Cultural Proximity and Trade*

Gabriel J. Felbermayr† and Farid Toubal‡

Jan., 2007

Abstract

Cultural proximity may influence bilateral imports through a preference and a trade-cost channel. In empirical gravity models, conventional measures such as common language or religion fail to separately identify those channels. We use bilateral score data from the Eurovision Song Contest, a huge pan-European television show, to construct a measure of cultural proximity that correlates strongly with conventional indicators. Its statistical properties allow to identify the trade-cost and preference channels. For trade in differentiated goods, we find evidence for both channels, with the former accounting for about 65% of the total effect. There is no preference effect for homogeneous goods.

(JEL Codes: F12, F15, Z10. Keywords: International Trade, Gravity Equation, Cultural Proximity, Identification.)

*Thanks to Hartmut Egger, Josef Falkinger, Wilhelm Kohler, Peter Neary, and Rudolf Winter-Ebmer, and to seminar participants at Göttingen University, the ETSG meeting 2006 in Vienna, the VfS meeting 2006 in Bayreuth, and a seminar at Zürich university for comments. Part of this research was undertaken while Felbermayr was visiting Zürich University. Benny Jung and Michael Böhm have provided excellent research assistance. The usual disclaimer applies.

†Department of Economics, University of Tübingen, Nauklerstraße 47, 72074 Tübingen, Germany. Tel.: +49 7071 29 78183. Email: gabriel.felbermayr@uni-tuebingen.de.

‡ Université de Paris I, Panthéon-Sorbonne, CNRS. Centre d'Économie de la Sorbonne, 106-112 Boulevard de l'Hôpital, 75647 Paris. Tel.: +33 (0)1 44 07 83 47. E-mail: toubal@univ-paris1.fr.

1 Introduction

Easy access to foreign markets is an important determinant of bilateral trade volumes and matters for countries' per capita income and welfare (Frankel and Romer, 1999; Redding and Venables, 2004). Usually, researchers model market access as a function of geographical distance. Cultural proximity has received less attention, although empirical trade flow models typically include some measures of it (Boisso and Ferrantino, 1997).

Cultural proximity affects bilateral trade flows through preferences (bilateral affinity) and/or trade costs. Two culturally close countries may trade a lot because they have strong tastes for each others products and/or because trade costs are low. However, only the second channel relates to market access and has a welfare theoretic interpretation. This paper attempts to disentangle the trade cost channel from the preference channel of cultural proximity in an empirical bilateral trade flow model.[1]

Following the recent literature (Anderson and van Wincoop, 2003; Combes, Lafourcade, and Mayer, 2005; Baier and Bergstrand, 2006), we start with a model of trade in differentiated goods. Our theoretical framework includes both the preference and the trade cost channels. We exploit information from a yearly pan-European televised show, the Eurovision Song Contest (ESC). Each participating country sends an artist to perform a song and grades the other competitors' performances according to a strict set of rules. The process gives rise to a matrix of bilateral votes. The grades are established either by popular juries, or more recently, by televoting.

The ESC data correlates strongly with conventional measures of cultural proximity, such as common language, religion or ethnicity. Contrary to these measures, the ESC data are both time-variant and asymmetric. We use this extra variance to identify the preference and the cost channels separately. In the context of bilateral trade, the slow-moving, symmetric component of cultural proximity–language, religion, ethnicity–can be associated to conventional transaction costs. The fluctuating, asymmetric component, in turn, has more to do with preferences.[2]

Various recent academic papers establish that cultural proximity shapes bilateral ESC scores (Ginsburgh and Noury, 2004; Ginsburgh, 2005; Clerides and Stengos, 2006). Through their voting behavior, countries cluster into clubs according to patterns of cultural closeness (Fenn et al., 2006). Interpreting ESC scores as a measure of cultural proximity poses two empirical challenges: First, scores need to be purged from the artistic quality of songs. We chose an agnostic strategy and use song-specific fixed effects. Second, ESC rules imply that countries

[1] In the received literature, authors acknowledge the importance to separately identify the trade cost and the preference channels; see Combes, Lafourcade, and Mayer (2005). However, to our knowledge, no systematic identification attempt exists.

[2] Guiso et al. (2006) distinguish between inherited and fast-moving aspects of culture; see also Manski (2000).

cannot establish complete rankings of their competitors' performances. We address the resulting measurement bias by a two-step Heckman procedure.

Recent economic literature defines culture as a set of "*customary beliefs and values that ethnic, religious, and social groups transmit fairly unchanged from generation to generation*" (Guiso, Sapienza, and Zingales, 2006). That research puts much effort into establishing a causal link between culture and economics, see Spolaore and Wacziarg (2006) or Giuliano, Spilimbergo, and Tonon (2006). Common instruments or proxies for the concepts of beliefs and values are common language, history, religion, ethnicity or genetic traits. We discuss endogeneity issues in a subsection, but focus our attention on hitherto unresolved identification problems.

In the trade literature, authors tend to use the above list of proxies as measures of cultural proximity without always analyzing the fundamental link to beliefs and values.[3] For example, Rauch and Trindade (2002) and Combes, Lafourcade, and Mayer (2005) emphasize the importance of ethnic ties across countries for the flow of information and hence for bilateral trade. There is also a growing literature that correlates attitudes, sentiments, or customary beliefs to the magnitude of bilateral trade (Disdier and Mayer, 2005; Guiso, Sapienza, and Zingales, 2004).

We view cultural proximity as the degree of affinity, sympathy, or even solidarity, between two countries. It is driven by the feeling of sharing a common identity and of belonging to the same group. In the sociological literature (Straubhaar, 2002), there is no doubt that cultural proximity can be asymmetric or fluctuate over time. A country can command huge respect and sympathy for its cultural, societal, and technological achievements without this feeling being reciprocal. For example, the ESC score data suggests that France has been relatively popular in Europe during the sixties and seventies, without these feelings being reciprocal. However, the relative attraction of France has declined since then.

We frame our analysis in a monopolistic competition trade model close to Hanson and Xiang (2004). However, in the specification of the utility function, we allow for an origin-specific preference term (Combes, Lafourcade and Mayer, 2005). That setup allows to study the trade cost and preference channel of cultural proximity for aggregates of homogeneous and differentiated goods. We propose and implement two alternative identification strategies. In the first, we assume that only the time-invariant elements of cultural proximity matter for trade costs. In the second, we assume that trade costs are symmetric, i.e., they affect exports and imports of a country from or to some trade partner in the same way. We discuss these

[3]Most cross-sectional gravity equations use time-invariant, symmetric measures of cultural proximity, e.g., a dummy for common language, or for colonial ties. Recent eminent examples include Alesina and Dollar (2004), who have used cultural proximity measures in the context of explaining international aid, and Rose (2004), who studies the effect of WTO membership on trade. Melitz (2002) and Hutchinson (2006) provide thorough discussions of the empirical relation between language and bilateral trade.

assumptions and find them largely in line with arguments presented in the theoretical and empirical literature.

Econometrically, we follow Baldagi, Egger, and Pfaffermayr (2003) and interact country and year fixed-effects. This strategy controls for multilateral resistance (Anderson and van Wincoop, 2003) and is a natural extension of the fixed effects approach discussed by Feenstra (2004) and used in Redding and Venables (2004) to a framework with time-varying variables.[4] We discuss the potential endogeneity of the time-variant, and/or asymmetric components of cultural proximity due to, for example, habit formation. We use dyadic fixed effects to effectively control for initial conditions, thereby reducing endogeneity concerns.

Our main empirical results can be summarized as follows. First, we argue that *quality-adjusted* ESC scores are good summary proxies of cultural proximity. They correlate positively and strongly with conventional measures of cultural proximity such as linguistic, genetic, religious, and legal system proximity and yield comparable overall predictions in empirical gravity equations. In line with expectations, the total cultural proximity effect is by an order of magnitude larger for differentiated goods than for homogeneous goods. Second, we use the adjusted ESC scores in two alternative econometric specifications that allow to disentangle the preferences and the trade-costs effects. With *aggregate* bilateral imports, the total effect of moving from the lowest to the highest possible degree of cultural proximity leads to trade creation of about 150%. Assuming that only time-invariant components of cultural proximity are relevant for the trade-cost channel, the preference effect accounts for about 50% of total trade creation, the remaining 100% made up by the cost channel. Focusing on *differentiated* goods (according to Rauch, 1999), overall trade creation is slightly smaller (due to the smaller elasticity of substitution), but the preference channel again makes up about a third of the total effect. Assuming that only symmetric components of cultural proximity are relevant for the trade-cost channel, we find no evidence for a preference effect in aggregate trade. However, focusing on differentiated goods, there is a significant preference channel again, which amounts to about half of total trade creation. We conclude that trade in differentiated goods is significantly affected through the preference channel along with the trade cost channel.

The remainder of this paper is structured as follows. In Section 2, we provide a thorough discussion of the data. In section 3, we propose a theoretical framework and discuss the empirical strategy. In section 4, we present the main results and provide some robustness checks. In Section 5, we conclude.

[4]Baier and Bergstrand (2006) interact country and year fixed effects, too.

2 Eurovision Song Contest score data

In this section, we discuss the ESC data and show that it is a meaningful measure of cultural proximity. We also discuss its statistical properties.

2.1 The setup of the contest

In 1955, a couple of European broadcasting stations represented in the European Broadcasting Union (EBU) founded the Eurovision Song Contest (ESC). A year later, the first contest took place in Lugano, Switzerland. The idea of the ESC is that each participating country selects an artist or a group of artists to perform a song, which is graded by other countries on a scale from 0 to 12 (see below). Hence, if N countries participate, there are $N(N-1)$ bilateral votes.

We focus on the period 1975-2003, in which the described grading rules have been stable. In 1975, the first year in our data, the number of participants was 19. Until 2003, the last year in our sample, the contest was organized in a single round. From 2004 onwards, there are two rounds to accommodate the rising number of participants. On average from 1975 to 2003, the number of participants was 21.6.[5] Contest participants are mostly European countries. However, Israel participates on a regular basis and Morocco has participated once. Each ESC is broadcast by television, and since 1985, this happens via satellite. In 2005, the contest was broadcast live in over 40 countries to over 100 million spectators. Until 1988, the scores were decided upon by a jury that is not necessarily consisting of experts. Nowadays, the scores are national averages obtained through a televoting process with huge popular participation.

Since 1975, each country scores the other countries' performances on a scale from 0 to 12. The scores 9 and 11 are not allowed and 12 is the highest possible score. This allows each voting country to give positive ratings to ten other countries. With an average of 21 countries to grade, a country awards non-zero points approximately to half of the performing countries, the rest is awarded zeros. The winner is the country that collects the largest sum of points.

2.2 Measurement issues

We argue that ESC scores are appropriate summary indicators of cultural proximity. However, this interpretation poses two difficulties. First, the scores may reflect the quality of the song along with cultural proximity. Second, the rules of the contest disallow ties for the ten most popular songs and force ties for the least popular ones.

Denote by S_{ijt} country $i's$ support of country $j's$ performance at time t. Let S_{ijt} depend positively on the quality of the song, Q_{jt}, on the degree of country $i's$ feeling of cultural proximity

[5]The number of participants never fell below 18 in the considered period.

towards j at time t, Π_{ijt}, and on normally distributed noise, u_{ijt}, i.e.,

$$S_{ijt} = S\left(\Pi_{ijt}, Q_{jt}, u_{ijt}\right). \tag{1}$$

We do not model the aggregation of jury members' or telespectators' preferences into S_{ijt}. However, we need to be more explicit on the mapping of S_{ijt} into actual scores, denoted by ESC_{ijt}. This mapping is shaped by the official rules of the competition.

Assume that at time t, $N_t \geq 11$ countries participate in the competition. Each country grades the other countries' performances; hence, for each $i \neq j$ and t, there are $N_t - 1$ realizations of S_{ijt}. According to ESC rules, each country must award strictly positive points to 10 songs, the remaining $N_t - 11$ songs receive zero points. Assuming that each country i ranks the competitors' performances according to the realization of S_{ijt}, we can introduce a function $G\left(S_{ijt}\right) : R^{+} \rightarrow \{0, 1, 2, 3, 4, 5, 6, 7, 8, 10, 12\}$ which maps support S_{ijt} into scores. $G\left(.\right)$ is a positive, monotonic, increasing function under the restriction that each non-zero score has to be used exactly once. That rule implies that it is impossible to express indifference between high-ranked alternatives. More importantly, the observer cannot infer anything about the ranking of the $N_t - 11$ lowest-ranked alternatives, except that they are weakly inferior to the 10^{th} best song. We can therefore write

$$ESC_{ijt} = \begin{cases} G\left(S_{ijt}\right) > 0 & \text{if} \quad S_{ijt} \geq \bar{S}_{it} \\ 0 & \text{if} \quad S_{ijt} < \bar{S}_{it} \end{cases}, \tag{2}$$

where $\bar{S}_{it}$ is a threshold value below which country i awards zero points.

ESC scores are only partly informative about cultural proximity if $ESC_{ijt} = 0$. For those observations, there is a negative correlation between measurement error and the true amount of cultural proximity, which would bias estimates in a way that is difficult to correct for. Hence, we use only observations for which $ESC_{ijt} > 0$ and implement a Heckman (1979) two-stage correction procedure to deal with the resulting non-random sample composition. In the first stage, we formalize the probability that country i receives a strictly positive rating by country j as a Probit model,

$$\Pr\left(S_{ijt} \geq \bar{S}_{it}\right) = \Phi\left(\chi_0 \ln \tilde{\Pi}_{ij} + \chi_1 \ln Q_{jt}\right), \tag{3}$$

where Φ is the c.d.f. of the standard normal distribution, and equation (1) has been linearized. We estimate the model using observable proxies of cultural proximities that are available for all country pairs $\tilde{\Pi}_{ij}$: common language, common religion and common legal origin. Since there is one singer per country, the information on the song quality is captured by using a country and time-varying specific effect, Q_{jt}.

We use the Probit equation to compute the inverse Mill's ratio $\lambda_{ijt} \equiv \lambda\left(\tilde{\Pi}_{ij}, Q_{jt}\right)$ that we then include into our gravity regressions.[6] λ measures the hazard of selection into the model; in our case, whether a performance with quality Q_{jt} of a country with cultural proximity Π_{ijt} will be awarded strictly positive scores from country i.[7] Countries do not participate permanently in the ESC. We have checked whether selection into participation is random. ESC rules determine that winners organize the following year's contest, so that they are very likely to participate. However, it does not turn out that past success systematically determines the probability to stand at the contest. Hence, we do not further pursue this potential selection issue.

To make use of the ESC measure in our gravity equations, for $ESC_{ijt} > 0$, we linearize (2) so that $ESC_{ijt} = \ln \Pi_{ijt} + \chi \ln Q_{jt} + \xi \lambda_{ijt} + u_{ijt}$ from where we compute a measure of cultural proximity

$$\ln \Pi_{ijt} = ESC_{ijt} - \chi \ln Q_{jt} - \xi \lambda_{ijt} - u_{ijt}. \tag{4}$$

As in the Probit equation, we account for unobserved song quality by including a set of song-specific fixed effects. Since we interpret the ordinal ESC score data in a cardinal way, in all our calculations we enforce an even spacing of scores. Moreover, to facilitate comparison with conventional measures of cultural proximity, we rescale the data so that $ESC_{ijt} \in (0, 1)$.[8]

2.3 ESC scores as summary indicators of cultural proximity

ESC data has been used in academic research, albeit not in trade empirics. Recently, in his empirical exercise, Ginsburgh (2005) shows that conventional measures of cultural proximity determine ESC outcomes to a large extent, refuting the alternative hypothesis of vote trading (logrolling). Clerides and Stengos (2006) and Spierdijk and Vellekoop (2006) similarly document the importance of culture for scoring outcomes. Haan, Dijkstra, and Dijkstra (2005) test whether the transition of the grading process from jury-based voting to generalized televoting has strengthened the role of cultural proximity and answer this question in the affirmative. Fenn et al. (2006) run a cluster analysis and find evidence for unofficial cliques of countries along lines of cultural proximity.

Table 1 shows pairwise correlation coefficients between quality-adjusted ESC scores and conventional measures and reports the P-values for the null-hypothesis that the correlation coefficient is zero. The quality adjusted ESC scores are the residuals from a regression of raw scores on song-dummies. We consider the following time-invariant and symmetric measures of cultural proximity used in the literature: (i) a common language dummy; (ii) the continuous Dyen et al.

---

[6] The inverse Mill's ratio is given by $\phi\left(\hat{\chi}_0 \ln \tilde{\Pi}_{ij} + \hat{\chi}_1 Q_{jt}\right) / \Phi\left(\hat{\chi}_0 \ln \tilde{\Pi}_{ij} + \hat{\chi}_1 Q_{jt}\right)$.

[7] We use the ESC score as a cardinal expression of S_{ijt}, while the scores really reflect ordinal preference rankings. This practice introduces additional measurement error which may bias our results towards zero.

[8] First, the original score of 10 is set to 9 and the score 12 is set to 10. Then, we divide all scores by 10.

(2002) measure of linguistic proximity; (iii) a dummy for common legal origin borrowed from La Porta et al. (1999); (iv) a continuous religious proximity measure based on Alesina et al. (2003)[9]; (v) a measure of ethnic links based on the stock of individuals born in country i but residing in country j; (vi) a measure of genetic similarity used by Giuliano, Spilimbergo, and Tonon (2006) or Spolaore and Wacziarg (2006). To capture geographical proximity, we use an adjacency dummy and geographical distance between main cities (in kilometers). Appendix A provides information on the summary statistics of the variables, their exact definition and the data sources.

Table 1: *Coefficients of correlation between different measures of cultural proximity*

	ESC score	Adjusted ESC score
Adjusted ESC score	79.34***	
Common language (dummy)	15.55***	13.67***
Linguistic proximity	29.91***	38.66***
Common legal origin	11.45***	20.33***
Religious proximity	13.34***	16.45***
Ethnic similarity	20.18***	20.63***
Genetic distance	-16.33***	-12.05**
Adjacency	18.63***	22.13***
Geographical distance	-13.81***	-18.76***

Number of observations: 940, except Linguistic proximity (176), and genetic distance (422). *** denotes that coefficient is different from zero at 1% level of significance, ** at 5 % level of significance.

The correlation between the quality adjusted ESC score and the raw ESC score data is 79.34. Hence, about a fifth of the variance in raw scores is accounted for by quality. The correlation coefficient between both ESC score measures and the conventional proxies of cultural and geographical proximity all have the right sign and are statistically different from zero. The adjusted ESC score displays larger correlation coefficients (with the exceptions of common language and genetic distance), signaling that quality adjustment improves our index of cultural proximity.

[9]Countries that once were part of the same larger political entity (such as, e.g., the Czech Republic and Hungary in the Austro-Hungarian Empire, England and Ireland, or Norway and Denmark) usually have the same legal system. Hence, that dummy also captures the broad institutional similarity of countries that have had a common past.

2.4 Properties of ESC scores

In this subsection, we briefly discuss two important statistical properties of the ESC scores. First, they exhibit some reciprocity, but they are far from symmetric. Second, they exhibit meaningful time variation.

Table 2 provides evidence for the first fact. The table reports the directed pair-specific intercepts $\hat{\nu}_{ij}$ derived from estimating $ESC_{ij} = \mu_{it} + \nu_{ij} + u_{ijt}$ by OLS, where μ_{jt} is the average score attained by country j at time t and u_{ijt} is an error term. We refer to the empirical estimates of the intercepts $\hat{\nu}_{ij}$ as $excess_{ij}$ since they measure grading behavior in excess of means. Table 2 shows selected country relations where excess-grading is statistically significant at least in one direction.[10] The expected country clustering readily emerges. For example, Cyprus awards to Greece an average of 7.41 points more than Greece receives on average; Greece reciprocates by awarding an excess of 6.26 points beyond the Cypriot mean score. Over-generous and reciprocal relationships can be found for many Scandinavian country pairs, and to a lesser extent for Mediterranean countries. However, relationships need not be reciprocal: Finland awards Italy an excess 3.10 points, but gets an average negative excess grade of -0.77 in return. France grades Great Britain 0.86 points below average, while the Brits treat France almost neutrally.[11] Scandinavian and Mediterranean countries tend to reciprocally award scores below the respective averages (with the pair Denmark-Yugoslavia the notable exception). Not surprisingly, Cyprus and Turkey stand out as two countries that systematically award each other grades below average. Germany (and Austria) over-grade Turkey without being compensated; this is likely to reflect ethnic ties due to migration.

The pattern found in Table 2 accords well with intuition: Reciprocal positive excess grades occur within pairs that share similar cultural traits; reciprocal negative excess grades appear where expected but are on average much smaller than positive excess grades. Moreover, non-reciprocal behavior seems to be quite frequent. The (simple) correlation coefficient between $excess_{ij}$ and $excess_{ji}$ is 35.07, the Spearman rank correlation coefficient is 27.81; both measures are statistically different from zero at the one percent level of significance.

Finally, we investigate the time behavior of ESC scores. Time patterns of multilateral scores are presented in Appendix B. The swings in bilateral and aggregate scores allocated to countries are not easy to interpret. Our claim is that the perceived patterns can be attributed to time-moving aspects of cultural proximity. Indeed, the ESC data seems to reflect political events that signal the cultural alienation of a country. For example, in 2003, the UK was punished

[10] See also Table 1 in Clerides and Stengos (2006).

[11] This confirms the Duke of Wellington: "*We always have been, we are, and I hope that we always shall be detested in France.* " (cited by Guiso et al., 2004, p. 1.)

Table 2: *ESC scores: Selected deviations from means*

Country i	**Country** j	$excess_{ij}$	**std. err.**	$excess_{ji}$	**std. err.**
CYP	GRC	7.41	0.65	6.26	0.65
ITA	PRT	3.95	0.63	0.02	0.63
DNK	ISL	3.27	0.72	2.05	0.72
FIN	ITA	3.10	0.65	-0.77	0.65
DNK	SWE	2.99	0.55	1.98	0.55
ISL	SWE	2.95	0.65	1.56	0.65
ESP	ITA	2.95	0.63	1.70	0.63
CYP	YUG	2.67	0.82	2.49	0.82
HRV	MLT	2.52	0.78	3.17	0.78
TUR	YUG	2.22	0.75	3.21	0.75
DNK	NOR	2.04	0.57	0.79	0.57
GER	TUR	1.79	0.53	-0.70	0.53
CYP	ESP	1.79	0.57	-0.27	0.57
ESP	GRC	1.77	0.54	2.62	0.54
ISL	NOR	1.71	0.65	1.08	0.65
NOR	SWE	1.56	0.50	2.64	0.50
BEL	NLD	1.33	0.55	0.30	0.55
AUT	TUR	1.12	0.55	-0.21	0.55
FIN	SWE	1.06	0.54	0.67	0.54
FRA	GBR	-0.86	0.49	0.18	0.49
ESP	NOR	-1.01	0.49	-1.37	0.49
ESP	SWE	-1.39	0.49	-1.20	0.49
ITA	SWE	-1.51	0.65	-0.93	0.65
DNK	ESP	-1.66	0.55	-0.39	0.55
ISR	ITA	-1.73	0.69	-1.73	0.69
HRV	SWE	-1.76	0.78	-3.06	0.78
DNK	HRV	-1.92	0.98	-2.53	0.98
CYP	TUR	-2.09	0.58	-1.22	0.58
DNK	YUG	-2.17	0.78	1.34	0.78

Pair-specific intercepts, means adjusted. All estimates are significant at the 10 percent level. All country pairs in the table occur at least 10 times in the data. $Excess_{ij}$ denotes the score awarded by i to j in excess to j's average score.

by European voters for her support of the war in Iraq. In that year, its total score was zero, which is a rare event given the numerics of the ESC rules. Similarly, when Austrians elected Kurt Waldheim, a person with an unclear record during World War II, as president, aggregate scores touched zero, too. Other examples are readily found. While the examples cited above mark short-term effects, the data also reveals long-run trends. For example, the popularity of countries seems subject to cycles that are difficult to reconcile with underlying movements in musical quality. For example, France scored high in the seventies, when France was seen as an example to follow and la Chanson Francaise enjoyed Europe-wide popularity. Since then, this popularity has faded; actually, it seems that France now suffers the drawbacks from its *"exception culturelle"*. The example of Italy offers a similar, but less spectacular, picture. It is well possible, that the popularity of celtic culture (Ireland) will suffer the same fate.[12]

[12] In the Appendix, we present the time profile of multilateral unadjusted ESC scores earned by frequently participating countries. Many of the swings in scores can be recognized as coinciding with important political or societal events. A detailed analysis, is, however, beyond the scope of the present paper.

3 Theoretical framework and empirical setup

3.1 Cultural proximity in the gravity model

We base our theoretical model on the multi-country monopolistic competition model of trade (see Feenstra, 2004, for an overview). Each country i is populated by a representative individual who derives utility from consuming varieties produced in different sectors, $s = 1, ..., S$, and possibly originating from different countries, $j = 1, ..., C$. Following Hanson and Xiang (2004), we assume a two-tier demand system. Consumers have identical Cobb-Douglas preferences, with θ_s the share of consumption spending on sector s and $\sum_s \theta_s = 1$. For each sector s, a continuum of varieties from different countries is aggregated using a standard CES utility function. We denote by z the index of a generic variety, by n_{sjt} the number of sector-s varieties produced in country j at time t and $\sigma_s > 1$ the sectoral elasticity of substitution between varieties. The quantity of consumption in country i of variety z from country j and sector s is $m_{isjt}(z)$. Moreover, as Combes, Lafourcade, and Mayer (2005), we allow for a specific weight $a_{isjt} \geq 0$ to describe the special preference of the representative consumer in country i for goods from country j. Hence, the utility function is given by

$$U_{it} = \sum_s \theta_s \ln \left\{ \sum_{j=1}^{C} \int_{n_{sjt}} \left[a_{isjt} m_{isjt}(z) \right]^{\frac{\sigma_s - 1}{\sigma_s}} dz \right\}. \tag{5}$$

This representation allows to consider groups of varieties ('sectors') with different degrees of within-group substitutability.

We assume that all varieties from the same origin bear the same f.o.b. (ex factory) price p_{sjt} (reflecting symmetric production technologies within sectors), and that iceberg *ad-valorem* trade costs payable for deliveries from j to i, $t_{isjt} \geq 1$, do not depend on the characteristics of the varieties within a sector. Hence, the c.i.f. price paid by consumers $p_{isjt} = p_{sjt} t_{isjt}$ is the the same for all varieties imported from j. It follows that consumed quantities m_{isjt} are identical for all z.

Maximizing (5) subject to an appropriate budget constraint, one derives country $i's$ demand quantity m_{isjt} for a generic variety. Calculating the c.i.f. value of total sector-s imports from country j at time t as $M_{isjt} = n_{sjt} p_{isjt} m_{isjt}$, we find

$$M_{isjt} = \left(\frac{a_{isjt}}{t_{isjt}} \right)^{\sigma_s - 1} \phi_{ist} \phi_{jst}. \tag{6}$$

We follow Redding and Venables (2004) and define $\phi_{ist} \equiv \theta_s E_{it} P_{ist}^{\sigma_s - 1}$ as country $i's$ market capacity for sector-s varieties, and $\phi_{jst} \equiv n_{sjt} p_{sjt}^{1-\sigma_s}$ as the sector-s supply capacity of the exporting country j. P_{ist} is the sectoral price index, $P_{ist} = \left[\sum_{j=1}^{C} \left(\frac{a_{isjt}}{t_{isjt}} \right)^{\sigma_s - 1} p_{sjt}^{1-\sigma_s} n_{sjt} \right]^{\frac{1}{1-\sigma}}$ and E_{it} denotes country $i's$ GDP.

We do not close the model by explicitly specifying supply side and equilibrium conditions, since this is not needed for our empirical investigation. However, we need to clarify the role of cultural proximity in shaping bilateral trade flows. Cultural proximity affects the bilateral trade equation (6) in two ways. On the one hand, it lowers direct trade costs, t_{isjt} : Costly translation and cultural advisory services are redundant when partners share a common language or/and interpret non-verbal communication correctly. Contracting costs are lower when buyers and sellers operate in similar legal environments, and trust builds up faster when there are ethnic links. Moreover, cultural proximity indirectly affects trade costs as it facilitates the formation of business and/or social networks. In turn, these networks help to overcome informational trade barriers (Rauch and Trindade, 2002). We refer to this channel as to the trade cost channel of cultural proximity.

On the other hand, cultural proximity influences bilateral trade through the bilateral affinity parameter a_{isjt}. A high value of a_{isjt} means that the representative consumer in country i puts a high value on products produced in country j. Equation 6 together with the assumption $\sigma_s > 1$ implies that this situation leads to larger sectoral trade volumes. We refer to this second channel to the preference channel of cultural proximity.

We now specify how country $i's$ cultural proximity to country j is related to bilateral affinity and trade costs. In both cases, we allow for this link to depend on sectoral characteristics: in particular, it is plausible that the preference channel is weaker in trade of homogeneous goods. We assume that country $i's$ preference for goods from i, a_{ijt}, depends on Π_{ijt} in the following way:

$$\ln a_{isjt} = \alpha_s \ln \Pi_{ijt}. \tag{7}$$

Concerning trade costs t_{isjt}, we assume that they are driven by three factors:[13] (i) Transport costs $K_{ijt} = I_t \cdot DIST_{ij}^{\delta_s} \cdot e^{\gamma_s (1-ADJ_{ij})}$, where $DIST_{ij}$ refers to geographical distance $DIST_{ij}$, ADJ_{ij} is a dummy that takes value of unity of two countries share a common country, and I_t captures the general state of transport technology. (ii) Trade policy $T_{ijt} = T_t \cdot e^{\varphi_s (1-FTA_{ijt})}$, where FTA_{ijt} takes the value of one if both countries are in the same free trade agreement.[14] (iii) Cultural proximity Π_{ijt}. All parameters are defined as positive numbers. For the sake of simplicity (and in line with the literature), we assume a log-linear relationship between t_{isjt} and its determinants

$$t_{isjt} = K \cdot I_t \cdot DIST_{ij}^{\delta_s} \cdot e^{\gamma_s (1-ADJ_{ij})} \cdot T_t \cdot e^{\varphi_s (1-FTA_{ijt})} \cdot \Pi_{ijt}^{-\beta_s}, \tag{8}$$

[13] See Anderson and van Wincoop (2004) for a classification of different types of trade costs.

[14] We have tried to use direct measures of trade policy, i.e., average tariff rates. This leads to a number of conceptual difficulties, but leaves our main empirical results unchanged.

where K is a constant. We expect that the trade cost elasticity of cultural proximity, β_s is smaller for homogeneous goods than for differentiated goods.

3.2 Two alternative identification strategies

We can now substitute expressions (4), (8) and (7) into the gravity equation (6). We neither observe song quality Q_{jt} nor the sectoral market and supply capacity terms ϕ_{ist} and ϕ_{jst}. We control for these variables by using a comprehensive set of interaction terms of exporter and importer fixed effects with year fixed effects and estimate *baseline* gravity equations *sector by sector*:

$$\begin{aligned}\ln M_{isjt} &= \eta_s^0 ESC_{ijt} - \bar{\delta}_s \ln DIST_{ij} + \bar{\gamma}_s ADJ_{ij} + \bar{\varphi}_s^0 FTA_{ijt} \\ &\quad + \bar{\xi}_s^0 \lambda_{ijt} + \nu_s + \nu_{st} + \nu_{ist} + \nu_{jst} + \varepsilon_{isjt},\end{aligned} \tag{9}$$

where $\eta_s^0 \equiv (\sigma_s - 1)(\alpha_s + \beta_s)$, $\bar{\delta}_s \equiv \delta_s(\sigma_s - 1)$, $\bar{\gamma}_s \equiv \gamma_s(\sigma_s - 1)$, $\bar{\xi}_s^0 \equiv \xi_s(\sigma_s - 1)$, and $\bar{\varphi}_s^0 \equiv (\sigma_s - 1)\varphi_s$. The term ν_s is a constant, ν_{st} is a set of year dummies, ν_{ist} is a comprehensive set of importer $\times$ year fixed effects that control for the demand capacity ϕ_{ist} and ν_{jst} collects exporter $\times$ year fixed effects to control for the supply capacity ϕ_{sjt} and song quality Q_{jt}. Finally, $\varepsilon_{isjt} \equiv \eta_s^0 u_{isjt}$ is the error term. This strategy makes the inclusion of GDP or price data redundant. It also frees us from the need to think about the choice of proper deflators for right and left-hand side variables of the gravity equation.[15]

Specification (12) can be estimated sector by sector using available data. The total effect of cultural proximity can then be identified using some external information on the value of σ_s. Anderson and van Wincoop (2004, table 7) consider a range between 4 and 10 for differentiated goods, and a substantially higher σ_s for homogeneous goods as reasonable. However, we do not need information on σ_s to assess the *relative* importance of the trade-cost versus the preference channels for given s. In the remainder of this section we discuss two strategies to identify α_s and β_s separately.[16]

In contrast to conventional measures of cultural proximity, quality-adjusted ESC scores are asymmetric in the sense that they are not necessarily reciprocal. And they are time-variant. This additional variance provides us with two natural identification strategies. These strategies build

[15] Baldagi, Egger, and Pfaffermayr (2003) first propose the use of interaction terms between country and time fixed effects in econometric gravity models, without, however, offering a theoretical rationale. The latter follows immediately from Anderson and van Wincoop (2003). Baier and Bergstrand (2006) use a similar strategy in their work on the effects of free trade agreements.

[16] Absent endogeneity concerns (which we address below), our empirical strategy allows consistent estimation of *average* effects; since changes in cultural proximity affect multilateral resistance terms (price indices) differently, the true coefficient is not constant across countries (Anderson and van Wincoop, 2003).

on a decomposition of country $i's$ cultural proximity to country j (Π_{ijt}) into three components,

$$\Pi_{ijt} = \pi_{ij}\bar{\pi}_{ijt}\pi_{ijt}, \tag{10}$$

where, for the sake of simplicity, we postulate a multiplicative relationship. The first component π_{ij} is time-invariant, but potentially asymmetric; the second is symmetric (i.e., $\bar{\pi}_{ijt} = \bar{\pi}_{jit}$), but potentially time-variant, while the third is both potentially asymmetric and time-variant.

In our first approach, we make the following assumption:

Assumption 1. *Trade costs depend on cultural proximity only through the time-invariant component* π_{ij}.

The motivation for this assumption is that the costs of doing business between two countries depend on linguistic, religious, or ethnic ties, which, if at all, change very slowly over time. To the extent that Assumption 1 holds true, one can unambiguously attribute the time variant component of cultural proximity, $\bar{\pi}_{ijt}\pi_{ijt}$, to the preference channel. To filter out time-invariant components, we estimate for each sector s the following *'dyadic fixed effects gravity'* (DFEG) model:

$$\ln M_{isjt} = \eta_s^1 ESC_{ijt} - \bar{\varphi}_s^1 FTA_{ijt} + \bar{\xi}_s^1 \lambda_{ijt} + \nu_s + \nu_{st} + \nu_{ist} + \nu_{sjt} + \nu_{sij} + \varepsilon_{sijt}, \tag{11}$$

where ν_{sij} denotes the comprehensive set of country-pair specific fixed effects. Running (11) for each sector s (and for aggregate data) we get first estimates of the preference channel $\hat{\alpha}_s^1 = \hat{\eta}_s^1 / (\sigma_s - 1)$ and of the trade-costs channel as $\hat{\beta}_s^1 = \left(\hat{\eta}_s^0 - \hat{\eta}_s^1\right) / (\sigma_s - 1)$.

The parameter η^1 is identified by drawing on time-variance in cultural proximity, which reflects fashions and fads and is therefore attributable to the preference channel. Note that we do not assume that the preference channel is driven *exclusively* by time-variant components of cultural proximity. Of course, we can use (11) also on aggregate data, since sectoral differences enter only through potentially different parameter values.

There is a natural way of validating our identification strategy. Country i's imports from j depend on the time variant component of cultural proximity $\bar{\pi}_{ijt}\pi_{ijt}$ through the preference channel. However, country i's exports to j should depend depend on $\bar{\pi}_{ijt}\pi_{ijt}$ only to the extent that cultural proximity is actually symmetric. We have seen in Table 2 that this is not systematically the case. Hence, we expect that Π_{ijt} affects imports more strongly than exports.

Assumption 1 is violated if the trade-cost related component of cultural proximity changes over time. We do not expect large shifts to occur in the time span and country sample that we study (1975-2003, European countries). However, linguistic, ethnic or religious proximity between countries can change, for example due to migration or due to learning (which might well be driven by trade itself; see below for more discussion). Then, $\hat{\eta}_s^1$ based on (11) may be a biased estimate of the preference channel, as part of the time change in Π_{ijt} also affects the cost channel. Regardless of the trend in time-variant component of cultural proximity, we would expect $\hat{\eta}_s^1$ to overestimate the true effect $(\sigma_s - 1)\,\alpha_s$. In the extreme case, where the time-variance of the trade-cost relevant component of cultural proximity cannot be restricted, we would actually estimate $\hat{\eta}_s^1 = (\sigma_s - 1)\,(\alpha_s + \beta_s)$.

We therefore work with a second strategy, which builds on the following assumption:

Assumption 2. *Trade costs are symmetric (but may be time-variant). I.e.,* $t_{sijt} = t_{sjit}$*. This is equivalent to stating that trade costs depend only on cultural proximity through the symmetric component* $\bar{\pi}_{ijt}$*.*

The prime motivation behind Assumption 2 is its wide-spread use in theoretical work. It implies that a decrease in trade costs through an increase in cultural proximity does not have a systematically different effect on imports and exports of a country. Under Assumption 2, and using equation (6) we can write country $i's$ exports to j relative to its imports from j (both measured c.i.f.)

$$\frac{M_{sijt}}{M_{sjit}} = \left(\frac{a_{sijt}}{a_{sjit}}\right)^{\sigma-1} \phi_{sit}\phi_{sjt} \tag{12}$$

where now $\phi_{sit} \equiv n_{sit}\,(p_{sit}P_{sit})^{1-\sigma_s}\,/E_{sit}$ and $\phi_{sjt} \equiv E_{sjt}\,(p_{sjt}P_{sjt})^{\sigma_s-1}\,/n_{sjt}$. Note that in deriving (6) we have not assumed that bilateral trade positions are balanced. Even if both countries share the same values for $n_{sit}, p_{sit}, P_{sit}$, and E_{sit}, and trade costs are symmetric, $t_{sijt} = t_{sjit}$, due to the asymmetric preference term, bilateral trade positions need not be balanced, so that the ratio M_{sijt}/M_{sjit} is not necessarily equal to unity. This argument holds also on the aggregate level.[17] The summary statistics in the Appendix show that bilateral trade positions display a large degree of variation across country pairs and time. Similarly, quality-adjusted measures of cultural proximity display asymmetry and time variation.

Substituting $\ln a_{jit} = \alpha \ln \Pi_{ijt}$ into (12) and using equation (4), we obtain the following empirical specification that we dub *'bilateral trade balance equation'*.

$$\ln M_{sijt} - \ln M_{sjit} = \eta_s^2\,(ESC_{ijt} - ESC_{jit}) + \xi_s^2\,(\lambda_{ijt} - \lambda_{jit}) + \nu_{sit} + \nu_{sjt} + \varepsilon_{sijt}, \tag{13}$$

where $\eta_s^2 \equiv (\sigma_s - 1)\,\alpha_s$, $\xi_s^2 \equiv (\sigma_s - 1)\,\xi_s$. The vectors ν_{sit} and ν_{sjt} collect interactions between country and year fixed effects which control for the difference of country i's and j's supply and demand capacities. Indirectly, those fixed effects also account for the real bilateral exchange rate, as they summarize the underlying stances of monetary and fiscal policy in both countries.

[17] Davis and Weinstein (2002) derive a bilateral trade balances model that holds for any trade model that implies perfect specialization. They discuss the possibility of bilateral trade imbalances. If supply and demand conditions are not symmetric across two countries–as in our case due to the bilateral preference term in the description of utility–bilateral trade positions need not be zero. Aggregate balance is then restored by triangular trade.

Again, we compute the size of the preference channel as $\hat{\alpha}_s^2 = \hat{\eta}_s^2 / (\sigma_s - 1)$ and the size of the trade-costs channel as $\hat{\beta}_s^2 = (\hat{\eta}_s^0 - \hat{\eta}_s^2) / (\sigma_s - 1)$. We can run model (13) separately for each sector. Note that this specification involves at most half the number of observations than in specification (11) since the unit of observation is the (undirected) country pair. A further significant reduction in observations results from the requirement that ESC_{ijt} and ESC_{jit} be simultaneously strictly positive.[18]

Literature on bilateral trade balances is rather scarce. The only recent empirical study that we are aware of is Davis and Weinstein (2002). Those authors ground their analysis in a Helpman-type gravity framework (Helpman, 1987). They estimate a standard gravity equation and compute predicted bilateral trade balances using those equations. They find that their model over-predicts trade balances. Moreover, predicted trade balances have the wrong sign in almost half of all cases. In contrast, our approach accounts for the fundamental drivers of bilateral trade positions (e.g., fiscal and monetary policies) by interacting exporter and importer fixed effects with year fixed effects. This strategy provides a robust shell for testing our hypothesis. Running equation (13) without including the ESC score results in a R^2 of about 55 percent; see below. The regression predicts the sign of the bilateral trade balance correctly in about 79 percent of all cases. The correlation between actual and predicted balances is about 74 percent.

3.3 Addressing endogeneity concerns

Can our results be interpreted as evidence about a causal link between cultural proximity an bilateral trade? In this paper, we are primarily interested in identification issues rather than in establishing causality. However, there is a growing theoretical literature on the endogenous emergence and evolution of cultural identities (see the recent survey by Guiso, Sapienza, and Zingales, 2006, and Bisin and Verdier, 2005). Hence, we discuss possible endogeneity issues to clarify the correct interpretation of our empirical results.

To that end, it is useful to recall equation (10), where we decompose cultural proximity into a time-invariant and a time-varying (possibly asymmetric) component. The time-invariant part can be identified with cultural features such as genetic and ethnic traits, or legal foundations, which are *inherited from the past*, and therefore at least weakly exogenous. The time-variant component is more difficult to operationalize and may be more strongly prone to endogeneity concerns. In the economic literature reviewed in the surveys cited above, cultural traits of groups usually depend on aggregate environmental conditions. If two groups face similar conditions, they may develop similar preference structures, institutions, and value systems. And if those

[18] Note that specification (13) is conceptually related, but not identical, to the 'odds'-specification used in Combes *et al.* (2005), who also bring in supply side considerations.

conditions evolve over time in different or similar directions, cultural proximity grows stronger or weaker. In our context, aggregate conditions are aptly controlled for by time-specific country-fixed effects. However, to the extent that cultural proximity between two countries is shaped by economic interactions between those two countries rather than by aggregate conditions, we do have an endogeneity issue. We discuss three possible sources of endogeneity.

First, the ethnic and linguistic composition of a country moves through time as a function of *mass migration*. However, apart from the mobility of people involved in the operation of trading activities (e.g., a company's sales staff in a foreign country), incentives to migrate depend on the comparison between the wage distributions across the two countries (see Borjas, 1987, and the ensuing literature). Those incentives then depend on multilateral and not bilateral trade volumes.[19] Hence, migration does not pose problems in our context.

Second, and more importantly, economic cross-border transactions require and condition *social interactions* between individuals, which may change their incentives to acquire certain skills or invest in the formation of certain networks. For example, interactions with foreign trading partners might lead to mutual learning, which may trigger convergence of cultural characteristics. Or, the sheer return to acquiring language skills is bigger if that language is useful in dealing with a larger trade volume. The joint determination of cultural characteristics and bilateral trade has not been studied systematically so far. However, the issue is certainly important and interesting. It concerns gravity equations in most circumstances: high bilateral trade volumes may trigger cultural convergence, which may then lower trade costs and/or affect specific preferences, thereby triggering more bilateral trade. To solve this problem econometrically, one would need to develop a workable dynamic gravity model. This is beyond the scope of this paper, but represents another worthwhile field of investigation.

A third source of endogeneity closely related to the second is *habit formation*. The specific preference for a country's varieties may grow stronger the more intensively a consumer is already exposed to products of that country. Again, this possibility would suggest a dynamic formulation. One way to deal with the second and the third sources of endogeneity is to include bilateral fixed effects. This helps, because it introduces a control for the unobserved initial bilateral trade volume, which may have led to the endogenous formation of social networks or consumption habits. Our specification (11) contains those effects, and is likely to deliver results robust to endogeneity concerns. Specification (13) can be straightforwardly augmented by dyadic fixed effects, and we will do so in the empirical analysis. Hence, we believe that our estimates of the preference channel are not grossly distorted by endogeneity bias. However, our baseline spec-

[19] The bilateral flow of migrants may well be favored by cultural proximity. This process may magnify the trade creating potential of cultural proximity, but does not induce correlation between the error term and cultural proximity.

ification is susceptible to a (positive) endogeneity bias and should be considered as an upper bound of the true overall effect of cultural proximity on trade.

We have experimented with applying instrumental variables techniques. First, we have used lagged ESC realizations as instruments in all specifications. Second, we have averaged the data over 1975-2003 and instrumented average ESC realizations by conventional measures of cultural proximity. Using that strategy in our baseline and in the bilateral balances specifications, one can identify the cost and preference channel of cultural proximity. Third, we have averaged our data over the period 1985-2003 and used ESC averages from 1975-1984 as instruments. All these instrumentation methods can be critizised on conceptual grounds. They also tend to delivernon-intuitive results: the instrumented effect of cultural proximity is typically of an order of magnitude larger than the uninstrumented. Hence, either the endogeneity bias is actually negative (i.e., higher bilateral trade volumes decrease cultural proximity), which seems implausible. Guiso, Sapienza, and Zingales (2006) and Combes, Lafourcade, and Mayer (2005) also discuss attempts to use IV strategies in similar setups. Both papers find (as we do) negative endogeneity biases. Both papers then come up with the conclusion that 'reverse causality is not a major problem'.

Before presenting the estimation results, it is worth summarizing. We have developed three theory-based specifications: (i) a baseline model, (ii) a dyadic fixed effects gravity (DFEG) model, and (iii) a bilateral trade balance (BTB) model. In this section we acknowledge that the total effect of cultural proximity on bilateral trade as estimated in model (i) may be distorted upwards due to endogeneity bias. We therefore interpret our estimates as upper bounds. However, since the default version of model (ii) includes dyadic fixed effects, the preference effect is probably estimated without bias as long as Assumption 1 is correct. Similarly, one can include dyadic fixed effects in specification (iii) to address endogeneity problems. Hence, we belief that our estimates of the preference channel are unbiased, while estimates of the trade cost channel (the difference between the total and the preference channel) may be biased upwards. Bearing that caveat in mind, results in the literature and own tentative IV regressions do not point towards massive overestimation. They signal rather the opposite.

4 Results

4.1 The *total* effect of cultural proximity

We estimate different versions of our baseline model (9) using aggregate bilateral trade data, and the sub-aggregates proposed by Rauch (1999). We distinguish between homogenous goods that are traded on organized exchanges and differentiated goods.[20]

Tables 3A reports our results for the baseline model, with the natural logarithm of total bilateral imports as the dependent variable.[21] Table 3B replaces the dependent variable by aggregates of homogeneous and differentiated goods, following the classification proposed by Rauch (1999); see the Appendix for details. By default, all regressions include comprehensive sets of year × exporter and year × importer fixed-effect interaction terms, which capture all country- specific time-varying variables such as GDP or the multilateral resistance index. Hence, our models contain only dyadic covariates: geographical distance, a dummy for adjacency, and a dummy that takes the value of unity if two countries belong to the same free trade area (FTA).[22] We consider the measures of cultural proximity discussed in section 2.3, namely common legal origin, linguistic proximity (as a dummy and a continuous variable), religious proximity, genetic distance, and ethnic ties, and compare them to results obtained using ESC scores.

Specification (S2) in Table 3A reports the results of a standard gravity model run on our data set of 10,560 observations. Exporter and importer GDPs are accounted for by the inclusion of country specific time varying effect effects. The elasticity of distance is close to unity, which is a typical result in this context. Hence, the value of bilateral trade increases by about 1 percent when distance increases by 1 percent. Adjacency boosts trade by about 41 percent, which is again a very conventional result. The FTA dummy is positive and of reasonable magnitude, but not statistically significant. This is also a usual finding and discussed at great detail in Baier and Bergstrand (2006). The adjusted R^2 is about 89%.

Specifications (S2) to (S4) report the results of regressions that use different measures of cultural proximity and that are directly comparable to specification (S1) in terms of sample. The cultural proximity in these specifications is approximated by dummy variables. The point estimates of all those measures are precisely estimated with high levels of significance. The beta coefficient associated to common legal origin, common language, and religious proximity are 0.0997, 0.0698, and 0.0545 respectively.[23] Compared to distance, which has a beta coefficient of

[20]We report findings for Rauch's conservative aggregation scheme (which minimizes the number of goods that are classified as either traded on an organized exchange or reference priced).Our results remain qualitatively similar if the liberal aggregation scheme is used. Results for Rauch's third category, reference-priced goods, are very similar to those for homogeneous goods and available upon request.

[21]Our data set being a sample of European countries only, we have only very little observations where the bilateral trade volume is zero. Hence, there is no need to estimate a corner solutions model. While this is true at a lesser extent when we look at trade in homogeneous and differentiated goods, we stick to the same econometric method.

[22]See the Appendix for summary statistics.

[23]A beta coefficient is defined as the product of the estimated coefficient and the standard deviation of its corresponding independent variable, divided by the standard deviation of the dependent variable. It converts the regression coefficients into units of sample standard deviations.

Table 3A: Baseline Model

Dep.var.: Aggregate bilateral trade

	(S1)	(S2)	(S3)	(S4)	(S5)	(S6)	(S7)	(S8)	(S9)	(S10)	(S11)
Joint FTA membership	0.081	0.005	0.127**	0.076	0.033	0.089*	0.248***	0.084*	0.084*	0.026	0.017
	(0.052)	(0.050)	(0.051)	(0.050)	(0.046)	(0.052)	(0.071)	(0.051)	(0.051)	(0.055)	(0.059)
Ln Distance	-0.986***	-0.941***	-1.010***	-0.902***	-0.746***	-0.557***	-0.265**	-0.961***	-0.962***	-0.943***	-0.956***
	(0.087)	(0.088)	(0.085)	(0.086)	(0.072)	(0.11)	(0.11)	(0.083)	(0.084)	(0.083)	(0.082)
Adjacency	0.412***	0.149	0.148	0.390***	0.114	0.603***	0.584***	0.412***	0.412***	0.319**	0.346**
	(0.14)	(0.14)	(0.16)	(0.14)	(0.11)	(0.12)	(0.15)	(0.14)	(0.14)	(0.13)	(0.13)
Common Legal Origin		0.649***									
		(0.076)									
Common Language			0.739***								
			(0.21)								
Religious Proximity				0.554***							
				(0.15)							
Ethnic Ties					0.258***						
					(0.028)						
Genetic Distance						-0.003**					
						(0.0014)					
Linguistic Proximity							0.616***				
							(0.16)				
ESC_{ij}								0.034***		0.022***	
								(0.0054)		(0.0052)	
ESC_{ii}									0.0327***		0.017***
									(0.0057)		(0.0058)
Mill's Ratio										5.774***	5.563***
										(0.63)	(0.63)
Observations	10560	10560	10560	10560	10455	6074	3584	10560	10560	4833	4833
Adjusted R^2	0.89	0.90	0.90	0.89	0.91	0.92	0.95	0.89	0.89	0.90	0.91

Standard errors (in brackets) have been adjusted for clustering around country pairs. *, **, *** denote statistical significance at 10%, 5% and 1% level. respectively. Each regression contains a constant and a comprehensive set of importer × year and exporter × year interaction of fixed effects (not shown).

about 0.24 (depending on the exact model), cultural proximity turns out an important determinant of bilateral trade volumes. In specifications (S2) to (S4), the distance elasticity is roughly the same as in specification (S1). Thus geographical distance is not a substitute measure for the measures of cultural proximity considered.

Specification (S5) uses ethnic ties as a measure of cultural proximity. The effect of ethnic ties is somewhat smaller that for the dummy variables discussed above, but is again statistically significant. Interestingly, including that variable reduces the coefficient of geographical distance quite substantially. It also comes with a beta coefficient that is an order of magnitude larger (0.2783). We conclude that in contrast to the measures employed in specifications (S2) to (S4), ethnic ties convey effects that are similar to distance.

In specification (S6), we find a negative and significant effect of genetic distance on bilateral exports. As with ethnic ties, the associated beta coefficient is above 0.2 and the distance coefficient is substantially lower than unity. However, the underlying sample is smaller. Specification (S7) uses a continuous variable to measure linguistic proximity between two countries. This reduces the sample to a third, and dramatically cuts the distance coefficient.

Specifications (S8) to (S11) report our results, when ESC scores are used as measures of cultural proximity. Specifications (S8) and (S9) use all ESC realizations, including zero scores. The two regressions differ only with respect to the direction of the scores: in specification (S8), ESC_{ij} measures the score given to j by i, while in specification (S9), ESC_{ji} measures the score given to i by j. One might conjecture that bilateral imports of i from j should depend more strongly on ESC_{ij} than on ESC_{ji}, but this expectation does not materialize. The regressions leave the distance coefficient fairly unchanged.

In specifications (S10) to (S11), we report Heckman-type baseline regressions. We use only strictly positive ESC scores. This strategy cuts the number of available observations from 10,560 to 4,833. The Mill's ratio is significantly different from zero and strictly positive, which signals that failing to control for sample selection would overestimate the ESC scores' coefficients. Both Heckman-type regressions yield coefficients for cultural proximity that are higher for ESC_{ij} than for ESC_{ji}, according with intuition. Beta coefficients for ESC scores are 0.029 and 0.023, respectively.

We have computed *ad valorem tariff equivalents* (AVTE), associated to a change from the lowest to the highest sample realization of the respective variable, assuming an elasticity of substitution of 5.[24] That number ranges from 18.4% for the common language dummy to 13.85% for religious proximity. Hence, sharing the same language is equivalent to a tariff reduction of

[24]We report *average* effects, and therefore simply compute the ad valorem tariff equivalents as the estimated coefficient divided by $\sigma - 1$.

about 18 percentage points. The AVTEs for the Esc_{ij} scores implied by specifications (S10) and (S11) are 8.48% and 5.48%, respectively. This is lower than for the conventional measures, pointing to substantial attenuation bias.

Table 3B reports the effect of cultural proximity on differentiated and homogenous goods. Each line in the table corresponds to two separate regressions. For sake of clarity, we do not report the full estimation results, which are available upon request but report the estimates and the ad valorem tariff equivalents of the cultural proximity variables.

Specification (S1) to (S6) use six conventional measures of cultural proximity. With the sole exception of the common language dummy, the point estimates are smaller for homogeneous goods than for differentiated ones. The AVTE calculations assume that the elasticity of substitution, σ is 5 for differentiated goods, and 20 for homogeneous goods (Anderson and van Wincoop, 2004). We find that the AVTEs that are of an order of magnitude smaller (and often close to zero) for homogeneous goods. Moreover, the effects are typically larger for differentiated goods than for aggregate trade (compare to Table 3A). This is a comforting result that accords well with intuition: cultural proximity matters little when transactions are performed anonymously on organized exchanges.

As in Table 3A, the results for the ESC scores are obtained on a smaller sample than those for the conventional measures. However, the ESC scores have strong effects on bilateral trade in differentiated goods, with AVTEs of 6.5% and 4.7%, respectively. There is no statistically discernible effect for homogenous goods. Moreover, the scores given from i to j matter substantially more for imports of i from j than the scores given from j to i. This is in line with the fact that ESC scores are imperfectly reciprocal.

4.2 Disentangling the trade cost and preference channels

4.2.1 Identification 1: The dyadic fixed effects gravity (DFEG) model

We present the results of our first identification strategy in Tables 4A and 4B. That strategy draws on Assumption 1, which states that the trade-cost related component of cultural proximity is time-invariant. In order to quantify the trade costs and preference channel, we proceed in two steps. First, we estimate equation (11) on aggregate bilateral trade data, including only strictly positive ESC scores and using a Heckman estimation technique. Second, we include the dyadic fixed effects in order to tease out the trade costs effect on trade. In Table 4A we use aggregate bilateral trade as the dependent variable, while in Table 4B we use trade in differentiated and homogeneous goods.[25]

[25] Notice that specification (S1) and (S3) in Table 4A corresponds respectively to the specification (S10) and (S11) in Table 3A.

Table 3B: Baseline Model

Dep.var.: Ln value of bilateral trade in differentiated versus homogeneous goods

		Differentiated goods $\sigma = 5$		Homogeneous goods $\sigma = 20$	
		Coef.	AVTE (percent)	Coef.	AVTE (percent)
(S1)	Common Legal Origin	0.637***	15.93	0.553***	2.91
		(0.085)		(0.12)	
(S2)	Common Language	0.753***	18.83	0.906***	4.77
		(0.15)		(0.23)	
(S3)	Religion Proximity	0.784***	16.66	0.468**	2.09
		(0.16)		(0.22)	
(S4)	Ethic Ties	0.281***	7.03	0.137***	0.72
		(0.026)		(0.043)	
(S5)	Genetic Distance	-0.836	–	-0.115	–
		(0.02)		(0.0023)	
(S6)	Linguistic Proximity	0.698***	17.45	0.122	–
		(0.18)		(0.38)	
(S7)	$Esc_{ij}^{(\ddagger)}$	0.0203***	5.08	0.00213	–
		(0.0070)		(0.011)	
(S8)	$Esc_{ji}^{(\ddagger)}$	0.0175**	4.37	0.00226	–
		(0.0077)		(0.010)	

AVTE refers to ad valorem tariff equivalents. Each row reports on one specific regression. Standard errors (in brackets) have been adjusted for clustering around country pairs. *, **, *** denote statistical significance at 10%, 5% and 1% level respectively. Each regression contains the FTA and adjacency dummies, ln Distance, a comprehensive set of importer × year and exporter × year fixed effects and a constant. Results shown in row (S1) through (S6) are obtained using OLS; (S7) and (S8) are obtained using the two-stage Heckman procedure described in the text. Numbers of observations are 7,826 and 7,161 for differentiated and homogeneous goods, respectively, except in (S2): 2,901 and 2,870 observations, respectively, and (S4): 5,033 and 4,881, respectively. Regressions (S7) and (S8) use 3,620 (differentiated goods) and 3,345 (homogeneous goods) observations, respectively. R^2 ranges from 0.73-0.78 for homogeneous goods, and from to 0.89 to 0.95 for differentiated goods.

Table 4A: Dyadic Fixed Effects Gravity Model

Dep.var.: Ln value of aggregate bilateral trade

	Imports		Exports	
	(S1)	(S2)	(S3)	(S4)
ESC scores	0.022***	0.007***	0.017***	0.001
	(0.0052)	(0.0024)	(0.0058)	(0.0023)
Both countries in same FTA	0.026	0.133***	0.017	0.095**
	(0.055)	(0.042)	(0.059)	(0.040)
Ln Distance	-0.943***		-0.956***	
	(0.083)		(0.082)	
Adjacency	0.319***		0.346**	
	(0.13)		(0.13)	
Mill's ratio	5.774***	-0.115	5.563***	-0.098
	(0.63)	(0.27)	(0.63)	(0.20)
Country pair FE	NO	YES	NO	YES
Observations	4833	4833	4833	4833
Adjusted R^2	0.90	0.81	0.90	0.81
RMSE	0.760	0.301	0.759	0.298
Number of pairs		797		797

Standard errors into brackets. Standard errors have been adjusted for clustering around country pairs. *, **, *** denote statistical significance at 10%, 5% and 1% level respectively. Each regression contains a constant, and a comprehensive set of importer x year and exporter x year fixed effects.

In Table 4A, the total effect of cultural proximity as captured by the ESC scores is 0.022. Specification (S2) reports the effect of including dyad-specific (directed) fixed effects.[26] The ESC coefficient remains significant at the 1% level of significance, but drops to 0.007. To the extent that Assumption 1 is correct, this coefficient measures the preference channel of cultural proximity. As the total effect reported in specification (S1) may be biased upwards[27] for aggregate data, the preference channel turns out to command *at least* 31.8% of the total effect (0.007/0.022).

Specifications (S3) and (S4) in Table 4A provide a natural check of our identification strategy. We know that ESC scores are asymmetric (but positively correlated) within country pairs ($0 < corr(\Pi_{ijt}, \Pi_{jit}) < 1$). Imports of i from j are increased by either Π_{ijt} through the trade cost and/or the preference channels. Exports, however, are increased by Π_{ijt} only through the trade cost channel. They could be affected by the preference channel if Π_{ijt} and Π_{jit} were sufficiently

[26]i.e., a different fixed effect is estimated depending on the direction of the trade relationship (from i to j or from j to i.

[27]See section 3.3

strongly correlated. Hence, if Assumption 1 is true, the effect of ESC_{ijt} (as a proxy of Π_{ijt}) on *exports* should be comparable to its effect on *imports*. However, the preference channel should be much smaller for exports than for imports.[28] Our results show that this expectation bears out, supporting the validity of our identification assumption.

Table 4B repeats the exercise performed in Table 4A using Rauch's conservative classification of goods (Rauch, 1999). We present the results on aggregates of differentiated goods in specification (S1) to (S4). Specifications (S5) to (S8) present the results for aggregates of homogeneous goods. Concerning differentiated goods, the preference effect accounts for *at least* 30.6% of the total effect (0.00623/0.0203). Again, and comfortingly, the preference channel is absent for exports.

We do not find any effect of cultural proximity on import and export of homogeneous goods. However, it is interesting to note that the coefficient of geographical distance seems larger for homogeneous goods compared to differentiated goods. Hence, the failure of ESC_{ijt} to pick up the trade cost effect might be due to the fact that distance captures the relevant aspects already. More interestingly, while we do not find any effect of FTA membership for trade in differentiated goods, we find an effect for trade in homogeneous goods. One reason might lie in the fact that protectionist policy makers cannot undo the elimination of tariffs on homogeneous goods as easily as on differentiated goods by introducing non-tariff barriers.

4.2.2 Assumption 2: The bilateral trade balance (BTB) specification

Finally, we turn to the bilateral trade balance equation (13). This equation exploits the assumption 2, which states that trade costs are symmetric between imports and exports. That assumption is arguably weaker than the assumption 1, hence we prefer it.

We present the results of the BTB model in Table 5. There are only 903 bilateral relationships for which both ESC_{ijt} and ESC_{jit} are strictly positive. The overall fit of the model is satisfactory (in particular for differentiated goods), it nevertheless provides a less tested shell for investigation than the standard gravity equation. For each trade classification, we present two specifications, one with country pair fixed effects and one without. As argued in section 3.3, we may account for habit formation by including dyad-specific fixed effects. This turns out to be important quantitatively.

Specification (S1) and (S2) report the results from using aggregate bilateral trade as dependent variable. In specification (S1), the preference channel effect is 0.010 and statistically different from zero at the 5% level of significance. However, when dyad-specific fixed effects are added, the estimate of the preference channel falls somewhat to 0.0079, remaining significant

[28] In fact, it should be zero if the theory is correct.

Table 4B: Dyadic Fixed Effects Gravity Model

Dep.var.: Ln value of bilateral trade in differentiated versus homogeneous goods

	Differentiated Goods				Homogeneous Goods			
	Imports		Exports		Imports		Exports	
	(S1)	(S2)	(S3)	(S4)	(S5)	(S6)	(S7)	(S8)
ESC scores	0.0203***	0.00623***	0.0175**	0.00101	0.00213	-0.00402	0.00226	-0.00878
	(0.0070)	(0.0024)	(0.0077)	(0.0027)	(0.011)	(0.0061)	(0.010)	(0.0061)
Both countries in same FTA	-0.0128	0.0169	-0.0482	0.0256	0.436***	0.320***	0.380***	0.397***
	(0.061)	(0.044)	(0.063)	(0.042)	(0.12)	(0.11)	(0.13)	(0.12)
Ln Dist	-0.712***		-0.703***		-0.917***		-0.967***	
	(0.095)		(0.092)		(0.13)		(0.13)	
Adjacency	0.569***		0.610***		1.038***		0.929***	
	(0.13)		(0.13)		(0.19)		(0.20)	
Mill's ratio	4.460***	-0.0581	4.267***	0.108	5.714***	-0.0444	3.720***	-0.342
	(0.55)	(0.20)	(0.52)	(0.22)	(0.94)	(0.57)	(0.85)	(0.54)
Country pair FE	NO	YES	NO	YES	NO	YES	NO	YES
Observations	3,064	3,064	3,064	3,064	3,064	3,064	3,064	3,064
Adjusted R^2	0.91	0.89	0.91	0.88	0.75	0.44	0.77	0.43

Standard errors (in brackets) have been adjusted for clustering around country pairs. *, **, *** denote statistical significance at 10%, 5% and 1% level respectively. All regressions contain a constant and exporter x year and importer x year fixed effects.

at the 10% level. Both estimates are roughly in line with the results obtained from estimating equation (11). Given the estimated coefficient of the ESC scores in specification (S10) of Table 3A, they signal that the preference effect makes up about 35% of the total trade creation due to cultural proximity. This finding is in line with the result found using the first identification strategy.

Table 5: Bilateral Trade Balance Model

Dep.var.: Ln value of bilateral trade

	All Goods		Differentiated Goods		Homogeneous Goods	
	(S1)	(S2)	(S3)	(S4)	(S5)	(S6)
$ESC_{ijt} - ESC_{jit}$	0.00994**	0.00786*	0.00864	0.0187*	0.0752	0.0181
	(0.0053)	(0.0048)	(0.006)	(0.010)	(0.044)	(0.025)
Inverse Mills i	1.084**	-0.892	1.588*	-0.781	1.186	1.340**
	(0.79)	(0.95)	(0.91)	(0.59)	(3.09)	(0.66)
Inverse Mills j	-1.390	-0.583	-2.386	-0.650	3.466	-1.505
	(0.99)	(1.35)	(1.75)	(1.17)	(2.79)	(1.78)
Country pair FE	NO	YES	NO	YES	NO	YES
Observations	903	903	588	588	588	588
Adjusted R^2	0.45	0.55	0.73	0.90	0.49	0.94
Number of clusters		330		201		201

Standard errors (in brackets) have been adjusted for clustering around country pairs. *, **, *** denote statistical significance at 10%, 5% and 1% level respectively. All regressions contain a constant and comprehensive sets of importer x year and exporter x year fixed effects.

Specification (S3) and (S4) use the bilateral trade position in differentiated goods as dependent variable. Here dyad-fixed effects turn out to be important in producing a statistically significant preference effect. The effect in specification (S3) is comparable in size to the preference effects detected earlier but does not turn out to be statistically significant. In specification (S4), which includes dyadic fixed effects, there is a substantially larger preference effect, which is estimated with some imprecision.

Finally, specification (S5) and (S6) report the results for homogeneous goods. The differential in ESC scores never turns out to affect the bilateral trade balance. Hence, the finding already reported in Table 4B is robust. There is no evidence for a preference effect in bilateral trade of homogeneous goods. While the bilateral trade balance model requires a weaker assumption than the dyadic fixed effects gravity model, it is more easily susceptible to measurement error in the data. However, the results obtained from the trade balance model are largely in line with those found in Table 4B.

4.3 Robustness checks

We have conducted a number of robustness checks, which are reported in detail in an earlier version of this paper (Felbermayr and Toubal, 2006). Here we restrict ourselves to briefly discuss the most important checks.[29]

First, we have tried to correct for song quality using a two-stage approach. In the first stage, we use a zero-inflated negative binomial model to regress scoring outcomes on the outcome of Google counts with the song's title and the performing artist/group as search elements, and on a host of observable song characteristics. In the second stage we use the residuals of this regression as quality adjusted measures of cultural proximity. It turns out that this procedure delivers results very much in line with those reported in Table 4A.

Second, we have divided the data into three subsamples 1975-1985, 1986-1995, and 1996-2003. Comfortingly, we find again results in line with those reported in Table 4A. This is important, since Assumption 1 is more likely to hold over shorter time periods. Finally, we cut the data into two subsamples along the data of the introduction of televoting in 1998. Again, the principal results reported in Table 4B remain unaffected. However, the estimated coefficient of cultural proximity increases slightly, signaling the reduction of measurement error as televoting reflects cultural preferences more closely.

5 Conclusions

Standard measures of cultural proximity, such as common language, common religion, genetic proximity, etc., do not allow to disentangle the channels through which bilateral trade volumes are affected: namely, through trade costs and preferences. The first effect obtains when a higher degree of cultural proximity makes culture-based misunderstandings less likely, leads to more trust, and therefore lowers transaction costs. The second effect appears when patterns of cultural proximity across countries are correlated with country-specific weights in the utility function. Both effects lead to trade creation, but only the former is relevant for welfare. In the present paper, we have studied the effect on cultural proximity on trade and isolated the channels through which it matters.

We exploit data from the Eurovision Song Contest (ESC). The ESC is a huge, televised show, in which European countries grade the songs of other countries. We argue that the matrix of bilateral votes, obtained in the ESC grading process, can be used as dyadic, time-variant information on European countries' cultural proximity. We control for song quality by

[29] Note that in Felbermayr and Toubal (2006), ESC scores are defined with the opposite direction as in the present paper.

including song-specific fixed effects and deal with the peculiarities of the ESC rules (not all countries are actually ranked). ESC grades correlate strongly with conventional measures of cultural proximity.

ESC scores have properties that allow to identify the trade-cost and the preference effects of cultural proximity. Assuming that the trade-cost related component of cultural proximity is largely time-invariant, the time dimension of the ESC data allows to separately identify the preferences effect. The validity of our identification strategy can be tested by exploiting the lack of systematic reciprocity in ESC scores. Our main results are corroborated in a second identification strategy, where we assume that trade costs are symmetric across a pair of countries, while preferences need not be.

We find that the total average trade-creating effect of moving cultural proximity from its lowest to its highest sample values amounts to a tariff equivalent of about 8% for total trade, and of about 6% for trade in differentiated goods. There is no trade-creating effect of cultural proximity with homogeneous goods. At least a third of the boost in trade is attributable to the preference effect.

Our evidence of a statistically and economically sizable preference effect of cultural proximity establishes an important conceptual difference between geographical proximity, which does not affect preferences, and cultural proximity. Remote countries, such as New Zealand or Australia, have more trade than their geographical location would suggest. However, to the extent that their extra trade is due to preference-driven trade creation, they are still disadvantaged by higher trade costs.

A Summary Statistics and Data Sources

Table A1: *Summary Statistics to Table 3A*

	Models S(1)-S(4), S(8), S(9) N=10,560				Models S(10), S(11) n=4,833			
	Mean	S.d.	Min	Max	Mean	S.d.	Min	Max
Ln value of agg. imports	19.10	2.54	7.08	25.08	19.36	2.43	9.31	24.84
ESC(ij)	2.67	3.40	0.00	10.00	5.40	2.87	1.00	10.00
ESC(ji)	2.67	3.40	0.00	10.00	5.40	2.87	1.00	10.00
Joint FTA membership	0.31	0.46	0.00	1.00	0.32	0.47	0.00	1.00
Ln distance	7.31	0.63	4.39	8.56	7.26	0.66	4.39	8.56
Adjacency	0.09	0.29	0.00	1.00	0.10	0.30	0.00	1.00
Common legal origin	0.18	0.39	0.00	1.00				
Common language (0,1)	0.06	0.24	0.00	1.00				
Religious proximity	0.21	0.25	0.00	0.85				
Ethnic ties	7.82	2.74	0.00	15.43				
Inverse Mills ratio (i)					0.59	0.12	0.30	0.79
Inverse Mills ratio (j)					0.59	0.12	0.30	0.79

Table A2: *Summary Statistics to Table 3B*

	Differentiated goods				Homogeneous goods			
	Mean	Std. Dev.	Min	Max	Mean	Std. Dev.	Min	Max
Model S(1)-S(3)								
Ln value of trade	18.49	2.41	8.01	24.28	16.61	2.40	6.91	23.07
Common language (0,1)	0.06	0.24	0.00	1.00	0.07	0.25	0.00	1.00
Ethnic ties	7.76	2.51	0.00	14.23	8.04	2.35	0.00	14.23
Religious proximity	0.22	0.26	0.00	0.85	0.22	0.26	0.00	0.85
Common legal origin	0.19	0.40	0.00	1.00	0.20	0.40	0.00	1.00
Joint FTA membership	0.33	0.47	0.00	1.00	0.35	0.48	0.00	1.00
Ln distance	7.32	0.62	4.39	8.56	7.29	0.62	4.39	8.56
Adjacency	0.09	0.28	0.00	1.00	0.09	0.29	0.00	1.00
Model S(7), S(8)								
Ln value of trade	18.64	2.35	9.21	24.28	16.70	2.43	6.91	23.07
ESC(ij)	2.87	3.49	0.00	10.00	5.44	2.86	1.00	10.00
ESC(ji)	5.41	2.85	1.00	10.00	2.93	3.50	0.00	10.00

a) N=7,826 for differentiated goods, and N=7,161 for homogeneous goods.
b) N=3,620 for differentiated goods, and N=3,356 for homogeneous goods.

Table A3: *Summary Statistics to Table 4A*

	Mean	Std. Dev.	Min	Max
Ln value of agg. imports	19.36	2.43	9.31	24.84
ESC(ij)	5.40	2.87	1.00	10.00
ESC(ji)	2.91	3.50	0.00	10.00
Inverse Mills ratio (i)	0.59	0.12	0.30	0.79
Inverse Mills ratio (j)	0.52	0.14	0.29	0.79
Joint FTA membership	0.32	0.47	0.00	1.00
Ln distance	7.26	0.66	4.39	8.56
				N=4,833

Aggregate imports. Data on aggregate bilateral import flows (measured c.i.f.) are taken

Table A4: *Summary Statistics to Table 4B*

	Mean	Std. Dev.	Min	Max
Ln imports, differentiated goods	19.12	2.13	11.42	24.28
Ln imports, homogeneous goods	16.90	2.40	6.91	23.07
Ln imports, reference-priced goods	18.35	1.98	10.31	23.17
ESC(ij)	5.42	2.84	1.00	10.00
Inverse Mill's ratio (i)	0.60	0.12	0.29	0.79
Joint FTA membership	0.36	0.48	0.00	1.00
Ln distance	7.24	0.66	4.09	8.30
				N=3,064

Table A5: *Summary Statistics to Table 5*

	Mean	Std. Dev.	Min	Max
Models S(1) and S(2), N=903				
Ln value of aggregate imports	-0.10	0.94	-5.74	3.17
ESC(ij) - ESC(ji)	0.13	3.92	-9.00	9.00
Inverse Mill's ratio (i)	0.60	0.12	0.30	0.79
Inverse Mill's ratio (j)	0.59	0.12	0.31	0.80
Joint FTA membership	0.33	0.47	0.00	1.00
Models S(3)-S(6), N=588				
Ln imports, differentiated goods	-0.04	1.12	-4.58	4.18
Ln imports, homogeneous goods	-0.22	1.91	-6.75	6.14
ESC(ij) - ESC(ji)	0.08	4.00	-9.00	9.00
Inverse Mill's ratio (i)	0.60	0.12	0.30	0.80
Inverse Mill's ratio (j)	0.59	0.12	0.31	0.79
Joint FTA membership	0.37	0.48	0.00	1.00

from the IMF's Direction of Trade Statistics (DoTS) CDROM (July 2006 edition) and cover the period 1975-2003.

Trade in differentiated and homogeneous goods. The data has been constructed by aggregating SITC4 rev. 2 trade data prepared by Feenstra et al. (2005) according to Rauch's (1999) 'conservative' classification of 4 digit SITC goods. The trade data is measured c.i.f. and covers the period 1975-2000.

Adjusted ESC scores. Adjusted ESC scores are the residual of a regression of raw ESC scores on a set of song-specific fixed effects.

Conventional measures of cultural proximity. The common language dummy takes value 1 if a language is spoken by at least 9% of the population in both countries. It is available from CEPII in Paris. The continuous variable 'linguistic proximity' from Dyen et al. (1992) is available only for a limited set of indogermanic languages. The measure 'common legal origin' is taken from La Porta et al. (1999) and distinguishes between a Latin, a Scandinavian, a British, and a Germanic system.

Following Alesina et al. (2003), 'religious proximity' is constructed as follows: Using mid-1990ies Census data[30], we compute the shares s_{ri} of adherents to religion r (Catholics, Protestants,

[30] The data are freely available on www.worldchristiandatabase.org and coincide closely with entries in the CIA fact book and other sources.

orthodox Christians, Muslims, Jews, atheists, and non-religious persons) in total population of country i. Then, we compute a bilateral index $\rho_{ij} = \sum_r s_{ri} s_{rj}$, with $\rho_{ij} \in [0, 1]$. This measure reaches a maximum value of 0.85 for the dyad Poland-Malta, and is minimum, 0.001, for the pair Poland-Turkey.

As a measure of ethnic ties, we use the stock of foreign born individuals by country of birth, provided by the OECD and discussed in Dumont and Lemaitre (2005). Finally, we use data on genetic distance computed by Cavalli-Sforza (1994) and made available by Paola Giuliano.

Geographical data. Data on geographical distance and on adjacency is taken from the CEPII in Paris.[31]

Free trade agreements. Data is from Baier and Bergstrand (2006) and updated for account for the pre-EU membership Europe agreements signed by East European countries with the EU (WTO website).

B Time behavior of multilateral ESC scores

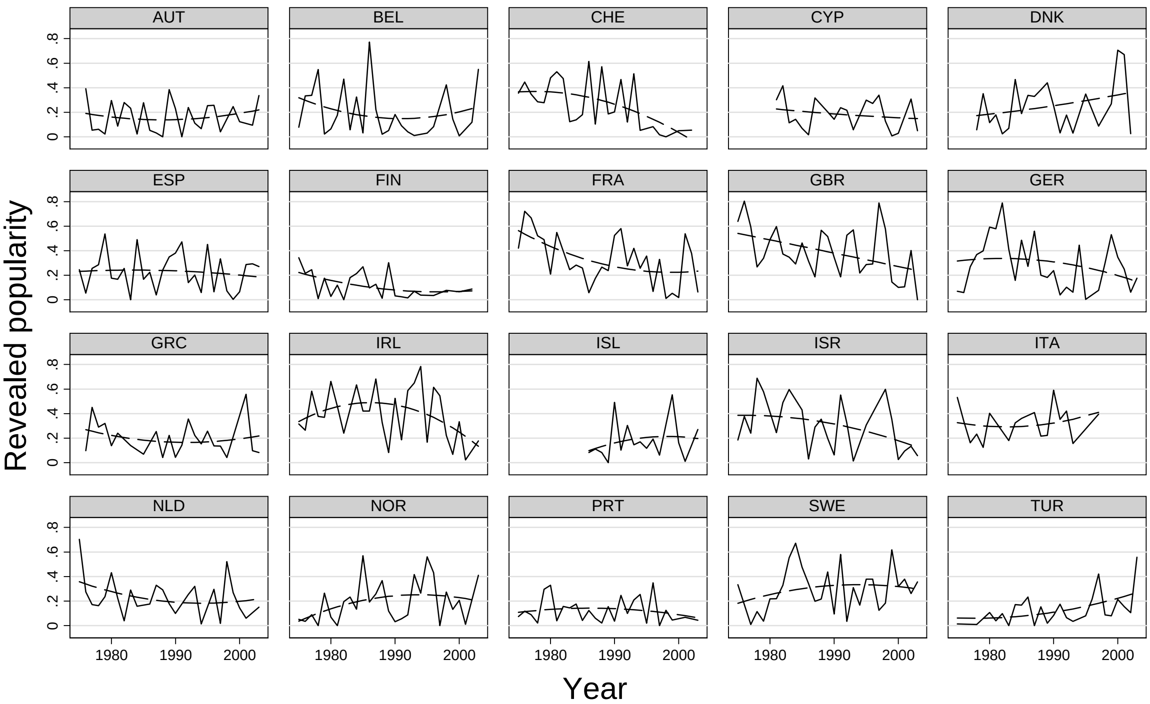

Figure 1: Time profiles of multilateral scores obtained by countries.

[31] The CEPII data can be downloaded from `www.cepii.fr/anglaisgraph/ bdd/distances.htm`.

References

[1] Alesina, A., Dollar, D., 2002. Who Gives Foreign Aid to Whom and Why. Journal of Economic Growth 5, 33-64.

[2] Anderson, J.E., Van Wincoop, E., 2003, Gravity with Gravitas: A Solution to the Border Puzzle. American Economic Review 03(1): 170-192.

[3] Anderson, J.E., Van Wincoop, E., 2004, Trade Costs. Journal of Economic Literature 32(3): 691-751.

[4] Baier, S.L., Bergstrand, J.H., 2006, Do Free Trade Agreements Actually Increase Members' International Trade? Journal of International Economics, forthcoming.

[5] Baltagi, B.H., Egger, P., Pfaffermayr, M., 2003, A Generalized Design For Bilateral Trade Flow Models. Economics Letters 80: 391-397.

[6] Bisin, A., and Verdier, T., 2005. Cultural Transmissions. New Palgrave Dictionary of Economics.

[7] Borjas, G., 1987. Self-Selection and the Earnings of Immigrants. American Economic Review: 531-553.

[8] Boisso, D., Ferrantino, M., 1997, Economic Distance, Cultural Distance, and Openness in International Trade: Empirical Puzzles. Journal of Economic Integration 12(4): 456-484.

[9] Casella, A., Rauch, J., 2003. Overcoming Informational Barrier to International Resource Allocation. Prices and Ties. Economic Journal 113(1): 21-42.

[10] Cavalli-Sforza, L., P.Menozzi, and A. Piazza, 1994, The History and Geography of Human Genes. Princeton, New Jersey: Princeton University Press.

[11] Clerides, S., Stengos, T., 2006. Love thy Neighbor, Love thy Kin: Voting Biases in the Eurovision Song Contest. Mimeo: University of Cyprus.

[12] Combes, P-P., M. Lafourcade and T. Mayer, 2005. The Trade Creating Effects of Business and Social Networks: Evidence from France, Journal of International Economics 66:1-29.

[13] Davis, D.R., Weinstein, D.E., 2002. The Mystery of Excess Trade (Balances). American Economic Review 92(2): 170-174.

[14] Davis, D., Weinstein, D., 2003. Market Access, Economic Geography and Comparative Advantage: An Empirical Assessment," Journal of International Economics 59(1): 1-23.

[15] Disdier, A.C., Mayer, T., 2006. Je T'aime, Moi Non Plus. Bilateral Opinions and International Trade. European Journal of Political Economy, forthocming.

[16] Dumont, J.-C. and Lemaitre, G., 2005, Counting immigrants and expatri- ates: A new perspective. OECD Social, Employment and MigrationWorking Paper 25.

[17] Dyen, I., Kruskal, J.B., and Black, P., 1992. An Indoeuropean Classification: A Lexicostatistical Experiment. Transactions of the American Philosophical Society 82(5).

[18] Feenstra, R. C., 2004. Advanced International Trade. Theory and Evidence. Princeton: Princeton University Press.

[19] Feenstra, R. C., Lipsey, R.E., Deng, H., Ma, A.C., and Mo, H., 2005. World Trade Flows: 1962-2000. NBER Working Paper 11040.

[20] Fenn, D., O. Suleman, Efstathiou, J., Johnson, N.F., 2006. How does Europe Make Its Mind Up? Connections, cliques, and compatibility between countries in the Eurovision Song Contest. Physica A 360(2): 576-598.

[21] Frankel, J.A., Romer, D., 1999. Does trade cause growth? American Economic Review 89(3): 379–399.

[22] Ginsburgh, V., Noury, A., 2004. Cultural Voting: The Eurovision Song Contest. Université Libre de Bruxelles, mimeo.

[23] Ginsburgh, V., 2005. Languages, Genes, and Cultures. Journal of Cultural Econonomics 29(1), 1-17.

[24] Giuliano, P., Spilimbergo, A., and Tonon, G., 2006. Genetic, Cultural and Geographical Distances. IZA Working Paper 2229.

[25] Guiso, L., Sapienza, P., and Zingales, L., 2004. Cultural Biases in Economic Exchange. NBER Working Paper 11005.

[26] Guiso, L., Sapienza, P., and Zingales, L., 2006. Does Culture Affect Economic Outcomes? Journal of Economic Perspectives 20(2): 23-48.

[27] Haan, M.A., Djikstra, G.S., Dijkstra, P.T., 2005. Expert Judgment versus Public Opinion. Evidence from the Eurovision Song Contest. Journal of Cultural Economics 29(1): 59-78.

[28] Hanson, G.H., and Xiang, C., 2004. The Home Market Effect and Bilateral Trade Patterns. American Economic Review 94(4): 1108-1129.

[29] Heckman, James J., 1979. Sample Selection Bias as a Specification Error. Econometrica 47(1):153-161.

[30] Helpman, E., 1987. Imperfect Competition and International Trade. Evidence from Fourteen Industrialized Countries. Journal of the Japanese and International Economies 1: 62-81.

[31] Hutchinson, W., 2006. Linguistic Distance as a Determinant of U.S. Bilateral Trade, 1970-1986. Southern Economic Journal 72(1): 1–15.

[32] La Porta, R., Shleifer, A., Vishny, R.W., Lopez Da Silanes, F., 1998. Law and Finance. Journal of Political Economy 106(6).

[33] Melitz, J., 2002, Language and Foreign Trade, CEPR Discussion Paper 3590.

[34] Manski, C.F., 2000. Economic Analysis of Social Interactions. Journal of Economic Perspectives 14(3): 115-136.

[35] Rauch, J.E., 1999. Networks Versus Markets in International Trade. Journal of International Economics 48(1): 7-35.

[36] Rauch, J.E., Trindade, V., 2002. Ethnic Chinese Networks in International Trade. Review of Economics and Statistics 84(1): 116-130

[37] Redding, S., Venables, A.J., 2004. Economic Geography and International Inequality. Journal of International Economics 62(1): 53-82.

[38] Rose, A.K., 2004. Do We Really Know that the WTO raises trade? American Economic Review 94(1): 98-114.

[39] Spierdijek, L. and Vellekoop, M., 2006. Geography, Culture, and Religion: Explaining the Bias in Eurovision Song Contest Voting. Mimeo: University of Twente.

[40] Spolaore, E., and Wacziarg, R., 2006. The Diffusion of Development. NBER Working Paper 12153.

[41] Straubhaar, J., 2002. (Re)asserting national television and national identity against the global, regional and local levels of world television. In J. M. Chan and B. T. McIntyre (eds.): In search of boundaries. Communication, nation-States and cultural identities. Westport, Connecticut: Ablex Publishing.