An Index Number Approach to the Measurement of Wage Differentials by Sex

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Communications

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I. Introduction

Over the last two decades, one of the most widely examined economic topics has been the earnings differential between men and women. If any common perception exists regarding the time path of male-female earnings differences, it is that little has changed over time.¹ This perception is due to the fact that the most commonly cited statistic representing the time series of earnings differences by sex, the ratio of female to male median annual full-time wage and salary income, is virtually unchanged over the 1955–85 period. A second widely reported aggregate statistic, the ratio of female-to-male median weekly wages, shows little improvement from 1967 until the 1980s.

However, these aggregate figures are seriously flawed as indicators of average wage differentials between men and women. Part of the problem is the lack of control for differences in average productivity characteristics between men and women. But these measures suffer from an even more fundamental problem: they do not reflect earnings differentials within individual labor markets but rather the difference in median earnings across labor markets. This problem arises since the median wage for

¹. In his survey of studies of discrimination, Cain (1986, 750–52) concludes that, “for my purposes the essential [fact] from time-series data that [pertains] to economic discrimination . . . is the near-constant ratio of women’s-to-men’s wages over a 40 year period, using the data on earnings of year-round, full-time workers.”
women is likely to be in a different labor market than the median wage for men.\textsuperscript{2} Furthermore, the pattern of growth in median wages for women is strongly affected by female earnings in the clerical occupation.

Such problems in aggregate price measurement have long been recognized in the development of consumer and input price indices. For this reason, price indices typically aggregate prices for individual commodities weighted by the share of the commodity in consumer budgets. In contrast, the typical practice used in measuring earnings differences by sex is analogous to measuring the change in consumer prices from one year to the next by looking at the change in the median price across all consumer goods. Clearly there is room to apply the standard procedures for creating price indices to the study of earnings differentials.\textsuperscript{3} In this paper, we provide such indices based on occupational groupings varying from 8 to 412 categories.

An alternative approach would be to compute earnings differentials from a micro data set such as the Current Population Survey (CPS) using an Oaxaca-type decomposition. Blau and Beller (1988) have very recently taken a first step in this direction using 1971 and 1981 data. However, repeating their procedure for each year of CPS data would be extremely burdensome computationally. Our approach using more aggregate data is easier to implement and provides the only available time series 1967–86 of female-male wage indices corrected for change in occupational composition. In addition, our approach provides insights into changes in relative wages for particular occupational markets such as clericals, which typically would not be analyzed using micro data.

In the next section, we examine the time paths of earnings differentials in individual occupations to determine if the trend of the aggregate ratio of median wages differs from the disaggregated trends. Next, we present the empirical methodology for creating an aggregate index of wage differentials from occupation level data. These indices are reported in section four. The final section reviews the implications of the results and suggests future directions for research.

\textsuperscript{2} Data on median year-round, full-time earnings by 8–12 major disaggregated occupations show that ignoring intra-occupation variation in earnings, the median occupation for women (determined by ordering the occupations from highest to lowest median earnings and selecting the median) is clerical work for each year examined between 1949 and 1982. For men, the median occupation would have been the clerical occupation in the early 1950s, sales starting in the late 1950s and early 1960s, and crafts from the late 1960s through 1982.

\textsuperscript{3} Sanborn (1964) made an early attempt to compute earnings indices. However, his index does not represent an index of differences in market level earnings by sex since it does not aggregate over individual market differentials between men and women.
II. The Trend in Earnings Differential by Occupation

The two most widely used measures of time series data on male and female earnings are the CPS data on annual earnings of year-round, full-time workers and the CPS data on usual weekly earnings of full-time workers. As O’Neill (1985) has pointed out, the annual earnings data is subject to greater measurement error due to recall bias and greater selection bias than the weekly data. The weekly earnings data may be subject to less error due to differences in weeks and/or hours worked per year between men and women although the weekly data still may imply shorter hours per week for women relative to men. Therefore, we concentrate on the analysis of the weekly wage data.

Because of the change in the occupational definitions starting in 1983, we examine the trends between 1967 and 1982. These female-to-male wage ratios are reported in Table 1. The first column, the ratio based on median weekly wages across all occupations, provides a reference against which the other series may be compared. It shows little movement from 1967 to 1979 and then a small increase in the 1980s. However, there is a wide variance in trends across the various occupational markets. The only occupation in which women’s wages fell relative to men’s wages over the 1967–82 period was the clerical occupation. It is interesting to note that the partial recovery for relative female clerical earnings from 1979–82 coincides with the aggregate relative gains observed for women during this more recent period. The professional occupation registered a very slight increase of one percentage point over the 1967–82 period. At the other extreme, relative wages in the sales, laborer and service occupations rose by 12, 15, and 17 percentage points respectively. The managerial and craft occupations posted increases of 9 and 5 percentage points. Information on operatives only is available beginning in 1973, but the relative gains by women operatives between 1973 and 1982 was only 3 percentage points. For the occupations that registered relative gains, the gains generally began in the early 1970s. However, these gains failed to appear in the aggregate data due, no doubt, to the fact that the decline in the clerical occupation had such a large impact on overall median female earnings. These results clearly show that wage gains for women were not uniform over all occupations and that the trend of occupational market wage differentials varies substantially from the overall trend of median wages.

4. Moulton (1986), using Social Security Data, is one exception.
Table 1
The Trend in Female-to-Male Ratios of Median Weekly Earnings by Occupation

<table>
<thead>
<tr>
<th>Year</th>
<th>Total</th>
<th>Prof</th>
<th>Mgr</th>
<th>Cler</th>
<th>Sales</th>
<th>Craft</th>
<th>Oper</th>
<th>Serv</th>
<th>Lab</th>
</tr>
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<tbody>
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<td>.70</td>
<td>.52</td>
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<td>.43</td>
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<td>.63</td>
<td>.68</td>
<td>.75</td>
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<tr>
<td>1980</td>
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<td>.59</td>
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<td>.63</td>
<td>.69</td>
<td>.76</td>
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<td>1981</td>
<td>.65</td>
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<td>.61</td>
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<td>.63</td>
<td>.71</td>
<td>.79</td>
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<tr>
<td>1982</td>
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<td>.61</td>
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<td>.55</td>
<td>.64</td>
<td>.64</td>
<td>.73</td>
<td>.83</td>
</tr>
</tbody>
</table>

The occupations include Prof = professional; Mgr = managerial; Cler = clerical; Sales; Craft; Oper = operatives; Serv = service; and Lab = laborers.

III. Empirical Methodology

Since the early work by Duncan and Duncan (1955), economists and sociologists have been constructing indices of occupational segregation as indicators of changes in the aggregate status of women in the labor force. The basic formulation for the index over n occupations is

\[ S_t = \frac{1}{2} \sum_{i=1}^{n} |m_{it} - f_{it}| \]

where \( S_t \) is the index of segregation at time \( t \), \( m_{it} \) is the proportion of all employed males in occupation \( i \) at time \( t \), and \( f_{it} \) is the proportion of all employed females in occupation \( i \) at time \( t \). Although this is not typically done, \( m_{it} \) and \( f_{it} \) could also be defined respectively as the ratio of total male earnings in occupation \( i \) to total male earnings across all occupations.

5. See, for instance, Blau and Hendricks (1979).
and the ratio of total female earnings in occupation \( i \) to total female earnings across all occupations. In either case, the index varies between zero and one with zero being perfect integration and 1 being perfect segregation.\(^6\)

One objective of this paper is to propose a related index for the dissimilarity in earnings between men and women. A natural extension of the formula in (1) is

\[
D_t = \frac{1}{2} \sum_{i=1}^{n} (m_{it} + f_{it})(\ln W_{it}^F - \ln W_{it}^M)
\]

where \( \ln W_{it}^F \) and \( \ln W_{it}^M \) are the occupation \( i \)-specific natural logarithms of median female and median male earnings respectively. In general, this index can be positive or negative. A positive value implies an aggregate earnings advantage for women and a negative value implies an aggregate earnings advantage for men.\(^7\) The index will equal zero if male and female earnings are equal in each occupation or if the sum of the weighted earnings advantages for men are equal to the weighted earnings advantages for women.

In order to transform the index into a form which corresponds directly to the typically reported ratio of female-to-male earnings, we need only exponentiate the index. The resulting form is

\[
\exp(D_t) = \frac{W_{it}^F}{W_{it}^M} = \prod_{i=1}^{n} \left( \frac{W_{it}^F}{W_{it}^M} \right)^{1/2} (m_{it} + f_{it})
\]

This measure can vary from zero (if women are not paid for their work) to \(+ \infty\) (if men are unpaid labor). Values less than one imply earnings advantages for men, whereas values greater than one imply earnings advantages for women. A value of one implies no aggregate wage differential between men and women.

The index form in (2) is known as a Tornqvist approximation to a Divisia Wage index if \( m_{it} \) and \( f_{it} \) are measured respectively as the share of total male earnings in occupation \( i \) in year \( t \) and the share of total female earnings in occupation \( i \) in year \( t \). Dievert (1976) has shown that for intertemporal comparisons of prices or quantities, the Tornqvist approximation is consistent with the commonly used translog second-order ap-

\(^6\) Note that because the \( m_{it} \) sum to one and the absolute value of the \((-f_{it})\) sum to one, perfect segregation would imply a total index value of 2. Hence, the parameter \((\frac{1}{2})\) normalizes the index so that it equals 1.

\(^7\) In contrast to Equation 1, we delete the absolute value notation precisely because we are interested in the size and direction of the earnings gap, and not solely interested in the extent of dissimilarity. The proportion employed \((m_{it} + f_{it})\) places more weight on the larger occupations.
proximation to an arbitrary aggregate production or price function. Caves, Christensen, and Diewert (1982) extended these results to interspatial comparisons. This research demonstrates that if aggregate male and female unit costs may be approximated by the translog form, then (2) and (3) may be derived exactly as the index of wage differentials between women and men under the assumption of cost-minimization by firms. In contrast, the difference in median or mean wages between men and women will be consistent with the true aggregate wage differential if women are perfect substitutes for one another and men are perfect substitutes for one another so that there is only one occupational market for women and one occupational market for men.

IV. Indices of Wage Differentials by Sex

In Table 2, we report several Tornqvist indices of wage differentials by sex, using various levels of occupational disaggregation and various controls for productivity characteristics. The first column presents the trend of female-to-male median annual earnings for full-time, year-round workers (U.S. Bureau of Census) and the second reports the trend of female-to-male median usual weekly earnings of full-time workers (U.S. Bureau of Labor Statistics, unpublished tabulations). Our analysis concentrates on the latter series for the reasons mentioned above, but the annual earnings ratios are included for the sake of comparison.

The trend in the annual earnings data is similar to the trend in the weekly wage data except that the movement upward in relative female annual earnings does not begin until 1982. By comparison, female relative weekly wages rose 7 percentage points between 1979 and 1986, after a

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8. See, for example, the discussion in Denny and Fuss (1983, 321–22).
9. To illustrate this point, we can consider the circumstances in which Equation 2 will equal \( \ln \bar{W}_f - \ln \bar{W}_M \) where \( \bar{W}_f \) is the mean wage across all females and \( \bar{W}_M \) is the mean wage across all males. This equality requires that

\[
\frac{1}{2} \sum (m_i + f_i) \ln W_f^i - \frac{1}{2} \sum (m_i + f_i) \ln W_M^i = \ln \bar{W}_f - \ln \bar{W}_M
\]

It is straightforward to show that this equality is consistent with an assumption that women are all in the same occupational market so that \( W_M^i = \bar{W}_M \) for all \( i \), and that all men are in the same occupational market \( (W_M^i = \bar{W}_M \) for all \( i \)). Such an occurrence of equal wages across all occupations is consistent with perfect substitutability among females and perfect substitutability among males. When this occurs, the average of the logs equals the log of the average. If wages are different across occupations, use of average wages will yield the true wage differential only by accident.

10. We performed some of the same procedure: on the annual earnings data and found results more similar to the uncorrected medians. These results are available from the authors upon request.
Table 2
Uncorrected and Corrected Female-to-Male Ratios of Median Earnings of Full-time Wage and Salary Workers

<table>
<thead>
<tr>
<th></th>
<th>Uncorrected</th>
<th></th>
<th>Corrected</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Annual</td>
<td>Weekly</td>
<td>Number of Occupations</td>
</tr>
<tr>
<td></td>
<td>Earnings</td>
<td>Earnings</td>
<td>8–12</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>2</td>
<td>3</td>
</tr>
<tr>
<td>1967</td>
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<tr>
<td>1986</td>
<td>.64</td>
<td>.69</td>
<td>.68</td>
</tr>
</tbody>
</table>

Column 1 is the ratio of female to male median annual earning of year-round full-time workers as taken from various issues of the U.S. Bureau of the Census, Current Population Reports.

Column 2 is the ratio of female-to-male median weekly earnings.

Columns 3 through 5 are computed using Equation (3). They are based on the weighted average of occupational wage differentials at differing levels of occupational aggregation.

Column 3 presents the index of market wage differentials using Equation (3). The level of disaggregation is limited by data availability. From 1967 through 1972, occupational wage data are available only for seven broad occupational groups as listed in Table 1. In 1973, nontransportation operative wages began to be reported. Then in 1983, the occupations were redefined into 12 major groups, adding farm laborers and transportation operatives, dividing professional/technical into two separate categories, and dividing service into protective service and other service. In addition, some occupational subgroups were shifted between major groupings. These occupational redefinitions may be responsible for the increase in the index between 1982 and 1983, but the indices should be consistent before and after that break. The results show that the aggregate wage index rose from a low of .62 in 1967 to a fairly steady ratio of about .64 in the 1970s and then rose an additional 2 percentage points by 1982.

Further disaggregation by occupations can only be done for a few years. Disaggregation into 35–36 occupations over the 1983–86 period (Column 4) increases the level of the female-male wage ratio by between 4 and 7 percentage points relative to Column 2. In addition, the Bureau of Labor Statistics has released data on median wages by still more finely defined occupations for the years 1979–84. This index of female to male wage ratios (based on between 375 and 412 occupations depending upon the year) is reported in Column 5. This index increases the female-to-male wage ratio relative to Column 2 by between 4 and 9 percentage points. However, it is difficult to assess whether this further disaggregation would affect the timing of the relative improvement in female wages because of the short time series available. The jump in relative female wages between 1982 and 1983 still occurs, even though the change in occupational definitions should have a smaller impact at the three digit level of occupational disaggregation. The implication is that the jump in relative female wages between 1982 and 1983 may be a true change and not an artifact.11

In principle, one should control for differences in productivity characteristics between men and women. As discussed above, this could be done at great expense if one had micro data on earnings and characteristics. Lacking such data and computational resources, an approximation

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11. An obvious candidate for an explanation of the increase in relative female wages during the early 1980s would be the severe recession in blue collar sectors in this time period which would presumably have lowered wage growth in the predominantly male craft, operative, and laborer sectors relative to other occupations.
would be to regress the log of median earnings on measures of various productivity characteristics across occupational groups. Using well-known procedures, such regressions could be run separately for one sex and then used to predict earnings given the characteristics of the other sex. These predicted earnings could generate yet another set of Tornqvist indices. We have, in fact, completed some explorations in that direction.\textsuperscript{12} Although space limitations prohibit detailed discussion of our methodology and results, these results also demonstrated that female wage rates started to improve relative to male wages early in the 1970s. By 1986, these indices indicated an unexplained gap of between 15 and 22 percent. These results must be regarded as preliminary given data limitations.

The pattern of estimates in Table 2 indicates that disaggregating to the less broadly-defined occupational distribution alone explains an average of 16 percent of the difference in median male and female weekly wages over the 1983–86 period.\textsuperscript{13} A portion of the remaining gap may be associated with shorter hours worked per week (and with less work experience) by women. Even in 1986, full-time women averaged only 41.1 hours per week versus 44.7 hours per week for men (U.S. Bureau of Labor Statistics 1987, 195).

A final question to be addressed is whether these relative wage gains by women can be attributed to changes in women’s relative occupational status or to wage gains within occupational markets. The results reported in Table 1 reveal some large increases in female relative wages within

\textsuperscript{12} See Goldin and Polachek (1987) for one such application of this procedure. We regressed log earnings on educational attainment, tenure, a time trend and dummy variables for high and low skilled occupations plus post-1983 earnings. Due to data limitations, our results should be regarded as approximations to a “true” earnings equation. Nevertheless, we obtained reasonable annual rates of return to an additional year of education and tenure (5 to 8 percent). Details of our regressions and related indices of wage differentials are available on request.

\textsuperscript{13} Clearly, all the ratios in Table 2 including the commonly used ratio of median wages may be subject to selection bias since we cannot hold constant the sample of men and women in the overall labor market over time. However, the size or direction of the selection bias is not likely to change across indices since all are based on the same sample of labor force participants. Thus conclusions based on comparisons of relative movements in the wage ratios across indices are unlikely to be altered because of sample selection. We are not able to control for selectivity bias. Using micro data for 1971 and 1981, Blau and Beller (1988) find that selectivity bias does not greatly change the returns to education and experience in their earnings equation but that lack of control for selectivity does tend to understate the size of the relative gains for white women. If this result also applies to our aggregate data, it reinforces our conclusion that women gained relative to men and suggests that our index understates the size of that gain. Also if one is interested in a measure of observed relative wages for men and women, the selected sample is appropriate. Only if one wishes to obtain ratios of wage offers (including rejected offers to those electing to stay out of the ranks of the employed) will the selection issue be of concern.
individual markets, but Census data also indicate some rapid movements of women (especially young women) into traditionally male occupations starting in the early 1970s. For this reason, we reanalyzed the index in Column 3 of Table 2, holding the values of $m_{it}$ and $f_{it}$ at their 1967 levels. This in effect fixes the occupational status of women and men at their 1967 levels and creates an index whose changes are due solely to within market changes in relative wages. The index was virtually unchanged, indicating that all of the improvement in female relative wages between 1967 and 1982 was due to within market increases in relative wages and not to a shift of women into occupations with higher relative wages. This suggests that if major shifts in women’s occupational status are expected to increase female relative wages, those gains had not been realized as of 1982. Unfortunately, the change in occupational definitions limit any interpretation of the effects of occupational status on the wage index after 1982 because of the overly short time period under the new definitions.

V. Comments and Conclusions

Our results from an examination of data on occupational weekly wages indicate that considerable variance in the size and trend of the female-to-male wage ratios exists across occupations. Furthermore, the relatively slow movement upward in aggregate relative median weekly earnings and in relative median annual full-time, year-round earnings may be due to the fact that there were large (and, until the late 1970's, increasing) proportions of women in the clerical occupation, the only major occupation exhibiting a decrease in relative female-to-male wages in the 1967–1982 period. Our results show that the gains by women are attributable to increases in relative female wages within occupational markets and not to shifts in female occupational status. Finally, while the extent of unexplained wage differentials has declined since 1967, an unexplained wage gap of 26 percent still existed as of 1986.

Two implications of our findings may be noted. First, we find reason to be optimistic that females have been making relative gains in earnings, at least since the early 1970s. We cannot say whether this has been due to market, legislative, or other causes. Second, interpreted literally, these gains appear to have come within occupational groups rather than as a

result of movements into male dominated occupations. However, this conclusion is based primarily on the most broad occupational groups. It may be that female gains within these broad groups have resulted from movements into more narrowly defined male occupational categories.

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