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Tourism Management (2014) 45, 226-233. Available

from: http://www.sciencedirect.com/science/article/pii/S0261517714000946

ABSTRACT

Notwithstanding the systematic inclusion of an exchange rate variable of some form in studies examining international tourism flows, only scant attention has been paid to testing for a possible exchange rate regime effect. Drawing from recent advances in exchange rate regime classifications, this paper begins to fill this gap by investigating the long-run impact of exchange rate regimes on international tourism flows. The study employs a system generalized methods of moments (SYS-GMM) estimation for tourist arrivals on a panel of 27 Organization for Economic Co-operation and Development (OECD) and non OECD countries for the period 1980-2011. The results identify multiple exchange rate regime effects and support the importance of maintaining a relatively stable exchange rate to attract international tourist arrivals.

Keywords: Exchange rate regimes; Tourism flows.

1. Introduction

The determinants of international tourism flows have been under intense theoretical and empirical scrutiny for several decades (for relevant surveys, see Crouch, 1994; Li, Song, & Witt, 2005; Lim, 1997; and Song & Li, 2008). Alongside other economic and social determinants, an exchange rate variable, in some form, is consistently used in modeling tourism demand (see, *inter alia*, Bond, Cohen, & Schachter, 1977; Dritsakis & Gialetaki, 2004; Quadri & Zheng, 2010; Webber, 2001; Yap, 2011) and typically found to have explanatory power in the determination of international tourism flows (e.g., Eilat & Einav, 2004; Patsouratis, Frangouli, & Anastasopoulos, 2005; Roselló-Villalonga, Aguiló-Pérez, & Riera, 2005).

Despite the above, the potential impact of alternative exchange rate regimes on the volume of international tourist arrivals is still severely under researched. For readers who may be unfamiliar with the concept of exchange rate regime, it is worth clarifying that it refers to the policy imposed on a currency by its issuing country in relation to other currencies and the foreign exchange market, hence affecting the prevailing exchange rate level. Various regime policy options are available, including: (i) a *fixed* exchange rate (an attempt to tie the currency to another); (ii) a *floating* exchange rate (where the market drives movements in the exchange rate); (iii) a *pegged float* (where a central bank keeps the rate from deviating too far from a set value or target band); and (iv) other more complex means of linking one currency to the value of another or to that of a basket of currencies (see International Monetary Fund, 2009).

The paucity of research on the role of exchange rate regimes on international tourism flows is conspicuous especially when one considers the research attention that the wider literature has already devoted to exchange rate regimes, in terms of their effects on trade (Adam & Cobham, 2007; Frankel & Rose, 2002; Rose, 2000; Rose & van Wincoop, 2001), price levels (Broda, 2006; Ghosh, Gulde, & Wolf, 2002), the transmission of terms of trade shocks (Broda, 2004; Edwards & Levy-Yeyati, 2005), growth (De Vita & Kyaw, 2011; Husain, Mody, & Rogoff, 2005) and foreign direct investment flows (Abbott, Cushman, & De Vita, 2012; Abbott & De Vita, 2011).

The gap highlighted above is important since although, to date, the number of the studies that have actually examined the impact of *exchange rate regimes* on tourism demand can be counted in one hand (Gil-Pareja, Llorca-Vivero, & Martínez-Serrano, 2007; Santana-Gallego, Ledesma-Rodríguez, & Pérez-Rodríguez, 2010; Thompson & Thompson, 2010), the evidence available suggests that this measure of the exchange rate too may play an important role in the determination of international tourism flows.

It is, of course, true that a Government would not choose an exchange rate regime policy solely on the basis of whether one exchange rate regime or the other would benefit the tourism industry the most. But it is equally true that knowledge of the extent to which the tourism sector (a sector of growing economic importance) responds to alternative exchange rate regimes provides extremely valuable information to policy makers on the relative merits of each regime option and, therefore, in guiding such policy decisions. Moreover, knowledge of the extent to which an exchange rate *regime* variable - as distinct from the typical exchange rate measure usually introduced in tourism demand equations - is a significant determinant of international tourism flows, is also useful to increase scholars' knowledge of the modelling of the international tourism demand function. This reasoning further strengthens the case for the study of the role of exchange rate regimes in the context of tourism flows.

This study extends the investigative route paved by the limited work that has gone before in several respects. First, unlike Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007), Ledesma-Rodríguez, Pérez-Rodríguez, and Santana-Gallego (2012) and Thompson and Thompson (2010), who only estimate the effect of Economic and Monetary Union (EMU) on tourism, the present study aims at identifying the impact of a wide menu of exchange rate regime policy options, with an analytical focus on binary exchange rate regimes by country pairs.

Second, unlike Santana-Gallego, Ledesma-Rodríguez, and Pérez-Rodríguez (2010), our contribution is distinguished by the comparative use of two *de facto* regime classifications, the ones developed by Reinhart and Rogoff (2004) and Levy-Yeyati and Sturzenegger (2005), in addition to the *de jure* classification published by the International Monetary Fund (IMF, 2012). Furthermore, for the regression using the latter exchange rate regime classification, our sample makes use of the most up-to-date data available, covering the period 1980-2011.

Third, the study benefits from a fairly comprehensive model specification which augments a gravity-type equation (our underlying theoretical framework of reference for our baseline model specification) with key variables typically found to have explanatory power in the determination of international tourism flows. Since the pioneering work of Tinbergen (1962), the gravity model has given rise to a myriad of studies to explain trade patterns and policy issues. In recent years its application has been profitably extended to explaining issues such as migration flows (e.g., Helliwell, 1997), bilateral equity flows (e.g., Portes & Ray, 1998), foreign direct investment flows (e.g., Abbott & De Vita, 2011) and tourism flows (e.g., Gil-Pareja et al., 2007). Though originally criticised for lacking a solid foundation, its theoretical 'gravitas' is now well established (see, for example, Anderson, 1979, and Anderson & Van

Wincoop, 2003), leading to another spur to its usage. In essence, the gravity model postulates a positive bilateral relationship to country mass, and an inverse relationship to distance. Moreover, augmenting this basic relationship can lead to further insights as this augmentation allows for both supply and demand forces to be represented, hence avoiding the omission of critical variables underlying the data generation process.

Finally, a fundamental advance of the methodology employed in the present study is that rather than assuming weak exogeneity of the regressors, the approach hereby employed explicitly controls for simultaneity bias. Controlling for potential endogeneity is particularly important within a gravity type model since variables such as tourist arrivals, income, bilateral trade and relative prices may be simultaneously determined and in some cases causation is likely to run both ways. Instrumental variable estimation of a dynamic panel model within a system generalized methods of moments (SYS-GMM) framework not only exploits the time series variation in the data while accounting for unobserved country specific effects, it also controls for a possible correlation between the regressors and the error term, measurement error and endogeneity bias.

2. Exchange rate regimes and tourism demand

That the exchange rate ought to affect the demand for tourism demand is, of course, fairly obvious, and its inclusion in the tourism demand equation, therefore, has never been in dispute. A devaluation of a country's currency makes inbound international tourism less expensive, and, consequently, increased tourist flows to that country should result. Conversely, an increase in the value of a country's currency will make international tourism more expensive and cause a reduction in inbound tourist flows.

Despite some ambiguities and inconsistencies on the specific measure of the exchange rate used in tourism demand studies (for a detailed discussion of this issue, see De Vita & Kyaw, 2013), by and large, the exchange rate has been found to be an important determinant of international tourism flows. The exchange rate is typically included in tourism demand equations either combined with relative prices as an effective price (real exchange rate) variable, or as a separate variable in levels. In general, the rationale for inclusion stems from either its expected influence on price or the proposition that ‘in practice’ tourists use the exchange rate as a proxy for destination prices (De Vita & Kyaw, 2013). However, several studies have also modeled international tourism flows by taking into consideration the change in the exchange rate as well as various measures of exchange rate volatility to proxy exchange value risk or uncertainty (see, *inter alia*, Arbel & Ravid, 1985; Bond, Cohen, & Schachter, 1977; Martin & Witt, 1987, 1988; Quayson & Var, 1982; Song & Witt, 2006; Webber, 2001; Witt & Witt, 1995).

Nevertheless, as noted earlier, the number of studies that have actually examined the impact of *exchange rate regimes* (rather than just a measure of the exchange rate) on international tourism demand can be counted in one hand and, by focusing almost exclusively on the effect of the euro, are fairly restricted in scope.

The groundbreaking study in this limited field is that by Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007) who, using a gravity model on a panel dataset of 20 OECD countries over the period 1995–2002, estimate the effect of the euro on intra-EMU tourist flows. They find that the euro has increased tourism, with an effect of just over 6%. Specifically, their results show that the euro’s effect on tourism amounts to 6.3% when the exchange rate volatility variable is not included in the regression, and 6.2% when it is added. By controlling separately for the impact of exchange rate volatility in the demand equation, they are also able to ascertain that the positive effect of EMU on tourism is attributable to factors other than the sole elimination of the volatility driven uncertainty. They go on to suggest additional factors underlying the significance of EMU in increasing tourist flows, such as the elimination of transaction costs stemming from currency exchange, increased market transparency and, more generally, the expansion of business tourism as a result of the positive impact of EMU on trade.

Santana-Gallego, Ledesma-Rodríguez, and Pérez-Rodríguez (2010) extend the analysis pioneered by Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007) by examining the influence of a set of exchange rate regimes, not only of the European common currency, on tourism. With this aim in mind, they use a series of bilateral dummy variables reflecting the exchange rate regime arrangement between pairs of countries (a currency union or currency board, a currency peg, a managed float, and a flexible exchange rate). These binary exchange rate regime dummy variables are constructed using the *de facto* exchange rate classification developed by Reinhart and Rogoff (2004). Their findings, based on a panel of OECD countries for the period 1995-2001, suggest that more fixity in the exchange rate arrangement generates a positive effect on tourism. Other intermediate exchange rate regimes (a currency peg or a managed float arrangement) also promote tourism but not as much as currency unions. In short, the less flexible the exchange rate regime is, the greater the positive impact on tourism. This is a plausible result although, just like the study by Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007), the euro effect on tourism is evaluated in the early stages of EMU.

In their research note, Thompson and Thompson (2010) focus on the effects of the real exchange rate and the switch to the euro on tourism revenues in Greece from 1974 to 2006. They employ a basic error correction framework since their static model does not show evidence of cointegration for the variables in levels. They find that the introduction of the euro had a large positive impact on Greek tourism revenues, which they quantify in the region of 18%.

Finally, Ledesma-Rodríguez, Pérez-Rodríguez, and Santana-Gallego (2012) revisit the question of the impact of the euro on international tourism in the euro zone using annual data on a sample of OECD countries for the period 1995-2008. Their results suggest that post-circulation of the single currency (2002-2008) the impact is much larger than in the period of irrevocable exchange rates (1999-2001), with estimated effects ranging between 21% and 43%.

Although insightful in their own right, all the contributions reviewed above (with the sole exception of Santana-Gallego, Ledesma-Rodríguez, & Pérez-Rodríguez, 2010), focus exclusively on the effect of the euro on international tourism.

3. Methodology

3.1 Model and data

The sample is based on an unbalanced panel of 27 OECD and non-OECD high income countries over the period 1980 to 2011, yielding over nine thousand country-year observations, across 345 country-pairs. The countries are Australia, Austria, Belgium-Luxembourg, Canada, Denmark, Finland, France, Germany, Greece, Hong Kong, Iceland, Ireland, Israel, Italy, Japan, South Korea, Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, United Arab Emirates, United Kingdom, and the United States. The sample period covers all the most up-to-date data available for estimations pertaining to the IMF exchange rate regime classification whilst due to limitations of exchange rate regime data under the Reinhart and Rogoff (2004) and Levy-Yeyati and Sturzenegger (2005) classifications, the estimations derived using the latter schemes cover the period 1980-2004.

Drawing from standard variables entering the gravity equation and key determinants of the demand for tourism, the estimating regression takes the general long-run form:

$$y_{ijt} = \delta_0 + \delta_1 \text{ trade}_{ijt} + \delta_2 \text{ gdp}_{it} + \delta_3 \text{ gdp}_{jt} + \delta_4 \text{ pop}_{it} + \delta_5 \text{ pop}_{jt} + \delta_6 \text{ ep}_{ijt} + \delta_7 \text{ dis}_{ij} + \alpha_8 \text{ LANG}_{ij} + \alpha_9 \text{ FTA}_{ij} + \alpha_{10} \text{ COL}_{ij} + \alpha_{11} \text{ COMLAN}_{ij} + \alpha_{12} \text{ rxrvol}_{ijt} + \alpha_{13} \eta' Z_{ijt} + \varepsilon_{ijt} \quad (1)$$

where lower case letters denote variables expressed in natural logarithms, *i* and *j* indicate destination and origin countries respectively, *t* is time, and the variables introduced are defined as:

y_{ijt} = total number of tourist arrivals to country *i* from country *j* at time *t* (source: World Tourism Organization), with such international tourists defined as persons visiting other countries for any reason other than making an income;

trade_{ijt} = total real bilateral trade for each country-pair, computed as the sum of exports and imports (source: IMF Direction of Trade Statistics);

gdp_{it} (gdp_{jt}) = denotes real per capita Gross Domestic Product (GDP) for country *i* (country *j*) (source: United Nations Common Database);

pop_{it} (pop_{jt}) = the population of country *i* (country *j*) at time *t* (source: *Centre d'Etudes Prospectives et d'Informations Internationales*, CEPII);

ep_{ijt} = real (effective) relative prices calculated as the natural logarithm of $\{[(\text{CPI}_{it} / \text{CPI}_{jt})] * [(1) / (\text{Exchange Rate}_{ijt})]\}$, where CPI is the consumer price index (source: IMF International Financial Statistics and OECD Main Economic Indicators);

dis_{ij} = geographic distance based on bilateral distances between the biggest cities of the two countries, with those inter-city distances being weighted by the share of the city in the overall country's population (source: CEPII);

$LANG_{ij}$ = dummy variable that takes the value of 1 when the two countries share a common official language (source: CEPII);

FTA_{ij} = equals one when both countries have a free trade agreement, to capture goods market integration (World Trade Organization);

COL_{ij} = dummy variable that switches on to indicate the former or current existence of a colonial relationship (source: CEPII);

$COMLAN_{ij}$ = dummy variable that equals 1 when the two countries share a land border (source: CEPII);

$rxrvol_{ijt}$ = measure of real exchange rate volatility, calculated as the annual standard deviation of the monthly percentage changes in the real bilateral exchange rate (source: IMF Financial Statistics and OECD Main Economic Indicators);

Z_{ijt} = a vector of binary (dummy) variables related to exchange rate regimes. More specifically, we consider the case of: (i) a common currency, $CU\text{-}CU_{ijt}$, when both countries, *i* and *j*, are members of the same currency union at time *t*; (ii) $FIX\text{-}FIX_{ijt}$ when two countries fix their exchange rate to each other; (iii) $CU\text{-}FLT_{ijt}$ when one country is in a currency union and the other floats its currency; (iv) $CU\text{-}FIX_{ijt}$ when a country is in a currency union and the other fixes its currency; and, finally, (v) $FIX\text{-}FLT_{ijt}$ when one country fixes its currency and the other lets it float freely; and (vi) the case of a double float ($FLT\text{-}FLT_{ijt}$), when both countries float their currency. This specification allows for the comparison of the level of tourist arrivals under each country-pair exchange rate regime combination.

Hence, although here interest centers on establishing whether exchange rate regimes have any impact upon international tourism flows, guided by previous literature we augment a gravity-type equation by including the most important variables that can be expected to exert a systematic influence on tourism demand.

The causal relationship between trade and international tourism flows is well documented in the literature and many empirical studies regularly report highly significant estimated coefficients for the trade variable (e.g., Goh, 2012). As per the model specification employed by Ledesma-Rodríguez et al. (2012), since the dependent variable refers to the unidirectional tourist flows from country *j* to country *i*, GDP per capita and population are introduced separately for the origin and destination country. Whilst the likely effect of these country of origin variables related to economic size is intuitively obvious, the potential role of GDP per capita and population of the destination country (the gravity mass variables) on tourist arrivals may be rationalized in terms of the argument that the richer and greater a country is, the greater its tourism infrastructure and capacity to supply better and more diversified tourist services, and thus attract more tourists.

The real (effective) relative price of goods and services is also expected to be an important decision criterion when selecting a tourism destination. Following the approach employed in most previous studies (e.g., Artus, 1970; Uysal & Crompton, 1984; Witt, 1980) our relative price variable (calculated adopting the specification employed by Martin & Witt, 1987), accounts for the cost of tourism in the destination market relative to the cost of tourism in the origin country.

Another component of tourism costs is the price of transportation. Yet, due to the complexities of the price structure of transportation, no completely satisfactory index exists for foreign transportation prices (Eilat & Einav, 2004; Stronge & Redman, 1982). Instead, researchers often include geographic distance as a proxy for these costs and for the forgone time spent and inconvenience of transportation (e.g., Marroco & Paci, 2013). The emergence of gravity models has made geographic distance even more popular as a regressor in equations aimed at explaining international tourism flows (see, for example, Eryiğit, Kotil, & Eryiğit, 2010). The greater is the proximity between two countries, the lower the transportation costs, but as distance increases, the costs of transportation also increase, with a subsequent decrease in tourist flows.

All the other standard gravity-type variables such as a former or existing colonial relationship, a common land border, joint membership of the same regional trade agreement (to capture goods market integration), are expected to promote the international flows of tourists. This also applies to a common language, which in the sample of this study relates to the English language for the USA, the UK, Canada, Australia, New Zealand and Singapore (though the Government of the latter recognizes four official languages: English, Malay, Chinese and Tamil), to French for Belgium-Luxemburg, Canada (Belgium-Luxemburg and Canada are officially bilingual), France and Switzerland, and to German for Austria and Germany.

Supported by the empirical literature reviewed earlier, to isolate the impact of exchange rate regimes, this study additionally controls for the independent effect of the volatility of the exchange rate. To avoid arbitrariness in the selection of the optimal volatility measure, following the non-nested testing tournament developed by the De Vita and Abbott (2004), alternative exchange rate volatility specifications (e.g., variance measure of volatility versus the Generalized Autoregressive Conditional Heteroscedasticity specification) were subjected to empirical scrutiny. The results of this pre-testing exercise revealed that the standard deviation provided the best fit to the data. The latter is specified as the standard deviation of the monthly percentage change in the bilateral exchange rate.

Unlike Santana-Gallego et al. (2010), who used exclusively the exchange rate regime classification developed by Reinhart and Rogoff (2004), three different exchange rate regime classifications to construct the binary regime dummies are employed in this study. The first classification is that published by the IMF in its *Annual Report on Exchange Arrangements and Exchange Restrictions* (IMF, 2012). Prior to 1999, the IMF's classification reflected exclusively countries' announced exchange rate regimes, thus failing to capture the extent to which actual policies conformed to countries' declared commitment (i.e., their official announcements). Since 1999 the IMF moved from a *de jure* classification to a hybrid one that combines data on the actual behavior of the exchange rate.

In response to Genberg and Swoboda's (2005) call for empirical investigations to make use of both types of classification, two of such alternative *de facto* schemes are adopted in this study, both for comparative purposes and as a robustness check. The one developed by Reinhart and Rogoff (2004) incorporates data on market determined exchange rates, is based on a 5-year horizon, and strips off the floating category of observations characterized by high inflation. The classification developed by Levy-Yeyati and Sturzenegger (2005) uses cluster analysis techniques to group countries' regimes on the basis of the volatility of the exchange rate relative to the relevant anchor currency, the volatility of exchange rate changes, and the volatility of reserves.

Table 1
Classifications of exchange rate regimes.

	Reinhart & Rogoff's classification	IMF's classification	Levy-Yeyati & Sturzenegger's classification	Classification for exchange rate regime dummies
1	No separate legal tender	Currency union	Currency union	Currency union
2	Pre announced peg or currency board arrangement.	Currency board/ Currency peg within horizontal band of ±1%.	Intermediate (dirty / crawling peg)	Fixed exchange rate
3	Pre announced horizontal band that is narrower than or equal to ±2%.			
4	De facto peg.			
5	Pre announced crawling peg.	Currency peg within crawling band of ±1%.		
6	Pre announced crawling band narrower than or equal to ±2%.			
7	De facto crawling peg.			
8	De facto crawling band narrower than or equal to ±2%.			
9	Pre announced crawling band wider than or equal to ±2%.	Currency peg within crawling band of at least±1%.		
10	De facto crawling band narrower than or equal to ±5%.			
11	De facto moving band narrower than or equal to ±2%			
12	Managed floating	Managed floating/ independently floating	Dirty float / managed float	Currency float
13	Freely floating			
14	Freely falling	N/A	N/A	N/A
15	Dual market in which parallel market data is missing			

Table 1 shows the exact correspondence between the original categories of the three classification schemes described above and those derived from them to inform the menu of exchange rate regimes applied in this study. It is worth pointing out that the extent of the 'compression' of the fine codes of the original regime classifications' categories within the categories used in the analysis that follows, was chosen to ensure empirical tractability and to derive a sufficiently informative menu of feasible exchange rate regime combinations to be examined.

In this study caution is exercised in declaring *a priori* expectations for the various exchange rate regime dummies since, to date, no full-blown theoretical framework has modeled the impact of exchange rate regimes on the behavior of international tourism flows. Nevertheless, intuitively, one could expect that the demand for tourism responds positively to a common currency (the CU-CU regime combination), due to both the absence of currency transaction costs and the elimination of the uncertainty associated with exchange rate volatility. Following the same logic, it could be hypothesized that the exchange rate regime combinations that entail a lower degree of uncertainty (e.g., FIX-FIX) would be more conducive to stimulating the flow of tourist arrivals than regime combinations that imply greater exchange rate uncertainty (e.g., FLT-FLT), though the arduous task of quantifying the multiple potential effects of exchange rate regimes on international tourism flows remains an empirical matter.

3.2 Econometric Approach

To appreciate the virtues of the SYS-GMM approach, it is worth to start by highlighting that in both fixed and random effects settings, estimation of autoregressive models (models in which the set of right-hand variables includes the dependent variable) can give rise to substantial problems. These problems stem from the fact that when included as a regressor, the lagged dependent variable will be correlated with the disturbance term, even if the latter is not itself autocorrelated.

The standard linear first-differenced GMM estimator addresses this issue but not without additional complications. To illustrate these complications, let us consider an autoregressive model of order one [AR(1)] with unobserved individual-specific effects:

$$y_{ijt} = \beta\, y_{ij,t-1} + \chi_{ij} + \nu_{ijt} \text{ for } i = 1, \ldots, N,\ \ j = 1, \ldots\ldots N-1 \text{ and } t = 2, \ldots, T \tag{2}$$

where $\chi_{ij} + \nu_{ijt} = \mu_{ijt}$ has the standard error component structure:

$$E[\chi_i] = 0;\quad E[\nu_{it}] = 0;\quad E[\chi_i\, \nu_{it}] = 0 \qquad \text{for } i = 1, \ldots, N; \text{ and } t = 2, \ldots, T \tag{3}$$

Assuming serially uncorrelated transient errors and that the initial conditions y_{ij1} are predetermined, the following m = ½ (T – 1) (T – 2) moment restrictions are obtained:

$$E\left(Z'_{ij}\Delta\nu_{ij}\right) = 0 \tag{4}$$

where Z_{ij} is the (T – 2) x m matrix given by

$$Z_i = \begin{bmatrix} y_{ij1} & 0 & 0 & \cdots & 0 & \cdots & 0 \\ 0 & y_{ij1} & y_{ij2} & \cdots & 0 & \cdots & 0 \\ . & . & . & \cdots & . & \cdots & . \\ 0 & 0 & 0 & \cdots & y_{ij1} & \cdots & y_{ij,T-2} \end{bmatrix} \tag{5}$$

and $\Delta\nu_{ij}$ is the (T-2) vector $\left(\Delta\nu_{ij3}, \Delta\nu_{ij4}, \ldots, \Delta\nu_{ijT}\right)'$. These are moment restrictions exploited by the standard linear first-differenced GMM estimator which entails the use of lagged levels as instruments for the equations in first-differences (Arellano & Bond, 1991). This yields a consistent estimator of β when N approaches infinity and T is fixed. However, in the presence of high persistence of the series or a large variance of the individual specific effect (relative to the variance of the remainder of the error term), lagged levels make weak instruments for the regression in differences. Instrument weakness, in turn, increases the variance of the coefficients and, in small samples, is likely to generate biased estimates.

To address the complications associated with the standard GMM estimator, the SYS-GMM model of Arellano and Bover (1995) and Blundell and Bond (1998) imposes the additional assumption:

$$E\left(\mu_{ij}\Delta y_{ij2}\right) = 0 \quad \text{for i = 1, …, N and j=1,……N-1} \qquad (6)$$

This assumption yields (T-2) further linear moment conditions:

$$E\left(\mu_{ijt}\Delta y_{ij,t-1}\right) = 0 \qquad \text{for i = 1, …, N, j=1,……N-1 and t = 3, 4, …, T} \qquad (7)$$

Such additional moment conditions allow the construction of a GMM estimator that uses the entire set of moment restrictions in a stacked system of lagged first-differences as instruments for T-2 equations in levels, in addition to lagged levels as instruments for T-2 equations in first-differences. The full set of second-order conditions can be expressed as:

$$E\left(Z_{ij}^{+'}\,\mu_{ij}^{+}\right) = 0 \qquad \text{where } \mu_{ij}^{+} = \left(\Delta\nu_{ij3},...,\Delta\nu_{ijT},...,\mu_{ij3},...,\mu_{ijT}\right)' \qquad (8)$$

The econometric procedure illustrated above suits the purpose of the analysis well. Tourist arrivals, factors influencing regime choice and variables typically entering the tourism demand equation and gravity models (e.g., bilateral trade, income of origin and destination country, and relative prices) may be simultaneously determined. With SYS-GMM every regressor is instrumented and including both level and first difference equations in a stacked system addresses potential issues of endogenity bias. Moreover, such an approach permits the investigation of whether time-invariant variables, such as distance (hereby included as a proxy for the costs of transportation), play an important role in the determination of tourist arrivals, something not possible from the standard GMM estimator. The specific linear dynamic model hereby used for estimation, therefore, can be defined as:

$$y_{ijt} = \alpha_0 + \sum_{k=1}^{p}\alpha_k\, y_{ijt-k} + \sum_{l=0}^{q}\beta_l\, x_{ijt-l} + \eta_i + \lambda_t + \nu_{ijt}$$

$$i=1,\ldots,n \qquad t=1,\ldots..T \tag{9}$$

where y_{ijt} is the total number of tourist arrivals at each destination i, from each origin country j, $y_{ij1},\ldots\ldots y_{ijp}$ represent the autoregressive structure to reflect habit/persistence in the tourist's choice of destination, and the rigidity of supply of tourism services, and $x_{ij0},\ldots x_{ijq}$ are the current and lagged values of the matrix of regressors that could be strictly exogenous, predetermined or endogenous with respect to ν_{ijt}, the error term. η_i are individual effects that estimate differences in the mean level of tourist arrivals across country-pairs. λ_t are time specific effects to capture the effect of common disturbances or spatial correlation across the units of the panel.

Transformation of equation (9) leads to a set of T_i equations being estimated across the country pairs. Provided that T>3, then for every period and with a lag length of q, (T_i-q) first-difference equations are estimated using (T_i-q) lagged level instruments, with (T_i-q) level equations estimated using (T_i-q+1) first-difference instruments. As T grows in size, the computational requirements of SYS-GMM rise significantly, resulting in a trade-off between gains in efficiency and greater bias in the estimates, due to over-fitting of the estimated equation with too many instruments. Accordingly, in the analysis that follows the instrument matrix is restricted so that only one lag is used for the first difference equation, while for the level equation, the contemporaneous first difference and its lag are used.

Consistency of the SYS-GMM estimator requires evidence of significant first-order serial correlation (by construction, from the first-difference equation, $\Delta\nu_{ijt} = \nu_{ijt} - \nu_{ijt-1}$ should correlate with $\Delta\nu_{ijt-1} = \nu_{ijt-1} - \nu_{ijt-2}$ due to the common element ν_{ijt-1}) but no higher order correlation. For this purpose, following Abbott and De Vita (2011), the (ARp) Arellano-Bond statistic (Arellano & Bond, 1991) is employed.

Moreover, to verify the validity of the chosen instruments, the Hansen's (1982) J-test is adopted.

4. Empirical results

As illustrated in the methodology section above, a short-run dynamic model was estimated using the two-step SYS-GMM econometric procedure in order to generate the long-run coefficients of interest across three alternative exchange rate regime classification specifications.[1] Although the short-run effects are beyond the scope of this investigation (for which interest centers upon the long-run impact of exchange rate regimes), this methodology requires checks on the dynamics of the underlying short-run model so as to increase the confidence on the thus derived long-run coefficients. [2] It is for this purpose that Table 2 reports the diagnostic test results of the underlying short-run specification.

Table 2
Diagnostic results from the underlying short-run model.

Diagnostics	IMF (1980-2011)	Reinhart & Rogoff (1980-2004)	Levy-Yeyati & Sturzenegger (1980-2004)

[1] Since SYS-GMM requires the stationarity of the process generating the initial conditions y_{ijt}, in the pre-testing phase, the time series (integration) properties of the dependent variable were tested for. Using the LM unit root test with up to two structural breaks (for a similar testing procedure see, for example, Abbott, De Vita, & Altinay, 2012), it was found that the tourist arrivals series is indeed I(0), i.e. stationary in levels, at the 5% significance level, with the significant break points located in 2001 and 2003, dates which coincide with the terrorist attacks on the US and the outbreak of the SARs epidemic, respectively.

[2] As a further preliminary check and for the sake of comprehensiveness, the existence of a long-run equilibrium relationship among the variables was also tested for using Pedroni's (2004) panel cointegration procedure which makes use of panel (parametric and non-parametric) v, ρ and t statistics. This method utilizes the residuals from the cointegrating regression given by $y_{i,t} = \alpha_i + \delta_i t + x'_{i,t}\beta_i + e_{i,t}$ for t=1,2,3…,T; i=1,2,3,…,N; where $\beta_i = (\beta_{1i}, \beta_{2i}, \ldots, \beta_{Mi})'$, and $x_{i,t} = (x_{1i,t}, x_{2i,t}, \ldots, x_{Mi,t})'$. This specification allows for considerable heterogeneity in the panel, since heterogeneous slope coefficients, fixed effects and individual specific deterministic trends are all permitted. Furthermore, no exogeneity conditions are imposed on the regressors in the cointegrating equation. Using the critical values of Pedroni (2004), the results indicated rejection of the null of 'no cointegration' with estimated coefficients of 11.23 for the panel v-statistic, -13.56 for the panel ρ-statistic, and -15.12 for the non-parametric panel t-statistic.

Wald test: all regressors~χ^2(df)	3127.66[a] (55)	2894.13[a] (40)	3438.35[a] (44)
Wald test: time dummies~χ^2(df)	71.57[a] (23)	72.83[a] (23)	70.55[a] (23)
Wald test: ER dummies~χ^2(df)	28.88[a] (6)	29.72[a] (6)	31.09[a] (6)
J-test~χ^2(df)	343.12 (482)	341.35 (344)	335.75 (341)
AR(1) test~N(0,1)	-3.71[a]	-3.68[a]	-3.87[a]
AR(2) test~N(0,1)	-0.28	-0.31	-0.33
Number of observations	9,380	7,390	7,390
R^2	0.73	0.65	0.69

Note: The Wald tests are for exclusion restrictions, with the number of restrictions reported in parentheses. J-test denotes the Hansen J-test for instrument validity. AR(1) and AR(2) are tests for first and second order serial correlation. The superscript "a" denotes significance at the 5% level.

As can be seen from Table 2, the model works well on many dimensions across all three classifications. The Hansen's J-test for instrument validity, and both the exclusion and serial correlation tests, do not reject the chosen econometric specification. Overall, the model displays high 'goodness of fit', explaining over two thirds of the variation in tourist arrivals, with an R-squared value ranging from 0.65 to 0.73 across the three regimes classification specifications.

Table 3

SYS-GMM long-run estimates across regime classifications.

Variable	IMF (1980-2011)	Reinhart & Rogoff (1980-2004)	Levy-Yeyati & Sturzenegger (1980-2004)
$trade_{ijt}$	0.31[a]	0.09[a]	0.26[a]
	(23.34)	(4.02)	(16.14)
gdp_{it}	0.20[a]	0.17[a]	0.18[a]
	(3.53)	(3.40)	(3.63)
gdp_{jt}	0.78[a]	0.53[a]	0.70[a]
	(19.58)	(13.37)	(15.95)
pop_{it}	0.25[a]	0.42[a]	0.27[a]
	(16.76)	(18.61)	(20.73)
pop_{jt}	0.75[a]	0.64[a]	0.72[a]
	(21.18)	(16.33)	(19.77)
ep_{ijt}	-0.48[a]	-0.39[a]	-0.40[a]
	(-4.81)	(-3.88)	(-4.33)
dis_{ij}	-0.78[a]	-0.76[a]	-0.81[a]
	(-34.11)	(-25.54)	(-30.43)
$LANG_{ij}$	0.30[a]	0.25[a]	0.28[a]
	(6.64)	(6.29)	(6.58)
FTA_{ijt}	0.68[a]	0.60[a]	0.63[a]
	(6.03)	(5.96)	(7.21)
COL_{ij}	0.71[a]	0.57[a]	0.64[a]

	(3.55)	(2.71)	(2.97)
$COMLAN_{ij}$	0.11	0.09 [a]	0.10
	(1.26)	(1.99)	(1.29)
$rxrvol_{ijt}$	-0.51 [a]	-0.37 [a]	-0.40 [a]
	(-2.33)	(-1.83)	(-1.99)
$CU\text{-}CU_{ijt}$	0.23 [a]	0.18[a]	0.21 [a]
	(5.37)	(3.14)	(6.93)
$FIX\text{-}FIX_{ijt}$	0.09 [a]	0.14[a]	0.09 [a]
	(3.21)	(3.25)	(2.16)
$CU\text{-}FLT_{ijt}$	0.13 [a]	0.13[a]	0.16 [a]
	(4.62)	(3.57)	(4.92)
$CU\text{-}FIX_{ijt}$	0.06 [a]	0.04 [a]	0.06 [a]
	(2.32)	(2.01)	(2.63)
$FIX\text{-}FLT_{ijt}$	0.07	0.02	0.05
	(0.58)	(0.49)	(1.18)
$FLT\text{-}FLT_{ijt}$	0.18	-0.01	0.08
	(1.55)	(-0.59)	(1.23)

Note: The above are long-run estimates derived from an underlying short-run dynamic model estimated using a two step SYS-GMM procedure (a constant term and time dummies were also included but not reported to conserve space). Robust t-ratios are reported in parentheses. The superscript "a" denotes significance at the 5% level.

The SYS-GMM long-run estimates are reported in Table 3. In comparing the performance of different exchange rate regime classification specifications, the first point to note is that although many observations are lost due to lack of data for some variables for some countries and years, the proportion of usable data remains fairly stable across exchange rate regimes. Despite few, rather marginal discrepancies, mostly pertaining to the Reinhart and Rogoff's classification, the estimates display a considerable degree of consistency, evidence which is, in itself, indicative of the reliability of the results in spite of the longer time span covered by the estimations using the IMF regime classification (1980-2011) and the inevitable heterogeneity of the sample countries, for example, in terms of income and market size. The slight misalignment of the estimates derived from the Reinhart and Rogoff's classification, particularly with regard to the CU-FIX, FIX-FLT and FLT-FLT estimated coefficients, can be explained by the differences in the distribution of the categorization of exchange rate regimes. A country-specific example of divergence relates to the observations for Greece in the years 1999-2000, which are categorized as 'fixes' by the IMF and the Levy-Yeyati and Sturzenegger's schemes, and as a 'common currency' by the Reinhart and Rogoff's classification.

Most of the estimated coefficients of the control variables are statistically significant at the customary level (5%). They all have the expected sign with sensible magnitudes, broadly in line with those found in previous studies. Consistent with the view that controlling for bilateral trade is crucial for a well-specified gravity model (e.g., Anderson, 1979), the estimated coefficient of the trade variable ($trade_{ijt}$) is found to exert a positive and significant effect, suggesting that total real bilateral trade promotes inbound tourist flows. In accordance with the rationale underlying the formulation of a gravity model, the 'economic mass' variables are also found to be statistically significant, and exert a positive influence on international tourism. Yet, like Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007), it is found that the population and real GDP per capita variables of the origin countries (pop_{jt} and gdp_{jt} respectively) have a larger impact on tourist arrivals than the population and real GDP per capita variables of the destination countries (pop_{it} and gdp_{it} respectively). That said, the statistical significance of the latter (gravity mass) variables, provides further empirical evidence of their relevance in estimation within a gravity framework, as their omission would have evidently resulted in model misspecification. This result also provides empirical support to the argument that the richer and greater a country is, the greater its tourism infrastructure and capacity to supply better and more diversified tourist services, thus attracting more tourists.

The estimated coefficient of the effective relative price variable (adjusted by the exchange rate) ranges from -0.39 under the Reinhart & Rogoff classification to -0.48 under the IMF classification, indicating that a 1% increase in the (real) price level of the destination country relative to the origin country, decreases the number of tourist arrivals by anything between 0.39% and 0.48%. Distance (dis_{ij}) too has the expected negative sign, showing that tourists prefer nearby destinations, possibly owing to the difference in the costs of transportation of near and far destinations in spite of the widespread availability of frequent intercontinental direct flights. A common language ($LANG_{ij}$), joint membership of a regional free trade agreement (FTA_{ij}) and the former or current existence of a colonial relationship (COL_{ij}) all have a positive effect on international tourism flows whilst sharing a land border ($COMLAN_{ij}$) is not statically significant in most specifications. The latter result would suggest that whilst geographic distance is a significant determinant of international tourism flows, the tourism destination decision is not necessarily influenced by the specific desire to cross the country's land border since the destination market's attractiveness (in terms of natural, historical, cultural and recreational elements) obviously remains a key factor in influencing tourists' destination preferences. Real exchange rate volatility ($rxrvol_{ijt}$) appears to discourage inbound tourism flows, with an estimated effect ranging from -0.37 to -0.51 across the three specifications.

Given the primary aim of the study, attention now centers upon the exchange rate regime dummies. Starting with the estimated coefficient of CU-CU, the effect of a common currency on tourist arrivals is found to be between 19.7% [($exp^{0.18}$ -1) *100] and 29.7% [($exp^{0.26}$ -1) *100] across the three regime classification specifications. This range of estimates is very plausible and so is the higher elasticity found for the regression based on the IMF classification which covers the full sample period and hence exploits seven years of additional observations (up to 2011) pertaining to the adoption of the euro. This range of estimates also sits comfortably across the spectrum of evidence available to date from previous studies. At one end of the spectrum, the estimated effect reported by Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007) is just over 6%, though their sample period was much shorter, ending at 2002, the year in which the euro started circulation. Moreover, the SYS-GMM estimation procedure hereby employed controls for endogeneity bias, and the model takes into account additional exchange rate regime combinations not considered by Gil-Pareja, Llorca-Vivero, and Martínez-Serrano (2007). At the other extreme, Santana-Gallego et al. (2010) reported a common currency effect on tourist arrivals ranging from 329% under their FE-2SLS estimation to 349% under OLS, estimates which do not appear to be plausible and that align poorly to the stylized facts on international tourism flows. However, it should be noted that Santana-Gallego, Ledesma- Rodríguez, and Pérez-Rodríguez (2010) did not control separately for the effect of exchange rate volatility. Moreover, their model specification omitted a relative price variable and a proxy for transportation costs, key variables in the tourism demand equation.

In interpreting the impact of a common currency, it is also of significance the fact that its effect on tourist arrivals appears not to stem solely from the elimination of exchange rate risk, since the model controlled separately for both the impact of exchange rate volatility ($rxrvol_{ijt}$) and for the role played by exchange rate movements in affecting relative prices (ep_{ijt}), both of which were found to be statistically significant across all specifications. Other factors, therefore, may be at work in driving the tourism creation effect of a common currency such as the elimination of exchange rate transaction costs and possibly greater market transparency (lower informational barriers).

Also the estimated coefficients of the remaining binary exchange rate regime dummies lend themselves to sensible interpretations. Membership of a currency union appears to promote tourist arrivals also from countries which fix and/or float their currency (CU-FIX$_{ijt}$ and CU-FLT$_{ijt}$ respectively), signaling an additional diversion effect of tourist arrivals from countries outside the euro zone to those adopting the single currency. The estimated coefficient of the FIX-FIX dummy is positive and significant, though its impact upon tourist arrivals is not as big as that of a common currency. Finally, the estimated coefficients of the FIX-FLT and FLT-FLT$_{ijt}$ dummies are found to be statistically insignificant, with no discernible effect on the number of tourist arrivals.

5. Conclusion

This study aimed at investigating the long-run impact of exchange rate regimes on international tourism flows, using data from a panel of 27 OECD and non-OECD countries for the period 1980-2011. Drawing from recent advances in exchange rate regime classifications, the study adopted three different *de jure* and *de facto* exchange rate regime classification schemes, and a SYS-GMM estimation technique that exploits the time series variation in the data, accounts for unobserved country-specific effects, and controls for a possible correlation between the regressors and the error term, measurement error and endogeneity bias.

Moving beyond the orthodox tourism demand framework, the study also benefited from a fairly comprehensive model specification which, in addition to making use of the most up-to-date data available, augmented a gravity-type equation with key demand and supply variables typically found to have explanatory power in the determination of international tourism flows.

The results indicate a significant effect of exchange rate regimes on inbound tourism flows. The effect of a common currency is found to exert the strongest positive impact on the volume of international tourist arrivals. Membership of a currency union also promotes tourist arrivals from countries fixing or floating their currency. The latter finding suggests that membership of a currency union generates a positive tourism diversion effect in addition to intra-union tourism creation. By way of contrast, other binary exchange rate combinations (namely, when one country fixes its currency and the other lets it float freely, and the case when both countries float their currency) appear to have no discernible effect on the number of tourist arrivals. By and large, these results are robust to whichever exchange rate regime classification is employed, irrespective of whether it is *de jure* or *de facto*.

These findings are of considerable significance since they identify multiple exchange rate regime effects and support the importance of maintaining a relatively stable exchange rate to attract tourist arrivals. In this respect it worth highlighting that whilst most other determinants of tourism demand are outside the control of policy makers (e.g., climate), the type of exchange rate regime directly results from a policy decision. At a time when many European countries have to deal with unprecedented recessionary pressures and growing skepticism over their membership of the currency union, this study highlights one of the most understated benefits of the single currency, the promotion of inbound tourism, which can play a key role for economic recovery and growth if aided by complementary policies in support of the sustainable development of the sector.

The findings also provide useful guidance for researchers willing to estimate the demand for international tourism using a panel data specification that can adequately account for the role that exchange rate regimes play on international tourism flows.

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