

# Tactical Prevention of Suicide Bombings in Israel

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Suicide bombings are the leading cause of death from terrorism in Israel. Counterterror tactics, such as the targeted killings or preemptive arrests of terror leaders or suspects, are meant to prevent such attacks. To investigate whether these tactics are successful, we estimated via maximum likelihood a family of shot-noise models from monthly data covering 2001 through 2003 to see whether we could predict the rate of suicide-bombing attacks as a function of prevention tactics over time. Although preventive arrests appear to lower the rate of suicide-bombing attacks, targeted killings seem to be followed by an increase in the number of suicide bombings. In addition, limited evidence suggests that the probability of intercepting a suicide bomber en route to an attack increases with the expected number of suicide-bombing attacks. Such an endogenous relationship could imply an upper limit on the rate of successful suicide bombings.

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Suicide terrorism has become a strategy of terrorists worldwide, who have conducted suicide attacks in Afghanistan, Egypt, England, India, Indonesia, Iraq, Israel, Lebanon, Pakistan, Russia, Spain, Sri Lanka, Saudi Arabia, Turkey, the United States (9/11), and Yemen (ICT, [http://www.ict.org.il/inter\\_ter/intnl\\_attacksearch\\_frame.htm](http://www.ict.org.il/inter_ter/intnl_attacksearch_frame.htm)). A growing number of studies address suicide bombing. Atran (2003, p. 1534) defines *suicide terrorism* as “the targeted use of self-destructing humans against noncombatant—typically civilian—populations to effect political change” and claims that the primary use of suicide terrorism is “as a weapon of psychological warfare intended to affect a larger public audience.” Berman and Laitin (2006, p. 8) claimed that suicide attacks around the world “were most often marked by a religious difference between the attackers and the victims.” Merari (1990) and Pedahzur et al. (2003) examined the characteristics of suicide bombers and their motivations. In a lead article in the *American Political Science Review*, Pape (2003, p. 1) argued that suicide terrorism follows a strategic logic, “one specifically designed to

coerce modern liberal democracies to make significant territorial concessions.” Pape also wrote that, over the past two decades, suicide terrorism has been rising “largely because terrorists have learned that it pays,” although this has been questioned (Abrahms 2004).

Countertactics used by governments around the world to fight suicide terrorism include nonlethal preventive measures (for example, intelligence-driven arrests of terror suspects) and lethal offensive measures (for example, targeted killings, or hits, of terrorist operatives or leaders). Ganor (2005) postulated that targeted killings can lead to a boomerang effect, causing the number of attacks to increase following such hits. In Ganor’s view, targeted killings increase suicide bombers’ motivation while reducing terrorist organizations’ operational capabilities.

Kaplan et al. (2005) developed a terror-stock model of suicide bombings based on theory proposed by Keohane and Zeckhauser (2003). In that model, suicide bombings derive from the number of terrorists actively involved in planning and conducting suicide bombings (the terror stock), and the intent of

counterterror measures, such as targeted killings or preemptive arrests, is to prevent suicide bombings by reducing the size of the terror stock. Within this modeling framework, Kaplan et al. (2005) found that, whereas preventive arrests were associated with reductions in suicide bombings, targeted killings were associated with increases in suicide-bombing attacks because the killings enhanced recruitment to the terror stock, consistent with Ganor's boomerang effect.

We revisited the Israeli–Palestinian case to investigate the effectiveness of targeted killings and preventive arrests in reducing the rates of suicide-bombing attacks. We did so by specifying and testing a series of shot-noise statistical models oriented toward estimating the downstream consequences of different counterterror events with monthly data on Palestinian suicide-bombing attacks and Israeli countermeasures from January 2001 to December 2003; these models are more flexible than those Kaplan et al. (2005) considered, and they do not rely on the terror-stock assumptions made previously. The models suggest whether hits and arrests are differentially associated with increases or decreases in the number of future suicide attacks, as well as the marginal impact of each intervention on the timing of downstream attacks. We also studied whether the likelihood of intercepting a dispatched suicide bomber en route, the prevention tactic of last resort, relates to the expected suicide-bombing hazard rate. It could if Israel's counterterror alertness increases when it anticipates an increase in suicide-bombing attacks, or alternatively if suicide-bombing attempts during periods with high attack frequency are poorly planned and of lower quality, making interception more likely. We also conducted sensitivity analyses to establish the robustness of our findings to the time period covered by our analysis and pressed our model to predict suicide bombings "out of sample" during the first four months of 2004. We emphasize that, as an observational statistical study, our results are suggestive, not definitive. Nonetheless, the good statistical fit and out-of-sample forecast obtained indicate that these suggestions should be considered seriously.

## Data Sources

We collected data on Palestinian suicide bombings within Israel's 1967 border (the green line), Israeli

arrests of terror suspects planning suicide-bombing operations, Israeli targeted killings of terror suspects, and Israeli interceptions of suicide bombers already en route (Mintz et al. 2004). We obtained data describing suicide bombings in Israel from Israel's Ministry of Foreign Affairs (<http://www.mfa.gov.il/MFA/Terrorism-+Obstacle+to+Peace/Palestinian+terror+since+2000/Victims+of+Palestinian+Violence+and+Terrorism+sinc.htm>) and from the International Policy Institute for Counter Terrorism's terror-attack database ([http://www.ict.org.il/arab\\_isr/mideast\\_attacksearch\\_frame.htm](http://www.ict.org.il/arab_isr/mideast_attacksearch_frame.htm)). We obtained data on targeted hits from B'tselem, the Israeli Information Center for Human Rights in the Occupied Territories (<http://www.btselem.org/index.asp>). We obtained published data on interceptions from the archives of the newspaper *Ha'aretz* (<http://www.haaretz.co.il>) and cross-validated them with data from Palestinian sources (for example, the Palestinian Human Rights Monitoring Group, <http://www.phrmg.org>, and the Palestinian Information Center, <http://palestine-info.info/index.html>), while we took data on Israeli preventive arrests from the Israel Defense Forces Web site (<http://www1.idf.il/dover/site/mainpage.asp?sl=EN&id=22&docid=37572.EN>). All the data we used to test our models were in monthly figures covering 2001 through 2003 (Mintz et al. 2004).

## Methods Overview

We defined a suicide-bombing attack as a realized or intercepted suicide bombing and assumed that the number of suicide-bombing attempts in a given month  $t$  follows a Poisson distribution with mean attack rate  $\lambda_t$ . The mean attack rate can be thought of as a strategic variable, in that terror organizations choose when (and hence how frequently) to attempt bombings (Pape 2003). However, the choices that terrorists make are themselves influenced by Israeli tactics. A series of arrests might dismantle a terror cell, reducing  $\lambda_t$  during the month in which the arrests occur and in subsequent months. A targeted killing might have the same effect, although it is also possible that such hits motivate terror organizations to intensify their efforts. Israeli preparedness and tactics themselves are the product of intelligence efforts that, presumably, derive from noisy estimates of the attack rate  $\lambda_t$ .

We did not attempt to model the deadly strategic game being played between suicide terrorists and Israeli intelligence and security forces. Our more modest goal was to see whether an empirical relationship exists between suicide-bombing attacks and Israeli tactics and also to see whether Israeli preparedness (as modeled by the probability of intercepting a suicide bomber) relates to the current expected risk of attack. We did not attempt to model Israeli tactics as a function of past terror attacks.

We modeled the mean suicide-bombing-attack rate in a given month as a function of Israeli tactics in the current and prior months. To do so, we adopted a flexible shot-noise specification that enabled us to estimate both the instantaneous and lagged effects of tactics on suicide-bombing attempts with a minimal number of parameters. Once we had a model for predicting suicide-bombing attacks, we asked whether predicted attack rates are related to the likelihood of intercepting suicide bombers (and the implications thereof).

## Modeling Suicide Bombing Attacks

To operationalize Israeli tactics, we let  $x_{it}$  denote the number of type- $i$  tactical events that occur in month  $t$ . In this analysis, there are only two types of tactics, hits (targeted killings) and preventive arrests, although the model easily generalizes to additional tactics should such data become available. We linked Israeli tactics to the mean suicide-attack rate via the log-linear equation

$$\log \lambda_t = \beta_0 + \sum_i \beta_i \gamma_{it}, \quad (1)$$

where

$$\gamma_{it} = \sum_{j=0}^t \theta_i^{t-j} x_{ij}. \quad (2)$$

In these equations,  $\gamma_{it}$  represents the discounted sum (or the effective stock) of type- $i$  tactical events from time 0 (the end of 2000 in our data) through month  $t$ , while  $\theta_i$  is the discount rate for type- $i$  tactical events (with  $\theta_i$  restricted to fall between 0 and 1). With this model, the impact of an additional type- $i$  tactical event in some month  $j$  is felt in month  $j$  as well as in all subsequent months:

$$\frac{\partial \gamma_{it}}{\partial x_{ij}} = \theta_i^{t-j}, \quad (3)$$

which shows that the weight attached to the marginal hit or arrest in month  $j$  in computing the effective stock of type- $i$  tactical events in future months  $t > j$  decreases geometrically in  $t$ . Values of  $\theta_i$  close to zero imply that type- $i$  tactical events in a given month affect only the mean suicide-bombing-attack rate in that same month, whereas values of  $\theta_i$  near one imply that type- $i$  tactical events in a given month affect the mean suicide-bombing-attack rate in all subsequent months.

Our choice of the log-linear function linking tactics to the mean attack rate is sensible on both statistical and theoretical grounds. Statistically, we required that the mean attack rate remain nonnegative in all months, a requirement obviously satisfied by Equation (1). Theoretically, a marginal change in the effective stock of type- $i$  tactical events in month  $i$  produces a proportional change in the mean attack rate, that is,

$$\frac{\partial \lambda_t}{\partial \gamma_{it}} = \beta_i \lambda_t. \quad (4)$$

Thus, in our formulation, the impact of a marginal increase in the effective tactical stock will have a greater effect in months in which the suicide-bombing-attack rate is high. For example, if targeted hits serve to increase rather than decrease the suicide-bombing-attack rate (as would be the case if the relevant  $\beta > 0$ ), retaliatory suicide-bombing attacks would occur with greater frequency during months that already exhibit high levels of terror attacks.

The constant term  $\beta_0$  serves to regulate the overall level of suicide-bombing attacks. Were there no relationship whatsoever between tactical events and suicide-bombing attacks,  $\beta_0$  would set the time-average attack rate (and all other  $\beta_i$ 's would equal zero). The constant term also absorbs the impact of tactical events that occurred prior to 2001 on subsequent suicide-bombing-attack rates.

To comprehend how all of the parameters of the model relate together, suppose that the mean monthly attack rate remained constant and equal to  $\bar{\lambda}$  because of constant monthly type- $i$  tactical-event rates. Starting from any month  $t$ , it then follows that the incremental change in the total expected number of attacks from month  $t$  on into the future resulting from an incremental change in  $x_{it}$  is given by

$$\frac{\partial}{\partial x_{it}} \sum_{j=t}^{\infty} \lambda_j = \sum_{j=t}^{\infty} \beta_i \theta_i^{t-j} \bar{\lambda} = \frac{\beta_i}{1 - \theta_i} \bar{\lambda}. \quad (5)$$

Thus, the marginal increase (or decrease) in the total number of suicide-bombing attempts over all time that would follow from a marginal increase in the type-*i* tactical-event rate in month *t* would equal the constant multiple  $\beta_i/(1 - \theta_i)$  times the (presumed constant) mean attack rate  $\bar{\lambda}$ . The roles of  $\beta_i$  and  $\theta_i$  are clear from Equation (5). Larger absolute values of  $\beta_i$  imply larger increases in the incremental number of suicide-bombing attacks if  $\beta_i > 0$  (decreases if  $\beta_i < 0$ ) than do smaller absolute values of  $\beta_i$ , whereas values of the discount rate  $\theta_i$  that are close to one imply many more (or fewer) incremental attacks (depending on the sign of  $\beta_i$ ) than values of  $\theta_i$  that are close to zero. The overall marginal increase in attacks is also proportional to the mean attack rate  $\bar{\lambda}$ , again reflecting the notion that incremental changes in terror activity are scaled by the prevailing level.

To estimate the parameters of this model, we employed the method of maximum likelihood. If  $a_t$  is the observed number of suicide-bombing attacks in month *t*, we maximize the log-likelihood function

$$\log \mathcal{L} = \sum_t (a_t \log \lambda_t - \lambda_t) \quad (6)$$

as a function of the  $\beta$  and  $\theta$  parameters, where  $\lambda_t$  follows Equation (1). We estimated standard errors and conducted goodness-of-fit tests using the standard methods associated with maximum-likelihood estimation (Table 1).

### Statistical Analysis

The suicide-bombing-attack rate was not constant over time (Figure 1); thus it is not a surprise that a model assuming only random (Poisson) variation in suicide-bombing attacks around a constant mean attack rate fails to fit the data (the goodness-of-fit  $\chi^2$  statistics are far too large to accept the constant-attack-rate model, as shown by the associated tiny *p*-values (Table 1)), even though the overall monthly suicide-bombing-attack rate can be estimated with precision at 3.33 attacks per month (95 percent confidence interval 2.74–3.93).

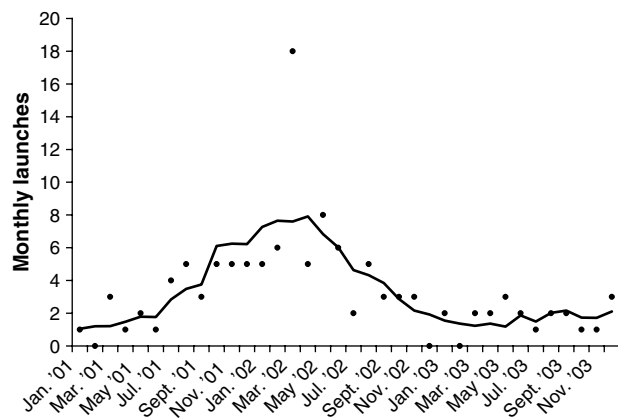
A shot-noise model incorporating targeted hits improves the fit of the model, although the goodness-of-fit statistics are still unacceptable. The hit slope  $\beta_H$  is positive, suggesting that a marginal increase in the effective stock of hits in a given month is associated

Parameter	Constant attack rate	Hit-dependent attack rate	Arrest-dependent attack rate	Hit- and arrest-dependent attack rate
Constant ( $\beta_0$ )	1.204 (0.091)	0.376 (0.309)	1.539 (0.133)	0.111 (0.389)
Hit slope ( $\beta_H$ )	—	0.084 (0.031)	—	0.096 (0.029)
Hit discount ( $\theta_H$ )	—	0.800 (0.048)	—	0.959 (0.044)
Arrest slope ( $\beta_A$ )	—	—	-0.006 (0.007)	-0.031 (0.014)
Arrest discount ( $\theta_A$ )	—	—	0.928 (0.167)	0.870 (0.059)
Log likelihood	24.48	29.21	31.21	49.86
Likelihood ratio $\chi^2$ ( <i>p</i> -value)	83.45 ( $7.8 \times 10^{-6}$ )	73.97 ( $5.6 \times 10^{-5}$ )	69.98 (0.0002)	32.68 (0.38)
Pearson $\chi^2$ ( <i>p</i> -value)	105.60 ( $5.2 \times 10^{-9}$ )	88.14 ( $6.5 \times 10^{-7}$ )	79.58 ( $1.0 \times 10^{-5}$ )	33.16 (0.36)

**Table 1:** In this table of the maximum-likelihood estimation of alternative shot-noise models, the cell entries are parameter estimates (standard errors) for the different models considered. We also show the log-likelihood (Equation (6)) and goodness-of-fit measures (likelihood ratio and Pearson  $\chi^2$  statistics).

with an increase in the suicide-bombing-attack rate by 8.4 percent (standard error 3.1 percent) over the current level. This effect is significant ( $z = 2.71$ , *p*-value = 0.003) but is conditional on a poorly fitting model.

With a model based solely on preventive arrests, we obtained a slightly improved but still unacceptably poor statistical fit to the data (Table 1). The arrest slope  $\beta_A$  is negative, suggesting that arrests are associated



**Figure 1:** This graph shows the actual suicide-bombing attacks (dots) and the expected suicide-bombing attacks (line) as estimated with our shot-noise model per month in Israel for 2001 through 2003.

with reducing the level of suicide-bombing attacks. However, again the poor fit of this model mitigates against taking this result too seriously, as does the fact that the estimated arrest slope is not statistically different from zero ( $z = -0.86$ ,  $p$ -value = 0.39).

Taken together, however, hits and arrests appear strongly predictive of suicide-bombing attempts (Table 1). The expected monthly suicide-bombing-attack rates from this model show striking agreement with observed attacks over time (Figure 1). The tactical event slopes and discount rates are well identified for both hits and arrests, while the goodness-of-fit statistics are extremely good (Table 1,  $\chi^2 \approx 33$ ,  $df = 31$ ,  $p$ -value  $\approx 0.37$ ), despite an obvious outlier in March 2002, the month with the most suicide-bombing attempts in Israel's history.

What does this model tell us? First, the evidence suggests that targeted hits are associated with increases in suicide-bombing attacks. A marginal increase in the effective stock of hits in a given month is associated with a 9.6 percent (standard error 2.9 percent) increase in the attack rate in that month. Moreover, the estimated hit discount of  $\theta_H = 0.959$  (standard error 0.044) is very close to one, suggesting that a targeted killing in any month is associated with suicide-bombing attacks for many, many months following that hit. Were the suicide-bombing-attack rate constant (and that is not the case), the incremental number of suicide-bombing attacks that would follow any given hit would equal  $\beta_H/(1 - \theta_H) = 0.096/(1 - 0.959) = 2.34$  times the prevailing monthly rate. For example, if a mean attack rate of 3.33 per month prevailed, our model would predict that a marginal hit would lead to an incremental  $2.34 \times 3.33 = 7.79$  or roughly eight additional suicide-bombing attacks over all time beyond what would have occurred had there been no additional hit.

On the other hand, our analysis also suggests that preventive arrests are associated with reduced suicide-bombing rates. A marginal increase in the effective stock of such arrests in a given month is associated with a 3.1 percent (standard error 1.4 percent) reduction in the rate of suicide bombings in that month. The estimated arrest discount of  $\theta_A = 0.87$  (standard error 0.059) is also close to one, although it is smaller than the estimated hit discount. Still, preventive arrests yield future benefits in this model; again assuming a

constant attack rate to help us interpret the results, the model suggests that a marginal increase in arrests in a given month would reduce the total number of future suicide-bombing attacks by  $0.031/(1 - 0.87) = 0.24$  times the prevailing attack rate. Again using the observed monthly attack rate of 3.33 as an example, a marginal arrest would prevent  $0.24 \times 3.33 = 0.8$  suicide-bombing attacks.

The model suggests that the negative impact of a single targeted hit far outweighs the benefits of a single preventive arrest, but we should remember that the number of arrests in our data (316) far outweighs the total number of targeted killings (75). Furthermore, the number of preventive arrests greatly increased with the launch of Operation Defensive Shield following the suicide bombing of a Passover seder in Netanya that killed 30 Israelis and wounded another 140 on March 27, 2002. Indeed, only 52 preventive arrests were reported from January 2001 through June 2002, followed by 264 from July 2002 through December 2003. This contrasts with the distribution of targeted killings over time: 38 hits from January 2001 through June 2002, and 37 from July 2002 through December 2003. The effective ratio of arrests to hits in the last 18 months of our study period was  $264/37 = 7.1$ , so although the preventive impact of arrests might appear slight, the sheer volume of such arrests had a substantial effect.

## Sensitivity Analysis

One potential concern with our analysis is that, although our data begin in January 2001, suicide bombings and associated counterterror tactical events began in earnest in September 2000. Given that our model highlights the downstream consequences of tactical events via the shot-noise formulation, it is fair to question the stability of our results as a function of the starting date of analysis. We earlier argued that the constant term  $\beta_0$  is responsible for absorbing the impact of pre-2001 tactical events on later suicide-bombing attempts, and we would expect to obtain different estimates of  $\beta_0$  for analyses that start at different times. However, the other parameters are intended to capture the relationship between tactical events and attack rates and thus should reflect the state of affairs both pre- and post-January 2001.

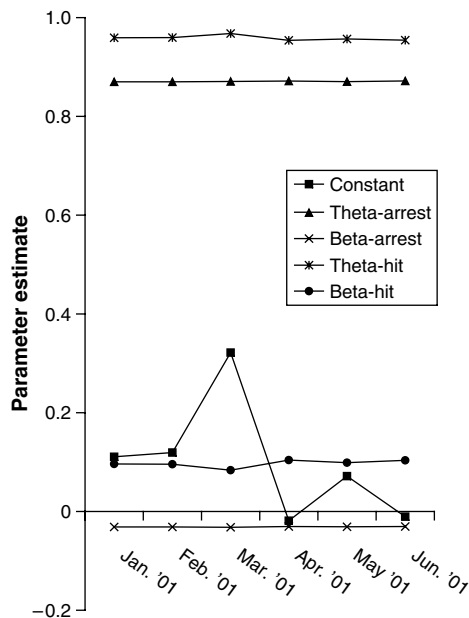


Figure 2: To analyze the sensitivity of our shot-noise parameter estimates to the starting date of analysis, we plotted the five parameter estimates resulting from analyses starting in each of the six months January through June of 2001.

To investigate, we reestimated our model by successively advancing the starting month of analysis from January through June of 2001 (Figure 2). The degree of stability among the estimated tactical-event slopes and discount rates was remarkable: only the estimated constant term changed over the six months. This suggests that the relationships between tactics and suicide-bombing attacks captured by our model were not unduly influenced by the start date of the analysis.

## Terror Preparedness, Attack Quality, and En-Route Interception

Having developed a model linking suicide-bombing attempts to counterterror tactics, we asked whether the likelihood of intercepting suicide bombers en route was higher in some time periods than in others. Intercepting suicide bombers who are already traveling to the target of attack is the prevention tactic of last resort (Atran 2003). Such interceptions are rarely the result of chance but rather the result of intelligence-driven search and pursuit that often ends with the capture of a would-be bomber at a checkpoint, roadblock, or

some other location en route to the target of attack. One would therefore suspect that such interceptions are more likely when the intelligence services have good reason to believe an attack is imminent. Alternatively, interceptions could be more likely when suicide attacks are spontaneous, less organized, or generally of lower quality. Given the relationship shown between targeted hits and suicide-bombing attempts, it could be that, although targeted hits spawn downstream attacks, the quality of such attacks is lower than it would be otherwise, making them easier to intercept.

Intelligence data that would enable a direct test of such propositions are not available for public scrutiny. However, if the intelligence and security services are operating effectively, it is reasonable to assume that the time periods during which the intelligence services believe suicide-bombing operations are underway correspond to those time periods with the highest expected suicide-bombing-attack rate, a variable already estimated via the shot-noise model. And surely the security services are in a heightened state of preparedness when the anticipated suicide-bombing-attack rate increases.

If the probability of intercepting a suicide bomber en route is in fact higher in months with higher expected attack rates, an interesting endogeneity arises. Let  $p(\lambda)$  denote the probability of intercepting a suicide bomber en route when the expected suicide-bombing-attack rate equals  $\lambda$ . Then, the expected rate of suicide bombings (attacks in which the terrorist actually detonates the bomb), denoted by  $b(\lambda)$ , is given by

$$b(\lambda) = \lambda(1 - p(\lambda)). \quad (7)$$

If  $p(\lambda)$  is increasing in  $\lambda$ , then  $b(\lambda)$  might possess an upper bound, which in turn would imply a worst-case suicide-bombing-attack rate. The idea is that as the suicide-bombing-attack rate increases, intelligence services become more attuned to such attacks and successfully intercept a greater fraction of them. There could thus exist a phenomenon whereby terrorists try too hard: if increasing the rate of suicide bombings serves to increase greatly the probability of interception, the number of actual bombings could decline even as suicide-bombing attempts increase. For terrorists, the intriguing possibility exists that there is an optimal suicide-bombing-attack rate that maximizes

the expected number of successful suicide bombings. This optimal attack rate  $\lambda^*$  is the root of the equation

$$\frac{db(\lambda)}{d\lambda} = 1 - p(\lambda) - \lambda \frac{dp(\lambda)}{d\lambda} = 0, \quad (8)$$

and the corresponding worst-case suicide-bombing rate would equal  $b(\lambda^*)$ .

To explore this with our data, we fit via maximum likelihood a logistic model for the probability of intercepting a suicide-bombing attack en route as a function of the expected attack rates estimated from our shot-noise model, that is, we specified

$$p(\lambda) = \frac{1}{1 + e^{-(\alpha_0 + \alpha_1 \lambda)}}. \quad (9)$$

The estimated coefficients (standard errors) are given by  $-1.6625$  ( $0.5024$ ) and  $0.1527$  ( $0.0872$ ) for  $\alpha_0$  and  $\alpha_1$ , respectively. The  $\alpha_1$  coefficient linking the expected attack rate to the probability of interception is positive and marginally significant ( $z = 1.75$ ,  $p$ -value =  $0.08$ ). This provides weak evidence consistent with our proposition. The graph of this relationship is nearly linear, with interception probabilities doubling from roughly 20 percent to 40 percent as the expected attack rate varies from roughly one to seven attacks per month, although the statistical fit is weak (Figure 3).

The optimal attack rate that solves Equation (8) in this case roughly equals 12 attempts per month, which

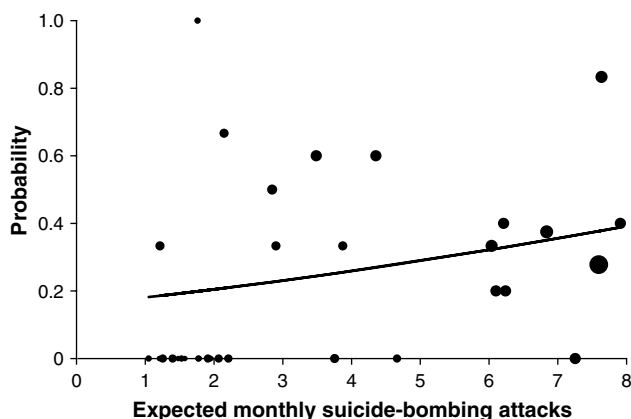


Figure 3: We plotted the observed (dots) and estimated (curve) probability of en-route interception as a function of expected suicide-bombing-attack rates based on the logistic model of Equation (9). The sizes of the dots are proportional to the number of suicide-bombing attempts associated with each data point.

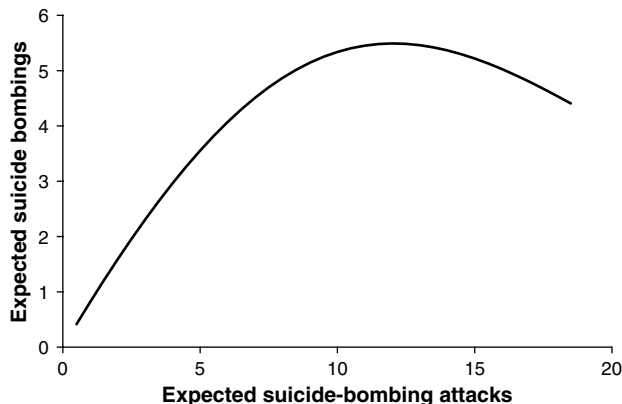


Figure 4: We calculated the number of successful suicide bombings expected monthly as a function of the monthly suicide-bombing-attack rate. The optimal suicide-bombing-attack rate is 12 attempts per month, and the worst-case rate of realized suicide bombings is 5.5 per month.

would result in a worst case of 5.5 expected bombings monthly (and an interception probability of 54 percent) (Figure 4). The optimal attack rate of 12 exceeds the expected monthly rates produced by the shot-noise model (the largest of which is about eight), while the monthly worst-case bombing rate of 5.5 exceeds the maximum estimated (which is 4.8).

Although we find these ideas intriguing, the evidence available to us (Figure 3) supports them only weakly. Nonetheless, the possible endogeneity of the interception probability with the expected suicide-bombing-attack rate is intriguing, given the implications for optimal terrorist behavior and worst-case suicide-bombing rates, and is worthy of further research. We also believe that our thesis is plausible in light of more recent events in Israel.

### Out of Sample in Spring 2004

From January through April 2004, only three suicide bombings occurred inside Israel’s green line (one in each of January, February, and March). There were three hits in February, three in March, and one in April. In particular, Ahmed Yassin, the head of the Hamas, was killed on March 22, and his successor, Abdel Rantisi, was fatally targeted on April 18. However, there were also nine preventive arrests in January, eight in February, 17 in March, and 14 in April.

By applying our best-fitting shot-noise model to the monthly hits and preventive arrests data, we obtained

monthly suicide-bombing-attack rates, and by filtering these rates through our estimated interception probability model (Equation (9)), we obtained estimated interception probabilities. By combining these two estimated quantities, we estimated suicide-bombing rates of 1.64, 2.13, 2.09, and 2.08 over the four months in question for a total of 7.94 expected successful bombings. Only three bombings were observed, a deficit of 4.94. This deficit is within the bounds of chance, however: if  $X$  is a Poisson random variable with mean 7.94, then

$$\Pr\{|X - 7.94| \geq 4.94\} = 0.105. \quad (10)$$

In other words, if the true expected number of successful suicide bombings was 7.94, at least 4.94 fewer (or additional) bombings would occur with a probability of 10.5 percent. In this light, seeing only three bombings cannot be viewed as unusual. We do not take this result as proof that our models have uncovered the truth. Rather, we more modestly contend that these out-of-sample data are not inconsistent with our previously estimated models.

## Discussion

We fit a parsimonious shot-noise model that appears to be consistent with observed suicide-bombing attacks over time in Israel. This model tells a simple story. Initially, Israel employed targeted hits of terror operatives and leaders in response to suicide terrorism after September 2000. However, these hits appeared to be associated with increased rather than decreased suicide-bombing-attack rates, and thus the number of suicide-bombing attacks continued to grow. After Israel launched Operation Defensive Shield and began arresting terror suspects in the West Bank on a large scale, suicide-bombing rates began to fall. Thus, it seems to us that preventive arrests, and not targeted killings, were more responsible for the decline in suicide-bombing attacks in Israel between March 2002 and April 2004.

Other contributing explanations are possible for events in Israel. For example, Israel began constructing a security fence in mid-2002, but by the end of 2003, it had finished only about one-quarter of the barrier. The partially completed security fence no doubt changed the geography of suicide bombings over the

latter part of our study by redirecting attacks from areas protected by the fence to unprotected cities, such as Jerusalem and Tel Aviv (Kaplan et al. 2005), but it could not plausibly have contributed greatly to reducing the overall level of attacks over the period of our study. Border closures and curfews are another tactic Israel employed to prevent terrorist infiltrations from the West Bank. However, the timing of curfews and closures likely correlates with the timing of operations intended to arrest leading terror suspects, and if so, it would be difficult to separate the preventive effects of closures from the preventive effects of arrests (Kaplan et al. 2005).

Some people may argue that targeting the political leaders of terror organizations may cause different outcomes than hitting operatives. With few exceptions, Israel did not begin targeting the political leaders of Hamas and Islamic Jihad until the summer of 2003 (near the end of our study period); thus, we cannot separate the effects of hits on political leaders from hits on operatives in this study.

We also found weak evidence consistent with an interesting endogeneity hypothesis, namely, that the probability of intercepting suicide bombers already en route to an attack is an increasing function of the expected attack rate. We hypothesized this relationship as an implication of effective intelligence and security operations: in periods with high expected attack rates, either preparedness increases, leading to an increased probability of interception or, alternatively, bombing attempts are of poor quality, making them easier to intercept. Either interpretation leads to the possibility of an optimal terrorist-attack rate and a worst-case (that is, maximal) rate of suicide bombings.

Together, our models provide a means of understanding some developments in Israel. On March 22, 2004, Israel assassinated Ahmed Yassin, the founder of the Hamas terror group responsible for the majority of suicide bombings in Israel. Less than one month later, Yassin's successor, Abdel Aziz Rantissi, was also killed. These hits provoked tremendous rage among Palestinians, and terror organizations swore to retaliate. Although attempted suicide bombings in Israel increased, Israel thwarted virtually all of these by making preventive arrests before suspected bombers could launch an attack or by intercepting would-be bombers en route. Indeed, the number of successful



suicide bombings in Israel during the first four months of 2004 is consistent with our model's predictions as calibrated through the end of 2003.

All of the elements of this story are present in our analysis: targeted killings are associated with increased suicide-bombing attempts; preventive arrests reduce the number of suicide-bombing attacks; and during periods when the expected number of attacks is high (such as the days and weeks following the Yassin and Rantissi hits), attacks are more spontaneous and less organized than usual, and consequently intelligence and security services on high alert for potential attacks are more likely to intercept suicide bombers.

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