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COMPOSITIONAL CHANGES OF THE LABOR FORCE AND THE INCREASE OF THE UNEMPLOYMENT RATE: AN ESTIMATE FOR THE U.S.

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by
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ABSTRACT.

This paper estimates how demographic changes in the composition of the labor force affect the unemployment rate in the U.S. In addition to the effect coming from changes in the weights of the individual groups comprising the total, the impact of compositional changes on the individual unemployment rates is estimated by using a dummy variables model and incorporated into the estimate upon the overall unemployment rate. The results suggest that the incorporation of the latter effect is significant. Indeed, the estimates presented in this paper are almost twice as high as those provided by the fixed-weight-method used in previous studies.

1. INTRODUCTION.

It has been argued that the principal causes of the recent high unemployment rates are aggregate factors, such as increases in the real price of oil. However, over the last decade the unemployment rate has increased irrespective to cyclical variations in economic activity. Since 1969 there has been a change for the worse in the relationship between unemployment rates and cyclical peaks, in that each downturn begins with a higher unemployment rate (O.E.C.D., 1978). Further, the relationship between the level of unemployment and unfilled vacancies has changed over time in most of the countries with available data. The data show that a strong shift occurred in almost every country around 1969-70, i.e. before the first oil shock: when unemployment rises, vacancies tend to fall less than in previous years or may even rise (O.E.C.D., 1978). These factors suggest that structural changes, rather than aggregate factors, may be an important part of the explanation of the rise in the unemployment rate over the past fifteen years.

Several explanations have been offered to the higher structural rate of unemployment (see Johnson and Layard, 1985, for a review). In this paper I shall focus on one such explanation, viz. the changing demographic composition of the labor force. In the United States the composition of the labor force has shifted towards groups with higher than average unemployment rates. In particular, there has been a large increase in the proportions of young people and females in the labor force. Whereas workers aged 16-24 constituted 16.5 percent of the labor force in 1960, they were 23.7 percent in 1980 and 21.8 percent in 1983. The proportion of median-age women in the labor force increased from 21.1 percent in 1960 to 27.9 percent in 1983. Because different age-sex components of the labor force have different skills and work expectations, a changing composition can cause upward pressure on the unemployment rate. For example, for young workers, the job-shopping process implies high rates of labor turnover and mobility as workers search for durable employment conditions. Topel and Ward (1985) have shown that for workers on their initial jobs, 63 percent of all first year job endings result in a transition to non-employment.

This paper provides an estimate of the effect of demographic changes in the labor force on the overall unemployment rate in the United States. There are two reasons to be interested in these estimates. First, the quantification of the change in the natural rate of unemployment ascribable to demographic factors can be useful in determining the extent to which employment policy initiatives must be directed to structural supply problems. Second, as Perry (1970) and Wachter (1976a) have argued, changes in the composition of the labor force and in the unemployment experience of different age-sex groups have made the aggregate unemployment rate an increasingly misleading proxy for labor-market tightness. Therefore, an estimate of the unemployment rate "purged" of shifts attributable to compositional changes could be a better indicator of the labor-market tightness than the unemployment rate itself.

To the best of my knowledge, Perry (1970) was the first person to address this subject. He computed an unemployment rate adjusted for the changing composition of the labor force. In Perry's calculation workers are heterogeneous in terms of productivity, and each worker is given a normalized weight equal to his relative productivity. More specifically, individuals are weighted according to the average hours worked and the average hourly earnings of employed persons in each demographic group.

Since Perry's study, other schemes have been suggested for adjusting the unemployment rate and measuring the compositional impact. Most of these studies use a fixed-weight-method, in which the actual unemployment rates of individual groups are weighted with a distribution of the labor force that is held constant from a given base period. Although different studies have used different procedures to hold the weights of the various labor force groups constant, all of them have a common failure: demographic variations in the composition of the labor force are only used as weighting factors and assumed to have no effect on group-specific unemployment rates[1]. This is a strong assumption in conflict with economic theory. Participation, employment and unemployment are interrelated (see Jovanovic and Shakotko, 1983) and compositional developments in the labor force affect the equilibrium relative wages and the unemployment rates of individual groups. Therefore, the fixed-weight-method fails to capture the entire compositional impact, providing measures that are clearly incomplete.

In this paper, I estimate the effect of compositional changes on individual unemployment rates, and propose a methodology to incorporate this effect in an estimate of the compositional impact on the overall rate. The paper is organized as follows. In section 2, I present a general methodological framework. Sections 3 and 4 describe the model and the estimation results. Section 5 contains the final estimates of the compositional impact on the unemployment rate, and section 6 concludes the study.

2. MEASURING THE COMPOSITIONAL IMPACT ON THE UNEMPLOYMENT RATE. GENERAL FRAMEWORK.

Dividing the working-age population into a certain number of age-sex groups, let u_t be the overall unemployment rate in period t. Then

$$u_t = \sum_i l_{it}u_{it} \tag{1}$$

where u_{it} is the unemployment rate in the demographic group i in period t and l_{it} is the number of persons in the labor force in the i^{th} group in period t expressed as a fraction of the total labor force. The aggregate unemployment rate is expressed as a weighted average of the unemployment rates in demographic groups. The weights are the proportions of the labor force corresponding to each group.

Since the objective is to decompose the change in the unemployment rate into (i) a component due to changes in labor force composition and (ii) a component due to other factors, a period of time needs to be taken as a benchmark (denoted by subscript t = 0). The change in the observed unemployment rate from period "0" to period "t" can be written as

$$\Delta u_t = u_t - u_0 = \sum_i (\Delta l_{it})u_{i0} + \sum_i (\Delta u_{it})l_{i0} + \sum_i (\Delta l_{it})(\Delta u_{it}) \tag{2}$$

Variations in the unemployment rate are due to (i) changes in the demographic percentage distribution of the labor force and (ii) changes in the unemployment rates of individual groups. The first term on the right hand side of equation (2), $\sum (\Delta l_{it})u_{i0}$, is the part of the total change in the unemployment rate directly attributed to changes in labor force composition. The second term, $\sum (\Delta u_{it})l_{i0}$, is the part of the total change attributed to variations in the unemployment rates of individual groups. The last term,

$\sum (\Delta l_{it})(\Delta u_{it})$, is the interaction or joint effect between the changes in the composition of the labor force and the changes in the unemployment rates of individual groups.

The objective of previous studies (Flaim, 1979, Kaufman, 1980 and Podgursky, 1984) has been to separate the effect coming from changes in the age-sex-specific shares of the labor force and the effect resulting from changes in the age-sex-specific unemployment rates[2]. The weakness of this method is that demographic variations in the composition of the labor force are assumed to have no effect on group unemployment rates. Since participation is one of the determinants of unemployment, group-specific unemployment rates will be affected by compositional changes in the labor force[3]. For example, as women come to represent a larger fraction of the labor force, an effect on the female unemployment rate can be expected. Therefore, the method above provides an incomplete estimate of the labor force composition component.

To estimate the compositional effect on the overall unemployment rate, two separate components need to be included. First, the direct effect, which is just the first right-hand term of equation (2). Second, the indirect effect of changes in labor force composition on the overall unemployment rate through their impact on group unemployment rates.

Rewrite equation (2) as

$$(2') \qquad \Delta u_t = \sum_i (\Delta l_{it}) u_{i0} + \sum_i (\Delta u_{it}) l_{it}$$

Let $\Delta\tilde{u}^L_{it} = (\tilde{u}^L_{it} - \tilde{u}^L_{i0})$ be the predicted change of the i^{th} group's unemployment rate ascribable to the change in labor force composition. Then,

$$(3) \qquad \Delta\tilde{u}^L_t = \sum_i (\Delta l_{it}) u_{i0} + \sum_i (\Delta\tilde{u}^L_{it}) l_{it}$$

becomes the overall estimate of the amount of change in the observed unemployment rate, from period "0" to period "t", attributable to compositional variations in labor supply.

The methodology employed to estimate $\Delta\tilde{u}^L_{it}$ is described in the next section.

3. THE DUMMY VARIABLES MODEL.

As equation (3) reveals, the problem has been reduced to one of decomposing the change in the unemployment rates of the individual groups into a predicted component ascribable to the change in labor force composition and a predicted component ascribable to the change in general economic conditions. General economic conditions include any other factors, cyclical or structural, affecting unemployment rates other than the change in the age-sex composition of the labor supply.

To carry out such a decomposition, the unemployment rate of group i in period t could be written as the sum of a group-specific component (a_i), a time-specific component (b_t), and a component depending on the fraction of the labor force comprised by group i in period t ($\lambda_i l_{it}$),

$$u_{it} = m + a_i + b_t + \lambda_i l_{it} + v_{it} \tag{4}$$

for $i=1,2,\ldots,I$ and $t=1,2,\ldots,T$, where v_{it} is the i^{th} group's disturbance term in period t.

This model is a dummy variables or analysis of covariance model, in which the coefficient of the continuous variable, λ_i, is allowed to vary over age-sex groups. The intercepts $m_{it} = m + a_i + b_t$ are assumed to be fixed parameters which, along with the slope coefficients, λ_i, need to be estimated. To carry out the estimation problem, the following dummy variables are defined:

W_i = 1 for the i^{th} age-sex group, $i = 1,2,\ldots,I$
0 otherwise

Z_t = 1 for the time period t, $t=1,2,\ldots,T$
0 otherwise

Also, restrictions such as $\sum a_i = 0$ and $\sum b_t = 0$ need to be imposed, and the equation must be reparameterized so that the restrictions are incorporated. One such reparameterization is to define the new variables $W_i^* = W_i - W_I$, for $i=1,2,\ldots,I-1$, and $Z_t^* = Z_t - Z_T$, for $t=1,2,\ldots,T-1$[4]. The model becomes

$$(5) \qquad u_{it} = \mu + \sum_{i=1}^{I-1} \alpha_i W_i^* + \sum_{t=1}^{T-1} \beta_t Z_t^* + \sum_{i=1}^{I} \lambda_i W_i l_{it} + v_{it}$$

where $u_{it} = (u_{1t}, u_{2t}, \ldots, u_{It})'$ is of dimension ($IT * 1$). The parameters to be estimated are: μ, the "general mean" that is constant over time and individual groups; α_i, representing the difference from the general mean for the i^{th} age-sex group; β_t, representing the difference from the general mean for the time period t; and λ_i, the slope coefficient relating the fraction of the labor force comprised by group i to that group's unemployment rate. After estimating equation (5), α_I and β_T can be calculated from the two restrictions imposed in the coefficients.

The linear formulation of the model in (5) permits the straightfoward decomposition of the predicted change in each group i's unemployment rate into the two desired effects, $\tilde{u}_{it} - \tilde{u}_{i0} = (\tilde{\beta}_t - \tilde{\beta}_0) + \tilde{\lambda}_i (l_{it} - l_{i0})$, i.e. a term ascribable to the change in general economic conditions, $(\tilde{\beta}_t - \tilde{\beta}_0)$, and a term ascribable to the change in labor force composition.

However, this linear formulation (i) allows for negative predicted unemployment rates and (ii) ignores the truncation of the dependent variable ($0 < u_{it} < 100$).

Therefore, the logit model

$$(6) \qquad u^*_{it} = \mu + \sum_{i=1}^{I-1} \alpha_i W^*_i + \sum_{t=1}^{T-1} \beta_t Z^*_t + \sum_{i=1}^{I} \lambda_i W_i l_{it} + \epsilon_{it}$$

with $u^*_{it} = \log [u_{it}/(100-u_{it})]$ was estimated. u^*_{it} satisfies $-\infty < u^*_{it} < \infty$, and thus is not truncated[5]. This transformation also excludes the possibility of obtaining negative predicted unemployment rates.

Once the parameters in (6) have been estimated, the predicted unemployment rate for group i in period t is

$$\hat{u}_{it} = \left[100 * EXP(\tilde{\mu}_{it} + \tilde{\lambda}_i l_{it})\right] / \left[1 + EXP(\tilde{\mu}_{it} + \tilde{\lambda}_i l_{it})\right]$$

where $\tilde{\mu}_{it} = \tilde{\mu} + \tilde{\alpha}_i + \tilde{\beta}_t$.

Hence,

$$(7) \qquad \Delta\tilde{u}^L_{it} = \hat{u}_{it} - \left[100 * EXP(\tilde{\mu}_{it} + \tilde{\lambda}_i l_{i0})\right] / \left[1 + EXP(\tilde{\mu}_{it} + \tilde{\lambda}_i l_{i0})\right]$$

is the predicted change in the unemployment rate of the individual group i, for each i, from period "0" to period "t", ascribed to the change in the composition of the labor force.

An interpretation of the parameters in equation (6) and a description of the estimation method are given in the next section.

4. ESTIMATION OF THE MODEL.

For estimation purposes the working-age population has been divided into eight groups, males and females of ages 16-19, 20-24, 25-54 and 55 and over. Equation (6) is estimated for the sample period 1948-83. Data are quarterly and seasonally adjusted. For simplicity the time-dummy variables are chosen to be annual.

Before analyzing the results, some features regarding the estimation method employed should be mentioned. In section 2 it was indicated that a feedback is expected between the unemployment rates of the individual groups and the fraction of the labor force comprised by each group. If this is the case the least squares estimator will be biased and inconsistent. To achieve consistency instrumental variables must be used. Since there is no other variable in the model that is highly correlated with l_{it} and at the same time contemporaneously uncorrelated with the disturbance, the choice of instruments is limited to lagged values of l_{it}. In addition, preliminary checks of the residuals indicate the presence of time-series autocorrelation in the pooled model. This creates an estimation problem in which two assumptions of the classical linear model are violated simultaneously -autocorrelated errors and a regresor contemporaneously correlated with the disturbances.

To deal with both problems a variant of the generalized least squares estimation procedure is applied. It is essentially an adaptation to this specific problem of both the three-stage-method suggested by Wallis (1967) and the method involving predetermined, but not exogenous, instruments due to Hayashi and Sims (1983). With this procedure asymptotically efficient and consistent estimators are achieved. Also, the autocorrelation coefficients of each autoregression are allowed to vary from individual group to individual group. The estimation technique can be summarized as follows:

In the first stage the entire pooled model is estimated using instrumental variables. The instruments chosen are the second and third lagged values of l_{it} and a time trend. The first lag has been dropped due to the presence of first order autoregression schemes in each of the time-series disturbances.

Since these parameter estimates are consistent, in the second stage they are used to calculate the regression residuals. Since the filter $(1-\rho_i B)$, where B is the lag operator, eliminates serial correlation in ϵ_{it}, the "forward" filter $(1-\rho_i B^{-1})$ will also eliminate serial correlation. Therefore, regressions such as $(1-\rho_i B^{-1})\epsilon_{it}=a_{it}$ have been run to estimate consistently the autocorrelation coefficients corresponding to each of the time-series autoregressive processes.

In the stage three, using these autocorrelation coefficients, the transformation of the dependent and independent variables is constructed by using the filter $(1-\rho_i B^{-1})$, and the generalized difference form of the pooled model is estimated.

Note that the "forward" filter will be an exponentially weighted sum of non-positive powers of B. Thus, a_{it} depends only on ϵ_{it+s} for non-negative s. Since the instruments used are the second and third lags of l_{it}, ϵ_{it+s} for non-negative s will be uncorrelated with the filtered instruments. Therefore, consistent estimates can be obtained by applying least squares to the transformed model.

The equation estimates are presented in Table A.1 in the Appendix. Table A.2 summarizes the results for the white noise diagnostic check of the time-series residuals. Table A.3 shows joint tests of significance for certain subsets of coefficients. The following should be emphasized:

(1) The coefficients of the time-dummy variables measure the differential effect of the particular economic and social conditions prevailing in each year upon the logit transformation of the unemployment rate. The differential effect is measured with respect to the average impact over the whole sample. Over the 1950s, a period of economic recovery, the differential effect is negative. It becomes positive in the early 1960s, and turns negative again during the Vietnam war period. After 1973, the coefficients of the time-dummy variables are positive, reflecting the impact of the recession on the unemployment rate.

(2) The effect associated with the different demographic groups relative to the general mean is measured by the coefficients of the group-dummy variables. These parameter estimates also have the expected sign. Male and female teenagers (16-19 years old) have a positive and significant differential effect. This indicates that, although unemployment rates of different demographic groups move together in response to fluctuations in economic activity, the levels about which they fluctuate are much higher for the younger workers. The coefficient for median age women, though positive, is not significant. The reason is that the unemployment rate for older females is close to the economywide level. Finally, the adult male groups have negative coefficients relative to the general mean.

(3) The λ_i coefficients are a measure of the effect of changes in the percentage of workers in group i upon the logit transformation of the unemployment rate of the same group. For male and female teenagers, females aged 20-24 and older male workers, the coefficients are positive and significant. The coefficient for women aged 55 and over is negative. This might be explained by the fact that the trend of the percentage of older females in the labor force and the trend of their rate of unemployment have

opposite signs over the whole sample period. Finally, it is interesting to note that the effect of the increasing proportion of adult females (25-54 years of age) on that group's unemployment rate is not significantly different from zero.

(4) In addition to the t-test associated with individual coefficients, F-tests for the two main effects have been carried out. First, the joint hypothesis H_0: $\beta_t = 0$, for all t, is tested. Accepting the null hypothesis means that the different economic conditions in each year do not significantly explain variations on the unemployment rates of the individual groups. The hypothesis has been rejected at the 1 percent level of significance. Second, the null hypothesis H_0: $\alpha_i = 0$, for all i, has also been rejected at 1 percent level of significance. Differences among demographic groups are significant in explaining unemployment rates.

5. ESTIMATES OF THE COMPOSITIONAL EFFECT ON THE OVERALL UNEMPLOYMENT RATE.

Once equation (6) is estimated, the predicted compositional changes in the unemployment rates of the individual groups, $\Delta\hat{u}^L_{it}$, for all i, are derived from equation (7). The estimate of the compositional impact on the overall unemployment rate is then computed as indicated in equation (3).

The computation of equations (7) and (3) requires a benchmark. Instead of a single initial year, decade averages have been used as benchmarks. This avoids the possibility of misleading results that would arise from using years with abnormally low or high unemployment.

Table 1 presents the estimates computed from equation (3) using the decade 1950-59 as a benchmark. The first column of the table shows the percentage-point change in the unemployment rate relative to the average rate that prevailed in the benchmark period. Estimates of the total compositional impact are presented in column 4. These estimates measure the impact of the change in the make-up of the labor force on the percentage-point change in the unemployment rate. For example, the estimate of the compositional impact in 1974 is 0.73, which means that a 0.73 percentage point of the increase in the overall unemployment rate from its average level in the 1950s to its level in 1974 is attributable to the change in the composition of the labor force. The total estimated effect presented in column 4 is the summation of the two partial effects reported in the second and third columns of the table. The second column shows the partial effect corresponding to the first-right-hand term of equation (3), which is in fact the incomplete fixed-weight-method estimate. The estimates of the second-right-hand term of equation (3), which correspond to the effect of demographic changes on the overall rate through their impact on group unemployment rates, are reported in column 3.

* (Table 1)

From the estimates presented in Table 1 the following points should be noted:

(1) The estimates of the compositional effect on the unemployment rate are all positive over the sample period.

(2) The compositional impact on the unemployment rate is very small in the early 1960s. During those years the distribution of the labor force did not change significantly in comparison with the 1950s. The impact increases slightly in the late 1960s when the maturing of the post-World-War baby boom started pushing up the percentage of the labor force comprised by the youth.

(3) After 1970, thé compositional effect rose considerably and peaked in the mid 1970s. The net effect was about 0.8 percentage point of the increase

TABLE 1. ESTIMATES OF THE DEMOGRAPHIC IMPACT IN THE CHANGE OF THE UNEMPLOYMENT RATE. BENCHMARK PERIOD: 1950-59.

	%-point change in u_t relative to the benchmark period	$\sum_i (\Delta l_{it}) u_{i0}$	$\sum_i (\Delta \tilde{u}^L_{it}) l_{it}$	effect of compositional changes on Δu_t
	Δu_t (1)	(A) (2)	(B) (3)	$\Delta \tilde{u}^L_t = (A+B)$
1960	1.2	.07	-.01	.06
1961	2.3	.08	-.01	.07
1962	1.2	.08	-.01	.07
1963	2.3	.11	.00	.11
1964	0.8	.14	.01	.15
1965	0.1	.19	.04	.23
1966	-0.6	.25	.08	.33
1967	-0.6	.25	.08	.33
1968	-0.9	.25	.08	.33
1969	-0.9	.28	.10	.38
1970	0.6	.32	.14	.46
1971	1.5	.35	.17	.52
1972	1.2	.40	.21	.61
1973	0.5	.43	.23	.66
1974	1.2	.45	.28	.73
1975	4.1	.49	.34	.78
1976	3.3	.45	.33	.78
1977	2.6	.47	.34	.80
1978	1.7	.47	.32	.79
1979	1.5	.45	.30	.75
1980	2.8	.42	.28	.70
1981	3.2	.39	.26	.64
1982	5.3	.34	.23	.57
1983	5.2	.30	.19	.49

Source: derived from equation (3).

Notes: (1) The average observed unemployment rate over the 1950s was 4.4 percent. The observed unemployment rate in each year can be calculated by adding 4.4 to the figures in the first column.

(2) Fixed-weight-method estimate of the effect of demographic changes on Δu_t.

(3) Estimate of the effect of demographic changes on Δu_t through their impact on group unemployment rates.

in the overall unemployment rate over the 1975-78 period, accounting for about a 25-30 percent of the total change in the rate. The reason for this result is that the percentage of the labor force comprised by young workers -16 to 24 years old- also peaked in the mid-1970s .

(4) After 1978 the compositional impact shows a relatively sharp decline, even though observed unemployment rates were higher in this period. This is due to the decline in the absolute number of youths, which in turn caused their share in the labor force to drop drastically. The compositional effect, however, continues to reflect the persistent rise in participation among young women.

(5) The pattern of both the estimates of the compositional impact over time and the percentage of the labor force comprised by teenagers and young workers indicates that a large part of the compositional impact on the unemployment rate over the last decade stems from the gradual entry into the labor force of the millions of youths born during the baby boom. The large increase in the labor force participation rate of adult women and their increasing relative position in the labor force, which often have been blamed for boosting the overall unemployment rate, did not, in fact, have much of an impact. The reason is that the unemployment rate for adult women has generally been slightly lower than the overall rate.

(6) The estimates obtained using the methodology presented in section 2 are almost twice as high as those obtained if demographic changes are assumed to have no effect on the unemployment rates of individual groups (column A in Table 1). The incorporation of the effect of compositional changes on the unemployment rates of individual groups is not just a theoretical issue. Indeed, over the sample period the incorporation of this effect accounts for up to forty percent of the whole compositional impact. Figure 1 shows plots of

both the fixed-weight-method figures and the complete estimate derived from equation (3). Obviously the fixed-weight-method leads to a considerable underestimation of the compositional impact. It is also important to note that the estimates provided by the fixed-weight-method are increasingly misleading over time. In the late 1960s the fixed-weight-method underestimation was only about 25 percent with respect to the estimate provided by my methodology while in the mid-1970s the underestimation increased to 40 percent.

* (Figure 1)

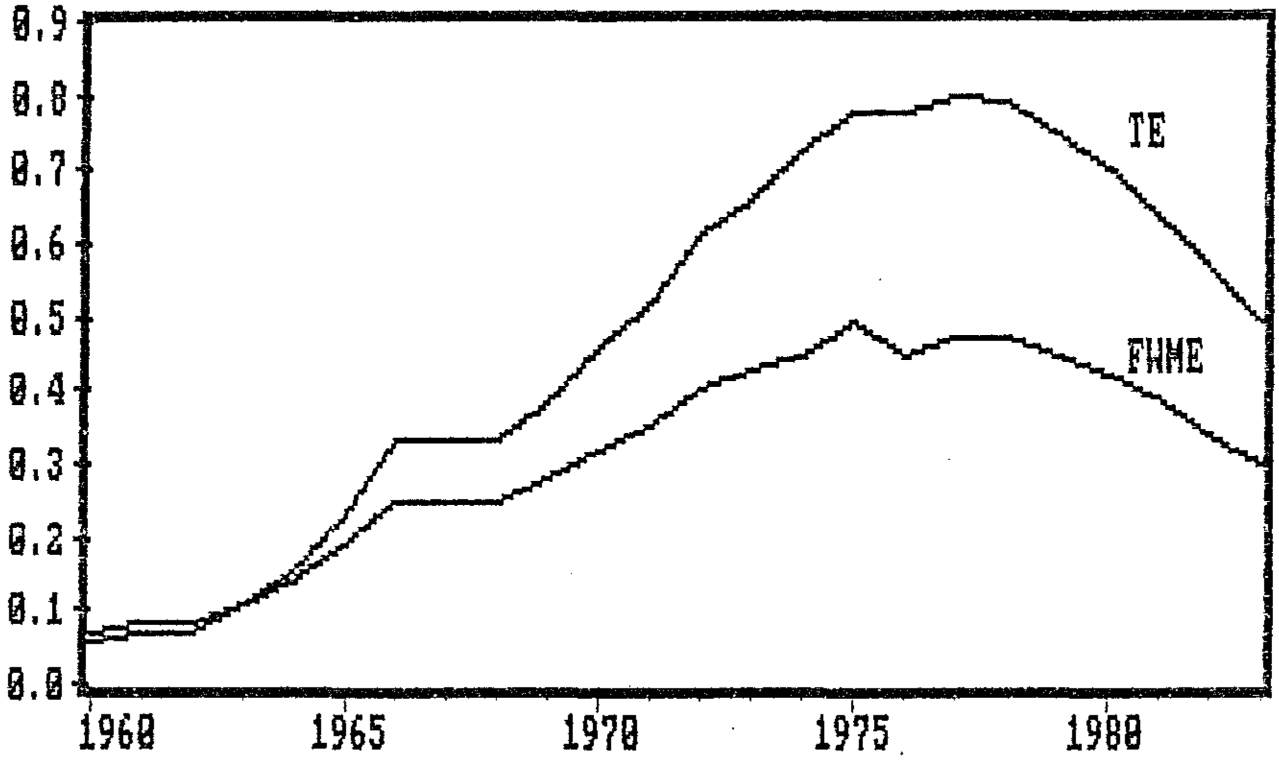
6. SUMMARY AND CONCLUSIONS.

In this paper I have estimated the part of the change in the unemployment rate that can be attributed to changes in the age-sex composition of the labor force in the United States for the period since 1960.

Since the late sixties two facts have characterized the changing makeup of the labor force: the increasing number of young males and females in the labor force as a direct consequence of the baby boom, and a dramatic rise in participation rates among adult women. Both of these have been blamed for exerting an upward pressure on the overall unemployment rate.

Previous studies provide an incomplete estimate of the compositional effect, since they restrict themselves to movements in the overall unemployment rate coming from changes in the weights of the individual groups comprising the total. Movements coming from changes in the unemployment rates of the individual groups, occurring as a consequence of changes in participation, have been totally ignored. This later effect is large and so should not be neglected. I have proposed a methodology by using a dummy variables model to incorporate the impact of demographic changes upon the

FIGURE 1. Total and fixed-weight-method estimates



TE: total estimate from equation (3).

FWME: Fixed-weight-method estimate.

individual unemployment rates in an estimate of the compositional impact on the overall rate. It has been shown that the incorporation of this effect is not just a theoretical issue and that the fixed-weight-method calculations substantially underestimate the compositional impact. Indeed, the effect of demographic changes on individual unemployment rates might account up to forty percent of the whole compositional impact on the overall unemployment rate.

The estimation results based on detailed data from the Department of Labor show that the compositional effect, although low during the 1960s, increased considerably in the early 1970s and peaked in the mid-1970s. The net effect over the 1975-78 period was about 0.8 percentage point of the increase in the overall unemployment rate. For example, if the average of the unemployment rate observed over the 1950s was 4.4 percent and in 1977 was 7.0 percent, 0.8 of the 2.6 percentage points of such increase is attributable to changes in labor force composition. After 1978 the compositional impact declines.

Only a relatively small part of the compositional impact on the unemployment rate stems from changes in participation rates of adult women. The demographic effect on the unemployment rate can be traced primarily to the gradual entry into the labor force of millions of youths born during the baby boom. This result is supported by two factors. First, the decreasing compositional impact after 1978 is accompanied by two observed facts, a sharp decline in the youth population in the labor force, and a continuing rapid increase in participation rates of adult women. Second, the unemployment rate for adult women has generally been slightly lower than the overall rate. These results suggest that the demographic effects on the unemployment rate may well have peaked in the late 1970s, and its declining trend is expected to continue.

FOOTNOTES.

1) See Kaufman (1980) and Podgursky (1984). Also the Bureau of Labor Statistics (BLS) has carried out detailed calculations subdividing the total compositional impact into its two principal components -the part stemming from changes in the age structure of the population and the part stemming from changes in the labor force participation rates of the various groups. For results of thr BLS study, see Flaim (1979). Wachter (1976a) and (1976b) has used a different approach. He focused on measuring the demographic impact on what he called the "noninflationary rate of unemployment". However, his method is based on two strong assumptions: (i) the unemployment rate among prime-age males -25 to 34 years of age- is held constant at a full-employment rate, and (ii) the "crowding" of the baby-boom youth into the labor force is considered to be the single cause of all the imbalances in the composition of the labor supply.

2) The general method used to calculate the impact of demographic changes on the overall unemployment rate has been to construct an alternative hypothetical "shift-free" rate. This is done essentially by applying the actual distribution of the labor force to the age-sex-specific unemployment rates which occurred in a given benchmark period. The difference between this hypothetical rate and the published overall rate in the initial period is attributed to the compositional impact.

3) And viceversa. The distribution of the labor force may be affected by the magnitude of the unemployment rates. When the unemployment rate of an individual group is high, workers in this group might become discouraged and drop out of the labor force. The same reasoning applies to encouraged workers when unemplyment rates are lower.

4) There are other alternative procedures. The advantage of the reparameterization imposed in equation (5) is that the interpretation of the coefficients become straightforward. Each of the dummy variable coefficients has intuitive meaning independent of the categories of the remaining dummy variables in the model.

5) An alternative formulation was also considered, $u_{it}^{**} = \log(u_{it})$, where $-\infty < u_{it}^{**} < \log 100$. Although this transformation does not solve completely the problem of truncation of the dependent variable, the exponential model has also been estimated. The estimation results were similar to those reported here.

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APPENDIX.

TABLE A.1 PARAMETER ESTIMATES FOR THE DUMMY VARIABLES MODEL.
Dependent variable: $u^*_{it} = \log[u_{it}/(100-u_{it})]$

COEFFICIENTS	PARAMETER ESTIMATE	t FOR H_0 PARAMETER = 0
μ	-2.92*	37.62
β_{1950}	-0.08*	-3.14
β_{1951}	-0.37*	-12.32
β_{1952}	-0.49*	-15.87
β_{1953}	-0.53*	-16.94
β_{1954}	-0.04	-1.56
β_{1955}	-0.15*	-4.89
β_{1956}	-0.16*	-5.53
β_{1957}	-0.15*	-5.06
β_{1958}	0.17*	5.93
β_{1959}	0.09*	3.00
β_{1960}	0.06*	2.13
β_{1961}	0.21*	7.25
β_{1962}	0.09*	3.39
β_{1963}	0.09*	3.29
β_{1964}	0.02	0.61
β_{1965}	-0.09*	-3.50
β_{1966}	-0.25*	-9.15
β_{1967}	-0.26*	-9.32
β_{1968}	-0.30*	-11.01
β_{1969}	-0.32*	-11.62
β_{1970}	-0.08*	-2.97
β_{1971}	0.03	1.12
β_{1972}	0.01	0.45
β_{1973}	-0.08*	-2.83
β_{1974}	0.02	0.97
β_{1975}	0.35*	11.75
β_{1976}	0.29*	9.64
β_{1977}	0.22*	7.49
β_{1978}	0.10*	3.47
β_{1979}	0.08*	2.88
β_{1980}	0.20*	6.41
β_{1981}	0.30*	9.71
β_{1982}	0.49*	15.10
β_{1983}	0.52	
$\alpha_{males(16-19)}$	0.87*	4.62
$\alpha_{males(20-24)}$	0.38*	2.11
$\alpha_{males(25-54)}$	-0.95*	-2.21
$\alpha_{males(55+)}$	-1.50*	-8.93

COEFFICIENTS	PARAMETER ESTIMATE	t FOR H_0 PARAMETER = 0
$\alpha_{females(16-19)}$	0.57*	3.71
$\alpha_{females(20-24)}$	0.20	1.95
$\alpha_{females(25-54)}$	0.09	0.56
$\alpha_{females(55+)}$	0.34	
$\lambda_{males(16-19)}$	7.14*	2.15
$\lambda_{males(20-24)}$	2.71	0.94
$\lambda_{males(25-54)}$	-1.41	-1.23
$\lambda_{males(55+)}$	10.19*	7.74
$\lambda_{females(16-19)}$	17.08*	3.67
$\lambda_{females(20-24)}$	6.35*	3.52
$\lambda_{females(25-54)}$	0.49	0.59
$\lambda_{females(55+)}$	-13.29*	-5.04
$\rho_{males(16-19)}$	0.46*	5.84
$\rho_{males(20-24)}$	0.74*	13.27
$\rho_{males(25-54)}$	0.81*	16.73
$\rho_{males(55+)}$	0.58*	8.54
$\rho_{females(16-19)}$	0.70*	11.13
$\rho_{females(20-24)}$	0.56*	7.51
$\rho_{females(25-54)}$	0.41*	5.11
$\rho_{females(55+)}$	0.51*	6.89
Adj. R^2	0.99	

Notes: (1) Estimation results from equation (6).
(2) Quartely unemployment and labor force data are from the U.S. Bureau of Labor Statistics. Handbook of Labor Statistics.
(3) One asterisk indicates coefficient significantly different from zero with 99 percent confidence.

TABLE A.2 ANALYSIS OF THE RESIDUALS. AUTOCORRELATION CHECK FOR WHITE NOISE. REGRESSION FOR THE DUMMY VARIABLES MODEL.

Null hypothesis: the time series residuals are white noise

cross-section units	calculated CHI square
group 1: males(16-19)	21.69
group 2: males(20-24)	29.92
group 3: males(25-54)	28.12
group 4: males(55+)	36.08
group 5: females(16-19)	35.99
group 6: females(20-24)	25.56
group 7: females(25-54)	33.55
group 8: females(55+)	13.90

Notes (1) The CHI square values are calculated for 24 degrees of freedom.

(2) The critical CHI square level is 36.4 at the 5 percent level of significance. The null hypothesis of white noise is accepted with 95% confidence for all the eight individual regressions.

TABLE A.3 TESTS OF HYPOTHESES REGRESSION FOR THE DUMMY VARIABLES MODEL.

Null hypothesis	Calculated F value	F value Prob.>F	Result
H_0: $\beta_t=0$, $t=1950,\ldots,1983$	39.02	.0001	H_0 rejected
H_0: $\alpha_i=0$, $i=1,2,\ldots,8$	16.91	.0001	H_0 rejected