ISSN 1032-3813
ISBN 0 7326 1025 7

MONASH UNIVERSITY

AUSTRALIA

STRIKE DATA WITH A CRISIS POINT

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Working Paper 2/97
March 1997

DEPARTMENT OF ECONOMETRICS
AND BUSINESS STATISTICS

Strike Data with a Crisis Point

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February 1997

Abstract

The paper is concerned with the analysis of strike data in which the distribution of short strikes differs from that of long strikes. It appears through visual inspection and asymptotic procedures that for Israeli strikes in the years 1965–1992, the hazard function is exponential for strikes lasting less than 40 days and that it is Weibull with a Weibull parameter greater than unity for longer strikes. The economic interpretation of the phenomenon is discussed. As there is typically only a small sample of long strikes available (Kiefer 1988), classical asymptotic tests are unlikely to convey the correct message. We suggest a test statistic for the hypothesis that a break does not occur. A new F-based expansion for the small sample distribution of the test is derived. The test rejects the hypothesis of no break for the data under investigation. While conventional approximations to the distribution of the test statistic are shown to break down catastrophically, the F-expansion appears to be highly accurate.

Keywords: F-expansion, Gumbel distribution, Pseudovariate, Strike Data.

JEL: C10, C41, J52.

*The paper was written while the author was visiting the Department of Econometrics, Monash University.

1 Introduction

The correct specification of the distribution of duration of strikes has been debated for at least three decades. It appears that the only agreement researches in this area were able to reach so far, is on the degree of its importance. It is well recognized that under misspecification of the hazard function, standard tests for neglected heterogeneity and functional forms are invalid (Jaggia 1991).

Horvath (1968) analysed strike durations in the U.S. prior to 1961 and found the Weibull distribution to be suitable. For strikes in Britain, Lancaster (1972) established that the Inverse Gaussian fits the data most. Morrison and Schmittlein (1980) reexamined Horvath and Lancaster's data set and argued that a Gamma-Weibull mixture better describes the duration-generating process. Kennan (1980) extended Horvath's (1968) model to include strikes from 1953–1974 and perceived differences between contract and noncontract strikes. Newbey and Winterton (1983) argued that the distribution of durations of Lancaster's (1972) data is not the same for official and nonofficial strikes. Lawrence (1984) preferred the lognormal specification for the previously analysed Weibull and Inverse Gaussian for the British data, reasoning that it accounts better for short strikes and that it offers a greater economy of parameters. In the aftermath, Newbey (1985) commented on Lancaster's (1972), Newbey and Winterton's (1983) and Lawrence's (1984) work ; his main criticism was that the inverse normal and the lognormal are heavily tailed distributions, and these are more affected by grouping and the truncation of the data.

The debate on strike duration specification continued with Kennan's (1985) paper. Kennan proposed the flexible beta logit model with a U-shaped hazard. He found that strike durations are countercyclical. The importance of his work is in the introduction of heterogeneity in which the hazard is a polynomial of strike age up to a random individual effect. Keenan's data has been consequently analysed by Kiefer (1988), Jaggia (1991) and DeJong (1993). Jaggia (1991) demonstrated that if the hazard is misspecified as exponential instead of Weibull, standard tests will fail to detect heterogeneity or misspecification of the functional form.

The literature is endowed with other characterizations of the hazard. In addition to the inverse normal, lognormal, exponential, generalized gamma, and Weibull hazards listed above, choices of F and normal distributions have been attempted as well. For further accounts, the reader is referred to Kalbfleisch and Prentice (1980), Cox and Oakes (1984), and Lancaster (1990).

Without exception, all existing specification tests are asymptotic in nature. These are either comprised of parametric likelihood-based procedures with known asymptotic chi-square distributions, or the array of nonparametric graphical techniques. See for instance, Diagostino and Stephens (1986) and Gan and Koehler (1990). A popular test uses generalized residuals, a useful concept developed by Cox and Snell (1968). The method essentially compares the empirical distribution of the integrated hazard evaluated at estimates, or an elementary transformation of it, to the theoretical distribution. This tool was utilized in constructing conditional moment tests (Jaggia 1991). Generalized residuals were also employed in testing for random heterogeneity (Lancaster 1985). Kiefer (1985) suggested an LM test for the null of exponentially distributed durations against a class of alternatives which can be approximated by an expansion in Laguerre polynomials. The test amounts to zero restrictions on the coefficients of these polynomials. The reliability of all of the aforementioned tests when only a small sample of durations is available is questionable and it is data dependent. It appears that specification tests with known small sample properties do not exist for this analysis.

The present paper is motivated by the ongoing debate on which hazard function fits best strike data. The literature accounted for hitherto is primarily concerned with strikes in the U.K. and in the U.S. It is unclear whether differences in economic environments result in different hazards or not. It is well known that strikes in the U.S. are, on average, narrow[1] and long, whereas in France, they are generally wide and short. Moreover, French unions often attempt to gain a political reward from a strike, whereas in the U.S., the chief consideration is an economic one. These economies appear to be on the extremes of the spectrum defined upon width, length and motivation for strikes. Most other countries have strikes with characteristics somewhere in between. These are, of course, key factors in generating strike durations. The historical findings for U.S. and British strikes were not supported by other countries in the main stream literature.

In this study, we suspect through graphical observations that short Israeli strike durations between 1965–1992 are exponentially distributed, whereas long strikes in this period are Weibully distributed with a Weibull parameter greater than unity. It appears that around the 40-day mark, a crisis point is reached and an acceleration towards a solution takes place. A possible explanation for this phenomenon is that when a crisis point is reached, greater efforts for settlements and compromise are made so that the rate at which disputes end increases.

We propose in this paper a formal test procedure for the detection of a crisis point. The exact distribution of the test is unknown and it is shown that Edgeworth and saddlepoint approximations for its distribution, now very popular in the literature (Barndorff-Nielsen and Cox 1979), break down. The contribution to distribution theory that the paper makes is in the provision of a new expansion to the distribution of the test with a much more natural F-based leading term. The performance of the new expansion is excellent in small samples.

The plan for the paper is as follows. The data, model and preliminary, indicative

[1]The width of the strike is defined as the number of strikers participating.

results for a break in the hazard are described in Section 2. A formal test for the detection of a change in the hazard function is suggested. In Section 3, we develop the distribution theory for the test statistic. Relations to conventional approximations are drawn. The F-expansion is applied to the strike data in Section 4. It confirms that at the crisis point the hazard transforms from exponential to Weibull, reflecting an acceleration towards settlements. The performance of the various approximations is compared using the data set. The new approximation is found best by far. Conclusions are drawn in Section 5.

2 Data and Preliminaries

2.1 Data

The empirical hazard and integrated hazard functions for 1106 completed Israeli strike durations over the period 1965–1992 are given in Figures 1 and 2. The method of construction of these functions parallels the detailed description given by Kiefer (1988 pp 657-659), accounting for the fact that the data under consideration is uncensored. The data were supplied by the Social Sciences Data Archives at the Hebrew University of Jerusalem. At a first glimpse, the empirical hazard function appears to be constant for durations shorter than 40 days and increasing for longer durations. This suggests an exponential distribution for short durations and a Weibull distribution, with a Weibull parameter greater than unity, for longer durations. A possible economic interpretation of this phenomenon is that strikes reach a crisis point. When this point is reached, the costs incurred on both sides of the dispute encourage agents to make compromises and reach a settlement at a faster rate.

We are interested in whether a crisis point exists or not and in devising a suitable test for the phenomenon. There are 1082 observations for strikes lasting less than 40 days and 24 for longer strikes. The Lagrange multiplier (LM) test for the hypothesis of exponentially distributed durations against the Weibull alternative is 2.834 for spells less than 40 days and it is 10.306 for spells which are at least 40 days. Under the null hypothesis, the LM test is distributed asymptotically chi-square with one degree of freedom. At the 5% level of significance, the critical value is 3.841. The LM test supports then the graphical observation that the hazard function has a break at about the 40-day mark, with exponential specification for short spells and Weibull for longer ($\geq$ 40 days) spells.

The problem with the LM test and related likelihood-based tests is that their small sample behavior is generally unknown. Kiefer (1988) and others noted that inferences about long durations are often based on fewer observations. The typical existence of a small sample of long durations raises doubts on the reliability of classical asymptotic tests, such as the LM test. In the next section, we shall consider a suitable test and derive a reliable approximation for its finite sample distribution.

2.2 A Proportional Hazard Model

Let δ be a random variable representing the duration of a strike and t be a point in time measured from the commencement of the strike. Our starting point is the proportional hazard model

$$-\ln \Lambda_0(\delta;\theta) = X\beta + \varepsilon \ , \tag{1}$$

where

$$\Lambda_0(\delta;\theta) = \int_0^{\delta} \lambda_0(t;\theta)dt$$

is the integrated baseline hazard, parametrized by θ, X is a set of k exogenous covariates, and ε is a type I extreme value (Gumbel) variate with known cumulants. The explanatory variables chosen for the model are: x_1, an intercept, x_2, a dummy variable taking the value of unity if the strike has been authorized and zero otherwise, and x_3, the number of participating strikers. Both variables are indicative of the 'degree of militancy' of the strike (Lawrence 1984). There are 675 observations for the variables in the model, 660 for strike durations less than 40 days, and 15 for strikes lasting at least 40 days. It is well known that the exponential and Weibull specifications yield

$$-\ln\delta = X^*\beta + \varepsilon^* \tag{2}$$

and

$$-\ln\delta = X^*\beta/\alpha + \varepsilon^*/\alpha \tag{3}$$

respectively. Here X^* is the same matrix as X, except that its first vector is a column of $\{\ln\gamma - \Gamma'(1)\}$'s, $\Gamma'(1) = -E(\varepsilon) = -0.57721$, γ is the exponential parameter, α is the Weibull parameter and ε^* is a centered Gumbel variate. The explanatory variables are measured in deviations from their means. When the Weibull parameter α equals unity, the model is exponential.

For short durations, we fit (2) by least squares to obtain

$$\begin{array}{lllll} (\widehat{-\ln\delta})_i = & -1.316 & -0.197x_{2i} & +0.000012x_{3i}, & n = 660 \\ & (0.038) & (0.077) & (5\times10^{-6}) & F = 5.871 \\ & & & & \bar{R}^2 = 0.0143. \end{array}$$

Standard errors are in brackets. The t-statistics are not t-distributed, due to the nonnormality of ε^*. Running the same specification for the longer strikes, we get

$$\begin{array}{lllll} (\widehat{-\ln\delta})_i = & -4.003 & -0.065x_{2i} & +0.000014x_{3i}, & n = 15 \\ & (0.061) & (0.146) & (4\times10^{-5}) & F = 6.889 \\ & & & & \bar{R}^2 = 0.457. \end{array}$$

It is observed that the standard errors from the second regression are larger than the corresponding ones from the first regression, raising further the suspicion that a Weibull model with $\alpha > 1$ should have been fitted for the longer durations.

In view of (2) and (3), a test for a constant hazard across all strikes' lengths against a Weibull alternative, is equivalent to a test of parameters' constancy. The Chow test (Chow 1960) for a structural break is

$$r = \frac{n-2k}{k} \quad \frac{\varepsilon^{*\prime} A \varepsilon^*}{\varepsilon^{*\prime} M \varepsilon^*} \,,$$

where n is the total number of durations,

$$\begin{aligned} A &= X^*(X^{*\prime}X^*)^{-1}R'\{R(X^{*\prime}X^*)^{-1}R'\}^{-1}R(X^{*\prime}X^*)^{-1}X^{*\prime} , \\ M &= I_n - X^*(X^{*\prime}X^*)^{-1}X^{*\prime} \end{aligned}$$

and

$$R = (I_k; -I_k) \,.$$

The matrix X^* is partitioned as

$$X^* = \begin{pmatrix} X_1^* & 0 \\ 0 & X_2^* \end{pmatrix} \quad ,$$

with X_1^* being $n_1 \times k$ and X_2^* being $(n - n_1) \times k$. n_1 is the sample size of 'short' durations, i.e., strikes which last less than 40 days.

The Chow test is probably the most indigenous candidate for the hypothesis $\alpha = 1$. The problem is that because ε^* is a centered Gumbel and not normal, r is not F-distributed. Below we derive an F-expansion for its distribution.

3 An F-Expansion to the Distribution of the Test Statistic

3.1 General

Let r_0 be a statistic distributed as $F_{k,n-2k}$. Consider the pseudovariate S, assumed to be independent of r_0 and defined as $r = r_0 + S$. The conditional density of r given

$S = s$, and evaluated at $r = \tilde{r}$, is $f_{k,n-2k}(\tilde{r} - s)$, i.e., the density of an f-variate with k and $n - 2k$ degrees of freedom. The unconditional density of r is

$$f_r(\tilde{r}) = E_S\{f_{k,n-2k}(\tilde{r} - s)\} \,. \tag{4}$$

Expanding (4) about $s = 0$ and taking expectations w.r.t. S, we obtain

$$\begin{aligned} f_r(\tilde{r}) &= f_{k,n-2k} - \mu_1'(S) f_{k,n-2k}'(\tilde{r}) + \frac{1}{2}\mu_2'(S) f_{k,n-2k}''(\tilde{r}) \\ &- \frac{1}{3!}\mu_3'(S) f_{k,n-2k}^{(3)}(\tilde{r}) + \frac{1}{4!}\mu_4'(S) f_{k,n-2k}^{(4)}(\tilde{r}) - \cdots, \end{aligned} \tag{5}$$

where $\mu_j'(S)$ is the jth raw moment of S, $j \geq 1$, and $f_{k,n-2k}^{(j)}(\tilde{r})$ is the jth derivative of $f_{k,n-2k}(\tilde{r})$. Because S is independent of r_0,

$$\kappa_j(S) = \kappa_j(r) - \kappa_j(r_0), \qquad j \geq 1 \,,$$

that is, the jth cumulant of S equals the difference between the jth cumulants of r and r_0. Upon conversion from cumulants to raw moments, the terms appearing in (5) are

$$\begin{aligned} \mu_1'(S) &= \mu_1'(r) - \mu_1'(r_0) \\ \mu_2'(S) &= \mu_2'(r) - \mu_2'(r_0) + 2[\mu_1'(r_0)]^2 - 2\mu_1'(r)\mu_1'(r_0) \\ \mu_3'(S) &= \mu_3'(r) - \mu_3'(r_0) + 6\mu_1'(r_0)\mu_2'(r_0) - 3\mu_1'(r)\mu_2'(r_0) \\ &- 3\mu_1'(r_0)\mu_2'(r) + 6\mu_1'(r)[\mu_1'(r_0)]^2 - 6[\mu_1'(r_0)]^3 \\ \mu_4'(S) &= \mu_4'(r) - \mu_4'(r_0) + 8\mu_1'(r_0)\mu_3'(r_0) - 4\mu_1'(r)\mu_3'(r_0) \\ &- 4\mu_1'(r_0)\mu_3'(r) + 6[\mu_2'(r_0)]^2 - 6\mu_2'(r)\mu_2'(r_0) \\ &- 36[\mu_1'(r_0)]^2\mu_2'(r_0) + 24\mu_1'(r)\mu_1'(r_0)\mu_2'(r_0) \\ &+ 12[\mu_1'(r_0)]^2\mu_2'(r) + 18[\mu_1'(r_0)]^4 - 24\mu_1'(r)[\mu_1'(r_0)]^3 \\ &+ 6[\mu_1'(r_0)]^2 + 6[\mu_1'(r)]^4 - 6[\mu_1'(r)]^2 \,. \end{aligned} \tag{6}$$

From standard distribution theory (e.g., Johnson and Kotz 1970), we know that

$$\mu_j'(r_0) = \left\{\Gamma\left(\frac{k}{2}\right)\Gamma\left(\frac{n}{2} - k\right)\right\}^{-1}\left(\frac{n}{k} - 2\right)^j \Gamma\left(\frac{k}{2} + j\right)\Gamma\left(\frac{n}{2} - k - j\right), \quad n - 2k > 2j.$$

Although the exact moments of r are unknown, we can make use of the Laplace approximation to these moments, writing

$$\mu_j'(r) \simeq \frac{\mu_j'(\varepsilon^{*\prime}A\varepsilon^*/k)}{[\mu_1'(\varepsilon^{*\prime}M\varepsilon^*/(n-2k))]^j} = \left(\frac{6}{\pi}\right)^j \mu_j'(\varepsilon^{*\prime}A\varepsilon/k), \quad j \geq 1,$$

where $\pi^2/6 = \mathrm{Var}(\varepsilon^*)$. See Lieberman (1994) for justification. Because $(\varepsilon^{*\prime}A\varepsilon^*/k)$ is quadratic in ε^*, its first four cumulants are (Lieberman 1997)

$$\begin{aligned}
\kappa_1(\varepsilon^{*\prime}A\varepsilon^*/k) &= \kappa_2 = \pi^2/6 \\
\kappa_2(\varepsilon^{*\prime}A\varepsilon^*/k) &= \left\{\kappa_4 \sum a_{ii}^2 + 2\kappa_2^2 k\right\}/k^2 \\
\kappa_3(\varepsilon^{*\prime}A\varepsilon^*/k) &= \left\{\kappa_6 \sum a_{ii}^3 + 12\kappa_4\kappa_2 \sum_{i,j} a_{ii}a_{ij}^2 + 2\kappa_3^2 \sum_{i,j} (3a_{ii}a_{ij}a_{jj} + 2a_{ij}^3) + 8\kappa_2^3 k\right\}/k^3 \\
\kappa_4(\varepsilon^{*\prime}A\varepsilon^*/k) &= \left\{\kappa_8 \sum a_{ii}^4 + 24\kappa_6\kappa_2 \sum_{i,j} a_{ii}^2 a_{ij}^2 + 24\kappa_5\kappa_3 \sum_{i,j} a_{ii}^2 a_{jj} a_{ij}\right. \\
&\quad + 32\kappa_5\kappa_3 \sum_{i,j} a_{ij}^3 a_{ii} + 8\kappa_4^2 \sum_{i,j} (3a_{ii}a_{jj}a_{ij}^2 + a_{ij}^4) \qquad (7) \\
&\quad + 48\kappa_4\kappa_2^2 \sum_{i,j,\ell} (2a_{ii}a_{ij}a_{i\ell}a_{j\ell} + a_{ij}^2 a_{i\ell}^2) \\
&\quad \left. + 48\kappa_3^2\kappa_2 \sum_{i,j,\ell} (2a_{ii}a_{ij}a_{j\ell}^2 + a_{ii}a_{jj}a_{i\ell}a_{j\ell} + 2a_{i\ell}a_{j\ell}a_{ij}^2) + 48\kappa_2^4 k\right\}/k^4,
\end{aligned}$$

where $\kappa_r = \Gamma^{(r)}(1)$, $r > 2$ are the cumulants of ε^* and $\Gamma^{(r)}(1)$ is the rth derivative of the gamma function evaluated at one. The first eight cumulants of ε^*, appearing above are

$$\begin{aligned}
\kappa_1 &= 0 \\
\kappa_2 &= 1.6449 \\
\kappa_3 &= -2.4041 \\
\kappa_4 &= 6.4939 \\
\kappa_5 &= -24.8863 \\
\kappa_6 &= 122.0812 \\
\kappa_7 &= -726.0115 \\
\kappa_8 &= 5060.550\ .
\end{aligned}$$

Conversion of the $\kappa_j(\varepsilon^{*\prime}A\varepsilon^*/k)$'s into the $\mu_j'(\varepsilon^{*\prime}A\varepsilon^*/k)$'s is immediate. To complete the specification of the expansion (5), the terms required are

$$
\begin{aligned}
f_{k,\omega}(\tilde r) &= \left\{\tilde r B\left(\frac{\omega}{2},\frac{k}{2}\right)\right\}^{-1}\left(\frac{\tilde r k}{\tilde r k+\omega}\right)^{\frac{k}{2}}\left(\frac{\omega}{\tilde r k+\omega}\right)^{\frac{\omega}{2}} \\
f'_{k,\omega}(\tilde r) &= \left\{2\tilde r^2(\tilde r k+\omega)B\left(\frac{\omega}{2},\frac{k}{2}\right)\right\}^{-1}\left(\frac{\tilde r k}{\tilde r k+\omega}\right)^{\frac{k}{2}}\left(\frac{\omega}{\tilde r k+\omega}\right)^{\frac{\omega}{2}} \\
&\quad (-2\tilde r k-2\omega+k\omega-\tilde r k\omega) \\
f''_{k,\omega}(\tilde r) &= \left\{4\tilde r^3(\tilde r k+\omega)^2B\left(\frac{\omega}{2},\frac{k}{2}\right)\right\}^{-1}\left(\frac{\tilde r k}{\tilde r k+\omega}\right)^{\frac{k}{2}}\left(\frac{\omega}{\tilde r k+\omega}\right)^{\frac{\omega}{2}} \\
&\quad (-8\tilde r^2k^2-16\tilde r k\omega+8\tilde r k^2\omega \\
&\quad -6\tilde r^2k^2\omega-8\omega^2+6k\omega^2-4\tilde r k\omega^2-k^2\omega^2+2\tilde r k^2\omega^2-\tilde r^2k^2\omega^2) \\
f^{(3)}_{k,\omega}(\tilde r) &= \left\{8\tilde r^4(\tilde r k+\omega)^3B\left(\frac{\omega}{2},\frac{k}{2}\right)\right\}^{-1}\left(\frac{\tilde r k}{\tilde r k+\omega}\right)^{\frac{k}{2}}\left(\frac{\omega}{\tilde r k+\omega}\right)^{\frac{\omega}{2}} \\
&\quad (-48\tilde r^3k^3-144\tilde r^2k^2\omega+72\tilde r^2k^3\omega \\
&\quad -44\tilde r^3k^3\omega-144\tilde r^3k\omega^2+108\tilde r k^2\omega^2-60\tilde r^2k^2\omega^2 \\
&\quad -18\tilde r k^3\omega^2+30\tilde r^2k^3\omega^2-12\tilde r^3k^3\omega^2-48\omega^3 \\
&\quad +44k\omega^3-24\tilde r k\omega^3-12k^2\omega^3+18\tilde r k^2\omega^3-6\tilde r^2k^2\omega^3 \\
&\quad +k^3\omega^3-3\tilde r k^3\omega^3+3\tilde r^2k^3\omega^3-\tilde r^3k^3\omega^3),
\end{aligned}
$$

and so on. In the preceding, $\omega \triangleq n-2k$ and $B(\cdot,\cdot)$ is the beta function. Finally, termwise integration of (5) yields the F-expansion to the cdf of r

$$
\begin{aligned}
F_r(\tilde r) &= F_{k,n-2k}(\tilde r)-\mu_1'(S)f_{k,n-2k}(\tilde r)+\frac{1}{2}\mu_2'(S)f'_{k,n-2k}(\tilde r) \\
&- \frac{1}{3!}\mu_3'(S)f''_{k,n-2k}(\tilde r)+\frac{1}{4!}\mu_4'(S)f^{(3)}_{k,n-2k}(\tilde r)-\cdots \qquad (8)
\end{aligned}
$$

3.2 Discussion

The general idea of connecting the distribution of a statistic of interest with a pseudovariate with known cumulants is due to Davis (1976), who was primarily concerned with deviations from normal theory. The method was further discussed by McCullagh (1987).

The expansion (5) can be potentially constructed around any approximating density. A normal-based expansion, for instance, leads to the Edgeworth expansion, now standard in statistical applications. The choice of base, or leading term, is basically dictated by the geometry of the statistic. The Edgeworth expansion with a normal leading term is suitable for a standardized sum of i.i.d. variables by virtue of the central limit theorem. For the Chow F-statistic though, which is not asymptotically normal, the F-based expansion is much more natural. By "natural", it is meant that the cumulant structure of the approximating density is similar to that of the desired density.

The numerical example in the next section confirms that conventional approximation methods, such as the Edgeworth and saddlepoint expansions, which are built around the normal density, break down. The principal reason for this breakdown is that under Gumbel $\varepsilon^{*\prime}$s, the terms in the expansions, comprising of the cumulants of the approximating densities, do not decay as their order increase. The new approximation performs well and is very reliable.

4 Accuracy

For the model and data presented in Section 2, the Chow test for parameter stability against the alternative of a break at the 40-day mark yields 37.98. The p-value of the statistic with the wrongly-used $F_{3,669}$-distribution is zero. Fortunately, for this case, using a second order F-expansion, the p-value turns out to be zero as well, supporting the hypothesis that a crisis point occurs at around the 40-day mark, as was suspected previously.

To assess the small sample accuracy of the F-expansion, a random sample of 33 short strikes (less than 40 days) were selected together with 15 long durations. For the design matrix described in Section 2, we compute $F_r(\bar{r})$ by: simulation (SIM) consisting of 5000 replications, the second order F-expansion (F2), and the Edgeworth (EDG) and saddlepoint (SP) expansions. We take SIM to be the benchmark. The results are described in Figure 3. It is seen that the Edgeworth expansion assigns negative values and in fact decreases on some parts of the range of the statistic. Its absolute error is in general unacceptably large. The saddlepoint approximation performs somewhat better than EDG but its relative error is still significant. The curve of the second order F-expansion tracks the simulated line closely, showing that the expansion is clearly very accurate.

5 Conclusions

The paper was motivated by the suspicion that the hazard function for Israeli strikes is broken at some point. Preliminary visual inspections and the classical asymptotic LM test support this conjecture and it is found that the rate at which disputes end increases for strikes lasting at least 40 days, whereas it is constant for shorter durations. A possible economic interpretation for these findings is that when a crisis point has been reached, the costs incurred on both sides of the conflict drive agents to make compromises and reach an agreement at a faster rate.

As samples of long strike durations are typically small (Kiefer 1988), classical likelihood based tests are not likely to convey the correct message to the researcher. For this reason we suggested the use of the Chow test (Chow 1961) and derived a new, F-based expansion to the distribution of the test under Gumbel variates. In comparison to existing approximations, the new expansion suits better the geometry of the statistic. The superiority of the F-expansion over conventional approximation methods was illustrated with the data set.

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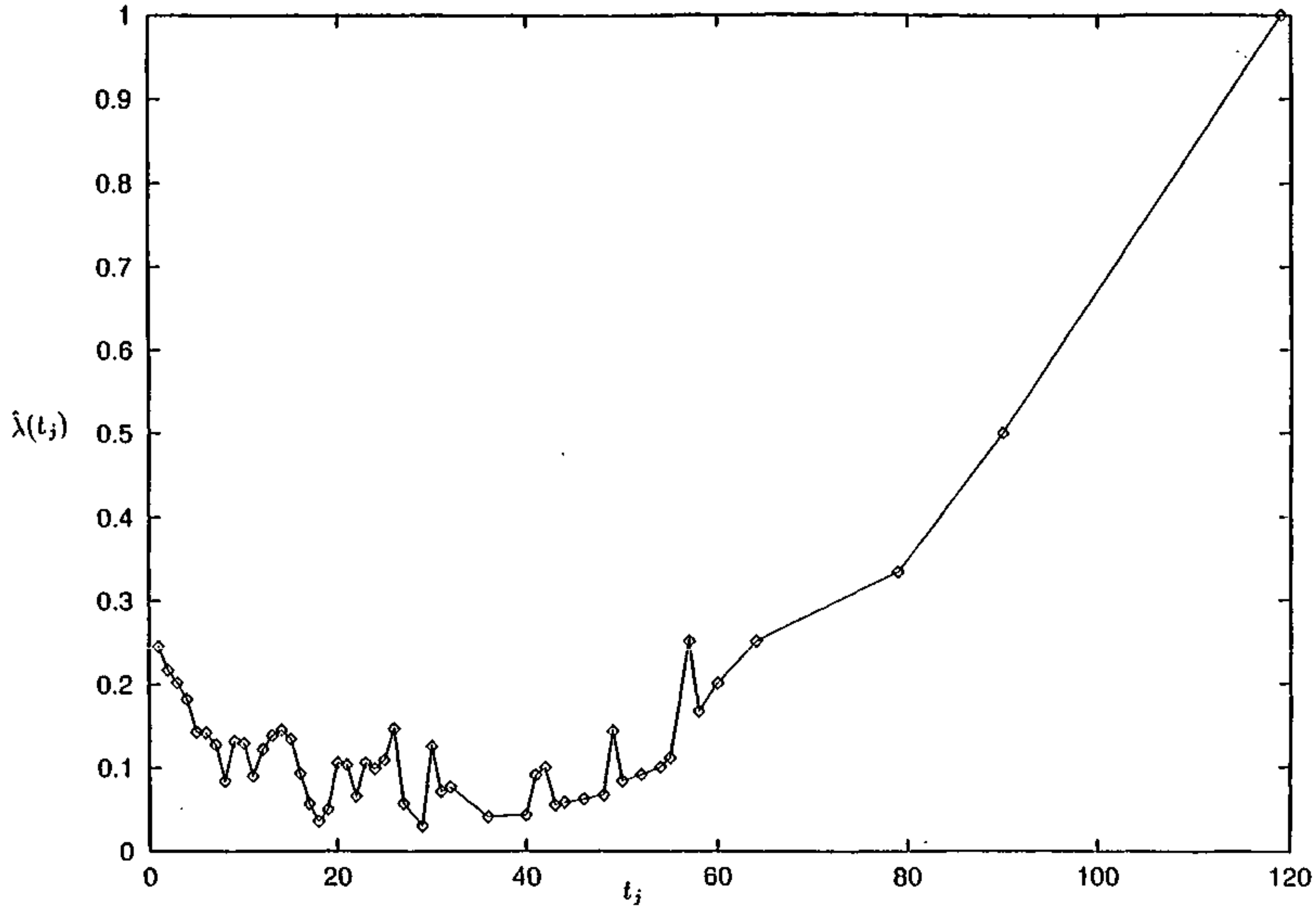
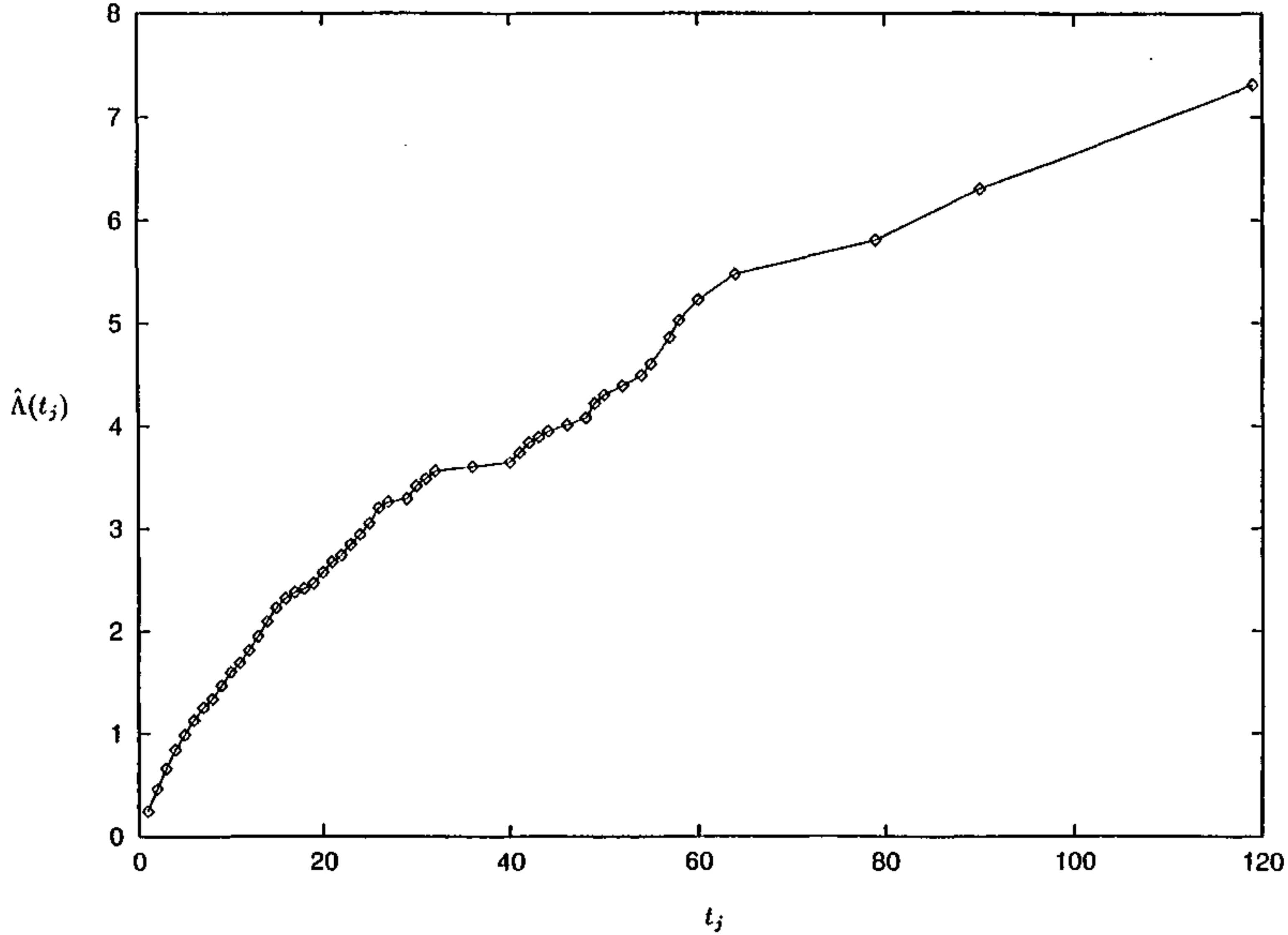
Figure 1: The Empirical Hazard Function



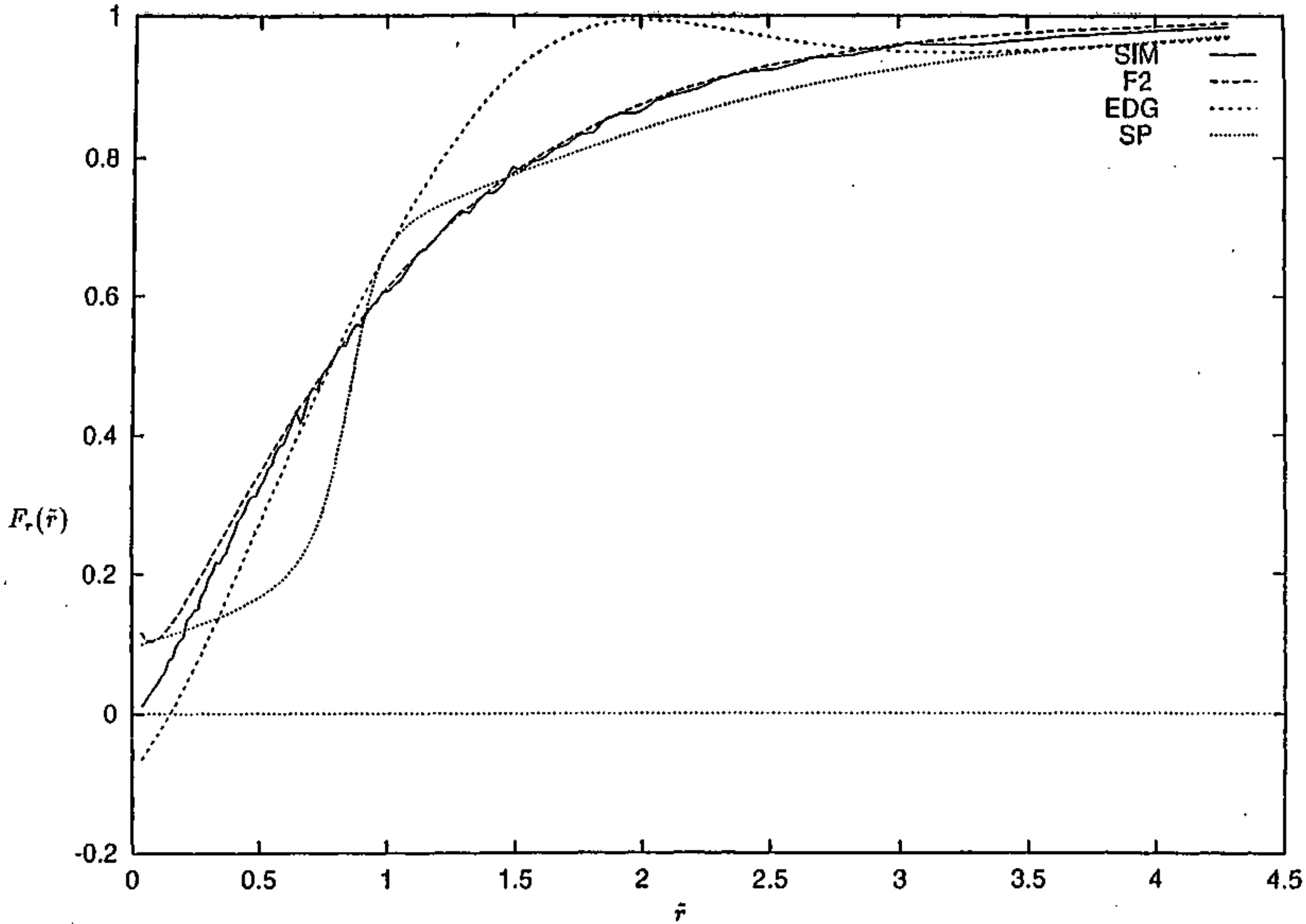
$\hat{\lambda}(t_j)$ is the empirical hazard function. It is calculated as the number of completed spells of duration t_j divided by the number of uncompleted spells before duration t_j.

Figure 2: The Empirical Integrated Hazard Function



$\hat{\Lambda}(t_j) = \sum_{i \le j} \hat{\lambda}(t_i)$

Figure 3: Approximations to the Distribution of the Chow Test in Gumbel Variates



SIM ≡ simulated cdf

EDG ≡ Edgeworth expansion

SP ≡ Lugannani-Rice saddlepoint expansion

F2 ≡ second order F-expansion