FORSCHUNGSINSTITUT FÜR EUROPAFRAGEN
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BUSINESS ADMINISTRATION VIENNA

Working Papers

IEF Working Paper Nr. 48

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What Has Determined the Rapid Post-War Growth of Intra-EU Trade?

February 2003

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Medieninhaber, Eigentümer Herausgeber und Verleger: Forschungsinstitut für Europafragen der Wirtschaftsuniversität Wien, Althanstraße 39—45, A—1090 Wien; Für den Inhalt verantwortlich: Univ.-Prof. Dr. Stefan Griller, Althanstraße 39—45, A—1090 Wien.

What Has Determined the Rapid Post-War Growth of Intra-EU Trade?

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February 2003

Abstract: Based on the gravity model by Baier and Bergstrand (2001), we use a static and dynamic panel data approach to estimate the relative contributions of income growth, income convergence, and the reductions in tariffs and trade costs to the growth of intra-EU trade over the period 1960 to 2000. The results suggest that income growth was the major force, accounting for approximately two third of total growth. Trade liberalization still had a sizeable effect, accounting de facto for the rest of growth, while income convergence played only a minor role. Reductions in trade costs had no significant effect on the growth of intra-EU trade. The results turn out as robust against several robustness checks and the use of alternative estimators.

JEL Classification: C23, F12, F14, F15
Keywords: European integration, gravity model, international trade

We wish to thank Gerhard Fink and Stefan Griller for a number of helpful comments.

Introduction

Over the period 1960 to 2000 intra-EU trade has grown by an impressive 1200 per cent in real terms (6.7 per cent per annum), compared with a more moderate 730 per cent growth of the EU countries' trade with the rest of the world (see Table A1 in the Appendix). This impressive growth performance indicates strong trade creating forces in Europe and offers a valuable source for investigating the determinants of the growth of trade in general and the role of European integration in particular.

Krugman's (1995) statement that the fundamental question of the determinants of the growth of trade has remained surprisingly disputed over a long time, triggered an ongoing debate on this issue. More recently, Baier and Bergstrand (2001) undertook a comprehensive investigation of the determinants of the growth of world trade. Estimating a cross-section gravity equation they conclude that income growth and tariff reductions were the major propelling forces. The reduction in trade costs still played some role, while income convergence was only of minor importance. An implicit conclusion is that the GATT/WTO-liberalization, accounting for a large part of the tariff reductions, was a propelling force of world trade. This (indirect) conclusion of the results was at least called into question by Rose (2002, p. 22), who does not find a significant effect of GATT/WTO membership (measured by zero-one dummies) in a gravity approach and concludes that „it is surprisingly hard to demonstrate convincingly that the GATT and the WTO have encouraged trade". In contrast he finds that "regional trade associations seem typically to have a much larger effect than the multilateral GATT/WTO system; nine of the ten RTAs have point estimates greater than .7 (all are statistically significant), indicating that trade at least doubles with membership" but that "Curiously, the outlier is the EEC/EC/EU." (Rose, 2002, Appendix, p. 10). Thus, Rose indirectly also challenges the trade creating role of European integration, although he admits himself the curiosity of this argument, leading him to the following qualification of his results: „One should not conclude the GATT and WTO have not increased trade (although I wish it was easier to see in the data). Rather, since common sense and conventional wisdom accord an important role to the GATT/WTO in creating trade, I prefer to view this negative result as an interesting mystery." (Rose, 2002, p. 22).

Against the background of these ambiguous results, in particular for the case of European integration, more empirical work on this issue seems warranted. The enormous growth of intra-EU trade, together with considerable variation across separate trade flows offers a potentially valuable source to re-examine these questions in a more comprehensive approach.

In particular, the availability of data over the comparably long period of 1960 to 2000 allows us to use a panel approach. Besides the obvious advantage of more observations and potentially more precise estimates, it allows us to control for country-specific and time-specific effects. Furthermore it enables us to conduct a more rigorous robustness analysis, including the extension to a dynamic specification and the consideration of potential endogeneity concerns. Finally, the focus on intra-EU trade allows us to asses the role of European integration, since the elimination of intra-EU tariffs is unambiguously linked to European integration.

The remainder of the paper is organized as follows. In section I we briefly review the model by Baier and Bergstrand (2001), on which our empirical analysis will be based. In section II we present the data used in the estimation of our empirical model in section III. In section IV we summarize the results and conclude.

I. Theoretical background and empirical model

The theoretical background for our study is provided by the model of Baier and Bergstrand (2001), which is a synthesis and generalization of previous work on the gravity equation. Accounting for expenditure constraints, emphasized by Anderson (1979), market structure, stressed by and Krugman and Helpman (1985), and distribution costs, emphasized by Bergstrand (1985) it provides the adequate formal framework for our research question.

We can only briefly sketch the essential features of the model here: On the demand side, a representative consumer in country i maximizes a constant-elasticity-of-substitution (CES) utility function over all available varieties of goods in every country subject to a budget constraint, where prices of the imported goods reflect iceberg transportation costs and ad valorem tariffs. The solution to this utility maximization yields an import demand function for the product of a representative firm in country j. This representative firm faces monopolistic competition and maximizes its profits subject to two technology constraints: First, the production of goods has fixed and constant marginal costs. Second, the presence of distribution costs lets firms face each potential market's supply as imperfect substitutes, which is reflected formally in a constant-elasticity-of-transformation (CET) function (see Powell and Gruen, 1968). This implies that firms incur cost when they substitute output between foreign markets due to tailoring the product, marketing and distribution to the particular market. The resulting long-run (supply-side) equilibrium is characterized by two conditions: i) Prices are a mark-up over marginal costs, which depends on the elasticity of substitution in consumption. ii) Under monopolistic competition, firms earn zero profits. From these conditions, bilateral export supply functions of the representative firm can be derived and set equal to the respective import demand functions. The model is closed by the assumption of full employment and a given factor endowment (labour), which determines the available varieties of goods. At the bottom line, the model ultimately yields an equilibrium solution for the bilateral exports from country i to j (X_i^j), which can be written in log-differences as follows:

$$\begin{aligned}\Delta \ln X_i^j &= b_1 \Delta \ln(Y_i + Y_j) + b_2 \Delta \ln(s_i s_j) + b_3 \Delta \ln(1+T_i^j) + b_4 \Delta \ln(1+TC_i^j) + \\ &+ b_5 \Delta \ln Y_j + b_6 \Delta \ln(P_i^F / P_j^C)(P_j^F / P_j^C)^{-1/s}\end{aligned} \tag{1}$$

Y_i (Y_j) is real *GDP* of country *i (j)*; $s_i s_j = Y_i Y_j/(Y_i+Y_j)^2$ is an expression for income convergence, whose trade stimulating role was stressed by Helpman (1987). According to the theoretical model, the parameters b_1 and b_2 should equal two and one, respectively. The interpretation of the next two variables is straightforward: both tariffs for exports from country i to country j (T_i^j) and trade costs for exports from i to j (TC_i^j) should yield a negative coefficient. The coefficient of ΔY_j depends on the elasticity of transformation of output across markets and should equal zero if production is perfectly substitutable between home and foreign markets; in the case of a finite elasticity of substitution, a negative sign would be expected. The last variable is a ratio of two Dixit-Stiglitz price indices[1]; according to the theoretical model, a negative coefficient is expected. As opposed to all other variables, however, these price indices are not observable and can only be proxied crudely, using the ratio of the countries' GDP deflators. Equation (1) will be the point of departure for our estimation. Before presenting the results of our estimations we give a brief description of the data.

II. Data

As opposed to the cross-section approach by Baier and Bergstrand we use a panel to estimate the gravity equation described above. This has some obvious advantages: First, we have more observations and potentially less multicollinearity, which should yield more precise estimates. Second, it allows to control for cross-section-specific time-invariant effects as well as time-

[1] P_j^C is an index of landed prices in country j of products from all markets (resulting from the CES utility function), and P_i^F is an index of the firm i's prices (resulting from the CET function).

specific cross-section-invariant effects. Third, it extends easily to a dynamic model and allows us to address potential endogeneity problems of the right hand side variables. Table 1 gives an overview of the definition of the variables and the data used in the estimation.

Table 1 – Description of variables and data

Variable	Definition / Data sources
$X_{i,t}^{j}$	real exports from country i to country j in million \$US (1990 prices, 1990 PPPs), taken from IMF: Direction of Trade Statistics and converted into real figures using the implied deflators of the position “imports (exports) of goods and services” from the OECD: *National Accounts*. 1950-1960: Series for deflator supplemented, using the relation between the GDP and the trade deflators for the period 1960-2000.
$GDP_{i,t}$	real gross domestic product of country i in \$US (1990 prices, 1990 PPPs), taken from OECD: *National Accounts*. Data for 50-60 supplemented accorrding to Maddision (1995).
$P_{i,t}$	implicit deflator of gross domestic product from real and nominal GDP series of OECD: *National Accounts*. Data for 50-60 supplemented accorrding to Maddision (1995).
$REER_{i,t}$	index of real effective exchange rate (1990 = 1); constructed as $REER_{i,t} = \sum_{k=1}^{16} w_{ik} ER_{ik} \frac{\sum_{k=1}^{16} w_{ik} CPI_k}{CPI_i}$ w_{ij} = share of exports to country k in total exports of country i, ER_{ik} = exchange rate from country i against country k, CPI_i = consumer price index (taken from IFS and transformed so that 1990 = 100), k = 1, . . . , 16: EU member states, JP, and rest of world (\$-exchange rate).
$T_{i,t}^{j}$	tariff (of country j) for exports from country i to country j (as fraction of one); time series constructed using the country-specific external tariffs in 1950 according to Breuss (1983) and accounting for the average tariff reductions following the GATT rounds (Source: WTO, 1995) and tariff changes as a result of European Integration, i.e. adoption of harmonized external tariff and elimination of intra-EC tariffs after EC (EU) accesion respectively in course of the free trade agreements between EC and EFTA in the 70s. See Table 2 for details on the assumed timing and size of the tariff reductions.
$TC_{i,t}$	multilateral trade costs of country i (as fraction of one), 100(CIF/FOB-ratio-1); CIF/FOB-ratio taken from International Financial Statistics Yearbook (1995). Data available for 1965-1994. Data for 1950 to 1964 were set to the 1965-value, data for 1995-2000 to the 1994-value, which seems justified, given their little variation over the observed period.

i(*j*) = 1, . . ., 14 (country index): Austria (AT), Belgium+Luxembourg (BLX), Denmark (DK), (West)-Germany (DE), Finland, (FI), FR (France), Greece (GR), Ireland (IE), Italy (IT), Netherlands (NL), Portugal (PT), Spain (ES), Sweden (SE), United Kingdom (UK). t = time index: 1950(60)-2000. Data were converted into \$US using 1990 PPPs from the OECD (EKS method). All data were taken from the database of the Austrian Institute of Economic Research (WIFO: http://www.wifo.ac.at/).

Our sample comprises the actual fifteen Member States of the European Union; since trade data for Belgium and Luxembourg are only available as aggregate, fourteen countries remain. Thus, the cross section dimension includes 182 bilateral trade flows, which in sum constitute total intra-EU trade.

Table 2 – Economic integration of EU Member States (1950-2000)

European integration	GATT-liberalization
1944: Benelux Customs Union (BE, LU, NL) a) elimination of tariffs between BE, LU, NL, b) harmonization of external tariff (1950: 9%), (assumed) implementation: 1945-1950.	1950: Individual External Tariffs (%) AT(20), BE(9), DE(16), DK(5), ES*(24), FI(13.5), FR(19), GR*(24), IE*(17), IT(24), NL(9), PT*(24), SE(6), UK(17).
1957: EC-6 (BE, DE, IT, NL, LU, FR): Customs Union a) elimination of intra-EC-6 tariffs, b) harmonization of external tariff (1968: 16.8%), implementation: 1957-1968.	
1960: EFTA-7 (AT, CH, DK, NO, PT, SE, UK) elimination of intra-EFTA-7 tariffs, (1961: free trade agreement FI-EFTA), implementation: 1960-1967.	1964-1967: Kennedy-Round average (relative) tariff reductions: 47%, assumed implementation: 1968-1972.
1973: First EC-enlargement (DK, IE, UK) → EC-9 a) elimination of tariffs between DK, IE, UK and EC-6, b) harmonization of external tariff, implementation: 1973-1978.	
1973: Free trade agreements between EFTA-6 and EC-9 elimination of tariffs between EFTA-6 members (AT, CH, IS, NO, PT, SE) + FI and EC-9, implementation: 1973-1978.	1973-1979: Tokyo-Round average (relative) tariff reductions: 30%, assumed implementation: 1980-1985.
1981: Second EC-enlargement (GR) → EC-10 a) harmonization of external tariff, b) elimination of tariffs between GR and EC-9, implementation: 1981-1985.	
1986: Third EC-enlargement (ES, PT) → EC-12 a) harmonization of external tariff, b) elimination of tariffs between ES and EC-10, implementation: 1986-1995.	1986-1993: Uruguay-Round average (relative) tariff reductions: 40%, assumed implementation: 1994-1999.
1993: Single Market (EU-12) 4 freedoms + flanking measures (common policies), instantaneous implementation assumed.	
1994: European Economic Area (EEA): partial implementation of four freedoms between EU-12 and EFTA-7' except CH (AT, FI, IS, LI, NO, SE), instantaneous implementation assumed.	
1995: Fourth EC-enlargement (AT, FI, SE) → EU-15 a) harmonization of external tariff, b) participation in Common Market, instantaneous implementation assumed.	

Monetary integration (1978: EMS, 1999: EMU) is not considered here. – Data on tariff levels, timing and structure of tariff reductions and tariff harmonization taken from Breuss (1983), El-Agraa (2001), WTO (1995). – * indicates missing values that were completed according to the relative position of a country at a later point of time, for which data were available.

The time period considered ranges from 1960 to 2000, including all major steps from European integration since the Customs Union. In order to smooth out cyclical fluctuations and short-run shocks, we use overlapping, five year periods (1960-1965, 1965-1970, . . . , 1995-2000). We have also data for the period from 1950 to 1960, but restrict our main analysis to the period as of 1960 for two reasons: First, data for 1950 to 1960 are less reliable with a number of missing observations, which had to be interpolated. Second, we also want to reserve two lagged observations, which we will need when running two-stages least squares with lags as instruments. Summing up, we have a sample of 182 cross-section units, each of which with eight observations, yielding us a total sample size of 1456 observations.

Note that only some of the variables are really trade-flow-specific, while some of them are only cross-section specific. Also note that *TC* and *REER* refer to economic relationships with the whole world, not only with the EU; due to the dominant share of EU-relationship, however, this slight deviation from theory seems admissible.

III. Estimation results

The empirical counterpart of the theoretical model above is given by

$$\begin{aligned}\Delta \ln X_{it}^{j} &= b_1 \Delta \ln(Y_{it} + Y_{jt}) + b_2 \Delta \ln(s_{it} s_{jt}) + b_3 \Delta \ln(1 + T_{it}^{j}) + b_4 \Delta \ln(1 + TC_{it}) + \\ &+ b_5 \Delta \ln Y_{jt} + b_6 \Delta \ln(P_i \; / P_j \;)_t + m_i + h_t + u_{it}^{j}\end{aligned} \qquad (2)$$

m_i = cross-section-specific fixed effect, h_t = time-specific effect, $u_{i,t}^{j}$ = error term (IID), *i* = cross-section unit 1, . . ., 182 (export flow for country *i* to country *j*), *t* = time period (t = 1, . . ., 8 (1960-2000).

Our estimation exercise proceeds as follows: First, we will present the estimation results for the static model as given by (2), using the standard least squares dummy variable estimator. Then we will extend model (2) to a dynamic specification, which offers a convenient framework to address endogeneity concerns, too. Since dynamic panels require other estimators than static panels, we will briefly discuss some econometric issues before turning to the estimation of the dynamic model. Finally, we will use our preferred models to simulate the growth of intra-EU trade in order to asses the relative importance of the respective variables.

1. Results for static models

Results from a least square dummy variable (LSDV) estimation of model (2) are given in Table 3.

Table 3 – Results of estimation: static models (2), (3) and variants, N = 182, T = 1-8 (1960-2000)

dependent variable: $\Delta \ln X_{it}$

	(a)	(b)	(c)	(d)
constant	0.087	0.087	0.003	0.002
$\Delta \ln (Y_i + Y_j)$	0.384	1.843***	1.874***	1.876***
	(0.99)	(6.15)	(6.28)	(6.28)
$\Delta \ln s_{ij}$	0.025	0.748***	0.703***	0.720***
	(0.10)	(3.87)	(3.63)	(3.76)
$\Delta \ln (1+t_i^j)$	-1.946***	-1.967***	-1.915***	-1.919***
	(-5.54)	(-5.53)	(-5.43)	(-5.44)
$\Delta \ln (1+tc_i^j)$	0.022	0.035	0.040	
	(0.66)	(0.98)	(1.14)	
$\Delta \ln Y_j$	1.458***			
	(6.00)			
$\Delta \ln RP_i$	0.170***	0.093*		
	(3.14)	(1.73)		
$\Delta \ln REER_i$			0.307***	0.307***
			(3.96)	(3.96)
R^2	0.260	0.237	0.243	0.243
Adj R^2	0.146	0.121	0.128	0.128
SEE	0.321	0.325	0.324	0.324

*, **, *** indicate significance at the ten, five and one per cent level. – t-values in parentheses, based on White heteroscedasticity-consistent standard errors. – All models estimated using the least squares dummy variable (LSDV) estimator, based on mean centred variables. – Static models were estimated using Eviews 4.0.

A look at the coefficients of $\Delta \ln(Y_{it}+Y_{jt})$ and ΔY_{jt} in columns (a) and (b) suggests a multicollinearity problem with these two variables; indeed, their correlation amounts to 0.80. Thus, $\Delta \ln Y_{\mathrm{j}}$ can be assumed to measure not the elasticity of transformation, which would suggest a coefficient smaller than or equal to zero, but also an income effect. After excluding the variable $\Delta \ln Y_j$ (see column (b)), the new coefficient of $\Delta \ln(Y_i + Y_j)$ corresponds approximately to the sum of its value and the coefficient of $\Delta \ln Y_j$ in column (a), confirming this presumption. Since the ratio of the *GDP* deflators, used as a proxy for a Dixit-Stiglitz price index, is only weakly significant and takes the wrong sign we decided to exclude this variable from the regression, too (see column (c)). At the risk of slightly departing from the theoretical model we introduce a new variable, the real effective exchange rate ($\Delta \ln REER$) in order to control for changes in competitiveness; an increase in *REER* is associated with real effective depreciation. Changes in the real effective exchange rate, however, turn out as de facto orthogonal to the other variables and thus hardly affect their coefficients. The results for this final model are given in column (d). All variables are significant at the one per cent level an show the expected sign. Furthermore the coefficients $\Delta\ln(Y_i + Y_j)$ and $\Delta\ln(s_i s_j)$ take values, which are close to the predictions of the theoretical model. The joint hypothesis that $b_1 = 2$ and that $b_2 = 1$ cannot be rejected (p-value of F-test: 0.40).

In a next step we check the sensitivity of the results with respect to changes in the estimation period. We estimate the model in column (d) for all possible sub-periods of the time 1960 to 2000 with at least 4 observations (i.e. 20 years), yielding us 14 different models. The detailed results are given in the Appendix (Table A2). The main results of the sensitivity analysis can be summarized as follows:

- The variable $\Delta\ln(Y_i+Y_j)$ always enters significantly at the one per cent level. While the estimated coefficient range from 1.57 to 2.69, the hypotheses that the coefficient of $\Delta\ln(Y_{it}+Y_{jt})$ equals two as postulated theoretically can only be rejected in two of the 14 models (5 and 10 per cent level).
- The variable $\Delta\ln(s_i s_j)$ is also significant in all but two models, at least at the 10 per cent level. The hypothesis of a unity coefficient as expected theoretically can be rejected only in two further of the 12 remaining regressions.
- The change in real effective exchange rate index is significantly different from zero in all but three regressions. Its exclusion hardly affects the other coefficients. Due to its small (relative) variation, it plays only a minor role in the explanation of the growth of trade, as will be seen more clearly below.
- The reduction in tariffs $\Delta\ln(1+T_i^j)$ is highly significant in all models except in that for the period 1980-2000. This is plausible, since most of the tariffs have been reduced until the end of the 70s within the EC and between the EC and the EFTA-countries (see Table 2) so that this variable shows only little variation in the period 1960-2000. The significant coefficients range from -1.1 to -2.2; the highest value is obtained for the period 1960-1980, when most of the tariff reductions were implemented (customs union 1957-1968, first EC enlargement in 1973 by UK, DK; IE; establishment of free trade between European Free Trade Area, FI and the EC (1973-1977). In the period from 1960 to 1980 only income growth and trade liberalization appear to have driven the growth of intra EC-trade. None of the other variables enters significantly in this period.

- Trade costs and relative prices (proxy for Dixit-Stiglitz price indices) are insignificant in more than two third of the models (b). Moreover, their coefficients are close to zero and generally show the wrong (positive) sign. This justifies their exclusion for the model.

Thus our preferred static model turns out to be

$$\Delta \ln X_{it}^{j} = b_1 \Delta \ln(Y_{it} + Y_{jt}) + b_2 \Delta \ln(s_{it} s_{jt}) + b_3 \Delta \ln(1 + T_{it}^{j}) + \\ + b_7 \Delta \ln REER_{it} + m_i + h_t + u_{it}^{j} \tag{3}$$

So far, our sensitivity analysis shows a very robust relationship between growth of trade and income growth, income convergence, tariff reductions and changes in the real effective exchange rate, with coefficients close to the predictions of the theoretical model.

2. Results for dynamic models

Despite the remarkable robustness of the results obtained so far, there are two further concerns: a) a possible dynamic mis-specification and b) the potential endogeneity of the right hand side regressors. Although the theoretical model does not suggest a dynamic specification, Baier and Bergstrand, in their cross section estimation, include the initial level of exports $\ln X_{i,t-1}$ in order to control for deviations from the long-run equilibrium level. In our case this yields the following *dynamic variant* of the preferred static model (3)

$$\Delta \ln X_{i,t}^{j} = b_1 \Delta \ln(Y_{i,t} + Y_{j,t}) + b_2 \Delta \ln(s_{i,t} s_{j,t}) + b_3 \Delta \ln(1 + T_{i,t}^{j}) + \\ + b_7 \Delta \ln(REER_{i,t}) + g \ln X_{i,t-1}^{j} + m_i + h_t + u_{it}^{j}$$

which is, of course, equivalent to the following conventionally formulated dynamic panel in levels

$$\ln X_{i,t}^{j} = b_1 \Delta \ln(Y_{i,t} + Y_{j,t}) + b_2 \Delta \ln(s_{i,t} s_{j,t}) + b_3 \Delta \ln(1 + T_{i,t}^{j}) \\ + b_7 \Delta \ln(REER_{i,t}) + (1+g) \ln X_{i,t-1}^{j} + m_i + h_t + u_{it}^{j} \tag{4}$$

It is well-known that the LSDV-estimator is biased in dynamic panels (Nickell, 1981). Although this bias tends to zero as the time dimension of the panel approaches infinity, it cannot be ignored in our particular panel with a large number of cross-section units and a short time range. In contrast to the LSDV estimates of (1+*g*), which is biased downward as shown by Nickell (1981), the pooled OLS estimator of (4) with a common intercept will produce an upward biased estimate of (1+*g*) in the presence of fixed effects (Hsiao, 1986). Thus, the LSDV and the pooled OLS estimator provide a range for a plausible parameter estimate.

In order to obtain consistent parameter estimates in dynamic panels, instrumental variable procedures have been suggested in the literature. Thereby, the fixed effects m_i are eliminated using first differences; this however, induces correlation between the differenced error term ($u_{i,t}$–$u_{i,t-1}$) and the lagged difference of the dependent variable ($y_{i,t-1}$–$y_{i,t-2}$). Consequently, an instrumental variable estimation is performed, where the lagged level ($y_{i,t-2}$) (or the lagged difference ($y_{i,t-2}$ – $y_{i,t-2}$)) can be used as instruments (Anderson and Hsiao, 1981). More recently, Arellano and Bond (1991) argued that if $y_{i,t-2}$ is a valid instrument, further lags ($y_{i,t-3}$, $y_{i,t-4}$, etc.) are valid, too and that by exploiting all moment restriction (E[$y_{i,t-s}$($u_{i,t}$–$u_{i,t-1}$)] = 0, s ≥ 2) more efficient parameter estimates can be obtained (First-Differences-GMM estimator). Validity of instruments requires the absence of second order serial correlation in the residuals; overall validity of instruments can also be tested using a Sargan test of overidentifying restrictions (see Arellano and Bond, 1991, for the more details on the test). The First-Differences-GMM estimator was criticized recently by Blundell and Bond (1998), who argue that levels may be valid, but poor instruments for first differences, in particular if data is highly persistent. An indication of such a poor quality may be that the First-Differences-GMM estimate of (1+g) is close to its (downward-biased) LSDV estimate (see Bond et al., 2001). Blundell and Bond (1998) suggest a system GMM estimator**,** which supplements the equations in first differences with equations in levels, where in the levels equations, lagged differences $\Delta y_{i,t-1}$ are used as instruments for $y_{i,t-1}$. This is based on the assumption that $E(m_i \Delta y_{i2}) = 0$ for i = 1,...,N, which (together with the standard assumptions for the first differences estimator) yields the additional moment conditions $E(u_{it} \Delta y_{i,t-1}) = 0$ for i = 1,..., N and t = 3, 4, ..., T, $u_{i,t} = m_i + u_{i,t}$.[2] Again, the validity of instruments can be checked by the Sargan test and the validity of the additional instruments by the Difference Sargan test. Using Monte Carlo studies, Blundell and Bond (1998) showed for the AR(1) model that the finite sample bias of the difference GMM estimator can be reduced dramatically with the system GMM estimator. Similar results were obtained for a model with additional right-hand side variables by Blundell et al. (2000). Other estimators were also suggested (see Baltagi, 2001), none of which can claim to be superior in all cases. We will thus restrict our attention to the first differences and the GMM estimator.

[2] Note that this requires the first moment of y_{it} to be stationary. Including time dummies in the estimation is equivalent to transforming the series into deviations from time means. Thus any pattern in the time means is consistent with a constant mean of the transformed series of each country (Bond et al. 2001).

The GMM framework, either in first differences or in a system framework, offers a convenient way to address also the problem of measurement error or potentially endogenous or predetermined right hand side variables. Since causality may also run from trade to growth, the variables $(Y_i + Y_j)$ as well as $s_i s_j$ in equation (3) or (4) are obvious candidates. For the sake of simplicity let's denote $\Delta\ln(Y_i+Y_j)$ and $\Delta\ln(s_i s_j)$ by Z_i. In our context this means that we would expect Z_{it} to be endogenous ($E(Z_{it}u_{is}) \neq 0$ for $i = 1, \ldots N$ and $s \leq t$), allowing for both contemporaneous correlation between the current shock u_{it} and Z_{it} as well as feedbacks from past shocks $u_{i,t-s}$ to the current value of Z_{it}. Alternatively, Z_{it} might be considered as predetermined, ruling out contemporaneous correlation between Z_{it} and u_{it}, i.e. $E(Z_{it}u_{is}) \neq 0$ for $i = 1, \ldots N$ and $s < t$. If Z_i is considered endogenous (predetermined), again lagged values of the variable level Z_{it} dated t-2 (t-1) and earlier can be used as instruments.

Table (4) shows the estimation results for models (4). Thereby we used different estimators and different assumption concerning the nature of $\Delta\ln(Y_i+Y_j)$ and $\Delta\ln(s_i s_j)$ in order to check the sensitivity of the results with respect to the estimation method. Columns (a) and (b) show the results for the estimation of model (4) using the pooled OLS and the LDSV estimator, respectively. As expected the estimate of $(1+g)$ is clearly lower in (b). Columns (c) and (d) shows the results using the GMM-estimator in first differences by Arellano and Bond (1991), where in column (d) the variables $\Delta\ln(Y_i + Y_j)$ as well as $\Delta\ln(s_i s_j)$ are treated as predetermined. Treating them as endogenous in the estimation, which differs from (d) only in the use of less instruments yielded similar results, but did not further improve the Sargan test for the validity of instruments; in fact it deteriorates if endogeneity is assumed. Thus we regard the assumption of predeterminedness as sufficient, but hasten to add that the size of the parameter estimates hardly differs, when endogeneity is allowed for.

Table 4 – Results of estimation: dynamic model (4) N = 182, T = 1-8 (1960-2000)

dependent variable: $\ln X_{it}$	(a)	(b)	(c)	(d)	(e)	(f)
	OLS	*LSDV*	*GMM-FD*1	*GMM-FD*2	*GMM-SYS*1	*GMM-SYS*2
constant	0.139*	2.506	0.210	0.206	-0.021	0.010
$\Delta\ln(Y_i+Y_j)$	1.929***	1.622***	2.378***	1.850***	1.945***	1.957***
	(6.75)	(5.96)	(7.08)	(7.94)	(3.50)	(7.50)
$\Delta\ln s_{ij}$	0.899***	1.424***	1.348***	0.938***	1.083***	0.809***
	(5.79)	(7.92)	(3.24)	(4.87)	(3.06)	(4.37)
$\Delta\ln(1+t_i^j)$	-1.946***	-1.317***	-1.409**	-1.686***	-1.372**	-1.739***
	(-5.31)	(-4.20)	(-2.34)	(-3.47)	(-2.47)	(-4.15)
$\Delta\ln REER_i$	0.318***	0.277***	0.263***	0.233**	0.327***	0.263***
	(4.08)	(3.92)	(2.59)	(2.47)	(3.05)	(3.18)
$\ln X_{i,t-1}$	0.979***	0.667***	0.942***	0.903***	0.964***	0.956***
	(150.73)	(26.09)	(21.23)	(32.72)	(13.42)	(52.43)
m_1			-6.45***	-6.53***	-6.73***	-6.52***
m_2			1.18	1.28	1.34	1.20
Sargan			59.64***	147.30*	72.72***	165.93
			(35)	(121)	(43)	(145)
Diff.-Sargan					13.80	18.63
					(8)	(24)
R^2	0.972	0.979	0.970	0.972	0.974	0.975
Adj R^2	0.971	0.976	0.965	0.967	0.970	0.971
SEE	0.318	0.294	0.353	0.342	0.328	0.325

[a] pooled OLS estimation with common intercept. – [b] LSDV-estimator, based on mean centred variables. – c) First Differences GMM-estimator by Arellano and Bond (1991), two step estimates. – [d] as (c), but with $\Delta\ln(Y_i+Y_j)$ and $\Delta\ln s_{ij}$ treated as predetermined variables. – [e] System GMM estimator in first differences and levels by Blundell and Bond (1998), two step estimates. – [f] as (e), but with $\Delta\ln(Y_i+Y_j)$ and $\Delta\ln s_{ij}$ treated as predetermined variables. – As recommended by Arellano and Bond, inference is based on (robust) one-step estimates. – Constant not directly comparable. – m_1 (m_2): test for first (second) order serial correlation. – Sargan-Test: Values in parentheses are degrees of freedom; chi-square distributed. – Models (a), (b) were estimated using Eviews 4.0, models (c)-(f) using the DPD98 software for Gauss by Arellano and Bond (1998).

Columns (e) and (f) show the estimates using the system-GMM estimator by Blundell and Bond (1998), again assuming predeterminedness in column (f). Note that the coefficients of $\ln X_{i,t-1}$ show the expected relationships; Both GMM estimators lie in the bound provided by the LSDV and the OLS estimate, the system estimator yielding a slightly higher coefficient. Since the GMM estimate of $(1+g)$ in first differences differs hardly from the system estimate, we appears to be no weak instruments problem in our case. A more relevant point is, whether exogeneity or predeterminedness is assumed, since the coefficients of $\Delta\ln(Y_i + Y_j)$ and

$\Delta\ln(s_i s_j)$ change substantially. The highly significant Sargan test in columns (c) and (e), where exogeneity is assumed, clearly gives rise to using instruments for $\Delta\ln(Y_i + Y_j)$ and $\Delta\ln(s_i s_j)$, given the absence of second-order serial correlation. Indeed, the Sargan test can be improved substantially by assuming predeterminedness of $\Delta\ln(Y_i + Y_j)$ and $\Delta\ln(s_i s_j)$, although it remains significant at the 10 per cent level in the case of the first-differences estimator. As already outlined above the assumption of endogeneity did not improve the Sargan test. Thus, our preferred dynamic estimates are given by column (f), using the system GMM estimator with $\Delta\ln(Y_i + Y_j)$ and $\Delta\ln(s_i s_j)$ treated as predetermined. Since the instruments used, contain those of column (d) as a subset, some doubts remain on the failure of the Sargan test to reject the null of valid instruments. It is interesting to note that the long-run coefficients of our preferred dynamic estimates (but also that of column (d)) differ only slightly from the results of our static model. Thus, we may conclude that our results do not differ substantially, when a dynamic structure of the models is controlled for.

3. Simulations

We go on to simulate the growth of intra EU-trade over the period 1960 to 2000 in order to identify the importance of the respective variables. Table 5 shows the results, when our preferred static model is used. Table 6 shows the results from using the preferred dynamic dynamic models for the simulation. We start from a scenario with purely exogenous growth of trade, where only fixed effects, the time specific effects and the residuals (and the lagged endogenous variable) are included, i.e. where b_1, b_2, b_3 and b_7 have been set to zero in equation (3) and (4), respectively. We then set the paramters of $\Delta\ln(Y_i + Y_j)$, $\Delta\ln(s_i s_j)$, $\Delta\ln(1+T_{ij})$ and $\Delta\ln REER_i$ to their estimate value, so that we end up with simulating the actual scenario.

Tables 5 and 6 show the simulation results, where the trade flows have been aggregated at the country level. Though the static and dynamic results differ somewhat in detail, the main conclusions are consistent: In terms of the aggregate EU, income growth was the driving force of intra EU trade, accounting for some 70 per cent of the growth of intra-EU trade over the period 1960 to 2000. European integration in terms of tariff reductions also played a prominent role in creating intra EU-trade, accounting for some 19 to 26 per cent of its total growth. Income convergence and real effective exchange rate changes, though statistically significant, played only a minor role in shaping the post-war growth of intra-EU trade. We do not go into the detail of the country specific results, which should not be overstressed in panel estimations generally.

Table 5 – Results of simulation (1960-2000) – static model (Table 3, column (d))

	X_{1960} [1]	X_{2000} [1]	$g_{1960-2000}$		relative contributions (% of total growth)					
	intra EU		%	% p.a.	exog.[3]	GDP	SIJ	T	REER	Σ
AT	2924	40807	1296	6.81	8.36	66.96	5.04	27.66	-8.02	100
BLX	11239	135928	1109	6.43	8.53	70.34	1.69	22.90	-3.46	100
DE	28448	285099	902	5.93	9.02	65.99	4.08	23.68	-2.76	100
DK	3447	21814	533	4.72	12.54	90.48	0.34	28.65	-32.01	100
ES	3347	90042	2591	8.58	6.83	79.77	7.36	34.95	-28.90	100
FI	2297	22117	863	5.83	8.86	70.38	7.19	23.80	-10.24	100
FR	10463	184784	1666	7.44	7.97	69.62	1.97	24.69	-4.25	100
GR	402	6378	1486	7.15	6.03	55.05	13.42	26.27	-0.77	100
IE	1289	48585	3669	9.50	5.14	43.67	31.49	26.45	-6.76	100
IT	7482	136185	1720	7.52	8.01	71.41	2.12	25.22	-6.76	100
PT	11374	180291	1485	7.15	9.72	80.47	4.37	26.17	-20.73	100
NL	2538	26477	943	6.04	7.23	73.95	18.09	35.29	-34.55	100
SE	4765	38653	711	5.37	10.31	72.05	-1.59	23.26	-4.04	100
UK	12773	140445	1000	6.18	11.24	65.59	10.40	27.89	-15.11	100
EU-15	102786	1357605	1221	6.66	8.79	70.23	5.42	25.93	-10.37	100

[1] Intra-EU exports in Mill $US, 1990 prices, 1990 PPPs. – [2] export growth from 1960 to 2000, total and per annum. – [3] exogenous: only fixed effects, time effects and residuals included.

Our main conclusions for the EU are comparable with that by Baier and Bergstrand (2001) for world trade. However, it is interesting to note that the trade costs reductions have no explanatory power in our model, though they turn out significant in the Baier-Bergstrand model. Accounting for only 8 per cent of the growth of world trade, however, Baier and Bergstrand downplay their importance. This difference in the results is plausible since Baier and Bergstrand have more distant countries in their sample which includes 16 OECD members (most EU countries, the USA, JP and Canada). Since intra EU trade flows constitute a large share of their sample, this suggests that their significant results for trade costs can be traced back to only a small share of their observations; thus, their (average) coefficient might substantially underestimate the role of trade costs in promoting growth of trade between more distant countries such as between the EU Members States, the USA and Japan.

Table 6 – Results of simulation (1960-2000) – dynamic model (Table 4, column (f))

	X_{1960} [1)]	X_{2000} [1)]	$g_{1960\text{-}2000}$		relative contributions (% of total growth)					
	intra EU		%	% p.a.	exog.[3)]	GDP	SIJ	T	REER	Σ
AT	2924	40807	1296	6.81	11.90	67.91	5.46	19.59	-4.86	100
BLX	11239	135928	1109	6.43	11.89	71.03	1.19	16.74	-0.85	100
DE	28448	285099	902	5.93	12.26	66.02	4.22	17.44	0.06	100
DK	3447	21814	533	4.72	15.34	86.59	-1.36	19.28	-19.85	100
ES	3347	90042	2591	8.58	8.54	77.56	4.85	26.77	-17.73	100
FI	2297	22117	863	5.83	11.98	73.56	4.17	15.10	-4.81	100
FR	10463	184784	1666	7.44	10.91	68.73	2.76	18.52	-0.94	100
GR	402	6378	1486	7.15	8.74	59.29	11.15	19.91	0.90	100
IE	1289	48585	3669	9.50	6.31	47.18	31.20	18.24	-2.93	100
IT	7482	136185	1720	7.52	10.95	70.73	2.66	18.59	-2.93	100
PT	11374	180291	1485	7.15	12.63	77.12	4.19	18.20	-12.15	100
NL	2538	26477	943	6.04	9.27	73.91	13.97	26.14	-23.29	100
SE	4765	38653	711	5.37	13.56	71.71	-1.34	16.57	-0.50	100
UK	12773	140445	1000	6.18	14.52	65.38	9.66	20.64	-10.20	100
EU-15	102786	1357605	1221	6.66	11.26	68.59	6.44	18.94	-5.22	100

[1)] Intra-EU exports in Mill $US, 1990 prices, 1990 PPPs. – [2)] export growth from 1960 to 2000, total and per annum. – [3)] exogenous: only fixed effects, time effects, residuals and lagged endogenous variable included.

The strong role of tariff reductions both in the Baier and Bergstrand (2001) and in our estimation contradict the results by Rose (2002), at least if one is willing to assume that the largest share of the tariff reductions is due to global economic integration (GATT), or – as more relevant in our context – regional European integration. Rose's negative results that appear to him as a mystery may have such a simple explanation as an improper measurement of GATT or European integration by zero-one dummies.

IV. Conclusions

What has determined the rapid growth of intra-EU trade over the period 1960-2000? Based on the gravity model by Baier and Bergstrand, we use a panel data approach to estimate the relative contributions of income growth, income convergence, and the reduction of tariffs and trade costs to the growth of intra EU trade. Our results indicate that the major force was income growth, accounting for 70 per cent. European integration and GATT liberalization, reflected in the reduction of tariffs also played a substantial trade creating role, accounting for approximately one quarter of the growth of intra EU trade. Increased income similarity had a positive but little effect, while the real effective appreciation of most countries slightly impeded the growth of trade. The reduction in trade cost played no role. These results turn out robust against various robustness checks such as changes in the estimation period, a dynamic re-specification, or controlling for the potential endogeneity of right hand side variables.

Rose's negative results for the trade encouraging role of GATT/WTO and the EU, interpreted as an "interesting mystery", may have such a simple explanation as an improper measurement. The conclusion arising from the significant results for the tariff reductions is straightforward: If the reduction of tariffs is a success of GATT/WTO, then the GATT/WTO has also significantly created trade; if European integration is assumed to be responsible for the elimination of intra-EU tariffs, it has also contributed significantly to the growth of intra-EU trade.

The generalized gravity-equation by Baier-Bergstrand (2001) turns our as fairly complete model, which survives several robustness checks in the estimation. Furthermore the estimates of the parameters which are central to the model are close to the theoretical predictions. The role of the real effective exchange rate in the estimation provides a rationale for incorporating real exchange rates into the model, though in the light of their little contribution to the growth of trade their omission seems justifiable.

Our insignificant result for the reductions in trade costs contrasts with other studies whose samples contain more distant countries. The role of trade costs as a determinant of the growth of world trade thus still deserves attention in future empirical studies, since their contribution to the growth of trade between distant countries like the EU Members and the US may have been substantial.

Finally, a further interesting direction for future research is the question, whether there are asymmetric gains from integration, as postulated by some theoretical models: An example is the model by Casella (1996), which postulates that the gains from enlarging a trade bloc fall disproportionately on its small Member States. Since these models generally assume economies of scale, more disaggregated sectoral data may be informative on whether such asymmetric gains exist in certain types of industries which exhibit increasing returns. Thus, the issue of potential asymmetries in the gains from integration (e.g. with respect to country size) remains an interesting and challenging line of both theoretical and empirical research.

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Appendix: Table A1 – Trade of EU Member States: Exports in 1950 (1960) and growth over the period 1950 (1960) - 2000

	trade 1950			trade 1960			trade 2000			cumulative growth 1950-2000 (%)			cumulative growth 1960-2000 (%)		
	intra-EC	RoW	total	intra-EC	RoW	total	intra-EC	RoW	total	intra-EC	RoW	total	intra-EC	RoW	total
AT	862	845	1708	2924	1988	4912	40807	25324	66131	4632	2896	3772	1296	1174	1246
BE	5398	3630	9028	11239	5900	17139	135928	50596	186524	2418	1294	1966	1109	758	988
DE	8276	5471	13747	28448	25401	53849	285099	222186	507285	3345	3961	3590	902	775	842
DK	2896	676	3572	3447	1750	5197	21814	12744	34559	653	1784	867	533	628	565
ES	1402	2272	3670	3347	2074	5421	90042	41670	131712	6321	1734	3489	2591	1909	2330
FI	1182	941	2114	2297	1408	3704	22117	17856	39973	1771	1798	1790	863	1168	979
FR	5856	8268	14124	10463	14464	24927	184784	121947	306731	3055	1375	2072	1666	743	1131
GR	229	124	334	402	410	801	6378	7980	14358	2687	6340	4196	1486	1846	1693
IE	1131	52	1147	1289	274	1551	48585	34490	83075	4196	66739	7145	3669	12499	5255
IT	3561	4394	7955	7482	8827	16309	136185	111713	247898	3725	2442	3016	1720	1166	1420
NL	5808	2983	8792	11374	5217	16590	180291	50171	230461	3004	1582	2521	1485	862	1289
PT	1037	1318	2316	2538	3464	6002	26477	6838	33316	2454	419	1338	943	97	455
SE	3014	2401	5415	4765	2998	7763	38653	34258	72912	1182	1327	1246	711	1043	839
UK	10946	29262	40208	12773	30349	43122	140445	134176	274621	1183	359	583	1000	342	537
EU-15	51598	62637	114130	102786	104524	207286	1357605	871951	2229555	2531	1292	1854	1221	734	976

Exports in Mill. $US (1990 prices, 1990 PPPs). – RoW ... Rest of World.

Table A2 – Sensitivity analysis of static model: Results for different estimation periods

	7 observations		6 observations			5 observations				4 observations				
Variable	60-95	65-00	60-90	65-95	70-00	60-85	65-90	70-95	75-00	60-80	65-85	70-90	75-95	80-00
$\Delta\ln (Y_i+Y_j)$	1.877^{***}	1.815^{***}	1.486^{***}	1.839^{***}	1.962^{***}	2.186^{***}	2.367^{***}	2.082^{***}	2.240^{***}	3.200^{***}	2.575^{***}	2.588^{***}	2.691^{***}	2.247^{***}
$\Delta\ln s_{ij}$	0.657^{***}	0.863^{***}	0.705^{**}	0.777^{***}	0.440^{**}	1.006^{***}	1.314^{***}	0.223	0.436^{**}	1.027^{*}	1.532^{***}	0.664^{*}	0.124	1.020^{***}
$\Delta\ln (1+t_i^j)$	-1.978^{***}	-1.745^{***}	-1.781^{***}	-1.832^{***}	-1.215^{**}	-1.995^{***}	-1.893^{***}	-1.116^{**}	-1.996^{***}	-2.196^{***}	-2.049^{***}	-1.071^{*}	-1.745^{**}	-1.192
$\Delta\ln REER_i$	0.244^{***}	0.380^{***}	-0.025	0.336^{***}	0.383^{***}	0.267^{***}	0.274^{**}	0.322^{***}	0.285^{***}	0.099	0.229	0.247^{*}	0.234^{**}	0.452^{***}

LSDV estimates of model (d) in Table 3 for different estimation periods.

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