SEASONALITY IN STOCK RETURNS:

SOME INTERNATIONAL EVIDENCE

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NOT FOR QUOTATION.
COMMENTS WELCOME.

Thanks are due to Grant Clowery, Don Keim, Allan Kleidon, and Paul Schultz. The usual caveat applies.

I. Introduction

A number of recent studies have documented seasonal patterns in common stock returns. For example, Officer (1975) finds pronounced autocorrelations in returns of common stocks listed on the Melbourne stock exchange over the 1958-1970 period. In particular, he finds significant autocorrelations at lags of six, nine, and twelve months. Rozeff and Kinney (1976) find that seasonality in NYSE stocks (from 1904 to 1974) is due primarily to high mean returns in January versus other months.

Keim (1982) provides evidence that this January seasonal is *much* more pronounced for small firms than for large firms, (where size is defined as the market value of common equity). In fact, Keim (1982) shows that much of the "size effect", documented in Banz (1981), is concentrated in the first few trading days in January.

A number of different explanations have been proposed for the "January" and/or "size" effects. Officer suggests that the opportunity cost of capital may exhibit seasonality, presumably due to seasonality in consumption/investment patterns. Other hypotheses include compensation for increased uncertainty in January because of large amounts of information released (e.g., accounting information); seasonality in "liquidity" risk premia (Rozeff and Kinney); transactions costs (Stoll and Whaley (1982)); and tax related portfolio rebalancing (Roll (1982)).

In this paper I present some preliminary evidence on stock market seasonality across countries. In addition to providing useful descriptive evidence on international stock portfolios, cross-country comparisons may be useful in discriminating between alternative explanations of the January effect. This evidence could be particularly useful in examining tax related hypotheses since tax laws vary across countries.

In the next section of the paper the "tax-selling" hypothesis is discussed along with some relevant results found in Roll (1982). In section III the data are described and evidence on seasonality is presented. A summary and conclusions are given in the fourth section of the paper along with some proposals for future work.

II. Tax-Selling Hypothesis

The essential argument behind the tax-selling hypothesis is as follows. Those stocks which have experienced price declines during the year are sold by investors near the end of the year in order to realize capital losses. These sales, the tax-selling hypothesis asserts, create downward pressure on the stock price. When this pressure is relieved at the beginning of the new tax year the price of the stock rebounds. Since small firms tend to have more volatile stock returns, they have a greater probability of achieving a large negative return over any given period and, hence, are more likely vehicles for the realization of capital gains, (see Roll (1982), p. 11). This is presumably the reason why small firms show a more pronounced January effect than large firms.

If the tax-selling hypothesis were true then there should be a negative correlation between the return on a security over the year and its return in the first month of the subsequent tax year. That is, those firms with the largest capital losses experience the greatest selling pressure and, consequently the largest rebound in January. Roll (1982) finds that this negative correlation is found in the NYSE and AMEX stocks on the CRSP tape (see his Table 3, p. 27). He also finds that the correlation is more negative (larger in absolute value) for price changes that occur later in the year. On the basis of his tests Roll concludes that the January seasonal "cannot be explained by data errors, listings, de-listings, or outliers. Instead, it is closely associated with tax loss selling induced by negative returns over the previous YEAR. Transaction costs and low liquidity probably prevent arbitrageurs from eliminating the return seasonality."[1]

If the January effect is induced by tax laws we should see

[1]Roll (1982, p.19). The evidence in Schultz (1982) seems to indicate that transaction costs are not large enough to explain the persistance of the January seasonal. In addition, since taxpayers can choose tax years which do not end on December 31st, they could avoid selling stock at depressed prices in December by choosing a non-standard tax year. There may be costs to such a strategy, however.

quite different patterns in stock market returns of countries with differing tax laws. For example, the United Kingdom and Australia have different tax years than the United States. We should see an April effect in the U.K. (corporate tax year ending March 31 and personal tax year ending April 5) and a July effect in Australia (June 30 tax year).[2] In the Netherlands and Norway we should not see any seasonal because these countries do not have capital gains taxes. There was no capital gains tax in Canada before 1972. Thus, under the tax-selling hypothesis, one should see a seasonal only after the change in the tax law.

In the next section of the paper preliminary evidence on stock market seasonality is presented. At this stage the data consist of monthly stock market indexes for the period from January 1969 through January 1982.

III. Data and Statistical Tests

The data used in this draft are monthly indexes published in Capital International Perspective (CIP). The period for which I have data is January 1969 through January 1982. There are a number of important shortcomings to this data set. First of all the time period is quite short (just over thirteen years). Thus, parameter estimates will be much less precise when compared to other studies (e.g., Rozeff and Kinney (1976) have 71 years of

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[2] The major source of information on tax laws is Taxation in Western Europe published by the Confederation of British Industry.

monthly data while Keim (1982) has seventeen years of daily data and 53 years of monthly data on individual stocks). In addition, the stock indexes in CIP are value weighted. The results of Keim imply that large firms have much less pronounced seasonals than small firms. Since a value weighted index, by definition, places more weight on large firms it will exhibit much less seasonality than an equally weighted index. Therefore, the CIP data is biased against finding a seasonal. Finally, it would be preferable to have data on individual securities rather than aggregated stock indexes. Some of the implications of different tax laws (under the hypothesis that the tax-selling argument holds) can only be tested on data for individual securities.[3]

In the near future the Center for Research in Security Prices (CRSP) will be supporting stock return tapes for Canada, Japan, and the United Kingdom. Once these data become available more extensive testing of the tax-selling hypothesis will be conducted. Until then the indexes will be used.

In order to test for a seasonal shift in mean returns, the percentage change in the stock index of each country was regressed on a constant and eleven dummy variables, one for each

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[3]For example, in West Germany, if assets are held longer than six months there is no capital gains tax. If Roll's finding that the negative correlation between the return in January and the returns in preceeding months (through February) is due to tax-selling then this negative correlation should be zero at lags greater than six months. One needs disaggregated data to test this since cross-sectional regressions regressions are used to estimate the correlation.

month from February through December. That is:

$$\tilde{R}_{it} = a_{i1} + \sum_{j=2}^{12} a_{ij}D_{jt} + \tilde{e}_{it} \tag{3.1}$$

where: i=country index

$$D_{jt} = \begin{cases} 1 \text{ if period } t \text{ is month } j \\ 0 \text{ otherwise} \end{cases}$$

$\tilde{e}_t$ =a random error term.

The general pattern of OLS coefficient estimates was a positive intercept, a_{i1}, and negative slope coefficients, a_{ij}, j=2,...,12. For each country the hypothesis $a_{i2}=a_{i3}=\ldots=a_{i12}$ was tested. The hypothesis was accepted (at the 5% level of significance) for fifteen out of sixteen countries.[4] Given this result the degrees of freedom can be increased by imposing the above constraint. Table 1 contains the results for the constrained regression:

$$\tilde{R}_{it} = a_i + b_iD_t + \tilde{u}_t \tag{3.2}$$

where: $D_t = \begin{cases} 0 \text{ if month is January} \\ 1 \text{ otherwise} \end{cases}$

The general pattern of the coefficient estimates is consistent with the January seasonal documented by others on U.S. data. However, the parameter estimates are not very precise. In addi-

[4]The test is rejected for Belgium.

tion the seasonal explains very little of the variance of stock returns, as can be seen by the small values of R^2. The t-statictic tests the hypothesis that the mean return in January is different than the mean return of the remaining eleven months. Six out of the sixteen values of $\hat{b}$ are significant at the 10% level. Five of these six display the pattern found in U.S. data while one country (Spain) shows the reverse. Although many of the values of $\hat{b}$ are statistically insignificant, economically they imply non-trivial differences between January returns and the returns of other months. The average, across countries, of $\hat{b}$ is -.0145 while the average t-statistic is -1.20. A 1.45% return differential <u>per</u> <u>month</u> represents a 18.9% annualized return. Thus there may be an economically significant seasonal effect even though there is not sufficient data to estimate the seasonality precisely.

While the lack of very precise estimates of b should lead one to be cautious about interpreting the results in Table 1, it seems reasonable to conclude that stock returns in many of the countries investigated here exhibit a January seasonal. One might attempt to obtain a more precise estimate of the seasonal effect by constraining $b_i = b_j$, for all i,j. However, an F-test rejected this restriction at the 1% significance level. Therefore, the difference between the mean returns in January and the other months is not constant across countries.

The regressions reported in Table 1 were also run on data over the period from January 1972 to January 1982. Over this subperiod the data show a much more pronounced January seasonal. Nine of sixteen estimates of b were significant at the 10% level. Of these nine, seven were significant at the 5% level. This increase in significance from dropping three years of data seems to indicate that some of the results may be due to outliers.

Figures 1 through 16 are plots of the residuals from the regressions reported in Table 1. The horizontal lines represent two sample standard deviations above and below zero. The plots give some indication that then variance of returns is larger in January than in other months. Also, the large positive return, for many countries, in January 1970 may be causing the seasonality. However, the January seasonal was found to be <u>more</u> significant when the period 1-1969 through 12-71 is excluded. In addition to the large January return in 1970 there seem to be many large (in absolute value) returns during the month of January. Bartlett's test was used to test the hypothesis that the variances are equal across months. Under the null hypothesis the statistic has a χ^2 distribution with eleven degrees of freedom.[5] The results are reported in Table 2. For only three countries (Canada, Italy, and U.S.) the null hypothesis of homoscedasticity is accepted.

Given this heteroscedasticity the standard errors reported in

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[5]See Brownlee (1965, p. 293).

Table 1 will be incorrect. In order to adjust for the heteroscedasticity weighted least squares (WLS) was used to estimate the model:

$$W_{it}\tilde{R}_{it} = a_i W_{it} + b_i W_{it} D_t + \tilde{\eta}_{it} \tag{3.3}$$

where:

$$W_{it} = \begin{cases} S_{i1}^{-1} & t = \text{January} \\ S_{i0} & t \neq \text{January} \end{cases}$$

S_{i1} = the sample standard deviation of the OLS residuals (from 3.2) for January data only

S_{i0} = the sample standard deviations of the OLS residuals from all other months

The parameter estimates from (3.3) are identical to those from (3.2) in Table 1. However, the reported standard errors increase when WLS is used. The average t-statistic for the slope coefficient is -.74 for WLS as opposed to -1.20 for OLS. Now only three values of $\hat{b}$ are significant at the 10% level (Italy, Belgium, and Denmark). As before, although the overall pattern of the parameter estimates is consistent with the January seasonal the parameters are measured with little precision. Differences in mean returns of 1.45% per month (18.9% per annum) are, in general, not statistically significant.

Table 3 contains estimates of (3.1) for Australia and the

United Kingdom. The regressions are estimated for two time periods (a) 1-1969 to 1-1982 and (b) 1-1972 to 1-1982. Over the full time period the results are statistically insignificant and give conflicting results regarding the tax-selling hypothesis. The U.K. regression shows a positive intercept and a positive slope coefficient for the April dummy variable. The remaining coefficients are negative. One might possibly take this as evidence for the existence of a "January" and a "Turn-of-the-Tax Year" effect. The Australian data yield negative estimates of the intercept and the July slope coefficient (with four of the ten remaining parameter estimates being negative).

Over the 1972-1982 period the results for both the U.K. and Australia are consistent with a January seasonal. In both cases the intercept is positive while all coefficients for the dummy variables are negative.

IV. Summary and Conclusions

Given the short time periods studied here and the fact that the stock indexes are value weighted (hence, biased against finding a January seasonal) it is not possible to produce strong evidence regarding the January effect. Dummy variable regressions give evidence generally supporting the existence of a January effect in the countries studied here.

Only two of these countries have tax years that do not end on

December 31st. One of these shows some weak (insignificant) evidence for a tax year effect while the other does not support this hypothesis. In addition, two other countries (Netherlands and Norway) which do not have capital gains taxes have January seasonals of approximately equal magnitude (2.3% per month). Only one of these is significant at the 10% level (Netherlands). This is some indication that the tax-selling hypothesis does not explain the January effect.

Clearly further work needs to be done. Future work will include analysis of stock return data at the individual security level for at least a subset of the countries studied here. The stock index data will be expanded. Finally a more detailed description of tax laws and their evolution is needed.

Table 1

Monthly Data 1-1969 to 1-1982

$$\tilde{R}_{it} = a_i + b_i D_t + \tilde{U}_{it}$$

Country	$\hat{a}$	$\hat{b}$	R^2	DW
United States	.007 (.58)	-.008 (-.61)	.00	1.93
Japan	-.004 (-.24)	.012 (.66)	.00	2.05
United Kingdom	.031 (1.52)	-.032 (-1.50)	.01	1.78
Canada	.021 (1.44)	-.018 (-1.21)	.01	1.94
Germany	.010 (.72)	-.013 (-.89)	.01	1.87
France	.023 (1.20)	-.026 (-1.27)	.01	2.15
Australia	-.003 (-.15)	.002 (.08)	.00	2.03
Spain	-.033* (-1.84)	.032* (1.70)	.02	2.01
Switzerland	.020 (1.39)	-.025* (-1.67)	.02	1.96
Netherlands	.020 (1.46)	-.024* (-1.70)	.02	1.84
Italy	.035* (1.86)	-.036* (-1.87)	.02	1.76
Belgium	.039** (3.77)	-.044** (-4.00)	.09	1.87
Sweden	.010 (.66)	-.009 (-.56)	.00	1.89
Denmark	.026** (2.32)	-.024** (- 2.01)	.03	1.44**
Norway	.023 (1.05)	-.022 (-.95)	.01	1.72
Austria	-.003 (-.33)	.003 (.40)	.00	2.38

* Significant at 10% level.
**Significant at 5% level.
DW = Durbin-Watson Statistic.
t- STATISTICS IN PARENTHESES

Table 2

Tests for Equality of Variance Across Months

Country	Test Statistic	Country	Test Statistic
United States	5.99	Switzerland	32.42***
Japan	55.27***	Netherlands	31.59***
United Kingdom	24.40**	Italy	15.62
Canada	14.83	Belgium	27.14***
Germany	48.84***	Sweden	42.94***
France	23.28**	Denmark	19.93**
Australia	29.56***	Norway	18.51*
Spain	68.53***	Austria	52.36***

$X^2(11)$ - Fractiles 90% = 17.3

95% = 19.7

99% = 24.7

*Significant at 10%
**Significant at 5%
***Significant at 1%

Table 3

$$\tilde{R}_{it} = a_{i1} + \sum_{j=2}^{12} a_{ij}D_{jt} + \tilde{e}_{it}$$

(a) 1969-1982

Country	a_1	a_2	a_3	a_4	a_5	a_6	a_7	a_8	a_9	a_{10}	a_{11}	a_{12}	R^2	DW
Australia	.003	-.001	.001	-.011	.014	.003	-.021	.000	-.026	.022	-.005	.042	.05	2.00
	(-.15)	(-.05)	(.03)	(-.38)	(.48)	(.09)	(-.72)	(.01)	(-.90)	(.75)	(-.15)	(1.43)		
United Kingdom	.031	-0.16	-.038	.003	-.052*	-.047	-.038	-.021	-.050*	-.038	-.037	-.015	.06	1.74
	(1.50)	(-.55)	(-1.29)	(.10)	(-1.77)	(-1.59)	(-1.28)	(-.70)	(-1.70)	(-1.29)	(-1.26)	(-.50)		

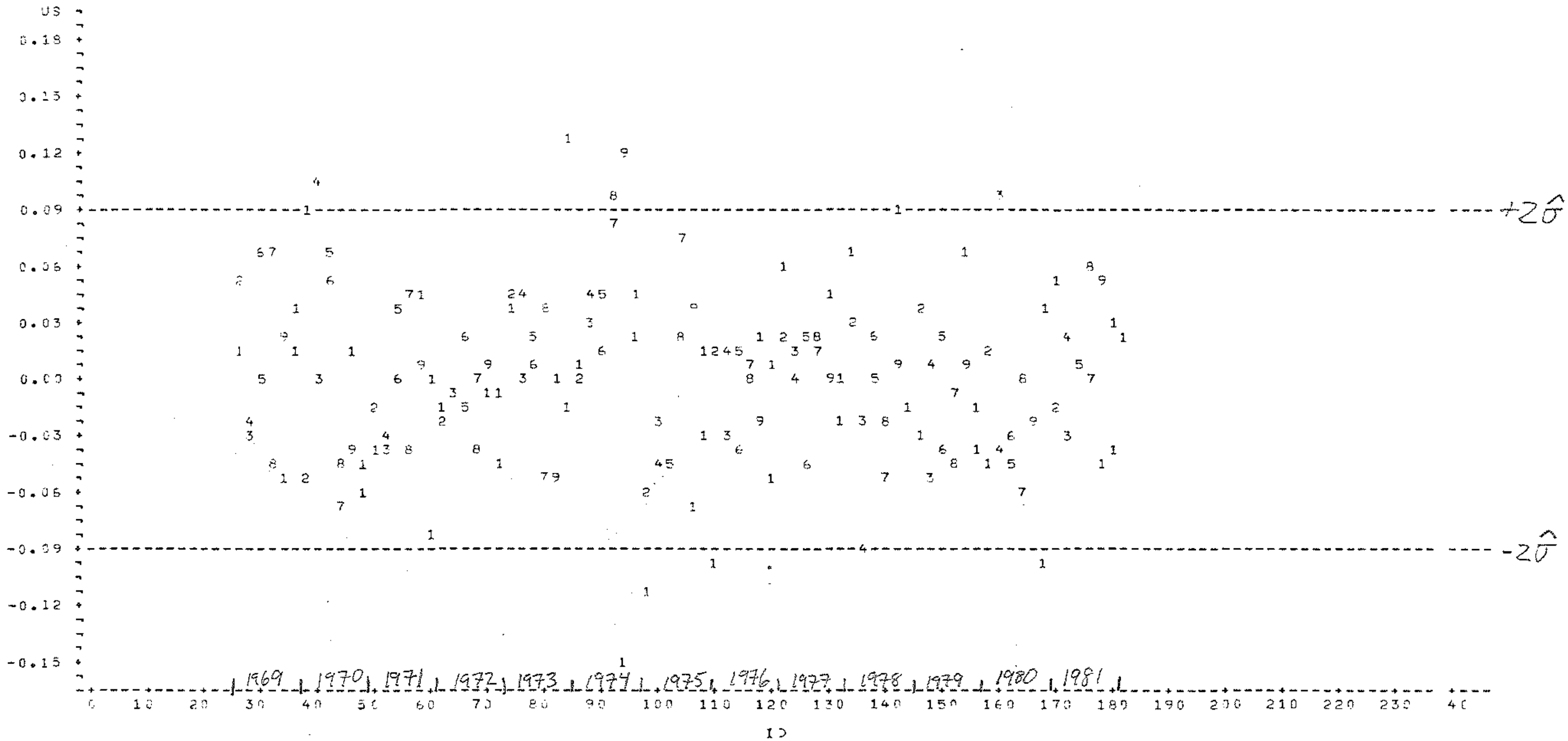
(b) 1972-1982

Country	a_1	a_2	a_3	a_4	a_5	a_6	a_7	a_8	a_9	a_{10}	a_{11}	a_{12}	R^2	DW
Australia	.030	-.028	-.037	-.034	-.000	-.041	-.056*	-.027	-.059*	-.001	-.032	-.012	.08	1.86
	(1.38)	(-.90)	(1.19)	(-1.09)	(-.02)	(-1.31)	(-1.79)	(-.88)	(1.91)	(-.04)	(1.03)	(-.37)		
United Kingdom	.056**	-.018	-.077**	-.006	-.080**	-.077**	-.066*	-.043	-.086**	-.056*	-.065*	-.050	.13	1.68
	(2.37)	(-.53)	(-2.26)	(-.17)	(-2.34)	(-2.26)	(-1.93)	(-1.27)	(-2.54)	(-1.66)	(-1.90)	(-1.48)		

*Significant at 10%
**Significant at 5%
t-statistics in parentheses

FIGURE 1
TIME SERIES PLOT OF ERRORS
US DATA 1-1969 TO 1-1982

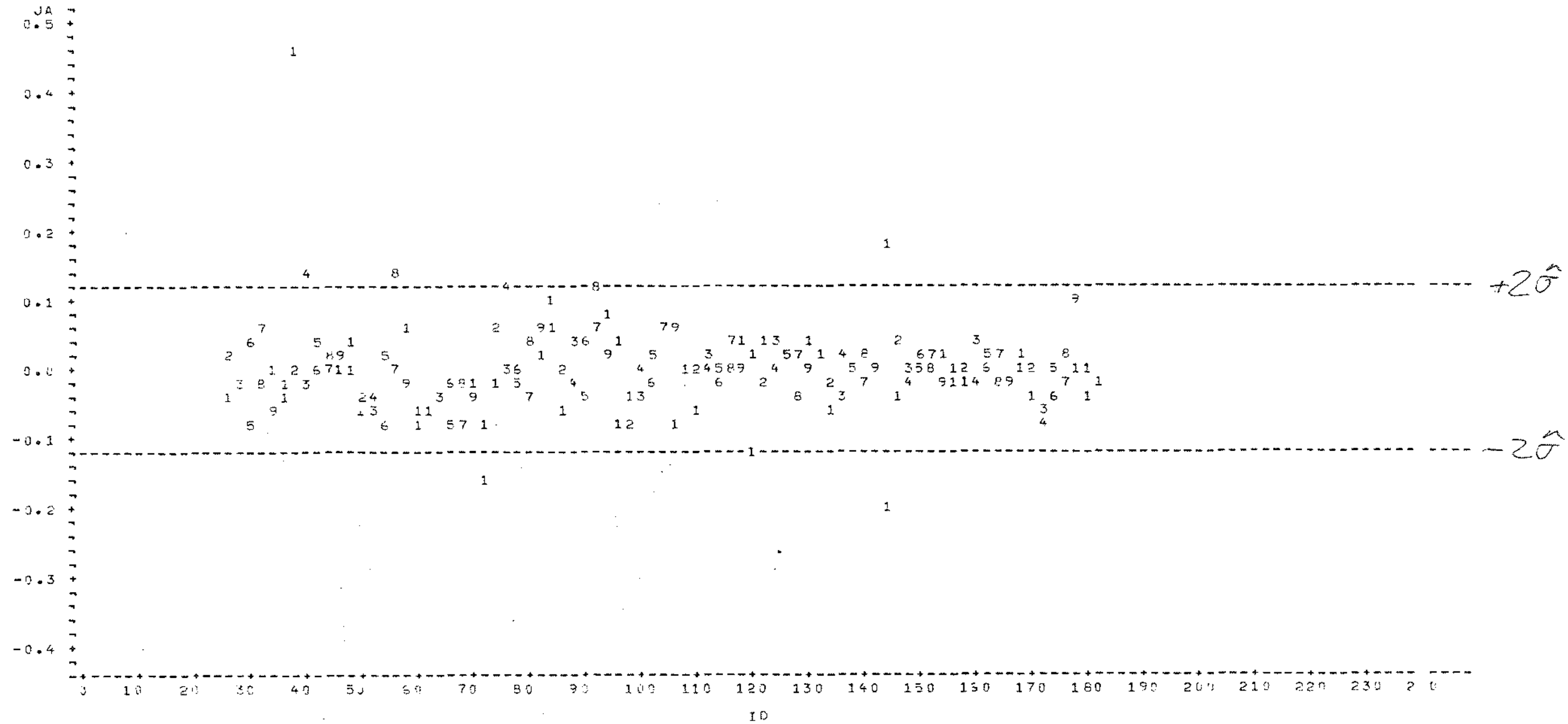
PLOT OF US*ID   SYMBOL IS VALUE OF MONTH



NOTE: 5 OBS HIDDEN

FIGURE2
TIME SERIES PLOT OF ERRORS
JA DATA 1-1969 TO 1-1982

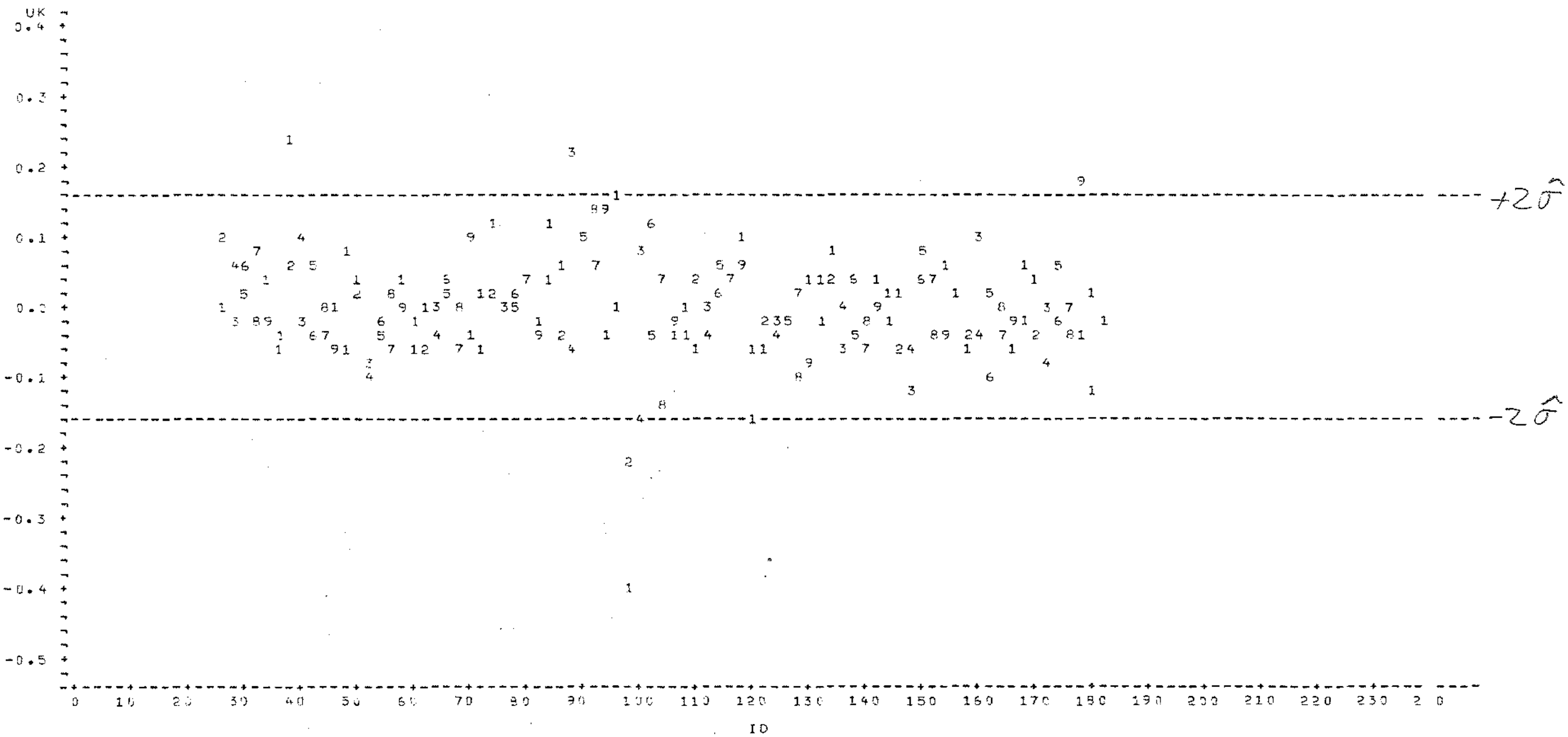
PLOT OF JA*ID   SYMBOL IS VALUE OF MONTH



NOTE: 10 OBS HIDDEN

FIGURE 3
TIME SERIES PLOT OF ERRORS
UK DATA 1-1959 TO 1-1982

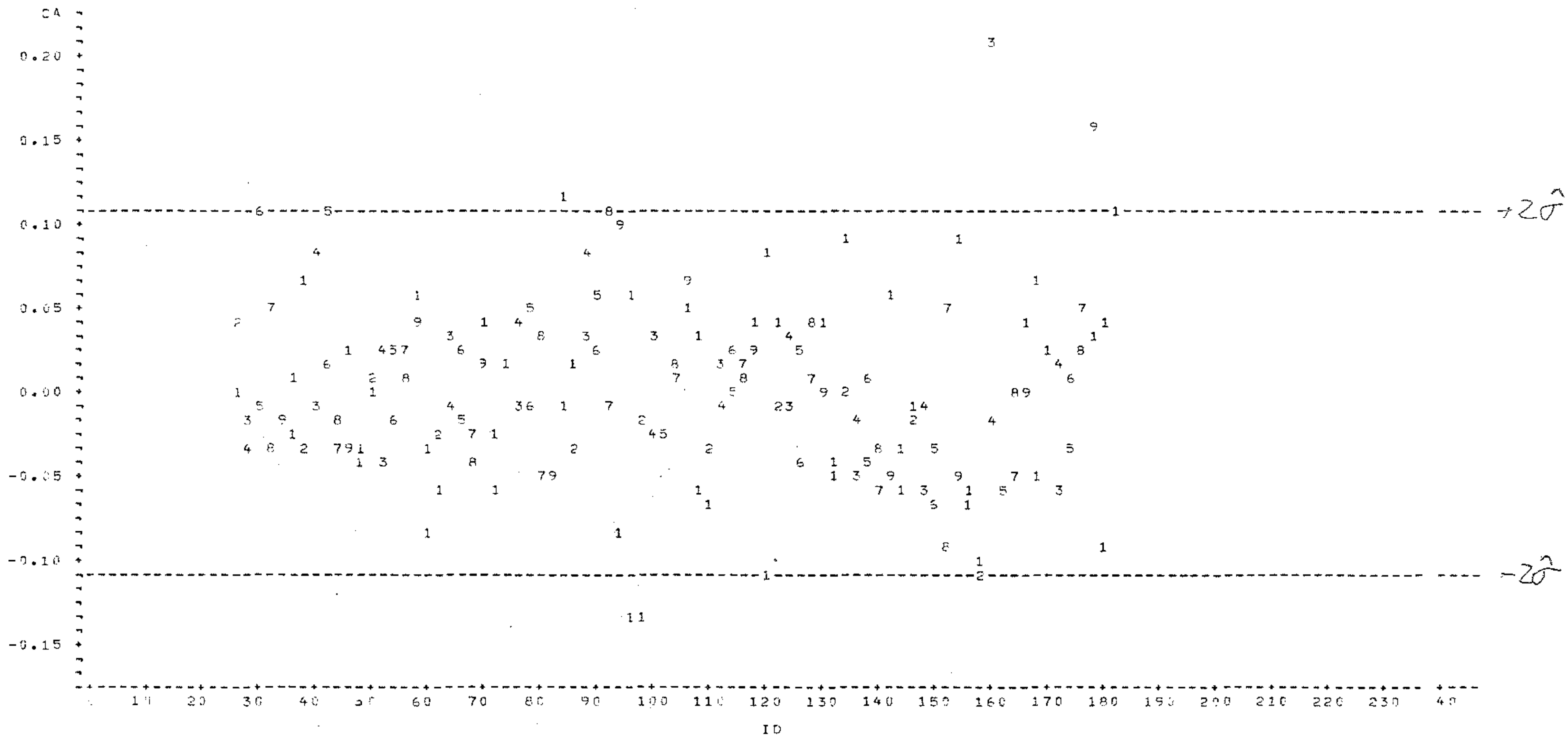
PLOT OF UK*ID SYMBOL IS VALUE OF MONTH



NOTE: 6 OBS HIDDEN

FIGURE 4
TIME SERIES PLOT OF ERRORS
CA DATA 1-1969 TO 1-1982

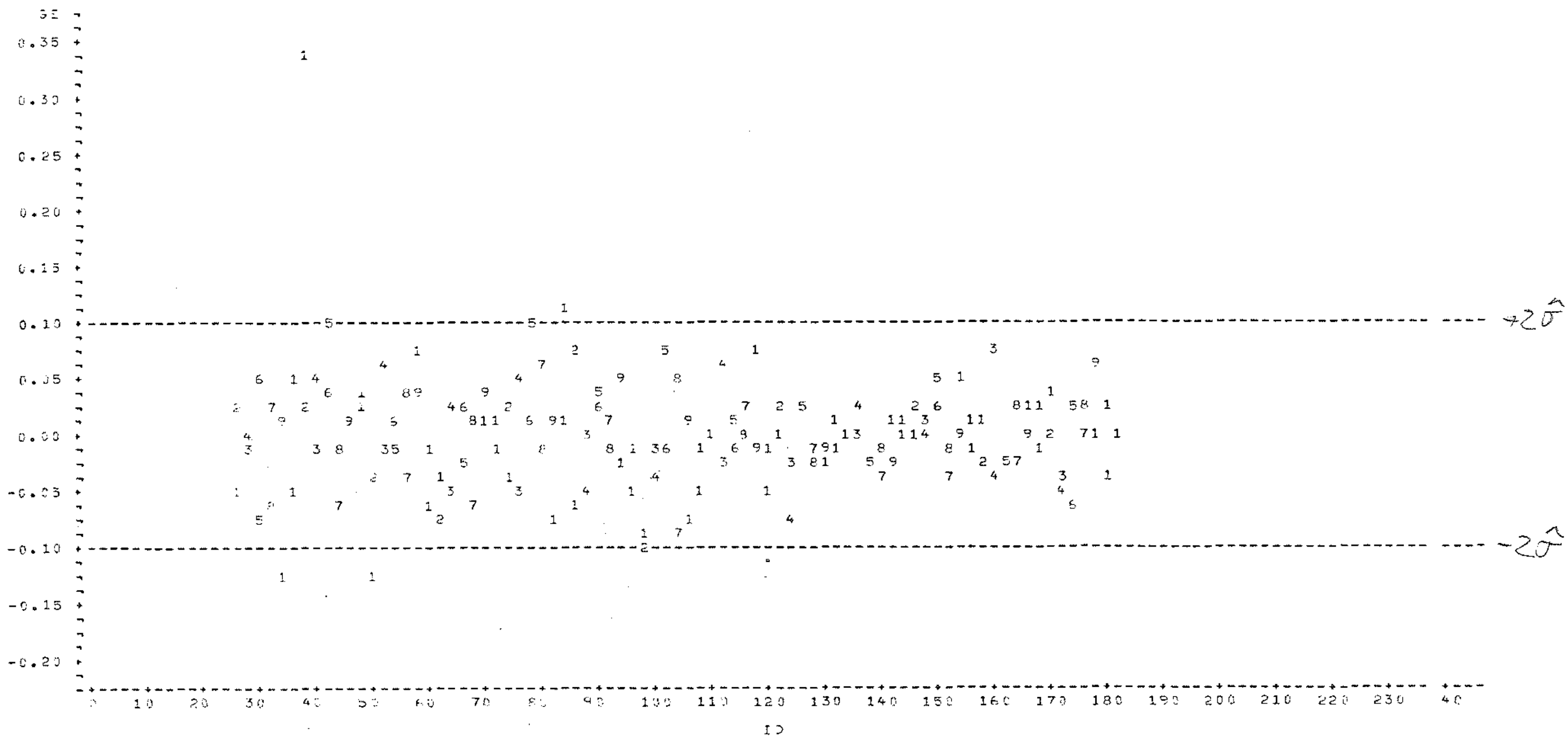
PLOT OF CA*ID   SYMBOL IS VALUE OF MONTH



NOTE: 6 OBS HIDDEN

FIGURE 5
TIME SERIES PLOT OF ERRORS
GE DATA 1-1969 TO 1-1982

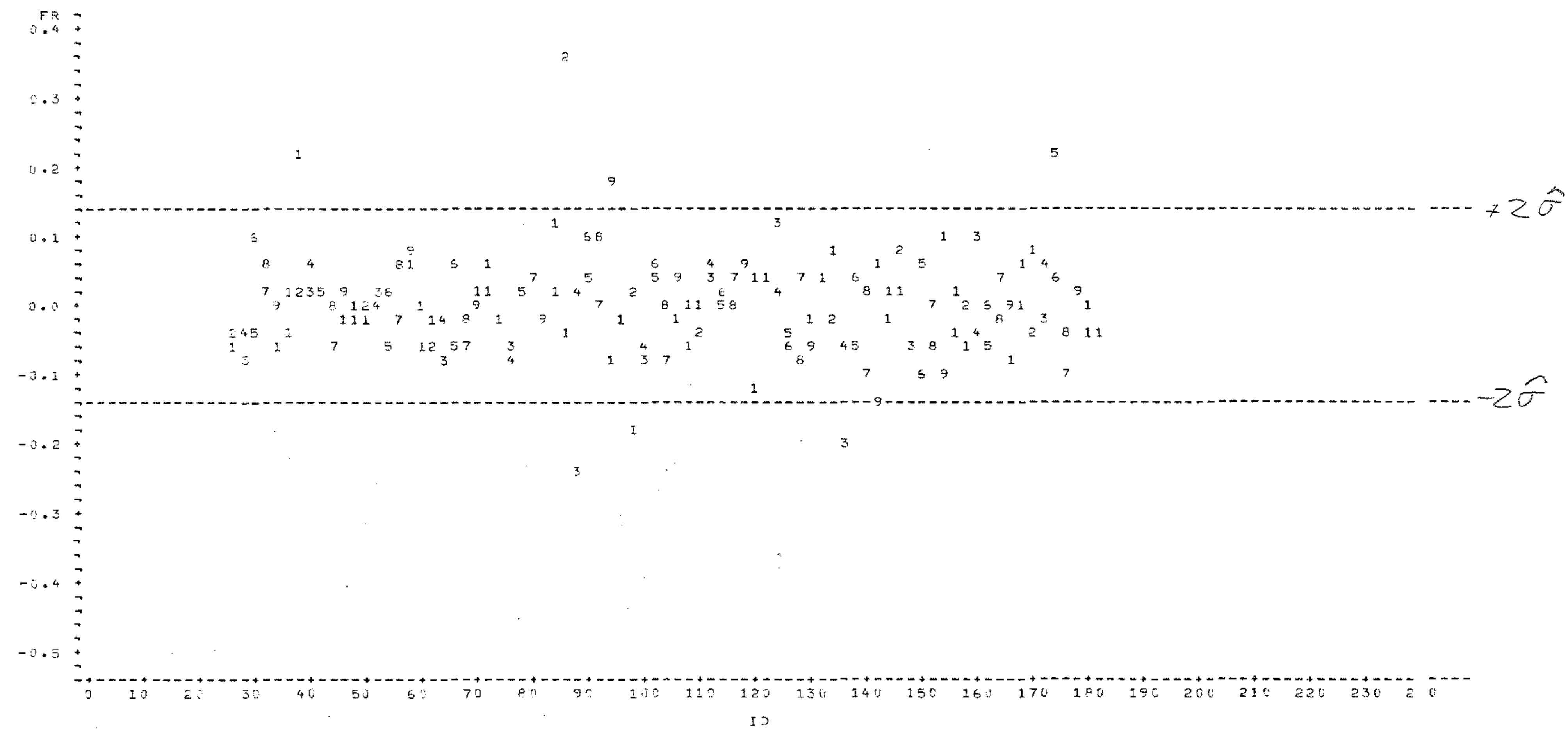
PLOT OF GE*ID   SYMBOL IS VALUE OF MONTH



NOTE: 6 OBS HIDDEN

FIGURE 6
TIME SERIES PLOT OF ERRORS
FR DATA 1-1969 TO 1-1982

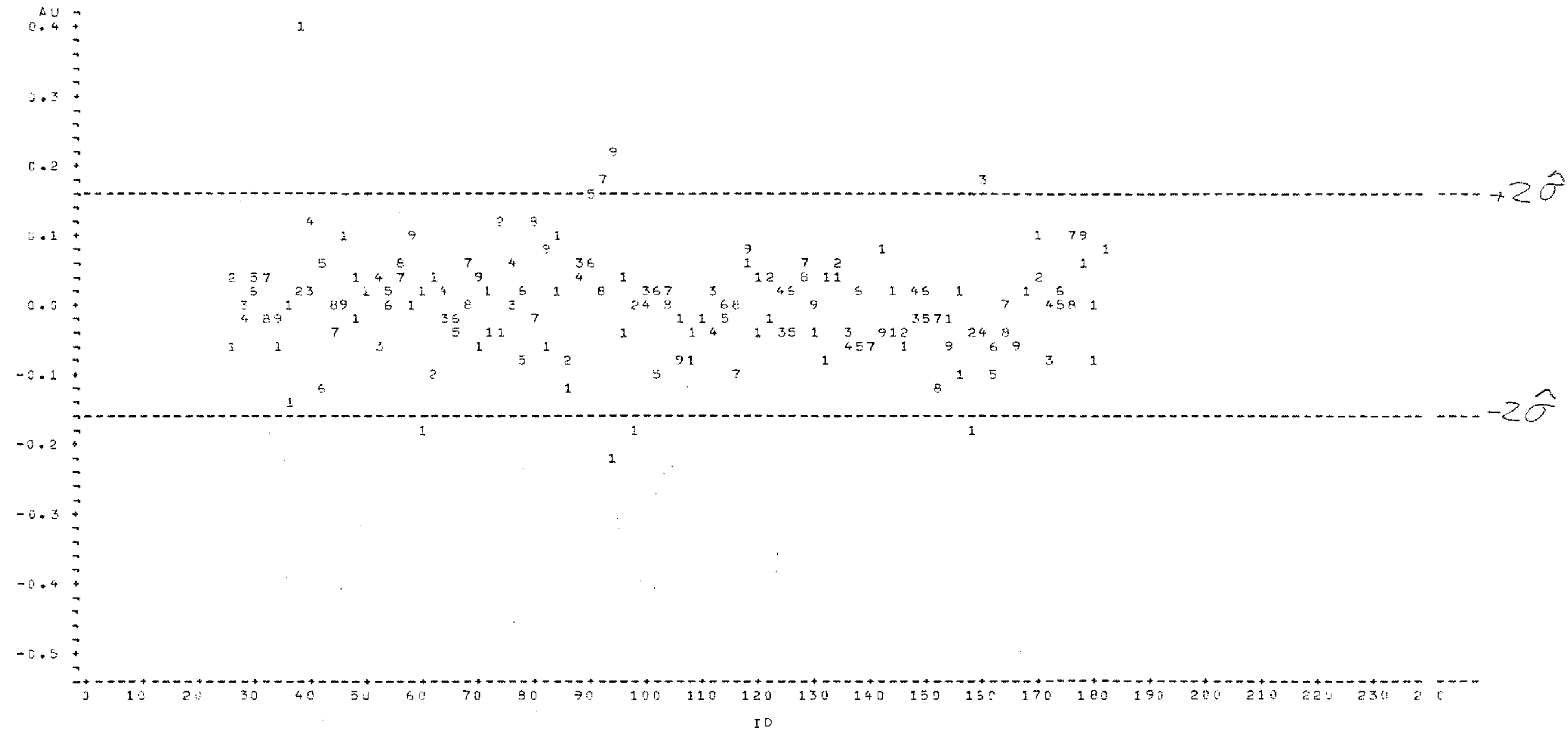
PLOT OF FR*ID   SYMBOL IS VALUE OF MONTH



NOTE: 11 OBS HIDDEN

FIGURE 7
TIME SERIES PLOT OF ERRORS
AJ DATA 1-1969 TO 1-1982

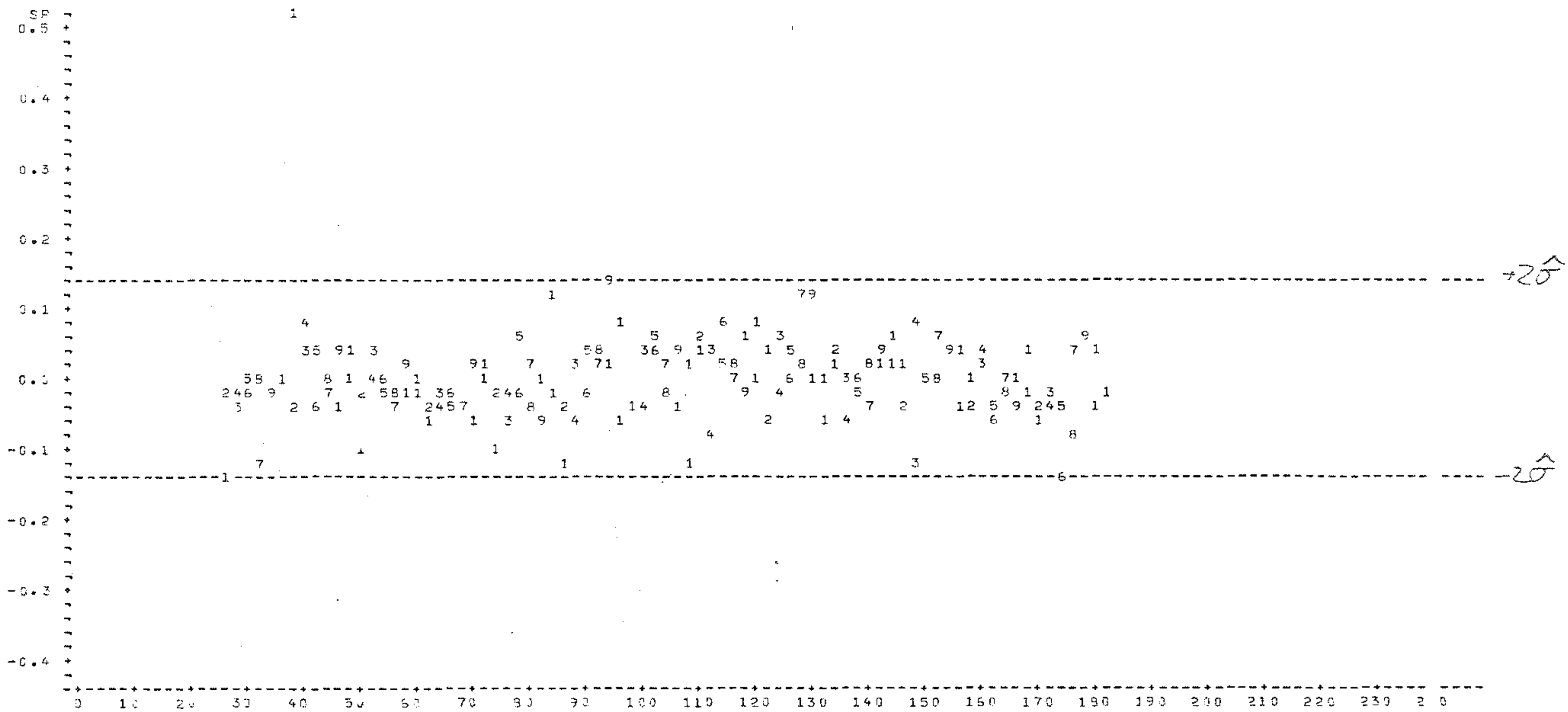
PLOT OF AU*ID   SYMBOL IS VALUE OF MONTH



NOTE: 5 OBS HIDDEN

FIGURE 8
TIME SERIES PLOT OF ERRORS
SP DATA 1-1969 TO 1-1982

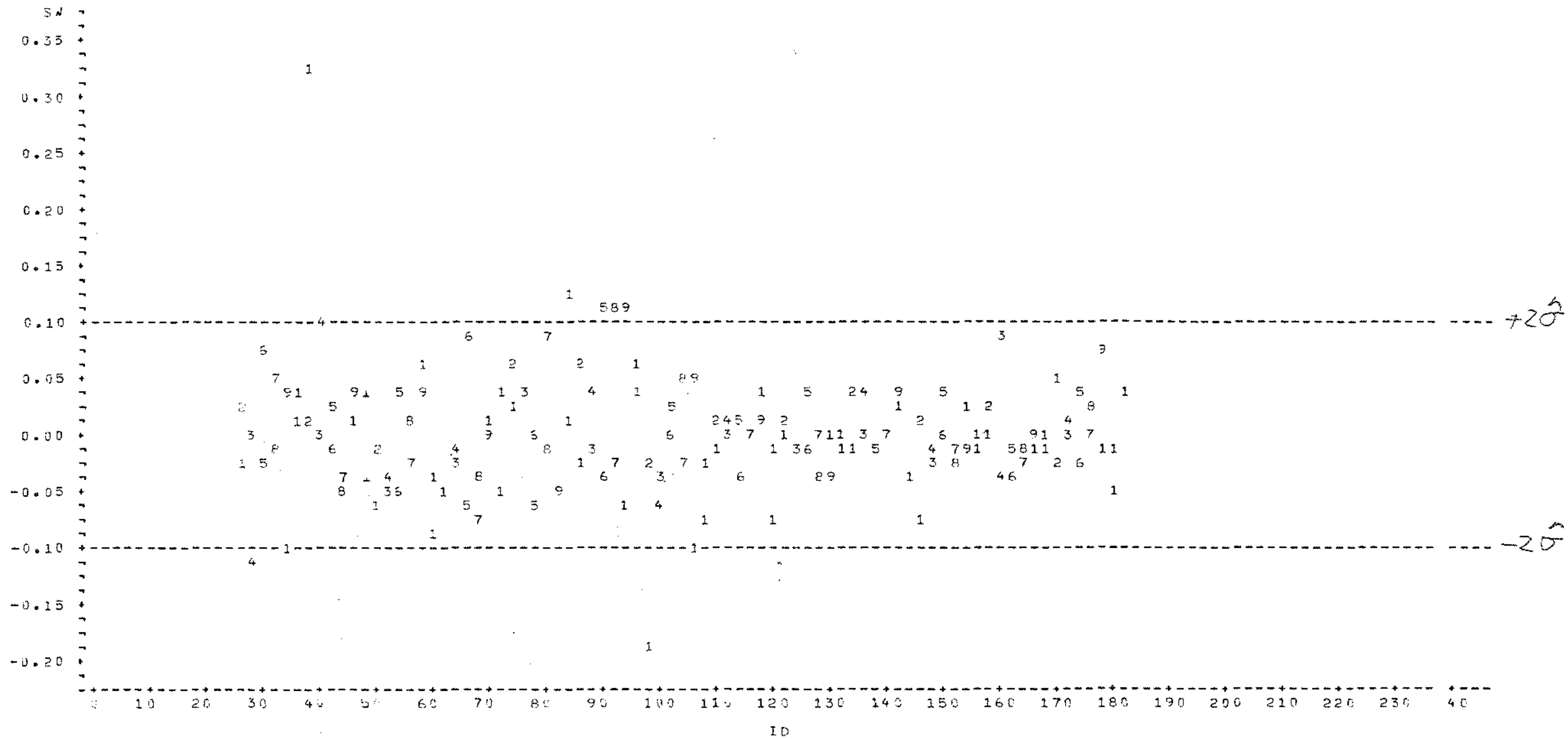
PLOT OF SP*ID SYMBOL IS VALUE OF MONTH



NOTE: 7 OBS HIDDEN

FIGURE 9
TIME SERIES PLOT OF ERRORS
SW DATA 1-1969 TO 1-1982

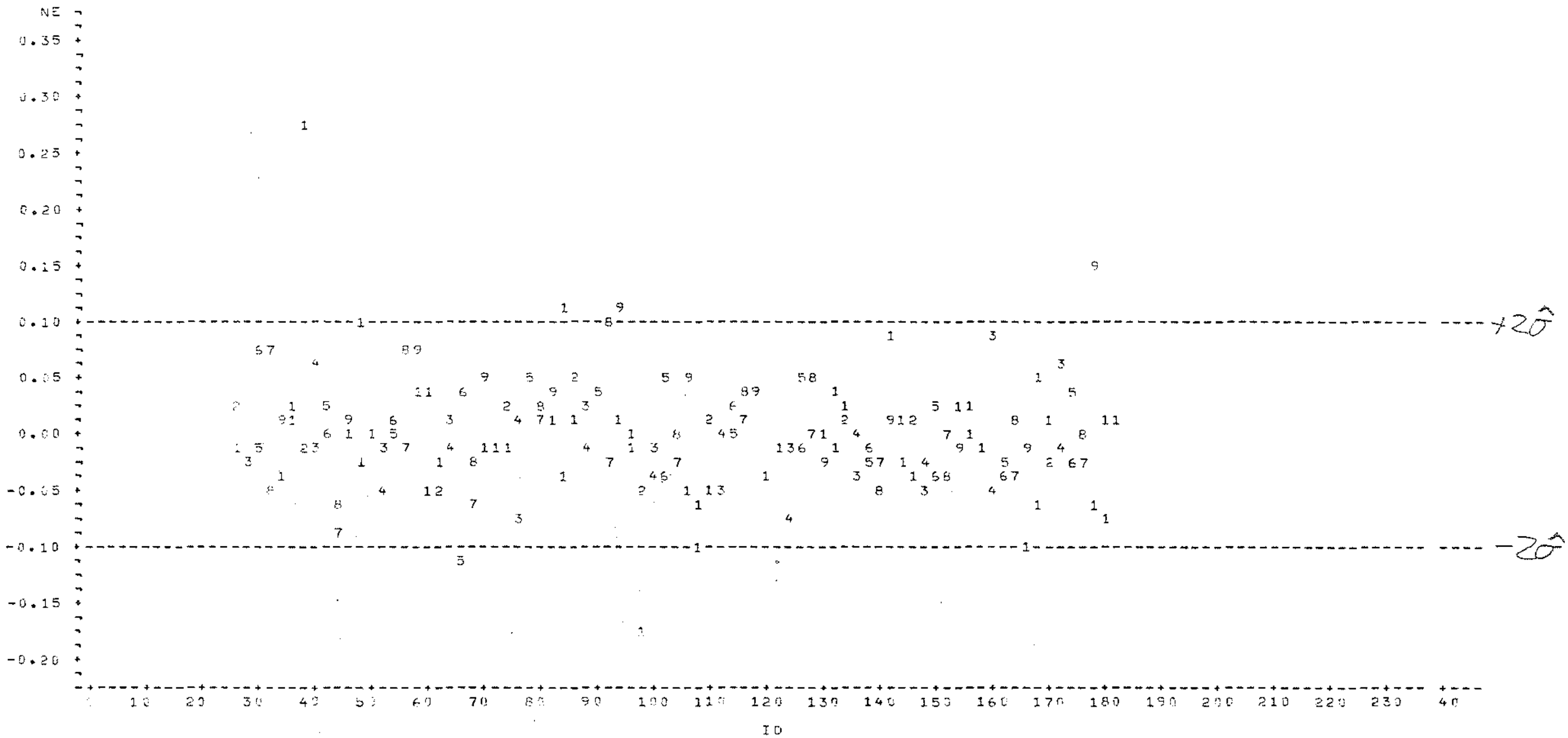
PLOT OF SW*ID   SYMBOL IS VALUE OF MONTH



NOTE: 3 OBS HIDDEN

FIGURE 10
TIME SERIES PLOT OF ERRORS
NE DATA 1-1969 TO 1-1982

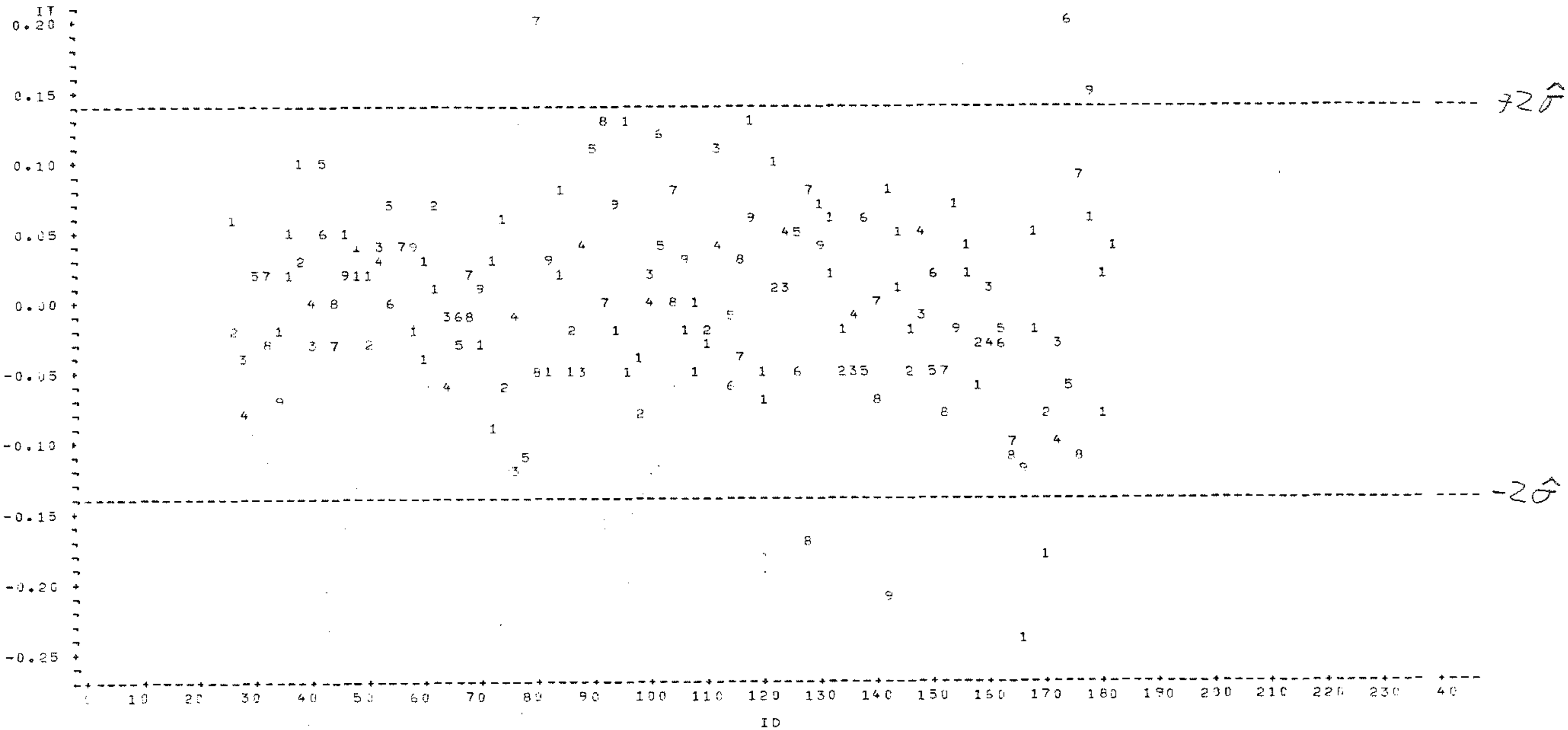
PLOT OF NE*ID   SYMBOL IS VALUE OF MONTH



NOTE: 9 OBS HIDDEN

FIGURE 11
TIME SERIES PLOT OF ERRORS
IT DATA 1-1969 TO 1-1982

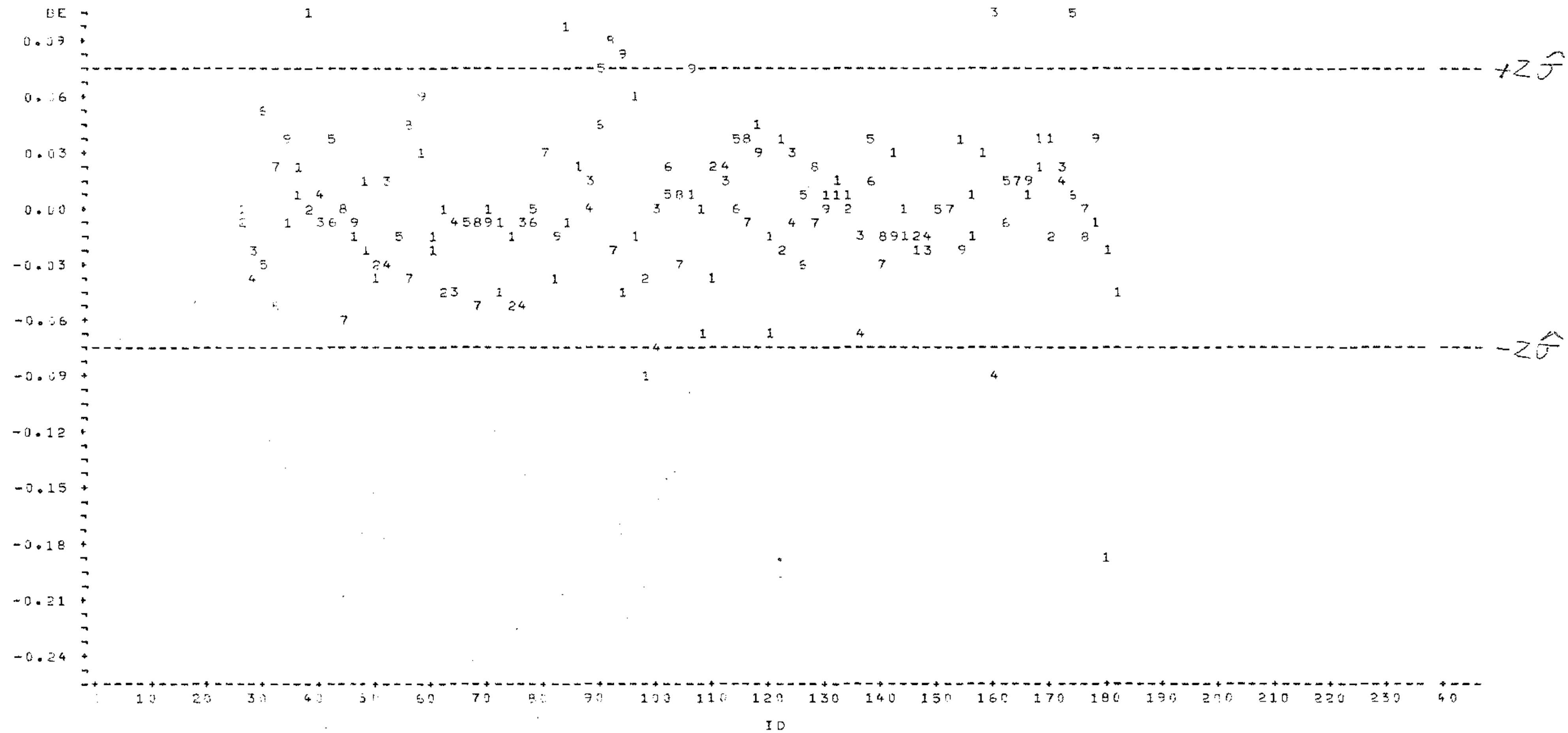
PLOT OF IT*ID   SYMBOL IS VALUE OF MONTH



NOTE: 4 OBS HIDDEN

FIGURE 12
TIME SERIES PLOT OF ERRORS
BE DATA 1-1969 TO 1-1982

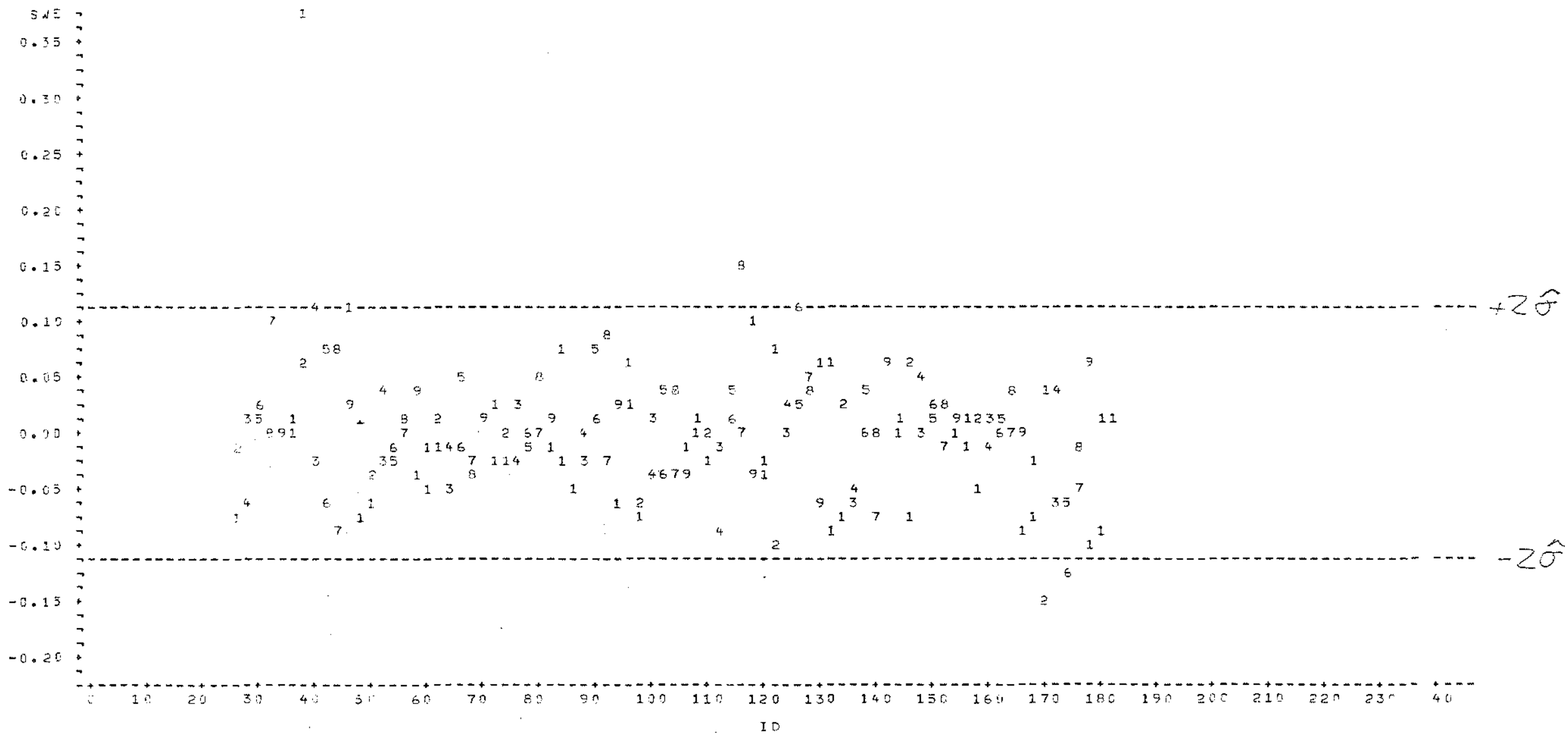
PLOT OF BE*ID SYMBOL IS VALUE OF MONTH



NOTE: 8 OBS HIDDEN

FIGURE 13
TIME SERIES PLOT OF ERRORS
SWE DATA 1-1969 TO 1-1982

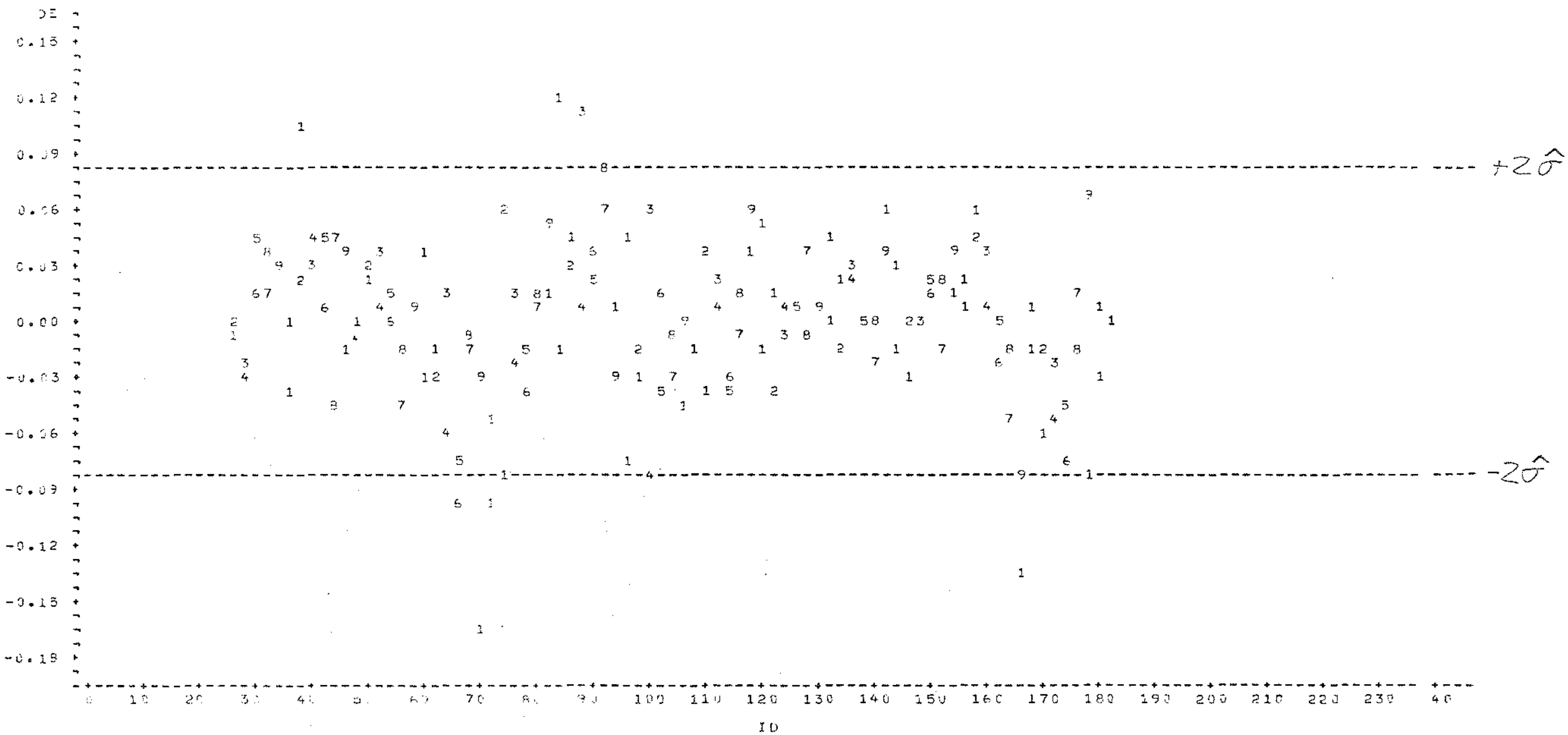
PLOT OF SWE*ID   SYMBOL IS VALUE OF MONTH



NOTE: 4 OBS HIDDEN

FIGURE 14
TIME SERIES PLOT OF ERRORS
DE DATA 1-1969 TO 1-1982

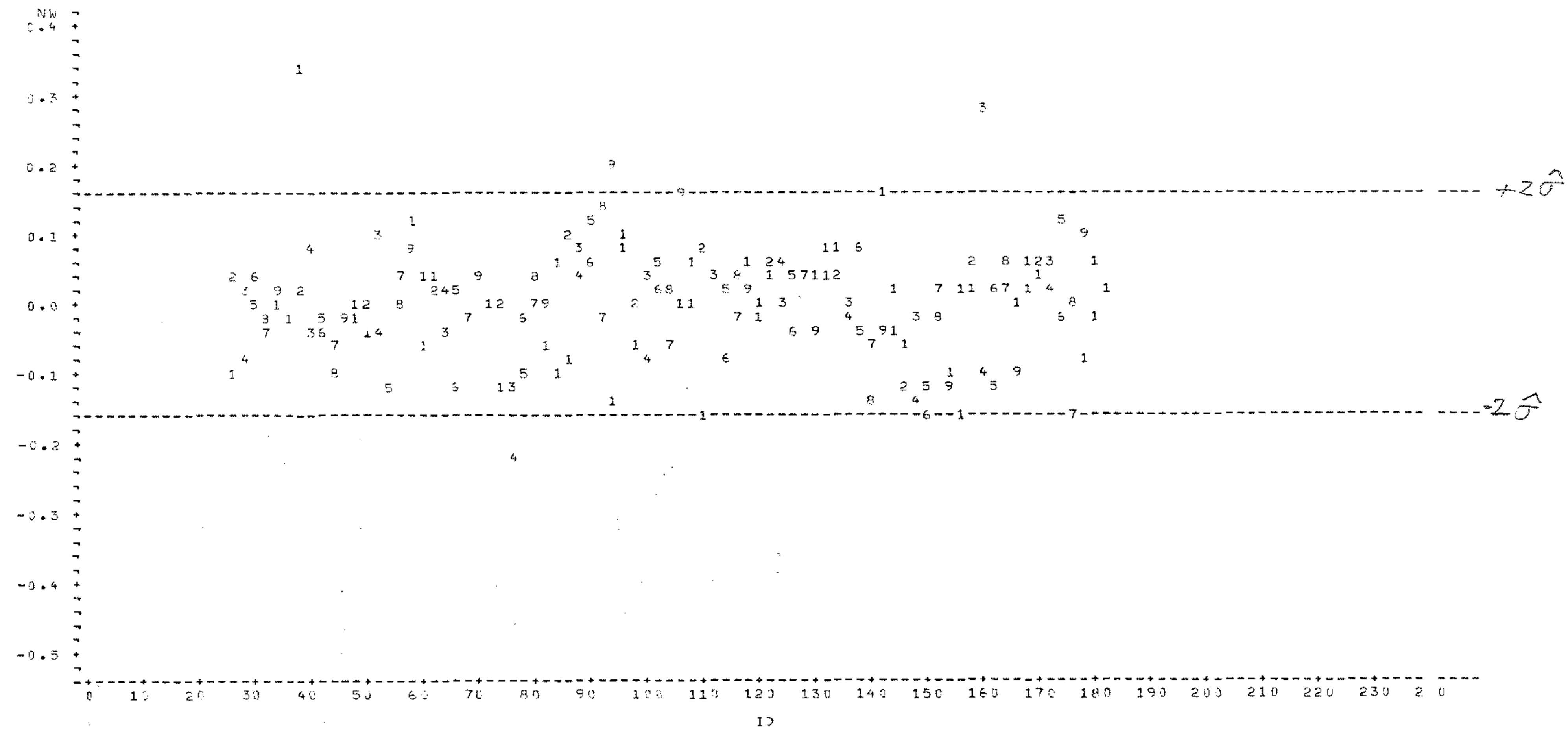
PLOT OF DE*ID   SYMBOL IS VALUE OF MONTH



NOTE: 7 OBS HIDDEN

FIGURE 15
TIME SERIES PLOT OF ERRORS
NW DATA 1-1969 TO 1-1982

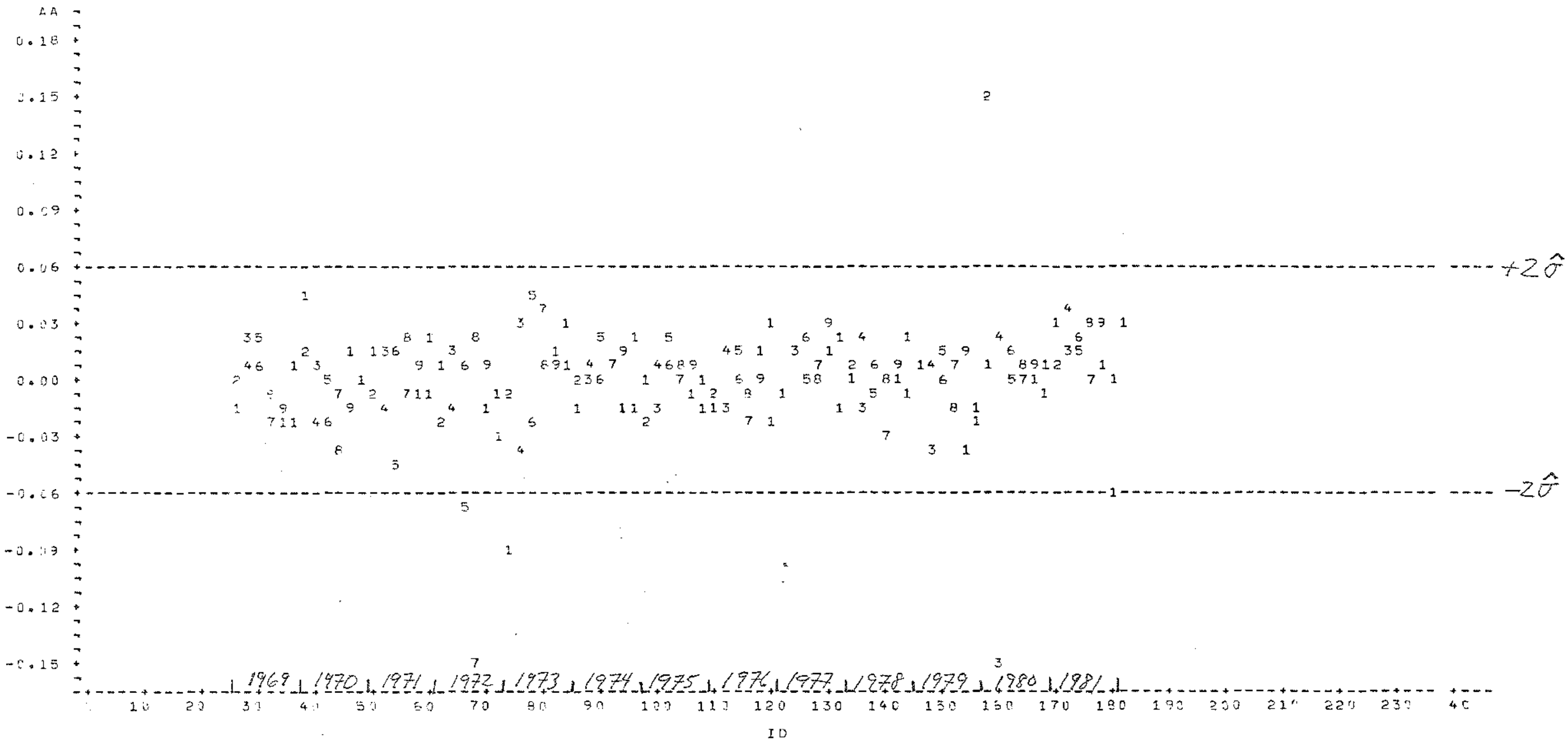
PLOT OF NW*ID   SYMBOL IS VALUE OF MONTH



NOTE: 8 OBS HIDDEN

FIGURE 16
TIME SERIES PLOT OF ERRORS
AA DATA 1-1969 TO 1-1982

PLOT OF AA*ID   SYMBOL IS VALUE OF MONTH



NOTE: 5 OBS HIDDEN

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