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The nature of the relationship between international tourism and international trade: the case of German imports of Spanish wine

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Christian Fischer and Luis A. Gil-Alana

Abstract

This paper deals with the relationship between international trade and tourism. In particular, we focus on the effect that German tourism to Spain has on German imports of Spanish wine. Due to the different stochastic properties of the series under analysis, which display different orders of integration, we use a methodology based on long memory regression models, where tourism is supposed to be exogenous. The results show that at the aggregate level, tourism has an effect on wine imports that lasts between two and nine months. Disaggregating the imports across the different types of wine it is observed that only for red wines from Navarra, Penedús and Valdepeñas, and to a certain extent for sparkling wine, tourism produces an effect on future import demand. From a policy-making perspective our results imply that the impact of tourism on the host economy is not only direct and short-term but also oblique and delayed, thus reinforcing the case for tourism as a means for economic development.

Keywords: International trade; tourism; long memory; Spanish wine.

JEL-classification: F14, C22, Q13, L83

1 Introduction

International tourism has grown strongly during the last decades. Worldwide tourist trips reached almost 700 million in 2000, as compared to about 25 million in 1950. Measured in relative terms, at 120 trips per thousand of world population in 2000, tourism activity has increased more than tenfold during this period (World Tourism Organisation, various issues). At the same time, international merchandise trade has also grown significantly. According to figures of the World Trade Organisation, world exports per head of world population have increased from 44 current US dollars in 1961 to almost 1,200 USD in 2003. Thus, because there seems to be correlation between the international flows of people and goods a growing body of studies has emerged, which, using different methodologies, have investigated the possibility of a causal relationship between these two phenomena.

From a policy perspective, the topic of possible significant and positive interdependencies between tourism inflows and exports of manufactures is important in at least two ways. First, industrial development officers and trade association officials may find it useful to better understand the dynamics and determinants of

industrial export success. While, in practice, it may be difficult to actively influence tourism arrivals, the knowledge about confirmed tourism-trade interdependencies may enhance the ability to predict exports by taking into consideration tourism data. Second, tourism development agencies could demonstrate that the positive impacts of international travel on a national economy may be multiple and lasting. If tourists can be shown of not only generating income and jobs while they are in the country, but also of creating significant economic impulses by means of resulting exports to their source countries later on, the attention given to tourism development may perhaps be raised.

Easton (1998) analysed whether Canadian total bi-lateral exports are complementary or substitutive to tourist arrivals, using pooled data regressions. The study finds "some evidence of substitution of Canadian exports for tourist excursions to Canada" (p. 542) by showing that when the relative price of exports goes up, the number of tourists visiting Canada increases. Kulendran and Wilson (2000) analysed the direction of causality between different travel and (aggregate) trade categories for Australia and its four main trading partners, using time series data. Their results show that travel Granger causes international trade in some cases and vice versa in others. Shan and Wilson (2001) replicate this latter approach and also find two-way Granger causality using aggregate data for the case of China. Aradhyula and Tronstad (2003) used cross-section data and a simultaneous bivariate qualitative choice model to show that cross-border business trips have a significant and positive effect on US agribusinesses' propensity to trade. Fischer (2004) explored the connection between bilateral aggregate imports and imports of individual products and bilateral tourist flows, using an error-correction model. The results show that trade-tourism elasticities are consistently higher for individual products. Table 1 summarises the previous research on the topic.

Table 1. Overview of previous studies analysing the relationship between international tourism and trade

<i>Re-search focus</i>	<i>Categories</i>	<i>Studies</i>
Country	Bilateral trade flows Country-world/region	Easton (1998); Kulendran & Wilson (2000); Fischer (2004). Shan & Wilson (2001).
Product	Aggregate (e.g., total country exports) Specific products	Easton (1998); Kulendran & Wilson (2000); Shan & Wilson (2001); Fischer (2004). Fischer (2004).
Trade categories	Exports (X) Imports (M) Total trade (X + M)	Easton (1998); Kulendran & Wilson (2000). Kulendran & Wilson (2000); Fischer (2004). Kulendran & Wilson (2000); Aradhyula & Tronstad (2003).
Travel categories	Total travel Business travel Holiday travel Visiting friends and relatives Other travel	Easton (1998); Kulendran & Wilson (2000); Fischer (2004). Kulendran & Wilson (2000); Aradhyula & Tronstad (2003). Kulendran & Wilson (2000); Aradhyula & Tronstad (2003). Kulendran & Wilson (2000). –
Direction of causality	TOUR → TRADE TOUR ← TRADE	Kulendran & Wilson (2000); Shan & Wilson (2001). Kulendran & Wilson (2000); Shan & Wilson (2001).
Nature of relationship	Substitutive Complementary	Easton (1998). –
Type of data	Time series - quarterly data - yearly data Cross section Panel data	Kulendran & Wilson (2000); Fischer (2004). Aradhyula & Tronstad (2003). Easton (1998).

Source: authors's compilation

One of the aspects which has not yet been explored according to our knowledge is the temporal nature of the relationship between tourism and trade. As Kulendran and Wilson (2000) put it when recommending further research directions: "... the lag structures between the travel and trade flows ... may require further attention" (p. 1007).

This paper deals with the case of the relationship between the German wine imports from Spain and the number of German travellers to that country. It aims at analysing empirically whether tourism has an effect on future imports and, if yes, at determining the length of this effect. We concentrate on the case of wine due to several reasons. First, wine has become a truly globalised industry with about 40% of production (in value terms) being exported worldwide in 2001 (Anderson, 2004). Second, in industrialised nations, wine is a commonly available commodity offered in a large variety mostly differentiated by production origin. Given that objective wine quality is hard to assess for non-expert consumers, the origin of a wine is often used as a short-cut quality indicator in cases where the country of origin is associated with a preferred holiday destination (Felzenstein et al., 2004). Last, wine imports have been shown to display the most significant connection with tourism activities among a range of investigated products in previous studies (Fischer, 2004). Using Spain and Germany as the two countries of investigation seems particularly interesting given the fact that Spain is both a significant exporter of wine and an important tourist destination, while Germany is an important wine import market and a main tourism source country. In 2001, Spain was the world's third biggest wine exporter (in value terms) while Germany was its third most important wine import market, behind the UK and the US (Anderson, 2004). In 2000, Spanish wine accounted for 14.8% of total German import value and imports were up 12% on the year earlier (Storchmann and Schamel, 2004). In Spain, the largest incoming tourist group were Germans from 1997 (7.8 million stays) to 2002 (11.3m), according to Eurostat data. Only in 2003, British tourists (12.2m) outnumbered the German ones (10.4m).

The organisation of this paper is as follows: in Section 2, we briefly describe the methodology employed in the paper and present the econometric model used in the empirical work. In Section 3 we describe the time series used in the analysis and examine their statistical properties. Section 4 contains the empirical application relating total German wine imports from Spain with the number of German travellers to Spain. In Section 5 we disaggregate the import wine series according to the different products, while Section 6 concludes.

2 The econometric model

As mentioned in Section 1, most of the time series work examining the relationship between international trade and tourism is based on cointegration techniques. However, that methodology imposes a priori the assumption that the individual series must share the same degree of integration, usually 1. In other words, the

series must be individually $I(1)$, and they will be cointegrated if there exists a linear combination of them that is $I(0)$ stationary.¹

In the context of the series analysed in the present paper (which are aggregate wine imports and total tourism), we face however with various problems. First the two series do not possess the same order of integration. In fact, as explained later in Section 3, the wine imports data is $I(0)$, while tourism is clearly nonstationary $I(1)$. Moreover, the later series presents a clear seasonal pattern, while the former does not. If the two series were in fact seasonally integrated, cointegration could still be the methodology used, along the lines of the procedure suggested by Hylleberg, Engle, Granger and Yoo (HEGY, 1990). (See, e.g., Kulendran and Wilson, 2000.) However, a simple inspection at Figure 1 shows that the aggregate wine import series is not seasonally integrated.

For many years, seasonality was considered as a component that obscured the time series properties of the data, and seasonal adjustment procedures were implemented to sort this problem out. However, these methods have been strongly criticised in recent years on the basis that seasonal data contains some statistical relevant information by themselves. Contributors to this view include Ghysels (1988), Barsky and Miron (1989), Braun and Evans (1995) and others. The first two articles point out that seasonal adjustment procedures lead to mistaken inferences about economic relationships between time series data. In this paper we deal with the seasonal problem in tourism by using two approaches. First, we deseasonalise the series by using seasonal dummy variables. As a second approach, we take first seasonal differences (on the logged series), such that the series then represents monthly growth rates. Looking at the orders of integration of the two deseasonalised series, we still face with the problem that both series are now $I(1)$, while wine import is $I(0)$, invalidating thus the analysis based on cointegration.

What we propose in this paper is to look at the relationship between the two variables (aggregate wine imports and tourism) by using a new methodology based on fractional integration.

We say that a time series $\{x_t, t = 0, \pm 1, \dots\}$ is integrated of order d (and denoted by $x_t \sim I(d)$) if:

$$(1 - L)^d x_t = u_t, \quad t = 1, 2, \dots, \quad (1)$$

where u_t is $I(0)$ and L is the lag operator ($Lx_t = x_{t-1}$). The polynomial above can be expressed in terms of its binomial expansion, such that for all real d ,

¹ For the purpose of the present paper we define an $I(0)$ process $\{u_t, t = 0, \pm 1, \dots\}$ as a covariance stationary process with a spectral density function that is positive and finite at the zero frequency. An $I(1)$ process is then defined as a process that requires first differences to get $I(0)$ stationarity.

$$(1 - L)^d = \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j = 1 - dL + \frac{d(d-1)}{2} L^2 - \dots$$

The macroeconomic literature has usually stressed the cases of $d = 0$ and 1 . However, d can be any real number. Clearly, if $d = 0$ in (1), $x_t = u_t$, and a “weakly autocorrelated” (e.g., ARMA) x_t is allowed for. However, if $d > 0$, x_t is said to be a long memory process, also called “strongly autocorrelated” because of the strong association between observations that are widely separated in time. As d increases beyond 0.5 and through 1 , x_t can be viewed as becoming “more nonstationary”, for example, in the sense that the variance of partial sums increases in magnitude. These processes were introduced by Granger (1980, 1981), Granger and Joyeux (1980) and Hosking (1981), and they were theoretically justified in terms of aggregation of ARMA processes with randomly varying coefficients by Robinson (1978) and Granger (1980). Ciozek-Georges and Mandelbrot (1995), Taqqu et al. (1997) and Chambers (1998) also use aggregation to motivate long memory processes, while Parke (1999) uses a closely related discrete time error duration model. More recently, Diebold and Inoue (2001) propose another source of long memory based on regime-switching models. Empirical applications based on fractional models like (1) are among others Diebold and Rudebusch (1989), Baillie and Bollerslev (1994), Gil-Alana and Robinson (1997), and Gil-Alana (2000).²

To determine the appropriate degree of integration in a given time series is important from both economic and statistical viewpoints. If $d = 0$, the series is covariance stationary and possesses “short memory”, with the autocorrelations decaying fairly rapidly. If d belongs to the interval $(0, 0.5)$, x_t is still covariance stationary. However, the autocorrelations take a much longer time to disappear than in the previous case. If $d \in [0.5, 1)$, the series is no longer covariance stationary, but it is still mean reverting, i.e., with the effect of the shocks dying away in the long run. Finally, if $d \geq 1$, x_t is nonstationary and non-mean reverting. Thus, the fractional differencing parameter d plays a crucial role in describing the persistence in the time series behavior: the higher the d , the higher the level of association between the observations.³

We now consider the following model,

$$y_t = \beta^t z_t + x_t, \quad t = 1, 2, \dots \quad (2)$$

² See also Baillie (1996) for an interesting review of $I(d)$ models.

³ At the other end, if $d < 0$, x_t is said to be “anti-persistent”, because the spectral density function is dominated by high frequency components. See Mandelbrot (1977).

where y_t is a raw time series; β is a $(k \times 1)$ vector of unknown parameters; z_t is a $(k \times 1)$ vector of deterministic (or weakly exogenous) variables, and x_t is given by (1).

Robinson (1994) proposed a Lagrange Multiplier (LM) test of the null hypothesis:

$$H_o : d = d_o, \quad (3)$$

in a model given by (1) and (2) for any real value d_o . Under H_o (3), the residuals are

$$\hat{u}_t = (1 - L)^{d_o} y_t - \hat{\beta}' w_t, \quad t = 1, 2, \dots$$

where

$$\hat{\beta} = \left(\sum_{t=1}^T w_t w_t' \right)^{-1} \sum_{t=1}^T w_t (1 - L)^{d_o} y_t; \quad w_t = (1 - L)^{d_o} z_t,$$

where T means the sample size. Thus, if $d_o = 1$, we are testing for a unit root, though other fractional values of d are also testable. The functional form of the test statistic (denoted by \hat{r}) is described in Appendix A.

Based on the null hypothesis (3), Robinson (1994) established that under some regularity conditions:⁴

$$\hat{r} \rightarrow_d N(0,1) \quad \text{as } T \rightarrow \infty,$$

and also the Pitman efficiency of the test against local departures from the null.⁵ Thus, we are in a classical large sample-testing situation: an approximate one-sided $100\alpha\%$ level test of H_o (3) against the alternative: $H_a: d > d_o$ ($d < d_o$) will be given by the rule “Reject H_o (3) if $\hat{r} > z_\alpha$ ($\hat{r} < -z_\alpha$)”, where the probability that a standard normal variate exceeds z_α is α . As in other standard large-sample testing situations, Wald and LR test statistics against fractional alternatives will have the same null and local limit theory as the LM tests of Robinson (1994). Sowell (1992) employed essentially such a Wald testing procedure but it requires an efficient estimate of d , and while such estimates can be obtained, the LM procedure of Robinson (1994) seems computationally more attractive.

⁴ These conditions are very mild, and concern technical assumptions to be satisfied by the model in (1) and (2).

⁵ This means that the test is the most efficient one when directed against local alternatives. In other words, if we direct the tests against the alternative: $H_a: d = d_o + \delta T^{-1/2}$, the limit distribution is normal, with variance 1 and mean that cannot be exceeded in absolute value by any rival regular statistic.

3 The time series data

The trade series (German imports of Spanish wine in euro) were obtained from two different Eurostat databases. First, aggregate imports were taken from “DS-016894 – EU trade since 1995 by HS2-HS4”. The source of the disaggregated data is the “DS-016890 – EU trade since 1995 by CN8” database. The latter database contains about two dozens of different wine categories. From these the eight most important ones (referred to as products A to H in our analysis) were chosen. These together represent about 62% of total German wine imports on average over the period of investigation. Mainly due to data availability, this period has been selected reaching from 1998m1 to 2004m11. All products are quality wines which are produced in certain Spanish areas and are sold with a controlled denomination of origin label. The different wine types are described in more detail in Table 2.

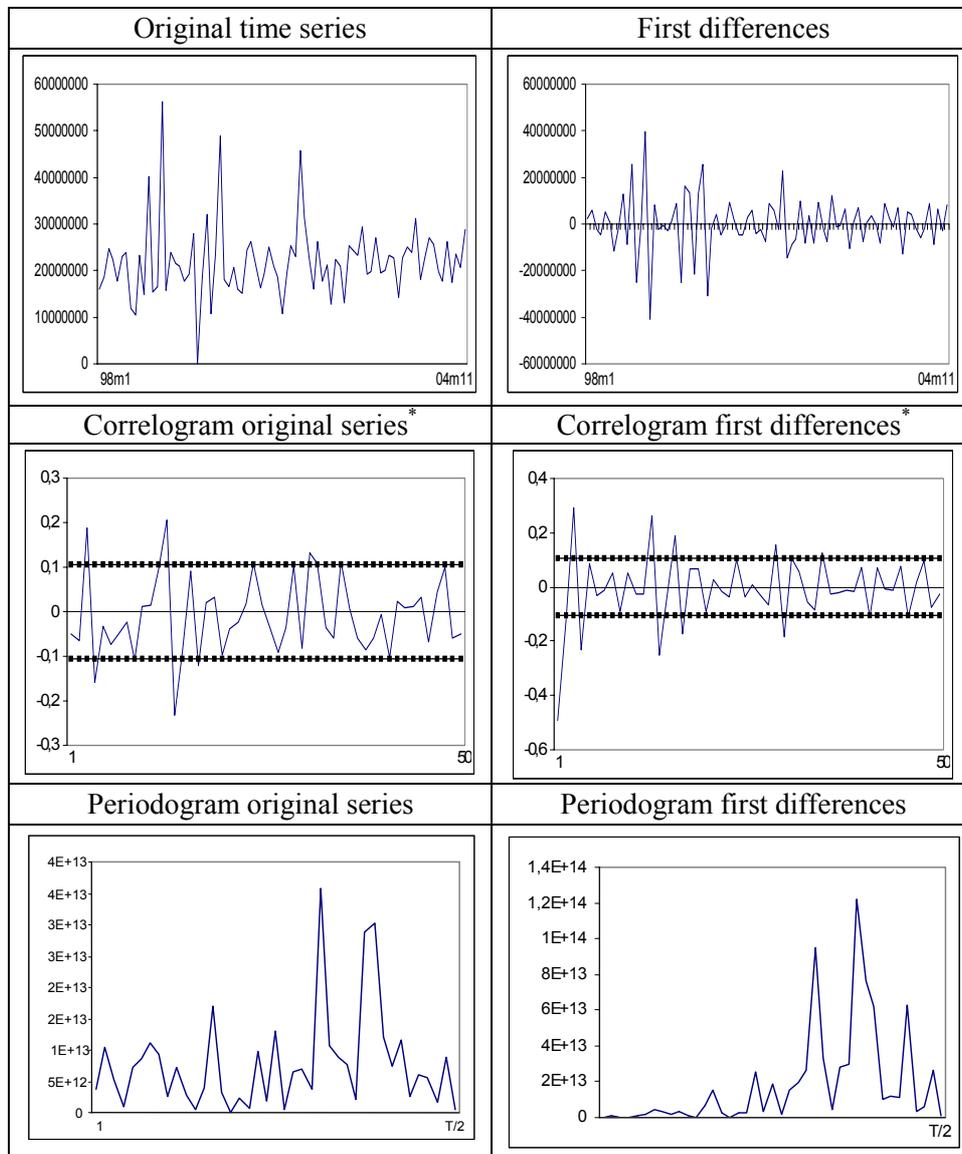
Table 2. Overview of products (German imports of Spanish wines) used in the empirical analysis

<i>Ab- bre- via- tion</i>	<i>Product code</i>	<i>Product description</i>
AWI	HS4 2204	Wine of fresh grapes, incl. fortified wines; grape must, partly fermented and of an actual alcoholic strength of > 0.5% vol or grape must with added alcohol of an actual alcoholic strength of > 0.5% vol
A	CN8 22041019	Sparkling wine of fresh grapes of actual alcoholic strength of $\geq 8.5\%$ vol (excl. champagne)
B	CN8 22042134	Quality white wines produced in Penedús, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $\leq 13\%$ vol (excl. sparkling wine and semi-sparkling wine)
C	CN8 22042136	Quality white wines produced in Rioja, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $\leq 13\%$ vol (excl. sparkling wine and semi-sparkling wine)
D	CN8 22042171	Quality wines produced in Navarra, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $\leq 13\%$ vol (other than sparkling wine, semi-sparkling wine and general white wine)
E	CN8 22042174	Quality wines produced in Penedús, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $\leq 13\%$ vol (other than sparkling wine, semi-sparkling wine and general white wine)
F	CN8 22042176	Quality wines produced in Rioja, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $\leq 13\%$ vol (other than sparkling wine, semi-sparkling wine and general white wine)
G	CN8 22042177	Quality wines produced in Valdepeñas, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $\leq 13\%$ vol (other than sparkling wine, semi-sparkling wine and general white wine)
H	CN8 22042192	Sherry, in containers holding ≤ 2 l and of an actual alcoholic strength by volume of $> 15\%$ vol to 18% vol

Source: Eurostat

Figure 1 displays plots of the aggregate wine imports and its first differences, along with their corresponding correlograms and periodograms. Starting with the original series, we observe that though there are some peaks at some probably seasonal values, it seems that there is no strong seasonal component. The correlogram and the periodogram of the original data seem to indicate that the series is

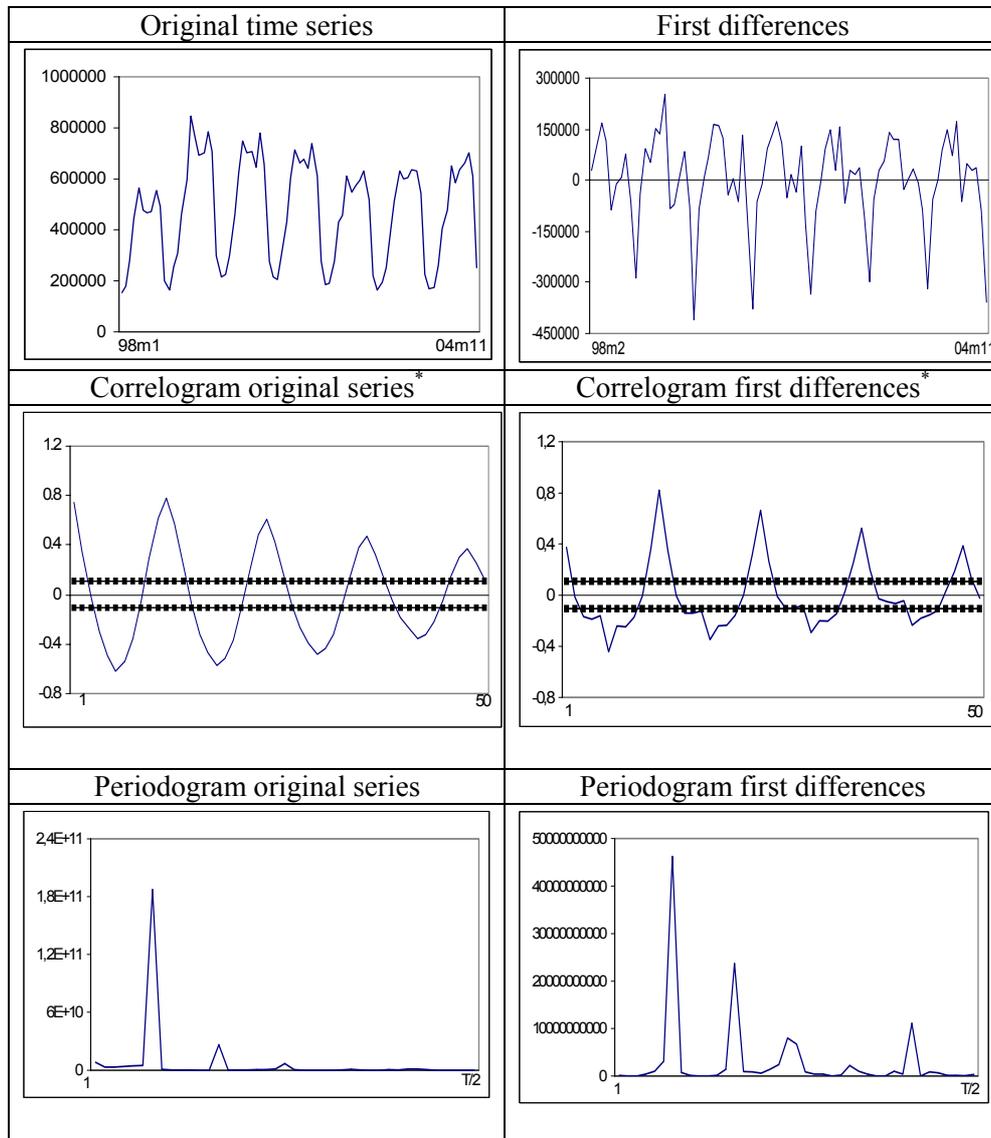
$I(0)$. In fact, if we take first differences both the correlogram and the periodogram show that the series is then overdifferenced with respect to the zero frequency.⁶



The large sample standard error under the null hypothesis of no autocorrelation is $1/\sqrt{T}$ or roughly 0.038. The periodograms were computed based on the discrete Fourier frequencies $\lambda_j = 2\pi j/T, j = 1, 2, \dots, T/2$.

Figure 1. German aggregate wine imports from Spain.

⁶ The periodogram is an asymptotic unbiased estimate of the spectral density function $f(\lambda)$. If a series is $I(0)$, $0 < f(0) < \infty$, and if it is overdifferenced, $f(0) = 0$. Thus, the periodogram should mimic that behaviour.



The large sample standard error under the null hypothesis of no autocorrelation is $1/\sqrt{T}$ or roughly 0.038. The periodograms were computed based on the discrete Fourier frequencies $\lambda_j = 2\pi j/T, j = 1, 2, \dots, T/2$.

Figure 2. German travellers to Spain.

For the tourism time series we use the number of German people travelling to Spain, monthly, for the same time period as before, obtained from the Instituto Nacional de Estadística (INE). Plots of the data, the first differences and their correlograms and periodograms are displayed in Figure 2. Contrary to the previ-

ous figure, the values here show a strong seasonal pattern and this becomes even clearer by looking at the correlograms and periodograms.⁷

Since we are mainly interested in the relationship between the two variables, the first thing we do is to analyze the statistical properties of each of the variables individually. For this purpose, we first implemented some classic methods to investigate if the series are stationary I(0) or nonstationary I(1). In particular, we use the tests proposed by Dickey and Fuller (ADF, 1979), Phillips and Perron (PP, 1988) and Kwiatkowski, Phillips, Schmidt and Shin (KPSS, 1992).⁸

The results of the above procedures for the Aggregate Wine Imports are displayed in Table 3. We observe that using no regressors, the tests cannot reject the hypothesis of a unit root. However, including an intercept and/or a linear trend, this hypothesis is rejected in all cases in favour of stationarity.

Table 3. Unit root tests for the for Aggregate Wine Imports

	<i>No regressors</i>	<i>With an intercept</i>	<i>With a linear trend</i>
ADF	-0.39 (-1.94)	-4.59 (-2.90)	-4.56 (-3.47)
PP	-1.42 (-1.94)	-10.0 (-2.90)	-9.99 (-3.47)
KPSS	---	0.076 (0.46)	0.075 (0.14)

The original series seems to displays no trend but has a non-zero mean. Therefore one would tend to use the 'intercept' test results, clearly implying AWI is I(0), i.e., stationary. In parenthesis the critical values at the 5% level.

Anyway, the use of these procedures for testing the order of integration of the series is too restrictive in the sense that they only consider integer values for the degree of differentiation. Moreover, it is a well-known stylised fact that the above unit-root procedures have very low power if the alternatives are of a fractional form (Diebold and Rudebusch, 1991, Hassler and Wolters, 1994, etc.). Across Table 4 and Figure 3 we display the results for the AWI series based on two approaches for estimating and testing the order of integration of the series from a fractional point of view.

The results in Table 4 refers to the parametric approach of Robinson (1994) described in Section 2, assuming that z_t in (2) is a deterministic component that might include a constant (i.e. $z_t = 1$) or a linear time trend (i.e. $z_t = (1, t)'$). In other words, we test the null hypothesis (3): $d = d_0$, for any real value d_0 in the model given by:

$$y_t = \alpha + \beta t + x_t, \quad (4)$$

⁷ As an alternative definition of German tourism in Spain we also employed the number of nights spent in Spanish hotels by German travellers. However, the series presents a similar pattern as the one used in the present article.

⁸ The first two methods (ADF, PP) test the null hypothesis of a unit root (I(1)) against the alternative of stationarity, while KPSS tests the null of stationarity against the alternative of a unit root.

$$(1 - L)^d x_t = u_t, \quad (5)$$

assuming that u_t is white noise and also autocorrelated. In the latter case, we use the Bloomfield (1973) exponential spectral model. This is a non-parametric approach of modelling the $I(0)$ disturbances, in which the spectral density function of u_t is given by:

$$f(\lambda; \tau) = \frac{\sigma^2}{2\pi} \exp\left(2 \sum_{r=1}^p \tau_r \cos(\lambda r)\right),$$

where p refers to the number of parameters required to describe the short run dynamics. Like the stationary $AR(p)$ model, the Bloomfield (1973) model has exponentially decaying autocorrelations and thus, we can use a model like this for u_t in (5). Moreover, this model accommodates very well to the functional form of the test statistic \hat{r} , and we can replace \hat{A} in Appendix A by the population quantity:

$$\sum_{l=p+1}^{\infty} l^{-2} = \frac{\pi^2}{6} - \sum_{l=1}^p l^{-2},$$

which indeed is constant with respect to the τ_r (unlike the AR case).

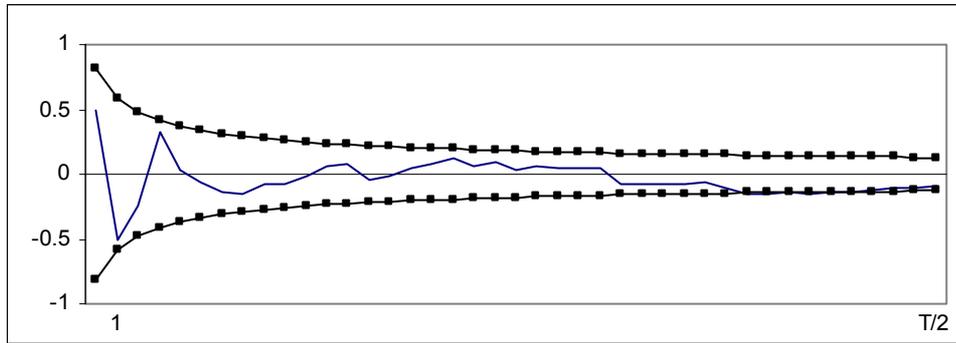
Table 4 displays the 95% confidence intervals of those values of d_0 where H_0 (3) cannot be rejected for the three cases of no regressors (in the undifferenced regression (4)), an intercept, and an intercept and a linear time trend. These confidence intervals were built up according to the following strategy. First, we choose a value of d_0 from a grid. Then, we form the test statistic testing the null for this value. If the null is rejected at the 5% level, we discard this value of d_0 . Otherwise, we keep it. An interval is then obtained after considering all the values of d_0 in the grid. We also report in the tables, (in parenthesis within the brackets), the value of d_0 (d_0^*) which produces the lowest statistic in absolute value across d_0 . That value should be an approximation to the maximum likelihood estimate.⁹ We observe that the intervals include the $I(0)$ null in all cases, the values of d ranging from -0.37 (Bloomfield u_t with an intercept and/or a linear time trend) and 0.39 (Bloomfield with no regressors). Moreover, the values of d producing the lowest statistics are in all cases negative, implying thus anti-persistent behaviour.

⁹ Note that the LM procedure of Robinson (1994) is based on the Whittle function, which is an approximation to the likelihood function.

Table 4. 95% confidence intervals of the non-rejection values of d for Aggregate Wine Imports.

	<i>No regressors</i>	<i>With an intercept</i>	<i>With a linear trend</i>
White noise	[-0.15 (-0.11) 0.06]	[-0.26 (-0.14) 0.02]	[-0.34 (-0.21) 0.02]
Bloomfield ($p = 1$)	[-0.16 (-0.09) 0.32]	[-0.33 (-0.08) 0.31]	[-0.35 (-0.26) 0.28]
Bloomfield ($p = 2$)	[-0.17 (-0.13) 0.39]	[-0.37 (-0.18) 0.36]	[-0.37 (-0.19) 0.34]

The values in parenthesis within the brackets refer to the value of d producing the lowest statistic. It is supposed to be an approximation to the MLE.



The horizontal axe refers to the bandwidth parameter number, while the vertical one corresponds to the estimated values of d . The dotted line refers to the 95% confidence interval for the $I(0)$ hypothesis.

Figure 3. Estimates of d based on Gaussian semiparametric estimate for AWI.

As an alternative approach to estimate d , we also use a semiparametric method proposed by Robinson (1995). It is semiparametric in the sense that we do not have to specify any particular model for the $I(0)$ disturbances u_t . It is basically a local “Whittle estimate” in the frequency domain, based on a band of frequencies that degenerates to zero. The proper form of the estimate (\hat{d}) is described in Appendix B. Under finiteness of the fourth moment and other mild conditions, Robinson (1995) proved that:

$$\sqrt{m} (\hat{d} - d_o) \rightarrow_d N(0, 1/4) \quad \text{as } T \rightarrow \infty,$$

where m is a bandwidth parameter number, d_o is the true value of d and with the only additional requirement that $m \rightarrow \infty$ slower than T .¹⁰ Robinson (1995) showed that m must be smaller than $T/2$ to avoid aliasing effects.

Figure 3 displays the estimates of d for values of m from 1 to $T/2$.¹¹ We also include in the figure the 95%-confidence intervals corresponding to the $I(0)$ case.

¹⁰ The exact requirement is that $(1/m) + ((m^{1+2\alpha}(\log m)^2)/(T^{2\alpha})) \rightarrow 0$ as $T \rightarrow \infty$, where α is determined by the smoothness of the spectral density of the short run component.

It is observed that practically all values of d are within the $I(0)$ interval, which is consistent with the results based on the parametric approach above.

As a conclusion, the results presented across Tables 3 and 4 and Figure 3 suggest that the aggregate wine imports series is stationary $I(0)$.

Next, we concentrate on tourism, and the first thing we do is to remove the seasonal component. Here we take two approaches. First, we assume that seasonality is deterministic and use seasonal dummies to remove that component. The plots of the deseasonalised series, though not reported, showed that the series may be nonstationary. As a second approach we assume that the seasonality in tourism has a stochastic nature, and use seasonal first differences on the logged series, creating thus a new series, which is the growth monthly rate. Similarly to the previous case, nonstationarity is found in this series.¹²

Table 5. Unit root tests for the deseasonalised travellers (DT) series (detrended by seasonal dummies)

	<i>No regressors</i>	<i>With an intercept</i>	<i>With a linear trend</i>
ADF	-2.13 (-1.94)	-2.19 (-2.90)	-2.38 (-3.47)
PP	-2.68 (-1.94)	-2.65 (-2.90)	-2.59 (-3.47)
KPSS	---	0.98 (0.46)	0.44 (0.146)

In this series there does not seem to be a trend but the series fluctuates around zero. One would therefore tend to accept the ‘no regressors’ test results, implying the series is $I(1)$, i.e., non-stationary. In parenthesis the critical values at the 5% level.

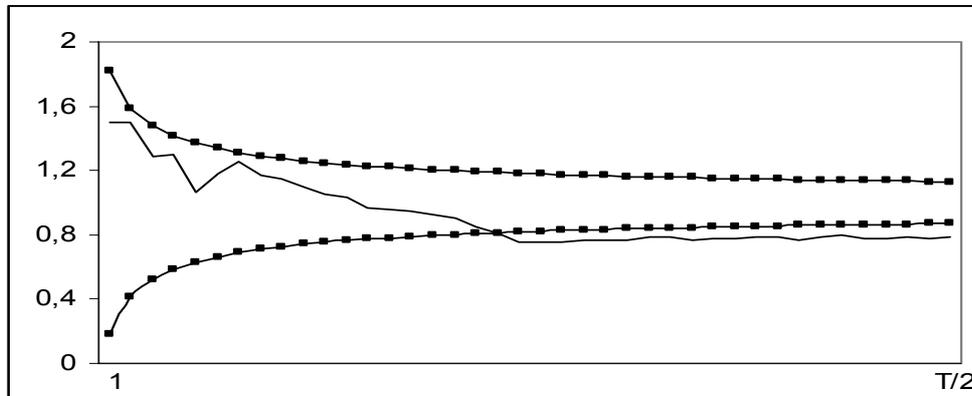
Table 6. 95% confidence intervals of the non-rejection values of d for the DT series

	<i>No regressors</i>	<i>With an intercept</i>	<i>With a linear trend</i>
White noise	[0.62 (0.75) 0.95]	[0.62 (0.73) 0.89]	[0.65 (0.74) 0.89]
Bloomfield ($p = 1$)	[0.40 (0.59) 0.91]	[0.45 (0.76) 1.02]	[0.61 (0.80) 1.03]
Bloomfield ($p = 2$)	[0.30 (0.61) 1.14]	[0.32 (0.98) 1.31]	[0.58 (0.99) 1.39]

The values in parenthesis within the brackets refers to the value of d producing the lowest statistic. It is supposed to be an approximation to the MLE.

¹¹ Some attempts to calculate the optimal bandwidth numbers have been examined in Delgado and Robinson (1996) and Robinson and Henry (1996). However, in the case of the Whittle estimator, the use of optimal values has not yet been theoretically justified. Other authors, such as Lobato and Savin (1998) use an interval of values for m but we have preferred to report the results for the whole range of values of m .

¹² For the two series the periodograms present peaks at the zero frequency, while the periodograms of the first differences have positive and finite values at such a frequency. Therefore, the two series seem to be $I(1)$.



The horizontal axe refers to the bandwidth parameter number, while the vertical one corresponds to the estimated values of d . The dotted line refers to the 95% confidence interval for the $I(1)$ hypothesis.

Figure 4. Estimates of d based on Gaussian semiparametric estimate for DT series

Table 7. Unit root tests for Monthly Growth Rate of the travellers series

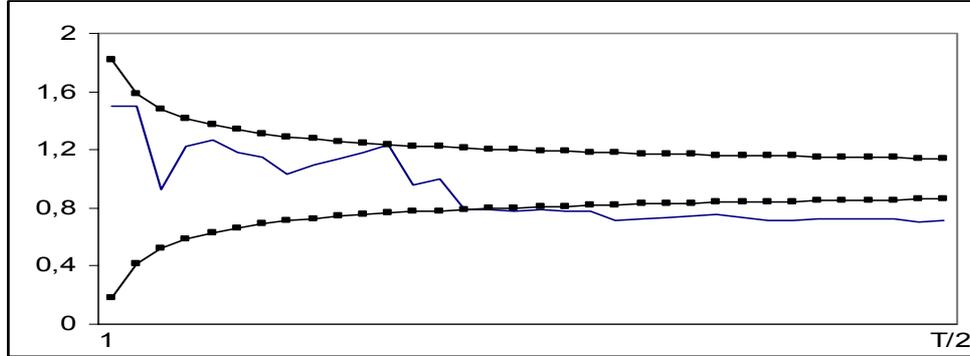
	No regressors	With an intercept	With a linear trend
ADF	-2.50 (-1.94)	-2.43 (-2.90)	-1.64 (-3.47)
PP	-3.02 (-1.94)	-2.91 (-2.90)	-2.44 (-3.47)
KPSS	---	0.99 (0.46)	0.23 (0.146)

In this series there is no trend but the series fluctuates around zero. Therefore, one would use the 'no regressors' test results, implying clearly the series is $I(1)$, i.e. non-stationary. In parenthesis the critical values at the 5% level.

Table 8. Unit root tests for Monthly Growth Rate of the travellers series

	No regressors	With an intercept	With a linear trend
White noise	[0.62 (0.73) 0.90]	[0.56 (0.79) 1.11]	[0.51 (0.77) 1.19]
Bloomfield ($p = 1$)	[0.56 (0.66) 0.80]	[0.57 (0.83) 1.11]	[0.52 (1.06) 1.39]
Bloomfield ($p = 2$)	[0.60 (0.69) 0.82]	[0.66 (0.87) 1.10]	[0.63 (1.07) 1.41]

The values in parenthesis within the brackets refers to the value of d producing the lowest statistic. It is supposed to be an approximation to the MLE.



The horizontal axe refers to the bandwidth parameter number, while the vertical one corresponds to the estimated values of d . The dotted line refers to the 95% confidence interval for the $I(1)$ hypothesis.

Figure 5. Estimates of d for the Monthly Growth Rate of travellers series

Across Tables 5 – 8 and Figures 4 and 5 we display the same type of analysis as the one performed before for the wine import series. The results are very similar in both series: using classic methods (Tables 5 and 7) evidence of a unit root is found in all cases when using the test statistic with most realistic assumptions. Using the fractional framework, the unit root is almost never rejected though fractional orders of integration, with values of d slightly below 1 are also plausible in most of the cases.

To conclude, we can summarize the results presented across this section by saying that the aggregate wine imports seem to be stationary $I(0)$, while tourism, once the seasonal component has been removed, is nonstationary $I(1)$.

4 An empirical application based on a long memory regression model

Denoting Aggregate Wine Imports by AWI_t and Deseasonalised Tourism as DT_t , we employ through the model given by (1) and (2), testing H_0 (3) for given values $d_0 = 0, 0.01, \dots, 2$, assuming that u_t is white noise and Bloomfield (with $p = 1$).¹³ However, in order to examine the dynamic structure underlying the two series, we use as a regressor lagged values of the tourism series.¹⁴ In other words, we test the null model,

$$AWI_t = \alpha + \beta DT_{t-k} + x_t, \quad (6)$$

¹³ Other values of p were also employed and the results were very similar to those reported in the paper with $p = 1$.

¹⁴ We conducted some tests for exogeneity of tourism in the wine imports equation. To establish evidence for non-causality, an unrestricted VAR was used. Weak exogeneity appeared to be satisfied in the dynamic equation because when entering the current value of DT in the equation it proved to be insignificantly different from zero. This finding supports the view that DT is weakly exogenous for the model.

$$(1 - L)^{d_0} x_t = u_t, \quad (7)$$

with k in (6) equal to 1, 2, ..., and 12. As an alternative approach, we could have employed a dynamic lag-structure for DT in (6) in line with the literature on dynamic regressions in standard models. However, that approach would impose the same degree of integration across the lags, while here we permit different values of d for each lag. First, we employ the deseasonalised tourism series based on the seasonal dummies. Table 9 reports the results for the case of white noise disturbances, while Table 10 refers to the Bloomfield model. In both cases, we report, for each k , the estimates for the coefficients (and their corresponding t-ratios), the value of d_0 producing the lowest statistic, its confidence interval (at the 95% level) and the value of the test statistic.

Starting with the case of white noise u_t , (Table 9), we see that β appears significant for $k = 1, 2, 3$ and 4, implying that tourism has an effect on wine imports that lasts at least the following four months. We also see that the interval of non-rejection values of d is relatively wide in all cases, ranging from -0.41 ($k = 8$) to 0.05 ($k = 6$). The case of $d = 0$ is included in all intervals but lowest statistics are obtained for negative d . Note that the estimates of α and β are based on the value of d producing the lowest statistic, which seems to be more appropriate from a statistical viewpoint. Imposing a weak dependence structure on the disturbances throughout the model of Bloomfield (1973), Table 10, the intervals are now wider, the values of d with the lowest statistics being still negative, and the slope coefficient is now significant for the first seven periods, implying a longer dynamic effect of tourism than in the previous case.

Tables 11 and 12 are similar to Tables 9 and 10 but using the monthly growth rates as the deseasonalised series. If u_t is white noise, only the first two lags appear statistically significant, however, using the model of Bloomfield (1973), the significant coefficients reach the lag 9.

We can therefore conclude this section by saying that there is some kind of dynamic behaviour in the effect that German tourism has on German imports of Spanish wines. This significant effect lasts less than a year though varies substantially depending on the model considered and the type of series used for measuring tourism.

Table 9. Estimates of parameters in AWI_t and $TRAV_{t-k}$ relationship, using deseasonalised data

$AWI_t = \alpha + \beta DT_{t-k} + x_t; \quad (1-L)^d x_t = u_t$					
With white noise disturbances (in parenthesis t-ratios)					
k	Alpha	Beta	d-95% confidence interval	d	Stat.
1	16.738 (1152.21)	0.507 (2.951)	[-0.37 0.01]	-0.22	-0.0245
2	16.751 (1126.75)	0.467 (2.713)	[-0.39 0.03]	-0.22	0.0445
3	16.751 (1056.77)	0.358 (1.996)	[-0.39 0.04]	-0.21	-0.0445
4	16.733 (1003.17)	0.352 (1.911)	[-0.37 0.04]	-0.20	0.0505
5	16.736 (992.33)	0.293 (1.592)	[-0.39 0.04]	-0.20	-0.035
6	16.755 (971.13)	0.268 (1.440)	[-0.36 0.05]	-0.20	0.0255
7	16.761 (910.11)	0.217 (1.147)	[-0.36 0.04]	-0.19	0.0053
8	16.770 (1190.02)	0.059 (0.345)	[-0.41 0.04]	-0.25	-0.0157
9	16.781 (1178.78)	-0.112 (-0.035)	[-0.38 -0.03]	-0.24	-0.0065
10	16.779 (1198.88)	-0.070 (-0.445)	[-0.40 -0.03]	-0.23	-0.0367
11	16.784 (1137.70)	-0.159 (-0.963)	[-0.39 0.02]	-0.24	0.0243
12	16.773 (1153.71)	-0.128 (-0.798)	[-0.41 0.01]	-0.22	0.0451

In bold, significant values at the 5% significance level.

Table 10. Estimates of parameters in AWI_t and $TRAV_{t-k}$ relationship, using deseasonalised data

With Bloomfield ($p = 1$) disturbances (in parenthesis, t-ratios)					
k	Alpha	Beta	d-95% confidence interval	d	Stat
1	16.734 (2645.81)	0.486 (4.363)	[-0.62 0.03]	-0.37	0.0065
2	16.713 (3555.17)	0.433 (4.650)	[-0.75 0.02]	-0.46	-0.0004
3	16.712 (3837.67)	0.396 (4.445)	[-0.77 0.02]	-0.44	0.0566
4	16.752 (3640.92)	0.368 (4.016)	[-0.85 0.01]	-0.43	-0.0987
5	16.754 (4838.55)	0.302 (3.911)	[-0.95 0.05]	-0.41	-0.0425
6	16.753 (4411.51)	0.260 (3.219)	[-0.97 0.05]	-0.49	0.0044
7	16.754 (2930.76)	0.219 (2.136)	[-0.94 0.05]	-0.49	0.0140
8	16.766 (2381.66)	0.053 (0.444)	[-0.75 0.08]	-0.47	-0.0447
9	16.787 (2519.90)	-0.105 (-1.111)	[-0.70 0.05]	-0.45	-0.0154
10	16.781 (2565.64)	-0.109 (-0.947)	[-0.75 0.04]	-0.45	-0.0156
11	16.790 (3225.34)	-0.198 (-1.150)	[-0.81 0.10]	-0.49	-0.0655
12	16.777 (3561.03)	-0.113 (-1.345)	[-0.82 0.11]	-0.44	-0.0165

In bold, significant values at the 5% significance level.

Table 11. Estimates of parameters in AWI_t and $TRAV_{t-k}$ relationship using Monthly Growth Rates

$AWI_t = \alpha + \beta DT_{t-k} + x_t; \quad (1-L)^d x_t = u_t$					
With white noise disturbances (in parenthesis, t-ratios)					
k	Alpha	Beta	d-95% confidence interval	d	Stat
1	16.771 (1227.33)	0.234 (2.031)	[-0.45 0.04]	-0.25	-0.0242
2	16.769 (1150.38)	0.223 (1.838)	[-0.44 0.09]	-0.25	-0.0354
3	16.761 (1058.18)	0.045 (0.342)	[-0.43 0.12]	-0.22	0.0254
4	16.764 (939.35)	0.135 (1.382)	[-0.41 0.14]	-0.17	-0.0235
5	16.765 (959.99)	0.131 (1.081)	[-0.44 0.13]	-0.19	-0.0235
6	16.766 (967.77)	0.070 (0.577)	[-0.43 0.12]	-0.20	-0.0153
7	16.769 (957.17)	0.067 (0.617)	[-0.42 0.11]	-0.19	0.0611
8	16.766 (904.16)	0.059 (0.863)	[-0.42 0.15]	-0.19	0.0783
9	16.769 (881.81)	0.035 (0.256)	[-0.44 0.18]	-0.19	0.0145
10	16.763 (940.87)	0.013 (0.045)	[-0.44 0.15]	-0.21	0.0246
11	16.763 (928.85)	-0.051 (-0.467)	[-0.45 0.11]	-0.22	-0.0265
12	16.763 (878.28)	-0.043 (-0.376)	[-0.44 0.15]	-0.21	0.0556

In bold, significant values at the 5% significance level.

Table 12. Estimates of parameters in AWI_t and $TRAV_{t-k}$ relationship, using Monthly Growth Rates detrended data

With Bloomfield ($p = 1$) disturbances (in parenthesis, t-ratios)					
k	Alpha	Beta	d-95% confidence interval	d	Stat
1	16.777 (6113.28)	0.212 (6.134)	[-1.33 0.17]	-0.53	0.0129
2	16.780 (8934.05)	0.255 (11.041)	[-1.31 0.28]	-0.55	-0.0545
3	16.766 (8500.14)	0.055 (2.400)	[-1.46 0.22]	-0.53	-0.0365
4	16.773 (10816.4)	0.106 (6.152)	[-1.47 0.15]	-0.58	-0.0654
5	16.773 (11433.3)	0.077 (4.688)	[-1.66 0.17]	-0.56	0.0655
6	16.772 (11261.3)	0.073 (4.344)	[-1.62 0.21]	-0.67	-0.0276
7	16.770 (9670.66)	0.057 (3.151)	[-1.62 0.21]	-0.66	0.0065
8	16.773 (10341.6)	0.087 (5.137)	[-1.55 0.24]	-0.66	0.0067
9	16.771 (14991.8)	0.090 (8.744)	[-1.64 0.23]	-0.71	-0.0869
10	16.773 (10229.8)	0.021 (1.911)	[-1.63 0.24]	-0.69	-0.0317
11	16.771 (12723.2)	-0.021 (-1.156)	[-1.71 0.19]	-0.67	0.0055
12	16.765 (12960.2)	0.005 (1.056)	[-1.72 0.15]	-0.63	0.0156

In bold, significant values at the 5% significance level.

5 Disaggregation by products

In this section we examine separately the different wine products and perform the same type of analysis as the one employed in Section 4. That is, we consider the same model as in (6) and (7), using specific types of wine rather than the aggregate flow.

Table 13. Slope coefficients in the regression using the DT (dummy variables) series and white noise u_t .

k	A	B	C	D	E	F	G	H
1	0.275	-1.682	-1.264	1.497	0.001	-0.327	1.432	-0.354
2	0.307	-0.821	-0.835	1.458	0.204	0.100	1.174	-0.736
3	0.287	-0.403	-0.311	1.229	0.529	-0.404	1.279	-1.049
4	0.519	-0.887	-1.048	1.073	0.651	-1.050	0.861	-1.520
5	0.608	-0.022	0.024	1.000	1.226	-0.548	1.567	-1.101
6	0.612	0.087	-0.694	1.101	1.614	-0.900	1.227	-1.120
7	0.617	1.234	-1.059	1.248	1.521	-0.829	1.532	-1.445
8	0.279	-0.673	-0.525	0.895	1.249	-0.560	1.332	-0.900
9	0.019	-1.980	-0.897	0.612	1.447	-1.074	1.100	-0.864
10	0.098	-1.560	-0.336	0.799	1.483	-1.335	1.008	-0.875
11	-0.045	0.293	-1.177	1.257	2.177	-1.198	1.256	-0.399
12	0.149	0.890	-0.782	0.482	2.287	-1.206	1.437	-0.273

A: Sparkling wine; B: White from Penedes; C: White from Rioja; D: Wines from Navarra; E: Wines from Penedús; F: Wines from Rioja; G: Wines from Valdepeñas; H: Sherry. In bold, significant coefficients at the 5% significance level.

Table 14. Slope coefficients in the regression using the DT (dummy variables) series and Bloomfield ($\rho = 1$) u_t .

k	A	B	C	D	E	F	G	H
1	0.274	-1.697	-1.112	1.674	-0.291	-0.033	1.226	-0.135
2	0.304	-1.583	-0.379	1.552	0.018	0.867	0.894	-0.701
3	0.308	-1.574	1.111	1.429	0.387	-0.008	1.250	-1.015
4	0.518	-1.470	-1.221	1.361	0.368	-1.074	0.602	-1.522
5	0.604	-1.225	1.953	1.028	1.157	0.344	1.534	-1.062
6	0.612	-1.188	-0.403	1.144	1.605	-0.739	1.119	-1.103
7	0.615	-1.073	-1.212	1.310	1.537	-0.480	1.539	-1.474
8	0.279	-1.326	0.284	0.924	1.455	0.622	1.265	-0.905
9	0.039	-1.398	-0.876	0.544	1.873	-0.547	1.034	-0.869
10	0.121	-1.194	0.690	0.805	2.012	1.206	0.947	-0.885
11	-0.011	-0.803	-1.485	1.263	2.309	-1.547	1.275	-0.404
12	0.196	-0.622	-0.698	0.468	2.353	-1.437	1.464	-0.310

A: Sparkling wine; B: White from Penedes; C: White from Rioja; D: Wines from Navarra; E: Wines from Penedús; F: Wines from Rioja; G: Wines from Valdepeñas; H: Sherry. In bold, significant coefficients at the 5% significance level.

In Tables 13 and 14 we use the DT_t series (with seasonal dummies), for the two cases of white noise (in Table 13) and Bloomfield disturbances (in Table 14).

We observe that the results are very similar in both cases, implying that the short run dependence (i.e., the type of modelling approach) is not very important when describing the behaviour of these two series. In general, we observe that only for two wine types (reds from Navarra and those from Valdepeñas) the coefficients are significant across the whole period of analysis.¹⁵ For sparkling wine and reds from Penedús, the significant coefficients start five periods after, and the effect of tourism lasts three periods for the former and 8 months for the latter wine type. Very similar results were obtained when using the monthly growth rate of tourism as a regressor.

6 Conclusions

In this paper we have examined the relationship between German imports of Spanish wines and German tourism to Spain. For this purpose, we first analysed each of the series separately, and it was found that wine imports was I(0) stationary, while tourism (once the seasonal component was removed) was nonstationary I(1). Due to the different orders of integration observed in the two series, we examine the relationship between the two variables by means of a long memory regression model, using tourism retarded k periods ($k = 1, \dots, 12$) as a weakly exogenous variable.

Table 15. Summary results from estimated regressions: relationship between German tourists to Spain and German imports of Spanish wine

	Product				
	Aggregate wine (AWI)	Sparkling wine (A)	Red wine from Navarra (D)	Red wine from Penedús (E)	Red wine from Valdepeñas (G)
Average sum of effects (euro per one percent increase of tourists)	2.07	1.83	12.02	13.65	13.93
Average lengths of effect (months)	5.5	3	10	8	11
Unit value (euro per litre), 2003	1.34	2.72	1.64	3.11	2.70

Unit values are calculated from Eurostat data. The 2003 unit values for the other analysed products are: white wine from Penedús (B): 2.78 euro per litre; white wine from Rioja (C): 1.89; red wine from Rioja (F): 2.31; Sherry (H): 2.32.

The obtained results are summarised in Table 15. The first row gives the total effect as the sum of the monthly effects in euro per one percent increase of tour-

¹⁵ In fact, they are significant up to lag 15 in case of the wines from Valdepeñas.

ists.¹⁷ On average, total monthly wine imports of Spanish wine into Germany have increased by about EUR 2 per every increase of roughly 5,000 tourists over the analysed period. For individual wine types, the impact has been mixed. While for sparkling wine the positive effect (about EUR 1.8) is lower than for the overall wine category, three wine types, all quality reds (from Navarra, Penedús and Valdepeñas), have experienced import-promoting effects of about EUR 12-14. Taken together, these three wine types accounted for about 7% of total wine imports during the analysed period.

We find that the connection between tourism and trade seems only to hold for red wines and sparkling wine but not for white wine. Moreover, there seems to be a possible connection between wine quality (as expressed by price) and the magnitude and length of the tourism effect. Table 15 lists unit values (import value/import quantity) as a proxy for import prices of the analysed wine types. The two most-expensive red wine types (Penedús and Valdepeñas) also display the strongest import-promoting effects. However, quality reds from Rioja seem to be an exception. Although the average import unit value at EUR 2.3 per litre is higher than the one for quality reds from Navarra (EUR 1.6), no significant relationship with the tourism series have been found. A possible explanation for this exception may be the fact that Rioja reds (accounting for on average 19% of imports during the period of investigation) comprise both some of the best, most expensive and internationally-appreciated Spanish quality wines and lots of lowly-priced bulk wine (mainly produced in the 'Baja' region) (Albisu, 2004). Given their long tradition, Rioja wines may thus be internationally received as the 'typical' Spanish wine, similar to Bordeaux wines in France or Chianti reds in Italy. Hence, Rioja wine exports may reflect both demand by quality-oriented international wine collectors and price-conscious mass retailers, both types of demand probably being little affected by international tourism flows.

The average lengths of the import-promoting effects is about 5.5 months for total wine imports, three months for sparkling wine and 9-10 months for the just mentioned quality reds. This result clearly shows that, at least in the analysed case, tourism has a positive impact on the travel destination economy, which lasts for many months after the tourists have already left the country.

From a methodology point of view, the approach employed in this paper solves at least partially the strong dichotomy produced by the I(0)/I(1) specifications by using fractional values for the orders of integration, and given the different stochastic nature of the two series considered here, the use of a long memory regression model where one of the variables is weakly exogenous allows us to examine the dynamic behaviour of the series in a much more flexible way. The

¹⁷ The numbers are the simple mean from the estimates given in Table 13 and 14. The interpretation of the estimates for the monthly growth rates series (Tables 11 and 12) are not directly comparable to the former ones, therefore they have not been included in the summary calculation of Table 15.

frequency domain formulation of the test statistic used here seems to be very unpopular with many econometricians and, though there exist time domain versions of the tests (Robinson, 1991, Tanaka, 1999), the preference here for the frequency domain approach is motivated by the somewhat greater elegance of formulae it affords, especially when the Bloomfield model is used.

Overall, this analysis has clearly shown that the export- (and thus economy-) promoting effects of international tourism, in some cases at least, are statistically significant, positive, relatively long-lasting and considerable in magnitude. Policy makers and industry as well as tourism development officials are therefore well-advised to consider these interactions in their planning and budget allocation decisions. Apart from the identification work to be done of which manufactured goods and tourist groups display significant connections, industrial and tourism development officer should work hand-in-hand to implement effective foreign marketing programmes and strategies to optimally exploit existing tourism-trade interaction effects.

Further research should focus on the identification of those goods and tourism groups (i.e., countries) which display the strongest links. In addition, more knowledge is needed on the causes of and determinants for these relationships.

Appendix A

The test statistic proposed by Robinson (1994) is based on the Lagrange Multiplier (LM) principle, and is given by:

$$\hat{r} = \frac{T^{1/2}}{\hat{\sigma}^2} \hat{A}^{-1/2} \hat{a},$$

where T is the sample size and

$$\hat{a} = \frac{-2\pi}{T} \sum_{j=1}^{T-1} \psi(\lambda_j) g(\lambda_j; \hat{\tau})^{-1} I(\lambda_j); \quad \hat{\sigma}^2 = \sigma^2(\hat{\tau}) = \frac{2\pi}{T} \sum_{j=1}^{T-1} g(\lambda_j; \hat{\tau})^{-1} I(\lambda_j);$$

$$\hat{A} = \frac{2}{T} \left(\sum_{j=1}^{T-1} \psi(\lambda_j)^2 - \sum_{j=1}^{T-1} \psi(\lambda_j) \hat{\varepsilon}(\lambda_j)' \times \left(\sum_{j=1}^{T-1} \hat{\varepsilon}(\lambda_j) \hat{\varepsilon}(\lambda_j)' \right)^{-1} \times \sum_{j=1}^{T-1} \hat{\varepsilon}(\lambda_j) \psi(\lambda_j) \right)$$

$$\psi(\lambda_j) = \log \left| 2 \sin \frac{\lambda_j}{2} \right|; \quad \hat{\varepsilon}(\lambda_j) = \frac{\partial}{\partial \tau} \log g(\lambda_j; \hat{\tau}); \quad \lambda_j = \frac{2\pi j}{T}; \quad \hat{\tau} = \arg \min_{\tau \in T^*} \sigma^2(\tau),$$

where T^* is a compact subset of the \mathbb{R}^q Euclidean space. $I(\lambda_j)$ is the periodogram of u_t evaluated under the null, and g above is a known function coming from the spectral density of u_t , $f = (\sigma^2/2\pi)g$.

Appendix B

The estimate of Robinson (1995) is implicitly defined by:

$$\hat{d} = \arg \min_d \left(\log \overline{C(d)} - 2d \frac{1}{m} \sum_{j=1}^m \log \lambda_j \right),$$

$$\text{for } d \in (-1/2, 1/2); \quad \overline{C(d)} = \frac{1}{m} \sum_{j=1}^m I(\lambda_j) \lambda_j^{2d}, \quad \lambda_j = \frac{2\pi j}{T}, \quad \frac{m}{T} \rightarrow 0.$$

where m is a bandwidth parameter number. Velasco (1999a, b) has recently shown that the fractionally differencing parameter d can also be consistently semiparametrically estimated in nonstationary contexts by means of tapering. See also Phillips and Shimotsu (2004, 2005) for recent refinements of this procedure.

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