WORKING PAPERS

Number 7801

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January 1978

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A CAUSALITY APPROACH TO LABOR FORCE PARTICIPATION: TESTS FOR STRUCTURAL CHANGE*

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Economists have long been interested in the labor supply decisions of various demographic groups in the United States because of the importance of these decisions to local labor market characteristics and to the success of income security programs. In addition, there is governmental concern for minimizing the overall rate of unemployment. With regard to the latter, the macropolicy arena, an impressive amount of labor supply research has focused on the role of women in the labor force and the degree to which policymakers need to regard the labor supply decisions of women as different from those of men.

This paper reexamines a labor supply issue that initially received extensive airing in the mid-1960's: the relationship between labor force participation rates and unemployment rates. Three aspects of this paper are novel: first, a Box-Jenkins causality approach is employed in place of the customary regression framework; second, the time period covered begins in 1954 and runs through 1976, a period in which large compositional changes in the labor force occurred; third, an explicit test for a change in time series models is applied to the participation rates of both adult white males and females.

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*The authors wish to thank various colleagues for helpful discussion and Pierretta Hughes for the careful typing of the manuscript.

The short-period labor force supply decisions of adult males and females generally appear to be unrelated to either the supply decisions of other groups or to their own or others' unemployment rates. Short-period supply decisions of teenagers, however, do seem related to teenage unemployment rates. These findings, based on a different technique, support a conclusion expressed by Bowen and Finegan [1969] regarding time series of this type. The second principal finding is that the labor force participation behavior of adult white males has changed over the 1954-1976 period, while that for adult white females has not. The meanings of "behavior" and "changed" will be made clear subsequently. Following the introduction of the causality methodology two primary findings are presented. The implications of the results are given in the last section of the paper.

THE CAUSALITY APPROACH

Let X and Y represent measurements of an input and an output, respectively, of an economic process across time (i.e., X_t, $X_{t + 1}$, $X_{t + 2}$, . . . $X_{t + k} = X$). Series X is said to cause series Y if a knowledge of X will reduce the error of predicting Y when only past values of Y are known. A model which describes the dynamic response of Y to changes in X and which can be used to study causality is called a transfer function model by Box and Jenkins [1970]. By presuming that the labor force participation rate of some group is an output of a decision process and that the labor force participation rate or unemployment rate of some other group (or the unemployment rate of the first group) is an input to that process, a test can be conducted to determine whether a transfer function model can relate two such time series.

In relating two time series, they are first represented as mixed autoregressive integrated moving average models as described by Box and Jenkins [1970]. These models are then used to obtain residual series, which are white noise (a sequence of uncorrelated error terms). After this process of prewhitening, the residual series are cross correlated to test for independence. By prewhitening, serial correlation up to a very high order is removed from the data, permitting more efficient estimates of the standard errors of the cross correlation function. Prewhitening can be done by two common procedures. Box and Jenkins report on the prewhitening of both data series by one filter, while Haugh [1976] presents a method of prewhitening each series separately. The Haugh method is more appropriate for testing for causality, and it is briefly discussed in Appendix One.

Using the Haugh technique, the filtered data series are cross correlated, showing the degree of association between x_t and y_{t+k}, where k is an integer. By comparing the size and pattern of cross correlation estimates, one of five possible relationships can be found between X and Y. X causes Y if the cross correlations between x_t and y_{t+k} are significant only for $k > 0$, Y causes X if significance is found for $k < 0$, and X and Y are contemporaneously correlated if significance is found only at $k = 0$. If the cross correlations are significant for lags on either side of $k = 0$, feedback is said to occur, and if no significant cross correlations are found, the two series are said to be independent.

One-way causality is a necessary and sufficient condition for a transfer function model to exist between two series. In this study no attempt has been made to fit transfer function models. Consequently, in those instances in which unidirectional causality was found between two time series, there is no way to sign the relationship. Short of fitting the transfer function model, the results can only be used to test the hypothesis that two time series are independent of each other and to indicate direction of causality, if present. Nevertheless, the inability to reject a hypothesis of independence between labor force series, such as those examined here, is a strong finding.

DATA AND HYPOTHESIS TESTS

Cross Correlations

Monthly labor force participation rates and unemployment rates were collected for white males and females, aged 16 to 19 years and 20 years-old and older. The data were not seasonally adjusted. Data for the adults pertain to the 1954-1976 time period, while teenage data covered the 1967-1976 period only. Separate autoregressive integrated moving average (ARIMA) models were fit to the adult data for the 1954-1966 and 1967-1976 periods. The data were broken into these periods to minimize the effect of changes in labor force definitions in 1967 and to obtain two periods of roughly equal size. In all, 12 different ARIMA models were fit: eight for adults and four for teenagers. The ARIMA models appear in Appendix Two.

To characterize the nature of the relationship between two series, two separate tests can be applied to the cross correlations. Individual cross correlations can be compared with their approximate standard errors--a test roughly equivalent to a t test in linear regression. Results must be interpreted carefully because there are 61 cross correlations for each pair of series, and about three would be significantly different from zero by chance at the 95 percent level, even if the series were independent. The entire set of cross correlations at positive and negative lags can be hypothesized as a group to be insignificantly different from zero--a test roughly analogous to an F test in linear regression. Technically, the group test statistic is approximately distributed as a chi-square. Both of these tests were applied to the cross correlations because each is sensitive to slightly different behavior of the cross correlations.

Independence Tests

Test results are reported by demographic groups. The groups examined are white males and females, aged 20 and over, and white males and females, 16 to 19 years. The results indicate which groups' short-run labor supply decisions can be usefully regarded as dependent, on an aggregate level, on the labor supply decisions of another group or as dependent on an aggregate measure of labor market tightness.[1]

The causality tests considered were of two kinds. The first kind was between unemployment rates and labor force participation rates and between participations rates for one group and participation rates for another group. Participation rates may be interpreted as evidence of labor supply decisions but unemployment rates are more difficult to interpret. The unemployment rate, after all, contains information on both labor supply and labor demand.

Labor force participation rates are not informative, per se, of labor market tightness; increases in labor force participation are consistent with increases in the unemployment rate for any group. Suppose that male participation rates are found to cause female participation rates. Even if the sign of the relationship were known to be negative, there could be two explanations. On the one hand, increases in male employment could be causing females to withdraw from the labor force. On the other hand, increases in male unemployment could be discouraging to females, causing them to withdraw from the labor force. To gain insight into the actual labor market flows and the causal relationships between them, the participation-participation tests must be examined in conjunction with the unemployment-participation tests.

In a study of time series relationships between the labor force and the overall unemployment rate Bowen and Finegan concluded that:

> We have found no convincing evidence in the postwar record that short-period changes in the overall rate of unemployment have had a large impact on the labor force participation rate of any population group other than teenagers and possibly males 65+.[2]

Cross correlation tests applied to the data series in this study point to the same conclusion: the cross correlations of the whitened series generally were independent of each other. In all but one instance, the exceptions occurred among teenagers.

The independence test results appear in Appendix Two.[3] For the 1954-1966 time period, the adult male unemployment rate was found to "cause" adult male participation rates. All other unemployment rates were found to be independent of participation rates. The participation rates of adult males and females were contemporaneously related.

For the 1967-1976 time period, all relationships between adult series were found to be independent. Only among teenagers were participation rates independent of each other; participation rates were found to respond to an unemployment rate in three out of four cases. The relationships between teenagers and adults were less clear. By and large these results indicate that adult participation rates and teenage participation rates are independent, while adult unemployment rates and teenage participation rates are related in complicated patterns. Wachter [1972] found that unemployment rates were not significant variables in participation rate regressions. Contrary to his expectations, Smith [1977] found a positive relation between labor market tightness and the probability of labor market withdrawal for prime-age males and females.

While many economists believe that market tightness and participation are positively related--the discouraged worker effect dominates--this contention is hard to support empirically. The findings of Wachter, Smith, and Fleisher and Rhodes [1977], combined with the findings presented here, indicate that the relationship between participation rates and unemployment rates cannot be described as one of cause and effect in any meaningful sense.

Care must be taken in interpreting these results of no causality because there are causality behaviors which cannot be determined by statistical tests. For example, if two series are only straight-line trends, there is no statistical method of proving that one series is causing the other because the trend in one series could be just a trend, independent of the other series, or it could be caused by the trend in the other series. Thus, the conclusions presented here hold for short run changes and not necessarily for long term trends.

These results of weak or no relationships between economic variables which were thought to be related are consistent with the results in Pierce [1977]. As demonstrated here, Pierce found that when the correct method of testing was used relationships were much weaker than had been believed.

Structures of Models

The results of the independence tests suggest that whatever conditions are associated with trend changes in the labor force participation of adult white males and females, the short-period labor force activities of the two groups are unrelated in general. But when the models for long-term participation rates of adult white males and females are compared for the 1954-1966 and 1967-1976 time periods, the participation rate models of adult white males have changed by more than those of adult white females.

Participation Rates. For the 1954-1966 period female LFPRs are modeled as $\nabla\nabla_{12}Z_t = (1 - .155B + .147B^{11} - .916B^{12})a_t$, while the 1967-1976 period yields $\nabla\nabla_{12}Z_t = (1 - .904B^{12})a_t$. These models are roughly identical except for the small first and eleventh order moving average parameters. The 1954-1966 estimation period provided a larger number of observations to fit than the later period and yielded sharper estimates of the parameters. The 1954-1966 model structure was fit to the 1967-1976 data with this result: $\nabla\nabla_{12}Z_t = (1 - .173B + .129B^{11} - .913B^{12})a_t$. The coefficients obtained in the fit to 1954-1966 data (.155, -.147, .916) are each in the 95 percent confidence intervals around the 1967-1976 estimates. The null hypothesis that the 1954-1966 model is also an acceptable model for the 1967-1976 period cannot be rejected.

Such a hypothesis can be rejected in the case of males over these two time periods. The male model for 1954-1966 is $\nabla\nabla_{12}Z_t = (1 - .33B - .87B^{12} + .35B^{13})a_t$ while the 1967-1976 period model is $(1 - .69B)\nabla_{12}Z_t = (1 - .67B^{12} - .29B^{13})a_t - .352$. These models are not close to being mathematically equivalent, so no attempt was made to fit the earlier period model to the later period data. For both males and females, however, the 1954-1966 period models were used to forecast the one-month-ahead LFPRs of the 1967-1976 period. The nature of the forecast errors which result from this procedure substantiate the idea that male behavior has "changed" in a way that female behavior has not.

First consider the mean squared error (MSE) of the various fits and forecasts involved (Table 1). For both men and women, there is an MSE associated with the fit to the 1954-1966 period, an MSE associated with the fit to the 1967-1976 period, and an MSE associated with the forecast of the 1967-1976 period made with the 1954-1966 model. In neither case was it possible to forecast over the 1967-1976 period from the earlier model better (in an MSE sense) than by using the fitted values from the 1967-1976 period model. For women, however, the 1954-1967 model yielded a smaller MSE over the latter period than over the period of its own estimation. This finding is due in part to the fact that the series structure is almost identical over the two periods and in part to the smaller series variance in the second period relative to the first. For men, the MSE associated with the 1954-1967 model forecasted into the 1967-1976 period is larger than either the MSE for the 1954-1966 fit or for the 1967-1976 fit.

Table 1

Mean Squared Errors for LFPRs

Period of Model	Period of Calculation	Males	Females
1954-1966	1954-1966 (fit)	.037	.093
1954-1966	1967-1976 (forecast)	.042	.083
1967-1976	1967-1976 (fit)	.022	.056

Another testing procedure that can provide insight on changes in series behavior uses the cumulative sum, or cusum, of the normalized forecast errors and the squared normalized forecast errors. Initially, the forecast errors of a series such as female LFPRs are calculated. For present purposes, consider the forecast errors obtained by forecasting the 1967-1976 white female LFPRs with the 1954-1966 model. Beginning with the first forecasted period (January 1967) plot the cusum of the normalized forecast errors. Where this plot remains flat, the prediction error has a mean of approximately zero; where the plot rises the predicted values are too low; where the plot falls the predicted values are too high.

The cusum of the detrended normalized squared forecast errors corresponds to the above example, except this cusum informs as to the forecast error variance relative to the residual variance (error variance of that model fitted to the earlier period).[4] Where this cusum rises, the forecast error variance associated with female LFPRs as predicted by the 1954-1966 model in the 1967-1976 period is greater than the error variance obtained from that model fit to the 1954-1966 period.

Figures 1-A and 1-B display the cusums of the normalized forecast errors for females and males, while Figures 2-A and 2-B display the cusums of the detrended normalized squared forecast errors. As can be seen from Figure 1-A, the cusum for females fluctuates in roughly a $\pm 2\sigma$ band, remaining relatively flat and indicating that the prediction error is approximately 0. The sum of squares in Figure 2-A primarily trends downward, indicating that the forecast error variance shrinks relative to the residual variance over the fitted period. This is an excellent forecasting performance.

For men, again the findings are different. Figure 1-B shows a steadily rising cusum--a sign of chronic underforecasting with the 1954-1966 model. The sum of squares, as shown in Figure 2-B, provides a clue to the underforecasting. This cusum is characterized by discrete, one-period jumps, followed by stable or downtrending periods. This pattern indicates that once the model gets on track, the forecast error variance shrinks admirably, but that after a short spell, the model jumps track for one month.

Closer examination reveals that the 1954-1966 model generally jumps track in January, performs well during the rest of the year, then misses the next January quite badly. The cusum of the normalized forecast errors for the 1967-1976 period for white adult males is 16.588, of which 16.575 can be attributed to January underpredictions. The 1970 calendar year marks the beginning of these large underpredictions. Examination of the data shows that in the 1954-1969 period, LFPRs for this group typically fell from December to January by 0.29 of a percentage point. In the 1970-1976 period, LFPRs for this group fell from December to January by 0.18 of a percentage point, on average. The difference, one-tenth of one percent, represents roughly 60,000 persons.

<u>Unemployment Rates</u>. Diagnostic tests again were applied to the forecasts of the 1967-1976 period made with the 1954-1966 models to determine if any changes had occurred in the series behavior. The two models of the female unemployment rates were very similar, as they were for participation rates. But in this instance the 1954-1966 model structure did not pass significance tests when estimated over the 1967-1976 period. The male unemployment rate models were not equivalent either.

For both males and females the MSEs of the 1967-1976 forecasts made with the 1954-1966 models were lower than the MSEs of the models estimated over the 1954-1966 period because of the lower variances of the latter period relative to the former and to the closeness--if not equivalence--of the second period models to their first period counterparts. Table 2 displays the MSEs of the fits and forecasts.

Table 2

Mean Squared Errors for URs

Period of Model	Period of Calculation	Males	Females
1954-1966	1954-1966 (fit)	.077	.101
1954-1966	1967-1976 (forecast)	.052	.095
1967-1976	1967-1976 (fit)	.034	.074

The cusum of the normalized forecast errors shows that unemployment rates, for females, were underforecasted during the economic slowdowns of 1970 and 1974. Other than these periods, the forecast errors had a zero mean. For men the 1970 and 1974 unemployment rates were also underpredicted, but in addition, the rates were overpredicted as the respective recoveries got underway. These cusums are shown as Figures 3-A and 3-B.

The cusums of the detrended normalized squared forecast errors are quite similar for male and female unemployment rates and reveal a great deal about the error variances of the two time periods. Figures 4-A and 4-B represent these cusums for females and males. The patterns are of almost monotonic decay until the 1974-1975 recession, when the cusum jumps sharply. Relative to the residual variance of the 1954-1966 model period, the forecast error variance associated with the 1967-1976 period continues to decline until the recession. Once the cusum has acknowledged this one-time shift in forecast error variance, it continues to decline as before.

SUMMARY

A causality approach to labor force participation suggests that among adult whites unemployment rates and cross-LFPRs are not useful information for predicting participation rates. Among teenagers, however, unemployment rates may be useful.

Economic policymakers are increasingly concerned about the determinants of labor force participation and possible changes in the structural relationships upon which the participation rates are based. The notion that the time series model for adult white female LFPRs has changed cannot be supported, while for adult white males that notion can be supported.

It is well known that labor market circumstances of individuals are strongly influenced by age. Therefore, a finer disaggregation of the data by age groups possibly would eliminate many of the findings of independence between various pairs of time series. In addition to the test performed in this study, one could test for changes in the individual parameters of the ARIMA models using the procedure discussed in Bagshaw and Johnson [1977].

Nonetheless, many theoretical issues remain to be considered. Further study could be undertaken to determine the amount of economic information that is removed from the data via the filtering procedure. To approach this issue one can consider the degree of residual variance that would prompt the consideration that the information contained in the white noise process is, in some sense, not worth examining. Another view would focus on the filter to determine whether the degree of differencing to achieve stationarity removes important information from the analysis.

FOOTNOTES

[1] Because the data are rates for groups as opposed to decisions of individuals, it would be inappropriate to discuss the results in terms of individuals.

[2] W. G. Bowen and T. A. Finegan, The Economics of Labor Force Participation (Princeton: Princeton University Press, 1969), p. 515.

[3] Detailed teenage data were unavailable before 1967. For the 1967-1976 period, no tests were done between teenage unemployment rates and adult participation rates. The authors assumed that adult participation decisions were not conditional on teenage market tightness.

[4] The cumulative sume of the square of the forecast errors is standardized by the residual variance from the earlier time period. One is substracted for detrending: $\sum_{t=1}^{n}\left(\frac{a_t^2}{\hat{\sigma}^2} - 1\right)$

APPENDIX ONE

The Haugh method of testing the independence of two time series is a two-stage process which involves fitting univariate models to each of the series and cross correlating the two resulting series. These cross correlations are then tested for significance to determine whether the two series are independent.

In the first stage, the univariate models used are integrated autoregressive moving average seasonal models. The general model of order (p, d, q) X (P, D, Q) is given by

$$\Phi(B^s)\phi(B)\nabla^d\nabla_s^D X_t = \Theta(B^s)\theta(B)a_t + \theta_o$$

where

$$\Phi(B^s) = 1 - \Phi_1 B^s - \ldots - \Phi_P B^{Ps}$$

$$\phi(B) = 1 - \phi_1 B - \ldots - \phi_p B^p$$

$$\Theta(B^s) = 1 - \Theta_1 B^s - \ldots - \Theta_Q B^{Qs}$$

$$\theta(B) = 1 - \theta_1 B - \ldots - \theta_q B^q$$

$$\nabla^d = (1 - B)^d$$

$$\nabla_s^D = (1-B^s)^D$$

θ_o is a constant, B is the backshift operator (e.g., $B^j X_t = X_{t-j}$), s is the seasonal period (e.g., s = 12 for monthly data), and a_t is a white noise series.

Thus, these models incorporate all the information contained in past values of the series with the resulting residual series, a, being a series of uncorrelated values. This procedure is in itself a three stage process involving model identification, model fitting and diagnostic checking. For a detailed discussion of this procedure, see Box and Jenkins [1970].

The second stage consists of calculating the cross correlations between the residuals from the two series. These estimates are given by

$$r_{a_1,a_2}(k) = \frac{C_{a_1,a_2}(k)}{\sqrt{C_{a_1}(o)C_{a_2}(o)}}$$

where

$$C_{a_1,a_2}(k) = \begin{cases} \frac{1}{N}\sum_{t=1}^{N-k} a_{1,t}a_{2,t+k} & k \geq 0 \\ \frac{1}{N}\sum_{t=1-k}^{N} a_{1,t}a_{2,t+k} & k < 0 \end{cases}$$

$$C_{a_1}(o) = \frac{1}{N}\sum_{t=1} a_{i,t}^2$$

and $a_{i,t}$ is the residual from series i at time t.

It can be shown that if the two original series are uncorrelated, then $r_{a_1,a_2}(k)$ will have an asymptotic variance of $(N - |k|)^{-1}$ and the covariances of the cross correlations at different lags will be on the order of N^{-1}. Thus, this method of transformation makes it much easier to interpret the cross correlations than the Box-Jenkins technique, which applies the same filter to each data series.

Using the definition of causality which says X causes Y if we can better predict Y using X than we can when using only past values of Y, these cross correlations can be used to test for causality as explained in the text of this paper.

APPENDIX TWO*

A. ARIMA Models

Series Group			Time	Structure
LFPR	Males	(A)	1954-1966	$\nabla\nabla_{12}z_t = (1 - .34B - .87B^{12} + .35B^{13})a_t$
LFPR	Females	(A)	1954-1966	$\nabla\nabla_{12}z_t = (1 - .16B + .15B^{11} - .92B^{12})a_t$
LFPR	Males	(A)	1967-1976	$(1 - .69B)\nabla_{12}z_t = (1 - .67B^{12} - .29B^{13})a_t - .35$
LFPR	Females	(A)	1967-1976	$\nabla\nabla_{12}z_t = (1 - .90B^{12})a_t$
UR	Males	(A)	1954-1966	$\nabla\nabla_{12}z_t = (1 + .18B^2 - .91B^{12} - .19B^{14})a_t$
UR	Females	(A)	1954-1966	$\nabla\nabla_{12}z_t = (1 - .23B - .93B^{12} + .20B^{13})a_t$
UR	Males	(A)	1967-1976	$(1 - .37B - .19B^2)\nabla\nabla_{12}z_t = (1 - .92B^{12})a_t$
UR	Females	(A)	1967-1976	$\nabla\nabla_{12}z_t = (1 - .95B^{12})a_t$
LFPR	Males	(T)	1967-1976	$\nabla\nabla_{12}z_t = (1 - .62B - .35B^5)a_t$
LFPR	Females	(T)	1967-1976	$(1 - .66B)\nabla\nabla_{12}z_t = (1 + .17B^4 - .50B^{12} - .36B^{24})a_t + 1.23$
UR	Males	(T)	1967-1976	$\nabla\nabla_{12}z_t = (1 - .14B - .81B^{12})a_t$
UR	Females	(T)	1967-1976	$\nabla\nabla_{12}z_t = (1 - .53B - .46B^{12})a_t$

B. Independence Test Results

Time Period	Series	Cross Correlated With	Finding
1954-1966	LFPR Males (A)	UR Males (A)	UR causes LFPR
1954-1966	LFPR Females (A)	UR Females (A)	Independent
1954-1966	LFPR Females (A)	UR Males (A)	Independent
1954-1966	LFPR Males (A)	UR Females (A)	Independent
1954-1966	LFPR Males (A)	LFPR Females (A)	Contemporaneous
1967-1976	LFPR Males (A)	UR Males (A)	Independent
1967-1976	LFPR Females (A)	UR Females (A)	Independent
1967-1976	LFPR Females (A)	UR Males (A)	Independent
1967-1976	LFPR Males (A)	UR Females (A)	Independent
1967-1976	LFPR Males (A)	LFPR Females (A)	Independent
1967-1976	LFPR Males (A)	LFPR Males (T)	Contemporaneous
1967-1976	LFPR Males (A)	LFPR Females (T)	Independent
1967-1976	LFPR Females (A)	LFPR Females (T)	Independent
1967-1976	LFPR Females (A)	LFPR Males (T)	Independent
1967-1976	UR Males (A)	LFPR Males (T)	Feedback
1967-1976	UR Males (A)	LFPR Females (T)	LFPR causes UR
1967-1976	UR Females (A)	LFPR Females (T)	Feedback
1967-1976	UR Females (A)	LFPR Males (T)	Independent
1967-1976	LFPR Males (T)	LFPR Females (T)	Independent
1967-1976	LFPR Males (T)	UR Females (T)	UR causes LFPR
1967-1976	LFPR Males (T)	UR Males (T)	UR causes LFPR
1967-1976	LFPR Females (T)	UR Males (T)	UR causes LFPR
1967-1976	LFPR Females (T)	UR Females (T)	Independent

*(A) refers to adults; (T) refers to teenagers.

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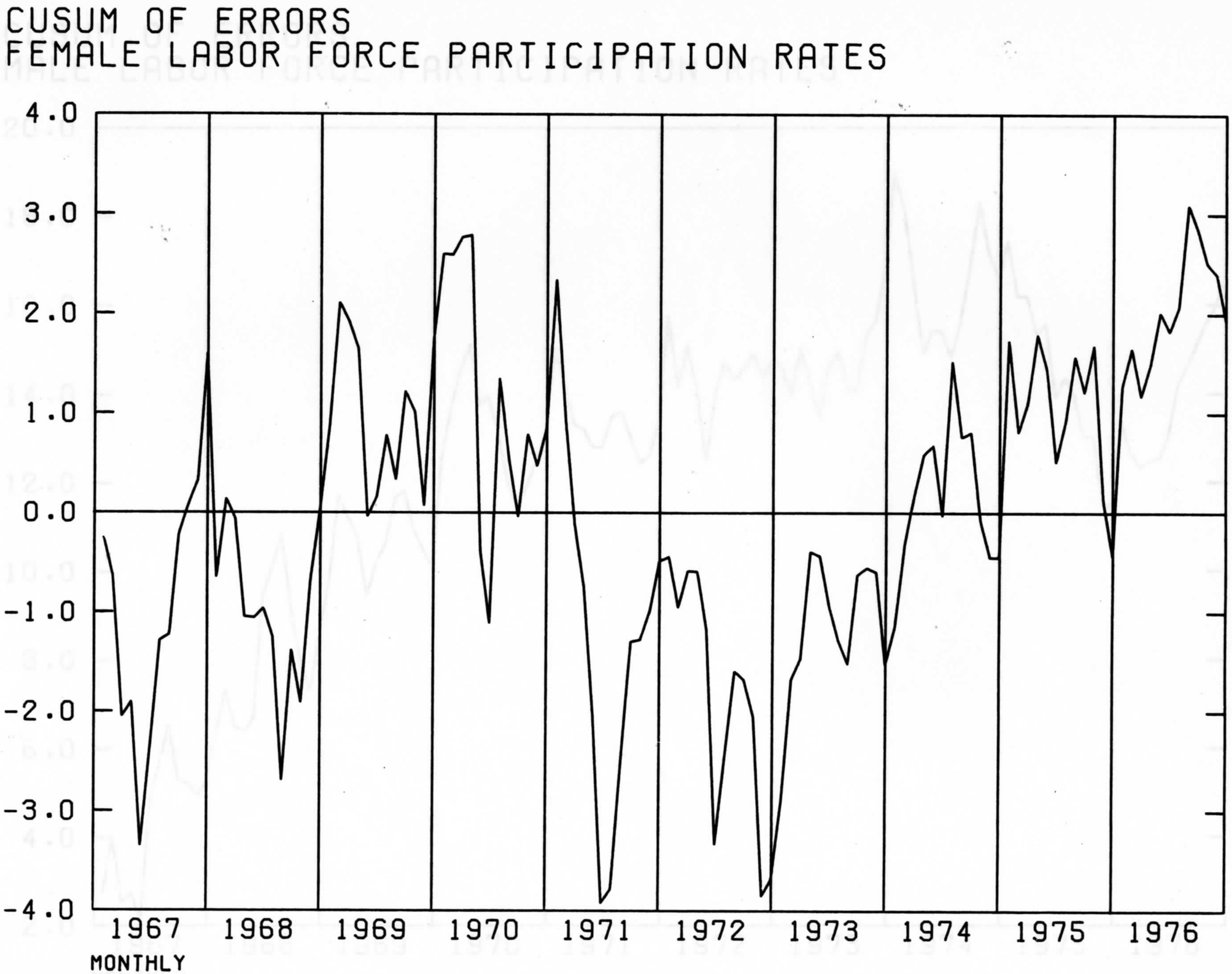
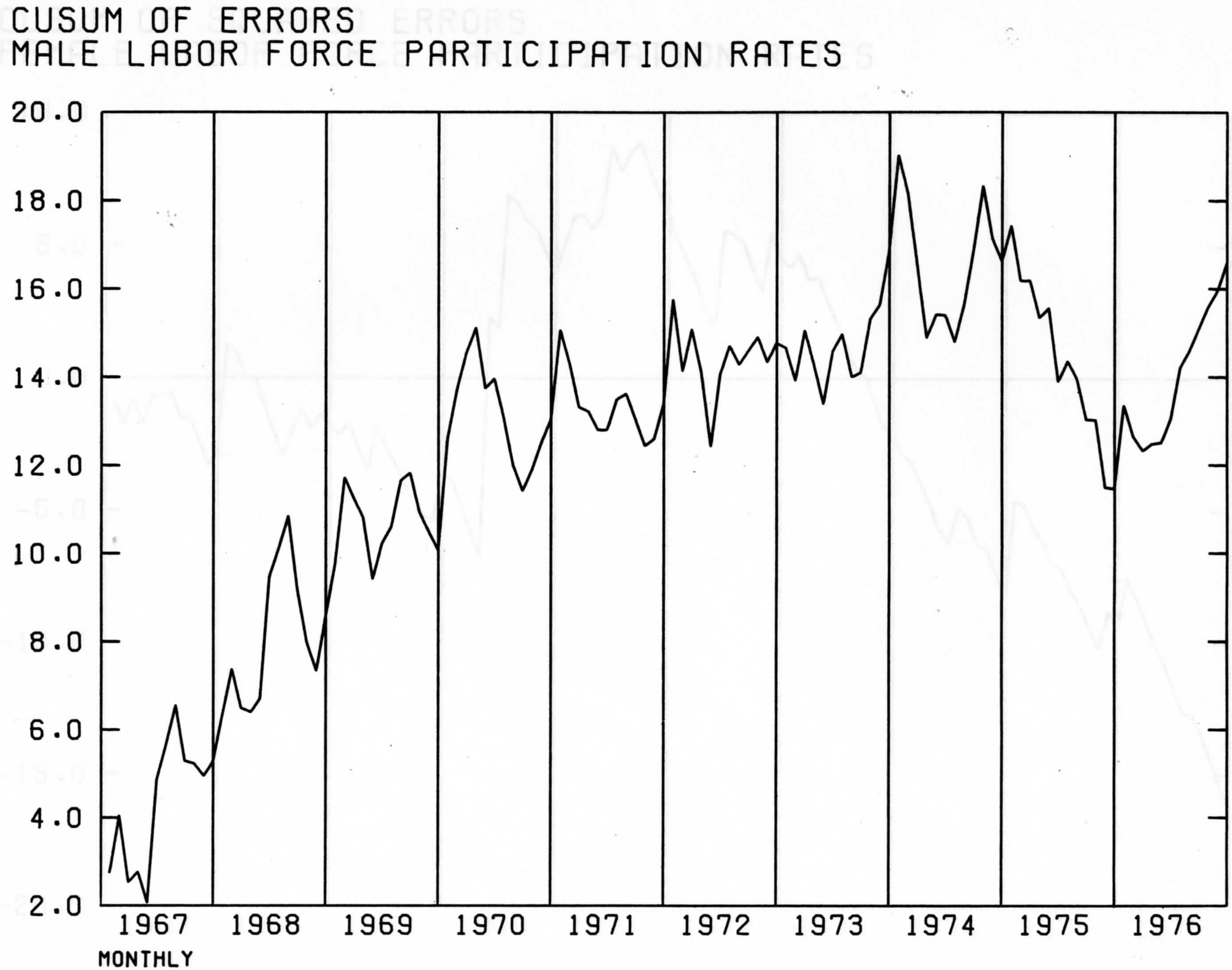


CUSUM OF ERRORS
FEMALE LABOR FORCE PARTICIPATION RATES
4.0
3.0
2.0
1.0
0.0
-1.0
-2.0
-3.0
-4.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976
MONTHLY

FIGURE 1—A


CUSUM OF ERRORS
MALE LABOR FORCE PARTICIPATION RATES
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18.0
16.0
14.0
12.0
10.0
8.0
6.0
4.0
2.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976
MONTHLY

FIGURE 1—B

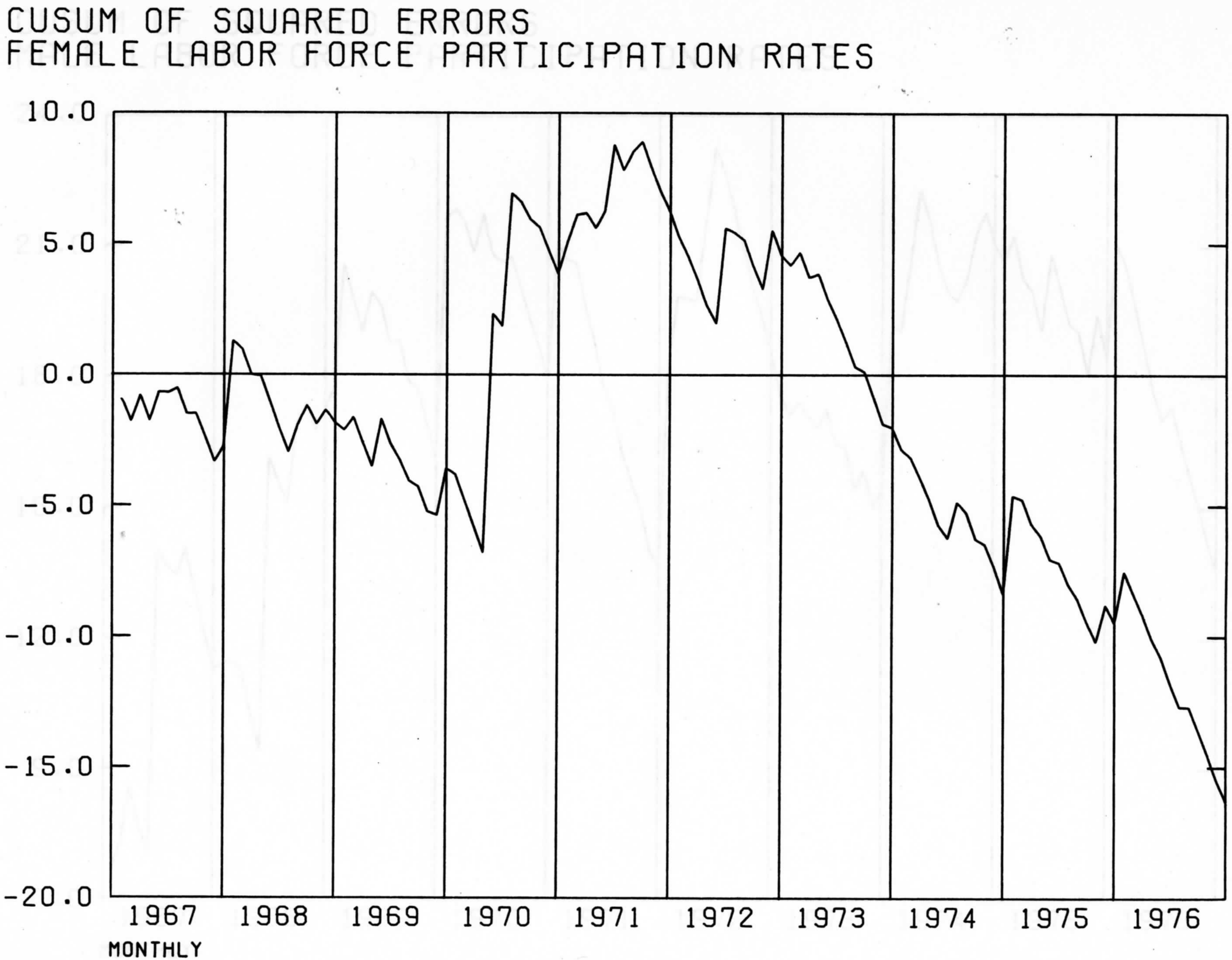
CUSUM OF SQUARED ERRORS
FEMALE LABOR FORCE PARTICIPATION RATES
10.0
5.0
0.0
-5.0
-10.0
-15.0
-20.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976
MONTHLY

FIGURE 2—A

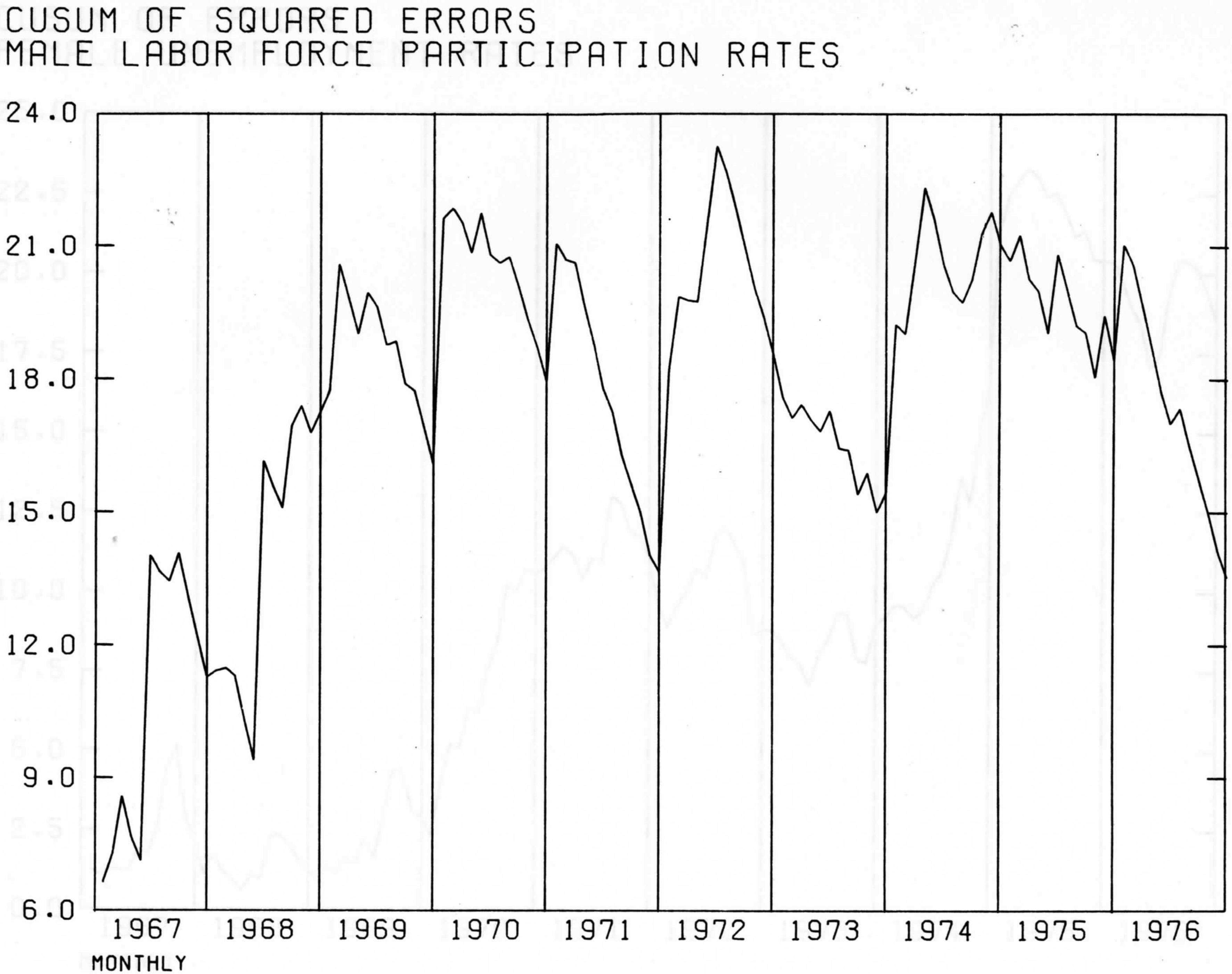
CUSUM OF SQUARED ERRORS
MALE LABOR FORCE PARTICIPATION RATES
24.0
21.0
18.0
15.0
12.0
9.0
6.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976
MONTHLY

FIGURE 2–B

CUSUM OF ERRORS
FEMALE UNEMPLOYMENT RATES

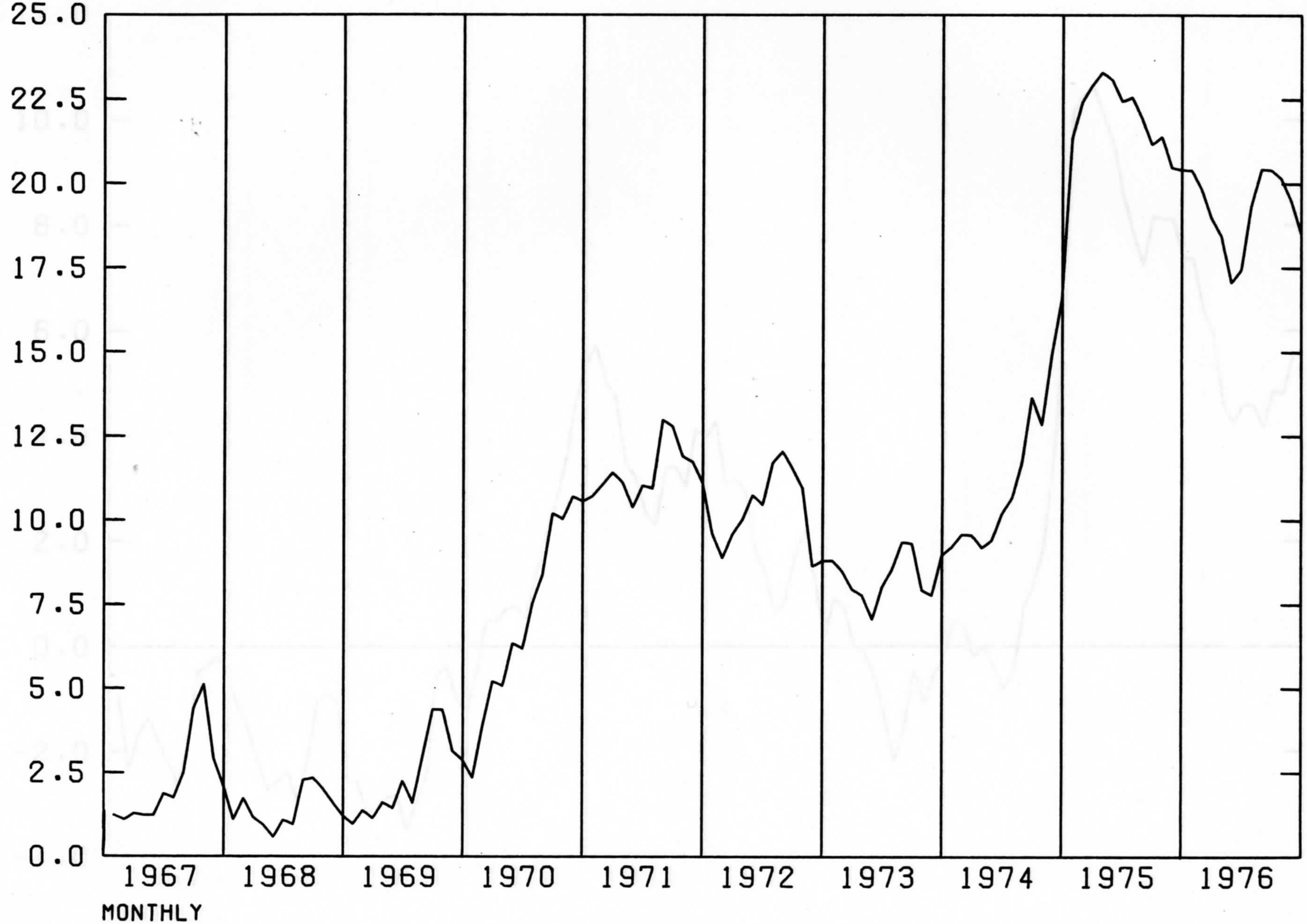

FIGURE 3—A


CUSUM OF ERRORS
MALE UNEMPLOYMENT RATES
12.0
10.0
8.0
6.0
4.0
2.0
0.0
-2.0
-4.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976
MONTHLY

FIGURE 3—B

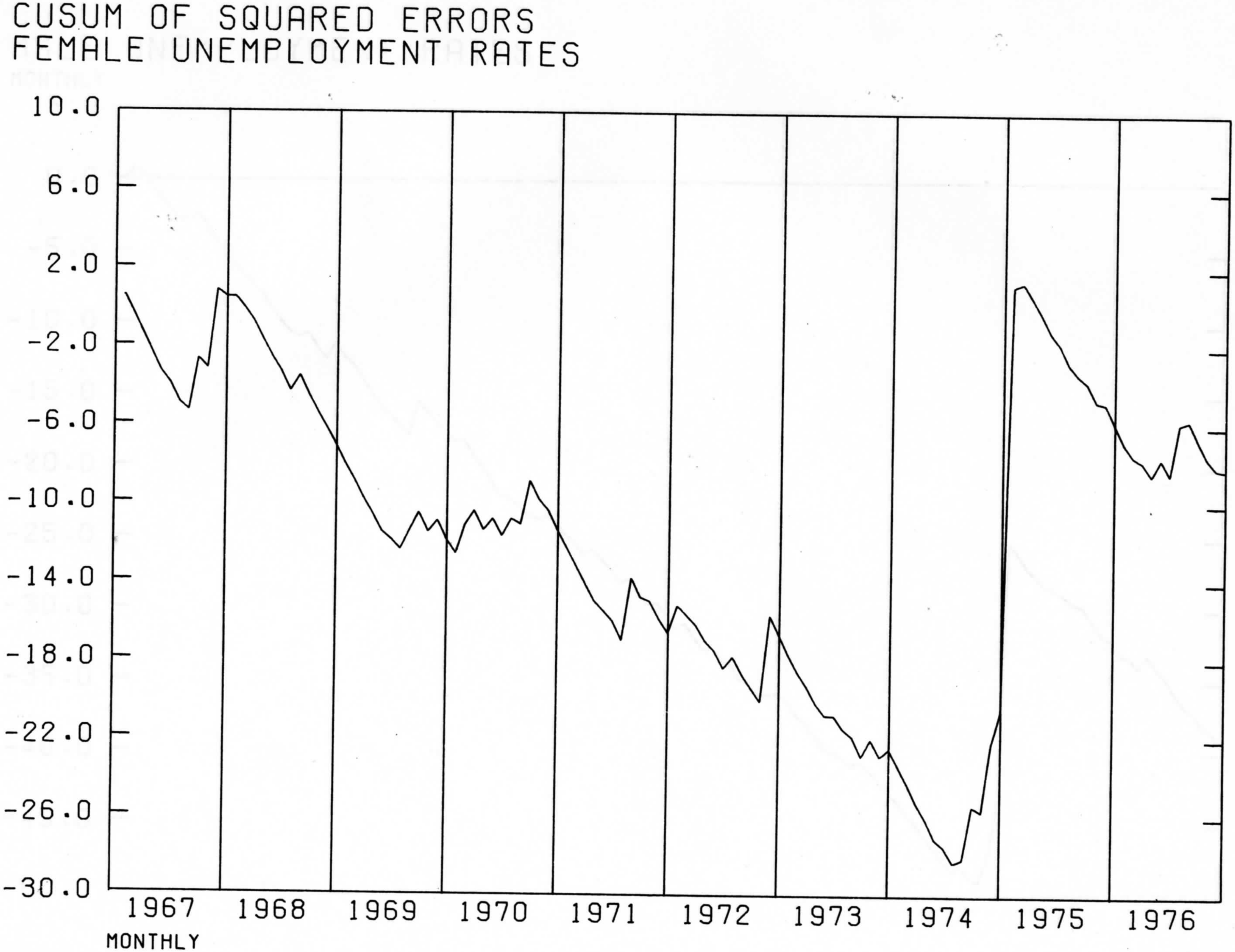
CUSUM OF SQUARED ERRORS
FEMALE UNEMPLOYMENT RATES
10.0
6.0
2.0
-2.0
-6.0
-10.0
-14.0
-18.0
-22.0
-26.0
-30.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976
MONTHLY

FIGURE 4—A

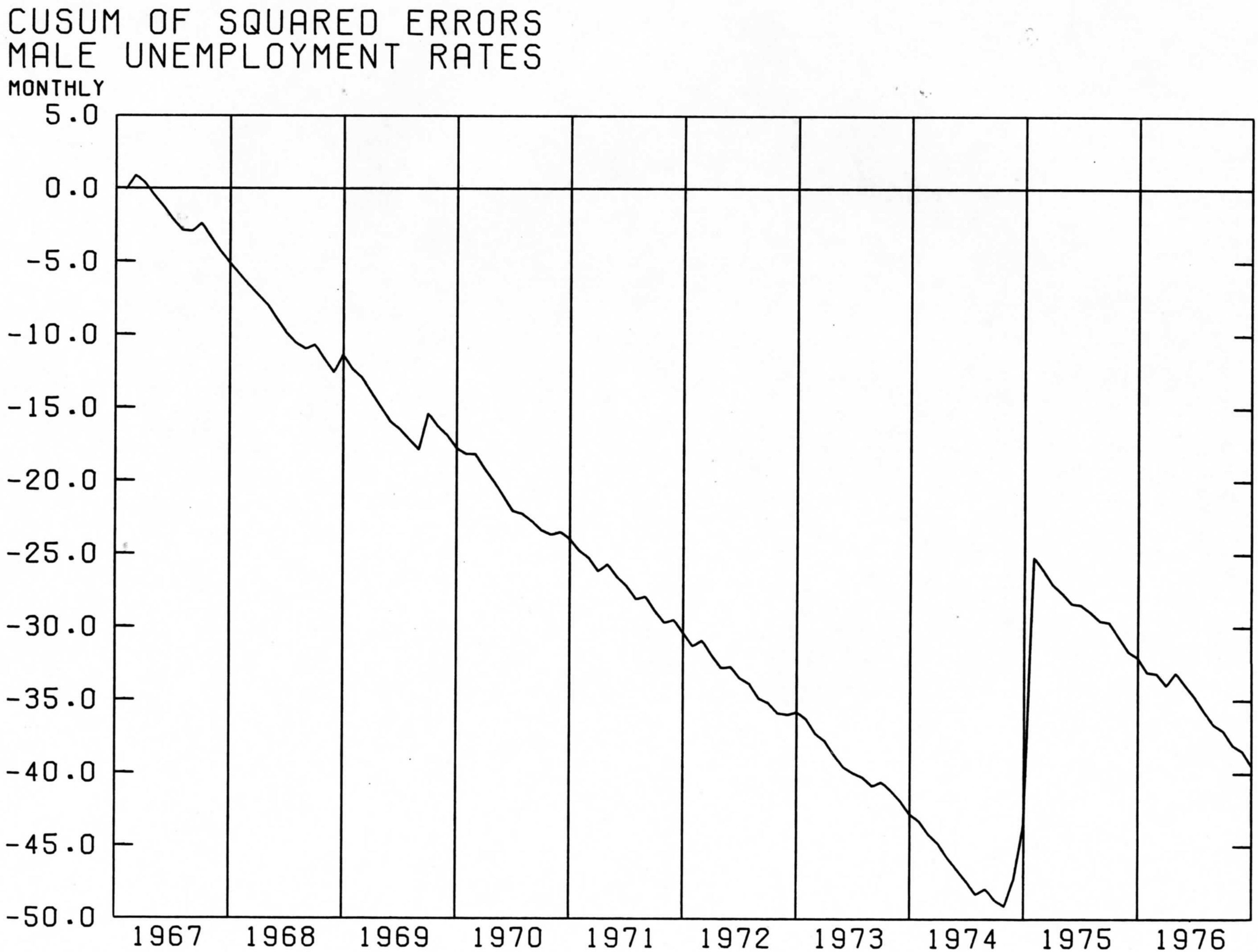
CUSUM OF SQUARED ERRORS
MALE UNEMPLOYMENT RATES
MONTHLY
5.0
0.0
-5.0
-10.0
-15.0
-20.0
-25.0
-30.0
-35.0
-40.0
-45.0
-50.0
1967
1968
1969
1970
1971
1972
1973
1974
1975
1976

FIGURE 4—B