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Migration and Regional Trade Agreement: a (new) Gravity Estimation

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Highlights

- Our econometric strategy controls for the multilateral resistance to migration (Bertoli and Fernandez-Huertas Moraga, 2013) and solves the zero migration flows problem by using a censored quantile regression approach.

- Our results suggest that the presence of a RTA stimulates the migration flows among member countries.

- The pro-migration effect of RTAs is magnified if the agreement includes also provisions easing bureaucratic procedures for visa and asylum among member countries.

Abstract

This paper investigates the role of Regional Trade Agreements (RTAs) on bilateral international migration. By increasing the information on the potential destination country, RTAs may favour bilateral migration flows among member countries. Building on the gravity model for migration by Anderson (2011), our econometric strategy controls for the multilateral resistance to migration (Bertoli and Fernandez-Huertas Moraga, 2013) and solves the zero migration flows problem by using a censored quantile regression approach. Further, the endogeneity problem of RTAs in migration settlement is addressed by using IV censored quantile regression (Chernozhukov and Hansen 2008). Our results suggest that the presence of a RTA stimulates the migration flows among member countries. The pro-migration effect of RTAs is magnified if the agreement includes also provisions easing bureaucratic procedures for visa and asylum among member countries. Finally, we find a non-linear effect of RTAs across the quantiles of the distribution of migration settlements.

Keywords

Migration, Gravity Equation, Censored Quantile Regression.

JEL

C13, C23, F22, F13.

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RESEARCH AND EXPERTISE ON THE WORLD ECONOMY

Migration and Regional Trade Agreement: a (new) Gravity Estimation[1]

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1. Introduction

Over the past 40 years, the gravity equation has been the workhorse for trade economists studying the effect of trade liberalization on bilateral trade flows. As recently highlighted by Head and Mayer (2014), given the high flexibility of the gravity micro-foundation, it can be applied to a wide range of other bilateral flows and interactions. Anderson (2011) extended the micro foundation of the gravity model to both FDI and migration flows. This paper focuses on the latter domain and points to use gravity model to explain international migration settlements across countries as a function of migration cost. In particular we focus on the role of Regional Trade Agreements (RTAs) as a possible source of information about destination countries for potential migrants in origin countries (which is expected to reduce the cost of migration). Beyond the information channel, RTAs may affect international migration by simply stimulating trade flows: in a standard factor content model of trade, RTAs boost trade flows and thus reduce wage inequalities among member countries; this reduces the incentive for international migration. If the latter were the prevailing effect, we should observe negative relation between RTAs and bilateral migration flows. However, previous evidence (Orefice 2013) shows that, all other determinants being constant, RTAs and bilateral migration flows are positively related: RTAs drive migration choice towards member countries.

This paper adds to Orefice (2013) in two different ways. First, we apply a fresh econometric technique to address recent problems observed in the Pseudo Poisson Maximum Likelihood (PPML) estimator. Indeed, PPML, being a non linear estimator, over-weighs large bilateral migration flows (Head and Mayer 2014). This is a crucial shortcoming in gravity for migration because very big bilateral migration flows usually involve the same country pairs across years (i.e., Mexico-USA, Morocco-Spain, Poland-Germany). Such specific migration cases are due

[1]We thank Matthieu Crozet for useful comments and suggestions. The views expressed are purely those of the authors and may not under any circumstances be regarded as stating an official position of the institutions they are affiliated to.

to historical and geographic factors and do not depend on time varying bilateral migration costs (i.e. the information cost); for this reason, we expect that the simple PPML estimation provides downward biased coefficient on the effect of RTAs on migration settlement. Moreover, Figueiredo, Lima and Schaur (2014) show that the PPML relies on identification conditions that may not hold in practice, leading to a significant bias in the estimates of the gravity equation. Our second innovation is to enlarge the sample of destination countries covered in the estimation of the gravity model. Thank to the recent World Bank data on bilateral migration stocks, here we cover also developing (destination) countries and thus also south-south migration flows. This is an important feature considered the recent tendency of developing countries in signing RTAs.[2]

The theoretical micro-foundation of the gravity equation for trade provided by Anderson and Van Wincoop (2003) highlighted the importance of controlling for the multilateral price resistance term (MRT) in gravity equation empirics (see Head and Mayer 2014 and Baltagi, Egger and Pfaffermayr 2014 for exhaustive survey of literature on this point). The MRT term can be captured by country specific fixed effects in cross section gravity model estimation, or by country-by-year fixed effects in case of panel gravity equation. In a parallel way, the micro-foundation of gravity equation for migration proposed by Anderson (2011) highlights the importance of migration cost and inward/outward migration frictions in predicting migration flows and their settlement across destination countries.[3] The importance or multilateral resistance to migration has also been highlighted by Bertoli and Fernández-Huertas Moraga (2013) who show that the multilateral resistance to migration is properly captured by country-year fixed effects in a standard random utility maximization (RUM) model for migration where the attractiveness of a destination country does not vary across origin countries (see also Beine, Bertoli and Fernández-Huertas Moraga 2014 on this point). Inward migration frictions – *Inward migration resistance term* - may be imagined as immigration policy restrictiveness; while outward migration frictions – *outward migration resistance term* – can be considered as origin country specific factor deterring emigration.[4] Failing to control for such migration frictions produces an omitted variable problem and consequently leads to biased gravity estimation (as in the case of gravity for trade). Thus, in our empirical gravity for migration we properly capture inward/outward migration frictions by including country-by-year fixed effects.

[2]The list of early announcement - under negotiation - RTAs, provided by the WTO RTAs information system, includes a lot of agreements involving developing countries, such as Canada - Dominican Republic launched the 7th June 2007, EFTA - Indonesia launched the 31th January 2011, EU - Thailand launched 6th March 2013, Ukraine - Singapore launched the 8th May 2007, and many others available here: `http://rtais.wto.org/UI/PublicEARTAList.aspx`.

[3]In Anderson (2011) the decision to migrate as an individual discrete choice model conditioned to country pair specific migration cost, and aggregated migration flows are determined by the relative size of origin and destination country, migration cost and inward/outward migration frictions (see eq. 20 in Anderson, 2011)

[4]Inward and outward migration resistance terms highlighted by Anderson (2011) mirror, to some extend, the more traditional *push* and *pull* factors explaining bilateral migration flows (Hatton, 2005; Mayda, 2010; Grogger and Hanson, 2011).

Although the inclusion of country (or country-by-year) fixed effects contributes to reduce the omitted variable problem, an endogeneity issue may rise due to the reverse causality argument. Ghosh and Yamarik (2004) and Baier and Bergstrand (2007) highlighted the endogeneity of RTAs in the context of gravity for trade.[5] However, the reverse causality argument might also apply in the context of gravity for migration when a RTA is signed to face high bilateral migration flows.[6]

Another econometric issue extensively discussed in the trade literature (but scarcely in the context of migration) is the zero migration flows problem. When the dependent variable has a large share of zeros (as in the case of bilateral migration), standard log-linear OLS estimations are biased for both loss of information and heteroskedasticity (see section 2.2 for more details).

This paper addresses the previous identification issues by using recent developments in instrumental variable censored quantile regressions proposed by Chernozhukov and Hansen (2008), Chernozhukov, Fernandez-Val and Kowalski (2011), Kowalski (2008), and Kowalski (2013). We use data from the World Bank for the 1960-2010 period and estimate the effect of RTAs on the entire distribution of bilateral migration flows. By using a quantile approach, we also test any potential asymmetry in the RTAs' effect on migration (the idea is that RTAs may have a different effect along the distribution of migration stocks). Finally, we study whether the inclusion of visa and asylum provisions in the RTA changes its effects on bilateral migration settlement.

We find that, failing to control for multilateral resistance to migration produces very unstable RTAs coefficients (in particular for OLS and PPML), and introduces a downward bias in the Censored Quantile Models - CQ Model (our preferred estimation). After the inclusion of country-by-year fixed effects the CQ model shows positive and significant RTA coefficient. According to our preferred specification (censored quantile regression procedure with country-by-year fixed effects), the presence of a common RTA stimulates bilateral migration by 36%. Further, the PPML estimator produces a null coefficient for the RTA dummy confirming the idea of downward bias of PPML estimator in gravity model for migration. Indeed, PPML, by over-weighting some specific bilateral migration stocks – which are big independently of the presence of a RTA – introduces a negative bias in the RTA coefficient. After controlling for endogeneity (IV on censored quantile regression model)[7] the effect of RTAs on migration stocks gets bigger, suggesting that RTA dummy is indeed endogenous with respect migration flows. Finally we show that the RTAs effect is increasing over the moments of the distribution: the signature of RTAs stimulates migration stocks by 9% at the 10^{th} percentile of the distribution and by almost 49% at the 90^{th} percentile of the distribution.

[5] Indeed, RTAs formation may be affected by trade flows (to secure current bilateral flows), implying endogeneity of RTAs in the estimation of gravity equations.

[6] In a standard Hecksher-Ohlin framework, trade in goods, by equalizing factor prices, is supposed to deter international migration flows. So in this case, government may decide to boost trade in goods to avoid migration flows.

[7] As an instrument for the RTA dummy, we follow the Baldwin and Jaimovich (2012) idea of domino effect in RTAs formation.

In the next section we describe in details the three possible shortcomings in estimating gravity equation for migration. Section 3 presents the details of our preferred estimator. In section 4 we present our results. Last section concludes.

2. Gravity Model for Migration and Bilateral Migration Costs

In this paper we point to estimate robust gravity model for migration focusing on the effect of bilateral Regional Trade Agreements (RTAs) on bilateral international migration.[8] The idea is that RTAs increase the awareness of the partner country among the potential migrants: the signature of a RTA implies improved diplomatic relations among signatory countries, and thus improves the information for all potential migrants in the destination country. This reduces the bilateral migration costs and, all other determinants being constant, favours migration among signatory countries. We rely on Anderson (2011) for the theoretical foundation of the gravity equation for migration. Here we want to stress the role of bilateral (time varying) migration cost in affecting the pattern of migration.

Bilateral migration costs are dyadic factors deterring migration flows, in particular, such costs may vary across years or being time-invariant. Time-invariant migration costs are country pair specific factors affecting migration, in general one may imagine such cost as geography driven. Linguistic barriers, colonial links, distance (as a proxy for travel cost) and cultural proximity belong to this kind of costs. Time variant migration costs refer to country pair factors that may change across years, information cost and policy barriers belong to this kind of costs. Recently, Bertoli and Fernandez-Huertas Moraga (2012) used bilateral visa policies as a proxy for bilateral policy measures.[9] Here we use data on the contents of RTAs (provisions easing the bureaucratic procedure to obtain a visa) to approximate the policy related cost of migration.

Information cost is a further, important, determinant of migration. The potential migrant chooses his/her destination on the basis of the information he/she has on the destination country. In the literature such information cost has been often approximated by past migration stock (network effect) - as the stock of immigrants in the starting year. However such proxy, being time-invariant, does not properly control for time varying information cost of migration (information about a destination country is likely to change over time). We consider the presence of a common RTA as providing additional information about the destination country to potential migrants in the origin country.

Considering the previous framework, and simply relying on Anderson's (2011) gravity equation,

[8] For a recent and updated practitioners' guide to gravity model for international migration see Beine, Bertoli and Fernández-Huertas Moraga (2014).

[9] Authors focus on Spanish immigration case.

bilateral migration settlements can be expressed as:

$$M_{ij} = \frac{N_i N_j}{N}\left(\frac{\delta_{ij}}{\Gamma_i \Omega_j}\right)^{1-\theta}, \quad (1)$$

where

$$\Gamma_i \equiv \sum_j (\phi_j/\delta_{ij}^{1-\theta})\Omega_j^{\theta-1}, \quad (2)$$

$$\Omega_j \equiv \sum_i (\phi_i/\delta_{ij}^{1-\theta})\Gamma_i^{\theta-1}, \quad (3)$$

and M_{ij} represents the number of migrants living in country i coming from j; N_i and N_j indicate respectively population size in destination and origin country, while N is the total population size. The Γ_i is the appropriate 'average' portion of migration costs borne by country i to all destinations, outward multilateral resistance, and Ω_j is the average portion of migration costs borne by j from all sources, inward multilateral resistance, with $\phi_i = N_i/N$, $\phi_j = N_j/N$. δ_{ij} represents bilateral migration costs.

The equation (1) is exactly analogous to the Anderson and van Wincoop (2003) gravity model for trade, while equations (2) and (3) respectively trace inward and outward multilateral resistance equations for trade (MRT), but applied to migration. Thus, the MRT are not observable but can be inferred along with δ_{ij}. The equation (1) can be split in two parts: $N_i N_j/N$ is the frictionless share of migrants in country j and $(\delta_{ij}/\Gamma_i\Omega_j)^{1-\theta}$ is the effect of migration frictions.

Following the literature, and extending equation (1) to a time variant setting we have to include t subscript and allow both inward and outward resistance term to migration to vary over time. For this reason in our empirical exercise we control for inward/outward resistance term by including country-by-year fixed effects. It must be noticed that the inclusion of country-by-year fixed effects properly control for the multilateral resistance to migration only under the assumption that the attractiveness of the destination country does not vary across origin countries (see Bertoli and Fernandez-Huertas Moraga 2013, and Beine, Bertoli and Fernandez-Huertas Moraga 2014 for more details on this point). However, the inclusion of country-by-year fixed effects to control for the multilateral resistance term to migration has already been used in the recent literature on the determinants of bilateral migration (among others Beine and Parsons 2012; Ortega and Peri 2013) and represents a compelling way to control for the multilateral resistance to migration.

Also bilateral migration costs can be imagined as varying over time. In particular we imagine the following exponential function for migration cost:

$$\delta_{ijt} = FC_{ij}^{\beta_1} VC_{ijt}^{\beta_2}, \quad (4)$$

where FC_{ij} is the time invariant component of bilateral migration cost as cultural diversity between origin and destination country (based on past migration settlement, common language or past colonial relationship) and, geographical cost of migration (such as distance and common border); VC_{ijt} is the time variant cost of migration based on the information cost. β_1 and β_2 are parameters.

In what follows, we capture time invariant migration costs (both geography and culture diversity) by using bilateral distance (travel cost), common language, common border, past colonial link dummies and past stock of migrants. Time variant migration cost here are approximated by RTA_{ijt}, our crucial variable, which is a dummy variable equal to one if country i and j share a common RTA at time t. As a proxy for the bilateral policy related (and time-variant) migration cost we use a dummy variable equal to one if the RTA contains a visa-and-asylum provision. The visa and asylum provision is meant to reduce the bureaucratic cost for a visa among signatory countries, so the inclusion of visa and asylum provision in the RTA is a further incentive for the potential migrant to choose his/her destination among signatory countries.[10] In our specification we include also two other control variables. A WTO dummy, being equal to one if country i and j are both part fo the WTO, is meant to isolate the "preferential" nature of the RTA; and the difference in per capita GDP between origin and destination country to capture the effect of the relative attractiveness of the destination.[11]

Most of the existing studies on the determinants of international pattern of migration use the flows of migration (IMD database) provided by the OECD. This allows to have annual data on migration flows from origin to destination country; but restrains the analysis to a short time period and especially on OECD only destination countries. This would be an important limitation for our purpose since it prevents us to cover south-south migration flows, for which the information channel is plausibly more important. For these reasons we decided to rely on bilateral stock data on migration provided by the World Bank. This reduces the number of observations across time, but guarantees deeper historical perspective (our database starts in 1960) and a wider set of destination countries (with respect to the OECD data). Some previous studies use the differences in stocks as a proxy for (long run) flows, however in this way the dependent variable may assume negative values (due to return migration, deaths and naturalization) creating problems from both estimation and intuition point of with (migration in the Anderson 2011 model is strictly positive). Thus, in this paper we follow Grogger and Hanson (2011) and Llull (2014) and simply use the stock of migrants in country i coming from j at time t.

[10]The presence of a Visa-asylum provision in the RTA is based on a recent mapping of 96 RTAs provided by the WTO and reporting the contents of each RTA. This dataset is not exhaustive of the all RTAs in force, that's why specifications using visa-and-asylum dummy have a reduced number of observations. Such database is available here: `http://www.wto.org/english/res_e/publications_e/wtr11_dataset_e.htm`.

[11]Controlling here for the relative attractiveness of countries is important since, as highlighted by Bertoli and Fernández-Huertas Moraga (2013), the multilateral resistance to migration is properly captured by country-year fixed effects if the attractiveness of a destination country does not vary across origin countries.

In the next section we focus on the importance to properly capture inward and outward migration resistance term and thus on the consequent potential endogeneity problem. Then we move to the zero migration flow problem.

2.1. Endogeneity Problem: Omitted Variable and Reverse Causality

The first crucial econometric issue in estimating gravity model for trade is the endogeneity related to the RTA variable due to omitted variables and reversal causality problems. The two previous problems hold also in a migration based gravity model.

The omitted variables problem comes directly from the theoretical micro-foundation of gravity for migration provided by Anderson (2011), where bilateral migration flows depends on destination and origin country specific migration frictions (outward and inward migration resistance term).[12] Such migration frictions are unknown to the econometrician and thus likely to be omitted in a gravity estimation model; ignoring outward and inward migration resistance terms generates bias in the estimated coefficients. Here we account for outward and inward migration resistance term by including country-by-year fixed effects. By comparing a model including only country fixed effect with a model including country-by-year fixed effects we may also provide an estimation of the bias in omitting outward and inward migration resistance term in gravity for migration.

The reverse causality problem is related to the possibility that RTAs are signed in response to migration pressure. We address this problem by using an instrumental variable approach, which is fully explained in section 3. Our instrumental variable is based on the idea of a domino effect in RTAs formation highlighted by Baldwin and Jaimovich (2012) – and already used in Orefice (2013) – who show that the probability that two countries join in a common RTA is positively affected by the number of RTAs that each potential partner has with the rest of the world (in order to avoid trade diversion effect). Following this idea, we use the total number of RTAs signed by origin and destination country (minus one if they have a RTA in common) to instrument the common RTA dummy. The number of RTAs signed by each country with the rest of the world can be thought as exogenous with respect to bilateral migration flows since having a RTA with a third country does not affect the inflow of migrants from the partner country j. The exclusion restriction for our instrument is related to the fact that having an existing RTA with a third country q does not affect the bilateral specific migration flows ij. In other words, no diversion effect in migration flows is needed for a valid instrument for the RTA dummy. Exclusion restriction is respected since inflows of migrants are not driven by differences in the efficiency level (terms of trade) across origin countries.

[12] Bertoli and Fernandez-Huertas Moraga (2013) define the multilateral resistance to migration as the contradictory effect that the attractiveness of alternative destinations exerts on the determinants of bilateral migration flows.

2.2. The Zero Migration Problem

The presence of a large amount of observations equal to zero in trade flows has been highlighted by Helpman, Melitz and Rubinstein (2008); Santos Silva and Tenreyro (2006); Head and Mayer (2014) and Anderson (2011). Indeed, Anderson (2011) argues that many potential bilateral trade flows are not active. The data presented to the analyst may record a zero that is a true zero or it may reflect shipments that fall below a threshold above zero. Helpman, Melitz and Rubinstein (2008) show that country pairs with zero trade account for about half of the observations. The presence of a high share of observations equal to zero in the dependent variable calls for two needs: (a) considering zero flows since they are important source of information; (b) robust estimation techniques in presence of a dependent variable that takes on zero frequently. The problem of zeros is crucial also in a migration setting where plausibly for a consistent share of country pairs does not exist any migration flow.

In taking logs of zero migration flows we drop such observations and incur in a systematic selection bias (see Head and Mayer 2014). Moreover, the OLS estimation on the log of migration might suffer the heteroskedasticity of the error term (see Santos Silva and Tenreyro, 2006). Among the several estimators proposed in the literature to solve both selection bias and heteroskedasticity of error terms in presence of zero trade flows, there is a fair understanding in the trade literature in considering a Poisson Pseudo Maximum Likelihood (PPML) as the most appropriate estimator in presence of zero flows.[13]

However PPML, being a non-linear estimation in levels, tends to over-weigh big bilateral flows, which is a relevant issue in the case of migration flows. Indeed, the largest migration flows are due to some specific country pairs (destination-origin), for example: (1) USA - Mexico; (2) Germany - Poland; (3) Spain - Morocco. So a PPML estimator will give high weight to some very specific migration flows which have historical and geographical (more than RTAs related) reasons. To avoid this problem we propose here a further estimator which produces unbiased and consistent estimator of RTAs dummy in presence of high share of zero flows and does not over weigh large bilateral flows. Additionally, the proposed estimator correctly accounts for endogeneity, heteroskedasticity and allow us to investigate the effect of RTA on the entire distribution of migration.

3. A New Gravity Estimation

3.1. The Econometric Model

In specific notation, we consider the following exponential model studied by Santos Silva and Tenreyro (2006)

$$M_{ij} = \exp\left(x_{ij}\beta\right)\eta_{ij}, \tag{5}$$

[13]See Head and Mayer (2014) for a survey on the estimation techniques proposed in trade literature to solve the zero trade flow problem.

where, in this paper, M_{ij} is the migration stock from i to j, x_{ij} represents the explanatory variables and the multilateral resistance terms, β is a vector of parameters and, η_{ij} is a non-negative random variable. Ignoring for a while the observations of M_{ij} equal to zero, we can linearize the model by taking logarithms of both sides of the equation to obtain

$$\ln M_{ij} = x_{ij}\beta + \ln \eta_{ij}, \tag{6}$$

where $\ln M_{ij}$ is now defined on the real line $\mathbb{R}$.

Heteroskedasticity can be included in this model by assuming that $\eta_{ij} = \exp\left[(x_{ij}\gamma)\,\varepsilon_{ij}\right]$, where ε_{ij} is i.i.d.. In this case, the above model becomes

$$\ln M_{ij} = x_{ij}\beta + (x_{ij}\gamma)\,\varepsilon_{ij}. \tag{7}$$

This is a location-scale model, in which the covariates x_{ij} affect not only the location (mean) of the conditional distribution of $\ln M_{ij}$, but also its scale and quantiles through $(x_{ij}\gamma)$.

Santos Silva and Tenreyro (2006) argued that if $E\left(\eta_{ij}|x\right) = 1$ (which implies that model (5) is identified) and $\gamma \neq 0$ (meaning that there exists heteroskedasticity), then the log-linear model (7) is severely biased. More recently, Figueiredo, Lima and Schaur (2014) showed that if we do the other way around by first assuming that model (7) is identified (which implies that $E\left(\varepsilon_{ij}|x\right) = 0$), then the exponential model considered by Santos Silva and Tenreyro (2006) will be severely biased. In other words, it is not possible to say which model is biased because we do not observe the identifying condition used by each model. More importantly, by the Jensen's inequality $E\left(\ln \eta_i|x_i\right) \neq \ln\left[E\left(\eta_i|x_i\right)\right]$ meaning that identification of the exponential model (5) does not lead to identification of the log-linear model (7) and vice-versa.

To address the above problem, Figueiredo, Lima and Schaur (2014) proposed using quantile regression to identify both models (5) and (7). Their idea relies on the fact that, unlike the mean function, the quantile function is invariant to monotone transformations. In other words, if $h(\cdot)$ is a nondecreasing function on $\mathbb{R}$, then for any random variable Y, $Q_\tau\left(h(Y)\right) = h\left(Q_\tau(Y)\right)$,where $Q_\tau(\cdot)$ is the $\tau - th$ quantile function. Based on this property, they show that identification of the exponential model leads to identification of the log-linear model and vice-versa without assuming any knowledge about the distribution function of η_{ij}. Because the quantile approach identifies both models, we can focus our estimation on the log-linear model to avoid the problem of overweighting large bilateral migration stocks.

However, as in the trade literature, the main problem related to the log-linear model (6) is a large amount of zero migration flows (see Orefice 2013, Ramos and Suriñach 2013, and among others). For these observations, taking the log of zero will automatically lead the computer to drop them. To address this issue, Dutt, Mihov and Zandt (2011) suggested estimating a Tobit model in which the observed dependent variable is censored at 1, implying that the log of the observed dependent variable was censored at zero. In our notation, this solution corresponds to defined the observed dependent variable as $z_{ij} = \max(1, M_{ij})$ and its log transform as $\ln(z_{ij}) = \max(0, \ln(M_{ij}))$. This implies that $\ln(z_{ij})$ is set equal to zero whenever the original observations are subject to censoring, or, whenever $M_{ij} < 1$. In this paper, estimation and inference is conducted by using a generalization of the Tobit model developed by Powell (1984, 1986) which, unlike the Tobit model, identifies the parameters of interest without imposing normalility and homoskedasticity. Finally, as shown in the subsection 2.1, a successful estimator for the log linear gravity model for migration must also address the endogeneity issue. This is carried out by applying to the Powell estimator the IV moment condition approach developed by Chernozhukov and Hansen (2008) (see Appendix 1 for technical explanations on the econometric modelling).

In sum, the estimating strategy adopted by this paper is able to address all the identifying issues discussed in the previous sections, which includes the presence of a large number of observations equal to zero, overweighting of large migration flows, endogeneity of the RTA, and heteroskedasticity. Moreover, the proposed estimator will allow us to investigate the effect of RTA not only on the mean, but also on the entire distribution of bilateral migration settlements. In what follows, we present the data used in our empirical exercise as well as our main results.

3.2. Data

We use data from World Bank Bilateral Migration Matrix 2010 (Özden et al, 2011). We have data for 200 countries (see Table 1) for a long time period starting in 1960 and ending in 2010 and it provides information on bilateral migration stocks for every 10 years: 1960, 1970, 1980, 1990, 2000 and 2010. As in Ramos and Suriñach (2013), we consider that the stock of migration can be interpreted as a representation of a long-term equilibrium, and that it is probably of higher quality than those that report annual immigrant flow.

Table 1 about here

There is a large number of observations already equal to zero and another large fraction less than one. We followed our strategy of assuming that any observation less than one must be treated as subject to censoring and therefore we rounded them down to zero. Behind this strategy there is the assumption that the existing observations equal to zero are indeed observations less than one that had already been rounded down to zero by the staff member responsible in organizing the data initially. After this adjustment 42% of our sample are zeroes.

The control variables are from the same sources as in Head, Mayer and Ries (2010). We collect the (logarithm of) GDP per capita for the origin and destination countries at time t, the origin's and destination's populations at time t, log of distances between origin and destination countries, and dummy variables indicating whether the countries share a common border, language or colonial link. Additionally, we consider informations about the RTA agreements and about WTO membership, they are equal to 1 in the case of a RTA (or WTO) in force between the origin and destination country and zero otherwise.

We also use dummy variables to indicate whether the RTA includes visas and asylum provision. Such dummy is based on a mapping of only 96 RTAs provided by the WTO, that is the reason why the number of observation dramatically reduces when one includes such a dummy variable in the regression specification (from 62,851 to 4,887 observations).[14]

4. Results

We estimate the gravity model (1) to examine the impact of trade agreements on the conditional mean and the quantiles of the bilateral migration stocks. To this end, we first use pooled data with the robust estimator presented in the previous section, which properly accounts for zero migration values and heteroskedasticity (see Appendix 1 for more details). Then we move to the second set of estimations addressing the potential endogeneity of the RTA dummy by using a IV moment condition approach.

4.1. Baseline Results

In our gravity specification, the RTA dummy is equal to one if origin and destination country share a RTA in the current year. Therefore, a positive coeficient on the RTA dummy means that RTA membership (occurred in the last decade) increases migration settlement among member countries (in the last decade). The behaviour of such coeficient across quantiles allows us to analyse of the effect of trade agreements on the different moments of the migration distribution. In this section our preferred specification is the simple censored quantile regression model (CS model) with no control for endogeneity (for more details see equations (10)-(11) in the appendix section) - we consider endogeneity in the next section.

We further compare this estimator with the PPML estimator to evaluate the bias due to the fact that PPML over-weighs big observations. Then, we compare specifications including country and year fixed effects (columns 1-3 in table 2) with specifications having country-by-year fixed effects (columns 4-8 in Table 2), in this way we have an idea of the potential bias due to the omission of inward/outward resistance term to migration. The OLS estimates are simply meant as a benchmark in which we will only consider the positive values of migration.

Table 2 shows our baseline results. According to our preferred specification, censored quantile regression estimation with country-by-year fixed effects (column 4), we find that signing a RTA stimulates the bilateral settlement of migrant by 36%. The rest of control variables have the expected sign. Distance negatively affects migration settlements, while past migrants stocks, common border and common language increase migration by reducing the bilateral (time invariant) cost of migration. Such results are robust to a bunch of robustness checks and estimators (see columns 1-8).

[14] See: `http://www.wto.org/english/res_e/publications_e/wtr11_dataset_e.htm`.

In columns 7-8 we include a further (potential) component of the time varying bilateral migration cost: a dummy being equal to one if the RTA (shared by origin and destination country) includes also a visa and asylum provision easing the bureaucratic cost of migration. Results in columns (7)-(8) confirm our previous findings on the RTA dummy (stimulating bilateral migration) and show a positive coefficient on visa-and-asylum dummy after controlling for the EU dummy (one if origin and destination countries belong to the EU). In particular, results in column (8) suggest that the presence of a common RTA boosts bilateral migration flows by 39%, such elasticity grows up to 55% if the RTA includes also a provision on visa and asylum.

All in all we can conclude that the signature of a RTA, by providing more information on the destination country, reduces the bilateral (time varying) component of migration costs and thus increase migration flows. Such effect is magnified by introducing in the RTA a provision on visa and asylum which reduces the bureaucratic cost of migration.

Table 2 about here

By comparing results in column (1) and (4) we have an idea of the potential bias due to the omission of proper controls for inward/outward resistance term to migration. After the inclusion country-by-year fixed effects,the coeficient on RTA increases by 50% suggesting the huge bias in estimations that do not control for the multilateral resistance term to migration.

Further, by comparing in turn columns (1) and (2) and then columns (4) and (5) we discover the bias in the PPML estimation. In column (2) and (5), PPML produces null coefficient on RTA dummy, implying zero effect of RTA on bilateral migration settlements. Conversely, censored quantile regression procedure gives positive and significant coefficients on RTA dummy. This is coherent with a downward bias in PPML estimations and also with intuition. Indeed PPML over-weighs some specific bilateral migration stocks (i.e. USA-Mexico, Germany-Poland, etc) that are big for historical and/or geographic reasons and not for sharing a RTA; which is the reason why PPML underestimates the importance of RTAs.

The Figure 1 confirms the goodness of fit of the proposed estimator (censored quantile regression) as compared with the OLS and PPML. We estimated a non parametric kernel for the predicted values of bilateral migration stocks by using all the three estimators, and compare these with the observed distribution of migration stocks.[15] We observe that the PPML's fit results to be more symmetric and with higher mean value than the censored quantile regression. If we consider the observed distribution, we notice that it is asymmetric with a mean value slightly above zero. This makes the censored quantile regression a more appropriate estimator for bilateral migration stocks.

Figure 1 about here

Same kind of evidence if we compare distribution properties of the observed stock of migrants with those coming from PPML, OLS and Censored Quantile estimations - Table 3; where the

[15] The censored quantile regression density was estimated by using the Gaglianone and Lima's (2013) technique.

10^{th} , 50^{th} and 90^{th} percentiles of the observed distribution are very close to those estimated by our Censored Quantile estimator (and very different from the ones estimated by OLS and PPML).

Table 3 about here

4.2. Controlling for Endogeneity of RTAs

As argued above, the RTA dummy may suffer reverse causality problem implying bias in its estimation. So, we aim to apply an IV moment condition approach to the censored quantile regression. As in Kowalski (2013), our endogenous variable is discrete and the IV moment condition approach can be implemented by following the procedure developed by Chernozhukov and Hansen (2008). Kowalski (2008) also showed that both control variable and moment condition approaches perform correctly with discrete endogenous variables.

Our instrument for the RTA dummy is the sum of RTAs that (respectively) origin and destination country has with the rest of the world. Indeed, Baldwin and Jaimovich (2012) show that the probability to have a RTA in common is positively related with the amount of RTAs that each country has with the rest of the world (to avoid trade diversion effects). This instrument is exogenous since RTAs with third countries should not affect bilateral migration settlement (exclusion restriction already discussed in section 2.1).

Table 4 reports the estimation results robust to endogeneity of RTAs. For all the three principal moments of the distribution (25^{th}, 50^{th} and 75^{th} quantiles) having a RTA in common boosts bilateral migration setting. Interestingly, Table 4 shows that the effect of RTA on the 25^{th} of the distribution is statistically lower than the effect on the 75^{th} percentile of the distribution, namely having a common RTA stimulates bilateral migration stocks by 22% for the 25^{th} percentile of the distribution and by 50% for the 75^{th} percentile. This suggests a strong asymmetric effect of RTAs along the distribution of the bilateral migration stocks.

Table 4 about here

To further explore such feature we use our robust estimator to show the RTA effect on various quantiles $\tau \in (0.1, 0.2, ..., 0.9)$ (results in Figure 2). The RTAs effect if monotonically increasing with the moments of the distribution: the higher the stock of bilateral migrants the stronger the effect of a reduction in migration costs through the signature of a RTA. In particular for very small migration stocks belonging to the 10^{th} percentile of the distribution, common RTAs has a small effect on migration (9%), while for large migration stocks belonging to the 90^{th} percentile of the distribution, common RTAs stimulate migration by 49%.

Figure 2 about here

Finally, we study the effect of the inclusion of visa and asylum provision in the RTAs using the IV moment condition approach (to solve the endogeneity issue also for the visa- asylum dummy).

The dummy standing for the presence of visa and asylum provision has been instrumented following the same approach as for the RTA dummy. Results reported in Table 5 show that RTAs still have their positive effect on migration stocks, with such effect magnified if the agreement includes visa and asylum provisions reducing the bureaucracy cost of migration. This last estimations are also meant as robustness check since we further include in turn the EU dummy. Specification in column (2) includes EU dummy and shows that having a common RTA stimulates bilateral migration stocks by 40% which increases up to 62% if the RTA includes also a provision on visa and asylum.

Table 5 about here

5. Concluding Remarks

This paper studies the effect of RTAs on the bilateral settlement of migrants; with the idea of capturing, through the signature of RTAs, the time varying bilateral migration costs.

To this end we estimated a structural gravity for migration model based on the seminal theoretical micro-foundations provided by Anderson (2011). The aim is to fill the gap between trade and migration empirical gravity estimation, by applying in a migration framework all the econometric advances made by recent literature on trade related gravity estimations (Santos Silva and Tenreyro 2006; Head and Mayer 2014; Baltagi, Egger and Pfaffermayr 2014).

First, we solve the zero migration flows problem by using the Powell's (1984, 1986) censored quantile regression, which, according to our results on migration, performs better than both OLS and PPML estimators. Second, we solve the endogeneity issue on RTA dummy by using the IV moment condition censored quantile regression approach. Using such new robust techniques we find two clear cut evidences: (i) RTAs stimulate bilateral migration settlements among member countries; (ii) the previous effect increases if the agreement includes provisions easing bureaucratic procedures on visa and asylum among member countries.

Although the main aim of this paper is to provide a practical toolkit for applied economists interested in estimating robust gravity models on migration, our paper suggests also interesting policy implications. RTAs might be used to regulate bilateral migration flows, and are informative for policy makers, who might use RTAs to increase migration inflows in the case of labour market shortages by leaving unchanged their migration policies. Finally, our results are particularly interesting for policy makers in developing countries. Indeed, RTAs represent for developing countries an opportunity to boost exports (and imports) but also - as showed by our results - a way to easy emigration toward (new) member countries.

6. Appendix 1. Econometric Modelling

In this section, we provice technical details on the econometric modelling used in the paper.

Figueiredo, Lima and Schaur (2014) considered the following generalization of model (5)

$$\begin{aligned} M_{ij} &= \exp\left(x_{ij}\beta\right)\eta_{ij}, \\ \eta_{ij} &= \exp\left[\left(x_{ij}\gamma\right)\varepsilon_{ij}\right], \\ \varepsilon_{ij} &\sim i.i.d.F_{\varepsilon}\left(\mu,\sigma^2\right). \end{aligned} \tag{8}$$

$F_{\varepsilon}(\cdot)$ is an unknown continuous distribution function of ε_{ij}, where $F_{\varepsilon}^{-1}(\tau) = Q_{\tau}(\varepsilon_{ij})$ is the $\tau - th$ quantile of ε_{ij} and $\tau \in (0,1)$. Let $Q_{\tau}(M_{ij}|x_{ij})$ denote the the $\tau - th$ conditional quantile of M_{ij}. Thus, the quantiles of M_{ij} can be written as

$$Q_{\tau}\left(M_{ij}|x_{ij}\right) = \exp\left(x_{ij}\beta\left(\tau\right)\right), \tag{9}$$

where $\beta + \gamma \cdot Q_{\tau}(\varepsilon_{ij})$. For instance, when $\tau = 0.5$, $Q_{\tau}(\varepsilon_{ij})$ becomes $Median(\varepsilon_{ij})$ and $\beta(0.5) = \beta + \gamma \cdot Median(\varepsilon_{ij}) = \beta_{median}$

Now, by the property of equivariance of the quantile function, if one assumes that $Median(\eta_{ij}|x_{ij}) = 1$, then this will imply that $Median(\varepsilon_{ij}) = 0$, then $\beta_{median} = \beta$ and $Median(M_{ij}|x_{ij}) = \exp(x_{ij}\beta)$. Therefore, the median estimator identifies the parameter β.[16]

The important consequence of $Median(\eta_{ij}|x_{ij}) = 1$ implying $Median(\varepsilon_{ij}) = 0$ is that, unlike the OLS or PPML based estimation, the identification of quantiles in the exponential model leads to the identification of quantiles in the log-linear model and vice-versa.[17] In other words, for any $\tau \in (0,1)$, equivariance gives $Q_{\tau}(\ln(M_{ij})|x_{ij}) = \ln[Q_{\tau}(M_{ij}|x_{ij})] = \ln[\exp(x_{ij}\beta(\tau))] = x_{ij}\beta(\tau)$, where $\beta(\tau) = \beta + \gamma \cdot Q_{\tau}(\varepsilon_{ij})$. Because the quantile approach identifies both, the exponential and the log linearized model, we can focus our estimation on the log-linear model to avoid the problem of overweighing large bilateral migration stocks.[18]

However, as in the trade literature, the main problem related to the log-linear model (6) is a large amount of zero migration flows (see Orefice 2013, Ramos and Suriñach 2013, and among others).[19] For these observations, taking the log of zero will automatically lead the computer to drop them. To address this issue, we adopt the solution proposed by Dutt, Mihov and Zandt (2011) who estimated a Tobit model in which the dependent variable is censored at 1, which implies that the log of the dependent variable was censored at zero. In our notation, this solution corresponds to set $\ln(z_{ij}) = \max(0, \ln(M_{ij}))$ and, therefore, $\ln(z_{ij})$ is set equal to zero whenever the original observations are subject to censoring, or, whenever $M_{ij} < 1$.

[16]The median estimator is just a special case of the quantile regression estimator proposed by Koenker and Bassett (1978).
[17]Figueiredo, Lima and Schaur (2014) conducted Monte Carlo simulations to assess the bias of the PPML when its identifying condition does not hold.
[18]Notice that we no longer have the bias problem associated to the log-linear model that was previously pointed out by Santos Silva and Tenreyro (2006).
[19]This problem has been discussed in the trade literature since the early 1980s. See, for instance, Head and Mayer (2014, section 5.2).

However, a Tobit model is used to estimate the conditional mean function and relies strongly on the assumptions of normality and homoskedasticity. The approach proposed in this paper can be seen a generalization of the Tobit model in the sense that it allows us to identify the quantiles of the conditional distribution of $\ln(M_{ij})$ (and M_{ij}) and does not rely on any distributional assumption such as normality and homoskedasticity. Given that $\ln(z_{ij}) = \max(0, \ln(M_{ij}))$, the equivariance property naturally leads to the censored quantile regression model

$$
\begin{aligned}
Q_\tau [\ln(z_{ij}) | x_{ij}] &= \max(0, Q_\tau [\ln(M_{ij}) | x_{ij}]) \\
&= \max(0, x_{ij}\beta(\tau)) . \qquad (10)
\end{aligned}
$$

The censored quantile model (10), devoloped by Powell (1984, 1986), provides a way to do valid inference in Tobin-Amemiya models without distributional assumptions and with heteroskedasticity of unknown form. The Powell's censored quantile regression is defined to maximize the objective function:

$$
L_n(\beta) = -\sum_{i,j=1}^{n} w_{ij}\rho_\tau[\ln(z_{ij}) - \max(0, x_{ij}\beta(\tau))], \qquad (11)
$$

where ρ_τ represents the traditional loss function of quantile regression developed by Koenker and Bassett (1978), w_{ij} is a weight. Chernozhukov and Hong (2003) show that the extremum estimator represented by (11) has optimization problems caused by the nonconvexity of the objective function. A robust solution to optimize this function is provided by Chernozukov and Hong (2003), in which the authors use the Markov chain Monte Carlo (MCMC) method to estimate a pseudo-quadratic objective function such as (11).

Nevertheless, as shown in the subsection 2.1, a successful estimator for the log linear gravity model for migration must address the endogeneity issue. This is carried out by applying the IV moment condition approach developed by Chernozhukov and Hansen (2008), which consist in adding to equation (11) an additional pre-step to handle endogeneity.[20]

In other words, let $x_{ij} = (x_{0,ij}, x_{1,ij})$ where $x_{0,ij}$ is the endogenous variable while $x_{1,ij}$ are exogenous, $b(\tau) = (\alpha_0, \beta_0)$ are the corresponding parameters for a given τ, and $v_{ij} = (x_{1,ij}, v_{0,ij})$, where $v_{0,ij}$ is a set of instrumental variables. Then, the estimating procedure goes as follows:

1. consider $\ln(z_{ij}) - x_{0,ij} \cdot \alpha_0 = x_{1,ij} \cdot \beta_0 + v_{0,ij} \cdot \zeta + \epsilon_{ij}$, with $Q_\tau [\epsilon_{ij} | v_{0,ij}, x_{1,ij}] = 0$. For a given value of $\alpha_k \in (\alpha_1, ..., \alpha_J)$, we run a quantile regression of $\ln(z_{ij}) - x_{0,ij}^T \cdot \alpha_k$ on $(x_{1,ij}, v_{0,ij})$, and compute the Wald statistic (W_n) corresponding to the test of $\zeta(\alpha_j) = 0$. Then we define

[20] Alternatively, we should consider the three-step estimator provided by Chernozhukov and Hong (2002). However, this approach uses the control function approach to endogeneity, and the assumptions necessary for the control function approach are less likely to be satisfied when the endogenous variable is discrete (see Kowalski 2013).

the estimator of α_0 as $\hat{\alpha} = \arg\min_{k=1,..,J} W_n(\alpha_k)$. In this paper we assumed that, for a given quantile τ, there is a searching grid with 200 values of α_k.

2. define $\tilde{x}_{ij} = (x_{0,ij} \cdot \hat{\alpha}, x_{1,ij})$ and replace it into the equation (11):

$$L_n(\beta) = -\sum_{i,j=1}^{n} w_{ij}\rho_\tau[\ln(z_{ij}) - \max(0, \tilde{x}_{ij}\beta(\tau))], \tag{12}$$

Finally, the vector of parameters from (12) is estimated by using the MCMC algorithm developed by Chernozukov and Hong (2003).

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Table 1 – List of Countries

Afghanistan	Dominican Republic	Lebanon	St. Lucia
Albania	Ecuador	Lesotho	St. Martin
Algeria	Egypt	Liberia	St. Pierre
Andorra	El Salvador	Libya	St. Vincent
Angola	Equatorial Guinea	Lithuania	Samoa
Antigua and Barbuda	Eritrea	Luxembourg	S. Tome Principe
Argentina	Estonia	Macedonia	Saudi Arabia
Armenia	Ethiopia	Madagascar	Senegal
Australia	Fiji	Malawi	Serbia
Austria	Finland	Malaysia	Seychelles
Azerbaijan	France	Maldives	Sierra Leone
Bahamas, The	French Guiana	Mali	Singapore
Bahrain	Gabon	Malta	Slovakia
Bangladesh	Gambia, The	Mauritania	Slovenia
Barbados	Georgia	Mauritius	Solomon Islands
Belarus	Germany	Mayotte	Somalia
Belgium	Ghana	Mexico	South Africa
Belize	Gibraltar	Micronesia	South Sudan
Benin	Greece	Moldova	Spain
Bermuda	Greenland	Mongolia	Sri Lanka
Bhutan	Grenada	Montenegro	Sudan
Bolivia	Guadeloupe	Montserrat	Suriname
Bosnia/Herzegovina	Guatemala	Morocco	Swaziland
Botswana	Guinea	Mozambique	Sweden
Brazil	Guinea-Bissau	Myanmar	Switzerland
Brunei	Guyana	Namibia	Syrian
Bulgaria	Haiti	Nauru	Taiwan
Burkina Faso	Honduras	Netherlands	Tajikistan
Burundi	Hong Kong	New Zealand	Tanzania
Cambodia	Hungary	Nicaragua	Thailand
Cameroon	Iceland	Niger	Timor-Leste
Canada	India	Nigeria	Togo
Cape Verde	Indonesia	Norway	Tonga
Cayman Islands	Iran	Oman	Trinidad and Tobago
Central African Rep.	Iraq	Pakistan	Tunisia
Chad	Ireland	Panama	Turkey
Chile	Israel	Papua N. Guinea	Turkmenistan
China	Italy	Paraguay	Uganda
Colombia	Jamaica	Peru	Ukraine
Comoros	Japan	Philippines	Unt. Arab Emirates
Congo, Rep.	Jordan	Poland	United Kingdom
Congo, Dem. Rep.	Kazakhstan	Portugal	United States
Costa Rica	Kenya	Puerto Rico	Uruguay
Cote d'Ivoire	Kiribati	Qatar	Uzbekistan
Croatia	Korea, Dem. Rep.	Reunion	Vanuatu
Cyprus	Korea, Rep.	Romania	Venezuela
Czech Republic	Kuwait	Russian Federation	Viet Nam
Denmark	Kyrgyzstan	Rwanda	Yemen
Djibouti	Lao PDR	St. Helena	Zambia
Dominica	Latvia	St. Kitts	Zimbabwe

Table 2 – Baseline Estimations

	CQ Model* (1)	PPML (2)	OLS (3)	CQ Model* (4)	PPML (5)	OLS (6)	CQ Model* (7)	CQ Model* (8)
Agreement dummie								
RTA	0.209[a]	0.031	-0.024	0.314[a]	0.104	0.001	0.391[a]	0.333[a]
Visa/Asylum	–	–	–	–	–	–	0.060	0.107[b]
Cost variables								
ln(GDPpcDest/GDPpcOrig)	0.201[a]	0.079	0.135[a]	0.572[a]	0.491[a]	-1.564	0.625[a]	0.086[a]
ln Pop. origin	1.512[a]	1.299[a]	0.618[a]	–	–	–	–	–
ln Pop. destination	-1.056[a]	-1.031[a]	-0.820[a]	–	–	–	–	–
WTO	0.376[a]	0.746[a]	0.257[a]	0.465[a]	0.462[b]	0.405[a]	0.362[a]	0.598[a]
ln Distance	-1.593[a]	-1.226[a]	-1.542[a]	-1.538[a]	-1.342[a]	-1.564[a]	-1.478[a]	-1.462[a]
ln Stock 1960	0.267[a]	0.399[a]	-0.070[b]	0.068[a]	0.396[a]	-0.052	0.079[a]	0.077[a]
Border	2.304[a]	0.908[a]	1.810[a]	2.314[a]	0.861[a]	1.702[a]	1.159[a]	1.132[a]
Colony	1.781[a]	1.433[a]	1.607[a]	1.845[a]	1.375[a]	1.584[a]	1.519[a]	1.640[a]
Common Language	0.986[a]	0.410[a]	0.967[a]	1.105[a]	0.419[a]	0.965[a]	0.925[a]	0.867[a]
EU	–	–	–	–	–	–	–	0.197[a]
Country Effects	yes	yes	yes	no	no	no	no	no
Year Effects	yes	yes	yes	no	no	no	no	no
Country-by-Year Effects	no	no	no	yes	yes	yes	yes	yes
Observations	62,851	62,851	39,263	62,851	62,851	39,263	4,887	4,887

Notes: (*) model with $\tau = 0.50$. ([a]), ([b]) and ([c]) denote statistical significance at 1%, 5% and 10%, respectively.

Table 3 – Descriptive Statistics: Observed and Fitted Models*

	Observed	CQ Model**	PPML	OLS
Min	0	0.0845	0.0067	0.0109
Max	1.16e+07	2.12e+07	6,290,064	7,999,113
Mean	6844.93	7334.00	6803.99	3592.15
Total	4.44e+08	4.85e+08	4.44e+08	2.25e+08
Centile 0.10	0	0.6008	9.6706	1.976
Centile 0.50	5	7.63	246.76	30.19
Centile 0.90	2,571	1,707.59	7,138.44	1,198.76
Recovered Values***	0	517,745.3	3.60e+07	0

Notes: (*) models from Table 1: columns (3)-(5). (**) model with $\tau = 0.50$. (***) Estimates for the migration stock when observed values are equal to zero.

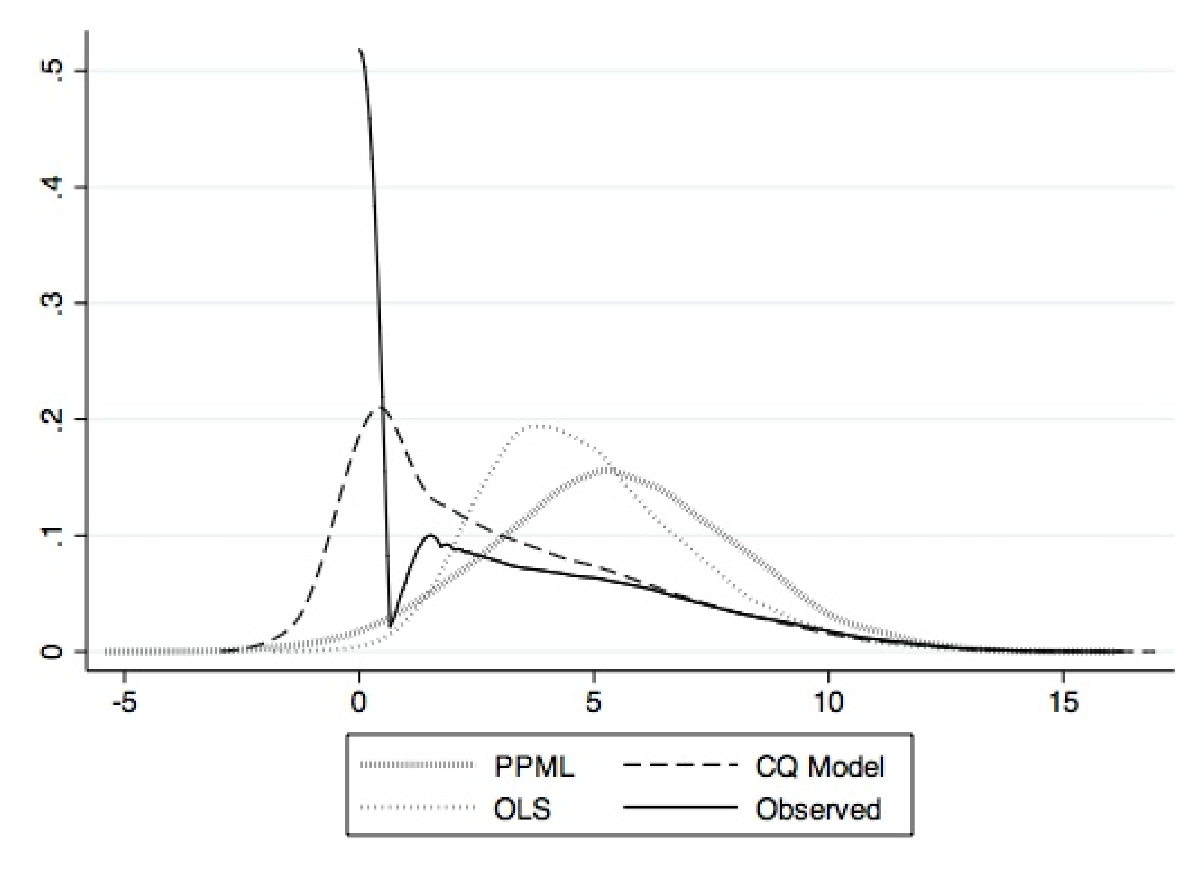

Figure 1 – Observed and Fitted Models

Table 4 – Gravity Models with Instrumental Variables

	Censored Quantile Models		
Specifications	$\tau = 0.25$ (1)	$\tau = 0.50$ (2)	$\tau = 0.75$ (3)
Agreement dummies			
RTA*	0.197[a]	0.361[a]	0.403[a]
Cost variables			
WTO	0.443[a]	0.456[a]	0.449[a]
ln(GDPpcDest/GDPpcOrig)	0.193[a]	0.151[a]	0.153[a]
ln Distance	-1.439[a]	-1.440[a]	-1.132[a]
ln Stock 1960	0.098[a]	0.208[a]	0.235[a]
Border	2.010[a]	2.269[a]	2.320[a]
Colony	1.018[a]	1.817[a]	1.770[a]
Common Language	1.018[a]	1.045[a]	1.098[a]
Country-by-Year Effects	yes	yes	yes
Observations	62,851	62,851	62,851

Notes: (*) Controlled for endogeneity. ([a]), ([b]) and ([c]) denote statistical significance at 1%, 5% and 10%, respectively.

Table 5 – Gravity Models with Instrumental Variables: Visa and Asylum

	Censored Quantile Models	
Specifications*	(1)	(2)
Agreement dummie		
RTA**	0.368[a]	0.342[a]
Visa/Asylum**	0.121[b]	0.142[a]
EU	–	0.174[a]
Cost variables		
WTO	0.305[a]	0.353[a]
ln(GDPpcDest-GDPpcOrig)	0.241[a]	0.232[a]
ln Distance	-1.678[a]	-1.454[a]
ln Stock 1960	0.084[a]	0.069[a]
Border	1.544[a]	1.623[a]
Colony	1.791[a]	1.701[a]
Common Language	1.142[a]	1.145[a]
Country-by-Year Effects	yes	yes
Observations	4,887	4,887

Notes: (*) Those specifications consider $\tau = 0.50$. (**) Controled by endogeneity. ([a]), ([b]) and ([c]) denote statistical significance at 1%, 5% and 10%, respectively.

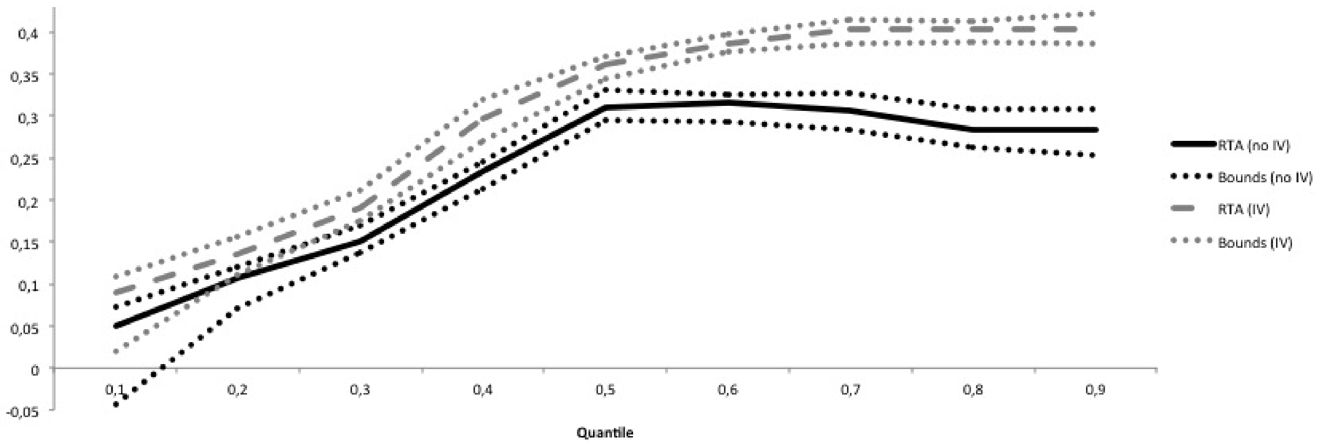

Figure 2 – RTA Effects Across Quantiles – Country-by-Year fixed effects estimation.