Trade in services in Spain. The effect of economic integration*

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Abstract

Services export is an increasingly important component of nations' export basket and it is growing as a share of the world economy. Therefore, it is important to understand the determinants of trade in services and its influence on economic growth. Moreover, it is crucial for policymakers to understand how economic policies, such as joining to a currency union, would affect this type of trade. The effects of economic integration on international merchandise trade have been extensively investigated, but its impact on trade in services has received low attention. This paper aims to fill this gap by exploring the impact of the EMU on international trade in services. Firstly, we highlight the relevance of trade in services for the Spanish case. To that end, we test the effect of export (in services) on economic growth. Results suggest that, although they are not conclusive, there seems to exist a short-run nexus between trade in services and economic growth. Secondly, we explore the impact of the Euro across a set of developed economies. Results suggest a substantial impact of the euro on intra-Eurozone trade in services. Specifically, for the case of Spain, the individual euro's effect is around 40 per cent, below the average for the rest of the Eurozone.

Keywords: *euro's effect, trade in services, trade-led growth hypothesis gravity model.*
JEL classification: *C500, F140, F150.*

Resumen

Las exportaciones de servicios son un componente cada vez más importante de la cesta exportadora de los países y tienen un peso cada vez mayor en la economía mundial. Por tanto, es importante entender los determinantes del comercio de servicios y su influencia en el crecimiento económico. Además, resulta crucial para los responsables políticos entender de qué manera las políticas económicas, como por ejemplo adoptar una moneda común, afectaría a este tipo de comercio. Los efectos de la integración económica en el comercio internacional de mercancías se han investigado a fondo, sin embargo, el impacto que tendría sobre el comercio de servicios ha recibido poca atención. Este trabajo pretende contribuir a esta literatura a través del análisis del efecto que la adopción del euro ha tenido sobre el comercio internacional de servicios. En primer lugar, resaltamos la relevancia del comercio de servicios para el caso español. Para tal fin, analizamos la influencia de las exportaciones (de servicios) en el crecimiento económico. Los resultados obtenidos sugieren que, aunque no son concluyentes, parece existir una relación a corto plazo entre el comercio de servicios y el crecimiento económico. En segundo lugar, exploramos el

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efecto del euro en un conjunto de economías desarrolladas. Los resultados sugieren un impacto sustancial del euro en el comercio de servicios dentro de la zona del euro. Específicamente, para el caso de España, el efecto del euro en el las exportaciones españolas de servicios es del 40 por 100, aunque está por debajo del promedio para el resto de la zona euro.

Palabras clave: *euro, comercio de servicios, hipótesis de crecimiento impulsado por el comercio, modelo de gravedad*

Clasificación JEL: *C500, F140, F150.*

1. Introduction

Services currently account for approximately two thirds of world GDP, World Development Indictors (WDI, 2017). Advances in technology and international trade agreements induced the increasing tradability of services. For this reason, trade in services is becoming a significant component of international trade. The share of trade in services on world GDP has raised from 8.5 per cent in 1995 to 12.7 per cent in 2016 (WDI, 2017). What is more, this trade in services seems to be underestimated due to the intangible nature of services, the interdependence of services and foreign direct investment flows that makes its measurement difficult and due to restrictions and barriers on trade in services (Karam and Zaki, 2013). For the Eurozone, services represent a 73.7 per cent of the region GDP in 2016 while trade in services represented a 23.7 per cent. Consequently, service sector is the main economic activity of the region, showing a growing relevance on total trade flows.

In spite of the increasing importance of trade in services, this type of flow is under-researched if compared to trade in goods. Kimura and Lee (2006) pointed out that reasons why trade in services has received lower attention than trade in goods are partially related to the lack of internationally comparable data on services. Both trade flows have been separated because the Standard International Trade Classification (SITC) applies only to goods and there has been no readily comparable classification of trade in services. Moreover, there are some concerns about the application of traditional trade models to explain trade in services.

In Spain, services represent a 66.9% of the country's GDP and trade in services (exports plus imports) represented a 15.9% of GDP in 2016 (WDI, 2017). In the present research the Spanish case is used to illustrate the relevance that this type of trade has on promoting economic growth and why it is crucial for policymakers of the region to understand the effects of economic policies, such as the economic integration into a monetary union, on services exports. To that end, we firstly use the case of Spain to explore the role of services exports on economic growth. This analysis would allow us to illustrate the relevance that this type of flows has on countries' economies. To do that, we test the link between exports in services and the Spanish economic growth by using annual data for the period 1975-2016. Secondly, we focus on the effect that the economic integration into the Economic and Monetary Union (EMU) has had on services exports, highlighting its effect in the Spanish economy. To achieve this objective, we explore the impact of the Euro across a set of 37 developed economies for the period 1995-2012.

To sum up, the aim of this research is twofold: (i) we explore the impact of service export on economic growth for the Spanish case and (ii) we analyse the effect of the EMU on intra-Eurozone trade in service flows. As the best of our knowledge, this is the first attempt to analyse the service export-led growth and to estimate the economic effect of the EMU on trade in services. Furthermore, this paper addresses some empirical problems that have arisen in the few existing papers on this issue by using a database with a longer time period, a proper control group and by including both country-year and country-pair fixed effects in the regression.

The rest of the paper is organized as follows. The second section briefly discusses some basic facts about trade in service in Spain and presents the literature review. The third section shows the analysis of the relationship between services exports and economic growth. The fourth section presents and discusses the economic effect of the EMU on the Eurozone countries, paying attention to its effect in Spain. Finally, some conclusions are drawn in the fifth section.

2. Motivation

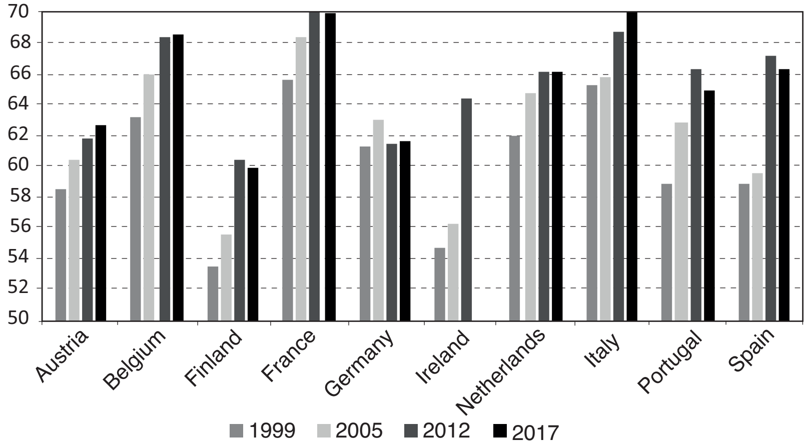
2.1. Trade in services in Spain

In this research, we use Spain as a case study to explore the relationship between services exports and economic growth, as well as to analyse the effect of economic integration on service exports. We aim to illustrate the relevance of the service sector on the country's economy. To do so, we compare the situation in Spain with other countries of the Eurozone. All data presented in this section are compiled from the World Development Indicators database (WDI, 2017) elaborated by the World Bank.

Figure 1 presents the share that service sector represents on the countries' GDP. It can be observed how France and the Netherlands present the largest share of service sector on GDP, being around 70 per cent in 2017. Indeed, service sector in all countries presented an increasing share since 1999, being relatively similar their weights in 2017 of around 60-70 per cent. For the case of Spain, the valued added of the service sector represented a 59 per cent in 1999 and a 66.4 per cent in 2017, which implies a growth of 13 per cent in this 18 years.

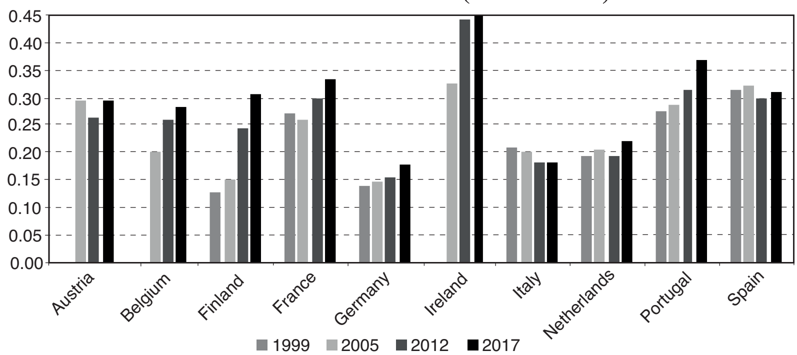
While the share of the service sector on EMU countries is relatively similar across countries and time, there are important differences on the relevance of the services exports on countries' GDP (Figure 2) and on total exports (Figure 3). The largest share of service exports as percentage of GDP is presented in Ireland, while in the rest of the countries this share is less than the half than the Irish one. Precisely, Spain presented one of the lowest share, where services exports represented a 10.6 per cent of the Spanish GDP and this percentage has barely changed since 1999. Regarding the relevance of service exports on total exports (goods and services), again Ireland

FIGURE 1
SERVICES, VALUE ADDED (% OF GDP)



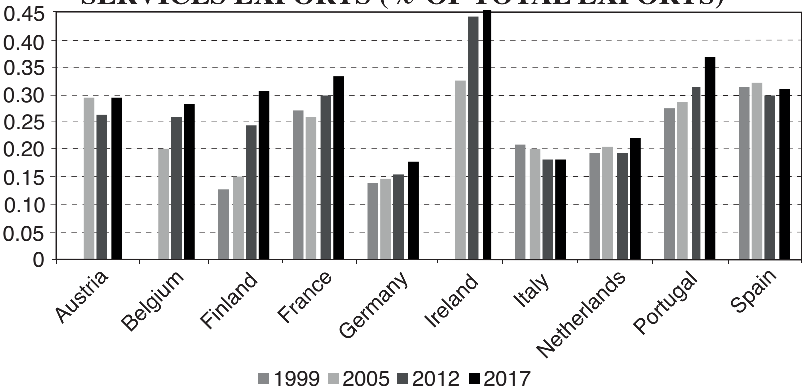
SOURCE: Worlk Bank Database and own elaboration.

FIGURE 2
SERVICES EXPORTS (% OF GDP)



SOURCE: Worlk Bank Database and own elaboration.

FIGURE 3
SERVICES EXPORTS (% OF TOTAL EXPORTS)



SOURCE: Worlk Bank Database and own elaboration.

present the largest share. In the case of Spain, its share has been around 31 per cent during the 18 years presented, although this magnitude is larger than for other EMU countries. Consequently, its seems that service sector in Spain presents a similar relevance on the country's economy as in other EMU countries, but is share on GDP and total exports has barely changed since the adoption of the euro.

2.1. Literature review

Since the early 1960s, there have been a great interest in exploring the possible relationship between exports and economic growth in the sense of accounting for the existence of such a relation, measuring its magnitude, sign, and time horizon (i.e., short or long-run) and the direction of the relationship. In the case of identifying such a positive relationship, it still remains to clarify if a strong output growth is a consequence of an increase in exports or the other way round[1]. From a demand-side perspective, a sustained growth cannot be maintained in domestic markets because of their limited size, but export markets do not involve such type of restrictions and can contribute to output growth through an expansion of aggregate demand.

In its original formulation, the Export-Led Growth Hypothesis (ELGH) predicts that exports also have an indirect effect on economic growth that goes beyond the mere change in export volume, namely, an effect on output through productivity. The productivity effects of exports can differ from country to country due to several factors, such as the level of primary export dependence, the degree of absorptive capacity, and the degree of business and labour regulations, and also can differ over time. The ELGH has been extensively studied in the empirical literature for both developing and developed economies[2]. However, this hypothesis is only tested for the case of merchandise trade or total exports. In the empirical literature, we can find studies that explore the tourism-led growth hypothesis (TLGH). Empirical research supporting the TLGH is extensive (Balaguer and Cantavella-Jordà, 2002; Eugenio-Martín et al., 2004; Santana-Gallego et al., 2010a,b; Castro et al., 2013; Pérez-Rodríguez et al., 2015; Wu and Wu, 2017; among others). Indeed, the evidence that tourism promotes economic growth is clearer than for the ELGH. However, as far as we are concerned, our research is the first attempt to explore the service export-led growth.

An extensive number of empirical studies have been conducted to investigate

[1] AHUMADA and SANGUINETTI (1995) and GILES and WILLIAMS (2000a,b) provide a comprehensive surveys of the empirical literature on this issue. At the theoretical level, it can be cited the seminal work by FEDER (1983) who proposes the so-called "classic" export-led growth model as the basis for examining the theoretical implications of the export-led growth hypothesis (ELGH) and for obtaining empirical evidence on its postulates

[2] For the case of developing countries, see for instance KÓNYA (2004a,b), DREGER and HERZER (2013) or EE (2016) for a group of countries, while for the case of a single country, we can cite, among many others, DHAWAN and BISWAL (1999) for India, MEDINA-SMITH (2000) and GOKMENOGLU et. al. (2015) for Costa Rica, or JIN and JIN (2015) for Korea. Regarding developed economies, MARIN (1992), MANESCHIÖLD (2008) and KONSTANTAKOPOULOU (2016) analyse ELGH for a group of economies, while for a single country analysis we can cite AWOKUSE (2003) for Canada, SHAN and SUN (1998) and MOOSA (1999) for Australia, or SILIVERSTOVS and HERZER (2006) for Chile.

the relationship between exports and economic growth for European countries and the results are not conclusive (Sharma et al., 1991; Thornton, 1997; Ramos, 2001; Konya, 2006; Pistoresi and Rinaldi, 2012; Konstantakopoulou, 2016). In general, their results vary, depending on the selected sub-period of their sample[3]. Balaguer and Cantavella-Jordá (2001, 2004a,b) and Balaguer et al. (2015) explore the ELGH for Spain. In general terms, their results suggest that for the trade liberalisation period, from 1960 onwards, there is significant empirical evidence on some feedback effects between economic growth and primary export activities, such as food, and a growing role of manufactured and semi-manufactured exports. So that, the structural transformation in export composition has also become a key factor for the economic development in this country.

Since the inception of the euro, substantial effort has been put into estimating the impact of the euro on international trade and its role in macroeconomic performance. Several papers have estimated an early effect of the euro on international merchandise trade that ranged from 5 per cent to 26 per cent (Micco et al., 2003; Faruquee, 2004; Flam and Nordstrom, 2006; Aristotelous, 2006; Baldwin, 2006; Bun and Klaassen, 2007). More recent papers, using a large time span since the common currency was adopted, estimate a larger effect of the euro on trade that ranged from 18 per cent to 92 per cent (Camarero et al., 2013; Sadeh, 2014; Glick and Rose, 2016). However, it is important to note that these different estimates of the euro's effects depend on the sample size, the countries considered in the analysis, the estimation techniques and the dependent variable used.

There are some important characteristics of services that clearly distinguish international trade in services from trade in goods. For instance, production and consumption of a service must appear simultaneously, and services have an intangible nature (Kimura and Lee, 2006). Sharing a common currency implies the elimination of exchange rate volatility and transaction costs. Moreover, the introduction of euro coins and notes in 2002 eliminated currency conversion between countries belonging to the Economic and Monetary Union (EMU). These issues become relevant to explain trade in services, however the effect of the euro on international trade in services has not yet been explored.

The only antecedents that exist in the empirical literature are the few papers that explore the effect of the EMU on international tourism movements. Gil-Pareja et al. (2007) estimate an effect of the euro on intra-Eurozone tourism flows of 6.5 per cent. De Vita (2014) considering the case of the euro up to 2011, obtain the effect of common currency on tourism is around 30 per cent. Santana-Gallego et al. (2016) obtain a substantial impact of the euro on intra-Eurozone tourism of between 44 and 126 per cent when proper estimation method, control group and definition of the Eurozone are used. Finally, Saayman et al. (2016) estimate that sharing the common currency increases tourism arrivals by 23 per cent and 88 per cent. All these papers use gravity model to explain bilateral tourism movements. While the effect of the

[3] The sub-period in which they observed a weak support of ELG is the post-WWII period.

EMU depends on the specification chosen, the positive effect of the Euro on tourism movements is undoubtedly strong and consistent.

3. Empirical evidence on the relation between exports and growth for Spain

3.1. Data and methodology

In this section, we aim to provide some empirical evidence on the relationship between exports and economic growth, paying special attention to the role of service exports. To that end, we use an up-to-date time span and recent econometric techniques for nonstationary variables, mainly in the framework of a single cointegrating regression model. Our final empirical analysis, based on a pure distributed lag (DL) model for the variables in first differences, tries to capture the possible positive short-run (but negative long-run) effects. This phenomenon comes from the situation where an increase in exports induces an expansion of sectors that do not exhibit positive externalities, and the associated productivity loss will offset the specialization gains in the long-run.

The theoretical model mainly follows the work by Dreger and Herzer (2013), where to capture the impact of exports on output through the productivity channel, the starting point is an AK-type (or simple neoclassical) production function of the form:

$$Y_t = A_t \cdot K_t^{\alpha} \cdot L_t^{\beta} \qquad [1]$$

where Y is the aggregate production (output) of the economy, K is the capital stock, L is the stock of labour, and A is the total factor productivity function. Given that the hypothesis to examine is how exports can affect economic growth via changes in productivity, it is assumed that $A_t = X_t^{\rho}$, being X total exports. Hence, combining these equations and taking natural logs yields the basic equation for output:

$$\log(Y_t) = \rho \log(X_t) + \alpha \log(K_t) + \beta \log(L_t) \qquad [2]$$

However, the estimate of ρ from this equation cannot be used to measure the productivity effect of exports on output, given that exports are part of output via the national account identity, inducing an almost inevitable empirical evidence of a positive and significant relationship between these variables. Therefore, to separate the impact of exports on output, it is proposed to replace the logarithm of total output, $\log(Y_t)$, with the logarithm of output net of exports (non-export output), $\log(N_t)$, where $N_t = Y_t - X_t$. Assuming a multiplicative relationship of the form $Y_t = X_t^{\gamma} N_t^{1-\gamma}$, where γ is the share of exports to GDP, we obtain the following specification:

$$\log(N_t) = \gamma_1 \log(K_t) + \gamma_2 \log(L_t) + \gamma_3 \log(X_t) + u_t \qquad [3]$$

where $\gamma_1 = \alpha/(1 - \lambda)$, $\gamma_2 = \beta/(1 - \lambda)$ and $\gamma_3 = (\rho - \lambda)/(1 - \lambda)$, which is zero if the coefficient of the export variable in the basic production function specification (ϱ) coincides with the share of exports on output (λ). If $\lambda_3 > 0$, the growth effect of exports exceeds the increase in export volume, suggesting that exports increase output through an increase in productivity. On the other hand, if $\lambda_3 < 0$, exports contribute less to output growth than the increase in export volume, suggesting that exports are productivity-reducing.

For the labour input (L), we can find different measures as, e.g., total number of employed people, active population, or hours worked by engaged persons. Given the very different time pattern of each of these series[4], we choose to proxy the labour input with a linear trend function as $\log(L_t) = \varkappa_0 + \varkappa_1 t$.

Our sample period covers from 1975 to 2016 at the annual frequency ($T = 42$ observations), on which the dependent variable is the logarithm of real output net of exports (N). This variable is generated as the difference between the GDP and total exports, both expressed in current US$.

Regarding the variable exports of goods and services (as a measure of aggregate exports, X) it represents an average of 22.1 per cent over the GDP (this percentage goes from 12.4 in 1975 to 32.95 per cent in 2016). Moreover, instead of only using this global measure of the exporting sector, we also consider the possibility that the productivity function depends on some of its components (as in Siliverstovs and Herzer, 2006), and particularly when $A_t = X_t^{\rho_1} \cdot S_2^{\rho_2} \cdot G_t^{\rho_3}$, where S measures service exports and G is the exports of goods both expressed in current US$.

All the variables used in this analysis, GDP, total exports (X), service exports (S), goods exports (G), and gross capital formation (K) for Spain, are deflated to be measured in real terms using the US GDP deflator (2010 as the reference year). All data are obtained from WDI (2017) databased[5]. The final specification of our empirical model is the following:

$$\log(N_t) = \varkappa_0 + \varkappa_1 t + \gamma_1 \log(K_t) + \gamma_2 \log(X_t) + \gamma_3 \log(S_t) + \gamma_4 \log(G_t) + u_t \qquad [4]$$

where u_t is the usual error term, $\gamma_2 = (\rho_1 - \lambda)/(1 - \lambda)$, $\gamma_3 = \rho_2/(1 - \lambda)$ and $\gamma_4 = \rho_3/(1 - \lambda)$, and the interpretation of the magnitude and sign of γ_2 is the same as for γ_3 in the above simplest model. As robustness check, equation [4] is also estimated using real output (Y) as the dependent variable. For the empirical analysis, we consider three cases depending on the structure of the deterministic component: (1) no deterministic ($\varkappa_0 = \varkappa_1 = 0$), (2) only constant term ($\varkappa_0 \neq 0, \varkappa_1 = 0$), and (3) constant term and linear trend ($\varkappa_0, \varkappa_1 \neq 0$), and also three specifications depending on the structure of the regressors: (1) full set of regressors (Model 1), (2) $\gamma_3 = \gamma_4 = 0$ (Model 2), and (3) exclusion of exports, $\gamma_2 = 0$ (Model 3).

4 See Figure A.1 in the Appendix for the series of average annual hours worked and employed people in Spain

5 Figure A.2 in the Appendix presents the time pattern of the series under analysis.

3.2. Empirical results

Given the apparent nonstationary behaviour of all the series used (Figure 2.A), the first step in the empirical analysis is the computation of a set of tests for the null hypothesis of a unit root against the stationary alternative. Among the wide set of possible tests, we use the efficient ones recently proposed by Perron and Qu (2007)[6], with good size and power properties in finite samples, and also the nonparametric test proposed by Breitung (2002), that does not require any choice of tuning parameters to account for the usual short-run dynamics of the error term driving the stochastic trend component, and it is very simple to construct. The results are shown in Tables 1 and 2.

Particularly, for the specification of the deterministic component including both a constant term and a linear trend component, all the results clearly agree with the nonstationary behaviour of the stochastic component underlying the generating mechanism of all these series. The same conclusion also holds from the results of Breitung's test in the case of including only a constant term. The empirical evidence on the nonstationary nature of all these variables implies that the analysis of our regression model requires the use of cointegration techniques.

Table 3 shows the results of the statistics proposed by Phillips and Ouliaris (1990) for testing the null hypothesis of no cointegration (spurious regression), against the stationary alternative (stable long-run linear relationship). In all the cases considered, for each specification of the deterministic component and number of regressors included, the results strongly support the absence of a stable long-run relationship between the GDP or non-export GDP (as measures of economic growth) and the regressors considered.

To complement these results, we also provide the estimated values of two tests for the null of cointegration against no cointegration. These are the tests proposed by Shin (1994) and Hansen's (1992) test for structural stability under cointegration.[7] In general terms, their results also supports the same conclusion. However, the results of the Box-Pierce statistic for testing the lack of serial correlation in the cointegrating error terms strongly contradicts this evidence (e.g. Phillips, 1986, Theorem 1(h)). Under no cointegration, this statistic will diverge at a rate given by the sample size times the number of sample autocorrelations used in its construction, and in our case we have very low estimated values, which can't be compatible with the nonstationarity of the regression error term under no cointegration, which shows very strong positive autocorrelation. The answer to this contradiction comes from

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[6] Neither of these statistics are completely new, but are modifications of the tests statistics proposed by ELLIOTT et al. (1996), PERRON and NG (1996) and NG and PERRON (2001). The novelty in its use proposed by Perron and Qu (2001) is the use of GLS residuals from demeaning and/or detrending to determine the number of lags required to compute the estimate of the long-run variance of the error term, which can improve their finite sample properties.

[7] In some sense, Hansen's test can be interpreted as a joint test of parameter constancy (structural stability) and cointegration, given that it displays nontrivial power against a structural change under cointegration and also against no cointegration, irrespective on the existence of time-varying parameters or not.

TABLE 1

EFFICIENT UNIT ROOT TESTS BASED ON GLS DETRENDING (PERRON AND QU, 2007)

Series in real terms (2010 US$)	GDP(Y)				Non-export GDP(N)			
	Demeaning		Detrending		Demeaning		Detrending	
	OLS	GLS	OLS	GLS	OLS	GLS	OLS	GLS
ADF-GLS	–0.992	–0.992	3.118	3.118	–1.390	–1.390	2.889	2.889
MZα-GLS	–2.014	–2.014	1.282	1.282	–3.766	–3.766	1.283	1.283
MSB-GLS	0.428	0.428	4.388	4.388	0.344	0.344	3.885	3.885
MZτ-GLS	–0.863	–0.863	5.623	5.623	–1.296	–1.296	4.983	4.983
MP-GLS	1.831	1.831	732.01	732.01	0.760	0.760	577.57	577.57
Series	**Exports of goods and services(X)**				**Gross capital formation(K)**			
	Demeaning		**Detrending**		**Demeaning**		**Detrending**	
	OLS	**GLS**	**OLS**	**GLS**	**OLS**	**GLS**	**OLS**	**GLS**
ADF-GLS	–0.022	–0.022	4.216	4.381	–1.476	–0.659	2.552	2.552
MZα-GLS	0.449	0.449	1.331	1.327	–4.323	–1.407	1.283	1.283
MSB-GLS	0.815	0.815	7.518	6.321	0.329	0.545	3.032	3.032
MZτ-GLS	0.367	0.367	10.005	8.389	–1.426	–0.767	3.889	3.889
MP-GLS	11.201	11.201	2201.82	1556.27	0.423	1.155	357.38	357.38
Series	**Service exports (S)**				**Goods exports (G)**			
	Demeaning		**Detrending**		**Demeaning**		**Detrending**	
	OLS	**GLS**	**OLS**	**GLS**	**OLS**	**GLS**	**OLS**	**GLS**
ADF-GLS	–0.006	–0.006	4.364	4.364	–0.087	–0.211	4.306	4.286
MZα-GLS	0.411	0.411	1.312	1.312	0.424	–0.085	1.365	1.360
MSB-GLS	0.784	0.784	5.843	5.843	0.811	0.628	7.147	5.813
MZτ-GLS	0.323	0.323	7.666	7.666	0.344	–0.053	9.757	7.907
MP-GLS	10.558	10.558	1317.12	1317.12	10.925	6.543	2042.13	1350.89

NOTE: The column labelled as OLS both in the case of demeaning and detrending indicates the result of the test statistics using the estimated lag truncation based on the MAIC selection procedure as described in PERRON and QU (2007), but with the auxiliary regression computed on the OLS residuals. A rejection of the null hypothesis of a unit root against the stationary alternative is registered for smaller values than the critical values (NG and PERRON, 2001, Table I, p. 1524).

the fact called sub-cointegration, where the integrated regressors are cointegrated among them[8].

This situation implies that the results from the cointegration and non-cointegration tests are no longer valid, and there is no known way to proceed further in the

[8] This case is the nonstationary analog to multicollinearity for a linear stationary regression. With regressors highly correlated among them, there is a loss of efficiency in the parameter estimation, and also could affect the outcome of many standard testing procedures, both in terms of size and power.

TABLE 2

UNIT ROOT TESTS BASED ON THE NONPARAMETRIC VARIANCE RATIO TEST OF BREITUNG

	Demeaned	**Detrended**
GDP	9.1426	0.01449
Non-export GDP	5.3049	0.02306
Exports of Goods and services	47.3446	0.00529
Gross capital formation	10.6331	0.09309
Service exports	34.079	0.0041
Goods exports	63.3122	0.0091

framework of a single equation model with the variables in levels. This argument is empirically supported with the results of the nonparametric cointegration rank test proposed by Breitung (2002) (Table 4), which has some advantages in terms of conservative size and good power over the Johansen's procedure and, again, it is simpler to construct.

The final step in our empirical analysis is based on a single equation dynamic model, particularly the ADL$(1, q)$ (autoregressive distributed lag) model, $q \geq 1$, given by

$$y_t = \alpha + \phi y_{t-1} + \sum_{j=0}^{q} \beta'_{k,j} x_{k,t-j} + e_{t,q}, \quad t = q+1, ..., T \qquad [5]$$

where y_t is the log of Y_t or N_t and $x_{k,t} = (k_t, x_t, s_t, g_t)'$, with the notation of the variables in lower case indicating the log transformation. Given the nonstationarity of all these series, with a slight manipulation we obtain the following equivalent error-correction model (ECM) representation for the variables in first differences:

$$\Delta y_t = \alpha + (\phi - 1)\left\{ y_{t-1} - \left(\sum_{j=0}^{q} \beta'_{k,j} \right) x_{k,t-j} \right\} + \sum_{j=0}^{q} \gamma'_{k,j} \Delta x_{k,t-j} + e_{t,q} \qquad [6]$$

where Δ is the differencing operator, $\gamma_{k,0} = \beta_{k,0}, \gamma_{k,0} = \beta_{k,0}$ and $\gamma_{k,i} = -\sum_{j=i+1}^{q} \beta_{k,j}$, $i = 1, \ldots, q-1$. Note that this formulation closely resembles the single-equation conditional error-correction model resulting from a vector autoregression of order q, to separate between long and short-run dynamics under possible cointegration (e.g. the works by Banerjee, 1998; Pesaran and Shin, 1999, and Pesaran et al., 2000, 2001). The term between brackets in [6] is the error-correction term that drives the long-run dynamics under the assumption that there exists at most one conditional level relationship between y_t and $x_{k,t}$, and corrects from eventual departures of the equilibrium. In this framework, only the testing procedures presented in Pesaran et al. (2001) to account for the existence of a single level relationship, remain valid in

TABLE 3

TESTS OF COINTEGRATION

3.1. The dependent variable is real GDP(Y)									
	No deterministic Component			Constant term			Constant term and linear Trend		
Tests	**Model 1**	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**
DFa	–12.335	–2.903	–3.595	–7.079	–6.536	–6.526	–6.879	–6.315	–6.254
DFt	–2.689	–1.218	–1.364	–1.964	–1.881	–1.877	–1.925	–1.837	–1.829
Q(12)	5.109	18.522	19.032	11.498	14.046	13.991	10.158	12.971	14.167
Za	–10.881	–4.903	–5.463	–7.736	–7.532	–7.618	–7.724	–7.528	–7.543
Zt	–2.555	–1.576	–1.671	–2.045	–2.008	–2.016	–2.031	–1.995	–1.997
ADFt(1)	–3.754	–2.228	–2.447	–3.048	–3.028	–2.926	–2.919	–2.862	–2.849
ADFt(2)	–3.624	–2.061	–2.568	–2.882	–2.841	–2.686	–2.717	–2.619	–2.58
ADFt(3)	-4.306^{2}	–2.348	–2.852	–3.009	–2.906	–2.666	–3.096	–2.863	–2.669
ADFt(4)	-4.016^{1}	–2.142	–2.545	–3.114	–3.126	–2.978	–3.104	–3.019	–2.952
ADFt(5)	-3.955^{1}	–1.751	–2.128	–3.13	–3.204	–3.013	–3.107	–3.049	–2.981
S(0)	0.316	1.931^{3}	1.522^{3}	0.283^{3}	0.285^{2}	0.297^{3}	0.293^{3}	0.302^{3}	0.306^{3}
$S_{BK}(q)$	0.107	0.325	0.292	0.084	0.084	0.085	0.084^{2}	0.085^{1}	0.086^{2}
$S_{AR}(p)$	0.103	0.299	0.334	0.088	0.094	0.095	0.090^{2}	0.095^{1}	0.098^{2}
L(0)	1.316	2.028	1.704	2.166	1.285^{3}	1.381	2.809^{3}	2.082^{3}	2.117^{3}
$L_{BK}(q)$	0.447	0.342	0.327	0.644	0.379^{1}	0.396	0.807	0.586^{1}	0.592
$L_{AR}(p)$	0.429	0.314	0.374	0.673	0.423^{1}	0.443	0.863	0.657^{2}	0.675

NOTES: (a) Superscripts 1, 2 and 3 indicate rejection at 10, 5 and 1 per cent level, respectively. (b) Q(m) denotes the Box-Pierce statistic for testing lack of serial correlation in the cointegrating regression term. Under cointegration, its asymptotic null distribution is $\chi^2(m)$. (c) S(a) and L(a) denote the tests by Shin (1994) and Hansen (1992), respectively, for the null of cointegration, where a = 0 indicates that the scale factor is the residual OLS estimate of the short–run variance, a = q (bandwidth) the estimate of the long–run variance (LRV) based on the Bartlett kernel (BK), and a = p (lag) the estimate of the LRV based on the autoregressive (AR) spectral density estimator at frequency zero.

TABLE 3 ***(continuation)***

TESTS OF COINTEGRATION

3.2. The dependent variable is real non-export GDP(N)									
	No deterministic Component			Constant term			Constant term and linear Trend		
Tests	**Model 1**	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**	**Model 1**	**Model 2**	**Model 3**
DFa	–11.722	–3.249	–3.957	–7.305	–7.204	–7.332	–7.035	–7.033	–6.819
DFt	–2.605	–1.298	–1.439	–1.999	–1.983	–2.004	–1.955	–1.955	–1.922
Q(12)	5.139	14.522	14.545	12.765	12.985	12.638	11.075	11.207	12.337
Za	–10.422	–4.979	–5.536	–7.878	–7.753	–7.883	–7.891	–7.899	–7.813
Zt	–2.481	–1.597	–1.691	–2.069	–2.051	–2.071	–2.061	–2.062	–2.047
ADFt(1)	–3.742	–2.271	–2.519	–3.11	–3.193	–3.117	–2.94	–2.932	–2.917
ADFt(2)	–3.599	–2.088	–2.583	–2.872	–2.989	–2.882	–2.639	–2.625	–2.602
ADFt(3)	-4.177^{2}	–2.367	–2.884	–2.905	–3.075	–2.922	–3.028	–2.985	–2.921
ADFt(4)	-3.895^{1}	–2.163	–2.581	–3.093	–3.221	–3.1	–3.051	–3.025	–3.008
ADFt(5)	-3.880^{1}	–1.864	–2.285	–3.139	–3.315	–3.146	–3.105	–3.069	–3.063
S(0)	0.336	1.871^{3}	1.475^{3}	0.271^{3}	0.266^{3}	0.271^{3}	0.285^{3}	0.286^{3}	0.287^{3}
$S_{BK}(q)$	0.111	0.319	0.283	0.082	0.082	0.082	0.082^{2}	0.082^{1}	0.082^{1}
$S_{AR}(p)$	0.103	0.277	0.299	0.085	0.086	0.085	0.086^{2}	0.086^{1}	0.087^{1}
L(0)	1.223	1.856	1.572	1.783^{3}	0.864^{3}	0.912^{2}	2.628^{3}	1.988^{3}	2.202^{3}
$L_{BK}(q)$	0.403	0.316	0.302	0.539	0.266	0.276	0.752	0.567^{1}	0.626
$L_{AR}(p)$	0.375	0.274	0.318	0.561	0.28	0.287	0.793	0.595^{1}	0.669

NOTES: (a) Superscripts 1, 2 and 3 indicate rejection at 10, 5 and 1 per cent level, respectively. (b) Q(m) denotes the Box-Pierce statistic for testing lack of serial correlation in the cointegrating regression term. Under cointegration, its asymptotic null distribution is $\chi^2(m)$. (c) S(a) and L(a) denote the tests by Shin (1994) and Hansen (1992). respectively, for the null of cointegration, where a = 0 indicates that the scale factor is the residual OLS estimate of the short-run variance, a = q (bandwidth) the estimate of the long-run variance (LRV) based on the Bartlett kernel (BK), and a = p (lag) the estimate of the LRV based on the autoregressive (AR) spectral density estimator at frequency zero.

TABLE 4
TESTING THE COINTEGRATION RANK BY THE NONPARAMETRIC TEST OF BREITUNG

	4.1. Output measured by real GDP		4.2. Output measured by real non-export GDP	
	Mean adjusted	**Trend adjusted**	**Mean adjusted**	**Trend adjusted**
$q_0 = n - r_0$ 1	10.023	70.871	10.016	74.126
2	97.327	220.608	103.676	216.617
3	265.026	483.897	260.711	483.876
4	531.144	804.316	533.489	790.862
5	852.725	1873.672	842.878	1856.159

the case where the process $x_{k,t}$ is mutually cointegrated of unknown order. Instead of using this approach in this paper, we simply proceed with the analysis of the dynamic model in [6] incorporating the nonstationary nature of the dependent variable to account for the structure of the short-run relationship with the regressors. Under $\phi = 1$, the model becomes a pure $DL(q)$, $q \geq 1$ for the first difference of the variables, such as

$$\Delta y_t = \alpha + \sum_{i=1}^{q-1} \gamma'_{k,j} \Delta x_{k,t-i} + e_{t,q} \qquad [7]$$

where the vector coefficients $\gamma'_{k,j}$ can be interpreted in terms of elasticities as $\Delta y_t/\partial x'_{k,j} = \gamma_{k,0}$, $\Delta y_t/\partial x'_{k,t-j} = \gamma_{k,j} - \gamma_{k,j-1}$, for $j = 1, \ldots, q-1$, and $\Delta y_t/\partial x'_{k,t-q} = -\gamma_{k,q-1}$, while the cumulated relative effect is given by $\sum_{j=0}^{m} \Delta y_t/\partial x'_{k,t-j} = \gamma_{k,m}$, $m = 0, 1, \ldots, q$. Given the stationarity of the error term, we can use the usual information criteria $AIC_T(q) = \log(\hat{\sigma}^2_{e,T}(q)) + 2\mathrm{N/T}$, and $SBIC_T(q) = \log(\hat{\sigma}^2_{e,T}(q)) + N \log(T)/T$, with $\hat{\sigma}^2_{e,T}(q)) = T^{-1} \sum_{t=q+1}^{T} \hat{e}^2_{t,q}$, and N the total number of estimated coefficients, to select the number of lags $q \geq 1$ to be included in the estimation. Given the relation *$AIC_T(q) < SBIC_T(q)$ between these two selection criteria for sample sizes $T > 7$, it is usual the case where the selected lag is smaller by the SBIC criterion than by the AIC criterion.*

Table 5 contains the results of the estimation of this dynamic model for the same three sets of regressors as before (Models 1-3), for the non-export GDP as the dependent variable and lags selected by the two information criteria. However, with the variables in first differences, Model 1 with the complete set of regressors cannot be estimated due to the almost exact linear relationship between two of these variables (Figure 3). Particularly, the correlation between the first difference of x and g is 0.99, while that between x and s is 0.94, and it is 0.89 between s and g. To avoid the distortions caused by the almost singularity of the sample moment matrix of the regressors in the OLS estimation of Model 1, we omit these results and only present the estimates of Model 2 and 3. For Model 2, we find a contemporaneous positive effect of total exports and capital on output, but no significant for exports, when the lag is selected by SBIC criterion. When selected by AIC criterion, we find negative contemporaneous and lagged effects of exports, but no significant except for lag $q = 4$, and positive and significant effects for lags 0 and 4 for capital. For Model 3, with total exports excluded from the regression, capital and service exports contribute with a positive and significant contemporaneous effect to promote economic growth in Spain (both for the results obtained when lags are selected by AIC and SBIC), while good exports does not have significant contemporaneous effect on growth in the model estimated without lags when selected by SBIC. All the short and medium-term significant effects of service exports on growth are positive up to lag 10, indicating a strong persistent effect of this variable, while that the dynamic significant effects of good exports on growth are mainly negative and this effect is also strongly persistent.

This wide variety of inconclusive results, with a no clear pattern of the effects, is mainly due to the high degree of cross-correlation between the regressors (multicollinearity).

TABLE 5

DISTRIBUTED LAG (DL) MODEL OF ORDER $q \geq 0$ FOR THE VARIABLES IN FIRST DIFFERENCES

	Model 2				Model 3			
	AIC(q)		BIC(q)		AIC(q)		BIC(q)	
	Est.	T	Est.	T	Est.	T	Est.	T
Const.	0.034	2.964	–0.001	–0.063	0.036	3.614	–0.003	–0.323
Δx_t	–0.091	–0.651	0.177	1.403				
Δx_{t-1}	0.217	1.624						
Δx_{t-2}	–0.139	–0.987						
Δx_{t-3}	–0.078	–0.527						
Δx_{t-4}	–0.571	–4.349						
Δk_t	0.824	9.206	0.701	8.834	0.760	10.462	0.652	7.901
Δk_{t-1}	–0.135	–1.322			–0.283	–3.013		
Δk_{t-2}	0.088	0.841			0.397	4.629		
Δk_{t-3}	–0.021	–0.186			–0.139	–1.568		
Δk_{t-4}	0.372	3.714			0.501	3.932		
Δk_{t-5}					–0.457	–3.748		
Δk_{t-6}					–0.063	–0.457		

NOTE: The column labelled as Est. shows the point coefficient estimate, and T indicates the T-ratio statistic for testing the null hypothesis of null significance.

TABLE 5 ***(continuation)***

DISTRIBUTED LAG (DL) MODEL OF ORDER $q \geq 0$ FOR THE VARIABLES IN FIRST DIFFERENCES

	Model 2				Model 3			
	AIC(q)		BIC(q)		AIC(q)		BIC(q)	
	Est.	T	Est.	T	Est.	T	Est.	T
Δk_{t-7}					–0.113	–0.843		
Δk_{t-8}					–0.077	–0.629		
Δk_{t-9}					0.221	1.786		
Δk_{t-10}					0.095	1.136		
Δs_{t}					0.132	1.747	0.253	1.843
Δs_{t-1}					0.302	3.769		
Δs_{t-2}					–0.133	–1.505		
Δs_{t-3}					0.158	1.512		
Δs_{t-4}					–0.001	–0.009		
Δs_{t-5}					0.321	2.689		
Δs_{t-6}					0.297	2.261		
Δs_{t-7}					0.035	0.294		
Δs_{t-8}					0.553	4.399		
Δs_{t-9}					0.036	0.368		
Δs_{t-10}					0.160	2.010		
Δg_{t}					–0.138	–1.874	0.005	0.042
Δg_{t-1}					0.308	4.052		
Δg_{t-2}					–0.219	–2.925		
Δg_{t-3}					0.037	0.424		
Δg_{t-4}					–0.919	–9.542		
Δg_{t-5}					0.143	1.463		
Δg_{t-6}					–0.109	–0.965		
Δg_{t-7}					0.211	1.798		
Δg_{t-8}					–0.336	–2.709		
Δg_{t-9}					–0.265	–2.036		
Δg_{t-10}					–0.796	–7.232		
R2	0.929		0.867		0.987		0.874	

NOTE: The column labelled as Est. shows the point coefficient estimate, and T indicates the T-ratio statistic for testing the null hypothesis of null significance.

On the basis of these results, including the nonstationary characterization of the variables analyzed and the existence of multiple cointegration relations, to obtain more conclusive and clarifying results on the characteristics of the relation between economic growth and export components for Spain, with particular attention to the contribution of service exports on growth, we leave for a future work the use of the general approach proposed by Pesaran et al. (2001).

4. Effect of economic integration on trade in services

4.1. Data and Methodology

The gravity model has been the workhorse for empirical analyses of the euro effect on merchandise trade flows[9]. However, until the 2000s the existing literature on the application of the gravity model to services trade is quite limited. Conversely, there is an increasing literature that use gravity models to explain tourism flows. Under the assumption of tourism as a particular type of trade in services, a gravity equation can be used to study the main determinants of tourism volume (see for instance Durbarry, 2008; Eilat and Einav, 2004; Khadaroo and Seetanah, 2008; Neumayer, 2010; Culiuc, 2014, or Morley et al., 2014).

The empirical analysis presented in this section follows the methodology proposed by Santana-Gallego et al. (2016). In their case the define a gravity model for tourism flows while we adapt their specification to explain trade in services. To that respect, Kimura and Lee (2006) show that trade in services is better predicted by gravity equations than trade in goods; Walsh (2008) obtains that gravity model fits service trade flows in a similar manner to trade in goods. For that reason, a gravity equation is an adequate model to evaluate the effect of the euro on trade in services. Our preferred specification is as follows:

$$\text{Ln } S_{ijt} = \beta_0 + \beta_1 EUijt + \alpha'^{E_{ijt}} + \lambda_{it} + \lambda_{jt} + \lambda_{ij} + u_{ijt} \qquad [8]$$

where Ln denotes natural logs, *i* and *j* indicate exporting and importing countries, respectively, *t* is time. The dependent variable is bilateral service exports (S_{ijt}) from the exporting country *i* to the importing country *j* in year *t*. The source of service exports is the International Trade in Services Statistics from the OECD (2017). This dataset covers trade in services from 1995 to 2012 for detailed services categories, including travel. Secondly, EU_{ijt} is a binary variable which is unity if *i* and *j* are both members of the European Union in year *t,* and it controls for the different

[9] ROSE (2009) surveys 26 studies and, taking together all these estimates, observes that EMU has increased trade by about 8 to 23 per cent in its first years of existence.

enlargement episodes of the European Union[10]. Regarding the variables of interest, E_{ijt} is a set of dummy variables measuring the effect of the euro on trade in services[11].

Anderson and van Wincoop (2003) show that the volume of trade between any two countries depends not only on their level of bilateral trade resistance, but also on how difficult it is for each of them to trade with the rest of the world, i.e. multilateral resistance. In our panel setting, we introduce exporter-year (λ_{it}) and importer-year fixed effects to account for any unobservable heterogeneity at the country level that vary with time. We also add country-pair fixed effects (λ_{it}) that controls for unobserved factors at the country-pair level. In this specification, time-invariant control variables are dropped due to perfect collinearity. Finally, u_{ijt} is a well-behaved disturbance term.

Sadeh (2014) argues that an appropriate control group must include enough countries that have not joined the euro area but that would have responded similarly to the launch of the euro had they joined it. Similar to Santana-Gallego et al. (2016) for tourism movements, the present study analyses trade in services flows between the 28-EU countries plus three non-EU countries (Switzerland, Norway and Iceland) that participate in the European Free Trade Association (EFTA), which is part of the EU's internal market. Moreover, to have enough countries in the control group, six non-European OECD countries (Australia, Canada, Japan, New Zealand, Turkey and United States) are also included[12].

4.2. Empirical results

Firstly, we focus on the overall impact of the euro on international trade in services. To that end, we distinguish three different specifications: Model (A) measures the effect of the euro on trade in services by using data from 1995 up to 2012 and controlling for the enlargement process of the EMU; Model (B) addresses differences in the effect of the euro depending on the date of inception, i.e. differences in the impact of the euro depending on whether the country initially adopted the euro in 1999 or joined later, and Model (C) takes into account the initial stage of the EMU when irrevocable exchange rates were set in 1999, and the second stage after the Euro started to circulate in 2002 for the Euro-11 countries. Results are presented in Table 6.

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[10] As suggested by BROUWER *et al*. (2007), dummy variables for both countries in the Eurozone or both countries in the EU are introduced separately as they represent two separate forms of economic integration: the first one, a first variable of interest, is an estimate of the marginal contribution of euro for participating countries, whereas the second is an estimate of the marginal contribution of EU for member countries. Note that Croatia joined the European Union in 2013, so although it is considered in the sample, it is not included in the EU_{ijt} dummy variable.

[11] Table A.1 in the appendix presents the countries included in the analysis as well as the date of the different enlargement episodes of the EU and the Eurozone used to define dummy variables.

[12] Mexico, Chile and Turkey are not included because of tourism data availability problems.

TABLE 6
EURO'S EFFECT ON TRADE IN SERVICES

	(A)	(B)	(C)
EU	0.384***	0.387***	0.367***
	–0,0447	–0,0436	–0,0437
Euro both	0.755***		
	–0,243		
Euro one	0.389***		
	–0,122		
Euro-11 both		0.763***	
		–0,243	
Euro-11 both (1999-2001)			0,121
			–0,341
Euro-11 both (2002-2012)			0.767***
			–0,244
Euro-11 one		0.388***	
		–0,122	
Euro-11 one (1999-2001)			0,0161
			–0,171
Euro-11 one (2002-2012)			0.409***
			–0,123
Euro-new both		0.648**	0.661**
		–0.289	–0,291
Euro-new one		0.298	0.309*
		–0.185	–0.185
Observations	23.758	23.758	23.758
Number of idpair	1.332	1.332	1.332
R-squared	0.791	0.791	0.792

NOTE: Significant at 1 per cent (***), 5 per cent (**) and at 10 per cent (*) level. Constant, EYFE, IYFE and CPFE are not reported. Standard errors appear between parentheses and p-values between brackets. Robust standard errors clustered by pair are computed.

In Model (A), a dummy variable that is unity when both countries in the pair belong to the EMU is defined (Euro both). This variable considers all the countries that belong to the EMU in a certain year. So, this variable jointly considers the initial countries that joined the EMU in 1999, as well as the new ones that joined during the various enlargements[13]. The coefficient of Euro both is positive and significant at 1 per cent level suggesting that the euro promotes intra-Eurozone service exports by

[13] For instance, Euro both takes the value zero for the pair France-Spain before 1999 and the value one since 1999.

a factor of 113 per cent[14]. This impact is larger than the one obtained by Sadeh (2014) or Glick and Rose (2016) for trade in goods.

Another relevant issue is to check whether adopting the euro has made the Eurozone more open to trade in services (trade creation) or, on the contrary, has led to more intense trade flows within the Eurozone at expense of diversion of trade in services with non-members (trade diversion). As defined by Sadeh (2014), a dummy variable that fully controls for trade with third party countries, whatever the direction, is included (Euro one). This variable takes the value one when only one country in the pair belongs to the EMU. Consequently, in Model (A) the excluded category is trade in service between two non-member states. If the estimated parameter of the variable Euro one is positive that provides evidence of trade creation (there is an increase on the volume of exports since a country substitute domestic production by imports from a foreign country when trade barriers are removed). On the contrary, if the estimated parameter is negative this suggests trade diversion (when a country substitute imports from a third country by imports from another member of the trade union after a change in relative bilateral resistances, i.e. the increase of relative costs with third-party countries could lead to tourism diversion). As presented in Table 6, the estimated coefficient shows that the euro's effect on trade in services with non-members is positive and around 47.5 per cent. Consequently, as for international trade, evidence of tourism creation is found[15].

In 1999, eleven countries joined the EMU, and afterwards six more countries adopted the euro at different stages. Model (B) addresses the different enlargement episodes on the effect of the euro depending on the date of inception, i.e. differences in the impact of the euro depending on whether the country initially adopted the euro in 1999, Euro-11, or joined later, Euro-new. In particular, Euro-11 both takes the value 1 if both countries in the pair joined the EMU in 1999, e.g. for the pair France-Spain for years 1999-2012. Euro-new both takes the value 1 when both of the countries in the pair are new members or when one of the countries in the pair is a new member and the other already belongs to the EMU. For instance, the pair Cyprus-Austria takes the value 1 for years 2008 to 2012. Euro-11 one and Euro-new one are accordingly defined to consider only one Euro-11 or a Euro-new country in the pair.

The estimated coefficients of both variables suggest that the impact of the euro on international trade in services is slightly higher for countries that initially joined the EMU rather than for those which incorporated afterwards. In particular, the impact of the euro on EMU-11 countries is 114 per cent, whereas the effect on new member states is around 91 per cent. For the trade creation/diversion effect, the impact on trade in services with third countries is 47.4 per cent for the EMU-11, suggesting

[14] The percentage effect is equal to $[\exp(\alpha) - 1] \times 100$, with α being the coefficient of the Euro dummy variable.

[15] Adopting the euro makes country members more open and therefore boosts their trade with third party nations (MICCO *et al*., 2003; FARUQEE, 2004, or CAFISO, 2011).

again evidence of trade creation. The trade creation effect for the new entrants is not significant.

Finally, Model (C) takes into account the initial stage of the EMU when irrevocable exchange rates were set in 1999, and the second stage when the Euro started to circulate in 2002. Two dummy variables are defined, Euro-11 both (1999-2001) that takes the value one if both countries in the pair belonged to the EMU-11 during the period 1999-2001, and Euro-11 both (2002-2012) that takes the value one when both countries are EMU-11 for the period 2002-2012. The former variable controls for the fixed irrevocable exchange rate between country members, although national currencies remained circulating, while the latter reflects the introduction of

TABLE 7

EURO'S EFFECT BY DESTINATION COUNTRY

Destination country	
Euro-11 both Austria	0.345**
Euro-11 both Belgium	0.366**
Euro-11 both Finland	0.769***
Euro-11 both France	0.335**
Euro-11 both Germany	0.321**
Euro-11 both Ireland	0.426**
Euro-11 both Italy	0.291**
Euro-11 both Luxembourg	0.413***
Euro-11 both Netherlands	0.350**
Euro-11 both Portugal	0.439***
Euro-11 both Spain	0.333***
Euro-new both Cyprus	0.0844
Euro-new both Estonia	0.524***
Euro-new both Greece	0.159
Euro-new both Malta	0.296*
Euro-new both Slovakia	0.399**
Euro-new both Slovenia	0.244**
Euro-11 one	0.417***
Euro-New one	0.250**
Observations	23.758
R-squared	0.793
Number of idpair	1.332

NOTE: Significant at 1 per cent (***), 5 per cent (**) and at 10 per cent (*) level. For simplicity standard errors are not reported. Constant, EYFE, IYFE and CPFE are not reported.

the euro as the national currency. From Model (C), it can be observed how the effects for the period 1999-2001 are not statistically significant. Therefore, the euro's effect on trade in service is concentrated since the common physical currency started to circulate.

Additionally, it is relevant to analyse the effect of the euro for each country and testing whether there are significant differences between them. Following Aristotelous (2006), we interact the Euro dummy variable with the destination country to obtain the impact of the euro for each member state. For instance, the variable Euro-11 both Austria takes the value one for the pair Austria and another Euro-11 country, since 1999. Similarly, Euro-new Cyprus takes the value one for the pair Cyprus and another Euro-11 or Euro-new country since 2008. Results are presented in Table 7 and it can be observed how the euro's effect on trade in services is relatively homogeneous across country members[16].

It is observed that the effect of the euro is significantly positive for all the initial members of the EMU, as well as for the new member apart from Cyprus and Greece. For the EMU-11 members, the largest estimated effects are found for Finland (115 per cent), Portugal (55 per cent), Ireland (53 per cent) and Luxembourg (51 per cent). For the case of Spain, the individual euro's effect is around 40 per cent, below the average for the rest of the Eurozone. Regarding the new members, the largest euro's effect is estimated for Estonia (69 per cent) and Slovakia (49 per cent).

5. Concluding remarks

In contrast to the extensive literature on trade in goods, trade in services has traditionally received fewer attention in the empirical literature. This lack of interest cannot be justified by its relevance in the world economy, since there is an increasing share of trade in service in world's GDP. For this reason, understanding the effect of trade in service on countries' economy and exploring the impact of economic integration on this flow are crucial for policymakers.

The objective of this paper is twofold. Firstly, we aim to empirically justify the effect of trade in service on economic growth. To that end, we use Spain as a case study. The empirical literature on the effect of merchandise exports on GDP is not conclusive. In our analysis, for the case of Spain, we also find a not conclusive result. In all cases where significant effects are found for merchandise trade, it is always negative and persistent. Conversely, for service exports we found a significant and positive effect on output in the short run. Consequently, it seems that trade in service is relevant to promote economic growth, at least in the short-run. In any case, the high degree of cross-correlation between the regressors seems to mask the true underlying relations.

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[16] This results are in contrast to the heterogeneous results for the euro's effects for the case of international trade in goods (GIL-PAREJA et al., 2007; MICCO et al., 2003; FARUQEE, 2004; ARISTOTELOUS, 2006).

Secondly, this paper contributes to the debate on the costs and benefits of monetary unions by analysing the role of the euro on intra-Eurozone trade in service flows. The estimated impact of the euro on tourism flows is 113 per cent. This effect is larger to the one estimated for trade in goods. Moreover, we find evidence of trade creation and the euro's effect on trade is limited to the period when the currency started to circulate (2002). Additionally, it seems that trade gains from adopting the euro have been evenly distributed among member states. For the specific case of Spain, the individual euro's effect is around 40 per cent, which is below the average for the rest of the Eurozone

These findings are relevant for demonstrating the effect of adopting the euro or for joining other currency union experiences. A better understanding of the euro's effect on trade in services contributes to the debate on the benefits of joining the Eurozone. In any case, this is only one dimension of the effect of the euro. Other economic consequences of the political integration need to be evaluated. Overall, our research provides policymakers of future and potential entrants with an additional argument in favour of joining the EMU.

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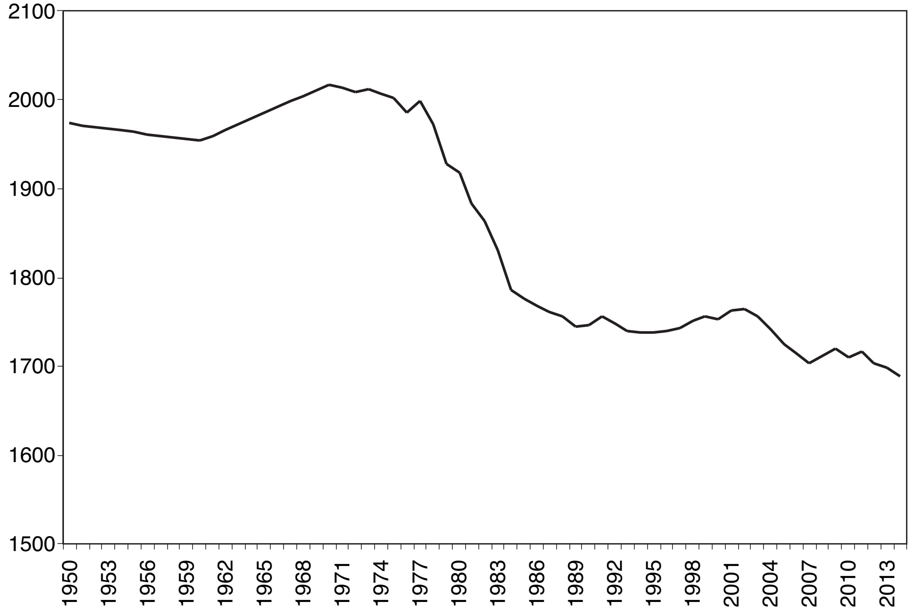
APPENDIX

FIGURE A1
ALTERNATIVE MEASURES OF LABOUR INPUT

A1.1. Average annual hours worked



A.1.2. Employment (millions of persons)

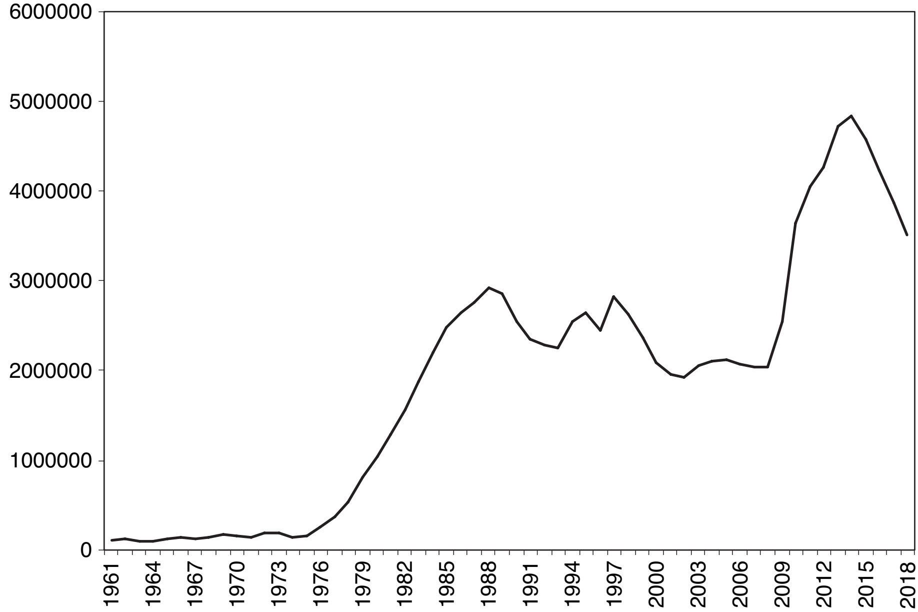
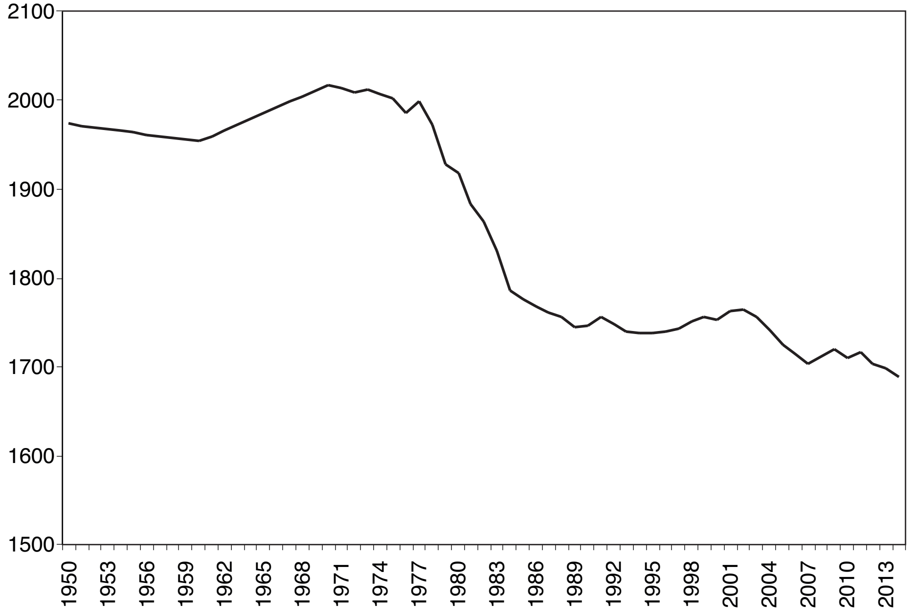


FIGURE 2
REAL SERIES (BASED ON US GDP DEFLATOR)

A.2.1. GDP(Y)

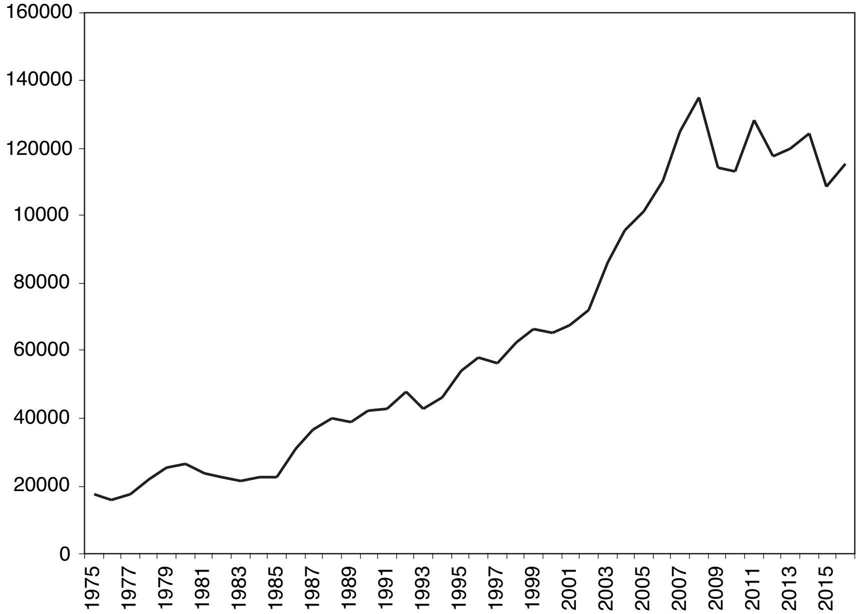


A.2.2. Exports of goods and services (X)

2.3. Non-export output (N = Y - X)

2.4. Gross Capital formation

2.5. Service exports (S)



A.2.6. Goods exports (G)

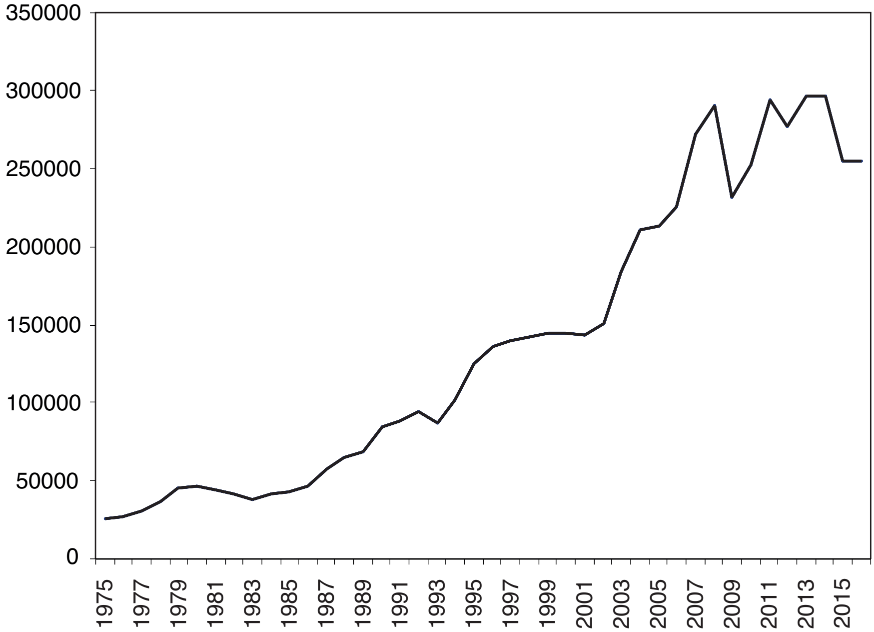


TABLE A.1
LIST OF COUNTRIES CONSIDERED IN THE DATASET AND DATE OF ADOPTION

	Euro	EU		Euro	EU
Australia			Latvia		2004
Austria	1999	1995	Lithuania		2004
Belgium	1999	1957	Luxembourg	1999	1957
Bulgaria		2007	Malta	2008	2004
Canada			Netherlands	1999	1995
Croatia*		2013	New Zealand		
Cyprus	2008	2004	Norway		
Czech Rep.		2004	Poland		2004
Denmark		1973	Portugal	1999	1986
Estonia	2011	2004	Romania		2007
Finland	1999	1995	Slovakia	2009	2004
France	1999	1957	Slovenia	2007	2004
Germany	1999	1957	Spain	1999	1986
Greece	2001	1995	Sweden		1995
Hungary		2004	Switzerland		
Iceland			Turkey		
Ireland	1999	1995	United Kingdom		1995
Italy	1999	1973	United States		
Japan					

NOTE: * Croatia joined the EU in 2013, so it is not included in the EUijt dummy variable since our sample covers from 1995-2012.