkin [5], Modigliani [19], and Sims [23] provide evidence of procyclical real wages.

2. An exception is Mehra [18], who examined the behavior of real wages at the disaggregated industry level.

experienced a real wage reduction. If in fact employment shares behave in such a fashion, the direction of the aggregation bias in previous real wage studies will be to find an inverse relationship between the real wage and the level of output and employment or, where studies have found a positive relationship, to understate the extent of the relationship. Of course, employment shares of different demographic groups may behave differently, although the above discussion is based on prevailing theory.

In this paper we first demonstrate the nature of this aggregation bias theoretically. Next, we demonstrate that employment shares of different subgroups are differentially affected by the business cycle. Evidence is then presented that real wages tend to move procyclically rather than countercyclically.

II. A Model of the Aggregate Real Wage

In this section, we develop a model of the aggregate real wage and we derive the conditions that are required for there to be no aggregation bias in aggregate real wage studies. To begin with, assume that the aggregate real wage W/P is a weighted average of the real wage of different subgroups where the weights, which sum to one, are the employment shares of the different groups:

$$W/P = \sum s_i (W/P)_i \tag{1}$$

Now suppose that the real wage of each subgroup is a linear function of a vector of exogenous factors that influence the wages of all subgroups (Z) and a vector of factors unique to each subgroup (X_i):

$$(W/P)_i = a_i + b_i Z + c_i X_i . \tag{2}$$

Included in Z are the state of the economy and trend. Factors in X_i will be discussed later. Therefore, the aggregate real wage may be written as

$$W/P = \sum s_i [a_i + b_i Z + c_i X_i]. \tag{3}$$

Now consider what happens to the aggregate real wage as the elements of Z and X_i change. (Implicitly, the employment shares are also functions of Z and X_i, but to keep the notation simple we will not specify an equation for s_i similar to equation (2) for $(W/P)_i$). Taking the total differential of equation (3) and allowing the employment shares to change, we obtain,

$$d(W/P) = \sum (W/P)_i \, ds_i + \sum s_i (b_i \, dZ_i + c_i \, dX_i). \tag{4}$$

The result of equation (4) shows that the aggregate real wage can change (1) when employment shares change even though each group's real wage remains unchanged and/or (2) when the elements of Z and X_i change, thus affecting real wages of the various subgroups.

The result in equation (4) may be compared with the result from aggregate real wage studies. These studies generally specify the aggregate real wage W/P as a linear function of Z, i.e., $W/P = a + bZ$. Differentiating this equation with respect to Z, we have that $d(W/P) = b\,dZ$. Note that equation (4) reduces to this simple result only under several rather restrictive conditions. First, since $\Sigma s_i = 1$, and $\Sigma ds_i = 0$, the first term in (4) will vanish only if the initial real wage $(W/P)_i$ is the same for every group (so that $\Sigma(W/P_i)\,ds_i = (W/P)\Sigma ds_i = 0$) or each $ds_i = 0$. Second, it must hold that $b_i = b$, $i = 1 \ldots n$. That is, the real wages of the different subgroups must respond equally to changes in Z. Third, since aggregate specifications do not allow special factors X_i to have an influence on the real wages of different subgroups, it must therefore hold from equation (4) that $\Sigma s_i c_i dX_i = 0$. All of these would appear to be rather restrictive assumptions. We therefore conclude that aggregate real wage studies are likely to be biased because (1) shifts in employment shares can cause the aggregate real wage to change even when the real wages of different groups remain unchanged, (2) the real wages of different groups are unlikely to be equally responsive to changes in the factors that affect all groups, and (3) special factors that affect the real wage of different groups are unlikely to cancel out at the aggregate level.

We now explore a slightly different specification of our model. Note that the linear specification of the real wage equations of the different subgroups may impose an overly restrictive condition for the existence of no aggregation bias. Recall that one of the conditions is that $b_i = b$, $i = 1, \ldots, n$. This condition requires that the real wages of the different groups respond by the same *dollar* amount to changes in Z. A less restrictive formulation would require that the real wages respond to changes in Z by the same *percentage* amount rather than the same dollar amount. Such a formulation follows if we specify the real wages of the different groups are exponential functions of Z and X_i:

$$(W/P)_i = e^{a_i + b_i Z + c_i X_i}. \tag{5}$$

Following our earlier procedure of substituting this real wage function into (1) and again taking the total differential of $W/P = \Sigma s_i(W/P)_i$, we obtain

$$d(W/P) = \sum (W/P)_i\, ds_i + \sum s_i b_i (W/P)_i\, dZ + \sum s_i c_i (W/P)_i\, dX_i. \tag{6}$$

If, in fact, the aggregate real wage is assumed to be $W/P = e^{a+bZ}$, with the total differential $d(W/P) = b(W/P)\,dZ$ then the absence of an aggregation bias requires (1) $\Sigma(W/P)_i\,ds_i = 0$, (2) $\Sigma s_i b_i (W/P)_i\,dZ = b(W/P)\,dZ$, and (3) $\Sigma s_i c_i (W/P)_i\,dX_i = 0$, conditions which are similar to those developed earlier. The second condition requires that $b_i = b$, $i = 1, \ldots, n$, which implies that the percentage change in the real wage for a change in Z be the same for all groups.

III. Empirical Analysis

The discussion in the previous section is the motivation for the empirical analysis. We argued that previous aggregate real wage studies may be biased because they neither allowed for the effect that changing employment shares have on the aggregate real wage or for the bias that results if the change in real wages due to a change in the state of the economy or other common exogenous variables is not the same for each employment group. In the next two subsections we first investigate the extent to which employment shares of different age-sex-race groups depend on the state of the economy, time, and on group specific factors discussed below. We then estimate the effect of these factors on the real wages of various age-sex-race groups, and we test whether there are statistically significant differences in the coefficients for exogenous variables common to all groups.

An Empirical Analysis of Employment Shares over the Business Cycle

In this section we report the results of the effect that the business cycle has on the employment shares of various demographic groups in the economy. To our knowledge, only Kosters and Welch [15] have attempted such a study. They looked at employment shares for eight groups corresponding to the interaction of the following three classifications: age (teenage-adult); color (white-non-white); and sex. Their results are consistent with the employment patterns of various groups of workers being differently affected by the business cycle. We extend their study by examining a more detailed breakdown by age, race, and sex. Specifically, we examine the age categories 16–19, 20–24, 25–34, 35–44, 45–54, 55–64, 65 and older, for white males, black males, white females, and black females. This 28 category breakdown allows us to examine the business cycle effect at a lower level of aggregation.

For each of the 28 subgroups, an employment share equation of the following form was estimated using annual data for the period 1954–1981.

$$\Delta s_i = \alpha_{0i} + \alpha_{1i} PDGNP + \alpha_{2i} PDPOP_i + \alpha_{3i} TIME + e_i . \tag{7}$$

In this equation, Δs_i measures the yearly change in the employment share of group *i*. *PDGNP* is the yearly percentage increase in real GNP minus the trend growth in real GNP, which was 3.5 percent per year for the period 1954–1981.[3] *PDGNP* measures the rate at which the economy is currently expanding relative to normal. For reasons discussed earlier we expect that positive values of *PDGNP* (the economy is expanding faster than trend) will increase the employment shares of young workers. On the other hand, the employment shares of older workers, white workers, and male workers will tend to increase when *PDGNP* is negative (the economy is expanding more slowly than normal).

In equation (7) $PDPOP_i$ measures the percentage change in the population share of group *i* from one year to the next. The predicted sign of the coefficient is positive. As the population of a group increases (decreases), its labor supply will also likely increase (decrease), and consequently so will that group's share of total employment.

Finally, we include *TIME* as a proxy for any other variables that systemically affect employment shares. No prediction is made as to the sign of this variable for the various subgroups.

Estimation of the employment share equations proceeded as follows. Because the errors may be correlated both over time and across equations, we adopted the procedure recommended by Kmenta [14, 529–530]. Each share equation was first estimated by Ordinary Least Squares and the Durbin-Watson statistic was obtained. The equations for which autocorrelation was a problem were transformed and then in a second stage all of the equations were estimated as a system using Zellner's seemingly unrelated regression technique.

The empirical results are in Table I. The variable *PDGNP* is statistically significant in 13 of the 28 employment share equations. *PDGNP* coefficients for 8 of these 13 groups (WM 16–19, WM 20–24, BM 16–19, BM 20–24, BM 35–44, BM 45–54, WF 16–19, and BF 20–24) are positive while *PDGNP* coefficients for the other 5 groups (WM 45–54, WM 55–64, WF 45–54, WF 55–64 and BF 45–54) are negative. Young workers comprise 6 of

3. Trend growth was estimated from a regression of the natural logarithm of real GNP on time for the period 1954–1981.

Table I. Regression Results for Employment Shares.

GROUP	INTERCEPT	*PDGNP*	$PDPOP_i$	*TIME*
WM 16–19	0.0005727	0.043059	−0.0000343	0.016431
	(1.29)	(5.45)[c]	(−1.11)	(3.07)[c]
WM 20–24	0.0011783	0.035090	−0.0000408	0.020671
	(1.74)[a]	(2.89)[c]	(−1.09)	(2.39)[b]
WM 25–34	−0.0030194	0.006407	0.0002666	0.002516
	(−4.25)[c]	(0.46)	(4.38)[c]	(0.23)
WM 35–44	−0.0013028	−0.016831	0.0000544	−0.001469
	(−2.29)[b]	(−1.52)	(0.82)	(−0.18)
WM 45–54	0.0007790	−0.028096	−0.0001844	−0.005824
	(2.04)[b]	(−3.44)[c]	(−5.74)[c]	(−0.89)
WM 55–64	0.0001963	−0.022473	−0.0000879	−0.005789
	(0.64)	(−3.55)[c]	(−3.26)[c]	(−1.17)
WM 65–up	−0.0011026	−0.002738	0.0000231	0.006712
	(−3.82)[c]	(−0.45)	(1.31)	(1.27)
BM 16–19	−0.0001087	0.006898	−0.0000181	0.003266
	(−1.58)	(5.44)[c]	(0.46)	(3.99)[c]
BM 20–24	0.0001011	0.009667	−0.0000075	0.004763
	(0.85)	(4.21)[c]	(−1.12)	(3.57)[c]
BM 25–34	−0.0003558	0.027080	0.0003182	0.000009
	(−3.10)[c]	(1.18)	(4.61)[c]	(0.26)
BM 35–44	−0.0000615	0.003938	0.0000001	0.004192
	(−0.62)	(1.95)[a]	(0.07)	(2.76)[c]
BM 45–54	0.0000252	0.002764	−0.0000080	0.001259
	(0.57)	(2.88)[c]	(−2.94)[c]	(1.67)
BM 55–64	0.0001184	−0.000658	−0.0000111	0.001180
	(1.36)	(−0.36)	(−2.09)[b]	(0.86)
BM 65–up	−0.0000536	0.000720	0.0000001	0.000241
	(−0.97)	(0.65)	(0.09)	(0.36)
WF 16–19	0.0005834	0.024929	−0.0000229	0.010632
	(1.24)	(3.00)[c]	(−0.70)	(1.66)
WF 20–24	0.0004341	0.005085	0.0000182	0.019280
	(1.29)	(0.75)	(0.71)	(3.79)[c]
WF 25–34	−0.0011659	−0.002611	0.0002466	0.008719
	(−2.92)[c]	(−0.34)	(6.55)[c]	(1.48)
WF 35–44	−0.0004626	−0.006174	0.0000792	0.005472
	(−1.23)	(−0.81)	(2.17)[b]	(1.00)
WF 45–54	0.0014181	−0.018769	−0.0001301	0.001206
	(4.17)[c]	(−2.75)[c]	(−4.56)[c]	(0.22)
WF 55–64	0.0099186	−0.009696	−0.0000827	−0.002108
	(3.85)[c]	(−1.81)[a]	(−3.72)[c]	(0.50)
WF 65–up	0.0000102	−0.001768	−0.0000067	0.00332
	(0.07)	(−0.55)	(−0.66)	(1.21)

Table I. (Continued)

GROUP	INTERCEPT	*PDGNP*	*PDPOP*$_i$	*TIME*
BF 16–19	−0.0000838	0.002500	0.0000030	0.003886
	(0.96)	(1.50)	(0.59)	(3.38)[c]
BF 20–24	0.0001171	0.006482	0.0000001	0.001519
	(1.21)	(3.30)[c]	(0.90)	(1.06)
BF 25–34	−0.0001565	0.002615	0.0000270	0.003695
	(−1.48)	(1.20)	(3.59)[c]	(2.22)[b]
BF 35–44	−0.0000159	−0.000189	0.0000060	0.001531
	(−0.16)	(−0.09)	(0.90)	(1.06)
BF 45–54	0.0001829	−0.004506	−0.0000109	0.002295
	(2.22)[b]	(−2.62)[b]	(−2.16)[b]	(1.62)
BF 55–65	0.0002230	−0.001488	−0.0000070	−0.000523
	(2.85)[c]	(−1.05)	(−2.09)[b]	(−0.38)
BF 65–up	0.0000565	0.005823	−0.0000028	0.000133
	(1.49)	(0.73)	(−1.42)	(0.17)

KEY: WM = white male, BM = black male, WF = white female, BF = black female.
a. Significance at the 10% level.
b. Significance at the 5% level.
c. Significance at the 1% level.

the 8 subgroups with positive coefficients for *PDGNP*. The other 2 subgroups are middle aged black males. These groups find their employment shares increasing during expansion periods and decreasing during recessions. Positive coefficients for these groups were expected. The groups with negative *PDGNP* coefficients are all middle age and older workers. Their employment shares increase during recessions and decrease during expansions. These results were also expected. Overall, our more detailed analysis is consistent with the earlier results of Kosters and Welch.

The variable *PDPOP*$_i$ was positive and statistically significant for all workers 25–34 years old and WF 35–44 years old. This means that percentage increases in the share of the population of these groups increases the employment shares of these groups. As the population of these groups increases, their labor supply is also likely to increase. An increase in the labor supply for these groups causes their employment share to increase. On the other hand, *PDPOP*$_i$ was negative and statistically significant for all workers 45–64 years old. Variables not standardized for in the equation could cause these negative coefficients. For example, the effect of Social Security has been to reduce the employment share of older workers. It is possible that *PDPOP*$_i$ may be picking up the Social Security effect and other variables which are influencing the employment share equations but are not taken account of.

With the exception of BM 35–44, *TIME* was only statistically significant, but always positive, for younger workers. Over time the work force has become younger.

To sum up, our results indicate that employment shares are not constant over time, but are influenced by the state of the economy, relative population growth, and time. Thus, we have found evidence of a bias in studies employing aggregate data to measure the direction of change in real wages over the business cycle. Furthermore, the evidence indicates that during expansions, the share of older workers decreases. Since younger workers tend to earn less than older workers, the increasing work force share of younger workers during cyclical expansion imparts a downward bias in existing studies of the real-wage business cycle relationship.

An Empirical Analysis of Real Wages over the Business Cycle

We now examine the behavior of real wages over the business cycle using disaggregated data. To our knowledge, the behavior of real wages over the business cycle has never been examined by age, race, and sex. It is our hypothesis that an examination of the real wage at a disaggregated level will provide a better picture of the behavior of real wages over the business cycle than examinations at the aggregate level. Furthermore, we test the second criterion of no aggregation bias, which requires equal coefficients on all common exogenous variables.

For each age-sex-race group for which data were available, we estimated a real wage equation of the following form:

$$\ln(W/P)_i = \beta_{0i} + \beta_{1i} PDGNP + \beta_{2i}\ POPS_i + \beta_{3i} TIME + u_i. \tag{8}$$

The *PDGNP* and *TIME* variables are the same as previously. Here $POPS_i$ is the population share of group i rather than the percentage change in the population share. This empirical model follows from the exponential formulation of our model of the aggregate real wage. The variables *PDGNP* and *TIME* correspond to variables in the vector Z of that formulation; $POPS_i$ is a group specific variable from the vector X_i.

Unfortunately, for the purposes of estimating these real wage equations, data on real wages of different age-sex-race groups are sorely lacking and only exist for full-time workers. The source and construction of our wage variable are discussed in detail in Appendix A. We were able to construct a wage series for only 18 of the 28 age-sex-race groups for which employment share regressions were estimated. Wage data were unavailable for the age categories 16–19 and 20–24 and for two of the oldest groups. Further, wage data for the 18 categories for which data were available only cover the period 1968–1981. Although the length of the period is fairly short, it should be recognized that this was not a period of steady growth. During this period two recessions, 1974–75 and 1980–81, occurred. There was also a minor decline in real GNP during 1970. Therefore the data should provide evidence of the effect of business cycles on real wages.

Once nominal wages were derived for each group for the period 1968–1981, two different measures of the real wage were computed using the Wholesale Price Index (WPI) and the Consumer Price Index (CPI) as alternative deflators. This was done because several earlier studies have found results to be sensitive to the deflator used.[4] We wanted to test the sensitivity of our results to different measures of price.

The possible effects of *PDGNP* on $\ln(W/P)_i$ have been discussed previously. The variable $POPS_i$ measures the ratio of the population of each group to the total population. The predicted sign of the coefficient of $POPS_i$ is negative. That is, holding everything else constant, as the population of a group increases relative to the total population, the labor supply of this group will also increase. Assuming the various groups are not perfect substi-

4. Studies that use the CPI as the deflator include Bodkin [5], Neftci [20], and Mehra [18]. Bodkin found evidence of procyclical real wages while Neftci and Mehra found evidence of countercyclical real wages. Studies that use the WPI as the deflator include Kuh [16], Geary and Kennan [11], Otani [22], Chirinko [8], and Leiderman [17]. The first two of these found evidence of procyclical real wages while the latter three found evidence of countercyclical real wages.

Table II. Regression Results for Real Wages (WPI).

GROUP	INTERCEPT	*PDGNP*	*POPS*$_i$	*TIME*	R^2	DW	*F* VALUE
WM 25–34*	1.2057	0.8973	−1.9546	−0.0130	.89	1.92	24.59
	(3.36)[c]	(4.14)[c]	(−0.18)	(−0.73)			
WM 35–44*	1.2855	0.9569	−2.2245	−0.0095	.78	2.54	11.08
	(6.99)[c]	(4.04)[c]	(−0.59)	(−3.45)[c]			
WM 45–54	−0.7233	1.2205	24.4481	0.0349	.65	1.62	5.51
	(−0.42)	(3.78)[c]	(1.37)	(1.29)			
WM 55–64*	0.3429	1.1215	16.8136	0.0067	.69	1.75	6.61
	(0.23)	(4.36)[c]	(0.51)	(0.68)			
WM 65–up*	2.1498	2.2531	−11.3023	0.0201	.95	2.34	60.73
	(3.03)[c]	(6.78)[c]	(−1.27)	(11.37)[c]			
BM 25–34	1.6838	0.4532	−204.3568	0.0857	.50	1.73	2.96
	(6.88)[c]	(0.82)	(−2.24)[b]	(2.19)[a]			
BM 35–44*	2.2242	0.9355	−189.1807	0.0199	.52	1.63	3.29
	(3.50)[c]	(1.83)[a]	(−1.88)[a]	(2.09)[a]			
BM 45–54*	1.1829	0.7479	−46.5071	0.0003	.17	2.09	0.61
	(0.61)	(1.10)	(−0.16)	(0.03)			
BM 55–64	3.9547	0.4620	−589.4132	0.0223	.60	2.14	4.44
	(2.88)[b]	(0.74)	(−2.21)[b]	(2.83)[b]			
WF 25–34*	1.5011	0.5655	−19.9573	0.0231	.50	2.87	3.06
	(2.58)[b]	(2.12)[a]	(1.37)	(1.15)			
WF 45–54*	2.4849	0.6009	−23.3020	−0.4689	.75	2.39	9.09
	(2.65)[b]	(3.50)[c]	(−1.87)[a]	(−2.10)[a]			
WF 55–64	−5.7981	0.4865	98.6701	0.0263	.86	2.88	18.63
	(−4.24)[c]	(2.72)[b]	(4.95)[c]	(4.52)[c]			
WF 65–up*	−9.4539	1.9846	150.1199	−0.0152	.63	1.98	5.18
	(−1.40)	(2.97)[b]	(1.47)	(−1.21)			
BF 25–34*	0.9202	0.3485	−207.4637	0.1146	.68	2.13	6.34
	(8.78)[c]	(1.33)	(−3.32)[c]	(9.26)[c]			
BF 35–44*	1.8083	0.3125	−113.0540	0.0122	.56	2.35	3.87
	(4.01)[c]	(0.85)	(−2.36)[b]	(3.10)[c]			
BF 45–54*	−3.3985	1.5065	458.4698	0.0197	.71	1.55	7.43
	(−1.06)	(2.94)[c]	(1.20)	(3.11)[c]			
BF 55–64*	5.8351	0.5937	−777.3221	0.0740	.76	1.62	9.36
	(1.64)	(0.76)	(−1.50)	(1.88)[a]			

KEY: WM = white male, BM = black male, WF = white female, BF = black female.
a. Significance at the 10% level.
b. Significance at the 5% level.
c. Significance at the 1% level.

tutes for each other, an increase in the labor supply then lowers the real wage of this group. We included *TIME* as a surrogate for other variables such as technological change that may affect real wages. Because we only have 14 observations and there are 18 different groups, it was not possible to utilize Zellner's technique. Instead, each regression was estimated by ordinary least squares, transforming where needed for autocorrelation utilizing the Cochrane-Orcutt procedure. In Tables II and III, asterisks denote transformed equations.

Table II presents the results for the real wage equations, in which WPI is used as the price deflator. The business cycle variable *PDGNP* indicates procyclical wages in all eighteen groups, with two of these groups statistically significant at the 10% level and ten of these groups statistically significant at the 5% level. Interestingly, in all of the white male and female groups, the *PDGNP* variable was positive and statistically significant, while only for black males 35–44 years old and black females 45–55 years old were the coefficients statistically significant. This appears to indicate that the real wages of white workers are more affected by the business cycle than those of black workers. However, because of data limitations, we think that more work is needed on this question before any definite conclusions can be reached.

The variable $POPS_i$ was statistically significant in seven out of the eighteen groups. Five of these groups were black worker groups and the other two groups were from the female categories. With the exception of white females aged 55–64, the coefficient on $POPS_i$ has the predicted negative sign. That is, increases in a group's population relative to the total population lowers the real wage for that group.

Of the ten groups for which the coefficient for *TIME* is statistically significant, eight of these groups have a positive sign. Most of these groups are black. The upward trend in real wages for these groups could be explained by a variety of factors, including improvements in education and government affirmative action programs.[5]

Table III presents the results when the CPI is used as the price deflator. The business cycle variable *PDGNP* is statistically significant in ten of the eighteen equations and the coefficient is positive in each regression just as when the WPI was used as the price deflator. The only differences between these results and those when the WPI are used is that one of the white subgroups, WF 25–34, and the only black male group that was previously significant, BM 35–44, lose significance.

The importance of these results is that regardless whether the WPI or CPI is used, the coefficient for *PDGNP* is always positive whenever statistically significant. Many of the previous studies of real wage behavior reported that their empirical results were sensitive to price deflator used. In some cases where both deflators were used, the regression estimates were statistically significant with only one of the deflators. Also, some of the studies reported that the sign of the business cycle coefficient depended on the price deflator. The fact that we found that real wages are procyclical regardless of which deflator is utilized may indicate the significance of looking at disaggregated rather than aggregated data.

Finally, recall from our theory that one requirement for the existence of no aggregation bias in real wage studies is that the variables in the vector Z have the same proportionate impact on the real wages of all subgroups, i.e., $b_i = b_2 = \dots b_n$. In our empirical analysis, this amounts to testing whether the variables *PDGNP* and *TIME* have the same impact on $\ln(W/P)_i$ in all 18 subgroups. To test these restrictions, we reestimated the regressions reported in Tables II and III imposing equality restrictions on the coefficients for *PDGNP*

5. Several recent studies based on micro data have reached contradictory conclusions concerning the source of rising black real wages. Freeman [10] attributes the rise to government antidiscrimination programs, while Butler and Heckman [6] attribute it to the declining labor force participation of older black males. Rising real wages of young blacks may be explained by improvements in education [24].

Table III. Regression Results for Real Wages (CPI).

GROUP	INTERCEPT	*PDGNP*	$POPS_i$	*TIME*	R^2	DW	*F* VALUE
WM 25–34*	0.3576	0.6262	22.0056	−0.4571	.80	1.37	12.17
	(0.87)	(2.66)[b]	(2.03)[a]	(−2.56)[b]			
WM 35–44*	2.3487	0.6469	−9.7297	−0.0059	.83	2.18	14.62
	(14.47)[c]	(3.33)[c]	(−4.71)[c]	(−3.39)[c]			
WM 45–54	−2.0760	1.0020	36.2456	0.0581	.61	1.55	4.69
	(−1.16)	(3.02)[c]	(2.05)[a]	(2.16)[a]			
WM 55–64*	3.2481	0.9408	−31.2265	−0.0008	.79	1.57	11.60
	(1.86)[a]	(3.54)[c]	(−1.10)	(−0.10)			
WM 65–up*	2.5615	1.9960	−15.3343	0.0250	.98	2.14	127.77
	(4.00)[c]	(7.29)[c]	(−2.14)[a]	(17.48)[c]			
BM 25–34	0.8924	0.2623	−100.2910	0.0435	.15	1.46	0.51
	(5.01)[c]	(0.46)	(−0.96)	(0.95)			
BM 35–44*	2.1554	0.5993	−223.5942	0.0791	.48	1.40	2.76
	(2.92)[b]	(1.02)	(−1.76)	(2.26)[b]			
BM 45–54*	−0.7966	0.7844	224.1980	0.0157	.31	1.91	1.37
	(−0.36)	(1.10)	(0.74)	(1.37)			
BM 55–64	5.2164	0.1257	−664.3229	0.0296	.83	2.34	14.48
	(4.14)[c]	(0.25)	(−3.39)[c]	(5.47)[c]			
WF 25–34*	0.4073	0.2913	4.8970	−0.0058	.17	2.45	0.61
	(0.80)	(1.23)	(0.38)	(−0.32)			
WF 35–44*	1.1708	0.4718	−4.1718	0.0038	.92	2.57	35.23
	(10.69)[c]	(3.71)[c]	(−3.37)[c]	(3.09)[c]			
WF 45–54*	−0.2602	0.4236	10.6708	0.0207	.60	2.32	4.52
	(−0.35)	(2.70)[b]	(1.52)	(1.69)			
WF 55–64	−0.6049	0.5411	19.7082	0.0093	.68	2.16	6.29
	(−0.34)	(2.51)[b]	(0.87)	(1.41)			
WF 65–up*	−8.9239	1.7154	139.5004	−0.0421	.71	1.95	7.36
	(−1.35)	(2.65)[b]	(1.41)	(−1.01)			
BF 25–34*	0.4151	0.1971	−60.5677	0.0317	.27	2.32	1.10
	(7.41)[c]	(0.87)	(−0.95)	(0.84)			
BF 35–44*	1.9441	−0.1016	−144.7246	0.0193	.71	1.94	7.40
	(4.01)[c]	(−0.03)	(−2.76)[b]	(4.46)[c]			
BF 45–54*	−8.2544	1.4884	884.6098	0.0314	.84	1.62	15.87
	(−2.31)[b]	(2.86)[b]	(2.42)[b]	(5.42)[c]			
BF 55–64*	9.2523	−0.0829	−1124.4500	0.1056	.89	1.77	24.33
	(2.72)[b]	(−0.12)	(−2.62)[b]	(3.26)[c]			

KEY: WM = white male, BM = black male, WF = white female, BF = black female.
a. Significance at the 10% level.
b. Significance at the 5% level.
c. Significance at the 1% level.

and *TIME*, across all 18 equations. Based on an *F* test for the validity of each of these sets of restrictions, the null hypothesis was rejected in every case at the .01 level. The likelihood that the real wages of all subgroups respond in a (proportionately) similar fashion to changes in *PDGNP* and *TIME* is highly doubtful. The second condition required for no aggregation bias in aggregate real wage studies does not hold.

V. Summary and Conclusion

The first part of this paper outlined the nature of a possible aggregation bias in previous studies of the effect of the business cycle on real wages. We derived the conditions under which such a bias would exist. Our empirical analysis confirmed the existence of this bias.

The first part of our empirical analysis examined the effect that business cycles have on the employment shares of twenty-eight demographic groups. We found that young workers are more affected by the business cycle than older workers and that black workers are more affected than white workers. On the other hand, older workers and white workers are less likely to be laid off. The significance of these results is that different employment shares are differentially affected by changes in the business cycle. This indicates that previous aggregate real wage analyses may have provided biased estimates since they failed to take account of changing employment shares.

The second part of the empirical work examined the behavior of real wages over the business cycle for eighteen demographic groups. The empirical results indicate procyclical real wages regardless of whether the WPI or CPI is used as the price deflator. Furthermore, results from a comparison of coefficients across equations provides further evidence of an aggregation bias.

While these results are suggestive, further work remains to be done. First analysis of the effect of the cycle of shares should be expanded to include an examination of how manhour shares and not just employment shares are affected by the cycle. That is, how the mix between full-time and part-time changes with cycle needs to be addressed. Second, real wage equations need to be estimated for part-time as well as full-time workers. Third, a more detailed analysis of the data to include occupation could be made. While these three tasks could not be performed with available data, we are exploring the possibility of such an analysis with the detailed microdata from the Current Population Survey.

Appendix

Yearly earnings data for different age-sex-race groups are available in the yearly *Current Population Reports: Consumer Income*, Series P-60. However, there are some gaps in the series. Therefore some interpolation was needed to derive hourly earnings for these groups. The mean income of year-round full-time workers for each demographic group was available over the period 1968 to 1981. Yearly earnings of these same year-round full-time workers were available from 1975 to 1979. The ratio of earnings to income was then calculated for each group for each year from 1975 to 1979. For each group there was very little variation in these ratios from year to year. So for each demographic group an average ratio was calculated from the five yearly ratios over the period 1975 to 1979. Then the mean income of year round full-time workers was multiplied by this earnings-income ratio for each year covering the period 1968 to 1981. This provided us with estimates of annual earnings of these workers from 1968 to 1981.

From these yearly earnings, we proceeded to derive the nomal hourly earnings for each subgroup. We assumed that year-round full-time employees work forty hours a week for fifty-two weeks a year for a total of 2,080 hours during the year. The average nominal wage was then derived by dividing yearly earnings by 2,080 hours.

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