THE WEALTH EFFECT AND TOURISM – ARDL MODELING AND GRANGER CAUSALITY IN SELECTED EU COUNTRIES

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Abstract
Purpose – The purpose of this paper is to examine the relationship between various forms of income/wealth and tourism departures in selected European Union (EU) countries.
Design – The design of this study is based on analysing quarterly data on incomes, house prices, net financial assets, financial derivatives and employee stock options in order to measure the link between wealth and tourism departures in selected countries—Austria, Belgium, Germany, Denmark, Spain, France, Ireland, Italy, Sweden, Slovenia and the United Kingdom (UK)—with individual-country time series that span the period from 2000 to 2018.
Methodology – Granger causality in the relationship between income and wealth inputs and tourism departures has been examined using the autoregressive distributed lag model (ARDL), the bounds test for cointegration, the vector error correction model (VECM), and long- and short-term (as well as joint) causality.
Findings – The findings showed the existence of cointegration and a direct effect on the relationship between various sources of wealth/income and tourism departures in both the long- and short-run, and jointly, only in the case of Austria. House price, net financial asset and stock option causality results essentially showed unidirectional causality that runs from that form of wealth to tourism departures in the case of Belgium, Germany, Spain, France, Slovenia and the UK, thus providing extensive support for the wealth effect–tourism link in certain countries.
Originality of the research – The originality of this paper stems from the fact that it offers the first analysis of the relationship between various forms of income/wealth and tourism departures in selected European Union (EU) countries. The findings of this study suggest that a bounds test for cointegration due to data constraints should be taken into account when examining the link between the wealth effect and tourism in country time series analysis.
Keywords Wealth effect, tourism departures, ARDL model, Granger causality, EU countries

INTRODUCTION

The importance of various sources of wealth for tourism consumption is undeniable. The essence of the debate on the relationship between income and wealth as a basis for spending and tourism consumption is whether the former cause the latter.

Alongside income, wealth plays an important role in determining tourism consumption to a greater or lesser extent, without an inherited hazard, depending on an enigmatic term, referred to—in academic economic jargon—as the ‘wealth effect’. The macroeconomic dissaving cycle and the positive wealth effect, with stock and housing market valuations rising dramatically (primarily in good times, or in the pre-Great Recession period), have encouraged numerous households to go on vacations abroad, some for the very first time. The psychological feeling of well-being by the average well-to-doer, based on rising financial and housing asset prices, along with incomes/savings, eventually triggers a higher propensity to buy tourism vacations abroad.

The consumption theory developed primarily from the absolute income hypothesis (AIH; Keynes 1936), and progressed to the relative income hypothesis (RIH; Duesenberry 1949), the life-cycle hypothesis (LCH; Modigliani and Brumberg1954), the permanent income hypothesis (PIH; Friedman 1957), the random-walk hypothesis of consumption under uncertainty (Hall 1978), and the precautionary saving theory (Fafchamps and Pender 1997). The existence of wealth effects is rarely disputed in theory, and as critical works add to existing knowledge, more innovative hypotheses emerge, but gauging the characteristics of such effects is, empirically, challenging. Since tourism consumption is just one fraction of household consumption, it is prudent to assume that unexpected transitory gains (arising from higher financial or housing asset valuations) can translate into higher tourism demand (i.e., a greater desire for foreign vacations).

Some transitory gains come as ‘manna from heaven’, quite unexpectedly, over time, and in our presumption boost tourism departure outflows. The digital edition of *The European economy since the start of the millennium: A statistical portrait* (Eurostat 2018) recently published an interesting set of data, which for the purposes of this paper was subjected to scrupulous econometric analysis. House prices, including purchases of both newly built and existing houses and flats, have fluctuated significantly since 2006 (i.e., since before the Great Recession); the annual growth rate in the EU as a whole was an average 11 per cent between 2010 and 2017, and 6 per cent in the Eurozone. The highest increases during this period were recorded in Estonia (73 per cent), Sweden (56 per cent), Austria (49 per cent), Latvia (47 per cent) and Luxemburg (40 per cent), while the largest decreases were in Spain (-17 per cent), Italy (-15 per cent) and Cyprus (-9 per cent) (Eurostat 2018). Some of these countries’ household wealth declined due to its being linked to real estate values. Financial derivatives and employee stock options include a variety of financial transaction dynamics. The common characteristic of stock options, motivated by the transfer of risk from the company to its employees, rather than the supply of funds or other resources to the company, is that they bring in wealth to working households. Legislation and political attitudes in EU countries differ as regards the augmentation of wealth in this manner (the UK is the absolute champion amongst EU countries). Stock options can generate significant wealth, despite fluctuations in value, or alternatively cause a loss of wealth if the company goes bankrupt. Due to their complexity, the delivery of transitory gains or losses through stock options cannot be foreseen from figures (nor is this the objective of the present paper), except in the case of the link to tourism consumption, as we will see in this paper. Wage earnings (income) and savings dynamics in families that habitually travel abroad, as a direct and more to earth (liquid) source of travel consumption, have undergone unpredictable changes too. In real terms, the disposable income of households in the EU grew in total by 16 per cent between 2000 and 2009. Following the financial crisis, it decreased by around 3 per cent from 2009 to 2013, and then rose by 5 per cent between 2013 and 2016. In total, the disposable income of EU households increased by around 18 per cent between 2000 and

2016 — an average growth rate of around 1 per cent per year. The household saving rate in the EU has been rather stable since the beginning of the millennium, fluctuating between 11 per cent and 13 per cent. The pattern is about the same in the Eurozone, but at a slightly higher rate. In 2016, the highest household saving rates were observed in Luxembourg (20 per cent), Sweden (19 per cent), Germany (17 per cent) and France (14 per cent), and the lowest in Cyprus (-2 per cent), Lithuania (0 per cent), Latvia (3 per cent) and Poland (4 per cent) (Ibid.).

Therefore, we have given serious attention to the study of 'income and wealth effect' factors that affect demand for tourism. To that end, we have used a different approach to investigate the effect of wealth factors, as well as income, on tourism departure growth, namely an autoregressive distributed lag (ARDL) cointegration test developed by Pesaran and Shin (1999) and Pesaran et al. (2001), and Granger causality analysis (Granger 1969), in eleven EU countries — Austria, Belgium, Germany, Denmark, Spain, France, Ireland, Italy, Sweden, Slovenia and the UK, based on data for the period between 2000 and 2018. These countries were chosen because they are amongst the largest generators of tourism demand in the EU.

Due to the aforementioned general hypothesis regarding a wealth–tourism consumption link, and in order to fill the literature gap (limited studies have been undertaken in this regard), we have examined the effects of wealth in selected EU countries.

The originality of this research is; (1) we consider a time series analysis of data in eleven EU countries, from 2000 to 2018. So far, this is the first empirical studies, which focused on the nexus between the wealth factors and tourism departures of the selected EU countries within this limited period. (2) We are using for the first time in research community a bound testing approach proposed by Pesaran et al. (2001) to investigate the cointegration between the wealth factors and tourism departures.

In this paper, we seek to prove a causal relationship between the wealth effect (and income) and tourism departures by exploiting quarterly data on tourism and wealth in selected EU countries for the period 2000–2018. The rest of this paper is organised as follows: Section 2 discussed the general model, constructs the estimation procedure (including the unit root test, ARDL approach and Granger causality analysis), and the data used in the paper; Section 3 explains the empirical results; finally, Section 4 offers concluding remarks and outlines the policy implications.

1. LITERATURE REVIEW

The existence of the wealth effect is rarely, in analysing tourism consumption, disputed, at least according to academic papers composed recently.

Past studies have showed that variations in one of the proposed wealth variables in certain regions of the world can positively affect tourism consumption. By using data from China Family Panel Studies for 2010 and 2012, Zhang and Feng (2018) found that changes in house prices had a positive and significant effect on tourism expenditure. In the study, housing stock was considered an illiquid form of wealth. Another study used quarterly data from Malaysia for the period 2000–2011 and the cointegration technique to examine the relationship between the wealth effect from real estate and outbound tourism, while controlling for other relevant outbound tourism determinants (Fereidouni, Al-Mulali and Mohammed 2017); the researchers found that the same variable (i.e., house prices) increased Malaysians' spending on international travel for leisure purposes. A paper written by Park et al. (2012) tested for the wealth effect on Korean outbound travel between 1989 and 2009; the results of this study favoured the possibility of a significant wealth effect from housing on outbound travel demand, but not from financial assets. Other studies have showed that redistributions of wealth, along with population growth, geopolitical changes and conflicts, fuel costs, climate change, new technologies and work patterns, and all forms of social fashion play a role in influencing consumers in terms of where they choose to travel, for how long, to do what, and at what prices (Buckley et al. 2015). In a theoretical paper, the wealth effect generated by revenues from overseas tourism taxation was modelled in a dynamic optimising macro model fashion (Chang, Lu and Hu 2011). Using a panel composed of Australian States, Dvornak and Kohler (2007) concluded that the housing and stock market wealth have a significant effect on consumption in a general sense. Meanwhile, a study by Bhadra (2017) investigated whether wealth had any quantifiable impact on U.S. air travel, controlling for all other relevant variables, such as current income, past wealth, fares, and credit availability. The paper found that a loss in household wealth of U.S. $17 trillion yielded a loss in air travel demand of 730,000 passengers, or a loss in revenue of $244 million. There is in general a U-shaped relationship between the period of time of owning residential property and tourism consumption, and the effects of housing wealth on tourism consumption are mainly not through the mortgage credit effect (the 'house slaves effect') but through the wealth effect (Zhang and Wang 2017). Combs (2015) is preoccupied with the topic of how to improve measurement of mobility, in particular the gap between observed travel and desired travel by resource constrained households in low-mobility contexts, in order to ensure that transit investment results in improvements in mobility and out of-home activity participation options. Forecasting long-haul tourism demand for Hong Kong using error correction models is a paper written by Lee (2011). The results indicate that the Permanent Income-Life Cycle (PI-LC) hypothesis based on the Engle–Granger (1987) approach produces more accurate forecasts for that country than other alternative forecasting models for all long-haul markets based on Mean Absolute Error (MAE) and Root Mean Square Error (RMSE) criteria. The tourism boom in Lijiang, South-West China, has increased wealth and widened economic opportunities, and led to the rapid growth of the tourist economy in Lijiang, but also increased infrastructure vulnerability of that region (Su et al. 2016).

2. GENERAL MODEL, METODOLOGY AND DATA

2.1. General model

We will distinguish between a conceptual model and the model that we have actually used for the estimate. The literature suggests a considerable range of measures of wealth-diversified inputs that can fulfil our metrics purpose in tourism consumption – the wealth nexus. One can theoretically model the wealth effect impact on tourism travel in any selected country with the help of the function, according to Zuo and Lai's (2019) formula, where consumption is related not only to income but also to tangible and intangible household assets. All of these combined variables are summed up into a wealth portfolio stock at a given time point, although some of them have a pure psychological intrinsic value for the tourism consumer. The feeling that one's house is gaining in value (e.g., thanks to bubble), or that the value of a financial bond or stock (unrelated to its annual yield) is rising can positively affect ongoing tourism-emitive markets.

depar = f (+*compemp*, +*rppi*, +*netasse*, +*stockin*)

where *depar* is international tourism departures (it is a proxy for consumption), *compemp* is compensation to employees based (a proxy variable for household income), *rppi* is the house price index, *netasse* is net financial assets, and *stockin* is financial derivatives and employee stock options. Such a formulation is in line with the aforementioned general consideration regarding the link between tourism departures and the household wealth portfolio consideration.

Accordingly, the hypotheses of this study are formulated as follows:

Hypothesis 1: An increase in income positively affects outbound travel demand.
Hypothesis 2: An increase in residential home prices boosts outbound travel demand.
Hypothesis 3: An appreciation in the value net financial assets (or savings) positively affects outbound travel demand.
Hypothesis 4: Capital asset price gains increase outbound travel demand.

2.2. Econometric methodology

2.2.1. ARDL cointegration and bounds tests

To analyse the long-term relationship between a set of variables, Pesaran et al. (2001) suggested the use of an autoregressive distributed lag procedure or bounds test that does not require stationary pre-testing, and which can be used regardless of whether the variables are I(0), I(1), or mutually cointegrated, given that none of the series is I(2). Despite these relaxing circumstances, we have made a verification to see whether a second-order integration in some time series exists by conducting a ADF unit root test in order to eliminate further exercises with data that encompass some of the wealth portfolio variables. If the dependent variable (*depar*) for some countries, after ADF unit root testing, indicates I(2), then the DF–GLS test and the Zivot–Andrews test with structural breaks in those time series will provide a final verdict regarding integration of the second order. Consequently, if those tests show that the *depar* time series variable is either I(0) or I(1), an analysis with that case country will continue with the bounds test. The bounds test is particularly appropriate for small samples, such as the one used in this paper, in which the order of integration of the variables of interest is not known or may not necessarily be the same. The bounds test is based on the following unrestricted error correction model (UECM):

$$\Delta Y_t = const + \sum_{i=1}^{k} \beta \Delta Y_{t-1} + \sum_{i=1}^{k} \gamma \Delta X_{t-1} + \omega Y_{t-1} + \theta X_{t-1} + \varepsilon_t \quad (1)$$

where Yt denotes tourism consumption measured in departures, and Xt denotes a specific wealth single input (as explained previously), with both expressed in natural logarithms. An appropriate lag selection is based on the Schwarz Bayesian Criterion (hereinafter "SBC"). The automated model selection process involves choosing the maximum lag for each regressor, and is set up to be 4 (because the data is quarterly). The ARDL procedure allows for the possibility that the variables may have different optimal lags (after the searching process has ended), whereas this is impossible with conventional cointegration procedures. The null hypothesis for no long-term relationship between tourism consumption growth and particular wealth variable growth is not rejected, by testing the F-statistic, when:

$H_0: \omega = \theta = 0$, against the alternative $H_0: \omega \neq \theta \neq 0$.

Instead of the conventional critical values, Pesaran et al. (2001) proposed a bounds test for two sets of critical variables. The first set assumes that all variables are (0), and the other set assumes that all variables are (1). If the tested F-statistic (or Wald statistic) value lies below the lower bound critical value, then the null hypothesis of a non-existent cointegration relationship cannot be rejected, and if it exceeds the respective upper bound critical value, the null hypothesis is rejected. If the tested F-statistic value falls within the lower and upper critical value bounds, inference is inconclusive. The set of the bound critical values for limited data was recently developed by Narayan (2005) (30 to 80 observations), and is the benchmark for F-statistic assessing (2005). Furthermore, because of the potential existence of a trend in the series (if the former case is unable to find cointegration between two series), estimations are completed to satisfy the unrestricted intercept and no trend case (as an auxiliary test). Estimations are completed using an ordinary least squares procedure with a White's test for cross-sectional heteroscedasticity-consistent standard errors, and a covariance matrix, appropriate serial correlation diagnostics (the Breusch–Godfrey LM test) and the Jarque–Bera statistic for the normality test.

2.2.2. Causality analysis

The bounds procedure is important for assessing how tourism consumption is affected by wealth input, by completing separate estimations of Eq. (1) using ΔYt as sole dependent variables, or to find the possibility of a cointegration link.

If there is a cointegration relationship between the variables, the next step is to assess the short-run and long-run dynamics of the series by examining the error correction model based on the Eq. (2), in which Granger causality is deduced in order to investigate the unidirection of causality that goes from wealth input to tourism departures in selected countries.

Namely, if international tourist departures as a metric of consumption and a single wealth input are regressed against one another in levels, the resulting residuals essentially represent error correction terms, which measure deviations in the long-run equilibrium between the two series. Hence, the ARDL Eq. (1) can be re-parameterised after replacing Yt−1 and Xt−1 with the lagged residuals, and become:

$$\Delta Y_t = const + \alpha ECT_{t-1} \sum_{i=1}^{n} \rho \Delta Y_{t-1} + \sum_{i=1}^{p} \sigma \Delta X_{t-1} + \mu_t \quad (2)$$

e.g., the error correction model *via* the two-step procedure of Engle and Granger.

These lagged residuals represent an error correction term, denoted in this paper by ECT (−1), which provides an insight into the speed of adjustment to a long-run equilibrium within a particular time frame from a change to one of the series. Furthermore, if the coefficient of ECT(-1) is statistically significant (by t-value), then it indicates long-run causality, as the first causality test. ECTt-1 should be between 0 and 1 with a negative sign, which implies convergence of the system back to the long-run equilibrium position.

The second type of Granger causality is short-run causality—the Wald test—which is applied for all the lag independent wealth variables using the joint F test. The lag of the individual coefficients is, in this manner, utilised to test the significance of the short-run relationship.

The third jointly lagged coefficients and the ECT are assessed to verify joint causality between wealth and tourism departures.

Additionally, μ_t represents the error terms and should be white noise and serially uncorrelated.

We will also assess reverse causality that goes from departures to wealth (rather bizarre and counterintuitive), only if the presumed direction of the short-run causality (from wealth to departures) does not occur. The same will not be interpreted. Namely, the Granger representation theorem states that if there is cointegration, then there is short-run Granger causality in at least one direction, i.e., the error correction term enters at least one of the equations of the error correction model. So, the vice-versa kind of causality, run out, test exercise shall be conducted to assess the validity of cointegration evidence (or implicitly the affirmation of the Granger representation theorem) for pure statistical curiosity reasons.

2.3. About the data

Income variability, in our study, is proxied by compensation that goes to employees as paid wages (*compemp*), the house price index (*rppi*) and its variation is set up to measure the other forms of household asset performance, along with net financial assets (*netasse*), financial derivatives, and employee stock options (*stockin*). These variables were sourced from Eurostat (2019A, 2019B, 2019C, 2019D).

Amongst the many potential positive side-effects of these variables on tourism consumption, repeatedly interpreted and re-interpreted in the tourism economics literature, we have included in our analysis only (quarterly) data regarding selected EU countries, namely Austria, Belgium, Germany Denmark, Ireland, Spain, France, Italy, Sweden, Slovenia and the UK, for the time span 2000–2018. This sample period was chosen purely for data availability in relation the measures of the desired variables. The tourism consumption profile, based on travellers by origin, fits well in our ambition to deduce the wealth effect impact on tourism consumption based on the intensity of tourism international departures.

International tourism departures (*depar)* are used as a measure of orientation towards tourism travelling, and in this study have been used as an alternative (substitutive term) to tourism consumption, which we were not able to deduce from the Eurostat site in any consistent or transparent manner.

A few extra words would be good to mention here, in this data section. The extraction of tourism expenditure data (from the Eurostat site), which is the first-rate expression of tourism consumption, depending on counts, was the first design solution. Alas, the breaks in and/or the short time span of that data hindered the compilation of the data sample, and may have brought this analysis to a swift and fatal end. Therefore, we were forced to use departures as an alternative, which were sourced from the World Bank's website. The last time series was in annualised form. The postulated requirement of designed studies, because of the ARDL time series methodology, is a disaggregation of their lower frequency value to higher (or quarterly unit) values. The tourism expenditure data and the inherited seasonality in that data enabled us to get quarterly tourism departure entries, with a prudent time span length (2000Q1–2018Q4), through temporal desegregation techniques, as explained by Chow-Lin (for more, see Sax and Steiner 2013).

All the variables used in this paper come in their natural log form.

The data series for the selected economies had 76 quartals (a moderate number); one of them, which expresses net assets, had just 46 observations, because the constitutive elements (assets – liabilities) began from 2006Q1. In the case of Germany (DE), the net financial assets time series were much shorter than is allowed by the ARDL procedure, and in the case of Slovenia (SI) we were not able to find consistent data regarding stocks and shares (the *stockin* variable). Hence, we dropped those variables in the wrangling initial phase of the data analysis process. We suppose that idiosyncratic outliers and structural breaks may be hidden in the Data Generating Process (DGP) of our time series. As pointed out above, an adequate technique to handle these handicaps is the ARDL bounds test approach.

3. EMPIRICAL RESULTS

3.1. Unit root test

Before conducting tests for cointegration, we had to make sure that the variables under consideration were not integrated at an order higher than 1. In the presence of I(2) or higher variables, the computed statistics provided by Pesaran et al. (2001) and Narayan (2005) are not valid (Ang 2007). Thus, in order to establish the integration properties of the series, we used quick ADF unit root tests to inspect that none of the series was I(2). Accordingly, the ADF test conducted was at level and at first difference, and the results are shown in Table 1 below.

Table 1: **Unit Root Test (Augmented Dickey–Fuller Test)**

		depar	**rppi**	**netasse**	**compemp**	**stockin**
AT	Levels	-1.280(7)	-0.410(1)	-4.276(2)***	-3.845(4)***	-2.123(1)
	First diff.	-0.788(7)	-5.422(1)***	-	-	-5.499(1)***
BE	Levels	-2.837(8)*	-1.521(4)	-2.586(1)	-1.174(4)	-2.644(2)
	First diff.	-	-1.570(3)	-2.721(1)	-0.752(3)	-4.668(1)***
DE	Levels	-1.52(10)	-1.748(5)	-	-1.264(5)	-2.712(1)
	First diff.	-1.265(9)	-3.033(4)**	-	-1.195(4)	-5.118(1)***
DK	Levels	-1.29(11)	-1.749(5)	-3.575(2)***	-1.567(4)	- 2.688(2)
	First diff.	-2.21(11)	-3.033(3)	-	-1.986(3)	-5.532(1)***
ES	Levels	-2.039(2)	-2.280(8)	-0,596(1)	-2.284(4)	-2.779(1)
	First diff.	-4.411(1)***	-1.776(4)	-3.468(4)**	-1.097(3)	-4.449(1)***
FR	Levels	-3.718(2)***	-0.952(5)	-1.833(1)	-1.228(1)	-2.198(2)
	First diff.	-	-3.229(4)***	-5.164(1)***	-0,963(3)	-4.898(1)***
IE	Levels	-1.634(1)	-2.192(5)	-2.431(1)	-2.538(4)	-2.180(1)
	First diff.	-2.721(1)	-2.055(4)	-4.003(1)***	-1.482(4)	-4.811(1)***
IT	Levels	-1.743(5)	-0.836(2)	-2.166(1)	-0.798(4)	-2.636(1)
	First diff.	-5.095(4)***	-2.730(1)	-4.482(2)***	-1.130(4)	-4.856(1)***
SE	Levels	-2.36(11)	-1.931(5)	-1.830(1)	-1.838(7)	-2.709(1)
	First diff.	-1.71(11)	-3.567(4)**	-2.176(4)	-1.59(11)	-4.479(1)***
SI	Levels	-2.781 (4)	-0.517(1)	-0,182(1)	-1.444(5)	-
	First diff.	-4.214(4)***	-4.079(1)***	-3.118(1)*	-1.498(4)	-
UK	Levels	-2.223(3)	-1.875(7)	-2.605(1)	-1.406(8)	-2.656(1)
	First diff.	-3.145(1)**	-1.850(5)	-5.904(1)***	-1.605(7)	-4.697(1)***

Source: Own calculation

Notes: Firstly, all the regressions include an intercept and a linear trend in the levels, and include an intercept in the first differences; secondly, the numbers in parentheses are the optimal lag orders and are selected based on Schwarz Bayesian; thirdly, *, ** and *** denote 10%, 5% and 1% level of significance, respectively.

As indicated in the table, the ADF test failed to reject the null hypothesis of the unit root at levels (unevenly for some variables), implying that the variables were non-stationary at levels. However, at first difference, the null hypothesis was rejected, implying that the variables became stationary at first difference, but not always. If first-order differences in such cases (without any adjoined asterisk above the value statistics as an orientation) do not eliminate the unit root of the variable under consideration, the maximum order of integration can be concluded to be greater than I(1). Table 1 shows that the *rppi*, *netasse* and *compemp* variables for Belgium; the *compemp* variable for Germany; *rppi* and *compemp* for Denmark/Spain/Italy/UK; and *compemp* for France and Slovenia were non-stationary, both in level terms and first differences.

Furthermore, the DF–GLS test and Zivot–Andrews unit root test results were processed in order to give final conclusive verdicts about the stationarity of the dependent variables (*depar*); the former tests have better power to detect non-stationarity, and these are given in Table 2.

Table 2: **Unit Root Test (DF–GLS Test and Zivot–Andrews Test Allowing for One Structural break)**

	depar	**DF-GLS test**	**Z-A test**
AT	Levels	-4.097(4)***	-10.420(2)*** at 2007q4
	First diff.	_	-
DE	Levels	-3.227(4)*	-3.726(2) at 2015q2
	First diff.	-3.485(4)**	-12.026(2)** at 2006q3
DK	Levels	-3.071(1)	-3.484(0) at 2016q1
	First diff.	-5.275(4)***	-7.838(2)*** at 2005q4
IE	Levels	-2.110(2)	-3.107(1) at 2005q4
	First diff.	-2.814(1)	-3.306(0) at 2009q2
SE	Levels	-2.327(4)	-3.230(2) at 2009q3
	First diff.	-2.185(3)	-3.797(2) at 2006q2

Source: Own calculation

Notes: Firstly, the lag length selected is based on the Schwarz Information Criterion (SIC), by the DF–GLS (Elliott, Rothenberg and Stock 1996) test statistics includes an intercept and a linear time trend in the levels and only an intercept in first difference; secondly, the critical values were obtained from Zivot and Andrews (1992). The null hypothesis is that a series has a unit root with a structural break in both the intercept and the trend (in levels) and intercept (in first difference); thirdly, *, ** and *** denote significance at 1%, 5% and 10%, respectively.

These results (Table 2) suggested that we drop Ireland and Sweden, exclusively for tourism departures and the wealth nexus the ARDL bounds testing, because of the revealed non-stationarity in the *depar* first difference time series after unit root re-testing. Our anxiety notwithstanding, we felt that we could confidently employ the ARDL bounds approach to our model for the other countries.

3.2. Results of the ARDL cointegration tests

In the first step in applying the bounds test, we specified the optimal lag length of the UECM, i.e., Equation (1), and checked the long-run level equilibrium relationship.

We tried to set up the best of the ARDL model, and fixed an optimal lag, which is crucial. With an initial lag of 4, the automated model selection, according to minimal SBC (Pesaran and Shin 1999), calculates the optimal lag length. Table 3 shows the estimated ARDL model that has passed several diagnostic tests, which indicates no evidence of serial correlation and heteroscedasticity, nor deviation from normal distribution.

Table 3: **Result of the cointegration test using ARDL Approach**

	Dependent variable	**Independent variable**	**F - test statistic**	**Coint-egration**	**LM-test**	**JB-test**	**HET**
AT	*depar*	*rppi*	101.235***	Yes	0.37	0.18	0.52
	depar	*netasse*	64.082***	Yes	0.42	0.32	0.10
	depar	*compemp*	264.651***	Yes	0.34	0.14	0.52
	depar	*stockin*	98.515***	Yes	0.46	0.77	0.54
BE	*depar*	*stockin*	27.396***	Yes	0.55	0.42	0.93
DE	*depar*	*rppi*	44.655***	Yes	0.52	0.21	0.57
	depar	*stockin*	43.377***	Yes	0.13	0.90	0.51
DK	*depar*	*netasse*	5.264	No	0.11	0.00	0.00
	depar	***netasse***	4.553	Inconclusive	0.08	0.00	0.00
	depar	*stockin*	4.670	No	0.06	0.00	0.00
	depar	***stockin***	2.148	No	0.09	0.00	0.00
ES	*depar*	*netasse*	7.933**	Yes	0.14	0.00	0.65
	depar	*stockin*	11.181***	Yes	0.23	0.34	0.30
FR	*depar*	*rppi*	13.507***	Yes	0.55	0.25	0.73
	depar	*netasse*	10.129***	Yes	0.29	0.11	0.85
	depar	*stockin*	8.620***	Yes	0.66	0.17	0.32
IT	*depar*	*netasse*	4.896	No	0.03	0.22	0.05
	depar	***netasse***	1.376	No	0.02	0.11	0.03
	depar	*stockin*	5.110	No	0.06	0.22	0.05
	depar	***stockin***	1.876	No	0.02	0.17	0.02
SL	*depar*	*rppi*	41.999***	Yes	0.35	0.05	0.98
	depar	*netasse*	43.182***	Yes	0.65	0.76	0.93
UK	*depar*	*netasse*	13.743***	Yes	0.34	0.72	0.45
	depar	*stockin*	5.430	No	0.13	0.82	0.42
	depar	***stockin***	5.505*	Yes	0.32	0.07	0.26

Source: Own computation

Notes: The critical values are derived from tables CI (V) and CI (III) (see Table 4). LM is the Lagrange multiplier test for serial correlation with a $x2$ distribution, with only one degree of freedom; J-B is the Jarque–Bera test for normality, HET is the White test for heteroscedasticity with a $x2$ distribution, with only one degree of freedom; Asterisks *, ** and *** denote statistical significance, respectively, at the 1%, 5% and 10% levels. Italic and bold labels for the variables indicate bounds testing repeats, according to case III.

The ARDL bounds test results show that there is no equilibrium relationship between tourism departures and the selected variables of wealth portfolio in Italy and Denmark. To be fair, the net asset variable for Denmark gave us an inconclusive result even in the relaxed case of unrestricted intercept and no trend equation frame. Therefore, we dropped both countries, in this stage of further Granger-causality testing.

Table 4: **Critical Values for the ARDL Modelling Approach Related to the Bounds Test**

	Case V		Case III	
	I(0)	I(1)	I(0)	I(1)
10% critical value	5.73	6.45	4.04	4.78
5% critical value	6.82	7.67	4.94	5.73
2.5% critical value	7.46	8.27	5.77	6.68
1% critical value	9.17	10.24	6.84	7.84

Source: Pesaran et al. (2001); case V and case III are related to ‘unrestricted intercept, unrestricted trend’, ‘unrestricted intercept, no trend’, and ARDL regression, respectively.

The Breusch-Godfrey Serial Correlation LM Test results presented in Table 3 show that there were no problems of serial autocorrelation (beside the dropped countries in few cases). In addition, the diagnostic test for heteroscedasticity and normality also showed the absence of such problem. This indicates that the model was good founded and suitable enough for the study of cointegration among the variables. There is plenty of evidence of a unidirectional relationship for the other countries, which goes from extracted (specific) wealth input to tourism departures, as shown in Table 3. The bounds F-statistic for the cointegration test statistics lies above the upper bound critical values, and therefore indicates a long-run relationship between the proceeding (or preselected) wealth portfolio and the tourism departures time series for Austria, and employee stock options that mimic wealth for Belgium. Furthermore, the findings indicate a relationship between house prices and employee stock options for Germany; net assets and employee stock options for Spain; house prices, net assets and employee stock options for France; and house prices and net assets for Slovenia. The related cointegration significance is at the 1 per cent significance level. What is more, the bounds F-statistic confirms a long-run link between tourism departures and net assets, at the 5 per cent significance level, but related departures to stock options at the 1 per cent significance level for Spain. In the case of the UK, the bounds test failed to report a long-run relationship with stock options in the conventional UECM framing, but succeeded by the unrestricted intercept and no trend form equation at the 10 per cent significance level. Net assets was the last variable that was cointegrated with departures, at the 1 per cent significance level for the UK.

3.3. Results of causality tests

The existence of an ARDL cointegration relationship between tourism departures and wealth variables, as depicted above, for Austria, Belgium, Germany, France, Spain, Slovenia and the UK provides that there should be Granger causality in at least one direction. The causality test results are shown in Table 5.

Table 5: **Results of Granger causality**

	Dependent variable	**Independent variable**	**t-Statistics (Long-Run)**	**F-Statistics (Probability) (Short-Run)**	**Joint (Short- and Long-Run)**
AT	*depar*	*rppi*	-0.367*** (-3.587)	4.975 ** (0.022)	18.886*** (0.000)
	depar	*netasse*	-0.939*** (-3.467)	1.582 3.465** (0.213) (0.042)	14.976*** (0.000)
	depar	*compemp*	-0.436*** (-2.765)	7.677*** (0.000)	18.047*** (0.000)
	depar	*stockin*	-0.619 (-2.345)***	4.788** (0.011)	23.525*** (0.000)
BE	*depar*	*stockin*	-0.280*** (-2.560)	3.577** (0.033)	12.856 (0.000)***
DE	*depar*	*rppi*	-0.581*** (-2.975)	3.206** (0.047)	6.287 (0.003)***
	depar	*stockin*	-0.442*** (-3.004)	3.636** (0.032)	6.725*** (0.002)
ES	*depar*	*netasse*	-0.791* (-1.856)	6.586** (0.003)	30.437*** (0.000)
	depar	*stockin*	-0.408*** (-2.452)	16.537*** (0.000)	53.076*** (0.000)
FR	*depar*	*rppi*	-0.145*** (-3.376)	25.308*** (0.000)	46.208*** (0.000)
	depar	*netasse*	-0.165*** (-3.345)	17.657*** (0.000)	41.004*** (0.000)
	depar	*stockin*	-0.632* (-1.804)	17.647*** (0.000)	31.477*** (0.000)
SL	*depar*	*rppi*	-0.225** (-2.006)	0.585 4.057** (0.567) (0.021)	10.555** (0.003)
	depar	*netasse*	-0.108* (-1.867)	1.186 3.745** (0.321) (0.033)	11.314** (0.004)
UK	*depar*	*netasse*	-0.195** (-2.005)	20.276*** (0.000)	61.585*** (0.000)
	depar	*stockin*	0.765** (-2.030)	32.065*** (0.000)	59.565*** (0.000)

Source: Own calculations

Notes:

- Long-run Granger causality is conducted by the t-statistics of α coefficient, which stands before the ETC term that measures how fast the deviations from the long-run equilibrium die out following changes in each variable, according to Eq. (2).
- Short-run Granger causality is conducted by testing $H0 : \sigma = 0$ that stands as a coefficient before the wealth variable for all p lags, according to Eq. (2). The figure in italics captures the wealth input as a dependent variable, F-stat. (objective is to inspect reverse causality).
- Joint or strong Granger causality is detected by testing $H0 : \sigma = \alpha = 0$ for each wealth variable respectively, according to Eq. (2).

There was long-run unidirectional Granger causality running from various sources of wealth to tourism departures. Specifically, from house prices, net financial assets, income and stock options in Austria; only from stock options in Belgium; from house prices and stock options in Germany; net assets and stock options in Spain; house prices, net assets and stock options in France; house prices and net assets in Slovenia; and net assets and stock options in the UK. The coefficient estimates for the lagged error correction terms (ECT-1) ranged between a low of 14.5 per cent for the housing price index for France and a high of 79 per cent for net financial assets for Spain, indicating the percentage of adjustment towards a long-run equilibrium that occurs within a quarterly interval. Meanwhile, the t-statistics of the coefficients of the lagged error correction terms (ECTt_1) indicate the statistical significance of the long-run causal effects.

Analysis also reveals reciprocal short-run unidirectional Granger causality running from various sources of wealth to tourism departures. Specifically, as was foreseen theoretically, in the same direction—from wealth to tourism departures—for all the mentioned cases except in a very few cases, where such causality got reversed (tourism departures increased Austria’s net assets and Slovenia’s house prices and net assets).

Amongst other things, there were strong (joint) unidirectional Granger causalities running from wealth to tourism departures for all the countries left over in this final stage of the analysis. The results can be copied here exactly, from the discussion about long-term causality cases, without any new particularity arising.

Concerning income hypothesis in estimating tourism departures, our results supported observations and conclusions of previous writers and researchers (e.g. Zuo and Lai 2019; Gržinić et al. 2017; Kim et al. 2012; Lim 1997; Zhou and Li 2010). As such, the second fragmentary context of this study is validated since previous research has shown that the rising housing wealth positively affects the demand for outbound tourism (Park et al. 2011); Kim et al. 2012; Fereidouni et al. 2017; Zhang and Feng 2018; Zuo and Lai 2019). Our result in this paper, which examines budgetary constraints (net financial assets plus employee stock options) in international tourism departures confirm the finding as in Wang, Y. S. (2014).

CONCLUSION

This paper was directed towards attaining a full understanding of the causal relationships between the wealth effect, along with income, and tourism departures, which is invaluable for the implementation of any relevant policy measures/instruments in advancing tourism departure flows in the selected EU countries. For the countries that passed the rigorous statistical testing—the unit root, cointegration and bounds testing (Pesaran et al. 2001)—Granger causality analysis was also particularly actual. Tourism consumption/international departures take place in unusually high structural-change environments concerning wealth restructuring (along with their composing particle complexities, ranging from real estate prices and stock volatility to net financial assets and income). Namely, means of the middle-class tourist visitor, who traditionally is heavily in debt, kind of tourist consumer. Uncertainty about the future (anxiety lurks behind every leisure consumption) for much of the past decade amongst the majority of potential travellers will not be lessened by this paper's messages. However, we hope that the identification of unidirectional causality (long-/short-run and joint), first of all for Austria (all wealth metrics if we exclude short-term causality, along with income), as well as France, Germany and other countries in minor cases, can produce some kind of causality (although in long-term and joint causality that causality runs interactively through in all cases). Some of our findings, that link wealth effect and tourism departures (comparatively amongst the countries) are at first sight curious, but they are based on obtained statistical results (which are, as ever, full of surprises).

Our main contribution in this paper consists of analysing long-run interactions between the wealth effect, along with income, and tourism departures for possible disparity in causality pattern among the selected EU countries. Concerning income hypothesis (based on wage magnitude) in assessing tourism departures casual direction, our results, to which we contribute to the literature, in the case of Austria, Spain, France, Slovenia and the UK supported findings of other authors (e.g. Zuo and Lai 2019; Gržinić et al. 2017; Kim et al. 2012; Lim 1997; Zhou and Li 2010). The rise in wealth stemming from the tourists' own homes price increase as a causal impulse to tourism departures in the case of Austria, Germany, France, and Slovenia, as a subsequent contribution, in this paper, it is a proven fact. The same is confirmed in a similar way but by other methods in papers written by Park et al. (2011); Kim et al. (2012); Fereidouni et al. (2017); Zhang and Feng (2018); Zuo and Lai (2019). Our contribution in support of the thesis that the budget constraint is valid factor in causing of tourist departures for a full set of chosen countries is alike to result revealed by Wang, Y. S. (2014).

Will the findings in this paper mitigate somehow the uncertainty regarding tourism demand arising from these countries? Policy measures, in this regard, referring to the short-term causality wealth effect (inner consistency of our findings), must contain better management of real estate prices (Austria, Germany, France and Slovenia). In addition, more ESOP stock option programmes that reward workers and redistribute financial wealth towards the less privileged (all the countries except Slovenia), furthermore, policies that stimulate saving (Spain, France, the UK), and higher income (Austria). If the household's wealth continue to growth and flourish smoothly, then it will increase the tourism revenue in the future.

Finally, it can be said that the tourism and wealth effect relationship is one of the neglected topic in tourism research, which currently is not fully explored. Despite our best intention, some weak points in research occurred; sample group in few cases do not feet well in ARDL approach, and those countries are predetermined to live no trace in testing our hypothesis about wealth effect-tourism causality. This shortcoming, in the future analysis can be bypass by panel cointegration (with more countries gathering in regression).

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