How Does Civil War Affect the Magnitude of Capital Flight? Evidence from Israel during the *Intifada*

David Fielding

Department of Economics, University of Leicester

and WIDER, United Nations University, Helsinki[§]

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Abstract

We use time-series data from Israel to investigate the dynamics of the causal links between the intensity of civil conflict and capital flight. The fraction of Israeli capital wealth held outside the country exhibits considerable variation over time. So also do indicators of the intensity of the Palestinian-Israeli conflict. Using quarterly time-series data, the paper shows that there is a high correlation between the two, conditional on economic conditions. This correlation is a consequence of a causal link that runs in both directions: more violence leads to more capital flight, but more capital flight is also a predictor of higher future levels of violence.

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[§] Address for correspondence (March-August 2003): UNU/WIDER, Katajanokanlaituri 6 B 00160, Helsinki, Finland. E-mail: DJF14 @LE.AC.UK; telephone: +358-9-61 59 92 04; fax: +358-9-61 59 92 04.

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1. Introduction

Interest in the links between economic performance and the political environment has recently been renewed. Economists have started to take seriously the possibility that political instability and unrest can have a large and lasting impact on the real economy. In part this is a by-product of the renaissance in growth theory, and much of the empirical work uses the cross-country panel data constructed by the World Bank. Time-varying data on political factors is seldom available on a monthly or quarterly basis, so the application of time-series econometric techniques is very rare.

In particular, attention has been drawn to the economic consequences of civil wars, and the factors that determine the magnitude of the economic costs that a war generates. For example, Collier (1999) develops a theoretical framework in which a key feature is the adjustment of the capital stock through capital flight. Conflict reduces the productivity of domestic capital and increases (expected) physical depreciation rates. Domestic investors will relocate their capital abroad, leading to a reduction in domestic output. The model is tested on data for all civil wars since 1960. After long civil wars (when the capital stock has had time to reach its lower steady state) the economy recovers rapidly, whereas after short wars it continues to decline.

A related body of literature pursues the empirical link between capital flight and a wide range of indicators of civil conflict and unrest. Indicators used in panel data regressions include the incidence of political protests, major government changes and civil rights indices. Papers taking this type of approach include Tabellini (1988), Lensink *et al*. (1998), Le and Zak (2001) and Ndikumana and Boyce (2002). There appears to be a consensus that cross-country variations in political stability do account for a substantial part of the variations in capital flows.

More generally, there is a body of work investigating the links between countries' long run growth rates and indices of political rights, political instability and violence. Alesina and Perotti (1994) survey many of the papers outlining these ideas; recent additions to the literature include Easterly and Levine (1997) and Fedderke and Klitgaard (1998). Different papers find different socio-political indicators to be significant in explaining variations in economic performance across countries, but there is a consensus that a substantial fraction of the variation is to be explained by the quality of a country's polity.[1]

[1] There is a closely related literature examining cross-country variations in gross investment performance. A paper by Alesina and Perotti (1993) explains such variations by using a "sociopolitical instability index" constructed using principal components analysis. The important factors in the index are indicators of the absence of democracy and the incidence of political violence. Kormendi and Meguire (1985) and de Haan and Siermann (1996) discover a negative cross-country correlation between the investment-GDP ratio and indices of political freedom.

These studies provide a wealth of international evidence on the links between political instability and conflict and the macro-economy. However, the use of panel data does embody certain limitations. Firstly, the slope coefficients on political variables in cross-country regressions should be interpreted as the mean effect on economic performance of a certain political characteristic, across all the countries in the sample.[2] We cannot say much about individual countries, or individual types of countries. Secondly, the panel data sets are based on annual data, so they cannot be used to explore the dynamics of the interaction between the political and economic variables. For example, to what extent do short-term variations in the level of conflict lead to changes in short-term capital flows? If conflict intensity increases, how much of a lag is there in the capital flow response? Here the potential value of econometric evidence on individual countries using time-series data – were it available – would be very high.[3]

One other country for which time-series data on indicators of conflict intensity do exist, and in which these indicators have exhibited a large degree of variability in recent years, is Israel.[4] In this paper we will use Israeli time-series data to investigate one link in the chain linking economy and polity. If Israeli investors are sensitive to the risks inherent in a volatile political environment, then the fraction of their portfolio made up by domestic assets should be sensitive to prevailing political conditions. In other words, the magnitude of capital flight should have an explanation that is partly political. Another advantage of using time series in this context is that it becomes feasible to address the potential endogeneity of the political variables. Variations in economic conditions might impact on the level of violence, so identifying the direction of causality will be an important aspect of the analysis.

However, the aim of this paper is not just to provide some time-series evidence to complement existing cross-section work. Econometric modeling of the impact of political instability and violence on the economy can also inform public policy. It can provide an estimate of the size of the "peace dividend", the magnitude of the economic benefits that are likely to ensue from an end to, or at least reduction in, political instability. The size of this expected dividend ought to be one factor in determining the amount of resources devoted to achieving political stability by the state (and also by

[2] One serious problem with the panel data regressions is the difficulty in producing an unbiased estimate of this mean value. See Pesaran and Smith (1995).

[3] To date there is (as far as we know) only one paper that looks at the links between political instability, investment and capital flight within a time-series context: Fedderke and Liu (2002), who use South African data. Proceeding from the assumption that changes in the South African polity are weakly exogenous to capital flight indicators, they find that capital flight is positively associated with an index of political instability. Reductions in stability lead to more capital flight.

[4] The term "Israel" will be used to denote the geographical area currently governed from Jerusalem, i.e. the State of Israel within its 1948 borders plus the West Bank, Gaza and the Golan Heights. All place names are purely geographical and have no geopolitical implications.

the international community).[5] Determining the amount of capital that investors locate outside of Israel because of the violence – capital that could be productively located within the economy, possibly at a higher rate of return – is one element of this accounting process.

In the next two sections, we review the political and economic time-series data for Israel that will be used in the econometric analysis in sections 4-5.

2. The *Intifada* and Indicators of Political Instability in Israel

As a consequence of the 1967 Arab-Israeli war, Israel currently governs territories outside its 1948 borders, including the West Bank, i.e., territory west of the River Jordan but east of the 1948 border, and the area around the city of Gaza. The majority of the population in these areas is made up of Palestinian Arabs, many of whom contest the legitimacy of Israeli rule and Jewish settlement of the territories. In December 1987 there was a sudden uprising (the first *Intifada*) among Palestinians in these areas (Peretz, 1990). The uprising consisted of strikes and public demonstrations, which often escalated to the point where protestors were shot dead by members of the Israeli Defense Forces (IDF); later there was an increase in the number politically motivated assassinations and attacks on Israeli targets by Palestinian paramilitary groups, particularly *Hamas*. The uprising continued up to September 1993, when the Israeli Government signed an agreement with the Palestine Liberation Organisation (the Oslo Peace Accord). This agreement included PLO recognition of the State of Israel and Israeli recognition of the need for Palestinian self-government in at least part of the West Bank and Gaza areas. The political structures envisaged by the Oslo Peace Accord have not yet been fully implemented, and the political violence and instability have not ceased. Over the 15 years since the start of the first *Intifada* the magnitude of political tensions and violence has varied considerably. Towards the end of 2000 there was another upsurge in the intensity of the conflict that has become known as the "second" *Intifada*. The purpose of our paper is to construct a macro-econometric model that uses this variation to estimate the extent to which the political instability has impacted on Israelis' decisions about where to locate their capital wealth.

In creating time-series measures of the intensity of the conflict, we aim to capture the factors that determine Israeli investors' perception of the degree of political risk associated with locating capital in Israel. For most investors, the direct threat to property from the current level of violence is very small: the fraction of the Israeli capital stock destroyed as a result of Palestinian bombs (or for that matter IDF actions that result in the destruction of property) is very small. However, variations in the current level of violence could alter investors' perceptions about the probability of a serious escalation in the conflict in the near future, in which either the rate of return to physical capital deteriorated

[5] Of course the consequences of the *Intifada* should be measured in terms of the ensuing human costs. The point of this paper is to show that these costs are not confined just to those suffering personal injury or loss of property as a result of

sharply, through a reduction in domestic demand, or capital was destroyed as a direct consequence of the violence.

Some time-series indicators of attitudes and beliefs about the Israeli-Palestinian conflict already exist: for example, the TSC Peace Index (Yaar and Hermann, 2003). However, this and other indices do not exist for periods before 1994. Also, the structure of the permanent questions in these surveys is designed to elicit responses about a peaceful resolution of the conflict: for example, about whether people expect the peace process to succeed. Questions explicitly about feelings of personal insecurity have not featured regularly. Opinions about the likelihood of success for the peace process do not map straightforwardly onto opinions about levels of perceived risk. Some parties in Israel claim that an abandonment of the peace process and a more draconian policy with respect to the Palestinians would increase Israeli security in the long term.[6]

In this paper we will restrict our attention to "objective" indicators of conflict intensity, that is, indicators that are not based on attitude surveys. We do not attempt to model the link between the objective indicators and attitudes, or between attitudes and capital flight. Rather we will fit a model that embodies both these links in reduced form, and estimate the relationship between the objective indicators and capital flight.

Conflict intensity can be measured along a number of dimensions. First, one can measure the number of casualties in politically motivated violent incidents each quarter. Figure 1 shows the number of Israeli and Palestinian deaths from the beginning of the first *Intifada* (in 1987) up to the end of 2001.[7] Both series exhibit a great deal of variability over the sample period. The most noticeable change is the large upsurge in violence associated with the beginning of the second *Intifada* in the last quarter of 2000; but this is by no means the only source of variation in the series. Both series together reflect the intensity with which the conflict is being pursued, although one possibility we will consider is that Israeli deaths have more impact on capital flight than Palestinian deaths, because most of the wealth is held by Israelis.

For a given level of violence, as indicated by total fatalities, there may be other factors that influence investors' perception of risk. The current security policy of the Israeli administration might also make a difference. One crude way of capturing this is to use dummy variables for different political administrations over the sample period.[8] However, such dummies turn out not to capture any

violent action by Israeli and Palestinian forces. Israel as a whole suffers from greater poverty.

[6] Questions relating to personal security do sometimes appear in the surveys, but not a regular or frequent enough basis to make use of them in a time-series model.

[7] All the political time series in the figures are provided by the Israeli human rights organisation B'Tselem. The general accuracy of the aggregate figures is not contested by other organisations such as the IDF, although interpretation and qualitative disaggregation of the series (for example into combatants and non-combatants) is highly controversial. We are using the latest estimates of the fatality statistics, which are periodically revised. However, using "real time" political data does not make any substantial difference to our econometric results.

[8] One could also include a dummy variable for the Gulf War. This is also statistically insignificant.

of the variation in the capital flight series discussed in section 3 below. One observable time-varying quantitative indicator of the severity of the government's security measures is the number of days in each quarter when the border between Israel proper and the West Bank or Gaza is closed. (In fact it is very rare for one border to be closed and not the other.) For a given level of violence, an increase in the number of closure days might increase Israeli investors' sense of security, because it is perceived that this will help to prevent future escalation in the conflict. On the other hand, an increase in closures might create the impression that the government is failing to contain Palestinian attacks, or that closures will provoke more violence in the future. So *a priori* the impact of closures on capital flight is uncertain. Figure 2 illustrates the closures time series. As with fatalities, the largest change is after the commencement of the second *Intifada*, after which the border has been closed virtually all the time. However, there are other periods in which the number of closures per quarter varies considerably. The frequency of border closures is not the only indicator by which the extensiveness of defensive Israeli security measures might be assessed. For example, qualitative indicators such as the intensity of government rhetoric might also matter. Nevertheless, we think that border closures are highly correlated with these more qualitative indicators, and so ought to work reasonably well as an overall indicator of the "toughness" of defensive Israeli security measures. Our main indicators of conflict intensity will be the fatality statistics (possibly disaggregated by nationality of the deceased) and the number of border closures.

When we are looking at the impact of political conflict on economic variables such as capital flight, it makes sense to measure conflict intensity in terms of conflict outcomes, such as the variables discussed above. However, we also wish to look at the possibility of causality running in the opposite direction. When we are looking at the determinants of conflict intensity, it is not so obvious that the focus should just be on outcomes *per se*. The impact of economic conditions on the intensity of the conflict, if such an effect exists at all, will be through an effect on the propensity of each side to engage in aggression. These propensities cannot be observed directly, nor can they readily be identified from the fatality data. In order to illustrate this point we refer to Tables 1-2.

The tables report some disaggregated statistics for fatalities during the second *Intifada*, i.e. the end of September 2000 to the present (end of November 2002). Table 1 disaggregates Palestinian fatalities according to whether the deceased was an adult, and whether (s)he died in an IDF assassination, or from other IDF gunfire, or from Israeli civilian gunfire.[9] Table 2 disaggregates Israeli fatalities according to whether the deceased was a member of the IDF, and whether (s)he died in a direct Palestinian assault (such as an ambush or bombing), or in other gunfire, or in some other way (such as stone-throwing). Some of these categories are readily identifiable as fatalities resulting directly from an aggressive action by the other side: for example, the IDF assassinations or

[9] Palestinian deaths in suicide bombing attacks are excluded.

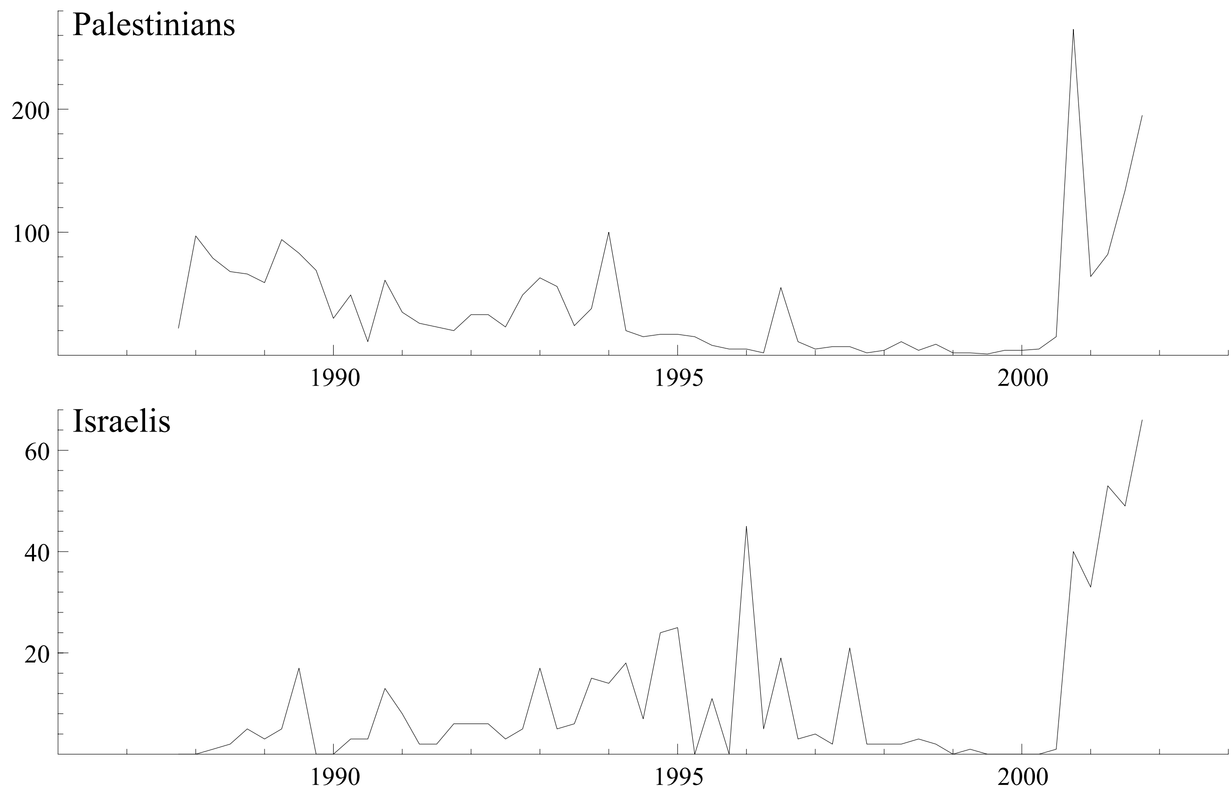

Figure 1: Quarterly Palestinian and Israeli Fatalities (Source: B'Tselem)

Figure 2: Number of Border Closure Days (Source: B'Tselem)

Table 1: The Total Number of Palestinian Fatalities During the Second *Intifada*

	Adults	Minors
From IDF Assassination	124	――
From other IDF Gunfire	1,197	299
From Israeli Civilian Gunfire	22	3

Source: B'Tselem (figures to the end of November 2002)

Table 2: The Total Number of Israeli Fatalities During the Second *Intifada*

	IDF	Civilians
From Palestinian Ambush/Bombing	99	364
From other Palestinian Gunfire	98	86
Other	12	22

Source: IDF (figures to the end of November 2002)

Palestinian bombings. But the majority of Palestinian fatalities and a large proportion of Israeli fatalities are in general fighting where it is not known with certainty – and is often hotly contested – which side instigated the conflict. Moreover, it is difficult to distinguish between combatants and non-combatants on each side. Most – but not all – of the Israeli civilians who died were unarmed at the time of the attack. The same can be said of most – but not all – of the Palestinian minors. How many of the Palestinian adult fatalities were combatants is also highly controversial: some, but certainly not all, were members of an armed militant group.

For this reason, any workable time-series model that explains or predicts variations in the intensity of the conflict will be in some sense "reduced-form". We will model the level of observed fatalities as a function of economic and political conditions, and interpret the results in terms of the impact of these conditions on the propensity of one or both sides to engage in aggression. However, such interpretation is necessarily speculative, since we do not observe "aggressiveness" directly.

Section 4 below discusses the detail of the model incorporating the fatality and bordure closure time series. The next section discusses the way in which we measure capital flight.

3. How Should Capital Flight be Measured?

A number of recent papers (for example, Collier *et al*., 1999; Le and Zak, 2001) have pointed out that it makes sense to model international capital flows as a consequence of a portfolio choice by the representative investor. That is, capital flows result from changes in the proportion of investors' *stock* of wealth made up by foreign assets. This argument applies equally to physical capital and to financial capital. Each period, domestic residents will make decisions about their preferred ratio of wealth in foreign assets to wealth in domestic assets. If this optimal ratio differs from the actual ratio, there will be some portfolio adjustment and a net capital inflow or outflow.

Figure 3 illustrates this concept for Israel. There are two quarterly time series in the figure; both represent the logarithm of an asset ratio. In each ratio the denominator is the same: the total net wealth of Israeli residents held in the form of domestic capital. This is measured as the total private sector Israeli capital stock (perpetually discounted private real domestic investment) less the fraction of this stock held by foreigners (perpetually discounted real net foreign direct investment in Israel). The numerator for the asset ratio F_1 is the stock of foreign physical capital held by Israeli residents (perpetually discounted private real net foreign direct investment by Israelis abroad). For the asset ratio F_2 it is the stock of foreign financial assets held by Israeli residents.[10, 11] The ratios F_1 and F_2 both have a positive trend over the sample period; that is, Israelis have been increasing the fraction of their wealth held abroad. Part of this trend might reflect increasing financial openness in the Israeli economy, but it might also be a consequence of trends in identifiable political or economic factors that have made Israel a less attractive place to invest. Over the *Intifada* period 1987-2001 the average quarterly growth rate in F_1 – the average rate of FDI capital flight as we measure it – is 2.68%. For F_2 the figure is 3.56%.

In the econometric analysis described in the following sections, we will make use of the portfolio share concept as illustrated by the asset ratios in Figure 3. Regression diagnostic statistics and information criteria suggest that the models we construct using the asset ratio approach are to be preferred to models based on deflated capital flows (such as real FDI by Israelis), and we will not report results for the latter. Nevertheless, the stylized facts that we infer from the asset ratio regressions also apply to the unreported capital flow regressions.

Before constructing an econometric model of capital flight, we should also look at evidence on the relative rates of return to domestic and foreign assets. One measure of the relative rate of return is depicted in Figure 4: the interest rate differential [i-i*-ds]. This is the difference between the domestic annualized three-month t-bill rate i and the corresponding US rate i*, controlling for the growth rate of the Sheqel-Dollar exchange rate ds. Most of the variance in this series is accounted for by movements in the exchange rate. As Figure 4 shows, there is considerable volatility in the interest rate differential, but as Table 3 below indicates the series is stationary. It also has a mean insignificantly different from zero, suggesting that there is interest parity between Israel and the USA in the medium/long term. There is no statistically significant risk premium in the t-bill interest rate differential.

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[10] *F2* measures gross foreign financial wealth; the general observations made about *F2* are also mostly true of net foreign financial wealth.

[11] The capital depreciation rates and investment deflators for all Israeli FDI abroad are assumed to be the same as for the US private business sector. Appendix A gives more details about data sources and data construction.

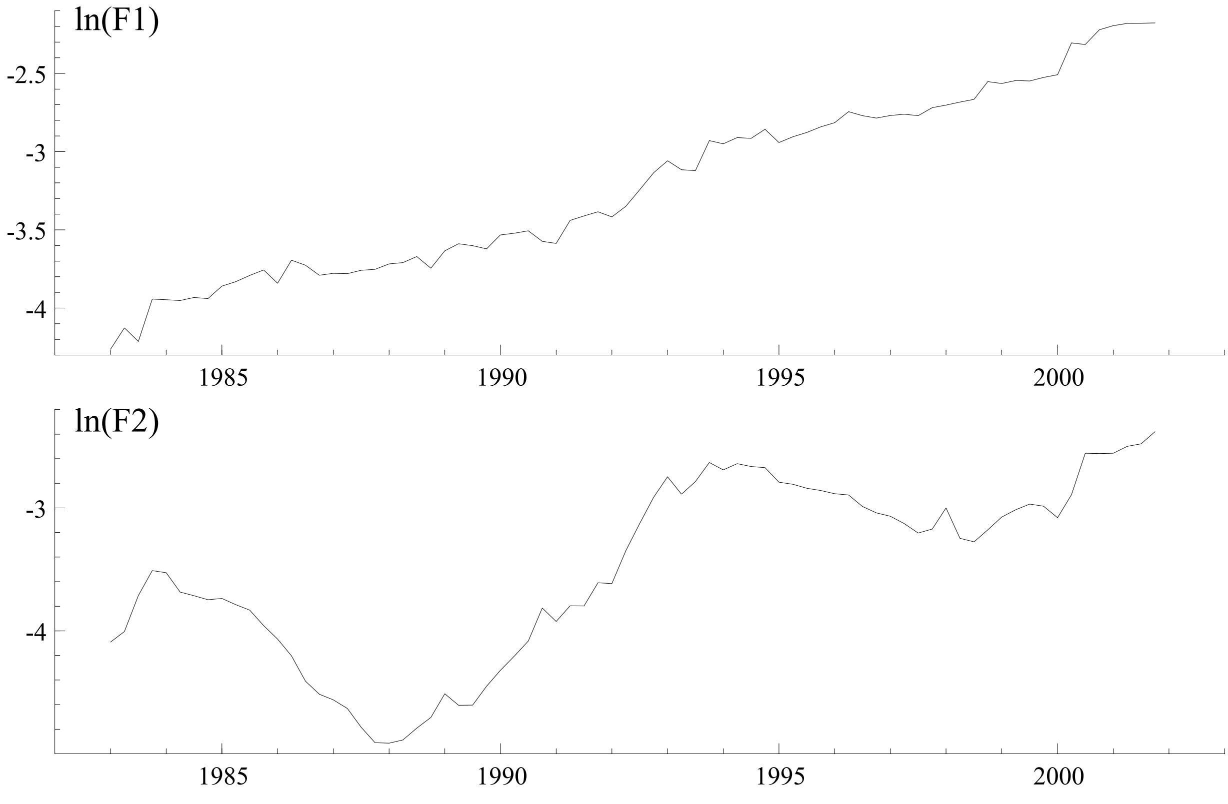

Figure 3: The Asset Ratios F_1 and F_2

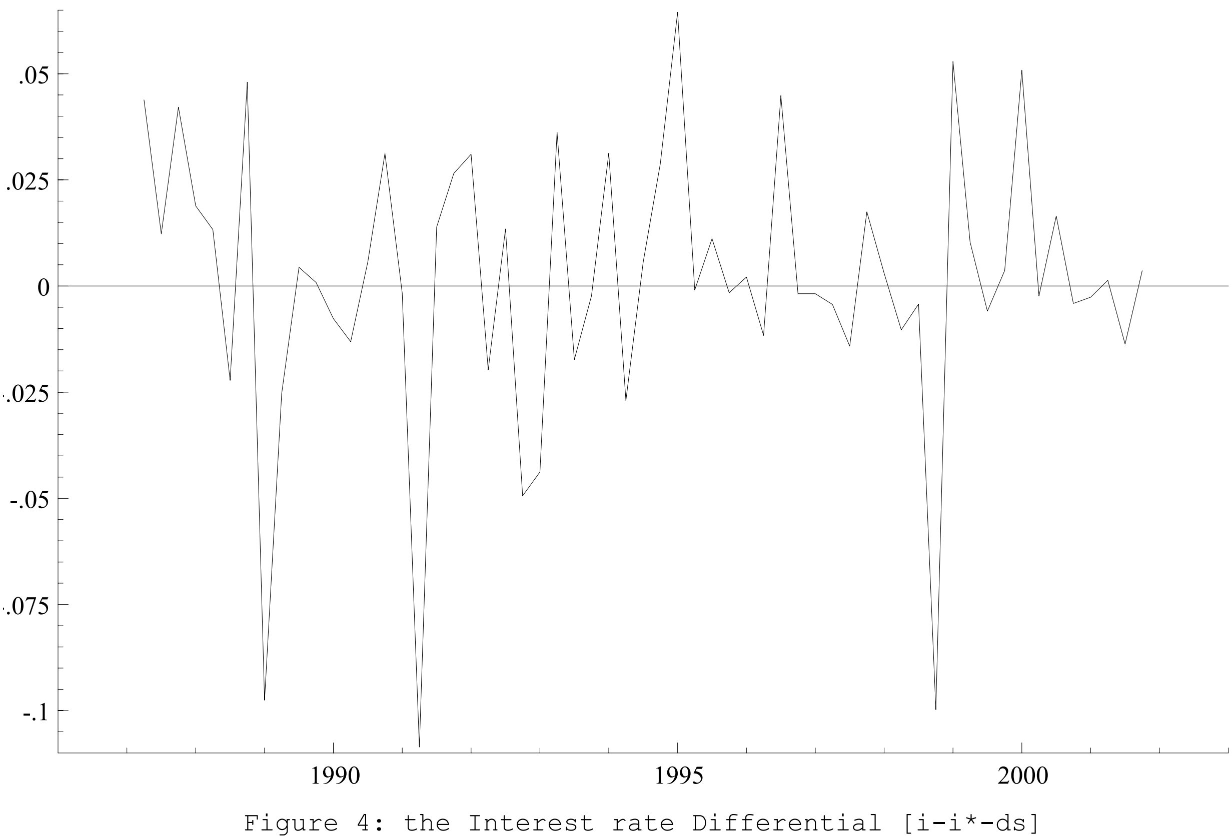

Figure 4: the Interest rate Differential [i-i*-ds]

One explanation for the absence of a risk premium in such a differential is that financial assets with a three-month maturity are liquid enough for the political risk factor not to be important. Given that the current conflict falls short of a full civil war (in which the rate of return would be seriously compromised), investing in local financial capital is not particularly risky: three months is long enough to relocate one's financial wealth offshore, if that becomes necessary. This suggests that we should focus on assets with less liquidity. It is not possible to disaggregate the financial asset ratio F_2 according to time-to-maturity, so in the analysis that follows we restrict our attention to the physical capital ratio F_1.[12]

4. The Model of Capital Flight

In the empirical model of capital flight we use quarterly time-series data, allowing the portfolio shares of domestic residents, as measured by our capital ratio F_1, to vary with both political and economic factors. The construction of a complete structural econometric model of the Israeli economy is beyond the scope of this paper, and certain macroeconomic variables that impact on capital flight (such as the real exchange rate and the industrial wage) are not modeled: when these variables appear as regressors, they are instrumented by their own lagged values. Nevertheless, we will identify the impact of conflict intensity on capital flight, and of capital flight on conflict intensity, conditional on macroeconomic conditions. For the sake of clarity, it may be useful to note at the outset that we will find conflict intensity to be weakly (but not strictly) exogenous to capital flight. This section of the paper focuses on the capital flight equation.

The first equation in our econometric model explains the share of capital wealth held in the form of physical capital abroad, F_1. The economic variables entering into the F_1 equation are as follows (sources are listed in Appendix A).

(i) The international interest rate differential, [i-i^*-ds]. With instantaneous domestic capital market perfection the interest rate differential would capture all there is to know about the current relative rate of return to capital in Israel and the US. However, investors in physical capital will base their decisions on forward-looking expectations about future relative rates of return that may depend on other macroeconomic factors.[13, 14] Moreover, it is possible that there are differences between the marginal cost of capital and the

[12] We did try some regressions with $F2$ as the dependent variable; but these did not produce any robust relationship between the asset ratio and indicators of political risk.

[13] One could in principle test theories about alternative models of expectations (for example, Rational Expectations), by imposing cross-equation restrictions on the interest rate and capital flight equations that are presented below. However, our sample is quite small, so such tests of non-linear restrictions would not have a great deal of power. We prefer to remain agnostic about the nature of expectations formation, so we do not impose any such restrictions.

[14] The Bank of Israel does report some rates of return on bonds with a longer maturity. However, there are many such rates, none of which is reported consistently for the whole of the sample period. We were not able to construct a robust model of capital flight using any of these alternative measures.

marginal return, at least in the short run. For this reason the model includes some other indicators of relative returns.

(ii) The first of these is the PPP real exchange rate, $p/(s.p^*)$. p is the retail price index for Israel, and p^* the corresponding index for the USA. A larger value of the real exchange rate indicates a relatively high return to goods sold domestically. If some commodities are not internationally tradable, or at least if international transport costs are greater than zero, a larger value of $p/(s.p^*)$ ought to encourage capital to locate in the domestic economy rather than abroad.[15]

(iii) The second is the real industrial wage ratio, $w/(s.w^*)$. w is the Israeli index of industrial wages, and w^* the corresponding US measure. A lower value of $w/(s.w^*)$ will be associated with relatively low labor costs, and ought to encourage capital to locate in the domestic economy rather than abroad.

(iv) Finally, there is the growth rate of output in the USA, $\Delta\ln(y^*)$, and of output in Israel, $\Delta\ln(y)$. With increasing returns to scale and imperfect competition, the rate of return to capital will depend on the overall level of economic activity in an economy (see for example Rama, 1993). Higher growth in the USA ought to encourage capital flight; higher domestic growth ought to discourage it.

In addition to the economic variables, we include the following indicators associated with the level of violence and political instability.

(i) The total number of fatalities in violent political incidents in Israel in each quarter, D. We anticipate that increases in the level of violence will intensify the perception that Israel is a risky place to invest, and so encourage capital to relocate abroad. It is possible to disaggregate D according to the nationality of the deceased (Palestinian, Israeli, other), their civilian / military status and the location of the incident (Israel, Gaza, West Bank). However, these disaggregations turn out not to have any explanatory power in the empirical model.

(ii) The total number of days in the quarter on which either the Israel-West Bank border or the Israel-Gaza border was closed, C. There are two potential reasons for the importance of this characteristic of the conflict. First, it might in itself be taken as an indicator of conflict intensity, so a rise in C is associated with greater capital flight. Alternatively, conditional on the level of D, higher values of C might reduce the perception of political risk in some investors' minds - because it is a quantifiable indicator of the "toughness"

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[15] Some authors, for example Fedderke and Liu (2002) and Collier *et al.* (1999), use the real exchange rate (or the transformation advocated by Dollar, 1992) as a measure of macroeconomic imbalance. A higher level of the real exchange rate is presumed to be associated with greater macroeconomic risk and more capital flight. This interpretation turns out not to be that plausible in the case of Israel, since increases in $p/(s.p^*)$ are associated with less capital flight.

of the Government response to Palestinian insurgency - and so reduce the intensity of capital flight. Here, it is the investors' subjective perceptions that matter; later on we will investigate what effect the border closures actually have on other conflict variables.

(iii) The commencement of the second *Intifada* at the end of September 2000 represents an enormous structural break in the intensity of the conflict. It is quite possible that the response of capital flight to conflict intensity exhibits some non-linearities, but we have only five observations for the period of the second *Intifada* (2000q4-2001q4), so such non-linearities are difficult to estimate with any precision. For this reason we include an intercept shift in the capital flight equation for the period after the onset of the second *Intifada*. The shift variable DUM_t takes a value of unity for 2000q4-2001q4, and a value of zero otherwise.

A number of existing papers using cross-section or panel data (for example Le and Zak, 2001) also include in their regression equation indicators of economic risk, such as measures of the variance of the domestic rate of return to capital. However, the economic variables in our time series data set do not exhibit any significant autoregressive heteroskedasticity: their conditional variance appears to be more or less constant over our sample period. For this reason we do not include any measures of economic risk in the model.[16]

In the absence of any compelling *a priori* model structure, we use a regression equation around an unrestricted trend with the following functional form:[17]

$$\begin{aligned} a_1(L)\ln(F_I)_t = {} & a_2(L)[i-i^*-ds]_t + a_3(L)\ln(p/s.p^*)_{t-1} + a_4(L)\ln(w/s.w^*)_{t-1} + a_5(L)\Delta\ln(y^*)_t \\ & + a_6(L)\Delta\ln(y)_t + a_7(L)\ln(1+D)_t + a_8(L)\ln(1+C)_t + b_0 + b_1\cdot t + b_2\cdot DUM_t + \varepsilon_t \end{aligned} \quad (1)$$

where ε_t is an i.i.d. residual and $a_i(L)x_t = \sum_{r=0}^{r=n} a_{ir}\cdot x_{t-r}$, with the lag order n to be determined by an empirical information criterion. Logarithmic transformations are used to ensure that the variables are approximately normally distributed; $(1+D)$ and $(1+C)$ are used instead of D and C because these variables are occasionally equal to zero. In the tables reported below, the $a_6(L)$ parameters have been set to zero, as $\Delta\ln(y)_t$ is never statistically significant in any regression equation. This makes no difference to the size or significance level of the other explanatory variables.

Table 3 reports ADF test statistics for the null that the variables of interest are difference stationary, against the alternative that they are stationary in levels, possibly around a deterministic

[16] One could perhaps include unconditional variance measures, but the theoretical grounds for doing so are dubious, and anyway such variables are not statistically significant when added to our regression equation.

[17] Note that contemporaneous values of the real exchange rate and the relative wage level do not appear in the regression. We do not want to assume that they are weakly exogenous to capital flight, and no convincing instrument (other than their own lagged values) is readily available.

linear trend. The table reports the sample size for each variable (the longest available, with the one exception noted below), the lag order of the ADF regression (chosen on the basis of the Schwartz Bayesian Information Criterion) and the p-value for the test of the null. Following Bierens (2002), who points out the magnitude of size and power distortions for the standard ADF test in small samples, this p-value is not taken from the standard tables. Instead we fit a model for each variable x_t under the null:

$$c(L)\Delta x_t = f_0 + \eta_t \tag{2}$$

and use the i.i.d. residual η_t to create 10,000 bootstrap samples on which to run the ADF regression:

$$c'(L)\Delta x_t = f'_0 - f_1 \cdot x_{t-1} + f_2 \cdot t + \eta'_t \tag{3}$$

Table 3: ADF Test Statistics

variable	*sample*	*lag order*	*trend*	*p-value*
ln(F1)	83(1)-01(4)	0	yes	0.0184
ln(F2)	83(1)-01(4)	6	yes	0.0006
Δln(y)	83(1)-01(4)	1	no	0.0000
Δln(y*)	83(1)-01(4)	1	no	0.0018
ln(p/s.p*)	87(4)-01(4)	4	yes	0.0142
ln(w/s.w*)	87(4)-01(4)	1	no	0.2746
[i-i*-ds]	87(4)-01(4)	0	no	0.0000
ln(1+D)	87(4)-00(3)	1	yes	0.0061
ln(1+C)	87(4)-00(3)	1	no	0.0116

The p-value indicates the fraction of bootstrap samples generating a value of f_1 greater than that for the true sample. In all cases except $\ln(w/s.w^*)_t$ we can reject the null at the 2% level. In these tests we have fitted the equation for the conflict variables $\ln(1+D)_t$ and $\ln(1+C)_t$ up to 2000q3, before the onset of the second *Intifada*. Figures 1-2 above indicate a definite structural break in these series from 2000q4 onwards, but this is too close to the end of our sample for a Perron (1989) unit root test to be sensible. We proceed on the assumption that the conflict variables are stationary around a break in 2000q4. Equation 1 is fitted in levels, without recourse to an error-correction isomorphism. Note that we do not take differences of $\ln(w/s.w^*)_t$, and the t-ratios on this variable should therefore be treated with some caution.[18]

Equation 1 cannot be estimated directly by OLS, because we cannot assume that the interest rate differential $[i-i^*-ds]_t$ is weakly exogenous to the capital flight variable $\ln(F_1)_t$. A larger capital outflow (i.e., a larger value of $\Delta\ln(F_1)_t$) might push up domestic interest rates or lead to depreciation

[18] The Schwartz Bayesian and Akaike information criteria for the regressions below indicate that levels of $\ln(w/s.w^*)_t$ are to be preferred to differences.

in the value of the Sheqel. Nevertheless, as noted above, we can reasonably assume long-run interest parity. In other words $[i\text{-}i^*\text{-}ds]_t$ is independent of $\Delta\ln(F_1)_t$ in the long run. So we can identify the parameters of equation (1) by employing the method of Blanchard and Quah (1989). Having fitted a conditional VAR of the form:

$$\begin{bmatrix} \ln(F_1)_t \\ [i-i^*-ds]_t \end{bmatrix} = \mathrm{A}\begin{bmatrix} g_1(L)\ln(F_1)_{t-1} \\ g_2(L)[i-i^*-ds]_{t-1} \end{bmatrix} + \mathrm{B}Z'_t + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix} \tag{4}$$

where Z_t is a vector of the other variables on the RHS of equation (1) and the u's are reduced-form residuals, and assuming that ε_t in equation (1) above is orthogonal to the innovation in the corresponding structural equation for $[i\text{-}i^*\text{-}ds]_t$, we can use the long-run independence restriction to recover the parameters in equation (1) from the parameters in the A and B matrices.

Table 4 reports the reduced-form parameter estimates for the $\ln(F_1)_t$ equation, along with corresponding standard errors and t-ratios. The lag order chosen on the basis of standard information criteria is 2. The quarterly sample runs from 1988 to 2001. The final column in the table shows the estimated coefficients in the structural equation (equation (1)). The diagnostic statistics in the table indicate that the coefficients are stable over time and robust to changes in sample size. This is confirmed by the recursive regression statistics shown in Figure 5.[19] In particular, there is no significant structural break during the Gulf War period (1990-91). Appendix B reports the corresponding $[i\text{-}i^*\text{-}ds]_t$ equation; none of the coefficients on the political variables is significant in the structural equation for $[i\text{-}i^*\text{-}ds]_t$, so we do not discuss it at any length here.

These estimates are based on the assumption that the political conflict variables are weakly exogenous to capital flight. It seems reasonable to assume that the level of violence during period t is independent of the level of the capital stock at the end of that period. Nevertheless, section 5 of the paper will present a model of conflict intensity. Using this model, it is possible to test for weak exogeneity in the capital flight equation using the approach of Engle and Hendry (1993). This test will be discussed later.

Table 4 shows that the coefficients on all of the explanatory economic variables take the expected sign, although the complexity of the dynamic response of $\ln(F_1)$ varies from one variable to another. A larger differential between domestic and foreign interest rates discourages capital flight and lowers $\ln(F_1)$, as does real exchange rate appreciation. An increase the in relative wage rate has the opposite effect. A higher rate of economic growth in the US also encourages capital flight. All of these effects are statistically significant at conventional confidence levels. Although there is some interaction between $\ln(F_1)$ and $[i\text{-}i^*\text{-}ds]$, the reduced-form coefficients are not all that

[19] Each point plotted for a given quarter in the figure indicates the test statistic for a regression on a sub-sample ending in that quarter. The Chow Tests are for forecasts from the end of each sub-sample to 2001q4.

Table 4: The Regression Equation for ln(F1) (1988(2)-2001(4))
The regression includes an intercept and trend.

Variable	***coeff.***	***s.e.***	***t ratio***	***h.c.s.e.***	***ptl. R^2***	***ins.***	***str. co.***
$\ln(F1)_{-1}$	0.36112	0.15211	2.374	0.16773	0.1387	0.04	*0.42552*
$\ln(F1)_{-2}$	-0.36322	0.15181	-2.393	0.15772	0.1406	0.04	*-0.33882*
[i-i*-ds]	———	———	———	———	———	———	*-0.28414*
$[i-i^*-ds]_{-1}$	-0.74027	0.29212	-2.534	0.29459	0.1550	0.09	*-0.69695*
$[i-i^*-ds]_{-2}$	0.22866	0.15196	1.505	0.13652	0.0608	0.13	*0.22870*
$\ln(p/s.p^*)_{-1}$	-0.23501	0.24286	-0.968	0.22209	0.0261	0.04	*-0.27248*
$\ln(p/s.p^*)_{-2}$	-1.47350	0.28555	-5.160	0.24216	0.4321	0.04	*-1.36790*
$\ln(w/s.w^*)_{-1}$	0.79851	0.20413	3.912	0.21863	0.3042	0.07	*0.74585*
$\ln(w/s.w^*)_{-2}$	0.10998	0.19697	0.558	0.16811	0.0088	0.08	*0.10064*
$\Delta\ln(y^*)$	4.11630	0.90216	4.563	0.98599	0.3730	0.09	*3.64030*
$\Delta\ln(y^*)_{-1}$	1.50990	1.18070	1.279	1.38370	0.0446	0.06	*1.18930*
$\Delta\ln(y^*)_{-2}$	3.31400	1.09000	3.040	0.93021	0.2089	0.20	*3.49720*
ln(1+D)	0.02238	0.00965	2.320	0.01053	0.1333	0.04	*0.02074*
$\ln(1+D)_{-1}$	0.01446	0.00888	1.629	0.00836	0.0705	0.04	*0.01374*
$\ln(1+D)_{-2}$	0.03261	0.00873	3.734	0.00822	0.2849	0.04	*0.02823*
ln(1+C)	-0.01440	0.00492	-2.923	0.00458	0.1963	0.04	*-0.01464*
$\ln(1+C)_{-1}$	0.00660	0.00564	1.170	0.00561	0.0377	0.06	*0.00449*
$\ln(1+C)_{-2}$	-0.01861	0.00461	-4.042	0.00461	0.3182	0.16	*-0.01523*
DUM	0.12452	0.05110	2.437	0.04518	0.1451	0.03	*0.11951*

h.c.s.e. = heteroskedasticity-corrected s.e.
ins. = Hansen (1992) parameter instability test statistic
σ = 0.03295 adjusted R^2 = 0.77440
Hansen (1992) variance instability test statistic: 0.12952
Hansen (1992) joint parameter instability test statistic: 3.66660
LM test for residual autocorrelation: F(1,34) = 0.60929 [0.4405]
LM ARCH test: F(1,33) = 0.00099 [0.9751]

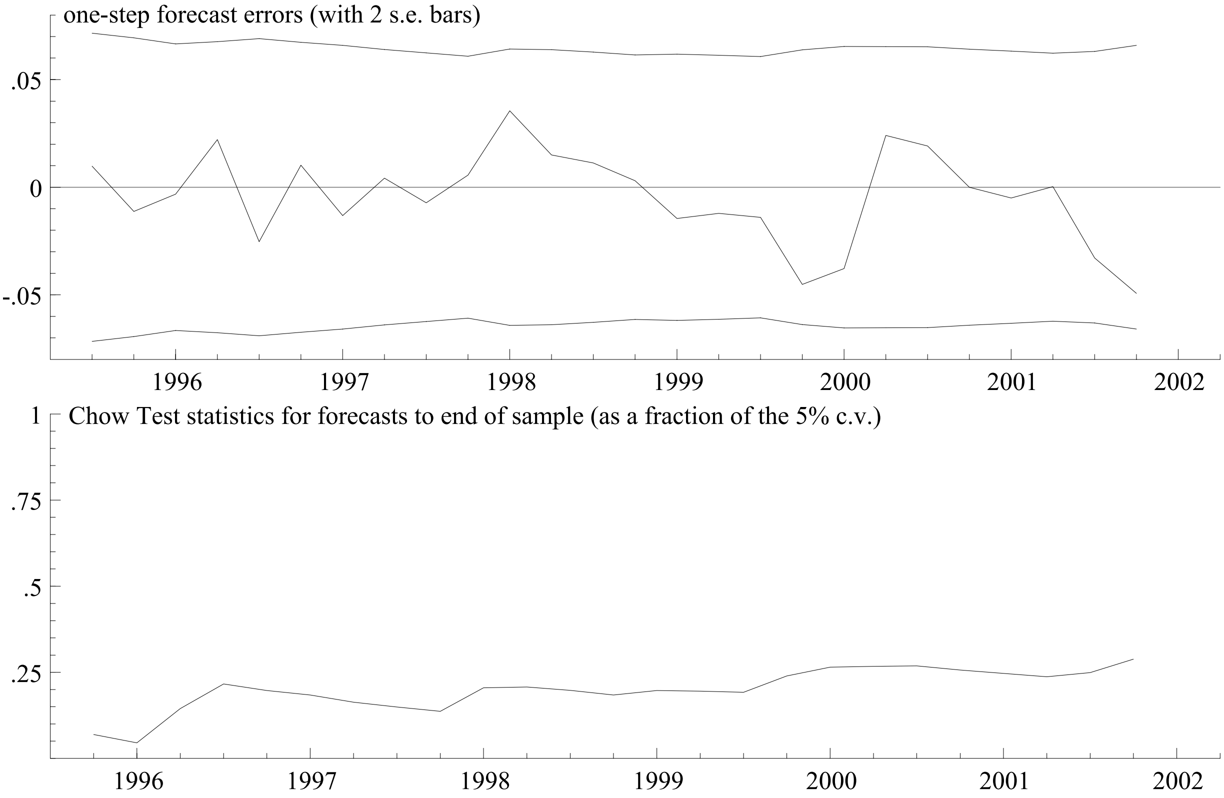

Figure 5: Recursive Regression Statistics for the ln(F1) Equation

different in size from the structural coefficients.

Moreover, increases in the level of political violence, as measured by the fatality variable D, induce a higher level of capital flight, *ceteris paribus*. Some of this increase happens within the quarter in which the intensification of violence occurs, although the significance of coefficients on lagged levels of $\ln(1+D)$ suggests that sustained periods of high violence have a much greater impact on capital flight than short-lived spikes in $\ln(1+D)$. However, increases in $\ln(1+C)$ lead to a reduction in $\ln(F_1)$. For a given level of violence, a more severe border closure policy seems to stem some of the capital flight. This suggests that investors perceive a more restrictive closure policy to be associated with greater security, perhaps in terms of the expected level of future violence. In section 5 below we will investigate whether such a perception is accurate.

The dummy variable for the second *Intifada*, last five quarters of the sample period, is also statistically significant. Capital flight has been over 10% higher than one might otherwise expect in this period, even taking into account the effect of fatalities and border closures. One explanation for this is that there are non-linearities in the effects of our conflict variables which cannot easily be estimated, because the period of greatest conflict happens just a few quarters from the end of our sample. However, it may also be that the effect of the dummy variable reflects general pessimism about the future stability of Israel stemming from increased hawkishness among leaders on both sides of the conflict, rather than the number of fatalities *per se*.

How large are the effects of the political conflict variables, relative to those of the standard economic variables? Table 5 gives some insight into this question, reporting the values of the solved-out long-run coefficients on each of the conditioning variables in the VAR. These coefficients should be treated with some caution: we are not claiming that all of these conditioning variables are strictly exogenous, so the table does not necessarily indicate the asymptotic response of $\ln(F_1)$ to shocks to each variable. Nevertheless, the table does indicate what would be the magnitude of the impact a of unit change in each conditioning variable in the long run with no feedback effects, and assuming that the other conditioning variables remain constant. Interpretation of these long-run coefficients makes more sense if we scale them by the sample standard deviation of the variable (around a trend); then we have a rough measure of the relative sizes of the long-run effects of "typical" changes in each of the conditioning variables. It turns out that these figures are roughly the same for each of the economic variables, a standard deviation change in each leading to a change in F_1 of around 6-7%. This is also the size of the impact of a standard deviation change in $\ln(1+D)$. In other words, a typical change in the level of violence has about the same impact on capital flight as a typical change in the real exchange rate or relative wages. The size of the impact of a typical change in the border closures level is somewhat smaller (about 3%) but still substantial.

Table 5: Solved-Out Long-Run Elasticities: ln(F1) Equation

variable	*std. dev. around trend*	*coeff.*	*s.e.*	*coeff. × std.dev.*
ln(p/s.p*)	0.04391	-1.55650	0.23474	-0.0684
ln(w/s.w*)	0.06345	0.87356	0.22986	0.0554
Δln(y*)	0.00745	8.60260	1.65460	0.0641
ln(1+D)	1.10186	0.07036	0.01411	0.0775
ln(1+C)	1.43474	-0.02403	0.01057	-0.0293
DUM	———	0.11351	0.06138	———

5. The Determinants of Conflict Intensity

Given that our measures of conflict intensity play such an important part in determining the level of capital flight, it is important to know whether variations in level of violence are entirely random, or whether they contain a deterministic, predictable component. There is a purely econometric reason for this exercise: to establish whether the level of violence is indeed weakly exogenous to capital flight. But also, the economic interpretation of the results presented in Tables 4-5 depends crucially on whether we think of changes in conflict intensity as strictly exogenous events that impact on the economic system, or whether there is some feedback from economic conditions to the level of violence.

At this point we should admit a limitation to our analysis: there is very little data on macroeconomic conditions in the West Bank and Gaza. Annual aggregate price and GDP figures have been constructed for some years in the late 1990s; but there are no consistent data prior to the inception of the Palestinian National Authority, even at annual frequency. Moreover, data collection for 2001 onwards has been severely disrupted by the increased conflict intensity. The one variable that is available on a quarterly basis, and that might reflect changes in economic conditions in the West Bank and Gaza, is the closures variable C. On those days when the border is closed, there is effectively a high unemployment level in the Palestinian labor force.[20] We do have very detailed quarterly macroeconomic data for Israel proper, but Israeli output, inflation and unemployment series have no explanatory power in regression equations for our conflict intensity measures. So there is a concern that any econometric model of the level of violence will suffer from omitted variables bias. Having said this, the regression results we discuss below turn out to be robust to changes in sample size, and the parameters appear to be stable over time, so we have some confidence that we have captured the most important factors influencing the intensity of the conflict.

[20] Some PNA data on number of person-days lost through closures are available for the late 1990s; these figures are of course highly correlated with C. However, such data are not available for the whole of our sample period. Since much of the employment of Palestinians in Israel proper is in the informal sector, official Israeli labor force statistics are of little use here.

Our conflict model embodies a combination of political and economic factors. First, we disaggregate total fatalities *D* according to the nationality of the victim, and construct two time series: ln(1+*P*) for Palestinian fatalities and ln(1+*I*) for Israeli fatalities.[21] We hypothesize that ln(1+*P*) and ln(1+*I*) depend on the following time-varying factors (the data sources are given in Appendix A). In each case, interpretation of an explanatory variable is subject to the caveats discussed in section 2 above.

(i) Each dependent variable's own past value. Shocks to the intensity of the conflict may persist for several periods, if past Palestinian deaths encourage greater Palestinian militancy now (causing a higher level of fatalities on both sides), and if past Israeli deaths encourage a more hard-line approach by the Israeli authorities now (causing a higher level of fatalities on both sides). In the absence of appropriate instruments, we will not include the contemporaneous level of ln(1+*I*) in the ln(1+*P*) equation, nor of ln(1+*P*) in the ln(1+*I*) equation, which should therefore be interpreted as reduced-form representations.

(ii) The number of border closures, ln(1+*C*), in previous quarters. The economic deprivation caused by closures might encourage greater Palestinian militancy, leading to conflict escalation. On the other hand, such deprivation might reduce some Palestinians' capacity to engage in militant activity: effective militancy might depend on a certain minimum income level.

(iii) The lagged rate of growth of Jewish settlements in the West Bank and Gaza, $\Delta\ln(B)$. If the rate of residential construction in the settlements increases, Palestinian militancy may also increase, leading to an increase in conflict intensity.

(iv) The lagged value of our FDI capital flight variable, $\ln(F_I)$. It is possible that the opportunity cost of violent activity is greater when more of people's physical capital is located in the conflict zone, and therefore at risk from Palestinian bombs or IDF bulldozers. If more capital is relocated abroad, perhaps because the current level of violence is higher, this may encourage even more violence in the future. In this case, it is the unconditional level of $\ln(F_I)$ that matters, and there is the potential for a hysteresis effect.

(v) The dummy variable for the second *Intifada*, DUM. We do not claim that the outbreak second *Intifada* can be explained in terms of the variations in variables (i-iv) above. Figure 1 above indicates that this represents are large structural break in the fatality series, due to political factors that cannot be quantified in a time-series model. A

[21] Greater disaggregation is possible, according to location and military status, but we have not been able to produce robust regression equations for more disaggregated series. In the analysis that follows we ignore the handful of fatalities

deterministic break in the intercept of the regression equations for 2000q4 onwards is required.

More formally, our regression specification is:

$$\begin{bmatrix} \ln(1+P)_t \\ \ln(1+I)_t \end{bmatrix} = \mathrm{M}\begin{bmatrix} h_1(L)\ln(1+P)_{t-1} \\ h_2(L)\ln(1+I)_{t-1} \end{bmatrix} + \mathrm{NY'}_t + \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} \tag{5}$$

where the v_{it} represent shocks to the level of violence and Y is a vector made up of variables (ii-v) above.

Table 6 summarizes our results. In our estimate of the parameters in equation (5), we have selected a lag order of 2 for all variables except $\ln(F_I)$ and $\Delta\ln(B)$, which are included with a lag order of 1; this minimizes the Schwartz Bayesian and Akaike information criteria. The diagnostic statistics in the table indicate that the coefficients are stable over time and robust to changes in sample size. This is confirmed by the recursive regression statistics shown in Figures 6-7. In both the ln(1+*P*) equation and the ln(1+*I*) equation, the dummy variable for the second *Intifada* is large and highly statistically significant, confirming the assumption that there is a deterministic break in the fatality series from 2000q4 onwards.

The table shows that the current levels of Palestinian and Israel fatalities both depend on the past level of Israeli fatalities. Interestingly, however, both Palestinian and Israeli fatalities appear to be more or less independent of past Palestinian fatalities: the lags of ln(1+*P*) are not significant in either equation at conventional confidence levels. *Ceteris paribus*, a 10% increase in Israeli fatalities is associated with an increase in Palestinian fatalities of around 2.5% within the next two quarters. Starting from the log-mean levels of *P* and *I* (22 and 5½ deaths respectively), this means that one extra Israeli death is on average associated with approximately one subsequent Palestinian death. One interpretation of these results (although in the absence of a fully identified structural model this is only a conjecture) is that the intensity of Palestinian activity in the conflict – protests or attacks on Israeli targets that lead to both Palestinian and Israeli fatalities – is driven by factors other than the magnitude of recent Palestinian fatalities. In contrast, the past level of Israeli fatalities does appear to influence current conflict intensity, in terms of both Palestinian and Israeli fatalities. A possible reason for this is that a higher Israeli fatality level provokes a subsequent increase in the intensity of IDF engagement with the Palestinians.

There are other factors that explain some of the variation in Palestinian fatalities, but that have no significant impact on Israeli fatalities. One of these is the rate of expansion of Jewish settlements in the Territories. A 10% increase in the rate of expansion is associated with a subsequent increase in Palestinian fatalities of about 5%. One interpretation of the result is that such

of foreign nationals during the conflict.

Table 6: The Fatality Regression Equations

Regression Equation for ln(1+P) (1988(2)-2001(4))
The regression includes an intercept and trend.

variable	***coeff.***	***s.e.***	***t ratio***	***h.c.s.e.***	***ptl. R^2***	***ins.***	***coeff. 2***
$\ln(1+P)_{-1}$	0.01927	0.12329	0.156	0.16428	0.0006	0.04	—
$\ln(1+P)_{-2}$	-0.00150	0.11176	-0.013	0.14190	0.0000	0.05	—
$\ln(1+I)_{-1}$	-0.02461	0.08020	-0.307	0.08620	0.0021	0.07	—
$\ln(1+I)_{-2}$	0.26362	0.08002	3.294	0.11132	0.1979	0.03	—
$\ln(1+C)_{-1}$	-0.13688	0.06162	-2.221	0.06797	0.1008	0.06	—
$\ln(1+C)_{-2}$	-0.14381	0.06178	-2.328	0.05501	0.1096	0.05	—
$\Delta\ln(B)_{-1}$	0.51072	0.15643	3.265	0.14347	0.1950	0.13	—
$\ln(F1)_{-1}$	4.37050	1.05170	4.156	1.25260	0.2818	0.04	2.9303
DUM	3.29360	0.49845	6.608	0.59964	0.4981	0.01	3.4046

h.c.s.e. = heteroskedasticity-corrected s.e.
ins. = Hansen (1992) parameter instability test statistic
σ = 0.49107 adjusted R^2 = 0.72902
Hansen (1992) variance instability test statistic: 0.11691
Hansen (1992) mean instability test statistic: 1.34241
LM residual autocorrelation test: F(1,43) = 0.09808 [0.7557]
LM ARCH test: F(1,42) = 2.00130 [0.1645]

Regression Equation for ln(1+I) (1988(2)-2001(4))
The regression includes an intercept and trend.

variable	***coeff.***	***s.e.***	***t ratio***	***h.c.s.e.***	***ptl. R^2***	***ins.***	***coeff. 2***
$\ln(1+I)_{-1}$	0.09294	0.13661	0.680	0.16548	0.0104	0.05	—
$\ln(1+I)_{-2}$	0.36324	0.13629	2.665	0.15881	0.1390	0.05	—
$\ln(1+P)_{-1}$	-0.27067	0.21000	-1.289	0.17538	0.0364	0.10	—
$\ln(1+P)_{-2}$	-0.32559	0.19036	-1.710	0.17006	0.0623	0.09	—
$\ln(1+C)_{-1}$	-0.07048	0.10495	-0.672	0.11580	0.0101	0.04	—
$\ln(1+C)_{-2}$	0.00162	0.10523	0.015	0.12755	0.0000	0.09	—
$\Delta\ln(B)_{-1}$	-0.27420	0.26644	-1.029	0.22813	0.0235	0.09	—
$\ln(F1)_{-1}$	5.31520	1.79140	2.967	1.39550	0.1667	0.08	6.0343
DUM	3.57550	0.84900	4.211	0.57904	0.2873	0.03	2.2168

h.c.s.e. = heteroskedasticity-corrected s.e.
ins. = Hansen (1992) parameter instability test statistic
σ = 0.83643 adjusted R^2 = 0.62909
Hansen (1992) variance instability test statistic: 0.129174
Hansen (1992) mean instability test statistic: 1.30387
LM residual autocorrelation test: F(1,43) = 0.56917 [0.4547]
LM ARCH test: F(1,42) = 0.07949 [0.7794]

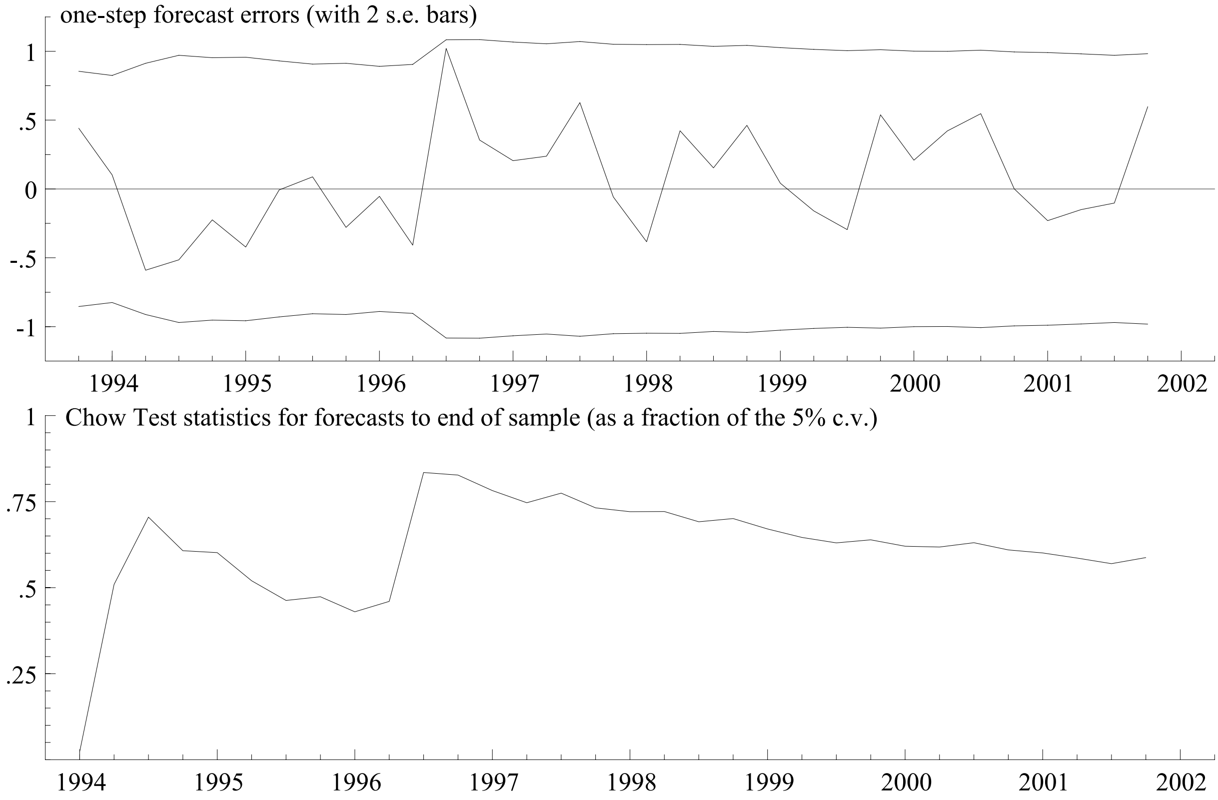

Figure 6: Recursive Regression Statistics: ln(1+P) Equation

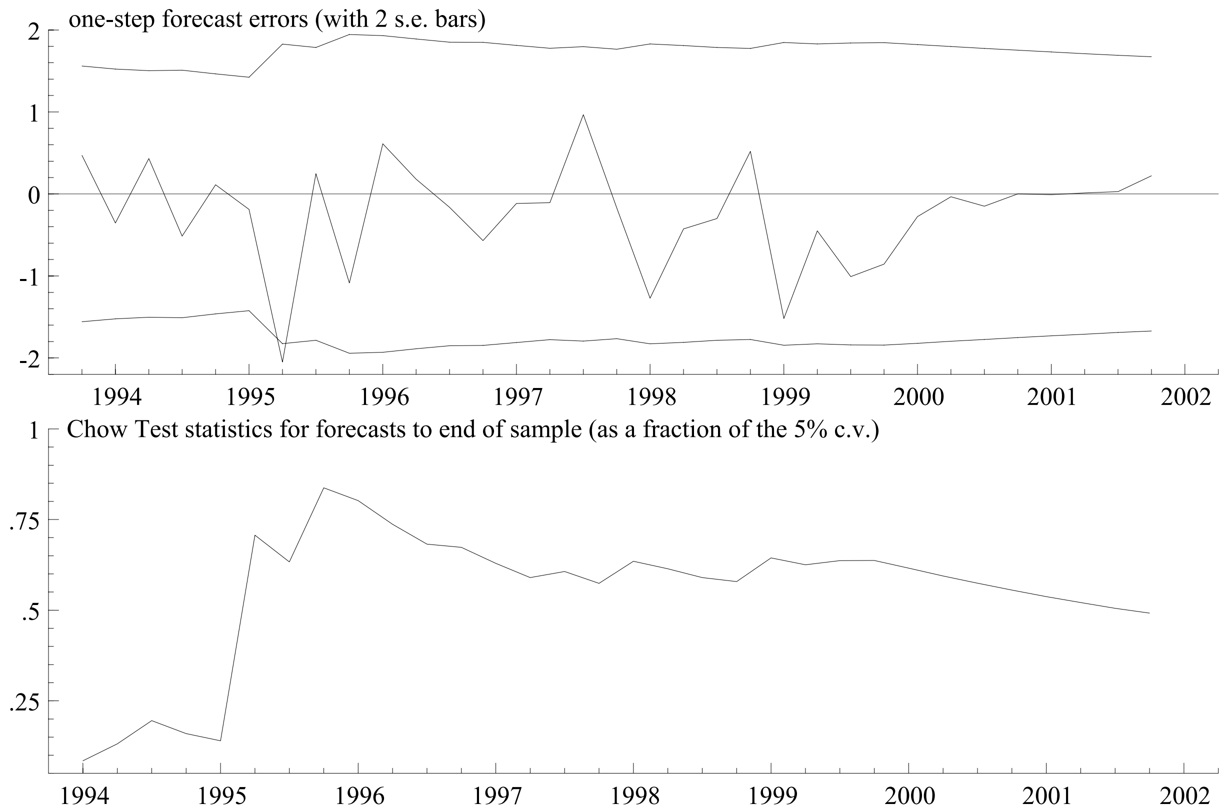

Figure 7: Recursive Regression Statistics: ln(1+I) Equation

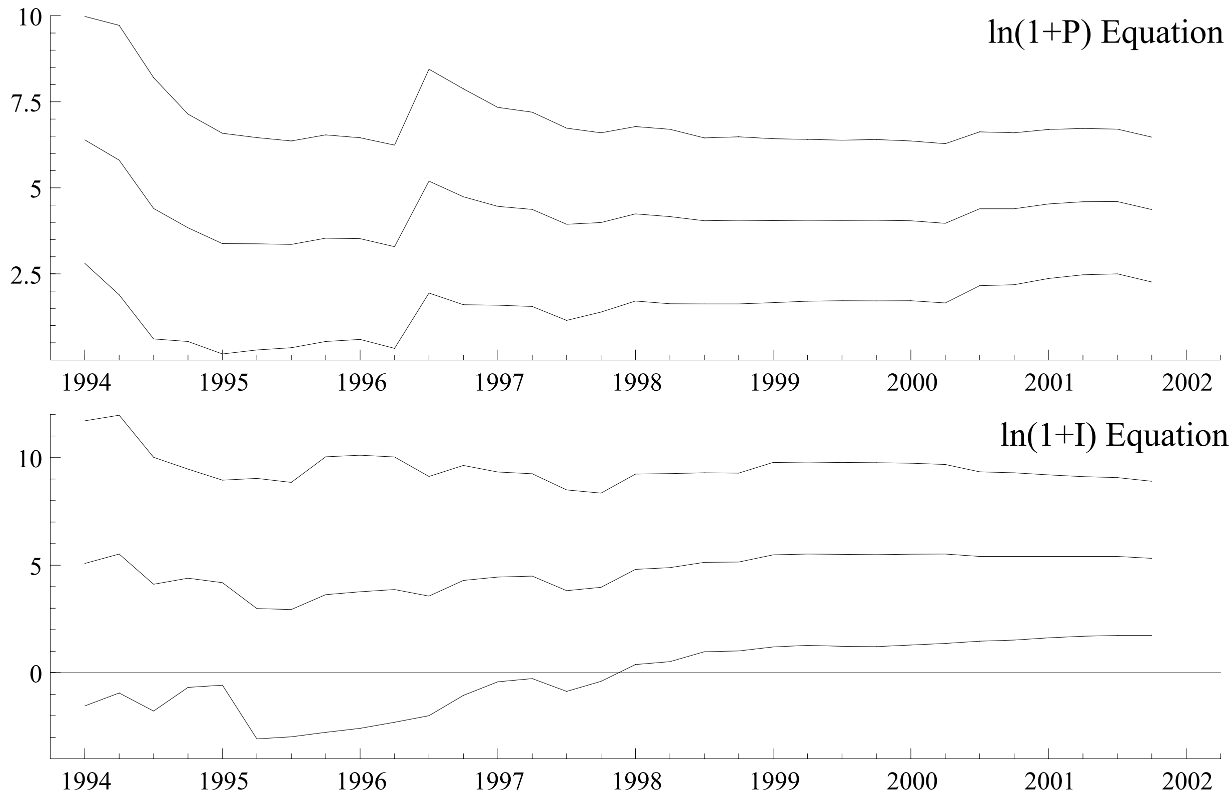

Figure 8: Recursive Estimates of the $\ln(F1)_{t-1}$ Coefficients (±2 Standard Errors) in the Fatality Regressions

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expansion provokes protest activity leading to more Palestinian deaths, but that it does not provoke significantly more attacks on Israeli targets *per se*. A second factor that affects the level of Palestinian fatalities is the number of border closure days in the quarter. A 10% increase in closures is accompanied by a *Palestinian* fatality level that is about 3% *lower* within the next two quarters. (The estimated impact on *Israeli* fatalities is also negative, but statistically insignificant.) One interpretation of this result is that the economic deprivation caused by border closures makes Palestinians less likely to engage in protests that lead to violent deaths, their greater sense of grievance being outweighed by the constraints of a lower income level.

Finally, the lagged capital flight variable $\ln(F_1)$ is highly significant in both equations. Both Palestinian and Israeli fatalities are higher after capital has relocated abroad. The estimated coefficients on this variable are 4.37 in the $\ln(1+P)$ equation and 5.32 in the $\ln(1+I)$ equation. The standard deviation of $\ln(F_1)$ around trend is 0.094, so a one standard deviation increase in $\ln(F_1)$ is associated with an increase in Palestinian fatalities of about 41% and in Israeli fatalities of about 50% in the next quarter. This large effect is robust to changes in sample size, and to the exclusion of all other non-deterministic explanatory variables in the regressions. (Figure 8 depicts recursive estimates of the coefficient on $\ln(F_1)$, and Table 6 includes the estimated coefficients on $\ln(F_1)$ –

"coeff. 2" – when other variables are excluded from the regression.) It appears that the level of violence in any given quarter is highly sensitive to the location of domestic residents' physical capital at the end of the previous quarter. The most straightforward explanation for this effect is that those engaged in conflict are more willing to escalate the level of violence when capital has been relocated abroad. For the reasons discussed in section 2 above, it is difficult to infer from the fatality data anything about the propensity of any one group to initiate a higher level of conflict; what we can say is that in aggregate, capital relocation impacts on the level of violence.

A potential alternative explanation of the significance of the lagged capital flight term in the regression is that it embodies information about the future level of the conflict not contained in our measured variables. Perhaps investors have access to informal, non-quantifiable information that helps them to predict the future level of violence, and which also affects the volume of capital flight. If so, then the lagged residual from the Table 4 model should be a more appropriate explanatory variable than lagged capital flight itself. However, if this residual is added to the Table 6 regressions it is statistically insignificant.

Having fitted a model of fatality levels conditional on predetermined regressors, it is possible to construct a test of the weak exogeneity of the fatality variable $\ln(1+D)$ in the capital flight regression reported in Table 4. The fitted values of $\ln(1+P)$ and $\ln(1+I)$ can be combined to create a fitted time-series for $\ln(1+D)$. This series can then be added to the Table 4 regression equation. The t-ratio on this additional regressor constitutes a test of the null hypothesis that $\ln(1+D)$ is indeed weakly exogenous to $\ln(F_1)$ (Engle and Hendry, 1993). The t-ratio in this case is –1.041, so the null cannot be rejected.

6. Summary and Conclusion

Using quarterly macroeconomic data, we have shown that there is a strong association between Israeli investors' decisions about where to locate their physical capital assets (inside or outside Israel) and the level of political violence arising from the Palestinian-Israeli conflict. Increases in conflict intensity – which appears to be weakly exogenous to the macroeconomic variables – lead to substantial capital flight: that is, Israeli investors increase the fraction of their physical assets located abroad. Moreover, capital flight in the current quarter leads to increased conflict intensity in subsequent quarters. These effects are both statistically significant and quantitatively important. The order of magnitude of the effects of the interaction of political and economic variables is at least as great as the "normal" economic interactions.

One consequence of this result is that any estimate of the economic costs of the *Intifada* that focuses solely on security expenditure and direct damage caused to property is likely to underestimate the true costs substantially. In order to put a precise figure on the costs associated with capital flight, we need to estimate the marginal product of capital within Israel (relative to the international rate of return on financial capital), which is beyond the scope of the current paper. We suspect that local physical capital productivity is higher than the international return, because the political conflict discourages some types of investment at the margin, but the magnitude of the wedge is still to be estimated. What we can say already is that even a small wedge between local capital productivity and the international return will imply a large value for conflict-related costs.

Even more striking is the sensitivity of the intensity of the conflict (as measured by the number of casualties of different kinds) to recent capital movements. Those engaged on one or both sides of the *Intifada* violence do appear to respond to changes in the opportunity cost of conflict escalation. So far, we have not found data that allow us to identify precisely how this opportunity cost impacts on the propensity of the combatants pro-actively to engage with the other side. All we can say is that there is a strong link between economic factors and the magnitude of the eventual conflict outcome, in terms of the number of casualties.

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Appendix A: Data Sources and Data Construction

variable	*construction*
F_{1t}	$\left\{s_t \cdot \sum_{\tau=0}^{\tau=t}(1-\delta_\tau *)\cdot(I_\tau * / P_\tau *)\right\} / \left\{K_t - \sum_{\tau=0}^{\tau=t}(1-\delta_\tau)\cdot(I_\tau / P_\tau)\right\}$
F_{2t}	$\left\{(s_t / P_t)\cdot \sum_{\tau=0}^{\tau=t} Z_\tau \right\} / \left\{K_t - \sum_{\tau=0}^{\tau=t}(1-\delta_\tau)\cdot(I_\tau / P_\tau)\right\}$

variable	*definition (all variables are quarterly)*	*source*
δ	rate of depreciation of private sector Israeli capital	BOI
δ*	rate of depreciation of US business sector capital	FRED
I	net foreign direct investment in Israel (in Sheqels)	CBS
I*	net foreign direct investment by Israelis (in Dollars)	CBS
P	Israeli private sector investment deflator (1995=100)	BOI
$P*$	US business sector investment deflator (1995=100)	FRED
Z	net foreign portfolio investment by Israelis (in Dollars)	CBS
K	real (1995) private sector Israeli capital stock (in Sheqels)	BOI
i	three-month Israeli treasury bill interest rate	BOI
$i*$	three-month US treasury bill interest rate	FRED
s	Shequels per Dollar spot exchange rate	BOI
p	Israeli consumer price index (1995=100)	CBS
$p*$	US consumer price index (1995=100)	FRED
w	Israeli aggregate industrial wage index (1995=100)	CBS
$w*$	US aggregate industrial wage index (1995=100)	FRED
$\Delta\ln(y)$	rate of growth of real Israeli GDP	CBS
$\Delta\ln(y*)$	rate of growth of real US GDP	FRED
D	total *Intifada* fatalities	B'Tselem
P	Palestinian *Intifada* fatalities	B'Tselem
I	Israeli *Intifada* fatalities	B'Tselem
C	number of days on which the Israel - West Bank or Israel - Gaza border is closed	B'Tselem
$\Delta\ln(B)$	Rate of growth of number of residential buildings completed in Jewish West Bank settlements in the previous quarter	CBS

BOI = Bank of Israel; CBS = Israeli Central Bureau of Statistics; FRED = Federal Reserve Bank

Appendix B: The Regression Equation for [i-i*-ds] (1988(2)-2001(4))

The regression includes an intercept and trend.

variable	***coeff.***	***s.e.***	***t ratio***	***h.c.s.e.***	***ptl. R^2***	***ins.***	***str. co.***
ln(F1)	———	———	———	———	———	———	*-0.31186*
$\ln(F1)_{-1}$	0.22663	0.13098	1.730	0.12467	0.0788	0.02	*0.33925*
$\ln(F1)_{-2}$	0.08588	0.13073	0.657	0.13431	0.0122	0.02	*-0.02740*
$[i-i^*-ds]_{-1}$	0.15246	0.25156	0.606	0.24618	0.0104	0.16	*-0.07841*
$[i-i^*-ds]_{-2}$	0.00014	0.13086	0.001	0.13633	0.0000	0.09	*0.07145*
$\ln(p/s.p^*)_{-1}$	−0.13188	0.20914	−0.631	0.19272	0.0112	0.04	*-0.20517*
$\ln(p/s.p^*)_{-2}$	0.37186	0.24590	1.512	0.27250	0.0613	0.03	*-0.08767*
$\ln(w/s.w^*)_{-1}$	−0.18531	0.17578	−1.054	0.18900	0.0308	0.03	*0.06371*
$\ln(w/s.w^*)_{-2}$	−0.03288	0.16962	−0.194	0.17883	0.0011	0.06	*0.00142*
$\Delta\ln(y^*)$	−1.67540	0.77688	−2.157	0.79189	0.1173	0.04	*-0.39171*
$\Delta\ln(y^*)_{-1}$	−1.12840	1.01670	−1.110	1.12870	0.0340	0.11	*-0.65751*
$\Delta\ln(y^*)_{-2}$	0.64463	0.93864	0.687	0.89330	0.0133	0.07	*1.67810*
ln(1+D)	−0.00579	0.00831	−0.697	0.00876	0.0137	0.03	*0.00119*
$\ln(1+D)_{-1}$	−0.00253	0.00765	−0.331	0.00794	0.0031	0.02	*0.00198*
$\ln(1+D)_{-2}$	−0.01541	0.00752	−2.049	0.00746	0.1071	0.03	*-0.00524*
ln(1+C)	−0.00086	0.00424	−0.203	0.00336	0.0012	0.08	*-0.00535*
$\ln(1+C)_{-1}$	−0.00740	0.00485	−1.525	0.00607	0.0623	0.07	*-0.00535*
$\ln(1+C)_{-2}$	0.01191	0.00397	3.004	0.00428	0.2050	0.06	*0.00611*
DUM	−0.01763	0.04400	−0.401	0.03447	0.0046	0.01	*0.02120*

$\sigma = 0.02838$ adjusted $R^2 = 0.78483$
Hansen (1992) variance instability test statistic: 0.45645
Hansen (1992) mean instability test statistic: 3.24159
LM residual autocorrelation test: F(1,34) = 0.42920 [0.5168]
LM ARCH test: F(1,33) = 0.05744 [0.8121]