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No. 04/10

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by

David Fielding and Anja Shortland

Centre for Research in Economic Development and International Trade, University of Nottingham

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The Authors

David Fielding is Professor, Department of Economics, University of Otago, Dunedin 9001, New Zealand (dfielding@business.otago.ac.nz) and External Fellow of CREDIT. Anja Shortland is a Lecturer, Department of Economics, University of Leicester, Leicester LE1 7RH (aks13@le.ac.uk).

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November 2004

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Abstract
Analysis of cross-sectional data on tourism to Israel during the *Intifada* period reveals some of the factors driving the behaviour of tourists from different countries. A large part of the heterogeneity in the observed response of different nationalities can be explained by socio-economic characteristics, some of which suggest differences in attitudes towards the risk associated with violence in Israel. Analysis of time-series data reveals the relative importance of different dimensions of violence in explaining the decline in tourism.

JEL classification: Z19, L83

Outline

1. INTRODUCTION

Israel and the Palestinian Territories are now one of several tourist locations severely affected by political violence. There is an increasing body of evidence to suggest that violent incidents resulting in only a handful of fatalities per year (and therefore representing only a very small risk to an individual tourist) have a substantial impact on tourist volumes and tourism revenues. Strong time-series evidence for such effects is reported in Enders and Sandler (1991) (Spain), Enders *et al.* (1992) (Austria, Italy and Greece), Drakos and Kutan (2003) (Greece, Israel and Turkey) and Sloboda (2003) (USA). Anecdotal evidence suggests that the effects of violence on tourism are equally large in developing country destinations such as Bali and Egypt.[1]

Whatever the true nature of the risk, many OECD governments actively dissuade their nationals from travelling to Israel. The following quotation from the US State Department website (August 3, 2004) is typical of advice given to Western tourists:

"The Department of State warns US citizens to… defer travel to Israel, the West Bank and Gaza due to current safety and security concerns."

Such violence has serious economic repercussions for a tourism destination like Israel, where since September 2000 there has been a marked increase in violent conflict between Israeli and Palestinian forces (the *Al-Aqsa Intifada*). Tourist arrivals are now less than half their pre-2000 level, and between 1999 and 2003 annual tourism revenue fell from \$4.3bn to \$2.3bn. This fall is almost equal in magnitude to the decline in the Israeli Balance of Payments in the same period, from a \$0.9bn surplus to a \$1.3bn deficit.

1 Frey *et al.* (2004) review the wider literature on tourism and political violence.

It is not at all surprising that the upsurge in violent conflict led to a dramatic fall in the number of tourists in Israel.[2] However, this simple statistic leaves many questions unanswered. Many people have chosen not to visit Israel any more, but a substantial minority has been undeterred by the violence. In this paper we will use time-series and cross-sectional data on tourism in Israel to explore the characteristics of these two groups of people. Along the way we will find out which dimensions of the violence affect tourists' choices, and whether variations in conflict intensity have more impact than variations the frequency of road traffic accidents. We will also explore the factors that drive the differences we will observe in the behaviour of tourists from different parts of the world. For example, some commentators insist that there are still large cultural differences between Americans and Europeans with respect to risk-taking. In the words of one Whitehouse spokesperson:[3]

"An American personality… prizes the calculated risk… Europeans often seem bent on preventing any chance of trouble arising."

If this is so, then *ceteris paribus* we should observe Americans to be less deterred than Europeans by the dangers of international tourism, and more inclined to ignore the advice of the State Department. Before we discuss our data and our model, the next section of the paper outlines in more detail the conceptual framework for the paper.

2. A CONCEPTUAL FRAMEWORK

In this paper we will use two slightly different Israeli datasets to address two key questions about the behaviour of tourists. First of all, we can ask a question about the distribution of attitudes towards the risk of death or injury faced by travellers. The fact that some – but not all – tourists are staying away from Israel these days suggests heterogeneous attitudes towards risk among the tourist population. For some – but not all – tourists the higher risk appears to

2 There is substantial evidence from studies of individual respondents that people's response to the risk of injury in a violent political conflict does not square with Expected Utility Theory, and that they place "excessive" weight on highly improbable states of the world with very low utility. Sunstein (2003) and Viscusi and Zeckhauser (2003) find that people assessing conflict risk are prone to deviations from EUT common in other risk perception contexts; so their behaviour might be better explained by, for example, Prospect Theory.

have raised the opportunity cost of travelling to Israel above the benefit. The fall in tourist numbers is consistent with the existence of discrete groups of people with different attitudes towards risk. Figure 1a illustrates this case. (Figure 1 represents a rough sketch of a model that will be outlined in much greater detail in section 3 below.) The frequency distribution in the figure indicates the number of people, g, who are just indifferent between travelling and not travelling at a certain level of risk, z. (Because the number of tourists is declining in the level of risk, we draw g as a function of $1/z$.) The fact that g is bimodal reflects the existence of two groups of people, a "timid" group clustered around the right-hand mode, and a "gung-ho" group clustered around the left-hand mode. The number of tourists will be the integral of g up to the current level of $1/z$. As the risk level rises from z_1 to z_2 between September and October 2000, the timid group drops out of the travelling population. Subsequently, small variations in the level of risk around z_2 have little or no impact on tourist numbers. However, the fall in tourist numbers is also consistent with a unimodal distribution, as illustrated in Figure 1b. In this case, there are no identifiable clusters with respect to attitudes towards risk. The rise in risk from z_1 to z_2 again leads to a substantial reduction in tourist numbers, but in this case subsequent variations around z_2 do have a substantial impact on tourist numbers.

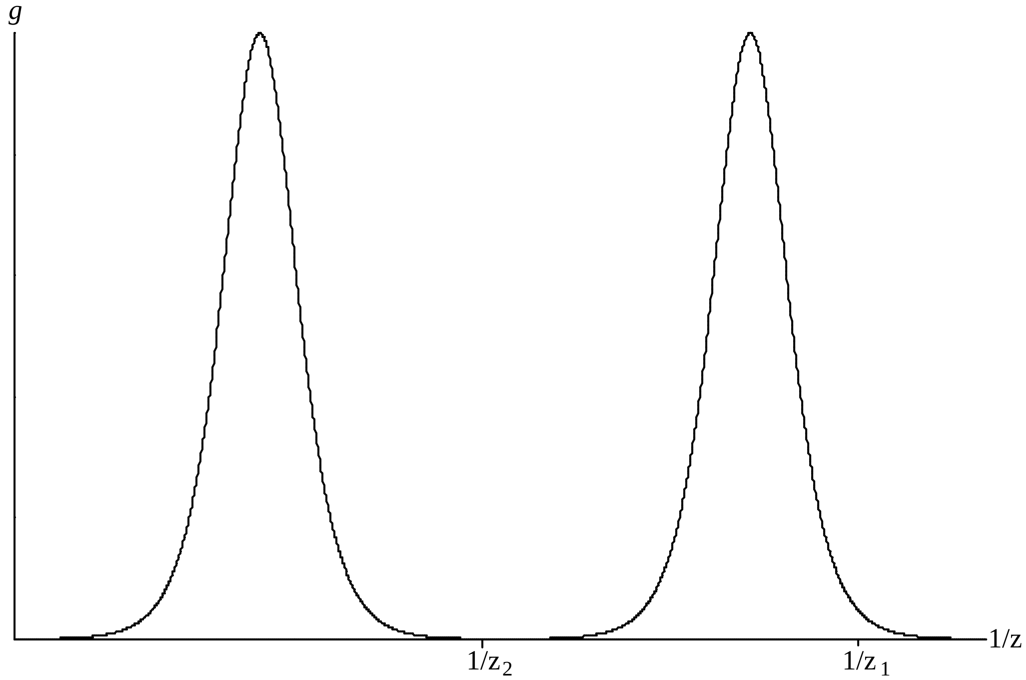


Figure 1a

3 The quotation is from a speech by M. Daniels, Office of Management and Budget, The Executive Office of the

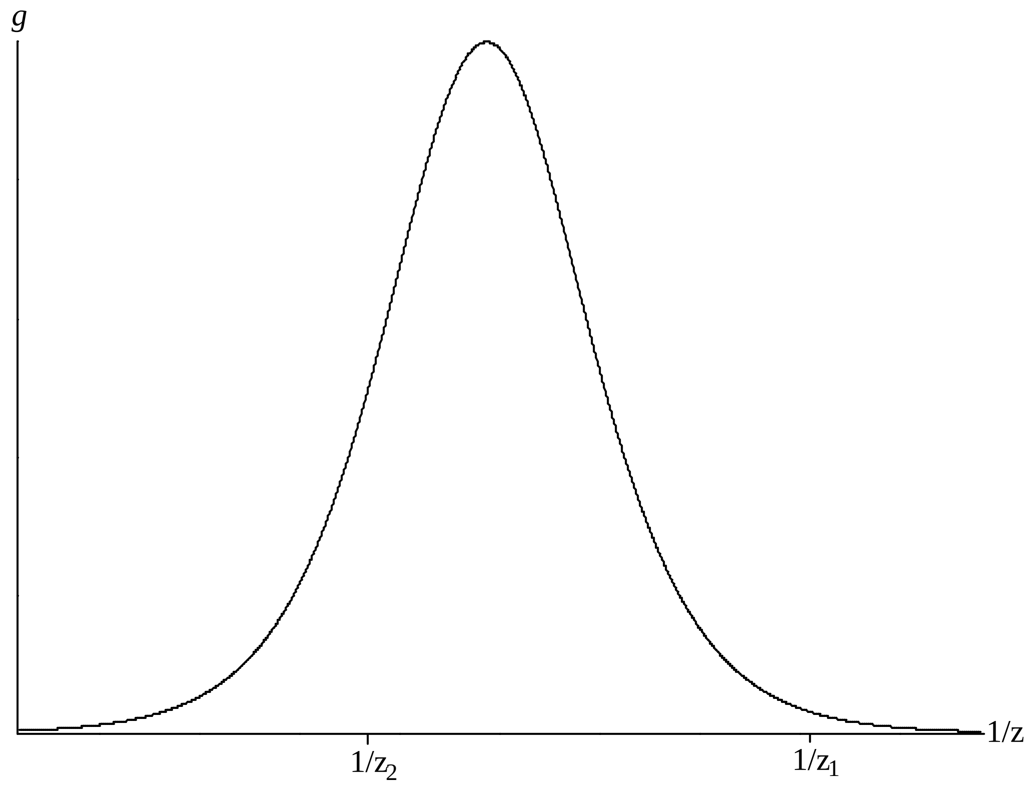


Figure 1b

Why does this matter? One reason is that the shape of g around z_2 affects the return to marginal improvements in the Israeli-Palestinian peace process. In Figure 1a nothing short of a complete return to peace will have any substantial impact on tourist numbers. Piecemeal measures that result in a partial reduction in violence cannot realistically be sold to the Israeli or Palestinian public on economic grounds. This makes a gradual return to normality very difficult. By contrast, in Figure 1b even a small reduction in violence yields an economic return, making partial peace agreements easier to sell to the public, and facilitating a gradual return to peace.

The second question relates to differences in the attitudes of tourists of different nationalities. The distributions in Figure 1 might vary from one part of the world to another, because of variations in the (net) benefits to the average tourist from visiting Israel, or because of variations in the costs associated with a certain level of risk. For example, the benefits might be higher in countries with a large Jewish population. The costs associated with risk might be lower in countries where people have learned better how to manage risk, or else have become

President, May 16 2003 (http://www.whitehouse.gov/omb/speeches/daniels051603.html).

acclimatised to it. People in some places might just be more gung-ho than people elsewhere. If there is cross-country variation in the size of the integral of g between z_1 and z_2, then an increase in violence will have markedly different effects on tourist numbers from different parts of the world.

If we can find correlates of the national characteristics that affect the shape of g, then we will be able to explain at least some of the cross-country variation in the decline in tourist numbers. As Table 5 below indicates, this variation has been substantial. This will provide evidence on some of the ways in which national characteristics affect attitudes towards risk and security.

In order to address the issues raised in the first question, we need to look at tourists' response to changes in the level of violence in Israel *after* September 2000, to see whether the relatively small fluctuations in conflict intensity during the *Intifada* have been associated with changes in tourist volumes. This requires the analysis of time-series data on tourist traffic. The Israeli Central Bureau of Statistics (CBS) reports consistent monthly data on the number of American tourists and on the number of European tourists checking into Israeli hotels each month. (Data on tourist numbers in the West Bank and Gaza, virtually zero since September 2000, are not included.) In the next section, we will outline a time-series model that is designed to explain variations in these data. If the month-on-month variations in tourist volumes in response to fluctuations in conflict intensity are substantial, relative to the large decline in tourism as a result of the start of the *Intifada*, then we are likely to be in the world of Figure 1b rather than that of Figure 1a. Small steps towards peace will yield an economic return, making a gradual return to normality more likely.

The hotels data are not well suited to answering the second question, because they disaggregate only between Americans, Europeans, and others. However, there are also annual CBS data on tourist arrivals into Israel, disaggregated by the nationality of the individual tourists. These data are not reported at a high enough frequency for time-series analysis, but we can construct a cross-section in which the dependent variable is the rate of decline in tourist arrivals from each country between 1998-9 (i.e., before the start of the *Intifada*) and

2001-2.[4] We can then look at the national characteristics associated with cross-sectional variations in the rate of decline. Section 3 below first outlines the modelling framework used to analyse the time-series data, then presents the results of our analysis. Section 4 deals similarly with the international cross-sectional data.

3. THE TIME-SERIES MODEL

3.1 The time-series data: concepts

Our time-series regression equations ought to be consistent with a plausible model of individual decision-making. In this section we expand on the ideas outlined in section 2, deriving a regression equation from the discrete choice theory outlined *inter alia* in Maddala (1983).[5]

The model concerns a population of people who have already decided to take a vacation, and are deciding where to go. Let the net utility an individual i derives from taking a vacation in location $m \in \{1,..., M\}$ in month t be designated v_{imt}. We will assume that each person's utility is of the form:

$$v_{imt} = m_{mt}(\boldsymbol{X}_{\boldsymbol{m}t}, e_{mt}) + u_{imt} \tag{1}$$

where m_{mt} is the average level of utility from visiting location m in month t for the vacationing population and u_{imt} is an individual's idiosyncratic deviation from this average. $\boldsymbol{X}_{mt}$ is a vector of identifiable time-varying factors that impact on one's net utility from a vacation in a particular location, and e_{mt} is a stochastic term reflecting the unpredictable component of the

4 The number of tourist arrivals is a little higher than the number checking into hotels, because some tourists do not stay in hotels; for example, some stay with friends or family. However, data from the two sources – hotels and immigration – are broadly consistent. Some monthly tourist arrival statistics reported by immigration are published in the CBS *Monthly Bulletin of Statistics*, but only for selected months.

5 The regression specification we end up with is similar in spirit to that of Fleischer and Buccola (2002), who analyze total foreign demand for Israeli hotel accommodation up to 1999, but differs from theirs in points of detail. They do not formulate an explicit discrete choice model, and do not disaggregate foreign hotel guests by nationality. They condition demand for hotel beds on a single lagged "terror index", and on foreign income and tourist expenditure outside Israel (rather than tourist volumes outside Israel). We contend that it is more appropriate to use tourist volumes outside Israel as a scale variable when modelling tourist volumes inside Israel. They also condition on Israeli hotel prices, using Israeli hotel wages as an instrument. We are not sure

average utility level (fads and fashions). We further assume that individual i chooses location m in period t if and only if:

$$v_{imt} = \max(v_{i1t}, \ldots, v_{iMt}) \tag{2}$$

It can be shown (Maddala, 1983) that if for any two locations (m, n) the distribution of u_{imt} is independent of that of u_{int}, and if each has a Weibull distribution, then the probability of any one randomly selected individual choosing location m in period t is:

$$p_{imt} = \exp(m_{mt}) \Big/ \sum_{j=1}^{j=M} \exp(m_{jt}) \tag{3}$$

r, in logarithmic form:

$$\ln(p_{imt}) = m_{mt} - \ln \sum_{j=1}^{j=M} \exp(m_{jt}) \tag{4}$$

(If the only factor influencing the m's were the level of risk, z_{mt}, associated with travel to location m, then our model would be as simple as the one depicted in Figure 1, with $g_{mt} = \mathrm{d}p_{imt}/\mathrm{d}(1/z_{mt})$. However, our model will not be so restrictive.) For a large population, the ratio of the number of people in period t visiting location m (p_{mt}) to the number visiting location n (p_{nt}) can therefore be written as:

$$p_{mt} / p_{nt} = \exp(m_{mt}) / \exp(m_{nt}) \tag{5}$$

and hence:

$$\ln(p_{mt}) - \ln(p_{nt}) = m_{mt}(\boldsymbol{X_{mt}}, e_{mt}) - m_{nt}(\boldsymbol{X_{nt}}, e_{nt}) \tag{6}$$

Location m here is to be interpreted as Israel; the identity of the reference location n will be discussed later. If we know the functional forms of m_{mt} (.) and m_{nt} (.), then we can fit equation (6) to time-series data. In what follows, we assume that for the data *after* September 2000 it is possible to find a linear specification such that:

$$\ln(p_{mt}) - \ln(p_{nt}) = [\boldsymbol{X_{mt}} - \boldsymbol{X_{nt}}]'b + e_t \tag{7}$$

that wages are really exogenous to the demand for hotel beds and anyway, as documented below, we find prices to be statistically insignificant in the post-2000 period.

where e_t is a linear function of e_{mt} and e_{nt}. (Note that we are not assuming linearity across the large change in conflict intensity following the onset of the *Intifada*, only linearity in the smaller changes observed since.)

We will begin with the assumption that $[\boldsymbol{X}_{mt} - \boldsymbol{X}_{nt}]$ had two major components: the anticipated relative enjoyability of the two locations and the relative chances of being a victim of a violent incident in the two locations this month (also the anticipated relative chances in the next month, since some tourist visits might straddle two consecutive months). So our regressions are based on an equation of the form:

$$\ln(p_{mt}) - \ln(p_{nt}) = b_1.E[\ln(w_{mt}) - \ln(w_{nt})] \tag{8}$$

$$+ b_3.[\ln(z_{mt}) - \ln(z_{nt})] + b_4.E[\ln(z_{mt+1}) - \ln(z_{nt+1})] + e_t$$

where w_{mt} is the enjoyability of location m in period t and z_{mt} is the probability of being a victim of a violent incident. An expectations operator is attached to w_{mt} because many tourists are likely to be first-time visitors who aren't sure whether they are really going to like the place they are staying until they get there. One might also wonder whether monthly variations in the relative cost of different locations make a difference to tourist numbers. However, as discussed in Appendix 1, our empirical measures of relative cost were never statistically significant in any regression equation. It seems that between 2000 and 2003, monthly variations in cost had no substantial impact on tourism to Israel.

Application of the model requires us to specify the expectations formation process. We will work with the following assumptions:

$$E[\ln(w_{mt}) - \ln(w_{nt})] = a_1(L)[\ln(p_{mt-1}) - \ln(p_{nt-1})] \tag{9a}$$

$$E[\ln(z_{mt+1}) - \ln(z_{nt+1})] = a_3(L)[\ln(z_{mt}) - \ln(z_{nt})] + a_4(L)[\ln(f_{mt}) - \ln(f_{nt})] \tag{9b}$$

where the a(L)s are lag polynomial operators. Equation (9a) builds some herding behavior into the model: if a destination has been popular in the past, people are more likely to consider it today.[6] Equation (9b) states that expectations about the future risk of violence are based on the past and current frequency of and current violent incidents. It also allows for extra dimensions of conflict intensity, other than the direct risk to tourists, to be used in predicting the future risk. These dimensions are captured by the variable f_{jt}. Substituting equations (9a)-9(b) into equation (8), we will have an ARDL equation of the form:

$$g_1(L)[\ln(p_{mt}) - \ln(p_{nt})] = g_2(L)[\ln(z_{mt}) - \ln(z_{nt})] + g_3(L)[\ln(f_{mt}) - \ln(f_{nt})] + e_t \qquad (10)$$

where the g(L)s are linear combinations of the a(L)s and the b's.

It is worth reiterating that the g_2 and g_3 parameters in equation (10) will be statistically significant only if in-sample variations in the level of violence are making the difference to the vacation location decisions of a substantial number of tourists. This

will be the case as long as there are a substantial number of people who are roughly indifferent between locations at the average level of violence *during* the *Intifada* period.

3.2 The time-series data: application

We first discuss the measurement of the variables in equation (10). The equation is to be fitted to the monthly hotels data for two tourist populations: tourists from America and tourists from Europe. In order to estimate the parameters of the equation, we need to construct a dependent variable in which the number of visitors to Israel from a certain population (America, Europe) is expressed relative to the number of visitors from that population to other locations. In order to focus on the effects of political violence within Israel, it will be convenient to use reference locations that are reasonably safe. In the case of American tourists the reference location will be Europe,[7] and in the case of Europeans it will be America. That is, for the American tourist sample, p_{mt} is interpreted as the number of American visitors to Israel in a particular month and p_{nt} is interpreted as the number of American visitors to Europe. For the European tourist sample, p_{mt} is interpreted as the

6 There is likely also to be some seasonality in w. Such seasonality should be taken as implicit in equations (8-10), and is accounted for in the regressions in section 3.2.

7 Not everywhere in Europe is safe, but most of the places we see American tourists are pretty quiet.

number of European visitors to Israel in a particular month and p_{nt} is interpreted as the number of European visitors to America.

For both samples, the monthly transatlantic tourism figures used to measure p_{nt} are taken from the dataset published by the ITA Office of Tourism and Travel Industries (http://tinet.ita.doc.gov), which reports both American tourists departing to Europe and European tourists arriving in America. The monthly Israeli tourism data are published by the Central Bureau of Statistics and are available online at http://www.cbs.gov.il. The Israeli tourism statistics used to measure p_{mt} are those for American and European tourists checking into tourist hotels.[8]

Measures of z_{mt} and f_{mt} are constructed from data provided by the Israeli NGO B'Tselem (http://www.btselem.org). Among other *Intifada*-related data, B'Tselem records: (i) the total number of Israeli fatalities within Israel proper, excluding the West Bank and Gaza (WBG); (ii) the total number of Israeli and Palestinian fatalities in WBG. The first of these series is used as a measure of z_{mt}. All – or almost all – fatalities within Israel proper, most of which are from bomb attacks, are in situations in which tourists are just as likely to be victims as Israeli residents. Of course, the vast majority of victims are Israelis, because the resident population is so much larger than the tourist population. The second series is used as a measure of f_{mt}. Violence in WBG does not pose a direct risk to tourists to Israel, who don't have to go there; but it could be used to forecast future levels of violence within Israel, if some of the violence in WBG "spills over" the Green Line. (Appendix 2 discusses the actual correlation between f_{mt} and z_{mt+1}.) It is worth noting at this point that disaggregation of WBG fatalities by nationality and by combatant status has no extra explanatory power in the regression equations reported in the next section.

8 The g_2 and g_3 parameters still need to be interpreted with some caution. For example, some "solidarity tourism" trips make use of tourist hotels. We are not suggesting that the responses of individuals to changes in the level of violence are entirely homogenous, and the parameters are to be interpreted as an estimate of the average response across the tourist population.

In addition to the fatality series in Israel, we will include a dummy variable (*DGW*) for the month of the Iraq War (2003m3). American and European tourists are likely to have thought travel in the Mid East to be more risky during the war, or at least during the first couple of weeks. The use of dummy variables in time-series regressions is never ideal: the interpretation of dummy coefficients is always open to question. But we have no other way of capturing the effect of the Gulf War, and removing the dummy from the regression does not substantially alter our estimated long-run elasticities.[9]

In all cases we assume that $z_{nt} = f_{nt} = 0$: there is no political violence in Europe or America. Our sample period does encompass September 2001; but the attacks in America made all overseas air travel more daunting for Americans, regardless of their destination. It also seems to have been perceived in Europe as increasing the risk of air travel generally, rather than air travel specifically to America. In any case, a dummy variable for 2001m9 is not statistically significant in the regression equations reported below.

Figures 2-3 depict the time-series $[\ln(p_{mt}) - \ln(p_{nt})]$, $\ln(z_{mt})$ and $\ln(f_{mt})$ in each of our two samples. Table 1 provides some descriptive statistics for the variables for our sample period (2000m9-2004m2, or 42 observations), as well as for the period before the onset of the *Intifada*. The table shows how much the mean values of both tourism and conflict intensity have changed: the *Intifada* represents a large structural break. As in many other empirical applications, it is unclear whether the variables are I(0) or I(1) over the 2000m9-2004m2 sample period: standard tests reject neither null at conventional levels of significance. So it is appropriate to re-parameterize equation (10) in error-correction form. Employing such a re-parameterization of equation (10) with the restriction that $z_{nt} = f_{nt} = 0$, and also allowing for the Gulf War dummy, we have:

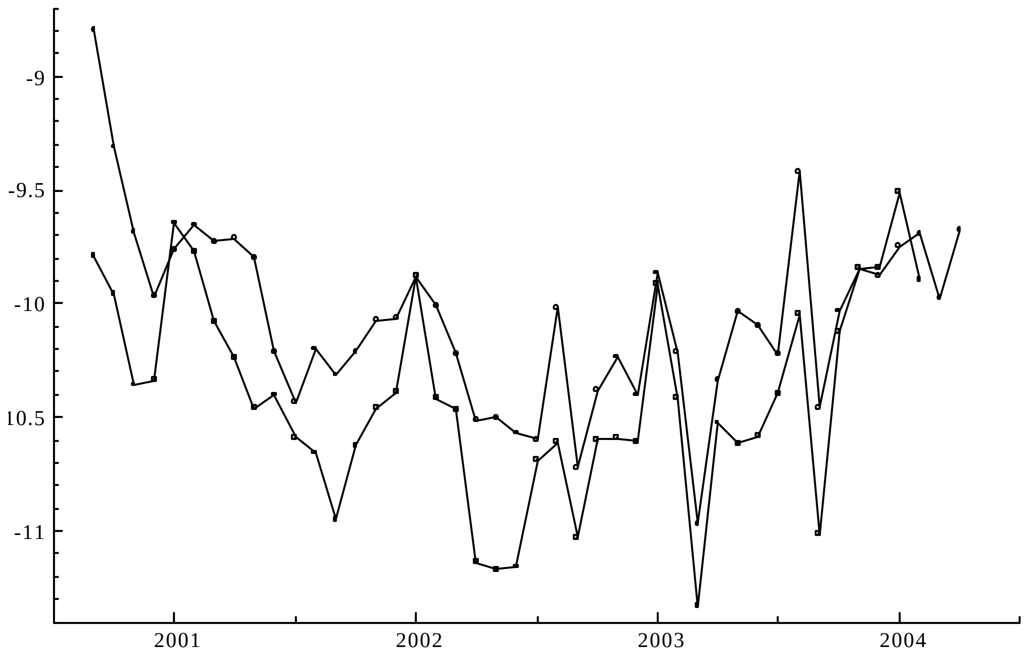
$$h_1(L)\Delta[\ln(p_{mt}) - \ln(p_{nt})] = h_2(L)\Delta\ln(z_{mt}) + h_3(L)\Delta\ln(f_{mt}) + j_1.[\ln(p_{mt-1}) - \ln(p_{nt-1})] + j_2.\ln(z_{mt-1}) + j_3.\ln(f_{mt-1}) + d.DGW_t + e_t \quad (11)$$

9 It does affect our estimated short-run dynamics. Further results are available on request.

where the j 's capture the long-run levels relationship between the variables.[10] In both the American and the European sample, the lag order used to fit equation (11) is 1. This lag order minimizes both the Schwartz-Bayesian and Akaike information criteria for the respective regressions equations. Pesaran *et al.* (2001) provide critical values for the F-statistic for the null that $\forall x(j_x = 0)$ under (i) the assumption that all variables are I(0) and (ii) the assumption that all variables are I(1). If the null can be rejected in both cases, then there is evidence that there is a long-run relationship between the variables.

Table 1: Sample Statistics 3½ Years Before and After the Al-Aqsa Intifada

	1997m3-2000m8		*2000m9-2004m2*		
	Mean	s.d.	Mean	s.d.	difference
In(P_{mt}/P_{nt}) (America)	-9.4967	0.2758	-10.4073	0.4416	-0.9106
In (P_{mt}/P_{nt}) (Europe)	-9.0252	0.3087	-10.0658	0.3993	-1.0406
In (f_{mt})	0.6885	0.7413	4.1661	0.6565	3.4776
In (z_{mt})	0.2142	0.5525	1.5756	1.2755	1.3614



10 The validity of this approach relies on the existence of a single levels relationship. Appendix 2 shows that there is no levels relationship between ln(z) and ln(f) as we measure them.

Figure 2:Tourism series In(p_{mt}/p_{nt})2000m9-2004m2:Americans (■) and Europeans (?)

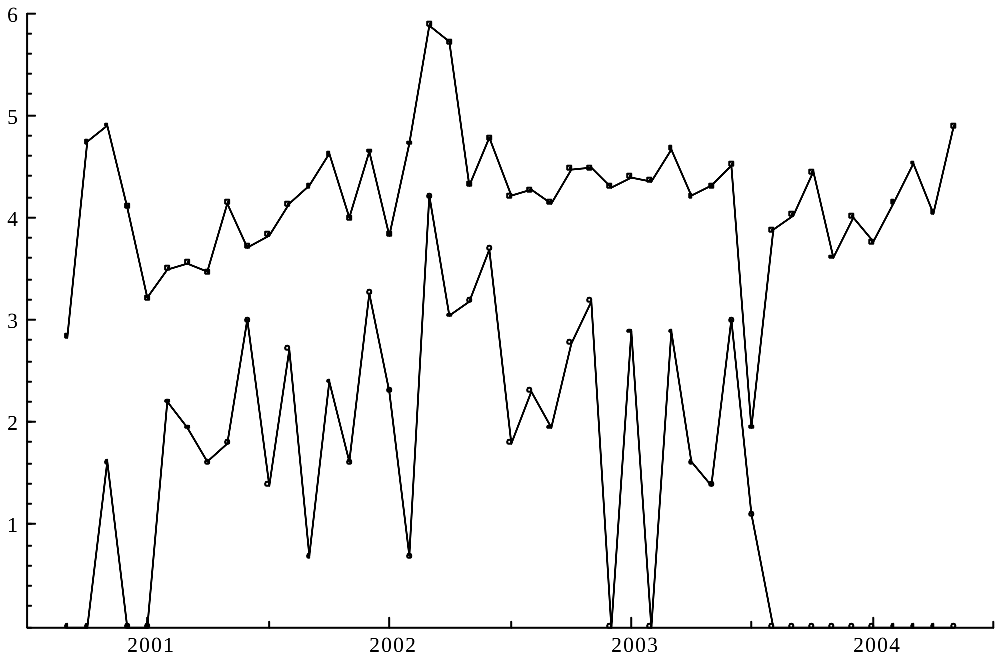


Figure 3: Fatality series 2000m9-2004m2:In(z_{mt})(?) and In(f_{mt})(■)

Table 2 reports the regression results for the American tourist sample and Table 3 reports the regression results for the European tourist sample.[11] In both cases there are two regression equations, an unrestricted one corresponding to equation (11), and a second equation that omits insignificant components of the dynamics. Note that there is a seasonally varying intercept in each equation. T-ratios on lagged level parameters (the j 's) should be treated with some caution, since the variables might not be stationary. However, the F-statistics for the joint significance of the j 's are always greater then the upper bounds reported in Pesaran *et al*. (2001), so there does seem to be a statistically significant long-run relationship between the variables. Recursive estimation of the equations over the last 12 months of the sample suggests that the key

11 These are OLS estimates; FIML estimates allowing for the non-zero equation residual correlations are very similar.

model parameters are stable over time, and that the results are robust to changes in sample size (Figures 4-5 depict the one-step ahead forecast errors for each equation, and Figure 6-7 depict the recursively estimated j coefficients.).

Rather than discussing individual coefficients in the dynamic regression equations, we will base our comments on the results depicted in Table 4 and Figures 8-9. Table 4 indicates the long-run coefficients on each of the explanatory variables implicit in the unrestricted equations in Tables 2-3. Figures 8-9 summarize the dynamics of these regression equations by plotting the response of the dependent variables to a temporary (one-month) unit increase in $\ln(z_{mt})$ and $\ln(f_{mt})$ over a six month-period.

Table 4 shows that a 1% rise in Israeli fatalities, as captured by z_{mt}, will, if sustained, result in a 0.31% decline in American tourists and a 0.44% decline in European tourists; the difference between these figures is statistically insignificant. It also shows that a 1% rise in WBG fatalities, as captured by f_{mt}, will, if sustained, result in a 0.17% decline in American tourists and a 0.07% decline in European tourists; the difference between these figures is statistically insignificant. The coefficients on the *DGW* Gulf War dummy are -0.99 and -1.23 respectively; again, the difference between these two numbers is statistically insignificant. So both American and European tourist numbers are sensitive to monthly variations in the magnitude of violence. This suggests that there are substantial numbers of people – both in America and in Europe – who are approximately indifferent between travelling to Israel and not, even at post-2000 levels of violence. The distribution in Figure 1b appears to be more appropriate here than the one in Figure 1a. The population (at least, the population of tourists) is not divided into discrete “gung-ho” and “timid” groups. As a consequence, small variations in conflict intensity do have a substantial impact on tourism.

Figures 8-9 show that the response to a change in conflict intensity is remarkably swift. For both Americans and Europeans, and for both dimensions of the conflict, the decline in tourism in response to a temporary (one-month) intensification of the conflict appears in the first three months following. From month 4 onwards, tourism numbers recover to their initial level.

It appears from the time-series data that there is no substantial difference between American and European responses to changes in the level of risk associated with visiting Israel since September 2000; certainly, there is no evidence that Europeans are more risk-averse. However, the cross-sectional data discussed in the next section will provide more detailed evidence on international differences in responses to the Israeli-Palestinian conflict.

Table 2: American Tourist Time-Series Regression Results

Standard errors are calculated using White's heteroskedasticity correction.

A. Unrestricted equation for $\Delta ln(p_{mt} / p_{nt})$

Variable	***coefficient***	***standard error***	***t ratio***	***partial R^2***
$\Delta\ln(p_{mt-1}/p_{nt-1})$	0.02831	0.08276	0.34207	0.0050
$\Delta\ln(z_{mt})$	-0.12125	0.04538	-2.67188	0.2349
$\Delta\ln(z_{mt-1})$	0.00951	0.05351	0.17772	0.0021
$\Delta\ln(f_{mt})$	-0.01666	0.02165	-0.76952	0.0179
$\Delta\ln(f_{mt-1})$	0.06924	0.02183	3.17178	0.2007
$\ln(p_{mt-1}/p_{nt-1})$	-1.04140	0.10229	-10.18090	0.7547
$\ln(z_{mt-1})$	-0.32112	0.04594	-6.98999	0.5640
$\ln(f_{mt-1})$	-0.17344	0.03714	-4.66990	0.5006
DGW_t	-1.03030	0.12059	-8.54383	0.6069

$R^2 = 0.94728$; $\sigma = 0.13469$
LM residual autocorrelation test: F(1,20) = 0.62967 [0.4368]
LM ARCH test: F(1,19) = 0.00001 [0.9931]
Residual normality test: $\chi^2(2) = 2.2101$ [0.3312]
F-statistic for joint significance of levels variables: F(3,21) = 22.768

B. Restricted equation for $\Delta ln(p_{mt} / p_{nt})$

Variable	***coefficient***	***standard error***	***t ratio***	***partial R^2***
$\Delta\ln(z_{mt})$	-0.12484	0.04101	-3.04414	0.2581
$\Delta\ln(f_{mt-1})$	0.07492	0.01693	4.42528	0.2685
$\ln(p_{mt-1}/p_{nt-1})$	-1.00750	0.07259	-13.87930	0.8475
$\ln(z_{mt-1})$	-0.31863	0.03648	-8.73438	0.6889
$\ln(f_{mt-1})$	-0.16160	0.02390	-6.76151	0.5190
DGW_t	-1.01760	0.10459	-9.72942	0.6011

$R^2 = 0.94601$; $\sigma = 0.12751$
LM residual autocorrelation test: F(1,23) = 0.52302 [0.4768]
LM ARCH test: F(1,22) = 0.19195 [0.6656]
Normality test: $\chi^2(2) = 1.0137$ [0.6024]

p_{mt} / p_{nt}	ratio of US tourists in Israel to US tourists in Europe
f_{mt}	1 + total Israeli and Palestinian fatalities in West Bank & Gaza
z_{mt}	1 + total fatalities in Israel
DGW_t	dummy variable = 1 in 2003m3, = 0 else

Table 3: European Tourist Time-Series Regression Results

Standard errors are calculated using White's heteroskedasticity correction.

A. Unrestricted equation for $\Delta ln(p_{mt} / p_{nt})$

variable	***coefficient***	***standard error***	***t ratio***	***partial R^2***
$\Delta\ln(p_{mt-1}/p_{nt-1})$	-0.10059	0.08268	-1.21662	0.0475
$\Delta\ln(z_{mt})$	-0.04037	0.05112	-0.78971	0.0248
$\Delta\ln(z_{mt-1})$	-0.00547	0.06146	-0.08900	0.0004
$\Delta\ln(f_{mt})$	0.00093	0.02454	0.03790	0.0000
$\Delta\ln(f_{mt-1})$	0.02278	0.02520	0.90397	0.0199
$\ln(p_{mt-1}/p_{nt-1})$	-0.68139	0.09645	-7.06470	0.5759
$\ln(z_{mt-1})$	-0.29780	0.06944	-4.28859	0.4544
$\ln(f_{mt-1})$	-0.04624	0.03411	-1.35561	0.0571

DGW_t	-0.84095	0.07174	-11.72220	0.4305

$R^2 = 0.90854$; $\sigma = 0.15492$
LM residual autocorrelation test: F(1,20) = 0.64525 [0.4313]
LM ARCH test: F(1,19) = 1.05240 [0.3178]
Normality test: $\chi^2(2) = 0.08624$ [0.9578]
F-statistic for joint significance of levels variables: F(3,21) = 10.395

B. Restricted equation for $\Delta ln(p_{mt}/p_{nt})$

variable	***coefficient***	***standard error***	***t ratio***	***partial R^2***
$\ln(p_{mt-1}/p_{nt-1})$	-0.69345	0.05716	-12.1317	0.75
$\ln(z_{mt-1})$	-0.27551	0.04857	-5.67243	0.6834
$\ln(f_{mt-1})$	-0.03583	0.02657	-1.34851	0.0626
DGW_t	-0.85763	0.06402	-13.3963	0.4350

$R^2 = 0.89462$; $\sigma = 0.14945$
LM residual autocorrelation test: F(1,25) = 0.00055 [0.9815]
LM ARCH test: F(1,24) = 0.01799 [0.8944]
Normality test: $\chi^2(2) = 0.23485$ [0.8892]

p_{mt}/p_{nt} ratio of Euro tourists in Israel to Euro tourists in the US
f_{mt} 1 + total Israeli and Palestinian fatalities in West Bank & Gaza
z_{mt} 1 + total fatalities in Israel
DGW_t dummy variable = 1 in 2003m3, = 0 else

Table 4: Long-Run Levels Elasticities in the Unrestricted Models

	$\ln(z_m)$	$\ln(f_m)$	*DGW*
US equation	-0.30835	-0.16654	-0.98934
Standard error	*0.04838*	*0.03255*	*0.20792*
European equation	-0.43705	-0.06787	-1.23420
Standard error	*0.09012*	*0.05803*	*0.36834*

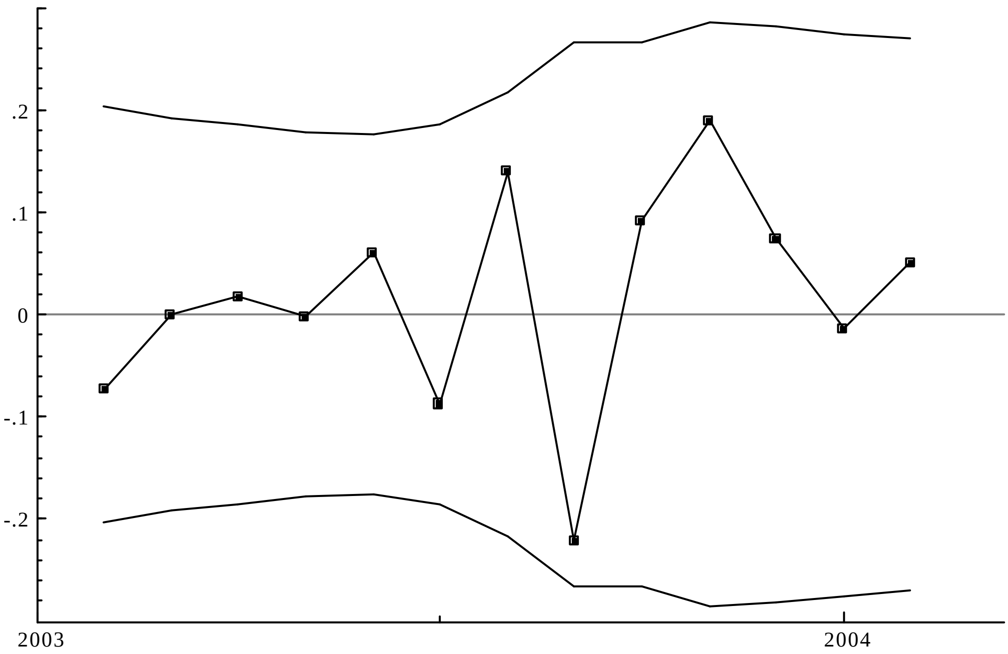


Figure 4:One-step American sample forecast errors with 2 s.e.bars (2003m2-004m2)

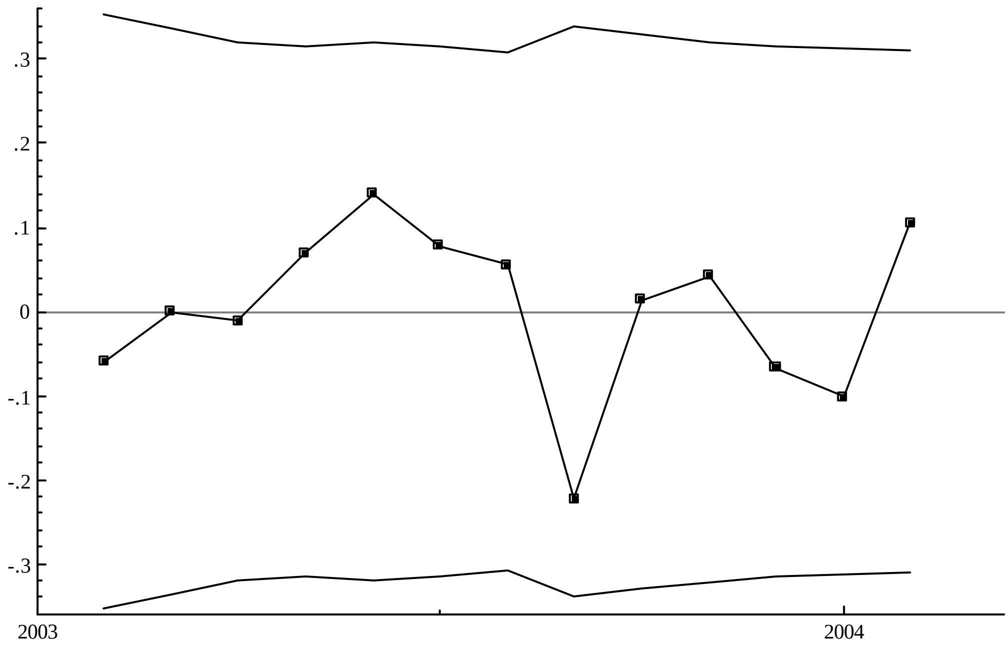


Figure 5: One-step European sample forecast errors with 2 s.e. bars (2003m2-2004m2)

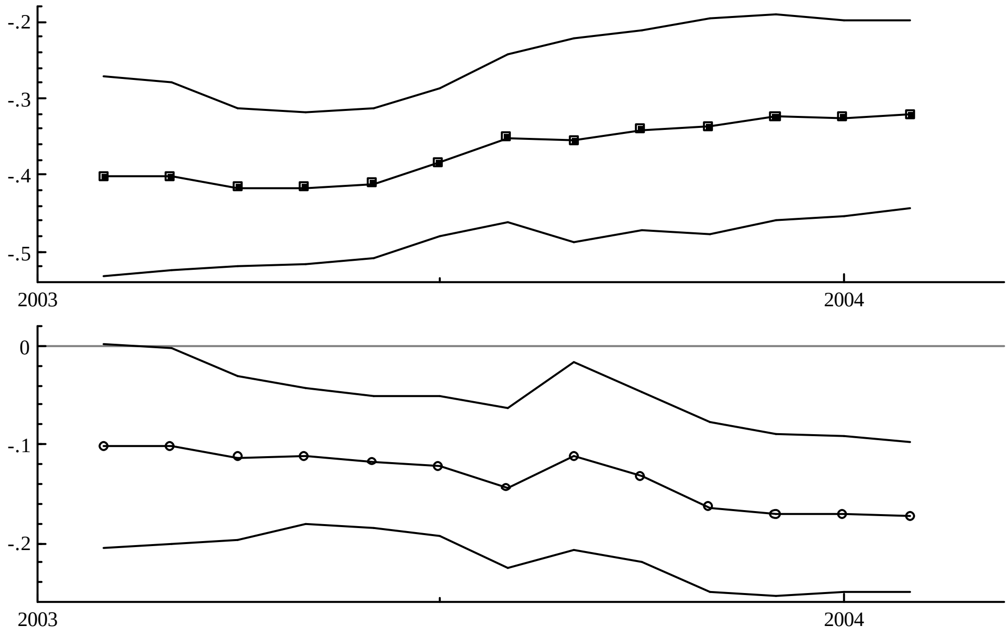


Figure 6: Recursive estimates of the In($z_{mt\text{-}I}$)(■)and In($f_{mt\text{-}t}$)(?)coefficients, American sample

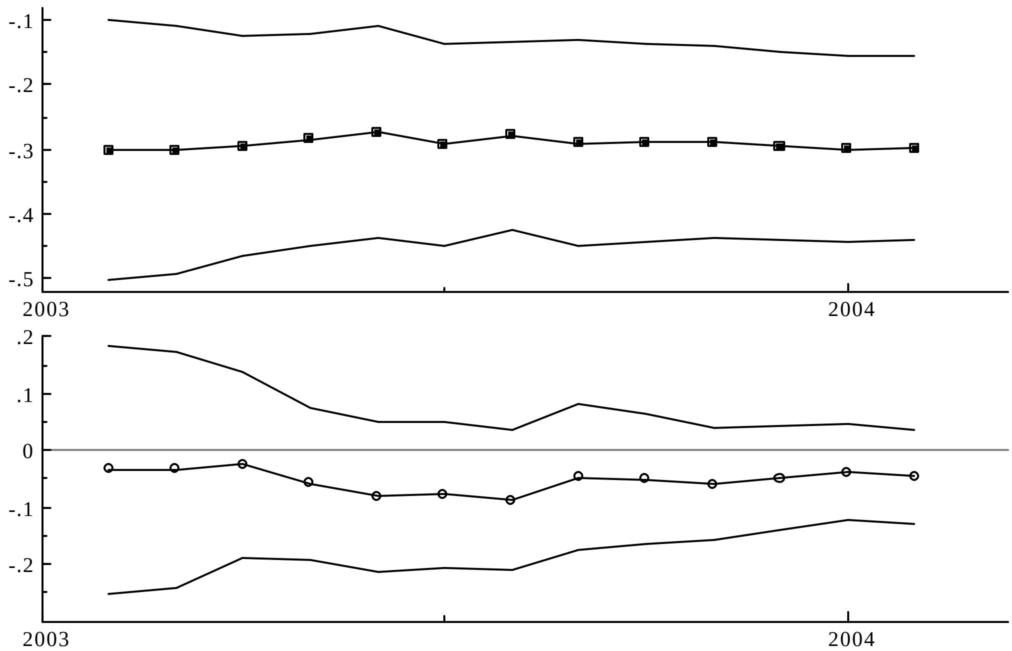


Figure 7: Recursive estimates of the In(z_{mt}-*1*)(■) and In($f_{mt\text{-}l}$)(?)coefficients, European sample

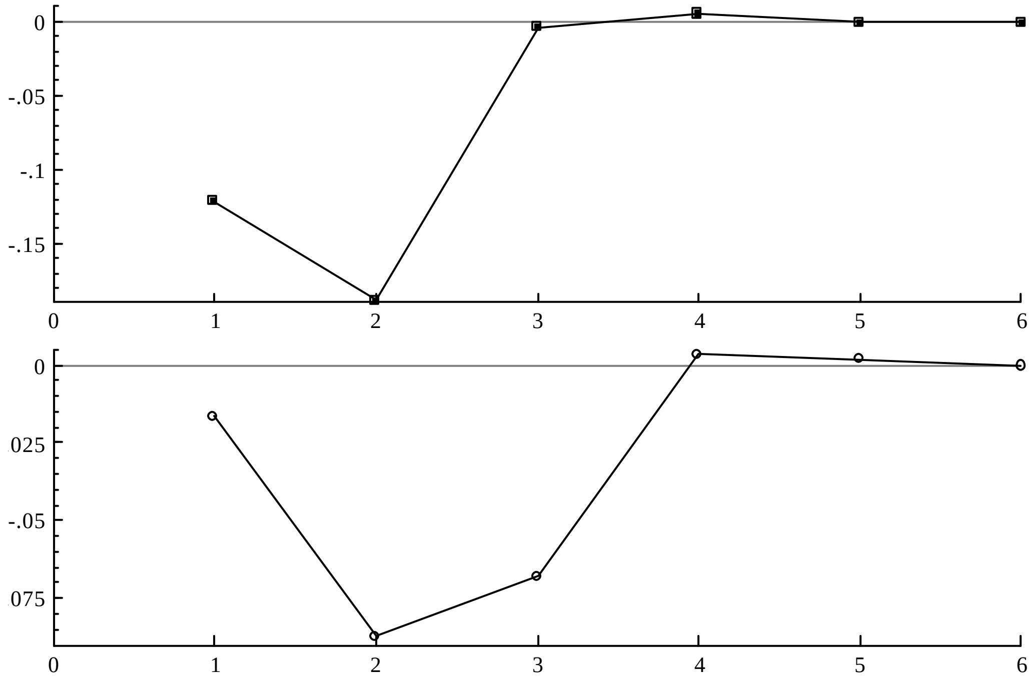


Figure 8: Response of In(p_{mt}/p_{nt}) to a unit increase in In(f_{mt})(?), American sample

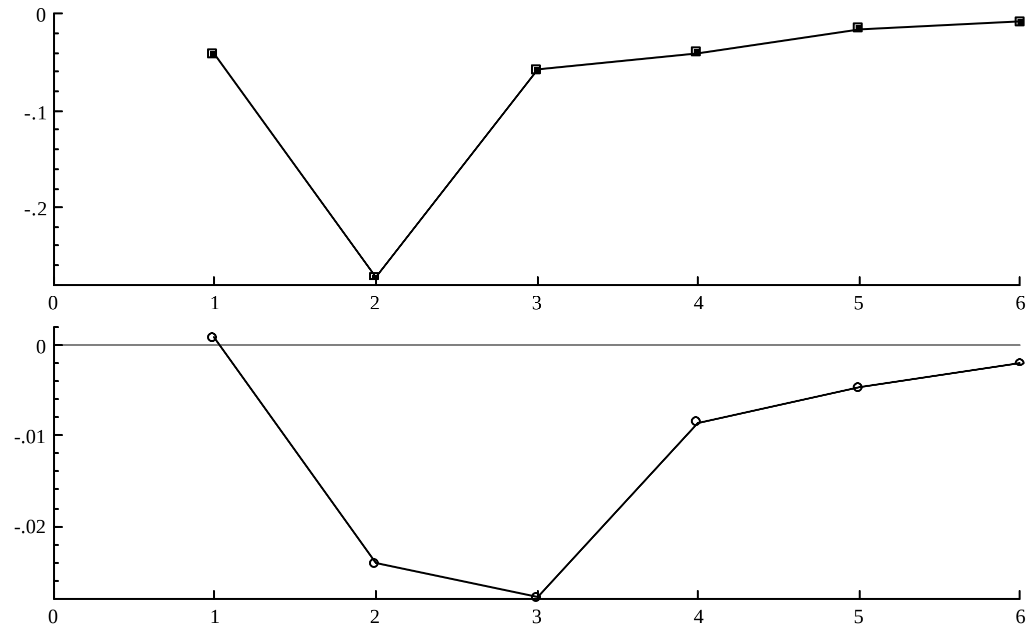


Figure 9: Response of In(p_{mt}/p_{nt}) to a unit increase in ln(z_{mt})(■) and In(f_{mt})(?), European sample

One often-quoted statistic in relation to the *Intifada* is that fewer people die in the conflict than in road-traffic accidents. In fact, if we look at the monthly CBS accident data, we see that there are indeed fewer conflict deaths *in Israel*, but there are more conflict deaths in WBG. Moreover, the variance of the log of road-traffic fatalities since September 2000 is less than half the variance of either of our conflict measures. In addition, there is no significant autocorrelation in deseasonalized monthly road fatalities, so last month's fatalities do not provide any extra information about how dangerous Israeli roads are this month (whereas there is some information in past levels of violence, as outlined in Appendix 2). Nevertheless, it is worth looking at the impact of changes in road-traffic fatalities on tourist flows. So we also fitted a regression equation of the form:

$$h_1(\mathrm{L})\Delta[\ln(p_{mt}) - \ln(p_{nt})] = h_2(\mathrm{L})\Delta\ln(z_{mt}) + h_3(\mathrm{L})\Delta\ln(f_{mt}) + h_3(\mathrm{L})\Delta\ln(RTF_t) +$$
$$j_1.[\ln(p_{mt-1}) - \ln(p_{nt-1})] + j_2.\ln(z_{mt-1}) + j_3.\ln(f_{mt-1})$$
$$+ d.DGW_t + j_4.\ln(RTF_t) + e_t \qquad (11a)$$

where RTF_t is the number of road-traffic fatalities per month. Since $\ln(RTF_t)$ is clearly stationary,[12] the t-ratio on j_4 on can be taken at face value, as an indicator of the significance of the long-run effect of road-traffic fatalities on tourist numbers. For the American sample this t-ratio is 0.18; for the European sample it is 0.09, so there is absolutely no evidence that *variations* in road-traffic fatalities have had any impact on tourism.[13] But given the low variance and insignificant autocovariance of $\ln(RTF_t)$, it would be somewhat rash to interpret the insignificance of j_4 as proof that tourists are more sensitive to the high-profile conflict risks emphasised in the media than they are to more mundane risks they face on the road.[14]

4. The Cross-sectional Model

12 Using data for the period 1988(2)-2004(5), the DF t-statistic for the seasonally adjusted ln(*RTF*) series is -12.19.

13 This is also true if Israeli road-traffic fatalities are scaled by foreign (for example, American) fatalities.

14 Nevertheless, such an asymmetry is consistent with the economic psychology of Kahneman *et al.* (1982), in which larger subjective probabilities are assigned to types of events, such as suicide bombings, that are more memorable.

4.1 The cross-sectional data: concepts

In the cross-sectional model, we intend to explain international variations in the decline in tourism to Israel between 1998-9 and 2001-2. By analogy with section 3.1, we will

focus on the growth in the probability that the i^{th} individual from a certain origin k will visit destination m, $\Delta\ln(p_{imk})$, and the corresponding growth in the actual number of tourists travelling from k to m, $\Delta\ln(p_{mk})$. We expect that these quantities will be negative for the majority of countries of origin. Note that the cross-sectional index k replaces the time-series index t. By analogy with equation (4) we might expect that:

$$\Delta\ln(p_{imk}) = \Delta m_{mk} - \ln\sum_{j=1}^{j=M}\exp(\Delta m_{jk}) \qquad (12)$$

where Δm_{mk} is the mean growth in the net utility of k-residents from visiting m. We will further assume that there is no substantial cross-sectional variation in the *change* in the desirability of locations other than Israel between 1998-9 and 2001-2.[15] In terms of equation (12) this means that if we take destination m to be Israel, $\sum_{j\neq m}\exp(\Delta m_{jk})$ is a constant. For no country does tourism to Israel make up more than 1% of total tourism, so $\sum_{j=1}^{j=M}\exp(\Delta m_{jk})$ is also likely to be approximately constant. Call this constant Δm. Our problem then reduces to modelling the determinants of Δm_{mk}. Again, we assume that it is possible to find a linear representation, this time of the form:

$$\Delta\ln(p_{mk}) = \Delta m_{mk} - \Delta m = \boldsymbol{W_{mk}}'z + e_k$$

where the $\boldsymbol{W_{mk}}$ are factors affecting the net utility from visiting Israel that vary according to the tourist's place of origin, and e_k is a cross-sectional residual.

The elements of $\boldsymbol{W}$ that appear in our regression equations are rather different from those in $\boldsymbol{X}$. First of all, we need to recognise that there is a substantial variation in the relative size of Jewish populations in different parts of the world. In the US, the Jewish population makes up about 2% of the total population, but in other countries the fraction is very much smaller.

15 Note that this assumption is consistent with some cross-sectional variation in the *level* of desirability of different locations; such variation is differenced out in our model.

Accurate data on the religious affiliation of tourists is not available, but it seems reasonable to suppose that a disproportionately large fraction of

tourists to Israel are Jewish. Moreover, these tourists might on average be more willing to visit Israel, even in the presence of violence, because of family ties or political commitment. So countries with a larger Jewish population might exhibit a smaller decline in tourism to Israel.

Secondly, the rate of decline in tourism might depend on the social and economic characteristics of the country of origin. In countries where there is a high level of violent crime residents may have learnt better how to avoid potentially dangerous situations, or they may have become less sensitive to the risks associated with living in a violent society. (Either they are not subject to the morbid and arguably irrational fears that beset tourists from very safe countries, or they are subject to cognitive dissonance regarding the risks they face, as in Akerlof and Dickens, 1982.) But even for a given level of crime, there may be a connection between the level of risk tourists are familiar with and the level of economic development in the country they come from. It is reasonable to assume that all international visitors to Israel are reasonably wealthy, relative to the world average: otherwise, they could not afford to travel. Those arriving from poor countries are atypically rich for their homeland; those arriving from rich countries are not especially wealthy, relative to the rest of their population. Because of their wealth relative to those around them, the first group may have more experience of being a potential target of criminals; so they may have become more acclimatised to a high level of personal security risk, and less sensitive to the risks currently involved in visiting Israel.

Thirdly, we ought to allow for changes in economic conditions in the country of origin between 1998 and 2002. (The insignificance of such effects in the time-series regressions does not mean that they will be insignificant in the cross-sectional regressions.) Two potentially important factors are the growth of the country's real exchange rate with respect to Israel – capturing changes in the cost of travel there – and the growth of it's real *per capita* income. A high rate of income growth might lead to a greater overall level of international tourist departures from the country and *ceteris paribus* a higher level of tourism to Israel.[16]

Allowing for all of these factors, the cross-section regression equation that we will estimate is of the form:

$$\Delta\ln(p_{mk}) = z_0 + z_1.\ln(1+PJ_k) + z_2.V_k + z_3.\ln(PCY_k) + z_4.\Delta\ln(C_k) + z_5.\Delta\ln(Y_k) + e_k \qquad (14)$$

where PJ_k is the proportion of the population of k that is Jewish, V_k is an indicator of lawlessness in k, PCY_k is a measure of *per capita* income in k in 2000, $\Delta\ln(C_k)$ is the growth of k's real exchange rate with respect to Israel between 1998-9 and 2001-2 and $\Delta\ln(Y_k)$ is the growth of its real income between these periods.

4.2 The cross-sectional data: application

$\Delta\ln(p_{mk})$ is measured using data reported by the Israeli Central Bureau of Statistics and available online at http://www.cbs.gov.il. For each tourist origin k we calculate the logarithm of the ratio of tourist arrivals in 2001 and 2002 combined to that in 1998 and 1999 combined. (Annual data for 2000 are difficult to interpret because the *Intifada* began in the middle of this year.) The ratios (not in logs) are reported in Table 5. We have excluded Arab countries from the data set, because tourists from the Arab world might be subject to varying visa requirements over the sample period. Otherwise, we report data from all countries listed by the CBS for which we can measure each element of $\boldsymbol{W_{mk}}$. It can be seen that there is substantial variation in the data. For three countries – Hong Kong, Malaysia and Italy – the ratio is less than 20%, but for another four –

16 No real income variable is included in the time-series model, which scales tourist volumes in Israel by tourist volumes in a reference location. This exclusion would be invalid only if the income elasticity of demand for international vacations varied with destination. When an income term is added to the time-series regression

Ukraine, Belarus, South Korea and Uzbekistan – the ratio is over 100%.[17] That is, there were a few countries from which tourist arrivals actually increased after the start of the *Intifada*. It turns out that the figures for Hong Kong, Malaysia and Italy are outliers in the distribution of $\Delta\ln(p_{mk})$. Inclusion of these three countries in the sample makes the distribution of $\Delta\ln(p_{mk})$ significantly non-normal.[18] At the very bottom end of the distribution there might be some non-linearity in the data generating process for $\Delta\ln(p_{mk})$. However, with only 57 observations in all we do not have enough degrees of freedom to model non-linearities in the tail. For this reason we adjust the figures for the three countries, raising them to -1.5 (implying a ratio of 22% in Table 5). A discussion of alternative ways of dealing with the non-normality is available on request: the results reported in section 4 are generally robust to the alternatives.

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equations it is statistically insignificant: there is no reason to suppose that there is any such variation in income elasticities.

17 The presence of South East Asian countries in both of these lists suggest that pure geographical factors are unlikely to be important determinants of $\Delta\ln(p)$. When regional dummιεσ αρε αδδεδ το τηε ρεγρεσσιον εθυατιονσ ρεπορτεδ ιν σεχτιον 4, τηεψ αρε ινδιϖιδυαλλψ ανδ φ οιντλψ ινσιγνιφιχαντ, ινχλυδινγ τηε φορμερ Σοϖιετ Υνιον δυμμψ. Νορ ισ τηε ρατε οφ δεχλινε ιν νυμ βερσ χορρελατεδ ωιτη τηε οριγιναλ σιζε οφ τηε τουριστ ποπυλατιον: τηε τ–ρατιο φορ τηε χορρελατιο ν οφ $\Delta\ln(p)$ and the log of tourist numbers in 1998-99 is -1.514.

18 There are no outliers at the other end of the distribution.

Table 5: Ratio of 2001/2 Tourist Arrivals to 1998/9 Tourist Arrivals

region	***ratio***	***Region***	***ratio***
Hong Kong§	0.108410	USA	0.488838
Malaysia§	0.122392	Singapore	0.497907
Italy§	0.180492	Latvia	0.501255
Sweden	0.262854	Thailand	0.531131
Slovakia	0.276851	Argentina	0.541050
Finland	0.282926	Iceland	0.544850
Denmark	0.293061	Venezuela	0.546745
Austria#	0.299289	Mexico	0.555017
Germany	0.319359	UK	0.603859
Brazil#	0.338326	Canada	0.624667
Portugal	0.339062	Croatia*	0.667670
Spain	0.352553	Colombia	0.668390
Japan	0.355187	France	0.680041
Greece	0.363988	China	0.698051
New Zealand	0.367148	Serbia/Montenegro#	0.698519
Estonia/Lithuania	0.377673	Turkey	0.711101
Norway	0.380121	Bulgaria	0.784864
Ireland	0.399542	Russia	0.792463
Netherlands	0.402578	Moldova	0.799424
Indonesia	0.420218	Georgia	0.827393
Australia	0.424272	India	0.830634
Chile	0.428402	Uruguay	0.863122
Belgium	0.432385	Romania	0.913361
Poland	0.435011	Philippines	0.934247
Czech Republic	0.442142	Ukraine	1.009040
Cyprus	0.443893	Belarus	1.088937
South Africa	0.451227	South Korea	1.098650
Hungary	0.475301	Uzbekistan#	1.140563
Switzerland	0.476993		

* Includes also Bosnia-Herzegovina, Macedonia and Slovenia.

No homicide data are available: this country is included in Sample B only.

§ In the regressions this country's observation is adjusted to exp(-1.5).

The Jewish population figures are taken from those published at www.jewishpeople.net; PJ_k is calculated by dividing these figures by the total population estimates published in the World Bank *World Development Indicators*. PCY_k is PPP adjusted *per capita* GDP in US Dollars, as reported in the United Nations *Human Development Report* 2001. C_k is calculated as the ratio of the GDP deflator in k to that in Israel, scaled by the value of the Sheqel in k-currency. Average figures for 1998-9 and 2002-2 are calculated, and $\Delta\ln(C_k)$ is the growth rate between the two periods. $\Delta\ln(Y_k)$ is constructed in an analogous way, with Y_k measured as real (2000) US Dollar GDP from *World Development Indicators* for 1998-9 and 2001-2.

Two alternative measures of V_k are considered. The first is the log of the number of reported homicides per 10,000 inhabitants in 2000, $\ln(H_k)$, reported in the UN World Crime Survey. This is available for 53 of our 57 countries. We expect $\Delta\ln(p_{mk})$ to be increasing in $\ln(H_k)$. The second, available for all 57 countries, is the 2000 *Rule of Law* measure described in Kaufmann *et al*. (2003) and here designated as ROL_k. This measure aggregates national scores awarded for the perceived level of crime in a country, the reliability of the judiciary and the enforceability of contracts. It is therefore a very much wider and more subjective indicator of the degree of lawlessness in society. Since higher scores are awarded to more lawful societies, we expect $\Delta\ln(p_{mk})$ to be decreasing in ROL_k.

Table 6 reports the results of fitting equation (14) to the data, first of all using the $\ln(H_k)$ measure, then using the ROL_k measure. The explanatory variables account for about half of the sample variation in $\Delta\ln(p_{mk})$. All variables except $\Delta\ln(C_k)$ and $\ln(H_k)$ are statistically significant at the 5% level, and all significant coefficients have the anticipated sign. The significance level for $\ln(H_k)$ is just above 10%, and it does not explain as much of the sample variation as the alternative measure ROL_k. However, with the exception of $\ln(PCY_k)$, the coefficients on other variables do not vary much between the two regression specifications. The $\ln(PCY_k)$ coefficient is sensitive to the specification because poor countries are much more crime-ridden, so $\ln(H_k)$, ROL_k and $\ln(PCY_k)$ are highly correlated. When $\ln(H_k)$ and ROL_k are replaced by their corresponding orthogonal components – i.e., the residuals from regressions of $\ln(H_k)$

Table 6: Cross Section Regression Results

The dependent variable is $\Delta\ln(p_{mk})$.

Standard errors are calculated using White's heteroskedasticity correction.

Sample A (53 observations)

variable	***coefficient***	***standard error***	***t ratio***	***partial R^2***
intercept	1.5214	0.6717	2.2651	0.0711
$\ln(1+PJ_k).100$	0.4305	0.1007	4.2764	0.1514
$\ln(PCY_k)$	-0.2729	0.0656	-4.1617	0.1952
$\Delta\ln(C_k)$	-0.4289	0.3631	-1.1811	0.0199
$\Delta\ln(Y_k)$	1.5398	0.7296	2.1106	0.0749
$\ln(H_k)$	0.0766	0.0462	1.6576	0.0452
R^2	0.4597			
σ	0.3228			
$\chi^2(2)$ residual normality test	2.3913			
RESET Test: F(3,44)	0.6622			

Sample B (57 observations)

variable	***coefficient***	***standard error***	***t ratio***	***partial R^2***
intercept	***0.8885***	***0.6646***	***1.3369***	***0.0267***
$\ln(1+PJ_k).100$	0.4526	0.1020	4.4363	0.1729
$\ln(PCY_k)$	-0.1826	0.0739	-2.4709	0.0846
$\Delta\ln(C_k)$	-0.2302	0.2395	-0.9611	0.0123
$\Delta\ln(Y_k)$	1.3228	0.5687	2.3260	0.0655
ROL_k	-0.1726	0.0515	-3.3544	0.1118
R^2	0.5258			
σ	0.3091			
$\chi^2(2)$ residual normality test	3.0845			
RESET Test: F(3,48)	0.7440			

$\ln(PCY_k)$ regression coefficients when $\ln(H_k)$ and ROL_k are orthogonalized

	coefficient	***standard error***	***t ratio***	***partial R^2***
Sample A	-0.3452	0.045497	-7.58687	0.4494
Sample B	-0.3405	0.051402	-6.62464	0.3896

and ROL_k respectively on $\ln(PCY_k)$ – the coefficients on $\ln(PCY_k)$ in the two specifications are almost identical. These coefficients are reported at the bottom of the table.

The table shows that if the fraction of local population that is Jewish is one percentage point higher (for example, 1% of the population instead of 2%) then the rate of decline of tourism over the sample period is on average 40% lower. This is consistent with large but unsurprising differences between Jews and non-Jews (on average) in terms of the deterrent effect of the violence. More interestingly, the regression equations with orthogonalized lawlessness indicators imply that a 10% increase in *per capita* income of the country of origin is associated with a rate of decline of tourism over the sample period that is around 3.4% higher. Part – but not all – of this effect is because a higher *per capita* income is associated with lower lawlessness in a country. Some of the *per capita* income effect has another source; one plausible explanation is that tourists from poor countries are more likely to be wealthier than their neighbors, and therefore more accustomed to being targets of violence.

Since the Rule of Law variable is an index, the coefficient on this variable is difficult to interpret *per se*. However, the sample standard deviation of ROL_k is 0.97, so the estimated coefficient shows, approximately, the effect on tourism decline of a one standard deviation change in the index. In more law-abiding societies the decline is greater, a standard deviation increase in ROL_k being associated with an additional 17% fall. The positive coefficient on the homicide variable $\ln(H_k)$ in the alternative regression specification is consistent with this effect, but the standard error on the homicide coefficient is very large, so it is not quite significant at the 10% level. Possibly this definition of lawlessness is too narrow.

Finally, despite the huge impact of the violence, tourists do seem to be sensitive to economic conditions at the margin. Countries with the largest real income growth have showed the smallest declines in tourism to Israel, *ceteris paribus*. Countries with income growth 1% higher have shown a rate of tourism decline that is about 1.5% lower on average. However, the coefficient on real exchange rate variable is insignificantly different from zero.

With regard to the potential differences between Europeans and Americans, the results in this section confirm those of the previous section, answering our original question in the negative.

Table 5 shows that the value of $\Delta\ln(p_{mk})$ for the USA lies in the middle range, only one observation away from the median. Some Western European countries (mainly Nordic and Southern Mediterranean ones, with lower crime rates and/or a smaller Jewish population) show far larger declines in tourism to Israel than does the USA. But others (notably France and the United Kingdom, with a lower *per capita* income) show substantially smaller declines. Once we have conditioned on a set of socio-economic characteristics, the remaining variation in the data (about half of the total variation) has no obvious socio-economic explanation and is uncorrelated with geographical location. In the *ROL* regression, the estimated value of e_k for the USA is -0.18, implying a *larger* decline in tourism than average, conditional on the RHS variables in equation (14). This compares with a German e_k of -0.11, a French e_k of 0.19 and a British e_k of 0.33; but the sample standard deviation of e_k is 0.30, so none of these differences is statistically significant. As Table 6 indicates, the null that e_k is normally distributed cannot be rejected.

5. CONCLUSION

Analysis of time-series and cross-sectional Israeli tourist data reveals some of the factors driving people's attitudes towards the risk associated with travel to a conflict region. Time-series analysis shows that since the onset of the *Intifada* even the relatively small variations in conflict intensity – as measured by the number of fatalities per month – have affected tourist volumes. This is true of both American and European tourists, with no significant differences between the two groups. It is consistent with a model in which, even at moderate levels of violence, a large number of people are approximately indifferent between travelling and not travelling. As a consequence, we can expect even a partial reduction in violent conflict in the region to boost tourism revenue, which could be grounds for optimism regarding a gradual resolution of the conflict.

It is also worth noting that tourists are sensitive not only to deaths within Israel, but also (to a lesser degree) deaths of both Israelis and Palestinians in the West Bank and Gaza. All dimensions of the conflict, and not only Israeli deaths in suicide bombings, have an impact on the Israeli economy. In our fitted model, an increase in monthly Israeli fatalities from zero to ten deaths, such as would be caused by a large suicide bombing, would reduce American tourist numbers by around 30% in the next month and 45% in the month following. (Thereafter tourist numbers would swiftly recover.) The estimated effects on European tourist numbers are of the same order of magnitude, implying to a total loss of tourist revenue in the order of $250mn. An equivalent increase in WBG fatalities would reduce American tourist numbers by around 15-20% in the second and third months following. Given that the monthly average number of fatalities in WBG is 64 (as opposed to five in Israel) Palestinian deaths cost the Israeli economy a substantial amount of money.

Analysis of cross-sectional data reveals more about the differences, and the absence of differences, between tourists of different nationalities. Some socio-economic characteristics (such as a high average income levels and a low crime rate) are associated with a larger decline in tourist numbers when the violence starts. Tourists from countries at lower levels of economic development are less sensitive to the violence. Once we have controlled for these characteristics there is no obvious geographical pattern to the variation in tourist behaviour. “Old Europe” demonstrates no more and no less risk aversion than the New World.

We ought to be cautious in inferring from these results about a sample of tourists conclusions about whole populations. In many countries international tourists might not be typical of the population in which they live. Nevertheless, the homogeneity of the time-series regression results across European and American samples, and the extent to which the international cross-sectional variation in tourist behaviour is associated with a few simple socio-economic characteristics, create a strong impression that, for a given level of social and economic welfare, people are pretty much the same everywhere.

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Appendix 1

Here we discuss briefly our measurement of the relative cost series, which turned out never to be statistically significant in the time-series regression equations. Data on hotel and restaurant prices in America, Europe and Israel are available, facilitating the construction of hospitality price real exchange rate series. However, such series are unlikely to be exogenous to total tourist volumes, and in this context there is no obvious instrument for hotel and restaurant prices. For this reason we measured relative costs as an aggregate consumer price real exchange rate. For American tourists this was the log of the ratio of the Israeli consumer price index to the Euroland consumer price index, scaled by the Shekel-Euro nominal exchange rate. For European tourists it was the log of the ratio of the Israeli consumer price index to the US consumer price index, scaled by the Shekel-Dollar nominal exchange rate. Nominal exchange rate and price indices are reported by the Israeli Central Bureau of Statistics (http://www.cbs.gov.il), the Federal Reserve Bank of St Louis (http://research.stlouisfed.org/fred2) and the European Central Bank (http://www.ecb.int). Substitution of a (probably endogenous) hospitality price real exchange rate for the aggregate consumer price real exchange rate made no difference to the insignificance of relative costs in the regression equations.

Appendix 2

In this section we explore the dynamics of the two conflict series, $\ln(z_{mt})$ and $\ln(f_{mt})$. First of all, we test for the existence of a long-run relationship between the two series by fitting a regression equation of the form

$$q_1(L)\Delta\ln(z_{mt}) = q_2(L)\Delta\ln(f_{mt}) + w_1.\ln(z_{mt-1}) + w_2.\ln(f_{mt-1}) + x_t \qquad \text{(A1)}$$

where x_t is a regression residual, and computing an F-test for the joint significance of the w parameters. With a lag order of 1, selected on the basis of Schwartz-Bayesian and Akaike information criteria, we compute $F(2,37) = 2.62$. This is not significant at the 5% level, even under the assumption that the variables are stationary, and certainly not if they are I(1) (Pesaran *et al.*, 2001). So there is no evidence for a long-run relationship. Next we test for the existence of a short-run relationship by fitting and equation of the form

$$\Delta\ln(z_{mt}) = y_0 + y_1\cdot\Delta\ln(z_{mt-1}) + y_2\cdot\Delta\ln(f_{mt-1}) + c_t \qquad \text{(A2)}$$

where (t is a regression residual. Our estimated value of (2 is 0.536 (t = 2.305), so there is some evidence for a short-run relationship. Changes in f_{mt} do help to predict changes in z_{mt+1}.

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