Estimation of the Effect of Operational Seeding on Rain Amounts in Israel

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ABSTRACT

During the period 1961–75, two cloud seeding experiments were carried out in Israel. The first Israeli experiment had a two-target crossover design. The results indicated a positive seeding effect of 15%, significant at 0.9%. The second experiment indicated an enhancement of 13%, significant at 2.8%, in rain amounts in the northern part of Israel. Since November 1975, seeding in the north has been operational.

The objective of this paper is to develop a method for estimating the seeding effect of the operational seeding. The proposed method is based on a historical logarithmic model of target precipitation on control precipitation.

It is argued that the validity of the proposed approach stems from the stability of the relation between target and control precipitation over time, and the robustness of the meteorological system in the region.

The seeding effect in the operational period 1976–90, on the annual rainfall, is estimated by a 6% increase in rain amounts, with a 95% confidence interval of (1.01, 1.12). An analysis of the sensitivity of the effect estimate to the choice of period shows a persistent indication of a positive seeding effect. Possible explanations for the reduction in seeding effect, in comparison to the Israel-2 experiment, are discussed.

1. Introduction

During the period 1961–75, two cloud seeding experiments were carried out in Israel. The first Israeli experiment (Israel-1, 1961–67) had a two-target crossover design of north versus center. The results indicated an enhancement of 15% in rain amounts, significant at an $\alpha = 0.009$ level (Gabriel 1967, 1970). In Israel-2 (1970–75), the northern target area was shifted eastward and a northern control area was defined along the coast. Analysis of the results for the north area alone indicated an enhancement of 13% in rain amounts, significant at an $\alpha = 0.028$ level (Gagin and Neumann 1981).

Since November 1975, seeding in the northern part of Israel is operational, that is, planes flying along an upwind seeding line (Fig. 1) seed clouds, with silver iodide (AgI) dissolved in acetone, on any day satisfying the following criteria: (a) a precipitating cloud system is observed by radar, (b) cloud-top heights, measured by the same radar, are at a temperature lower than −8°C, and (c) no eastern wind component is observed at the cloud level. The location of the seeding line combined with the above criteria guarantee that the control area is kept unseeded at all times.

As the contribution of the operational seeding to rain amount was never evaluated, it is the objective of this paper to develop a method for estimating the seeding effect of the operational seeding. Such an evaluation is needed for follow-up purposes as well as for the planning of future rain-enhancement activity.

Since seeding in an operational period is not random, there is a methodological problem in providing a valid estimate of the seeding effect. The suggested approach is based on careful use of a historical model for prediction of the "natural" (without seeding) target rainfall. Such a model incorporates control area data and historical (unseeded) data. It will be argued that the validity of this approach stems from the stability of the relationship between target and control precipitation over time.

The organization of the paper is as follows. The problem and the proposed approach are discussed in section 2. Following a description of the database in section 3, the limitations of a standard linear regression model of target precipitation on control precipitation are described in the first part of section 4. As an alternative to the standard model, a logarithmic model with a shift parameter is introduced. The method of estimation of the shift parameter and analysis of the model's properties are discussed in the second part of section 4. An estimate of the seeding effect and a sensitivity analysis of the effect are presented in section 5. Finally, the validity of the historical comparison is discussed in section 6.

2. Problem and approach

The problem of valid estimation of the effect of operational seeding is well known. Whereas in a controlled seeding experiment seeding days are allocated

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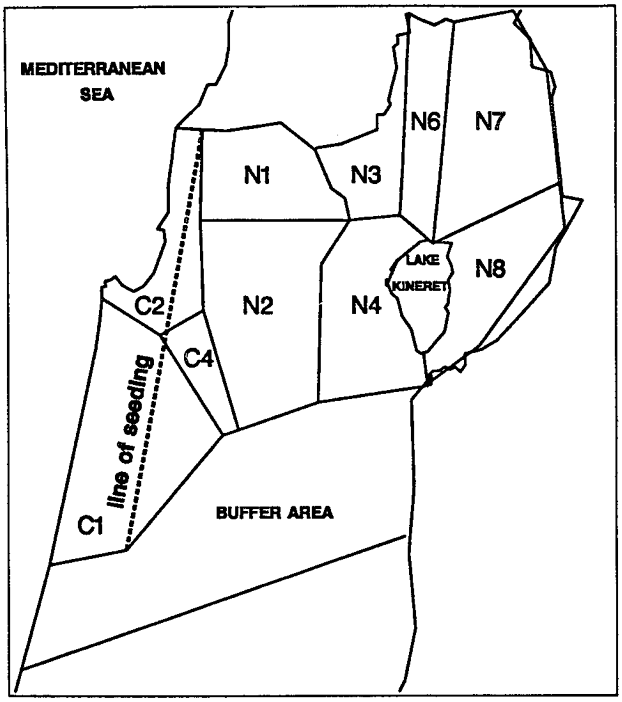

FIG. 1. Target and control areas and subareas in the operational period.

randomly, in operational seeding the randomization mechanism is absent. Randomization is the basis for the unbiased estimation of the seeding effect and for the validity of the significance level of the test of the null hypothesis of no effect against the alternative of some effect. These properties hold since under the null hypothesis the rain amounts in days allocated to be seeded and days allocated not to be seeded are statistically independent and identically distributed (i.i.d.).

Since in an operational period all rainy days are seeded, the seeding effect may be estimated by comparing seeding data with data of unseeded days in previous (historical) periods. In such a comparison, the i.i.d. properties do not necessarily hold. Examples of possible causes for a change or a trend in rainfall distribution over time are not hard to find and include industrialization and urbanization processes or climatic fluctuations.

The suggested approach is based on rainfall data in a period preceding the operational period and on rainfall in the control area. A model of target precipitation Y on control precipitation X is constructed for the unseeded days. Using control precipitation data in the operational period, the model estimates the target natural (without seeding) precipitation. This approach is often referred to as the historical regression method (Brier et al. 1967; Flueck 1976; Neyman 1977; Gabriel 1979). The seeding effect is then estimated by a double-ratio statistic based on the actual rainfall in the target area and the rainfall estimated by the model. Thus, the historical model approach is based on the stability of the *relationship* between target and control precipitation over time rather than on the stability of *rain amounts*.

The debate over the scientific benefit of such analysis is not limited to the field of weather modification, but is discussed also in the medical and social sciences literature. Attitudes span from determined opposition to cautious approval. Neyman (1977) criticized historical regression and brought examples of changes in rainfall distribution and in the relationship between areas over time. A similar attitude is expressed by Green and Byar (1984) and Dupont (1985). A more supportive attitude was presented by Gehan (1984), who claimed that the relevance of each historical comparison should be judged separately and viewed by its consistency with other projects in the same field. Begg and Pilote (1991) suggested a method for incorporation of experimental and nonexperimental "projects" in a metanalysis of a treatment effect.

Many researchers in weather modification support analysis of operational projects, while emphasizing the caution and special care that should be paid to the possible "hazards" (Tukey et al. 1978; Kruskal 1979; Flueck 1979; Braham 1979). Flueck (1976) and Hsu et al. (1981) emphasized three factors that are essential for a valid assessment of an operational project. First, the decision to start the project should be independent of the rainfall at the time of decision making. Second, the project's targets and plan should be stated in advance. Third, the definitions of the analysis' components should be uniform over time. We conclude that many researchers agree with Gabriel (1979) that "Suitable non-randomized operations should be evaluated, but such analyses must be done with great care."

In view of the above discussion, a valid historical model should satisfy the following attributes:

- The model parameters, which reflect the relationship between target and control precipitation, should be stable over time.
- The model parameters should be robust to outliers, that is, not unduly influenced by single observations.
- The model should predict rainfall in the target area as accurately as possible.

For these demands to be satisfied, the model will incorporate methods suitable for the complicated characteristics of the rainfall distribution, on the one hand, and the sensitivity of the parameter estimates to changes over time and to the influence of single observations will be examined, on the other. In addition, the stability of the relation between target and control precipitation over time will be examined through meteorological variables that determine this relationship, excluding local influences like seeding and urbanization.

3. Data

Analysis was based on data of daily rainfall amounts for the seasons 1950–90 (1950 stands for the rain season November 1949–April 1950) in subareas of the north and coast areas. These subareas were defined in the Israel-2 period (Gagin and Neumann 1981; Gabriel and Rosenfeld 1990). The data included 1204 unseeded days in the period 1950–75 and 1382 seeded days in the period 1970–90 (of which 202 days were allocated to be seeded in Israel-2). Only rainy days with at least 0.1 mm of rain in the control subareas C1 or C2 (see Fig. 1) were included.

Definitions of the analysis' components were kept uniform over time, following these lines:

1) The time unit was defined as a 24-h period beginning at 0800 LST. Data in periods with other time-unit definition were deleted (the seasons 1962 and 1963).

2) Mean rain amounts were computed from a fixed network of about 70 rain gauges (see appendix C of Gabriel and Rosenfeld 1990).

3) The target area included subareas N1–N6. Subareas N7 and N8 could not be included in a historical regression analysis since they were not part of Israel before 1967.

4) The control area was defined according to the following considerations. As the seeding line was slightly changed throughout the operational period, the "historical seeding line" was defined as the most westerly line in the period. Therefore, the control area included subareas C1 and C2. Subarea C4 was not included since it was seeded in some of the seasons. As seen in Fig. 1, the historical line crosses subareas C1 and C2. However, none of the rain gauges positioned in C1 were located to the east of the seeding line. Two of the rain gauges positioned in C2 were located about 3 km to the east of the line, but the effect of seeding at such proximity should be negligible. Moreover, if these gauges had been affected by the seeding, the absolute value of the effect, measured by DRG in (5), would have been reduced.

In addition, radiosonde measurements at eight atmospheric levels were available from January 1957. In the period 1957–74, the radiosonde was released daily from Bet-Dagan at 1300 LST, and in the following years the frequency was increased to two and then four measurements per day. The measurements include dewpoint temperature, wind direction (WD), and wind speed (WS in knots). From these measurements the amount of precipitable water from the ground to the 600-mb level (PW) and the westerly component of the wind at the 850-mb level (U850) were computed. The orographic component of the rain (WLIFT) was defined as the product of the wind speed at the 850-mb level (WS850) and PW. WLIFT and U850 can be associated with two different meteorological factors, strongly affecting the ratio between target and control precipitation in the following ways:

1) Rain clouds in Israel are formed over sea mostly as a result of the interaction between cold air and the relatively warm seawater. The clouds decay as they move inland. Stronger westerly wind will drive the rain clouds farther inland before decaying, thus increasing the ratio between target and control precipitation at stronger U850. The U850 is highly correlated with the cloud movement.

2) The target area is mostly hills, reaching a maximum height of 1200 m above sea level, while the control area lies mostly in the coastal plain. Alpert and Shafir (1989) have shown that the orographic component of the rainfall can be almost fully explained as the product of WLIFT with the sine of the hill's slope.

Both processes show that WLIFT and U850 are major meteorological factors determining the relationship between target and control precipitation, and are not influenced by local factors. Therefore, any historical changes in the target–control ratio that may be due to climatic fluctuations are likely to be revealed by historical changes in these factors.

4. Fitting the model

a. Standard linear regression

The simplest model for description of the relationship between target Y and control X precipitation is the standard linear regression model, satisfying

$$y_i = \gamma + \delta x_i + \theta_i, \quad i = 1, \cdots, n, \qquad (1)$$

with

$$E(\theta_i) = 0, \ \operatorname{var}(\theta_i) = \eta^2, \ \operatorname{cov}(\theta_i, \theta_j) = 0 \ \text{if} \ i \neq j,$$

where y_i, x_i are the daily rainfall in target and control areas, respectively, γ is the intercept, δ the slope, and θ_i the uncorrelated identically distributed error terms with variance η^2. Normality of θ_i is frequently assumed for applicability of tests based on the F and t distributions.

As the model's stability over time is of interest, unseeded days in the period preceding the operational period were divided into four subgroups, each comprising five to six seasons (of which the last two correspond to Israel-1 and Israel-2). The regression coefficients and residual statistics are presented in Table 1, and the joint distribution of rainfall in the target and control areas are presented in Fig. 2. Note that in order to keep the distinction between parameters and their estimates clear, the estimates are denoted hereinafter by putting a "hat" over the corresponding letter. The main features of Table 1 and Fig. 2 are as follows:

1) The rain distribution varies over time. This is most evident in variable frequencies of high rain amounts.

TABLE 1. Coefficients of the regression model of target precipitation on control precipitation and moments of the residuals in four periods. The analysis is based on unseeded days.

Period	n	Intercept (std err)	Slope (std err)	R^2	Residuals Variance	Skewness	Kurtosis
1950–55	469	0.54 (0.21)	0.90 (0.02)	0.85	13.09	−0.02	10.73
1956–60	373	0.23 (0.24)	0.89 (0.02)	0.84	12.66	0.65	5.02
1961–67*	168	0.35 (0.30)	0.87 (0.03)	0.87	8.79	−0.23	4.71
1970–75**	202	0.78 (0.42)	0.84 (0.04)	0.74	20.56	0.83	5.91

* Israel-1.
** Israel-2.

2) The slope of the regression line δ and the percent of explained variance R^2 are influenced by high rainfall values. In consequence, the estimates of δ changes from $\hat{\delta} = 0.84$ to $\hat{\delta} = 0.90$, and R^2 changes from $R^2 = 0.74$ to $R^2 = 0.87$. Fluctuations in the estimate of η^2 are even more extreme.

Examination of the model assumptions through analysis of the residuals show that

3) The absolute value of the residuals are seen to increase with the predicted rain amounts in the target $\hat{y}$ and with control precipitation, indicating higher residual variance for higher precipitation.

4) There is some indication of an asymmetrical residual distribution. For example, analyzing the residuals of the regression line fitted for all unseeded days, it was observed that for $30 \leqslant X < 40$, 57% of the residuals were positive, with mean 0.41 standard errors; whereas for $40 \leqslant X$, only 27% of the residuals were positive with mean −0.87 standard errors.

5) The residual fourth moment (kurtosis) is high. (The expected value for the standard normal distribution is zero.)

The third and fourth features suggest that the assumptions of homogeneity of the residual variances and of linearity of the relationship between Y and X are not satisfied, at least in some of the periods. These assumptions are vital for the adequacy of the model. The fifth feature is connected to the normality assumption and implies that tests and confidence intervals based on this assumption are not valid.

It is concluded that the stability and accuracy of the standard regression model may not be satisfactory when a historical model is considered.

b. The logarithmic model

A common remedy for the problem heterogeneity of the residuals' variance is a logarithmic transformation of Y and X. Logarithmic models are quite common in the atmospheric sciences, in general (see Crow and Shimizu 1988, chapter 13), and in evaluations of rain enhancement activity, in particular (Bradley et al. 1980; Weisberg 1980, p. 152).

To prevent the logarithmic function from going to infinity for zero rainfall, a shift parameter is added to the model. The proposed model is given by

$$\ln(y_i + \mu) = \alpha + \beta \ln(x_i + \mu) + \epsilon_i, \qquad (2)$$

with

$$E(\epsilon_i) = 0, \ \mathrm{var}(\epsilon_i) = \sigma^2, \ \mathrm{cov}(\epsilon_i, \epsilon_j) = 0 \ \text{if} \ i \neq j,$$

where y_i, x_i are the same as in (1), α is the intercept, β the slope, μ the shift, and ϵ_i the uncorrelated identically distributed error terms with variance σ^2.

1) ESTIMATION OF THE SHIFT PARAMETER

The problem of estimating the shift parameter in (2) is discussed in the literature. Hill (1963) showed that the global maximum likelihood (ML) estimate of μ is $-\min(y_i)$ (which equals zero for the rain data), and that the ML estimates of the other parameters in this environment are infinite. However, in many situations a local maximum likelihood (LML) estimate exists and maintains the optimal properties of efficiency and consistency of the ML estimate (Cohen 1951; Harter and Moore 1966; Calitz 1973).

The existence of a LML estimate of the shift parameter for the rain data was examined graphically by plotting the residual profile log-likelihood [denoted by $L^*(\mu)$] against μ. The residual profile log-likelihood is extracted by substitution the ML estimates of $\hat{\alpha}$, $\hat{\beta}$, and $\hat{\sigma}$ in the residual log-likelihood function. Assuming normality of the residuals, it is given by

$$L^*(\mu) = -n \ln(\hat{\sigma}) - \sum_{i=1}^{n} \ln(y_i + \mu), \qquad (3)$$

where n is the number of observations. An example of $L^*(\mu)$ for the rain data is shown in Fig. 3. The different curves represent five subperiods of the period 1950–61 (the first 1, 3, 6, 9, and 12 seasons of this period). It is seen that the position of the local maximum is stable (consistency of the LML estimate) for the different subperiods, and becomes more pronounced as the number of seasons increases (efficiency of the LML estimate). It is also evident that it is preferable to analyze periods longer than six seasons, since the shift parameter may be unidentifiable for short periods.

After the existence of the LML estimate is verified, the estimated value is found graphically by focusing on a neighborhood of the local maxima. The accuracy of this estimate is usually sufficient in practice. The graphic method of estimation also enables a simple construction of an approximate confidence interval for μ given by

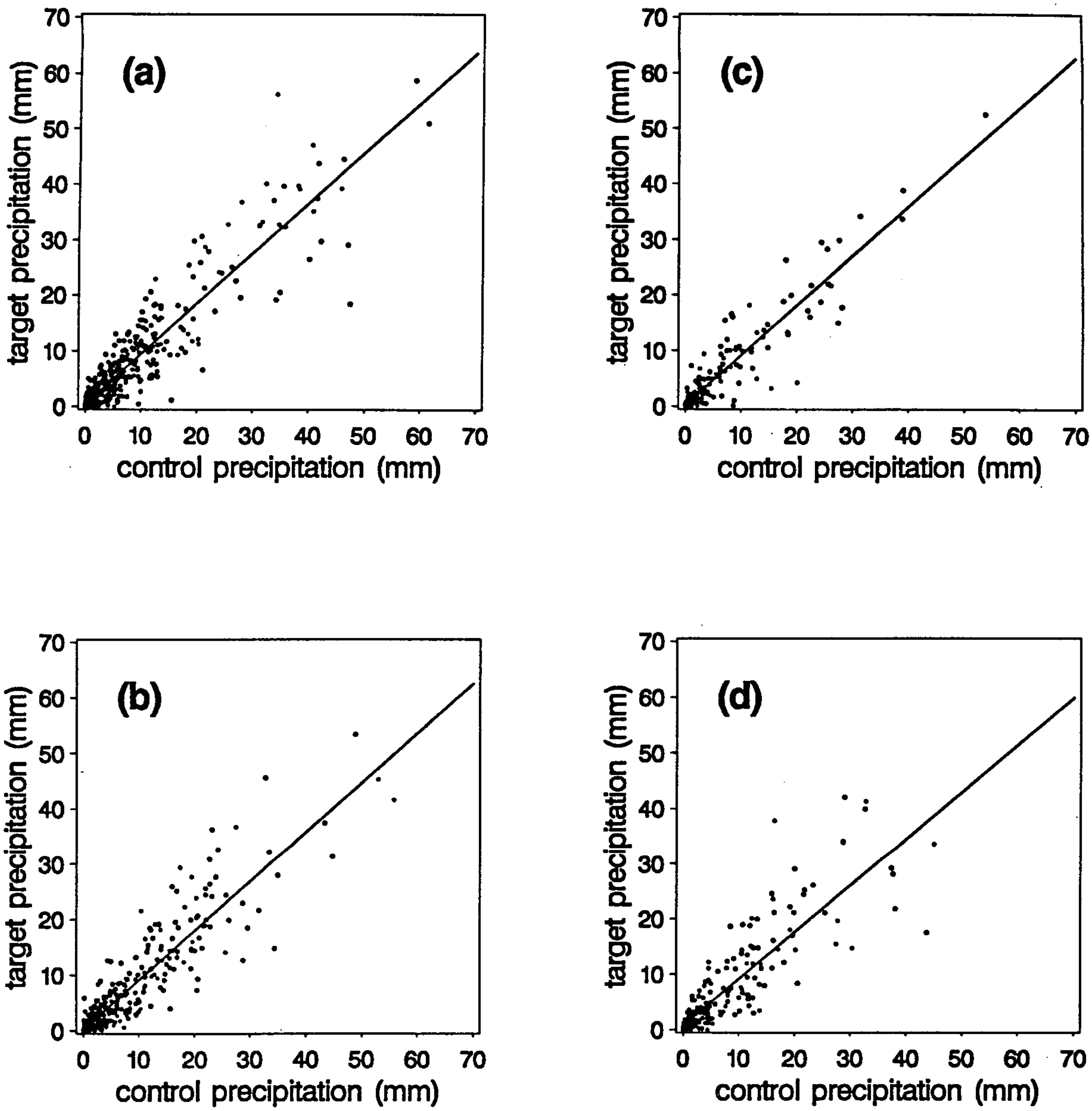

FIG. 2. Distribution of target and control precipitation and regression line in four periods. The analysis is based on unseeded days: (a) 1950–55, n = 469, (b) 1956–60, n = 373, (c) 1961–67, n = 168, (d) 1970–75, n = 202.

$$L^*(\hat{\mu}) - L^*(\mu) \leqslant \chi^2_{(\alpha,1)}, \quad (4)$$

where $\chi^2_{(\alpha,1)}$ is the α percentile of the χ^2 distribution with one degree of freedom (Box and Cox 1964). Thus, for the rain data in the period 1950–75, the shift parameter is estimated by $\hat{\mu} = 0.28$ mm, with a 95% confidence interval of (0.21, 0.37).

2) EXAMINATION OF THE MODEL

For the estimated shift of $\hat{\mu} = 0.28$, the least squares estimates of the other parameters in (2), based on data of 1204 days in the period 1950–75, are $\hat{\alpha} = 0.047$, $\hat{\beta} = 0.899$, and $\hat{\sigma}^2 = 0.377$. The percent of variance of target precipitation accounted for by the model is $R^2 = 0.87$.

As it is desirable that the historical model will not be unduly influenced by single observations, observations with a studentized residual of absolute value greater than 3 were excluded from the calculation of the above estimates. This group of outliers consists of less than 1% of the observations (11 observations out of 1204) and accounts for less than 1% of the total rainfall amount. An interesting feature of the outliers is that since they are outliers in a logarithmic scale, they are characterized by small to medium rain amounts in one area, and close to zero rainfall

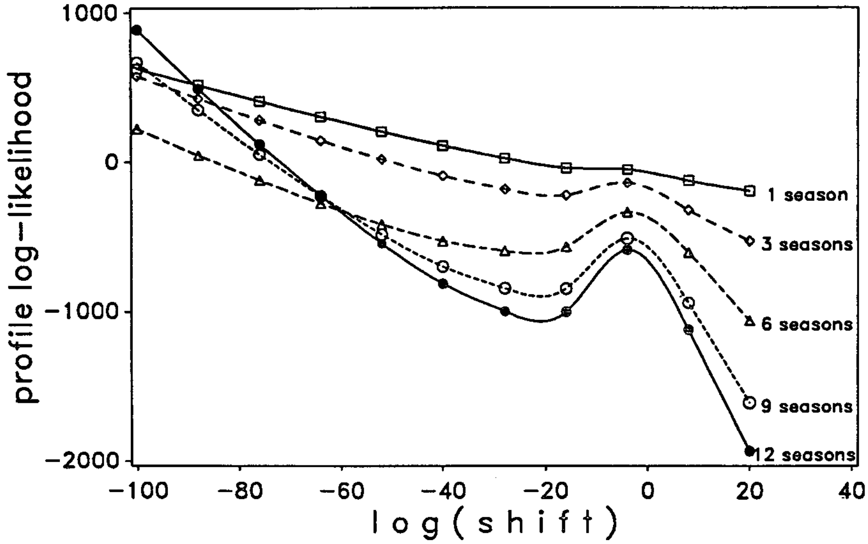

FIG. 3. Residual profile log-likelihood against the logarithm of the shift parameter, for the first 1, 3, 6, 9, and 12 seasons in the period 1950–61 (879 unseeded days).

amount in the other area. In comparison, the group of outliers satisfying the same criterion in the standard regression model of the previous section consists of 3% of the observations and accounts for 7% of the total rainfall amount (see example in Fig. 4).

The adequacy of the model, the stability of the estimates, and the sensitivity of the parameters to the shift estimate were examined for three 10-season periods: 1950–59, 1956–67, and 1964–75. The main results are

1) The parameter estimates are stable over time and differ by one standard error or less from each other (Table 2).

2) Diagnostic analysis of the residuals show a good fit to a linear model and homogeneity of the residual variance (see example in Fig. 5a). It is seen in the example presented in Fig. 5b that when the shift is too small, the residuals are no longer distributed randomly around zero for all values of $\hat{y}$.

3) The model is not sensitive to small fluctuations, of up to 0.2 mm around the shift estimate (Table 3), but larger departures cause the assumption of linearity to fail (see example in Fig. 5b).

4) The residual distribution is found to fit a normal distribution according to several standard tests (Table 2), although a slight negative skewness of the distribution is apparent.

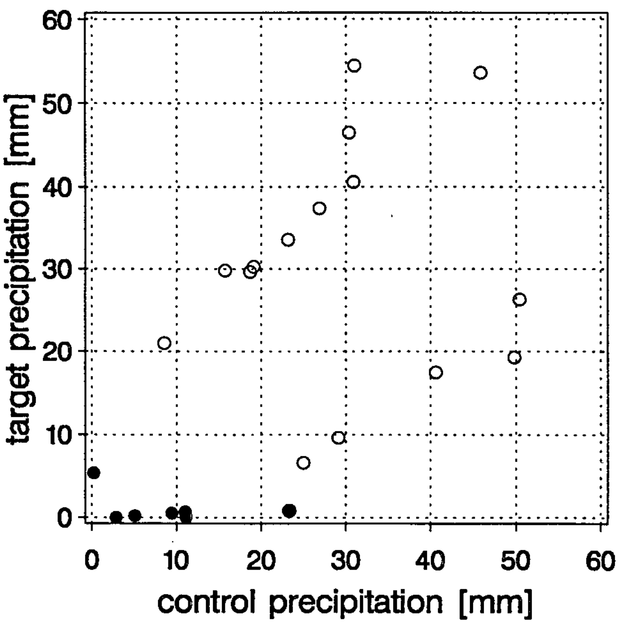

FIG. 4. Rainfall on days with studentized residuals with absolute value greater than 3, for the standard regression model of original rainfall values (circles) and for the logarithmic model (full dots). Based on the period 1950–59 (783 unseeded days).

5. Estimation of the seeding effect

The statistic used for the estimation of the seeding effect in Israel-2 north was the double-ratio statistic DR = $(\bar{Y}_s/\bar{Y}_u)(\bar{X}_s/\bar{X}_u)^{-1}$, where $\bar{Y}_s$, $\bar{Y}_u$ are the mean precipitation in the target area on seeded and unseeded days, respectively, and $\bar{X}_s$, $\bar{X}_u$ are defined similarly for the control area. A generalization of this estimator, where $\bar{X}$ is replaced by the mean of the predicted natural rainfall $\bar{\hat{Y}}$ in the target area, was suggested by Gabriel and Rosenfeld (1990). Predicted natural rainfall $\hat{Y}$ is computed from a regression model of target precipitation on control precipitation. This estimator has

TABLE 2. Coefficients of the logarithmic model and the moments of the residuals in three 10-season periods. The shift parameter is estimated by 0.28 mm. Outliers are defined as observations with studentized residual with absolute value greater than 3. The analysis is based on unseeded days.

Period	n	No. of outliers	$\hat{\alpha}$ (std err)	$\hat{\beta}$ (std err)	R^2	Residuals $\hat{\sigma}^2$	Skewness	Kurtosis	S–W*
1950–59	777	6	0.04 (0.03)	0.90 (0.02)	0.78	0.44	−0.18	0.24	0.19
1956–67	530	3	−0.03 (0.04)	0.91 (0.02)	0.78	0.44	−0.16	0.49	0.66
1964–75	330	3	0.03 (0.05)	0.89 (0.02)	0.80	0.38	−0.32	0.33	0.20

* P value of the Shapiro–Wilk's test of normality. Values greater than 0.05 imply that the normality hypothesis is accepted at the 0.05 level.

a lower variance than the DR and will be called in the context of operational seeding DRG:

$$\mathrm{DRG} = (\bar{Y}_s/\bar{\hat{Y}}_s)(\bar{Y}_u/\bar{\hat{Y}}_u)^{-1}, \quad (5)$$

where $\bar{Y}_s$ is the mean rainfall in the target area on seeded days (in the operational period 1976–90), $\bar{Y}_u$ is the mean rainfall in the target area on unseeded days (in the historical period 1950–75); $\bar{\hat{Y}}_s$ and $\bar{\hat{Y}}_u$ are defined similarly for the predicted rainfall in the target area, whose estimation is discussed hereinafter.

Note that because there is an interest in estimation of seeding effect in actual rain amounts, $\hat{Y}$ is expressed in original rainfall units of millimeters per day. An unbiased estimate for the retransformation of Y from logarithmic scale to original values was suggested by Bradu and Mundlak (1970). A simplification of their estimate of $\hat{y}_i$, for large samples, is given by

$$\hat{y}_i = \exp(\hat{\alpha} + 0.5\,\hat{\sigma}^2)(x_i + \hat{\mu})^{\hat{\beta}} - \hat{\mu}. \quad (6)$$

It should be noted that since empirically R_u will probably deviate from one, R_u represents a "calibration factor" of the model correcting for empirical biases. In fact, $R_u = 0.96$ for all unseeded days.

The estimated seeding effect in the operational period 1976–90 is DRG = 1.06, with a 95% confidence interval given by (1.01, 1.12). The respective estimate for Israel-2 north, when seeded days in the period 1970–75 are compared to unseeded days in the period 1950–75, is DRG = 1.09. The estimate of the effect using both seeded and unseeded days of the experiment (1970–75) is DRG = 1.13. This result is identical to the one presented by Gagin and Neumann (1981) in the formal analysis of the experiment.

The sensitivity of the estimate to the choice of period is demonstrated in a plot of the estimated ratios $R = \bar{Y}/\bar{\hat{Y}}$ for moving 10-season periods (Fig. 6). It is seen that the values of R in all unseeded periods R_u is below 1, and in all seeded periods R_s above 1. The range of the values of DRG corresponding to all combinations R_s/R_u is between 1.01 and 1.14. As the double ratio may be highly influenced by a few observations, the sensitivity of the estimates to deletion of observations with high rain amounts was studied. Deletion of three observations with rain in target or control, in either seeded or unseeded days, greater than 60 mm influenced the fourth digital place of the effect estimate. These results consistently indicate a positive seeding effect in the operational period, regardless of the choice of periods.

6. Validity of the historical comparison

As mentioned before, the validity of the historical model approach is mainly dependent on the stability of the relationship between target and control precipitation over time. The time dependence of this relationship will be examined through the estimates of the logarithmic model parameters (α, β, and σ^2) and the medians of WLIFT and U850. The model parameters reflect target–control relationship in the historical period (1950–75), while the meteorological variables reflect this relationship during both historical and operational periods (1958–90).

Two types of time series were inspected: annual estimates of the above properties and 10-season moving estimates (TSME). Annual estimates can be assumed to be independent of each other and can therefore be examined by regression over time. The TSME are formed by taking first the estimate based on the first 10-season period, then the next 10-season period, beginning a season later, and so forth. Inspection of the TSME series enables an insight into a variety of "time windows."

Table 4 summarizes the results of a regression analysis of the annual estimates on time. It is seen that the slopes of all the regression lines are slightly negative and are not significant at levels of at least $\alpha = 0.10$. Less than 9% of the variability of the above properties is accounted for by the time factor.

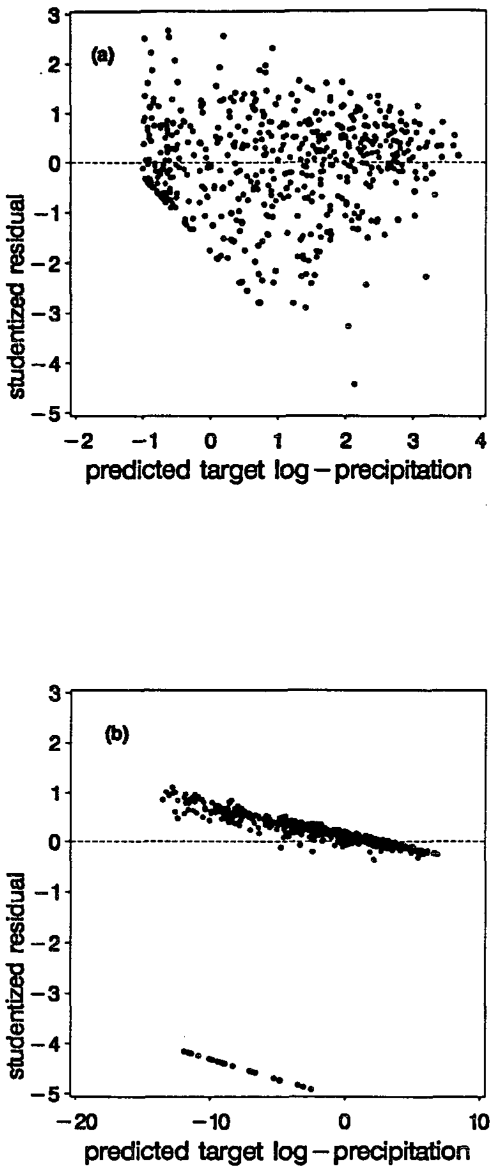

FIG. 5. Studentized residuals of the logarithmic model versus the predicted target log-precipitation. Based on the period 1956–67 (533 unseeded days): (a) $\mu = 0.3$ mm, (b) $\mu = 10^{-27}$ mm.

Examples of two TSME series are presented in Fig. 7: for the estimates of β and for the median of WLIFT. The estimates of β vary from 0.89 to 0.91, with estimated standard errors between 0.018 and 0.026. Thus, every comparison of the $\hat{\beta}$ values in two nonoverlapping periods is not significant at the 0.05 level. The medians of WLIFT range from 254.8 (in the period 1970–79) to 284.0.

The temporal trends examined above were found statistically nonsignificant at the 0.05 level. However, for the variables examined, the power of detection of a trend, when it actually exists, is only 30%–40%. Therefore, the implications of a time trend on the results should be discussed. It is seen in Table 4 that all model parameters (α, β, and σ^2) decrease slightly in time, thus contributing together to a decrease in the predicted target rainfall in the operational period [see (6)]. Furthermore, the indicated temporal changes in U850 and WLIFT provide a physical explanation to a decrease in target rainfall as compared to the control.

Therefore, seeing that $\bar{Y}$ is in the denominator of DRG, the effect cited above is, if anything, an underestimate of the "true" effect. Computation shows that if the indicated trend in the period 1950–75 is assumed to continue at the same rate in the operational period, then $\hat{\beta}$ would decrease from 0.90 to 0.88, and the effect estimate would be 11% instead of 6%.

TABLE 3. Sensitivity of the parameters of the logarithmic model to small deviations in the shift estimate. The analysis is based on 1204 unseeded days.

Shift	$\hat{\alpha}$ (std err)	$\hat{\beta}$ (std err)	R^2	$\hat{\sigma}^2$
0.1	−0.02 (0.03)	0.89 (0.02)	0.73	0.73
0.2	0.02 (0.03)	0.89 (0.01)	0.75	0.56
0.3	0.04 (0.03)	0.89 (0.01)	0.76	0.47
0.4	0.06 (0.03)	0.88 (0.01)	0.77	0.41
0.5	0.07 (0.02)	0.88 (0.01)	0.77	0.36

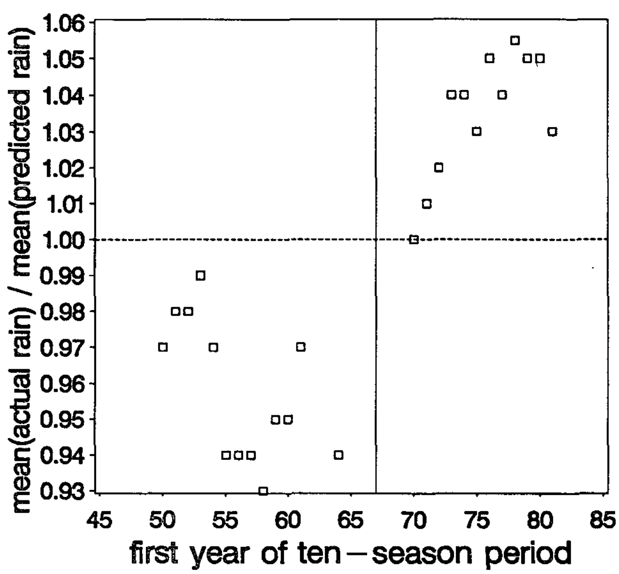

FIG. 6. Ratio of the mean precipitation in the target area to the mean natural precipitation predicted by the logarithmic model. Each square represents the ratio for the 10 seasons which begin in the year appearing in the abscissa.

In addition, if a temporal trend exists, its influence on the effect estimate would be reduced if the time gap between the historical and operational periods is shortened. For instance, if the first operational decade is compared to the first historical decade, or the last decades of each period are compared, then both estimates indicate an 8% enhancement in rain amounts. A comparison of the adjacent first operational decade and last historical decade indicates an increase of 11% in rain amounts.

7. Summary and conclusions

The decision to start the operational seeding in Israel was not influenced by weather conditions but followed the second Israeli experiment. The method of seeding did not change throughout the project, and the definitions of target and control areas and of the time unit were uniform over time. Thus, the preliminary requirements for a valid evaluation of the effect of seeding of an operational project are fulfilled.

The evaluation of the effect is based on a historical logarithmic model that uses historic data of unseeded days and control area data of seeded days. The indicated seeding effect in the period 1976–90 is of a 6% increase in rain amounts, significant at a 2% level.

A fallacy of the historical model may be caused by a time trend in the relationship between precipitation in the target and control areas. Therefore, this relationship was examined through time series of the model parameters and of relevant meteorological variables. No significant time trend was found for all the series that were examined.

All estimates of the effect, calculated for all comparisons of 10 seeded and 10 unseeded seasons, indicated a positive seeding effect. Furthermore, consideration of a historical trend of target–control relationship showed, that if anything, the estimated effect is an underestimate of the true effect by 2%–5%, indicating a seeding effect of 8%–11%.

This indicated effect is lower than the 13% effect found in the Israeli experiments. A possible (statistical) explanation is that both estimates are within the variability of the statistic. Another possible explanation is that the lower effect is related to the 42% reduction in the mean seeding time in the operational period compared to the Israel-2 period. The mean daily seeding time was defined as the flight time (in minutes) per kilometer of seeding line.

The methodology introduced in this work extends beyond the immediate application to the estimation of the seeding effect to domains such as hydrology, ecology, and climatology. Such an extension seems natural because of the similarity between distributional properties of rainfall and some of the variables in those areas. It is also expected that the methodology presented here can be used for detection of a temporal change in climatic and ecological variables due to local factors such as urbanization and pollution.

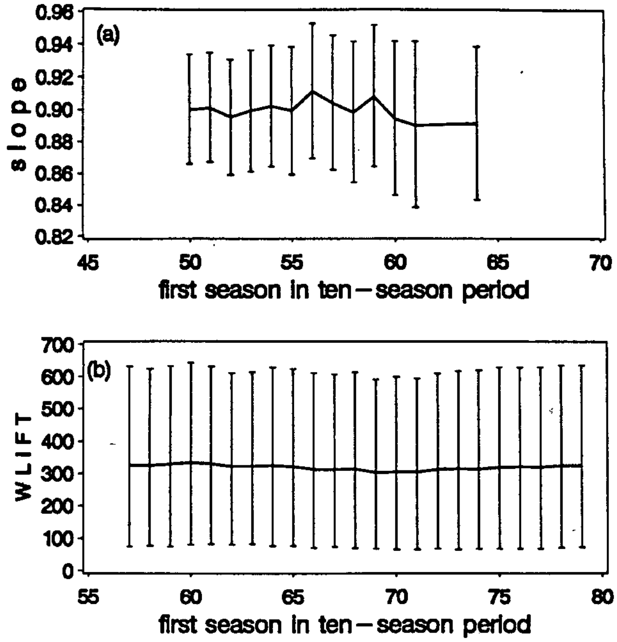

FIG. 7. Ten-season moving estimates. (a) The slope of the logarithmic model, the whiskers extend to a distance of two standard errors. (b) The median of WLIFT, the whiskers extend to a distance of the 5th and 95th percentiles.

TABLE 4. Coefficients of the regression over time of annual parameters of the logarithmic model (22 seasons) and of medians of meteorological variables (32 seasons): U850 is the speed of western component of wind in 850-mb level, WLIFT is the orographic component of rain amounts.

Estimator	Intercept	Slope (p value)	R^2
α	0.0558	−0.0009 (0.85)	0.022
β	0.9027	−0.0006 (0.66)	0.001
σ^2	0.4449	−0.0033 (0.26)	0.063
U850	13.3908	−0.0500 (0.10)	0.089
WLIFT	279.8516	−0.5564 (0.30)	0.036

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