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Visa waivers, multilateral resistance and international tourism: some evidence from Israel

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Abstract This paper tests the visa-led tourism hypothesis which contends that easing of visa restrictions increases international tourism. Israel acts as a natural laboratory in this case with clear before and after junctures in visa restrictions. We use panel data on tourism to Israel from 60 countries during 1994–2012. In contrast to previous work we take account of nonstationarity in the data and test for the effect of multilateral resistance on tourism. Partial waivers of visa restrictions are estimated to increase tourism by 48 % and complete waivers increase tourism by 118 %. Other results include the adverse effect of Israel's security situation on tourism, the beneficial effect of real devaluation on tourism, and the fact that the elasticity of tourism to Israel with respect to tourism to all destinations is very small.

Keywords International tourism · Visa arrangements · Security · Multilateral resistance

JEL Classification C23 · R23

1 Introduction

Visa restrictions can impose substantial costs to economic interaction. Although the effects of visa requirements on migration, investment and trade have been investigated

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(Bertoli and Fernandez-Huertas Moraga 2012, 2013; Neumayer 2006, 2011), the literature on their impacts on tourism is surprisingly thin (Neumayer 2010; Artal-Tur et al. 2013). If tourism serves as an important conduit for social and economic change (Marrocu and Paci 2011), the imposition of visa restrictions may have wider social and developmental effects beyond the simple screening of visitors. Conversely, the introduction of visa waivers may be responsible for jump-starting tourism-based economic growth.

This paper examines this under-researched driver of tourism. Common to earlier work, our interest lies in identifying the catalytic effect of visa waivers on tourism and estimating its magnitude. However, we progress beyond current research by addressing some of the methodological issues involved in the relation between visa restrictions and tourism. These include the nonstationarity inherent in tourism data, identification problems that arise when using panel data, and multilateral resistance (Anderson and Van Wincoop 2004) or third-country effects on bilateral tourism. Multilateral resistance (MLR) implies that pairwise tourism from origins to destinations does not just depend on the attributes of the destinations but also on the attributes of alternative destinations. We use appropriate econometric methods for these challenges.

Visa restrictions can be considered as a screening mechanism that hampers the free flow of tourism. They would seem to deflect demand to countries with lower barriers to entry. However, things may not be so simple, since the effect of visa waivers may be dependent on the outcome. If visa restrictions are eased for those origins where tourism is large in the first place, the causal effect of visa waivers on tourism will not be identified.

We use Israel as a case study to investigate this relationship. The Israeli tourist industry directly accounts for about 2 % of GDP and its total contribution (direct and indirect) is estimated at 8 % of GDP (WTTC 2013; Freeman and Sultan 1997). The number of overseas tourists grew from 1.06 m in 1980 to 2.88 m in 2012, or 3.2 % per year, which is less than the rate of growth of the economy. Additionally, the composition of visitors over this period underwent a significant change. In the 1980's Western European (EU15) and North American tourists accounted for 27 and 56 % of all visitors respectively. By 2012, these shares were 34 and 23 % respectively. In contrast, the share of non-EU European tourists rose from 4.5 to 25 % over the corresponding period. Of this latter share, Russia and Ukraine alone contributed 70 %. Finally, Asian tourism grew over this period from 3.5 to 9 %.

During this period bilateral visa waivers doubled from 24 to 49. While we are only interested in tourism to a single destination, it should be noted that the removal of visa restrictions is generally characterized by reciprocity. Neumayer (2011) finds that in 48 % of country-pairs visa restrictions are imposed bilaterally. In contrast, in 35 % of country-pairs the restriction is unilateral. More developed countries impose greater restrictions on visitors from less developed countries than vice versa. Visitors from developed countries face visa restrictions in 90 foreign destinations while their counterparts from less developed countries face visa restrictions in 156 countries.

We hypothesize that the removal of visa restrictions explains changes in the level and composition of tourism. For testing this visa-led tourism hypothesis (VTH), Israel provides a setting with clear before and after junctures as bilateral agreements on visa waivers come into force. Unlike, for example, EU countries, Israel is not part of a multilateral visa system, and determines its visa policy on a purely bilateral basis. Using econometric methods designed for nonstationary panel data, and treating visa waivers as structural breaks, we show that visa waivers have very large causal effects on tourism to Israel. On the other hand security problems in Israel have major adverse effects on tourism, especially from western countries, the elasticity of tourism to Israel with respect to total tourism turns out to be small, and the demand for tourism increases when the real exchange rate is devalued.

2 Literature review

Visa regulation is a mechanism for managing the trade-off between the security concerns of full accessibility on the one hand and the economic benefits of full accessibility, on the other hand. Restrictions on visas tend to be more stringent in developed countries wary of illegal immigration or terrorism and less restrictive in developing countries more dependent on trade and tourism. In some national contexts, visas are a source of income generation, financing the operation of embassies in host countries. Visa arrangements are not necessarily dichotomous (waiver or no waiver) and various intermediate screening processes exist. These include visa permits after screening processes by local embassies or consulates, or more stringent screening leading to visa authorization by the relevant agency at the destination. While embassies and consulates would be expected in this instance to inhibit tourism, Gil-Pareja et al. (2007) claim this is outweighed by their positive effect on generating information and promoting visitors. Using a cross sectional gravity approach, they estimate that this effect accounts for 15–30 % of tourism. They do not, however, deal with the latent identification issue here. It may be that embassy and consulate activity is a result rather than a determinant of the demand for tourism.

The effect of visa restrictions has been examined in relation to trade, foreign direct investment (Neumayer 2011) and immigration (Bertoli and Fernandez-Huertas Moraga 2012, 2013). In these studies efforts have been made to estimate the net effects of visa restrictions discounting correlated but unobservable factors that also shape the visa regime, such as the need to scrutinize illegal immigrants and security risks. Bertoli and Fernandez-Huertas Moraga (2012) estimate that roughly half of the 'cliff' that restricts cross border immigration can be attributed to visa restrictions. In Bertoli and Fernandez-Huertas Moraga (2013) third-country effects are addressed in the context of multilateral resistance. This is an issue in both migration and tourism research because visa restrictions in third countries can distort immigration, trade or tourism. For example, difficulty in getting a visa to Israel may deflect potential tourists to Cyprus or Egypt where visa restrictions may be lower. Bertoli and Fernandez-Huertas Moraga (2013) go beyond Anderson and Van Wincoop's (2004) use of dummy variables to capture multilateral resistance by allowing for cross-section dependence between panel units using Pesaran's (2006) common correlated effects (CCE) estimator and specify a common factor model that relates pairwise flows (in migration or tourism) from origins to destinations.

Neumayer (2010) is possibly the first empirical study of visa regulations on tourism. Due to the time invariant nature of visa permits, the estimation strategy adopted is to deal with identification problems by using country specific dummy variables and dyadic covariates likely to be correlated with visa constraints. Using a panel gravity model with country fixed effects, he estimates the effect of visa restrictions in reducing bilateral tourism at between 52–63 %, with a larger effect on tourism to/from developing countries. Neumayer controls for unobservable heterogeneity across pairs of countries, but does not account for nonstationarity in the data, or the effect of multilateral resistance on international tourism.

Artal-Tur et al. (2013) also use a fixed effects gravity model and estimate a significant but smaller impact of visa restrictions—a 23 % reduction in tourism. They find no difference in this effect between developed and developing countries. This more modest result is attributed to the treatment of unobservable heterogeneity between pairs of countries by controlling for country pairwise fixed effects. However they only use two points in time thereby ignoring the timing of the lifting of visa restrictions on tourism.

In the present study we confront a number of methodological challenges in estimating the causal effect of visa policy on tourism. First, the use of panel data with specific effects implicitly identifies the causal effect of visa policy on tourism by using differences-in-differences. It compares changes in tourism from treated countries for which visas are waived relative to changes in tourism from untreated countries for which visa policy remained unchanged. Second, since tourism is nonstationary we use panel cointegration (Pedroni 1999, 2004) to estimate the treatment effect on tourism of visa policy. Third, we allow for multilateral resistance in tourism by hypothesizing dependence between the origin countries in the panel using the observed common factor model of Pesaran (2006). Fourth, we treat the easing of visa restrictions as a structural break at a known date in the panel relationship (Westerlund and Edgerton 2008). However, because the common factor is nonstationary we use critical values for panel cointegration calculated by Banerjee and Carrion-I-Silvestre (2011, 2013). Fifth, we take account of the effects of terror-related threats to personal security that adversely affect tourism, and which may confound the estimation of the effect of visa policy. The extant literature on the effects of terror on tourism in Israel suggests that they are short term and non-uniform. International tourism to Israel is price elastic, income inelastic and moderately sensitive to terror over the short run (BOI 2014). Consequently international tourism demand can be met by alternative destinations. In contrast, domestic Israeli tourism demand is price inelastic, income elastic and less sensitive to security issues (Fleischer and Buccola 2002).

3 Data description

We compiled annual data for a panel of 60 tourist origins to Israel over the period 1994–2012. The source and construction of the variables used in the analysis can be found in the data appendix (Appendix1). Trends in this data are now described.

Over the long term (1970–2012) international tourism to Israel has a distinct trend (Fig. 1). The period investigated here captures the sharp dip in tourism due to the Second Intifada (Palestinian uprising) during 2000–2005, which smokescreens the overall long-term trend. As can be seen, by the end of the period international tourism

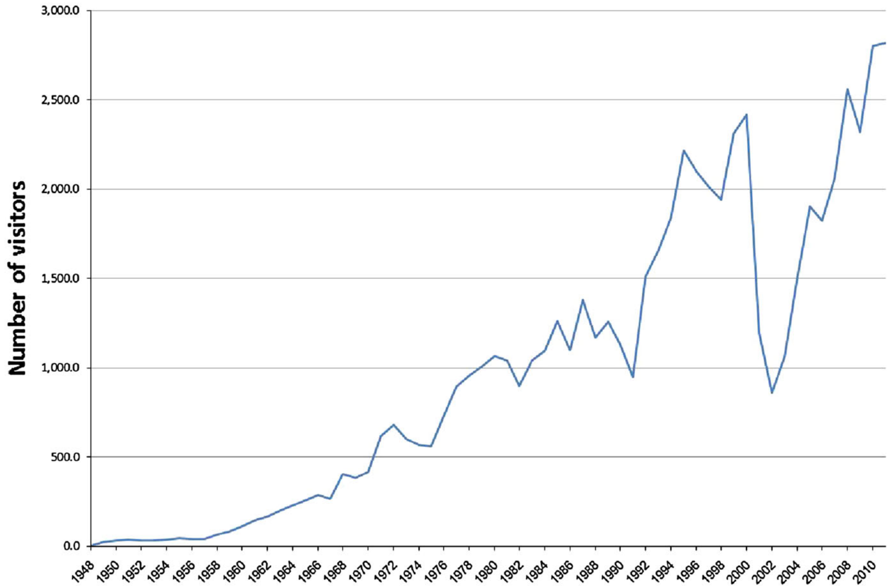

Fig. 1 Foreign tourists to Israel 1949–2012 *Source* CBS (2012)

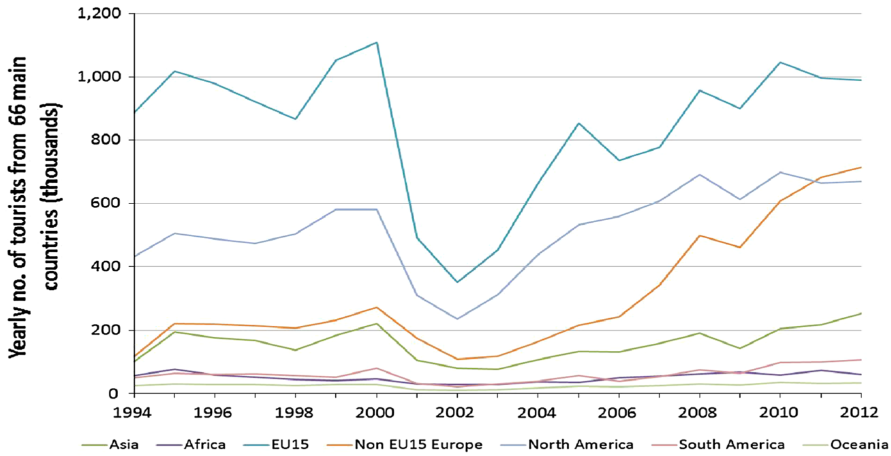

Fig. 2 Foreign tourism to Israel by Continent, 1994–2012

recovered to, and in the aggregate surpassed, its pre-2000 level (Fig. 2). The volume of international tourism in the origin counties is also upwards trending and is stationary in first differences (Fig. 3).

Cumulatively visa restrictions have eased over time. Figure 4 shows the growth in bilateral visa agreements since 1994. There are three different levels of visa arrangements: full waiver (48 countries), visa authorization issued in the place of origin by the local Israeli consulate (12 countries) and visa issued at the place of destination (5 countries). The latter involves the most stringent level of screening. The number of

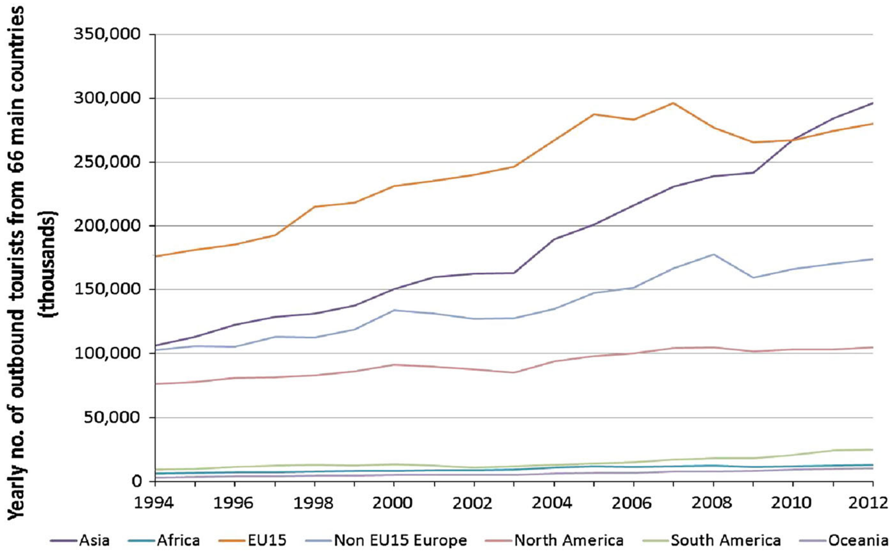

Fig. 3 Foreign tourism by Continent 1994–2012

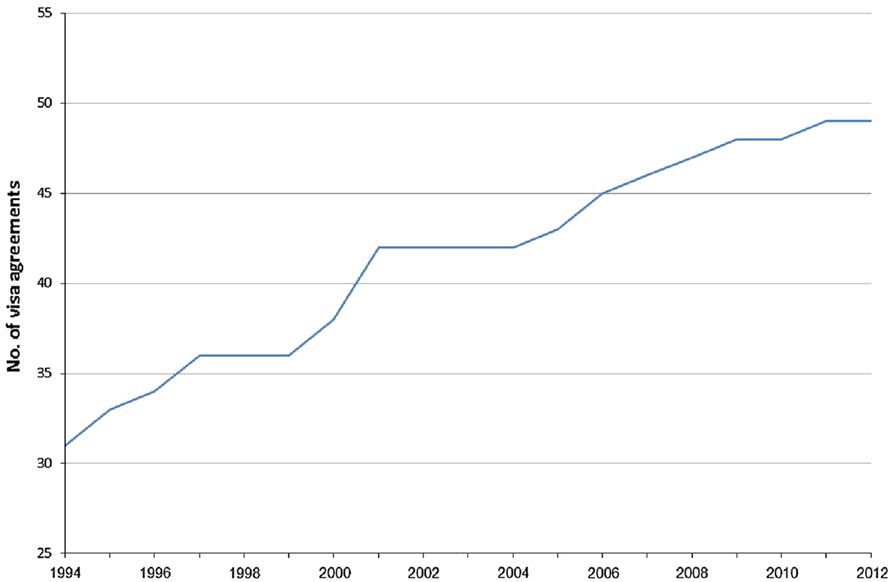

Fig. 4 Cumulative bilateral visa arrangements 1994–2012

bilateral visa agreements signed per year is charted in Fig. 5. Visa agreements have been characterized by two distinct waves of activity. The first occurred during 1964–1971 when 17 countries signed visa waivers with Israel, 10 of these prior to the Six Day War 1964–1967. The second period is 1993–2001 when a further 15 visa waivers were endorsed. During 1973–1993 there was very little visa waiver activity. Therefore,

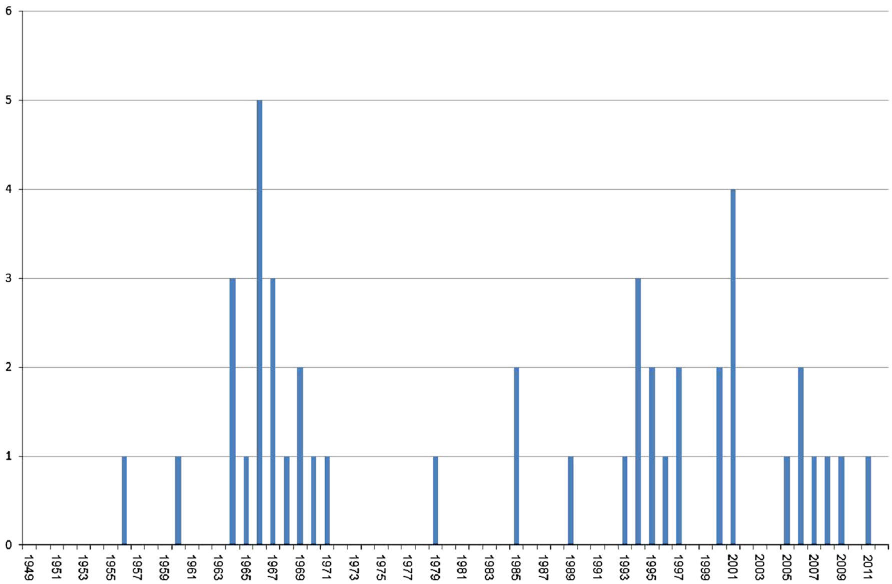

Fig. 5 Number of bilateral visa waivers signed per year, 1949–2012

our observation period includes, as appropriate, countries which received visa wavers and countries which did not. However, there are no countries in the data for which visa restrictions were increased. Therefore, strictly speaking the visa effect refers to waivers rather than visa restrictions.

In the absence of price indices for tourism, we use the real exchange rate to measure price competition in tourism. The real exchange rate is hypothesized to increase the demand for tourism because real devaluation of the shekel with respect to origin currencies lowers the relative price of tourism in Israel. We use terror or security related incidents recorded by the Israel Defense Force to capture the effects of the security situation on the demand for tourism. Most definitions of "terror" are based on fatalities (Gould and Klor 2010; Jaeger and Paserman 2008). By contrast the measure we use has a broader coverage and includes all incidents including those in which nobody is killed (see also Eckstein and Tsiddon 2004). This variable is expected to be inversely related to tourism. As can be seen, this variable 'erupts' from time to time and therefore cannot be considered as having a natural order of integration (Fig. 6). It may not be regarded as a structural break, but it may confound the estimation of the parameters of interest. In any case the effect of terror on tourism is of interest in its own right.

Finally, we use data on tourism to all destinations from the 60 origins to scale the demand for tourism to Israel. Given everything else, tourism from these origins to Israel is expected to vary directly with this variable. A parameter of interest is the elasticity of demand for tourism to Israel with respect to this scale variable. If Israel's share of tourism is constant over time, this elasticity is expected to be unity.

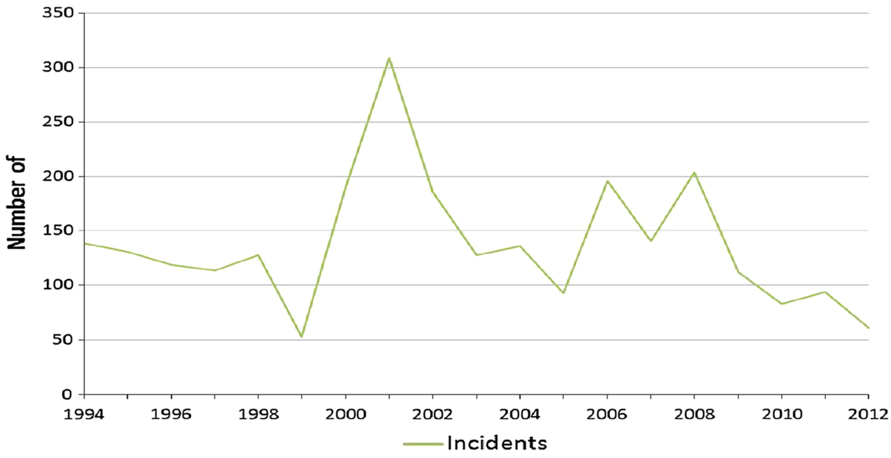

Fig. 6 Security incidents, 1994–2012

4 Methodology

The main hypothesis is:

$$\ln Y_{it} = \phi_i + \pi \ln T_{it} + \beta \ln r_{it} + \lambda S_t + \gamma V_{it} + \theta_i F_t + u_{it} \tag{1}$$

where Y_{it} denotes tourism to Israel from origin i in year t, ϕ is an origin fixed effect, T_i denotes tourism to all destinations from origin i, r_i denotes the real exchange rate between the shekel and the currency of origin i, S_t denotes the number of security incidents in Israel in year t, V_i is a dummy variable which is zero if no visa requirements apply to country i, rising to 2 if full visa requirements apply. F denotes a vector of observed common factors that induces dependence between the panel units. In Pesaran's common correlated effects (CCE) estimator (Pesaran 2006) these common factors include the cross-section averages of the panel data used to estimate Eq. (1), i.e. Y, T and r. Finally, u denotes model errors, which may be cross-section dependent, with correlation ρ_{ij}. Theory predicts that π and β are positive whereas λ and γ are negative.

We use Bartlett and Breusch–Pagan chi-square test statistics for cross-section dependence in the model errors, which are based on the squares of the 1/2N(N-1) = 1,770 pairs of correlations in the latter and the determinant of the correlation matrix in the former. The smaller these test statistics the greater is the degree of cross-section correlation. Cross-section dependence may be weak or strong (Chudik et al. 2011). In the former case it is spatial and localized whereas in the latter case it is pervasive. We use Pesaran's (2015) CD test for weak cross-section dependence, which is based on the average correlation (ρ-bar) between pairs of residuals (ρ_{ij}). Strong cross-section dependence does not necessarily mean that ρ-bar is large because pervasive correlations might still be small. For example, in our data in which N = 60 and T = 18 the critical value of ρ-bar is only 0.015.

In principle, strong cross-section dependence should be absorbed by the specification of common factors such as F. However, in practice these common factors might not absorb all of the cross-section dependence. Weak cross-section dependence should be absorbed by the specification of spatial dynamics (Anselin 1988). Cross-section dependence may be strong and weak at the same time. For example, in Eq. (1) the common factors induce strong cross-section dependence while the residuals are spatially autocorrelated.

We show below that Y, T and r are I(1). V may be regarded as a structural break the date of which is known. Since it refers to partial and full visa requirements there may be two structural breaks. The data generating process for security incidents (S) is naturally complex. It increases sharply during periods of geopolitical tension, otherwise it remains stable and small. We do not classify it as an integrated variable.

Equation (1) is panel cointegrated when the residuals are stationary, i.e. $u \sim \mathrm{I}(0)$. If $\lambda = \gamma = \theta = \boldsymbol{\rho} = 0$ the grouped augmented Dickey Fuller (GADF) statistic (Pedroni 1999, 2004) may be used to test panel cointegration since the panel units are independent and that there are no structural breaks. Banerjee and Carrion-I-Silvestre (2011) have calculated critical values for panel cointegration for the CCE model in which $\lambda = \gamma = 0$ in Eq. (1). They have also calculated critical values for the case in which $\lambda = \theta = 0$, and the cross-section correlation is strong (Banerjee and Carrion-I-Silvestre 2013). Cross-section correlation (ρ), structural breaks (γ) and common factors (θ) naturally increase the severity of panel cointegration tests, i.e. the critical value of GADF (average of ADF statistics for u) becomes more negative.

Residual-based panel cointegration tests are calculated using the formula:

$$z = \frac{\sqrt{N}(\bar{t} - E(\bar{t}))}{\sqrt{v}} \sim N(0, 1) \tag{2}$$

where N denotes the number of panel units, $\bar{t}$ denotes the average of the ADF statistics of the residuals for each panel unit, $\mathrm{E}(\bar{t})$ is the expected value of $\bar{t}$ and v is its variance. In the CCE case Banerjee and Carrion-I-Silvestre (2011) calculate $\mathrm{E}(\bar{t}) = -1.535$ and $\mathrm{v} = 0.341$ and the critical value ($\mathrm{p} = 0.05$) of $\bar{t} = -2.29$ when $\mathrm{N} = 50,\ \mathrm{T} = 20$ and there are two covariates. They suggest that these critical values also apply to the case in which Y is trend stationary, the covariates (T and r) are driftless random walks, the cross-section dependence in the residuals is strong, and there is a single structural break which occurs at a common known date (Banerjee and Carrion-I-Silvestre 2013). We refer to these critical values by BCS1. If there are two equidistant structural breaks $\mathrm{E}(\bar{t})$ is approximately -1.822 and $\mathrm{v} = 0.339$. We refer to these critical values by BCS2, which have been calculated for the case in which there is only one covariate.

Unfortunately, neither BCS1 nor BCS2 match the particular requirements of the estimates of Eq. (1). In our case the data are difference stationary rather than trend stationary. However, we do not expect that this will make much difference to BCS2. The difference between BCS1 and BCS2 is induced by structural breaks in the latter and observed common factors in the former. Therefore, below we use these critical values for indicative purposes only.

Since the parameter estimates of cointegrating vectors generally have nonstandard distributions, we do not report their standard errors. Since t-tests, F tests and chi square tests are misleading under these circumstances, hypotheses tests may be carried out by investigating the implications of restrictions on cointegration tests. For example, if a model ceases to be cointegrated when a variable is omitted, the variable in question is statistically significant. Or if it continues to be cointegrated but the p-value of the model increases, it should not be omitted. If, however, the p-value remains unchanged the variable should be omitted.

Had the data been stationary it would have been necessary to take account of cross-section dependence in the residuals to compute the standard errors of the parameter estimates for purposes of hypothesis testing. When the data are nonstationary cross-section dependence is only important for determining the critical values of panel cointegration tests such as BCS1 and BCS2.

5 Results

5.1 Unit root tests

The data depicted in Figs. 1, 3 and 4 suggest that international arrivals, outbound tourism and real exchange rates have positive time trends and are therefore nonstationary. We use panel unit root tests for independent (IPS) and dependent (CIPS) panel data to show (Table 1) that the log levels of these variables are nonstationary. The former are based on Im et al. (2003) and the latter is based on Pesaran (2007). The IPS statistic clearly shows that these variables are stationary in first differences. Surprisingly, the CIPS statistic indicates that these variables are not difference stationary. In what follows we assume these variables to be difference stationary, and therefore estimate Eq. (1) using panel cointegration method.

5.2 Panel cointegration tests of VTH

In Table 2 we present estimated variants of Eq. (1). Model 1 serves as a baseline estimated by seemingly unrelated (SUR) regression with country fixed effects. SUR is used because there is substantial cross-section dependence between the residuals. The average correlation of residuals is 0.388 and some correlations are negative. The Bartlett and Breusch–Pagan statistics clearly show that there is extensive cross-section

Table 1 Panel unit roots tests for difference stationarity

Logarithms	IPS		CIPS	
	d = 0	d = 1	d = 0	d = 1
International tourism to Israel	0.61	−8.84	−1.33	−2.16
Outbound tourism from origins	2.22	−6.94	−1.80	−1.82
Real exchange rate	1.81	−12.3	−0.71	−1.86

2 augmentations. Critical values ($p = 0.05$): IPS $= -2.1$, CIPS $= -2.2$. Order of differencing denoted by d

Table 2 Tourist arrivals to Israel: panel estimation of Eq. (1)

	Model 1	Model 2	Model 3	Model 4
Constant	3.221	1.597	0.491	0.239
Visa restrictions	−0.571	−0.392	−0.10	−0.348
Outbound tourism (ln)	0.097	0.078	0.003	0.012
Security incidents (ln)	−0.336	−0.339	−0.012	−0.032
Real exchange rate (ln)	0.273	0.007	−0.133	0.236
Common factor	No	Global tourism	CCE	Restricted CCE
Average correlation	0.388	0.495	0.007	0.051
Breusch–Pagan	0.267	0.338	0.154	0.252
Bartlett	0	0	0	0
CD	71.08	90.80	0.65	9.25
$GADF_2$	−2.22	−2.38	−3.44	−2.33
BCS1	−9.09	−9.21	−25.27	−10.54
BSC2	−5.29	−7.42	−21.5	−6.76

Dependent variable—ln International tourist arrivals from 60 countries during 1994–2012. Estimated by EGLS with SUR. $GADF_2$ = group ADF with 2 augmentations

dependence between the residuals, and the CD statistic overwhelmingly rejects the null hypothesis of weak cross-section dependence. There is also evidence of spatial cross-section dependence with the highest correlations (> 0.85) recorded between contiguous countries such as the USA with Canada, Germany with Austria, the Czech Republic, Netherlands and Italy, Italy with Spain and the UK with the Netherlands.

Model 1 shows that the elasticity of tourism with respect to outbound tourism to all destinations is only 0.097. Since a neutral elasticity is 1, this suggests that Israel is an inferior destination as far as international tourism is concerned. It also means that Israel's share of global tourism varies inversely with global tourism. In addition, the elasticity of demand with respect to the real exchange rate is 0.27 suggesting that foreign tourism to Israel is price sensitive. Recall that outbound tourism and the real exchange rates are integrated variables, as is tourism to Israel.

Next we consider variables that are not integrated. Tourism is adversely sensitive to security incidents. During the Second Intifada the number of incidents doubled which induced a decrease in tourism of 26 % according to Model 1. Finally, Model 1 implies that a partial visa waiver raises tourism by 77 % and a full waiver raises it by 213 %.[1] Both measures of BCS indicate that Model 1 is panel cointegrated. BCS2 is particularly indicative since its assumptions closely match the specification in Model 1.

Model 2 is the same as Model 1 except it includes a common factor, global tourism, which is assumed to affect each of the 60 countries differently. Note that this common factor includes tourism between countries not included in the study. The specification

[1] As the dummy variable for visas is 0, 1 or 2, a partial waiver reduces this variable by 1 and a full waiver by 2. This is multiplied by the estimated coefficient and the antilog of the result is taken.

of the common factor is expected to reduce the degree of cross-section dependence in the residuals, and to capture effects induced by multilateral resistance in tourism among the countries included in the study as well as between these countries and other countries not included in the study. Surprisingly, however, cross-section dependence strengthens rather than weakens. In Model 2 the parameter estimates are similar to those in Model 1 except for the real exchange rate effect. The effect of visa waivers is smaller than in Model 1; a partial waiver increases tourism by 48 %, and a full waiver increases it by 118 %. The average ADF statistic is more negative in Model 2 than in Model 1. However, the common factor has used 60 degrees of freedom. BCS2 is less informative than in Model 1 because it assumes that there is no common factor.

Model 3 is for the CCE specification proposed by Bertoli and Fernandez-Huertas Moraga (2013) to allow for multilateral resistance. It is estimated with three common factors (tourism from origins to Israel, outbound tourism from all origins, average real exchange rate). Since there are 60 origins Model 3 uses an additional 180 degrees of freedom. These common factors wipe out the effects on tourism to Israel of visa arrangements, real exchange rates, security incidents and outbound tourism. The residuals of Model 3 continue to be cross-section dependent, but as indicated by the CD statistic this dependence is weak. The average ADF statistic is smaller (more negative) than in Models 1 and 2, and the model is clearly cointegrated according to BCS1 and BCS2. But for the presence of the visa effect the appropriate test statistic would have been BCS1.

Model 4 is the same as Model 3 but with the omission of outbound tourism and real exchange rates as common factors. It therefore uses 120 fewer degrees of freedom. In contrast to Model 3, the visa effect is similar to its estimate in Model 2, but cross-section dependence in the residuals is greater in Model 4 than in Model 3. The appropriate critical values for Model 4 should be slightly less severe than for Model 3 because it specifies one common factor instead of three.

5.3 Holyland effect

Roughly 6 % of tourists to Israel also visit Jordan as part of a Holyland package. Since for Holyland tourists Jordan and Israel are complements, Jordanian visa arrangements with the origin countries might indirectly affect tourism to Israel. For example, if Jordan eases visa restrictions on Australians, not only might this increase Australian tourism to Jordan, it might also increase Australian tourism to Israel. To test this possibility we added to Model 1 Jordanian visa requirements as well as Israeli visa requirements. However, we were unable to detect any evidence of the Holyland effect.

6 Conclusions

We use tourism data to Israel from 60 countries during 1994–2012 to estimate the effect of visa waivers on tourism. The treatment effect on tourism is estimated from panel data, which is identified by the method of differences-in-differences. Potential confounders such as the effect of Israel's security situation on tourism, the increase in tourism in the countries of origin, and the effects of the real exchange rate on tourism are taken into consideration. Because most of these variables are non-stationary, the model is estimated and tested using panel cointegration in which changes in visa requirements are treated as structural breaks at known dates.

The estimated treatment effect is much larger (by far) than previous estimates. Complete elimination of visa requirements may almost triple tourism to Israel. At the same time, the elasticity of demand for tourism with respect to foreign tourism from the origin countries is of the order of only 0.1, suggesting that Israel is an inferior destination (by far). Also, the demand for tourism to Israel is price sensitive; although the elasticity of demand for tourism with respect to the real exchange rate is small (0.25).

We also tested the multilateral resistance model of tourism by specifying global tourism as a common factor and by using the CCE estimator. The former reduces the estimated effect of visa waivers on tourism, which nevertheless remains large. However, the latter implies that the visa effect is just a statistical artifact induced by ignoring multilateral resistance. This may be true. However, the CCE specification also implies that the adverse effect on tourism to Israel induced by the security situation just happens to be a statistical artefact due to multilateral resistance. We think that the latter implication is unreasonable. For example, it is unlikely that the decline in tourism of 32 % following Operation 'Protective Edge' in the summer of 2014 just happened to be induced by multilateral resistance. Therefore, we are inclined to discount the results obtained using CCE and conclude that the visa effect is large and statistically significant.

7 Data appendix

We use panel data (1994–2012) for the econometric analysis. This is assembled from a variety of sources as follows:

International tourist arrivals (Y): Annual data on tourists entering Israel is published by the Israeli Central Bureau of Statistics (CBS 2012Felsenstein). http://www.cbs.gov.il/publications12/1503_tayarut_2011/pdf/e_print.pdf. Since 1994, this is available by country of origin, for selected countries. Our data relate to 60 countries that serve as origins over this period.

International outbound tourism from origin Countries (T): the source of this panel data is World Bank Data (http://data.worldbank.org/indicator/ST.INT.DPRT). This variable captures the extent of international travel at the origin.

Visa agreements (V): this is panel data constructed from the Israeli Ministry of Foreign Affairs Bilateral agreements data base: http://mfa.gov.il/MFA/AboutTheMinistry/LegalTreaties/Pages/Bilateral-Treaties.aspx. This source flags all bilateral visa agreements signed between Israel and other countries and the year in which the visa waivers went into effect. This data relates to 60 countries and is coded as 0=automatic visa waiver, 1=visa approved in local consulate, 2=visa approved in Jerusalem only.

Security situation (S): The source of this annual data is the National Insurance Institute. http://www.btl.gov.il/English%20Homepage/Publications/AnnualSurvey/Pages/default.aspx. The data counts number of security incidents, some of which might result in damage and not result in persons injured or killed. It thus measures magnitude rather than intensity.

Real exchange rate (r): The source is: http://data.worldbank.org/indicator/PA.NUS.PPP/countries?display=default. The ratio of this index to the Shekel-Dollar exchange rate gives the PPP value of one shekel in the tourists' own currency.

Global tourism (GT): The source is the World Bank as above for International Outbound Tourism.

Global tourism receipts (GTR): Global receipts from tourism (1994–2012 in constant \$US (bn) where 2005=100. The source is UNWTO(2012) *World Tourism Barometer, Tourism Highlights*, http://mkt.unwto.org/en/publication/unwto-tourism-highlights-2013-edition.

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