Milestones of European Integration: Which matters most for Export Openness?

Sanne Hiller and Robinson Kruse

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Abstract

The European integration process has removed barriers to trade within Europe. We analyze which integration step has most profoundly influenced the trending behavior of export openness. We endogenously determine the single most decisive break in the trend, account for strong cross-country heterogeneity and propose a new measure for the strength of trend breaks. Highly open economies gain from both, monetary and real integration. In sharp contrast, less open economies do not benefit from real integration and even suffer from monetary integration. The major milestones for France, Germany, Italy and the Netherlands are the Euro introduction, the Maastricht Treaty, the Exchange Rate Mechanism I and the merge of EFTA and EEC to the European Economic Area, respectively. Our empirical results have important implications for inner-European economic development, as export openness feeds back into growth, unemployment and income convergence.

Keywords: European Integration; Export Openness; Trends; Structural Breaks

JEL-Codes: C22, F02, F15, F41

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1 Motivation

Export openness measures a country's de-facto integration in the world economy, which exhibits ample effects: It might stimulate economic growth (Frankel and Romer 1999, Buch and Toubal 2009, Felbermayr et al. 2010), reduce unemployment (Dutt et al. 2009, Felbermayr et al. 2009) or raise productivity (Alcalá and Ciccone 2004). Moreover, both the factor-price-equalization theorem in a Heckscher-Ohlin world, as well as extensions to Solow's growth model advocate that trade liberalization fosters convergence (see Ben-David 1996 for a detailed discussion). Empirically, Ben-David (1993) finds evidence for this link between income convergence and openness in the European Economic Community (see Slaughter 2001 for a summary of related empirical literature; compare Ben-David 2001 for a comment).

In light of the goal to boost inner-European trade and thereby to strengthen economic ties - for its own merit and for the sake of political and economic stability - the stepwise removal of tariff and non-tariff barriers to trade within Europe may have affected export openness fundamentally. Export openness generally exhibits a time trend. Events of European integration may have fundamentally changed this trend: Sequentially, separating economic, political and institutional walls have been torn down starting out with the Treaty of Rome 1957. Early economic integration steps have established a customs union in 1968, paved the way to monetary integration via the Exchange Rate Mechanism I 1979, and confirmed the intention to build a unified Europe in periods of skepticism like the Single European Act 1986. Later, the Maastricht Treaty 1992 has founded the European Union, the common market has been extended with the establishment of the European Economic Area 1994. Finally, the Euro has been launched as common currency in 1999. Moreover, several waves of enlargement have taken place, the latest being the entry of Bulgaria and Romania in 2007. A priori, it is not clear which event has most crucially affected export openness in Europe.

Our main contribution is to take a bird's eye view on the major steps of European integration and to analyze which milestone brings about the most crucial structural change in export openness of France, Germany, Italy and the Netherlands. Inspired by

the seminal work of Ben-David and Papell (1997) on trend breaks in openness, we endogenously determine the single most decisive break in the trend function. We accommodate country-pair specific trending behavior and heterogeneous effects of European policy on exports.

We apply a state-of-the-art testing procedure proposed by Perron and Yabu (2009). The testing procedure is custom-tailored to our needs: First, the test is valid in the case of stationarity or non-stationarity of openness. As our empirical results confirm, both cases are relevant. Thus, the test outcome is not influenced by pre-tests or unrealistic assumptions. Secondly, the test proves to be powerful, while exhibiting good size properties also in small samples, see Perron and Yabu (2009).

In order to assess the major milestone for openness, we propose a new and simple measure for the strength of a trend break. This measure remedies shortcomings of competing approaches. It is based on a comparison of the trend functions with and without an estimated break. It allows us to quantify the periodical openness gain (or loss) which is due to a structural change in the trend and thus, we are able to rank the importance of milestones.

Section 2 outlines the core integration events and their relation to export openness. Section 3 discusses general methodological issues. Section 4 provides a description of the econometric methodology, while section 5 presents and discusses empirical results. Main conclusions are drawn in section 6.

2 Major Steps of European Integration and Export Openness

The Treaty of Rome 1957 establishes the European Economic Community (EEC), also called "Common Market". It sets a road map to the European integration, which aims at the foundation of a common market with free trade in goods, services and factors. Eleven years later, in 1968, the Customs Union - a free trade area with a common tariff applied to third countries - is completely established. Early evidence points to an

increase in trade within the EEC due to the Common market, with both trade creation due to the reduction of intra-area tariffs as well as trade diversion, i.e., a reallocation of trade with non-members to member states (Balassa 1967).

Eleven years later, in 1979, the ERM I establishes the link of European currencies through a semi-pegged system as the first major step of monetary harmonization. It is set up to path the way to a single common currency, and to bring down exchange rate volatility within Europe and thereby to stabilize the monetary environment. Thus, already 30 years before the final launch of the Euro, a crucial monetary channel for a potential openness gain is opened. Recently, Gil-Pareja et al. (2007) find evidence for a trade-creating effect of ERM I. According to their results, this effect occurs on top of the trade expanding effect of exchange rate stability. They argue, that this effect is due to anticipation of additional real integration steps yet to come. Moreover, they find that especially periphery countries benefit from ERM, whereby the benefit manifests itself in an increased intra-EU trade volume. Using cointegration techniques, Fountas and Aristotelous (1999) find that growth in intra-EU trade is almost independent of the prevailing exchange regime. Still, they find that ceteris paribus trade is greater when the exchange rate exhibits a lower volatility. Tenreyro (2007) studies the effect of exchange rate stability on trade volumes, and finds no evidence for a trade-creating effect.

The Single European Act is the first major revision of the Treaty of Rome in response to the existence of considerable de-facto barriers to trade within Europe. Together, the two demarcate the transition period of the EEC (Ben-David 1993). In light of persisting de-facto barriers to trade within Europe, the Act's importance derives from the confirmation of the member states' conviction to form a European Union. Thus, after a couple of years of Euro-pessimism, it increases trust among economic agents that integration steps will not be reversed but rather be extended in the future. For an early discussion of the importance of the SEA, see McAleese and Matthews (1987).

As promoted in the SEA, the Maastricht Treaty 1992 accomplishes the common market within Europe with the establishment of the European Union and it paves the way to the currency union. Subsequently, in 1994, the European Economic Area is established, allowing members of the European Free Trade Area (EFTA) to participate in the single

market as established in 1993, whereby maintaining their currencies.

The most recent integration step has been the launch of Euro in 1999, which has been set in the Maastricht Treaty. On top of the elimination of exchange rate uncertainty, the introduction of a common currency implies substantially lower transaction cost due to the same money used for transactions. Presumably, the Euro introduction is the most studied among the core events, see Baldwin (2006) for a comprehensive survey. As in one of the early studies by Micco et al. (2003), a positive effect of the Euro is the prevalent finding. The estimated size of the effect varies substantially across studies (Berger and Nitsch 2008, Bun and Klaasen 2007, Glick and Rose 2002; compare Rose and Engel 2002, Rose and van Wincoop 2001 and Rose 2000). Still, as emphasized by Berger and Nitsch (2008), the Euro effect on trade has rarely been put into perspective with preceding integration efforts.

3 Methodological Aspects

The broad perspective we take is essential: Each of the milestones may have shaped European export openness through the removal of barriers to trade. Limiting the view to one event may lead to negligence of more crucial ones. In this case, empirical results for the effect of a certain event on trade are likely to be biased. We want to explicitly model these events if they matter and to rank the role they play in the process. Thus, we consistently estimate the single most decisive breakpoint in export openness series. Since the breakpoint is determined endogenously, we detect the most crucial event which has brought about a trend break - even if it is unrelated to the European integration process. If for example, the oil crisis has triggered the most decisive structural change in openness within Europe, this shows up in our results. This is a strength of our approach, since the test for trend breaks and the breakpoint estimation is independent of the number of potential breakpoints in our analysis (seven) and their location.

Clearly, our focus lies on the founding members of the EU, since they are directly affected by all integration events since the Treaty of Rome. Of course, the number of

participating countries increases with the accession waves. We do not include them in the set of milestones, because the accession to an organization is different from a modification of the organization itself. Still, as our approach is independent of the specification of the events and our sample comprises the founding members' exports to EU-27 countries, the accession waves are allowed to matter.

Our testing procedure does not allow to draw conclusions on what caused the structural change in the trend, which is a typical feature of a structural break analysis (compare Ben-David and Papell 1997, Kočenda 2005 among many others). For this reason, we use a control group to assess whether the association of breaks with EU events indeed reflects EU policy and not general changes in the economic environment. A good control group should be characterized by a substantial amount of exports to EU-27 countries, and share the conjectured exposition to global economic trends. This makes the US an appropriate control group. Our choice is supported by an argument of Serlenga and Shin (2007) who use the US reference variables to capture observable common time specific factors explaining within-EU trade.

Our object of interest is multilateral openness within Europe. Still, we want to account for the heterogeneity of bilateral trade relations since their trending behavior is pair-specific (compare Bun and Klaasen 2007). Micco et al. (2003) find considerable variation of the effect of the Euro on participating countries; a priori, there is no reason to assume homogeneity in intensity and timing for other events either. Therefore, inspired by Cheung and Lai (2000), we model each bilateral trade relationship individually, and aggregate our findings to the multilateral level as described in section 5.

The data set contains bilateral export openness y_t^{ij} which is defined as the ratio of exports (EXP) from country i to country j over the exporter's Gross Domestic Product (GDP), i.e. $y_t^{ij} = \text{EXP}_t^{ij}/\text{GDP}_t^i$. Annual export data originates from the Directions of Trade Statistics provided by the International Monetary Fund. Nominal GDP data from the International Monetary Fund is converted in national currency to US dollar using the periodical average exchange rate from the same source. We analyze an unbalanced panel of export openness from four founding members to 18 trade partners
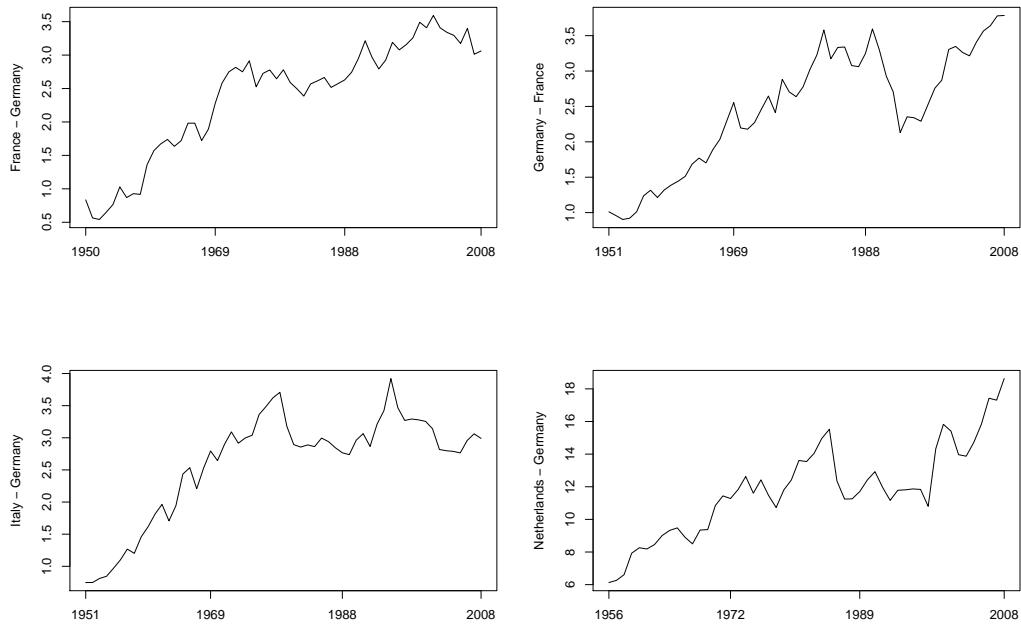
Figure 1: Export Openness of EU-4 countries with their major trading partners.

(72 bilateral series). As a control group, we additionally analyze US export openness. A typical sample length is $T = 58$.[1]

Figure 1 shows export openness time series for the four exporting countries to their major trade partner in 2008. The four plots suggest that export openness is positively trending over time, but the trending behavior is changing. The locations, directions and magnitudes of trend breaks are heterogeneous.

4 Econometric Approach

This section deals with the Wald-type test for trend breaks suggested by Perron and Yabu (2009). We consider a popular trend model which contains a linear trend and allows for breaks in both, the intercept and the trend slope. This specification is common in the related literature and has been successfully applied in modeling many economic

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[1]Full information about the data set is available from the authors.

time series like US real GDP, see Ben-David and Papell (1995). Moreover, it fits well with the nature of data shown in Figure 1.

The trend break model for bilateral export openness[2] between countries i and j is given by

$$y_t = \mu_0 + \mu_1 DU_t + \beta_0 t + \beta_1 DT_t + u_t \tag{1}$$

with $DU_t = 1(t > T_B)$ being a step dummy and $DT_t = 1(t > T_B)(t - T_B)$ being a trend dummy. This specification permits a simultaneous change in the intercept and the slope of the trend at the breakpoint $T_B = \text{int}(\lambda T)$. The intercept changes from μ_0 to $\mu_0 + \mu_1$ while the trend slope changes from β_0 to $\beta_0 + \beta_1$. The relative breakpoint λ is restricted to be an element of the interval $\Lambda = \{\lambda; \ \epsilon \leq \lambda \leq 1 - \epsilon\}$. Common choices for the trimming parameter ϵ are values between 0.01 and 0.20. In our application, we specify $\epsilon = 0.05$ since our sample size is small. We aim to avoid power losses from specifying the value of ϵ too low, therefore $\epsilon = 0.05$ is a reasonable compromise. This means for a case of $T = 58$ where annual data ranges from 1951 to 2008 that the interval of potential breakpoints is given by $(1954, 2005)$ which is large enough in order to cover the major events of European integration and small enough to maintain good power.

The null hypothesis of no trend break is given by $H_0 : \mu_1 = 0, \beta_1 = 0$. Hence, under the validity of the Null, y_t is modeled by a linear trend $\mu_0 + \beta_0 t$ with noise u_t. The noise process is allowed to be autocorrelated and even non-stationary. Moreover, the testing procedure suggested by Perron and Yabu (2009) does not even require the knowledge of the degree of integration of u_t. It is well known that the limiting distribution of test statistics differ with the degree of integration of u_t in general. The test by Perron and Yabu (2009) is robust in both cases, that is $u_t \sim I(0)$ and $u_t \sim I(1)$, in the sense that the test is correctly sized and powerful in both situations. Therefore, reliable inference can be conducted by using this test.

For a given breakpoint T_B, the test for the null hypothesis is carried out in a five-step procedure: As a first step, the parameters of the trend model are estimated by OLS.

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[2]The ij superscript is omitted to increase the readability.

The residuals $\widehat{u}_t$ can therefore be interpreted as de-trended data. In order to estimate the persistence in the residual series, which is measured as the sum of autoregressive coefficients, the autoregression

$$\widehat{u}_t = \alpha \widehat{u}_{t-1} + \sum_{i=1}^{k} \rho_i \Delta \widehat{u}_{t-i} + e_t \tag{2}$$

is considered as a second step. Here, α represents the sum of autoregressive coefficients. The lag length k is chosen using the Bayesian Information Criterion (BIC) with a maximal value of $k_{\max} = \text{int}(4(T/100)^{1/4})$. Since the estimate of α can be heavily biased in small samples, the bias-correction method suggested by Roy and Fuller (2001) is applied as a third step. This procedure leads to an approximately median-unbiased estimate of α, see Roy and Fuller (2001).

Once the persistence of the residual series $\widehat{u}_t$ is estimated, a truncation is applied in order to make the estimator for α super-efficient in step four. This truncation has the general form

$$\widehat{\alpha}_S = \widehat{\alpha} \cdot 1\left(T^\delta |\widehat{\alpha} - 1| > d\right) + 1 \cdot 1\left(T^\delta |\widehat{\alpha} - 1| \leq d\right)$$

with $\delta \in (0,1)$ and positive d. If $\widehat{\alpha}$ is in a $T^{-\delta}$ neighborhood of 1, then it is assigned a value of one. As suggested by Perron and Yabu (2009), we specify $\delta = 1/2$ and $d = 1$. Since the super-efficient estimator will be based on the bias-correction in our applications ($\widehat{\alpha}_M$), it is denoted and defined as

$$\widehat{\alpha}_{MS} = \widehat{\alpha}_M \cdot 1\left(|\widehat{\alpha}_M - 1| > T^{-1/2}\right) + 1 \cdot 1\left(|\widehat{\alpha}_M - 1| \leq T^{-1/2}\right). \tag{3}$$

This truncation bridges the gap between $I(0)$ and $I(1)$ noise and ensures that the test for H_0 has approximately the same limiting distribution in both cases.

The fifth step applies a quasi feasible GLS (QFGLS) estimator to the specified trend model from the first step. The trend break model (1) is considered with regressors $x_t = (1, DU_t, t, DT_t)'$ and parameters $\Psi = (\mu_0, \mu_1, \beta_0, \beta_1)'$: $y_t = x_t'\Psi + u_t$. The QFGLS

regression is then given by

$$(1 - \widehat{\alpha}_{MS}L)y_t = (1 - \widehat{\alpha}_{MS}L)x_t'\Psi + (1 - \widehat{\alpha}_{MS}L)u_t \tag{4}$$

for $t = 2, \ldots, T$ and $y_1 = x_1'\Psi + u_1$ for $t = 1$. L denotes the lag operator, i.e., $Lz_t = z_{t-1}$. The QFGLS estimator of the trend parameters is labeled as $\widehat{\Psi}$. The Wald test statistic for the null hypothesis $H_0 : R\Psi = (0,0)'$ is then given by

$$W(\lambda) = \left(R\widehat{\Psi}\right)' \left(\widehat{h}_v R(X'X)^{-1}R'\right)^{-1} R\widehat{\Psi} \tag{5}$$

with $X = \{(1 - \widehat{\alpha}_{MS}L)x_t\}$ for $t = 2, \ldots, T$ and $X = \{x_t\}$ for $t = 1$; $\widehat{h}_v$ is an estimate of the spectral density function at frequency zero of $v_t = (1 - \alpha L)u_t$ capturing the nature of the errors, $I(0)$ or $I(1)$. For details, see Perron and Yabu (2009). As indicated, this Wald statistic depends on a certain value of λ. Since we shall not assume a known breakpoint, the Wald statistic $W(\lambda)$ is computed for a sequence of values of $\lambda \in \Lambda$. Hence, we obtain a sequence of Wald test statistics $W(\lambda \in \Lambda)$. The individual test statistics are merged into a single statistic. As suggested by Perron and Yabu (2009), an exponential transformation is used for this purpose, see also Andrews (1993) and Andrews and Ploberger (1994):

$$\text{Exp-}W = \ln\left(\frac{1}{T}\sum_{\lambda \in \Lambda} \exp\left(\frac{1}{2}W(\lambda)\right)\right). \tag{6}$$

Critical values for the Exp-W test statistic are reported in Perron and Yabu (2009). The simulation evidence therein shows that the power of test is very satisfying even in small samples unless the break size is too small. The true and unknown breakpoint T_B is estimated by minimizing the sum of squared residuals obtained from QFGLS regression (4). Hence, the break date estimator $\widehat{T}_B$ is given by

$$\widehat{T}_B = \arg\min_{\lambda \in \Lambda} \sum_{t=1}^{T} \widehat{v}_t(\lambda)^2 \tag{7}$$

with $v_t(\lambda) = (1 - \widehat{\alpha}_{MS}(\lambda)L)u_t(\lambda)$.

Table 1: **Breaks in the Trend and Major European Integration Steps**

		France	Germany	Italy	Netherlands	USA
Total	Number of Breaks	14	12	14	13	11
Total	Event Breaks	7	7	5	10	3
1957	Treaty of Rome					2
1968	Customs Union	1				
1979	ERM I			2	2	
1986	SEA	2		2	3	
1992	Maastricht		4			
1994	EEA	1	2		3	
1999	Euro	3	1	1	2	1

This table depicts the estimated break dates of export openness with respect to 18 EU-27 countries which coincide with a major European integration event allowing for ± 1 year deviation.

5   Empirical Results

5.1   Trend Breaks in Export Openness

We test for trend breaks on a nominal significance level of 10% and find strong evidence for breaks in the trend of export openness. As reported in Table 1, we find 14 trend breaks in French export openness, 12 in Germany, 14 in Italy, 13 in the Netherlands and 11 in the USA. Out of these estimated 64 breaks in total, 47 are significant even at a 1% level and 57 on a 5% level.

Trend breaks are a prevailing feature of bilateral export openness within Europe (see Table 1): Almost 3/4 of the analyzed 72 series exhibit a structurally changing trend (53/72). More than half of these breaks (28/53) occur in association with the seven major European integration steps. An estimated break is considered as associated to an event if estimated either one year before, after or at the event date. Since no breaks are estimated in 1993, discrimination between the Maastricht Treaty and the European Economic Area is not of empirical concern.

As can be seen from Table 1, the picture is different across countries. In case of France, 78% of bilateral openness series exhibit (14/18) breaks. Half of these are associated

Table 2: **Estimated Non-EU Event Breaks**

	France	Germany	Italy	Netherlands	USA
Total	7	5	9	3	8
Accession*	2	2	2	2	0
Break with accessor°	0	0	2	1	0
	1954	1974*	1955	1972*°	1955
	1962	1976	1973*°	1996*	1975
	1966	1989	1975	2001	1975
	1971	1990	1982*		1976
	1981*	1996*	1988		1976
	1984		1989		1984
	1996*		1990		1989
			1990		1989
			2003*°		

This table reports estimated break dates that are not associated to major European integration steps. Asterisks indicate that at least one country accesses the EU in the respective year ($\pm$ 1 year). A circle signifies that the break occurs with an accessing country ($\pm$ 1 year)

with one of the seven key events of the European integration process (7/14). Next, we briefly consider the distribution of estimated breakpoints. Under uniformity of breakpoints and conditional on existence of a break, the probability that a break is estimated to coincide with one of the seven milestones is equal to 40% for a representative sample length. This benchmark for a purely random association with EU events is hit by France (50 %), Germany (58%) and the Netherlands (77%). Italy and our control group remain below the benchmark (36% and 27%, respectively). This implies that the results for Italy should not be overinterpreted.

A close look at the non-associated break dates as depicted in Table 2 does not suggest a systematic clustering of break dates in Germany, France and the Netherlands. In the case of Italy, a substantial cluster can be found in the end of the 1980s. Differently, trend breaks of the US bilateral export openness cluster strongly around 1975 and in the 1980s, presumably reflecting effects of the first and second oil crises. For all four EU countries, two non-associated breaks occur at the same time of an EU enlargement, while none are found for the US. Importantly, only for Italy and the Nether-

Table 3: **Estimated Persistence of the Noise Term**

	France	Germany	Italy	Netherlands	USA
$I(1)$	10	7	10	12	5
$I(0)$	4	5	4	1	6
$I(0)$, Ave $\widehat{\alpha}_{MS}$	0.59	0.46	0.57	0.65	0.57

This table shows how often the truncated and bias-corrected estimator $\widehat{\alpha}_{MS}$ indicates a non-stationary $I(1)$ or stationary $I(0)$ noise term. For the latter case, the average (Ave) of the estimated persistence is given. In the former case, it equals always one.

lands, breaks occur with countries which access the EU at the break dates (UK: 1973 and Cyprus: 2003; UK: 1973).

Table 3 confirms that the robust test by Perron and Yabu (2009) is indeed needed for this data set: Bilateral export openness does not generally exhibit a stationary or a non-stationary noise component. Thus, a unique testing approach should not impose assumptions of (non-)stationarity on the noise term.

5.2 The Most Decisive Milestone for Multilateral Export Openness

We are interested in measuring the strength of the effect a trend break has. This enables us to judge which milestone is most important. A possible approach would compare the average level of export openness before and after the break as done in Ben-David and Papell (1997). Given a trending time series with a trend break, a potential problem is the negligence of time-dependent means in each sub-sample. The means are time-varying not only before and after the break, but also within each sub-sample. If the time series was not trending at all, such a comparison would be admissible. Alternatively, one may compare the trend slope in the pre- and the post-break sub-sample, respectively. These are given by $\widehat{\beta}_0$ and $\widehat{\beta}_0 + \widehat{\beta}_1$, respectively, see

$$y_t = \mu_0 + \mu_1 DU_t + \beta_0 t + \beta_1 DT_t + u_t \ .$$

Although this approach accounts for time-varying means, it ignores the mean shift from $\widehat{\mu}_0$ to $\widehat{\mu}_0 + \widehat{\mu}_1$. In order to circumvent the mentioned shortcomings, we propose a new and simple measure for the strength of trend breaks. We propose to measure the strength of a trend break on bilateral export openness as follows:

$$\mathcal{S} = \sum_{t=\widehat{T}_B+1}^{T} \widehat{e}_t \tag{8}$$

where $\widehat{e}_t$ is the difference between the estimated deterministic trend under (i) a break (B) and (ii) no break (NB) during the post-break sub-sample[3], i.e.,

$$\widehat{e}_t = \widehat{y}_t^B - \widehat{y}_t^{NB}, \qquad t = \widehat{T}_B + 1, \widehat{T}_B + 2, \ldots, T. \tag{9}$$

The estimated trend function with a break is given by $\widehat{y}_t^B = \widehat{\mu}_0 + \widehat{\mu}_1 + \widehat{\beta}_0 t + \widehat{\beta}_1 \widehat{DT}_t$, while the extrapolated trend function without a break is $\widehat{y}_t^{NB} = \widehat{\mu}_0 + \widehat{\beta}_0 t$. The latter is a special case of the former one ($\widehat{\mu}_1 = \widehat{\beta}_1 = 0$). Inserting for $\widehat{y}_t^B$ and $\widehat{y}_t^{NB}$ gives the difference between the actual trend function (under a break) and the hypothetical one (under no break):

$$\widehat{e}_t = \left(\widehat{\mu}_0 + \widehat{\mu}_1 + \widehat{\beta}_0 t + \widehat{\beta}_1 \widehat{DT}_t \right) - \left(\widehat{\mu}_0 + \widehat{\beta}_0 t \right). \tag{10}$$

The measure $\mathcal{S}$ can be interpreted as the cumulated differences between the actual and the hypothetical trend function. Even though this measure is related to Balassa (1967) as it compares the differences between *actual and hypothetical* values, it is not our goal to establish a fully-fledged ex-post counterfactual of export openness that would have occurred without the considered integration step (compare Magee 2008). Instead, we want to measure the strength of the break.

Simple (and omitted) steps of calculation allow us to write $\mathcal{S}$ as

$$\mathcal{S} = (T - \widehat{T}_B)\widehat{\mu}_1 + \frac{\widehat{\beta}_1}{2} \left(T(T+1) - \widehat{T}_B(\widehat{T}_B + 1) \right). \tag{11}$$

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[3]The trends are the same in the pre-break sub-sample.

Table 4: **Breaks in the Trend and Related Gains/Losses**

		France	Germany	Italy	Netherlands	USA
1957	Treaty of Rome					-0.78
1968	Customs Union	-0.01				
1979	ERM I			-5.80	0.57	
1986	SEA	1.57		2.02	-4.25	
1992	Maastricht		6.22			
1994	EEA	-0.001	4.80		5.83	
1999	Euro	-1.58	1.45	-0.45	2.27	-0.01

This table reports periodical gains (or losses) $\widetilde{\mathcal{S}}^{ij}$ which are associated with a trend break. If more than one break is associated to a milestone, compare Table 1, then the sum of $\widetilde{\mathcal{S}}^{ij}$ is reported.

For comparability across different pairs with potentially different break dates, we normalize $\mathcal{S}$ by number of post-break periods such that the scaled measure, $\widetilde{\mathcal{S}}$, can be interpreted as the percentage gain (or loss) per period in openness induced by the trend break,

$$\widetilde{\mathcal{S}} \equiv \frac{\mathcal{S}}{T - \widehat{T}_B} = \widehat{\mu}_1 + \frac{\widehat{\beta}_1}{2}\left(T + \widehat{T}_B + 1\right) . \tag{12}$$

Intuitively, $\mathcal{S}$ and $\widetilde{\mathcal{S}}$ can be either positive or negative, depending on the estimated parameters $\widehat{\mu}_1, \widehat{\beta}_1$ and the estimated breakpoint $\widehat{T}_B$. $\mathcal{S}$ and $\widetilde{\mathcal{S}}$ approach zero as $\widehat{\mu}_1$ and $\widehat{\beta}_1$ both tend to zero.[4]

At this stage, we calculate $\widetilde{\mathcal{S}}^{ij}$ for bilateral trade relations ij which exhibit a trend break. Table 4 reports the results for $\widetilde{\mathcal{S}}^{ij}$. If we find more than one bilateral trade relation to exhibit a trend break at a milestone (see Table 1), the sum of individual values of $\widetilde{\mathcal{S}}^{ij}$ is reported. The summation is applied in order to interpret the results at a multilateral level. As multilateral openness is the sum of bilateral openness, summation is appropriate. It can be seen from Table 4 that EU events not only affect the number of export relations differently across countries, but this heterogeneity also translates into distinct openness gains and losses for the individual countries.

For France, the most incisive event has been the Euro introduction, and thus a mone-

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[4]In a few rare cases a correction of $\mathcal{S}$ and $\widetilde{\mathcal{S}}$ is needed in order to ensure positivity of $\widehat{y}_t^B$ and $\widehat{y}_t^{NB}$.

tary integration step. It brings about a periodical openness loss of -1.58%. The loss in overall periodical export openness does not result from structural changes with France's major trade partners (Germany, Italy, Spain, UK, Belgium), but arises from countries located in the European periphery, namely Ireland (-0.14%), Portugal (-1.33%) and Cyprus (-0.11%). This finding is in line with Aristoteleous (2006) who finds a negative trade effect of the Euro in France. Empirical evidence suggests that French exports are not responsive to exchange rate volatility (Fountas and Aristotelous, 1999). This closes down one beneficial channel of monetary integration for France, such that the launch of the Euro has led to a relative disadvantage for France. Moreover, the negative Euro effect may reflect what has been highlighted by Bayoumi and Eichengreen (1997): France's economic readiness for monetary integration is limited, and its participation rather driven by political than by economic considerations.

In contrast to France, the most decisive break for Germany occurs in conjunction with the Maastricht Treaty. The effect is large and positive amounting to 6.22%. Not surprisingly due to the prominent role in the integration process, the geographical location and its long tradition as exporting economy, Germany is the main beneficiary of the most comprehensive integration step. The openness gain results from trade with Austria (2.89%), France (-0.20%), Italy (1.03%) and Spain (2.50%). Notably, all four countries driving the openness gain belong to the top ten German export destinations. However, the break in the bilateral openness with Austria may be also be attributed to the German Reunification as both countries share the same language and have an intensive trade relation.

The outstanding event shaping Italian export openness is the ERM I. The effect is negative and quite large: -5.80%. It fully derives from exports to other EU founding members and participants in the ERM I, namely its most important export destination Germany (-5.41%) and the Netherlands (-0.39%). The beneficial effect from lower exchange rate volatility does not manifest itself for Italy. This corresponds to Fountas and Aristotelous (1999): the ERM I has not fostered exports due to supply-side problems of participating economies, restrictive fiscal policies and the remaining presence of trade barriers. Still, the association of Italian export openness breaks with European integra-

tion event is very low in general. This points to the existence of more important factors than the European integration which drive the path of Italy's export openness, like for example Italian industrial restructuring taking place.

The most decisive event for Dutch export openness is the EEA bringing about a 5.83% periodical gain. The Netherlands have strongly benefited from the widening of the single market to EFTA countries. Unsurprisingly, the increase in openness per year fully emanates countries which are not members of the EU in 1994. It stems from former EFTA countries, Austria (1.16%) and Finland (1.35%), as well as from Poland (3.32%). None of these is a major export destination for the Netherlands. As Creusen and Lejour (2009) point out, the major rise in Dutch exports originates from long-standing trade partners as Germany, France, Belgium, and the United Kingdom. Still, the most crucial change due to European trade policy relates to non-traditional export destinations: this highlights the particular flexibility of the small open economy par excellence.

Importantly, empirical evidence for the control group is fundamentally different. The Treaty of Rome stands out as the most decisive event with a minor loss of 0.78%. This shows that the ties to events which seem to matter for four European countries are weak. This confirms our approach: If the European breaks were caused by global economic events unrelated to European policy, they would be found in American export openness to Europe as well, since the economic ties between the USA and Europe are strong.

Across events, France does neither benefit nor lose from European integration: Its largest loss from the Euro is almost entirely compensated by the Single European Act. In particular, France does not benefit from real integration steps undertaken in the 1990s. On the contrary, Germany's openness enjoys substantial gains in this decade from all milestones. However, the gains arising from real integration are larger. Differently, in the 1990s, Italian openness loses its connection to the European integration process reflecting the uncertainty to participate in the EMU. Similarly to Germany, the Netherlands strongly benefit from EEA and the Euro introduction, but they suffer from the previously installed SEA.

Early integration steps do not matter. Although different in size and magnitude, SEA is the first milestone which hits all countries except of Germany. On the contrary, the Treaty of Maastricht affects only Germany. This is the strongest effect we find overall. Not only do the Netherlands strongly benefit from the EEA, but also does their neighbor Germany. ERM I only hits Italy. Most importantly, we now compare the effect of the Euro introduction on export openness: Notably, for no other country than France, the Euro has most decisively shaped the long-term trend of export openness. On top of this, when compared to the gains and losses for the major event in the other three European countries, the effect of the Euro introduction is the smallest in absolute value. This supports the main finding by Berger and Nitsch (2008) of no positive significant effect of the common European currency on exports which prevails on top of the preceding and continuing trade liberalization, economic harmonization and stabilization.

More open countries, Germany and the Netherlands, benefit from real integration, whereas the less integrated ones, France and Italy, do not. On top of this, they incur losses in export openness from monetary integration. In sharp, both, Germany and the Netherlands also benefit from the common currency and do not lose from ERM I. This highlights a certain structure: The main beneficiaries of both, monetary and real integration, are highly open economies. Our results reveal that the existence of a common long-run trend in European export openness is highly questionable. The distinct trends indicate divergence of inner-European export openness, which feeds back into economic growth and unemployment, thereby increasing inner-European disparities. The long-run aim of the European Union is to promote balanced economic and social development. The different trending behavior is not supportive or may be even counterproductive.

6 Conclusion

Our paper provides an answer to the following question: Which European integration event has most decisively shaped the trend of export openness of France, Italy, Germany and the Netherlands? This question derives its importance from the feedback

that export openness has on economic growth and unemployment in the individual countries and Europe as a whole. The long time span of the European integration process advocates a bird's eye view on different monetary and real integration steps. To this end, we apply a recently proposed test for trend breaks by Perron and Yabu (2009) and endogenously determine the single most decisive breakpoint.

We find strong evidence for trend breaks. A new and simple measure for their strength is introduced. We find four distinct country-specific most incisive milestones. Their effects differ in size and sign: Highly open economies gain from both, monetary and real integration. In sharp contrast, less open economies do not benefit from real integration and even suffer from monetary integration. These distinct trends in export openness may dangerously foster inner-European disparities in employment and growth.

The launch of the common currency is the most analyzed milestone. However, the Euro triggers the most decisive change only for France. Moreover, it is the least important event across the four major milestones. Therefore, the ERM I, the Maastricht Treaty and the merge of EFTA and EEC deserve at least as much research attention as the Euro introduction.

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