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# Explaining procyclical male–female wage gaps<sup>☆</sup>

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## Abstract

Our analysis based on the National Longitudinal Survey of Youth for the 1978–1999 period concludes that men's greater representation in cyclical occupational groups, such as craftsmen, operatives, and laborers, more than accounts for the gap between men and women in the cyclicalities of real wages.

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## 1. Introduction

A number of early studies that are based on aggregate time series data concluded that real wages are nearly noncyclical.<sup>1</sup> However, as first pointed out by Raisian (1979) and rigorously demonstrated by Bils (1985) and Solon et al. (1994), aggregate wage series are countercyclically biased by their tendency to weigh low-skilled workers more heavily in expansions than in recessions. More recent analyses of longitudinal micro data, which avoided this composition bias by tracking the same individuals over time, have found that real wages are much more procyclical than they appear in aggregate time series data.

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<sup>1</sup> See Solon and Barsky (1989) for a detailed summary of the time series evidence.

In an effort to explain why real wages are strongly procyclical, economists have investigated heterogeneity in the cyclicity of real wages across different economic and demographic groups. One notable finding is that real wages are much more procyclical for men than women. For example, on the basis of the Panel Study of Income Dynamics (PSID) data for the 1967–68 to 1986–87 period, [Solon et al. \(1994\)](#) found that a one percentage point reduction in the national unemployment rate is associated with a rise in real wages by 1.40% for men and 0.42% for women. Their results are quite consistent with those of [Blank \(1989\)](#) and [Tremblay \(1990\)](#).

Despite the repeated findings that men's real wages are more responsive to the cycle than women's wages are, no previous study explained empirically why they are so. This study is apparently the first that conducts a quantitative assessment of the procyclical wage gap between men and women. Explaining sources of men's greater wage procyclicality gives us a deeper grasp of why real wages are procyclical as well as how male–female wage gaps are different depending on the phase of the business cycle.

## 2. Data and econometric methods

As noted by [Abraham and Haltiwanger \(1995, p.1259\)](#), point-in-time wage measures such as survey week wages tend to select workers with strong labor market attachment, and, therefore, are countercyclically biased. On the other hand, annual wage measures such as average hourly earnings defined as the ratio of annual earnings to annual hours are less subject to this sample selection bias, because most workers are likely to work at some point in the year. However, given that we want to explain men's greater wage procyclicality based on occupational segregation between genders, assigning an occupation code to each average hourly earnings observation is often complicated when a worker has more than one job in a calendar year, which belong to different occupational categories.

To avoid the difficulty, this paper relies on the National Longitudinal Survey of Youth (NLSY), which began in 1979 with a national sample then between the ages of 14 and 22, reinterviewed the sample each year until 1994, and then switched to biennial interviews. The detailed job-specific information contained in the job history file enables us to tell how many jobs are held by each respondent in a calendar year and which occupation category each job belongs to. We use average hourly earnings observed for an individual in a calendar year only when all jobs held by the respondent in that year belong to the same occupation.<sup>2</sup>

The model for real wage changes is

$$\ln \frac{W_{it}}{W_{i,t-s}} = \beta_1 + \beta_2 s + \beta_3 (s \cdot t) + \beta_4 (s \cdot X_{it}) + \beta_5 (U_{it} - U_{i,t-s}) + (\varepsilon_{it} - \varepsilon_{i,t-s}) \quad (1)$$

where  $W_{it}$  is real average hourly earnings of individual  $i$  in year  $t$ , and  $s$  equals 1 or 2 according to whether the most recent interview before the year  $t$  interview was one or two years earlier.<sup>3</sup> As annual earnings and hours in year  $t-1$  are reported in the year  $t$  interview,  $s=2$  for the 1993–95, 1995–97, and 1997–99 observations. A linear trend is included, and potential experience,  $X_{it}$ , is measured as age minus years of schooling minus 6. Our cyclical indicator,  $U_{it}$ , is the average unemployment rate computed over the

<sup>2</sup> [Park and Shin \(2003\)](#) found that eliminating all within-year-between-occupation changers produces little bias in estimated wage cyclicality. They did find, however, that using survey week information (e.g., industry, occupation, or union status) in

<sup>3</sup> See [Shin \(1994, p. 139\)](#) for the justification of this specification.

Table 1  
Wage cyclicality by gender

|                    | Men                                     | Women                                 | Men–women          |
|--------------------|---|---------------------------------------|--------------------|
| All workers        | –0.0138*** (0.0026) [ <i>N</i> :53,921] | –0.0045* (0.0027) [ <i>N</i> :49,446] | –0.0093** (0.0038) |
| Occupation stayers | –0.0136*** (0.0040) [ <i>N</i> :20,131] | –0.0045 (0.0037) [ <i>N</i> :23,128]  | –0.0090* (0.0055)  |

Standard error estimates and numbers of differenced year–person observations are in parentheses and brackets, respectively. \*, \*\*, and \*\*\*Significant at the 10%, 5%, and 1% significance levels, respectively.

working months of individual  $i$  in year  $t$ .<sup>4</sup> Naturally, the average of monthly Consumer Price Indices computed over the working months is used to arrive at real wages.  $\beta_5$  is greater than, equal to, or less than 0 as real wages move countercyclically, acyclically, or procyclically. Note that all the time-invariant individual-specific characteristics that influence wages in levels are “differenced out” in the measurement of year-to-year change.

Because, as will be shown in Section 3, 100% of the gender gap in overall wage cyclicality is attributed to the gender gap in the wage cyclicality of occupation stayers, we further decompose the latter as follows.

$$\beta_s^m - \beta_s^f = \sum_{j=1}^8 \beta_{sj}^m (S_j^m - S_j^f) + \sum_{j=1}^8 (\beta_{sj}^m - \beta_{sj}^f) S_j^f \quad (2)$$

$$= \sum_{j=1}^8 \beta_{sj}^f (S_j^m - S_j^f) + \sum_{j=1}^8 (\beta_{sj}^m - \beta_{sj}^f) S_j^m \quad (3)$$

where  $\beta_s^g$  is the cyclicality of stayers’ wages for each gender  $g$ ,  $S_j^g$  the  $j$ th occupation’s employment share within gender  $g$ , and  $\beta_{sj}^g$  represents the wage cyclicality of gender  $g$  in occupation  $j$ . For each gender,  $\beta_{sj}^g$  is estimated by applying Ordinary Least Squares (OLS) to Eq. (1) with occupation dummies and their interactions with unemployment changes included as additional regressors. The first term in the right-hand side of Eq. (2) (called the between-occupation component) represents the amount of the gender gap in stayers’ wage cyclicality explained by occupational share gaps assuming that occupation-specific cyclicalities are equal between genders in all occupations. The second term (the within-occupation component) represents the amount explained by the cyclicality differences within occupations assuming that occupation shares are equal for both genders in all occupational categories. Note that the assumed equal occupation-specific cyclicalities in the first term and the assumed equal occupational shares in the second term are indexed as men and women, respectively. To check the robustness of the results, Eq. (3), which uses the opposite gender index, is also estimated.

### 3. Empirical results

For the 1978–79 to 1997–99 period, estimated wage procyclicality are –0.0138 for men and –0.0045 for women (row 1, Table 1),<sup>5</sup> implying that a one percentage point reduction in the unemployment rate

<sup>4</sup> This correct specification is motivated by Keane et al. (1988, pp. 1245–46).

<sup>5</sup> In most cases, the homoskedasticity hypothesis is rejected at any conventional significance level against the alternative that the error variance is greater when  $S=2$  than when  $S=1$ . Correcting for the heteroskedasticity, however, makes little difference in the final estimates, because the difference between the two types of error variances is empirically unimportant.

Table 2

Wage and hours cyclicalities by occupation and by gender

|                       | Men    |                    |                     | Women  |                    |                     |
|-----------------------|--------|--------------------|---------------------|--------|--------------------|---------------------|
|                       | Share  | Wage cyclicalities | Hours cyclicalities | Share  | Wage cyclicalities | Hours cyclicalities |
| Professional          | 15.60  | −0.0212** (0.0098) | −0.0210** (0.0084)  | 18.31  | −0.0121 (0.0090)   | 0.0168* (0.0090)    |
| Managers              | 8.36   | −0.0107 (0.0135)   | −0.0053 (0.0121)    | 6.04   | −0.0270* (0.0164)  | 0.0157 (0.0171)     |
| Sales                 | 3.37   | −0.0042 (0.0201)   | −0.0017 (0.0174)    | 2.32   | 0.0061 (0.0213)    | 0.0107 (0.0214)     |
| Clerical              | 6.41   | 0.0147 (0.0136)    | 0.0171 (0.0117)     | 42.22  | −0.0002 (0.0052)   | −0.0003 (0.0052)    |
| Craftsmen,<br>Foremen | 20.00  | −0.0170** (0.0085) | −0.0184** (0.0074)  | 0.62   | −0.0150 (0.0485)   | 0.0190 (0.0506)     |
| Operatives            | 21.39  | −0.0178** (0.0078) | −0.0160** (0.0067)  | 9.59   | −0.0076 (0.0113)   | −0.0190* (0.0113)   |
| Laborers              | 7.85   | −0.0305** (0.0126) | −0.0362*** (0.0108) | 0.47   | −0.1275** (0.0570) | 0.0762 (0.0555)     |
| Service               | 17.02  | −0.0052 (0.0087)   | −0.0190*** (0.0074) | 20.80  | −0.0034 (0.0074)   | −0.0013 (0.0073)    |
| <i>N</i>              | 20,131 |                    |                     | 23,128 |                    |                     |

Standard error estimates are in parenthesis. \*, \*\*, and \*\*\*Significant at the 10%, 5%, and 1% significance levels, respectively.

leads to a rise in real wages by 1.38% and 0.45% for men and women, respectively. The equal cyclicalities hypothesis between genders is rejected at the 5% significance level (last column).<sup>6</sup> These estimates are quite consistent with the results of Solon et al. (1994). Estimates for occupation stayers (row 2) virtually replicate respective estimates for all workers. In particular, approximately 100% of the gender gap in estimated wage procyclicality is explained by the gender gap among occupation stayers.<sup>7</sup>

The first and the fourth columns of Table 2 present gender differences in the occupational distribution of workers. Each share is computed as a simple average of yearly percent distributions of occupational workers among all workers within each gender. Clerical is the most female-dominated occupational category, while Craft, Operatives, and Laborers are male-intensive ones. Moreover, estimates in the second and the fifth columns reveal much greater wage procyclicality in these male-dominated occupational categories than the female-dominated one. On the basis of the men's sample, *t*-tests reject the null hypothesis of equal cyclicalities between Craft versus Clerical (*t*-value=3.97), between Operatives and Clerical (*t*-value=4.36), and between Laborers and Clerical (*t*-value=5.98) at the 5% significance level. For women, the null is rejected between Laborers and Clerical (*t*-value=4.96) at the 5% level.

In the third and the sixth columns are reported estimated cyclicalities of annual work hours by gender and by occupation. Just like real wages, hours are much more procyclical in male-intensive occupational groups than in the female-dominated one, with the tendency much stronger in the men's sample. For men, an *F*-test rejects the null hypothesis of inter-occupation homogeneity in hours cyclicalities at the 10% significance level (the *F*-value=1.96), and a share-weighted correlation of estimated occupation-specific wage cyclicalities and hours cyclicalities is 0.85, which is statistically significant even at the 1% level. Considering that overtime is included in both wage and hours data, these results suggest that greater wage procyclicality in male-dominated occupational categories is explained at least in part by greater procyclicality of overtime shares in these groups.

<sup>6</sup> To test for the potential bias in standard error estimates that may arise by neglecting a year-specific random component in the error term of Eq. (1), we conduct consistent covariance matrix estimation that is robust to within-year clustering as well as heteroskedasticity. All the test results in the current paper remain unchanged from this exercise.

<sup>7</sup> See Solon et al. (1997, p.412) for detailed explanation of the underlying logic.

Results in Table 2 produce estimated between-occupation components of  $-0.0132$  and  $-0.0125$  for Eqs. (2) and (3), respectively, with respective standard error estimates of 0.0056 and 0.0128. Therefore, the between-occupation components in both equations exceed the estimated gender gap in wage procyclicality,  $-0.0093$ . It is concluded that men's greater wage procyclicality is more than accounted for by men's greater representation in cyclical occupational categories which exhibit great wage procyclicality.

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