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Northwest Atlantic Fisheries Organization

Serial No. N845

NAFO SCR Doc. 84/VI/57

SCIENTIFIC COUNCIL MEETING - JUNE 1984

Temperature-Yield Relationships for the Maine American Lobster (*Homarus Americanus*) Fishery: A Time Series Analysis Approach

by

Michael J. Fogarty

National Marine Fisheries Service, Northeast Fisheries Center
Woods Hole Laboratory, Woods Hole, Massachusetts, USA 02543

ABSTRACT

Autoregressive-integrated moving average (ARIMA) models were developed for annual Maine lobster landings for two time periods: (1) 1928-81 and (2) 1945-81. During the latter period, landings were assumed to be less dependent on fishing effort due to trap saturation. Monthly Maine lobster landings for the period 1968-81 were also examined and a seasonal model was developed. Univariate time series models for annual yield provided reliable forecasts for 1982 catch levels, differing by no more than 4% of the actual 1982 yield. Monthly forecasts for 1982 were characterized by a mean absolute error of 12% when compared to actual monthly catch. However, the cumulative monthly forecast for 1982 differed from total 1982 landings by less than 1%. Development of transfer function models including a lagged temperature effect were found to be warranted only for the 1945-81 annual yield series. A significant temperature effect at a lag of six years resulted in a reduction in residual variance of approximately 13% relative to the corresponding univariate model. Various possible mechanisms of temperature dependence are evaluated.

INTRODUCTION

Variability in yield and productivity of harvested fish and invertebrate populations has frequently been attributed to environmental influences. Cushing (1983) provided a recent comprehensive summary of environmental effects on trends in abundance and/or yield for a number of marine species.

Temperature-related influences in particular have been examined in considerable detail, undoubtedly due to the availability and consistency of extensive water temperature records for several locations. Although temperature may have direct effects on survival, growth, and distribution of marine organisms, it may also have latent influences (e.g., effects on primary or secondary production) or may merely be correlated with other variables of greater importance.

The relationship between temperature and yield of fish and invertebrate stocks in the Gulf of Maine has been examined in detail (Taylor et al. 1957; Dow 1964, 1977; Sutcliffe et al. 1977) with particular emphasis on the American lobster, Homarus americanus (Dow 1964, 1969, 1976, 1977, 1978; Flowers and Saila 1972; Orach-Meza and Saila 1978). Flowers and Saila (1972) developed multiple regression predictors for Maine lobster yield based on present and lagged temperature values. Orach-Meza and Saila (1978) constructed a sophisticated polynomial distributed lag model for forecasting Maine lobster catches. Earlier studies were based primarily on correlation or simple linear regression.

In the present report, an alternative modeling and forecasting approach is explored based on recent developments in the statistical analysis of time series. Specifically, the use of univariate and multivariate autoregressive-integrated-moving-average (ARIMA) models (Box and Jenkins 1976) for modeling Maine lobster landings is considered. Boudreault et al. (1977), Saila et al. (1980), Mendelssohn (1981), and Kirkley et al. (1982) provide recent applications of ARIMA models to fishery forecasting problems. Stocker and Hillborn (1982) and Roff (1983) discuss the use of autoregressive models for predicting catch levels.

In general, ARIMA models are a flexible and powerful class of linear stochastic difference equation predictors. The reader is referred to Box and Jenkins (1976) and Anderson (1975) for an overview of the methods and philosophy employed in the Box-Jenkins approach.

METHODS

Data Base

An uninterrupted series of annual catch records was available for the Maine lobster fishery for the period 1928-81 (Figure 1; Dow 1976; Thomas

1983), providing the basis for a long-term analysis of trends in coastal Maine lobster catches. Time series models were examined for the entire period and for a more restricted period (1945-81). Continual escalation in the number of traps fished resulted in sharp increases in yield prior to 1945. Flowers and Saila (1972) suggested that trap saturation after about 1945 reduced the effect of fishing intensity on yield. Accordingly, the 1945-81 series was considered to be less dependent on fishing effort. Due to the highly nonlinear relationship between yield and effort, it was not possible to directly incorporate effort into the time series model.

A monthly series of catch records from January 1968 - December 1981 (Thomas 1983) was used to develop a seasonal model for the Maine lobster fishery.

Water temperature has been monitored daily at the fishery research station at Boothbay Harbor, Maine since 1905 (Welch 1983). Prior to October 1950, thermometer readings were used to record surface temperature several times daily; since then, continuous thermister readings have been taken at a depth of 5.5 m (Welch 1983). Despite the potential bias introduced by changes in methodology, the Boothbay Harbor temperature series agrees well with other temperature records for this period in the Northeastern United States (Sissenwine 1974; Skud 1982). Mean annual temperature for the period 1928-81 is illustrated in Figure 1.

Analysis

Statistical procedures followed the iterative Box-Jenkins approach of model identification, estimation, and checking (see Saila et al. 1980; and Mendelssohn 1981 for fishery-related examples). Model identification is based on diagnostic characterisics of the autocorrelation and partial autocorrelation functions of a stationary time series (Box and Jenkins 1976). The series must be stationary both in level and variance; transformation may be required to ensure homogeneity of variance. To check for violations of the assumption of variance stationarity, each series was first arbitrarily divided into segments and mean-range and mean-variance plots were examined; transformation was not required.

For series which are nonstationary in level, it may be necessary to difference the series (i.e., form the series $z_t = x_t - x_{t-n}$ where z_t is the differenced series, x_t is the original variable, and (n) is the period of differencing)[1]. Once stationarity is assured, an appropriate model containing autoregressive and/or moving average terms may be specified. The general model form is:

$$z_t = \phi_1 z_{t-1} + \phi_2 z_{t-2} \cdots \phi_p z_{t-p} + a_t - \theta_1 a_{t-1} - \theta_2 a_{t-2} \cdots \theta a_{t-q} + \theta_0$$

where z_t is the time series variable (after differencing and/or transformation if necessary); a_t are independent normally distributed random shocks (error terms) with zero mean and constant variance; ϕ_i are autoregressive parameters and θ_i are moving average parameters subject to bounds of stationarity and invertibility respectively (Box and Jenkins 1976); and θ_0 is a deterministic trend parameter.

Univariate models were first fit for the two annual catch series (1928-81 and 1945-81) and for the monthly catch series (1968-81) to provide a basis for further comparison. Following model identification and estimation, extra parameters were added and tested for significance to ensure that a full model was specified. Catch data for 1982 were held in reserve to check forecasted values.

In general, analyses were performed on the "centered" series (i.e., the mean was subtracted from each observation). Following tentative model identification, maximum liklihood estimates of the parameters were made and the residuals of the model checked to ensure independence and normality. Independence of residuals was checked using the portmanteau test of Box and Jenkins (1976). Normality of residuals was examined using the Kolmogorov-Smirnov statistic (D) for large sample sizes ($N > 50$) and the Shapiro-Wilk statistic (W) for $N < 50$.

Development of multivariate models depends on model identification based on the cross-correlation between the stationary dependent and independent variables. It is recommended that the series be "pre-whitened" prior to analysis (Box and Jenkins 1976); in the bivariate case, a univariate model is developed for the independent (input) variable, inverted, and applied to both variables to remove any within-series correlation. Thus the problem of correlating two series, each of which is autocorrelated, is eliminated (see

[1] For notational convenience, it is possible to define the "backshift" operator, B^n, which indicates lagged variables ($B^n x_t = x_{t-n}$). The backshift operator may be used to indicate a differenced series viz: $(1-B^n)x_t = x_t - x_{t-n}$.

Gulland 1965 and Ricker 1975 for discussion of the dangers of relating two serially correlated variables with specific reference to fishery-related problems.

The general form of a bivariate model is

$$z_t - \delta_1 z_{t-1} - \delta_2 z_{t-2} \cdots \delta_r z_{t-m} = \omega_o y_{t-b} - \omega_1 y_{t-b-1} \cdots \omega_n y_{t-b-n} + n_t$$

where the δ_i and ω_i are model parameters and n_t is an error term which may be modeled as an ARIMA process. Examples of multivariate ARIMA models in biology are provided by Hacker et al. (1975), Jenkins (1976), Poole (1976) and Mendelssohn (1981). The reader is urged to consult Box and Jenkins (1976) for a detailed description of the method.

RESULTS

Annual Series (1928-1981)

To provide a basis for comparison with multivariate models, a univariate model was first constructed for the 1928-1981 catch series. The slow decay of the autocorrelation function of the original series clearly indicated non-stationarity; accordingly, first differences were taken (Figure 2). Examination of the autocorrelation and partial autocorrelation functions of the differenced series indicated that a first order autoregressive model was appropriate (Figure 2). Parameter estimates for the final model are provided in Table 1. No significant autocorrelation in the residuals was detected and the assumption of normality was not rejected (χ^2 = 26.45; df = 23; P = .280) (Table 1). Attempts to "overfit" the model by inclusion of additional autoregressive and moving average terms indicated that a more detailed model was unnecessary. The final model was of the form:

$$(1 - \phi_1 B)\ z_t = a_t$$

where z_t is the differenced series. A comparison of the observed and fitted series is provided in Figure 3. Despite the simplicity of the model, it is clear that the dynamic behavior of the catch series is reasonably well represented. As a further check on the adequacy of the model, the observed 1982 catch (10,263 mt) was compared to the forecast of 10,341 mt (95% CI:

8716.1, 11965.8). The predicted yield differed by less than 1% from the observed catch.

Development of a transfer function model incorporating temperature effects required prewhitening to remove within-series correlation. A first order autoregressive model was developed for the differenced temperature series (x_t) based on the autocorrelation and partial autocorrelation functions. The final temperature model was

$$(1 + 0.3916B)\, x_t = a_t$$

Residuals of the temperature model did not differ significantly from white noise ($Q = 22.15$; $df = 22$; $P > .51$). The temperature model was applied to both series and the crosscorrelation between the residuals of both series examined. The crosscorrelation between the original (undifferenced) series was characterized by large positive correlations, however, the crosscorrelation between the prewhitened series showed only marginally significant correlations at lags 1 and 9 (Figure 4). Parameter estimates for temperature effects at these lags, however, were non-significant and no appreciable reduction in the residual variance was achieved. It was therefore concluded that inclusion of temperature terms in the model was unwarranted.

Annual Series (1945-1981)

In contrast to the 1928-81 annual series, the stationarity of the shorter series was not clearly defined. For simplicity, the undifferenced series was used in subsequent analyses. It was recognized that some analysts might choose to work with the differenced series. The relatively short time series available undoubtedly contributed to this difficulty.

The pattern of autocorrelations and partial autocorrelations (Figure 5) suggested a second order autoregressive model. Estimation and subsequent diagnostic checking of the second order model indicated that the residuals were independent, normally distributed random variables (Table 2). A comparison of the observed and fitted catch series is provided in Figure 6. The 1982 forecast estimate of 9875.0 mt (95% CI: 8417.1-11332.9) underestimated the actual 1982 catch by 4%.

A second order autoregressive model was found to be appropriate for the undifferenced 1945-81 temperature series. The final temperature model was:

$$(1 - 0.4454B - 0.3287B^2)\, y_t = a_t$$

The model residuals were found to be independent (Q = 15.62; df = 21; P = .79) The temperature model was used to filter both series and crosscorrelations were computed for the prewhitened series. Significant correlation at lags 0 and 6 were observed for the differenced series (Figure 7). A transfer function model incorporating the lagged temperature effects was specified and the noise component was modeled based on the autocorrelation function of the residuals of the preliminary model. On further testing, the temperature term at lag 0 was found to be unnecessary. The final model included temperature effect at lag 6 yr and first order autoregressive terms were used to model the noise component. Incorporation of the temperature term resulted in a 13.2% reduction in residual variance relative to the univariate model. Comparisons between the fitted univariate and transfer function models and the observed 1945-81 series are presented in Figure 6. Parameter estimates for the model:

$$(1 - \phi_1 B)\, x_t = \omega_0 y_{t-6}$$

are presented in Table 3.

Monthly Series (1968-1981)

A clearly defined seasonal pattern was observed for the 1968-81 monthly catch series (Figure 8). Catches typically peaked during September-October. Interestingly, a minor peak often occurred in May. It is possible that reduced catchability during the June molting period resulted in depressed catches relative to May levels.

The undamped sinusoidal pattern of the autocorrelation function for the monthly catch series indicated the need for seasonal differencing with a period of 12 months (Figure 9). Further differencing and transformation was found to be unnecessary. Autocorrelation and partial autocorrelation functions of the seasonally differenced series (Figure 9) suggested a moving average model of the form

$$z_t = (1-\theta B_1)\,(1-\Theta B^{12})\, a_t\ .$$

Examination of the model residuals indicated no departure from the assumptions of independence (Table 4). The hypothesis of normality of residuals, however, was rejected; although the predictive capability of the model is relatively robust to non-normality, it should be noted that hypothesis tests might be adversely affected.

Comparison of observed and fitted observations for the period January 1967 - December 1981 are provided in Figure 10; observed and forecasted values for the period January - December 1982 are depicted in Figure 11. Forecasted values differed from observed 1982 monthly catches by a maximum of 20% and the mean absolute forecast error was 12%.

Following seasonal differencing of the temperature series, an autoregressive model of the form:

$$(1-\theta B)\ (1-\theta B^{12} - \theta B^{24})\ z_t = a_t$$

was identified and fit. The prewhitened monthly catch and temperature series were then crosscorrelated. Although crosscorrelations between the original series were relatively high, no significant crosscorrelations were observed between the prewhitened series (Figure 12). Accordingly, further development of a seasonal transfer function model was deemed inappropriate.

DISCUSSION

The Box-Jenkins approach to time series analysis and modeling is an extremely flexible and adaptable technique with important applications in fishery forecasting (Saila et al. 1980; Mendelssohn 1981). Relatively simple autoregressive-integrated-moving-average models are capable of capturing complex dynamic behavior and of providing forecasts with relatively high levels of precision. In the present analysis, forecasts for the 1982 total annual Maine lobster catch differed by no more than 4% of the realized yield. In addition, a model based on monthly catches predicted the seasonal pattern of catches with reasonable accuracy. The annual total of the monthly forecasts for 1982 (10,864 mt) compared well with the 1982 catch of 10,263 mt. Forecast accuracy provided by these models may have been somewhat fortuitous, nevertheless, the general utility of the Box-Jenkins approach to modeling fishery dynamics is clear, as previously indicated by Saila et al. (1980) and Mendelssohn (1981).

Results of the present analysis differ from previous studies with respect to the relative importance of temperature on coastal Maine lobster yield. Highly significant correlations between Maine lobster yield and lagged temperature values have been reported (e.g., Dow 1976, 1977, 1978). Unfortunately, autocorrelation in both the catch and temperature series considerably complicates the analysis. Significance levels were undoubtedly overestimated in these correlative studies.

In the present study, an attempt was made to adjust for within-series correlation. The lack of significance of temperature terms in the 1928-81 annual yield model and the seasonal model does not imply that the apparent relationship between temperature and yield noted in previous studies is necessarily spurious. There does appear, however, to be little additional information in the temperature series relative to the past catch history.

It should be noted that prewhitening may in fact result in an underestimate of crosscorrelation significance levels (Box and Pierce 1970). In the present analysis, it was recognized that prewhitening results in a generally conservative approach. This would appear to be an acceptable forecasting strategy since reasonably accurate predictions are possible based on yield alone. Nevertheless, environmental inputs may clearly play an important role in forecasting and, more importantly, may allow a mechanistic interpretation of variability in abundance levels.

The time series examined in this report were stationary in variance and transformation was unnecessary. However, it might be noted that since environmental effects are expected to be multiplicative (Ricker 1975), a logarithmic transformation of catch and temperature variables might be desirable. Since earlier investigations of lobster yield-temperature relationships have been made on untransformed variables, this approach has been maintained in the present analysis for comparison.

Flowers and Saila (1972) developed several predictive multiple linear regression equations for the Maine lobster fishery based on combined mean annual temperature lagged 6-8 yrs and mean temperature for the months of January-March, at Boothbay Harbor. Equations were also developed from Nova Scotia lobster landings using various lagged mean annual or monthly water temperatures measured at Lurcher Lightship. Interestingly, the previous year's landings were included in one of these equations, anticipating the use of autoregressive terms in yield prediction equations. Orach-Meza and Saila (1978) extended the multiple regression approach to a polynomial distributed lag model in which the effect of the independent variable were distributed over time, with some fraction apportioned to each lag period. In the present analysis, higher order transfer functions would be necessary to allow complex behvior of this type. For the 1945-81 annual catch series, a single temperature effect at lag 6 yrs was found to be sufficient.

It is vital that mechanisms underlying proposed environmental effects on fish or invertebrate populations be clearly identified. Direct temperature effects on lobster yield can in fact be reasonably postulated. It is known that temperature directly influences activity levels and hence catchability of the American lobster (McLeese and Wilder 1957; Paloheimo 1963). Short-term increases in yield may therefore accompany increased temperature levels. Further, increased annual probability of molting of sublegal-sized lobsters has been associated with increased water temperature (Ennis 1982); short-term increases in yield may therefore be expected due to increased recruitment into the harvestable size range. The effects of elevated activity and growth levels would be expected to occur at short lag times (primarily 0-1 year) and to have a direct effect on catch (assuming relatively constant levels of effective effort). In the present study, evidence for the importance of temperature at lags of 0-1 year was equivocal. For the 1928-81 series, a marginally significant temperature effect at a lag of one year was noted but did not appreciably increase the predictive capability of the model.

A significant temperature effect at a lag of six years was observed for the 1945-81 series. Previous investigations have suggested that importance of previous temperature levels with lags centered at about 6 years (Flowers and Saila 1972; Dow 1977, 1978; Orach-Meza and Saila 1978). It has been inferred that this represents a temperature mediated effect on larval or early juvenile survival since the expected time to reach legal size is approximately 6 years in the Gulf of Maine. A mechanistic interpretation of the temperature-yield relationships at a lag of 6 years implicitly requires that catch be a measure of recruitment. Total mortality rates for lobsters in coastal Maine waters are indeed extremely high (Anthony 1980) and recruits dominate the Maine lobster catch (Thomas 1983). However, even if catch estimates were known to be free of error, annual variation in mortality rates would result in errors in estimated recruitment based on catch alone. Despite this source of observational error, further consideration of possible mechanisms underlying a lagged temperature effect on recruitment may be instructive. Scarratt (1964) and Caddy (1979) have suggested that the pelagic larvae phase must be completed prior to winter when further molting will be prohibited. Mortality of larvae which do not reach the benthic stage prior to winter is assumed to be complete. Since molting rates are highly temperature dependent (Templeman 1936), increased temperature may increase the probability of successful completion of the larval phase. Alternatively, given the synchrony between the molting and reproductive cycles of female lobsters (Aikin and Waddy 1980), it is possible that increased molt probability may also result in increased reproductive output.

Time series analysis methods have been shown to be a useful adjunct to standard fishery assessment methods such as surplus production modeling (Saila et al. 1980; Mendelssohn 1981). The enhanced forecasting capability of autoregression methods relative to the surplus production approach has been considered in some detail (Stocker and Hillborn 1981; Roff 1983). It is clear, however, that these methods address quite different requirements. Surplus production models attempt to model stock dynamics and to provide insight into optimal harvesting strategies. In contrast, the time series approach need not embody any a priori assumptions relative to the nature of population growth or the relationship between yield and fishing intensity, except to the extent that these are reflected in the catch history. The flexibility of the ARIMA methodology and the capability of including exogenous variables (e.g., environmental effects) certainly contribute to the forecasting success of this approach.

It is of interest that two recent advances in general fishery models (Schnute 1977; Deriso 1980) incorporate autoregressive components, thus linking the time series approach with more heuristic methods. Indeed, the advantage of these two models relative to earlier surplus yield models can be largely attributed to inclusion of autoregressive terms. It is anticipated that further fusion of population models with more sophisticated statistical methods will result in enhanced modeling capability.

ACKNOWLEDGEMENTS

I would like to express my appreciation to Hannah Goodale and Jean Klemm for word processing and to Uvetta Dozier for assistance with graphics.

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Table 1. Parameter estimate and associated statistics for first order autoregressive model for 1928-1945 lobster yield. T is the t-ratio for the parameter, Q is the χ^2 statistic for autocorrelation in the residuals (through lag 24) and D is the Kolmogorov-Smirnov statistic for normality of residuals.

Parameter	Estimate	S.E.	T	Q	D
ϕ	-0.3556	0.1291	2.75[1]	26.45[2]	.06716[3]

[1] Significant at P<.05
[2] Nonsignificant
[3] Nonsignificant

Table 2. Parameter estimates and associated statistics for second order autoregressive model for 1945-81 lobster yield. T is the t-ratio for the parameter, Q is the χ^2 statistic for autocorrelation in the residuals (through lag 24) and W is the Shapiro-Wilk statistic for normality of residuals.

Parameter	Estimate	S.E.	T	Q	W
ϕ_1	0.3888	0.1550	2.51[1]	26.70[2]	0.974[3]
ϕ_2	0.3822	0.1568	2.41[1]		

[1] Significant at P < .05.
[2] Nonsignificant
[3] Nonsignificant

Table 3. Parameter estimates and associated standard errors for transfer function model relating annual 1945-81 lobster yield with mean annual water temperature lagged by 6 yr. T is the t-ratio for the parameter, Q is the χ^2 statistic for autocorrelation in the residuals (through lag 24), and W is the Shapiro-Wilk statistic for normality of residuals.

Parameter	Estimate	S.E.	T	Q	W
ω_0	740.086	177.517	4.17[1]	20.68[3]	0.963[4]
θ_1	0.3336	0.1826	1.83[2]		

[1] Significant at P < .05
[2] Nonsignificant but retained in model
[3] Nonsignificant
[4] Nonsignificant

Table 4. Parameter estimates and associated standard errors for second order moving average model for monthly lobster catches during 1967-81. T is the t-ratio for the parameter, Q is the χ^2 statistic for autocorrelation in the residuals (through lag 30), and D is the Kolmogorov-Smirnov statistic for normality of residuals.

Parameter	Estimate	S.E.	T	Q	D
θ_1	-0.4317	0.0728	-5.93[1*]	19.64[2]	0.120[3]
θ_2	0.3417	0.0808	4.23[1*]		

[1*] Significant at P < .05
[2] Nonsignificant
[3] Significant at P < .05

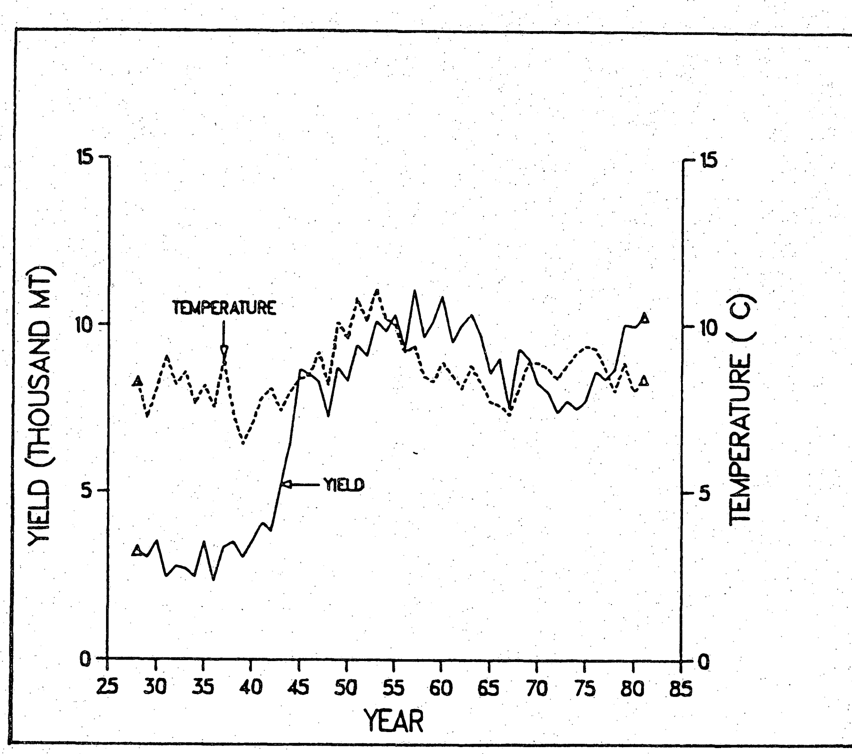

Figure 1. Maine lobster yield (mt x 10^3) and mean annual water temperature at Boothbay Harbor, Maine, for the period 1928-81.

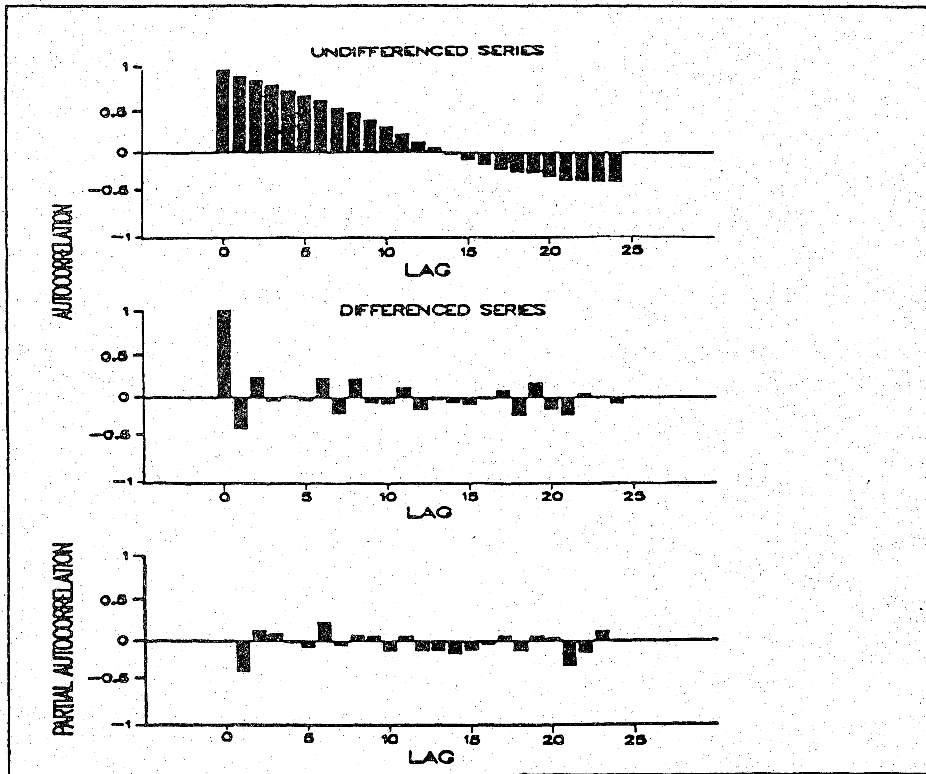

Figure 2. Autocorrelation function for the undifferenced 1928-81 catch series (upper), and the differenced series (center); and the partial autocorrelation function for the differenced series (lower).

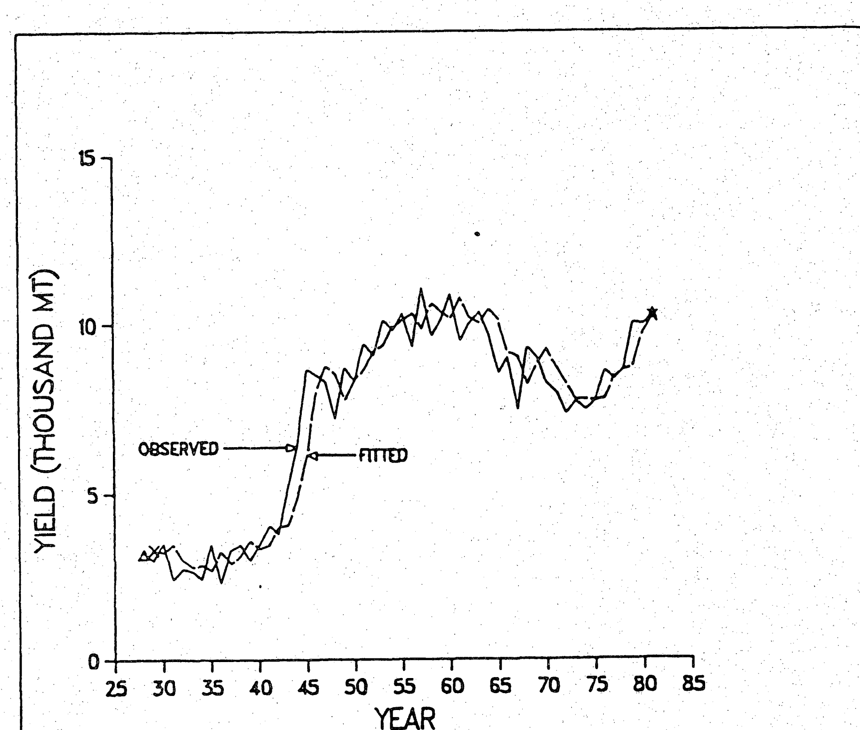

Figure 3. Comparison of the observed (______) and fitted (- - - -) lobster yield series for the period 1928-81.

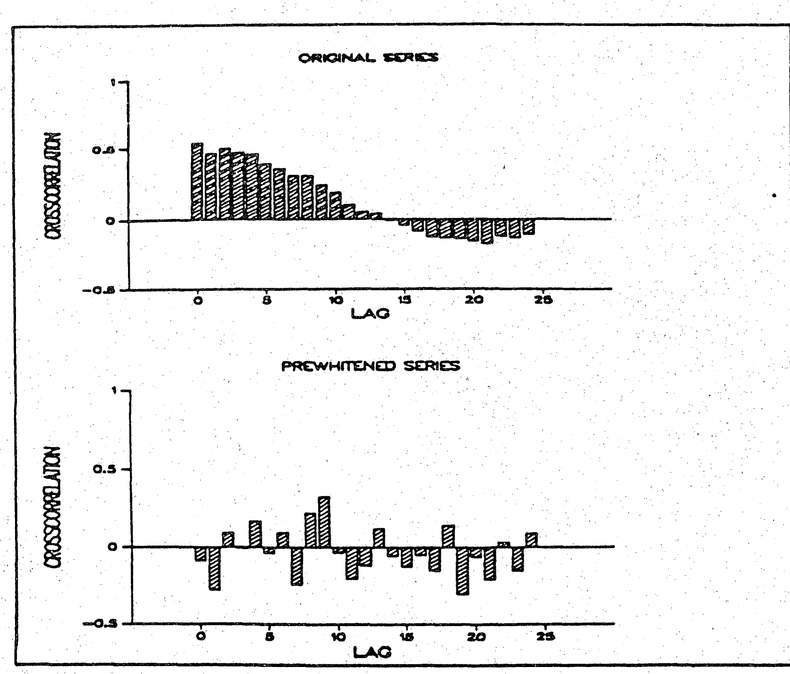

Figure 4. Crosscorrelation between the 1928-81 yield and temperature for the original (upper) and prewhitened (lower) series.

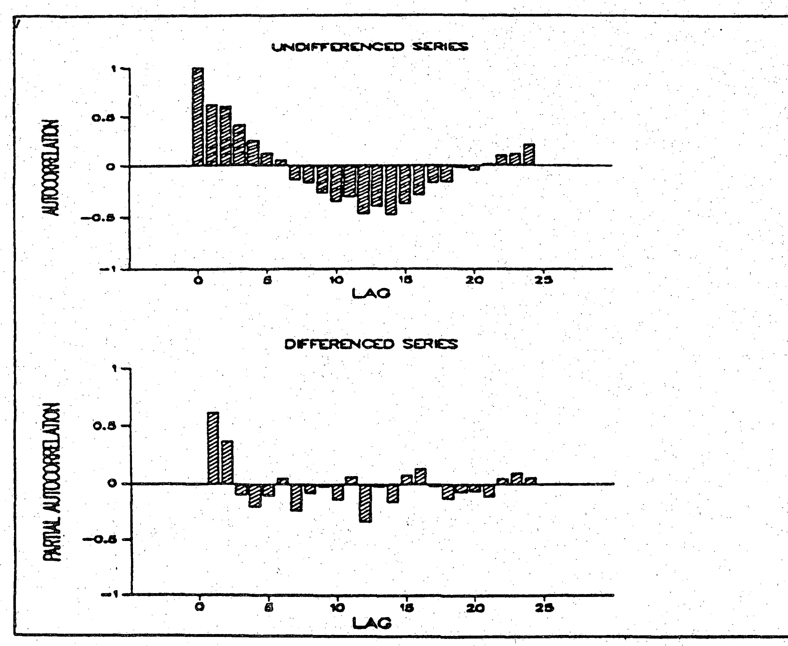

Figure 5. Autocorrelation function for the undifferenced 1945-81 catch-series (upper) and the differenced series (center), and the partial autocorrelation function for the differenced series (lower).

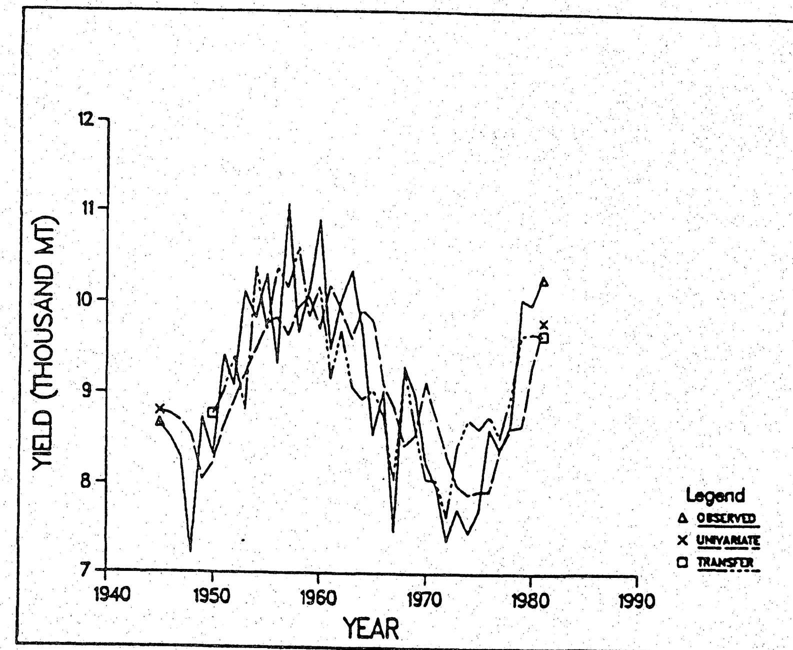

Figure 6. Comparison of the observed (———) fitted univariate (- - - -) and fitted transfer function (—— - -) series for the period 1945-81.

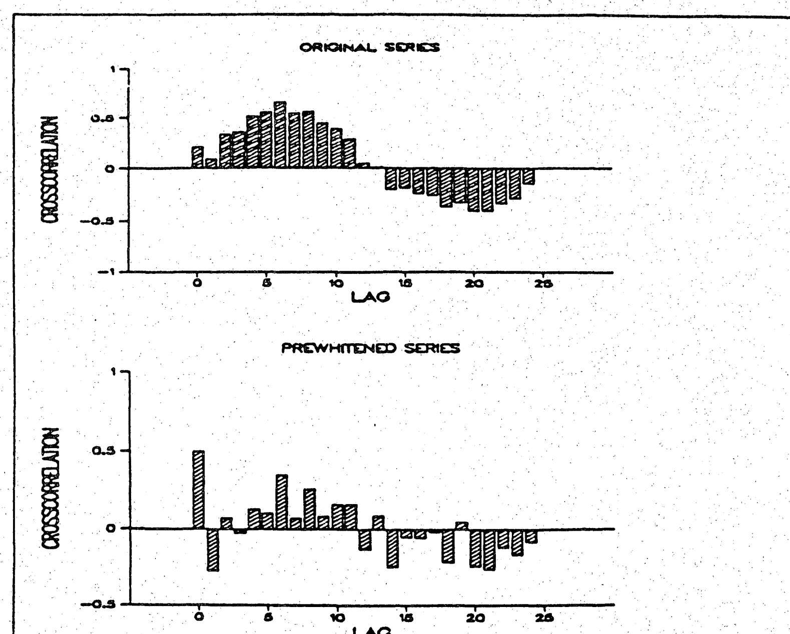

Figure 7. Crosscorrelation between the 1945-81 lobster yield and temperature for the original (upper) and prewhitened series (lower).

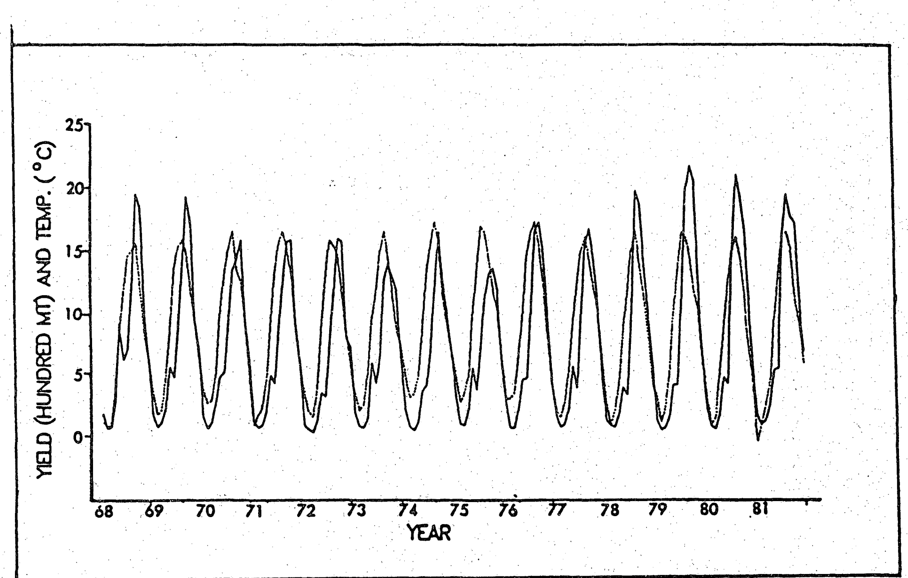

Figure 8. Monthly Maine lobster yield (mt x 10^2) and mean monthly water temperature at Boothbay Harbor for the period 1968-81.

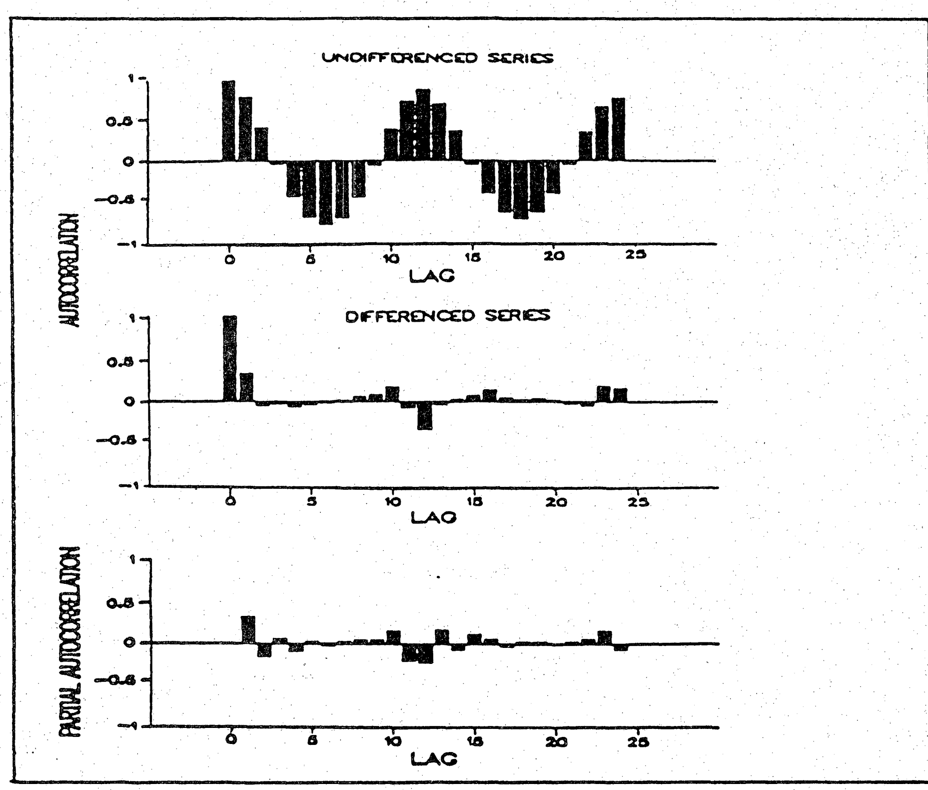

Figure 9. Autocorrelation function for the undifferenced 1968-81 monthly catch series (upper), the differenced series (center), and partial autocorrelation function for the differenced series (lower).

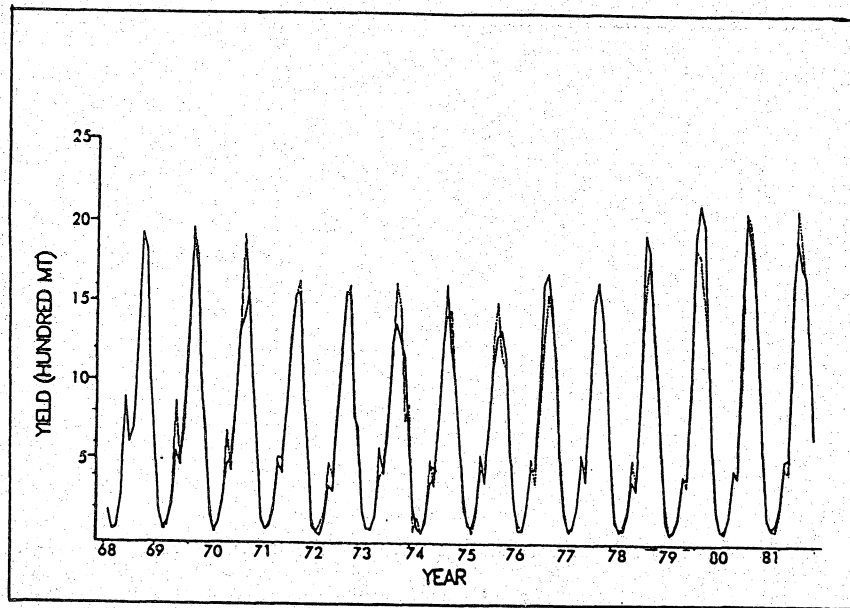

Figure 10. Comparison of observed (———) and fitted (- - - -) monthly lobster yield series for the period 1968-81.

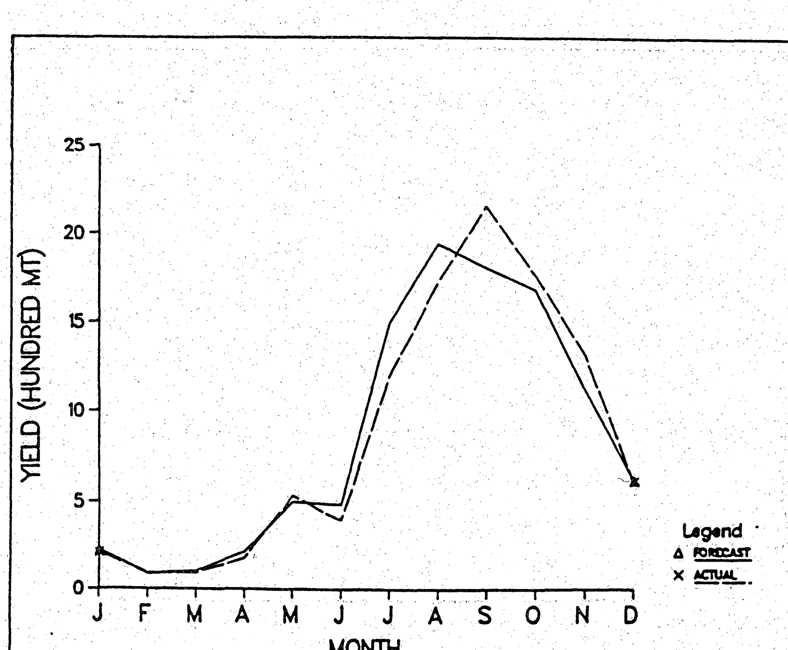

Figure 11. Comparison of actual (———) and forecasted (- - - -) 1982 monthly lobster yield.

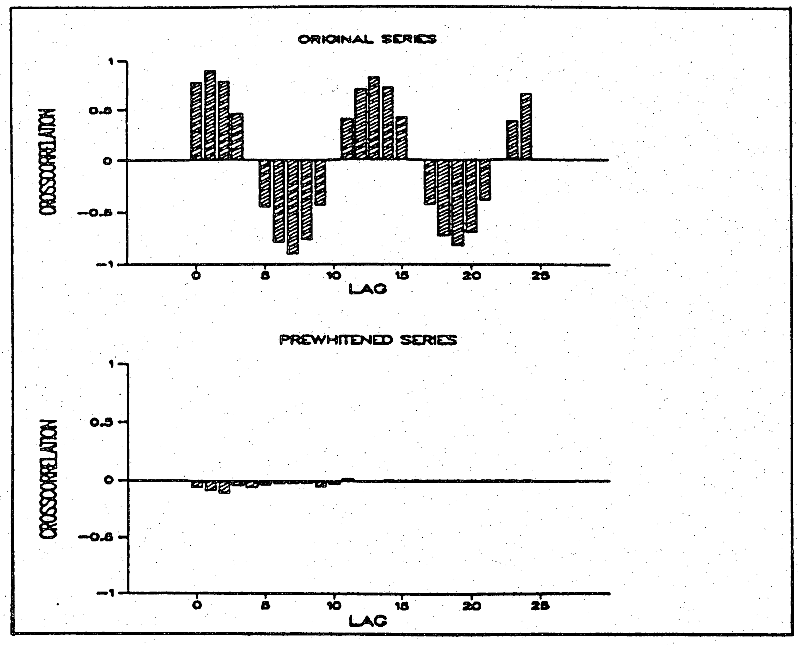

Figure 12. Crosscorrelation between monthly lobster yield and temperature for the original (upper) and prewhitened (lower) series.