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The impact of exchange rate uncertainty on exports: a panel VAR analysis*

El impacto de la incertidumbre del tipo de cambio sobre las exportaciones: un análisis de panel VAR

Leonel Muinelo-Gallo**
Ronald Miranda Lescano***
Gabriela Mordecki****

ABSTRACT

In this paper we analyze the impact of exchange rate uncertainty on export flows among a panel of 27 countries throughout the 1994/01-2014/12 period. In order to do this, we apply a panel vector autoregressive model approach. By dividing the panel into two subgroups that involve manufacturing-exporting and commodity-exporting economies, we observe a different effect of exchange rate uncertainty on exports. This has a negative impact in manufacturing-exporting countries, but does not affect commodity-exporting countries. This result appears to be explained by countries' economics characteristics, involving the flexibility or rigidities of the export adjustment arising exchange rate uncertainty.

Key words: *Exchange rate uncertainty, exports, panel vector autoregressive, manufacture-exporting economies, commodity-exporting economies.*

JEL Classification: *C33; F31; F41.*

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Resumen

En este artículo analizamos el impacto de la incertidumbre del tipo de cambio sobre los flujos de exportaciones en un panel de 27 países durante el período 1994/01-2014/12. Para ello, aplicamos un enfoque de vectores autoregresivos con datos panel. Dividiendo el panel en dos subgrupos que incluyen economías exportadoras de manufactureras y economías exportadoras de productos básicos, observamos un efecto diferencial de la incertidumbre del tipo de cambio sobre las exportaciones. Esta incertidumbre tiene un impacto negativo en los países exportadores de manufacturas, pero no afecta de forma significativa a los países exportadores de productos básicos. Este resultado parece explicarse por las características económicas de los países, las cuales involucran la flexibilidad o rigidez del ajuste de las exportaciones a la incertidumbre del tipo de cambio.

Palabras clave: *Incertidumbre del tipo de cambio, exportaciones, panel de vectores autorregresivos, economías exportadoras de manufacturas, economías exportadoras de productos básicos.*

Clasificación JEL: *C33; F31; F41.*

1. Introduction

The collapse of the exchange rate system adopted at the Bretton Woods Conference had as one of its consequences the free floating of the majority of currencies in the world, generating concern about the effects that exchange rate uncertainty could have on international trade flows. From that moment forward, an extensive literature has analyzed the economic effects of this exchange rate uncertainty (Arize *et al.*, 2008; Kandilov, 2008).

Since the concept of uncertainty is difficult to quantify precisely, exchange rate volatility (*i.e.* variability) has commonly been used by the literature on trade as a proxy of it and refers to the risk associated with unexpected exchange rate movements (McKenzie, 1999). However, there is no consensus in the empirical literature about which statistical measure to use to measure exchange rate volatility (Arize, 1997; Hall *et al.*, 2010) and also about the significance and sign of the impact of exchange rate volatility on exports (Bahmani-Oskooee and Hegerty, 2007; Bouoiyour and Selmi, 2016; Bayar, 2018).

The main aim of this paper is to analyze how exchange rate uncertainty impacts exports among a novel panel of European, South American and Oceanian countries throughout the 1994/01-2014/12 period. While prior empirical studies (Sauer and Bohara, 2001; Grier and Smallwood, 2007; Hall *et al.*, 2010) generally consider a group of developed countries and a group of developing countries, in this paper we focus on a novel classification of countries in a panel of manufactures-exporting economies (mainly European countries) and a panel of commodity-exporting economies (South American and Oceanian countries).

This novel categorization allows us to more deeply analyze the effects of exchange rate uncertainty on exports in countries with different production structures.

The empirical methodology applied is a Panel Vector Autoregressive model (P-VAR) developed by Abrigo and Love (2016). Specifically, the macroeconomic analysis consists of studying the dynamics of short- and medium-term relationships between exports and real effective exchange rate volatility, as well as the dynamics of a set of macroeconomic variables. In addition, impulse-response functions (IRF's) and Granger causality are examined.

The contributions of our study are fourfold. First, we provide novel empirical evidence about the effects of real effective exchange rate volatility on exports for a set of countries with different production patterns. Second, given that there is no consensus about exchange rate uncertainty specification, this paper aims to contribute to the debate by using and arguing in favour of specific measures of real effective exchange rate volatility. Third, the dynamics between exports and real effective exchange rate volatility are studied over a long period of time (1994/01 to 2014/12) with monthly data frequency, covering several macroeconomic shock events, which enables us to discuss episodes such as the Great Recession (the international financial crisis of 2008/2009). Finally, we applied the P-VAR methodology which is not yet explored in the international trade literature and which allows us to analyze the dynamics of short- and medium-term relationships between exports and real effective exchange rate volatility.

We observe a group-specific impact of real effective exchange rate volatility on exports. Manufactures-exporting economies show a negative effect of real effective exchange rate volatility on exports, while this effect is not significant for the commodity-exporting economies. There is also evidence that the Great Recession of 2008/2009 negatively impacts exports flows among manufactures-exporting economies but does not significantly affect commodity-exporting countries.

The paper is organized as follows. Section 2 reviews the related literature. Data and variables are presented in Section 3. The measures of volatility of real exchange rates are discussed in Section 4. The methodology is presented in Section 5. Section 6 presents the main findings. Finally, Section 7 presents conclusions and policy implications.

2. Exchange rate volatilities and its impact on international trade: background

There is an extensive and inconclusive literature about the impact of exchange rate volatility on international trade.[1] Theoretical literature has not been able to consistently support a strong relationship between these variables and

[1] Literature reviews of such studies are provided by Ozturk (2006), Bahmani-Oskooee and Hegerty (2007), Coric and Pugh (2010), Bouoiyour and Selmi (2016) and Bayar (2018).

this continues to be a controversial issue.[2] The most common hypothesis is a negative effect of real exchange rate (RER) volatility on exports. In this sense, some scholars have argued that RER volatility affects the behavior of traders in response to the risk of their international trade activities through the uncertainty of benefits and costs denominated in foreign currency (Ethier, 1973; Clark, 1973; Gagnon, 1993; Aftab *et al.*, 2012; Nazlioglu, 2013). More specifically, risk averse traders respond negatively to unanticipated exchange rate fluctuations and move to less risky activities (e.g. agents choose internal trade instead of foreign trade), leading to change in the size/contribution of economic activities to relevant macroeconomic variables such as the trade balance or the balance of payments (Were, 2015). However, other researchers point out that there can be positive effects on international trade, because some agents see exchange rate variability as an opportunity to increase benefits from international trade. Specifically, De Grauwe (1988) argues that the increase in foreign exchange risk can be decomposed into a substitution and an income effect. Due to an increase in risk, the substitution effect operates by reducing export activities in favor of less risky local activities. However, the income effect operates in the opposite direction: if producers are sufficiently risk-averse, an increase in exchange rate risk raises the expected marginal utility of exports revenue and therefore induces them to increase their export activity. Consequently, if the income effect is high and dominates, an increase in foreign exchange risk has a positive effect on export trade flows. Similarly, Broll and Eckwert (1999) point out that the effect will depend on the firm's behavior vis-à-vis the risk, which is why they conclude that volatility may increase exports since an increment of the exchange risk can enhance the potential gains of trade. Sercu (1992) shows that exchange rate volatility can in some cases increase the volume of trade rather than penalize it. If, on average, high volatility increases the probability that the price received by exporters exceeds the costs of tariffs or transportation in trade, trade is likely to be stimulated. Dellas and Zilberfarb (1993), using a theoretical asset market approach, explain a positive effect of exchange rate volatility on exports based on the risk aversion parameter of the traders. Finally, Serenis and Tsounis (2013) point out the existence of studies that suggest that the effect might be expected to be insignificant due to use of futures markets instruments to hedge the uncertainty associated with exchange rate movements (Willett, 1986; Nazlioglu, 2013).

The empirical literature has also undergone significant evolution. Earlier studies used simple regression methods to assess the effects of exchange rate volatility on exports and employed standard measures to model exchange rate volatility.[3] For example, Hooper and Kohlhagen (1978) find no evidence that exchange rate volatility, measured as the standard deviation of the nominal

2 See, for comprehensive surveys, McKenzie (1999) and Bahmani-Oskooee and Hegerty (2007).

3 See Bayar (2018) and Bahmani-Oskooee and Hegerty (2007) for a comprehensive surveys of these empirical studies.

exchange rate, affects bilateral and multilateral exports in developed countries between the mid-1960s and mid-1970s. However, Cushman (1983) follows the work of Hooper and Kohlhagen (1978) to analyze the impact of exchange rate variability, in this case measured as the standard deviation of the RER, on US bilateral trade with five other industrialized countries (Canada, France, Germany, Japan and the United Kingdom) over 1965-1977. For this group of countries, unexpected movements in the RER generally have a significant and negative impact on international trade. Akhtar and Hilton (1984) find a negative relationship when analyzing the impact of exchange rate volatility, measured as the standard deviation of the nominal effective exchange rate, on bilateral trade between the United States and Germany over 1974-1981.

Both the techniques for measuring volatility and also the available sources of information have evolved significantly in recent decades, enabling a significant evolution in the empirical trade literature. For example, Kroner and Lastrapes (1993) estimate exchange rate volatility using a Generalized Autoregressive Conditional Heteroskedasticity (GARCH) multivariate process, while Chowdhury (1993) estimates the volatility using a moving standard deviation of the RER. Arize (1997 and 2008) examines the volatility of the real effective exchange rate (REER) using an ARCH (Autoregressive Conditional Heteroskedasticity) process. Most of these studies find a negative effect of exchange rate volatility on exports flows. More recently, some studies tackle this issue using panel data analysis. For instance, Sauer and Bohara (2001) empirically analyze the effect of real effective exchange rate volatility (REERV) on exports for a panel of 91 developed and developing countries during the 1966-1993 period. They estimate exchange rate volatility using an ARCH process and two variants to the moving standard deviation of the REER. They find a negative effect of the REERV on exports. When the sample is divided into developed and developing countries, the impact for developing countries (Latin American and African countries) is negative.[4] However, they find no effect in advanced economies. Situ (2015) considers the bilateral trade of the United States with two groups of countries with different characteristics, developed and least-developed export-oriented countries, for two periods: 1994-2007 and 2008-2014. Using panel data techniques and modeling the volatility of the RER through a GARCH process, he finds a negative impact of RER volatility on exports (except for the first period for least developed countries), with a larger result for the developed countries, mainly in the 2008-2014 period. This is explained by the fact that firms in advanced countries have a greater capacity relative to export-oriented developing economies to adjust exports when facing variability in the RER. Furthermore, Vilela and MacDonald (2016) analyze the effect of REERV, estimated as the moving standard deviation and also using a GARCH process, on exports for a panel of 106 countries over 2000-2011. They find a negative impact for the sample as a

[4] See also a study for African countries by Bahmani-Oskooee and Gelan (2018).

whole and also for the developing and emerging economy sub-samples, which is attributed to the oil-exporting economies.

Other closely related literature provides insight into the relationship between exchange rate movements and trade balance, based on the so-called J- and S-curve concepts (Bahmani-Oskooee and Ratha, 2004; Bahmani-Oskooee and Hegerty, 2010). While the J-curve depicts the potential time path of a country's trade balance after a change in the exchange rate, the S-curve reflects what happens before and after a change in the exchange rate. A depreciation or devaluation should make imports more expensive in the short run and increase a country's exports in the long run (due to some delays in adjusting consumption and producers' contracts). That worsens the trade balance first and improvement comes afterward; this is a J-curve pattern. As the trade balance improves, the initial depreciation is reversed (a negative correlation), and it might always lead to a second period of depreciation (a positive correlation), *i.e.*, an S-curve pattern. However, the empirical evidence about the effect of exchange rate movements on a trade balance is still an unanswered question (Arize *et al.*, 2017; Yazgan and Ozturk, 2019).

As a general observation, we can state that although most of the empirical studies reviewed show that negative effects of REERV on exports prevail, they are difficult to compare and generalize since they differ in terms of sample periods, the variables used, the countries considered, the volatility specifications, the type of exports (aggregated, bilateral or sector-specific), the exchange rate (nominal, real or effective), and methodologies and estimation methods. In addition, the previous empirical evidence describes the importance of economies' characteristics; however, this issue has not been sufficiently examined. In this context, this paper pursues analysis on this important issue.

3. Data and variables

The sample used in this paper consists of a monthly frequency panel dataset of 27 countries, including 15 European (E-15), 10 South American and two Oceanian, over 1994-2014.[5] The panel selection criteria pertained to the export-related macroeconomic characteristics of the economies, in order to analyze different effects of REERV on exports for a sample of countries (Figure 1, panel a). In so doing, we distinguish between manufactures-exporting (MXE) and commodity-exporting (CXE) countries following the criteria established in the World Economic Outlook's Statistical (IMF, 2015). Therefore, we categorize a country as a commodity exporter if it satisfies two conditions: 1) at least 35% of total goods exports are classified as commodities; 2) net commodities exports represent at least 5% of total trade in goods, on average, during the 1994-2014 period. By using 1994-2014 averages and according

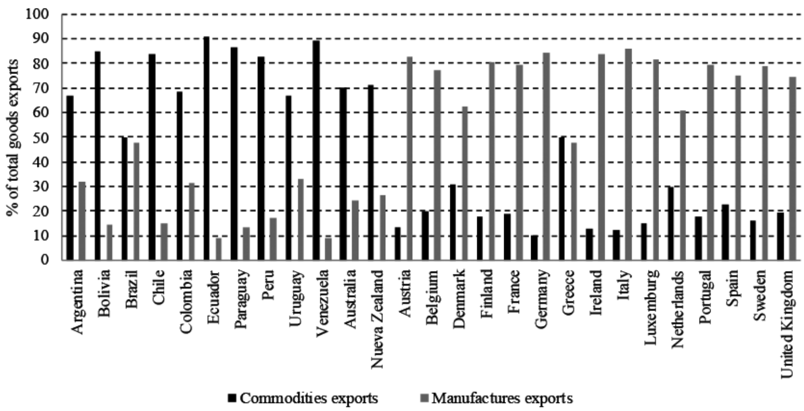
[5] See Table A.1 in the appendix for the list of countries considered.

to IMF data, the E-15 countries are classified into the MXE group and the South American and the two Oceanian countries are classified into the CXE group (Figure 1, panel b).

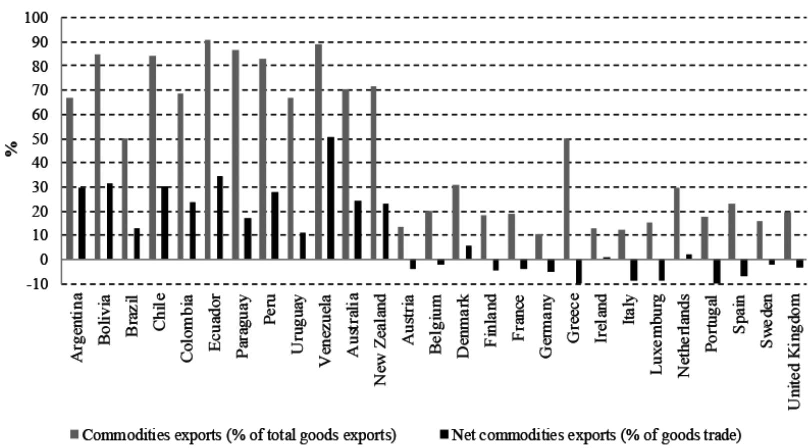
Following Miranda and Mordecki (2019), the main series used in this study correspond to: total goods exports (*X*), world goods imports (M^*), international commodity prices indices disaggregated into non-fuel prices (*P*) and fuel prices

FIGURE 1
CHARACTERISTICS OF COUNTRIES' EXPORTS OVER 1994-2014

(a) Manufactures and commodities exports countries



(b) Commodities and net commodities exports countries.



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(P^*), and REER is used to calculate the different measures of REERV.[6] The international trade literature typically uses GDP as a proxy of economies' demand at the country level; however, as monthly world GDP is not available to approximate world demand, in this paper we use world goods imports. The other series considered are the commodity prices indices, disaggregated into non-fuel and fuel commodities price indices. Both indices are relevant to explain export earnings in South American and Oceanian countries, while the fuel commodities price index is relevant in explaining E-15 export costs.

4. Measures of real exchange rate volatility

We considered two groups of univariate measures to quantify the REERV. First, a measure of historical volatility, quantified as the sample moving standard deviation of the growth rate of real effective exchange rate (REER). Second, a measure of conditional variance, specified as the squared residuals of the ARIMA model.

4.1. Historical volatility

As a measure of historical volatility, we consider the moving standard deviation:

$$Vm_t = \sqrt{\frac{1}{m}.\sum_{i=1}^{m}[\ln(REER_{t+i-1}) - \ln(REER_{t+i-2})]^2} \tag{1}$$

where V_{mt} is the moving sample standard deviation of the growth rate of REER, m refers to the order of the moving averages at m = 4, 8, 12 and 24 months and t represents time.[7] This type of measure allows the average of the series to vary, and will indicate different sensitivity of exports to exchange rate volatility depending on which moving average is used. In this sense, the longer the time used for the moving average of the standard deviation, the more difficult to capture variability, and vice-versa. Given the impact of exchange rate volatility on a macroeconomic variable such as exports, a relatively short time period for the moving average *a priori* would be meaningless in the export decision, since it is difficult to respond to a phenomenon of very short-term volatility. Analogously, a longer period for the moving average may not reflect such variability. For these reasons, in order to eliminate arbitrary selection of m, in this study we evaluate: m = 4, 8, 12 and 24 periods.

6 Table A.2 in appendix presents the definitions and sources of all variables; while and Table A.3 and Table A.4 provide the main summary statistics.

7 Similar procedures for obtaining a measure of exchange rate volatility are presented in Koray and Lastrapes (1989) and Arize (1997).

4.2. Conditional variance

In traditional time series models,[8] it is common to assume that the distributions of the conditional and unconditional variance are heteroscedastic. For this reason, and based on a linear function of the expected square of the lagged value of the error term from an ARIMA regression of the REER (Engle, 1982), we introduce a GARCH process in order to estimate the REERV:

$$\varepsilon_t/\psi_{t-1} \sim N(0, V_t), \tag{2}$$

$$V_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i \varepsilon_{t-i}^2 + \sum_{i=1}^{p} \beta_i V_{t-i} = \alpha_0 + A(L)\varepsilon_{t-i}^2 + B(L)V_{t-i} \tag{3}$$

Equation 2 denotes the distribution of the error term, ε_t, with a mean of zero and conditional variance V_t. Equation 3 specifies the conditional variance of a GARCH process (p, q), where $q > 0$ is the number of ARCH terms and $p > 0$ is the number of GARCH terms. In this sense, the conditional variance is represented by three terms: a) the mean of the conditional variance, α_0; b) the ARCH term, which measures the volatility of the previous time period as the squared residuals of an autoregressive process (ε_{t-1}^2); c) the GARCH term, which captures the prediction error of the variance of the previous period (V_{t-1}). Thus, the GARCH process (p, q) expressed in Equation 3 will be stationary in the broader sense if and only if $A(L) + B(L) < 1$.

A substantial number of works have also used this type of measures. In this sense, Bollerslev *et al.* (1992) argue that it is common to find, in the empirical evidence, a certain persistence of the variance over time in GARCH processes estimations. That is, the autoregressive polynomial has a unit root, which means that the GARCH process is integrated and not stationary, I (1), in which case it is called an Integrated GARCH (IGARCH). In consideration of this context, we also use this IGARCH process to specify the REERV for all countries in the sample.

Nelson (1991) introduced a nonlinear process called an Exponential GARCH model (EGARCH). In contrast with a GARCH model, that ensures positive conditional variance by employing a linear combination of positive random variables, it adopts an alternative specification, which does not restrict the α and β parameters to be non-negative, but ensures that the conditional variance is non-negative. This procedure gives us an alternative measure to estimate REERV when GARCH or IGARCH models do not satisfy the conditions described.

Considering the above, we then estimated the conditional volatility of the REER by country over 1994/01-2014/12 and plot them in Figures 2 and 3.[9]

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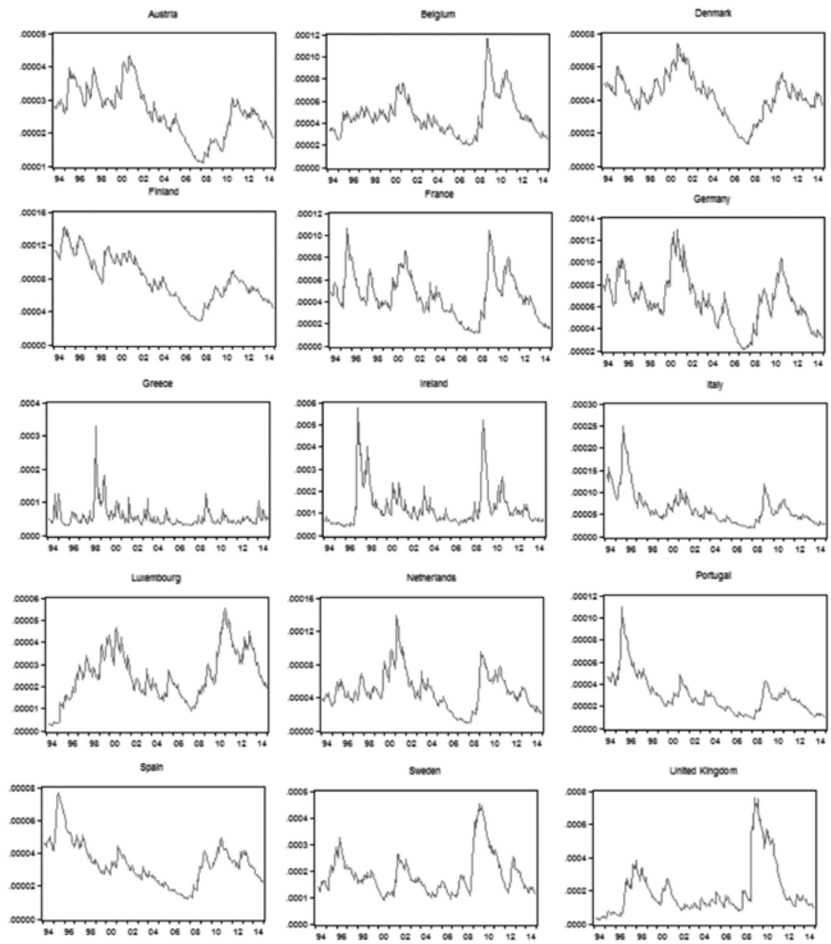
[8] See, for example, Bollerslev (1986).

[9] Tables A.5 and A.6 in appendix show the selected specification of estimate the conditional volatility of the REER by country.

FIGURE 2
CONDITIONAL VARIANCE OF THE REER OF THE MXE COUNTRIES OVER 1994-2014
(monthly data)

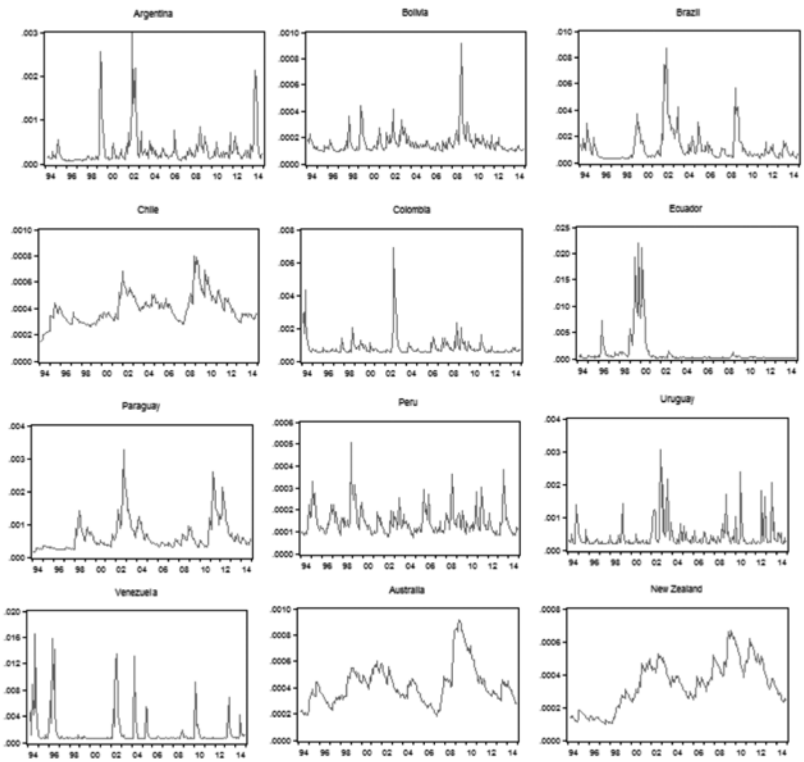


Source: Developed by authors based on IMF data.

The peaks and troughs that occur in the progression of the series represent the episodes of high or low volatility in the sample period.

From a visual inspection of both figures, we identified the main international crisis episodes that occurred in the sample period: the 1994/1995 Mexican crisis, the 1997/1998 Asian crisis, the 1999 Brazilian crisis, the 2001/2002 Argentinean crisis, the Great Recession (or 2008/2009 international financial crisis) and contagion effects. In addition, we also see the effects of the incorporation of Austria, Finland and Sweden into the European Union in 1995 and circulation of

FIGURE 3
CONDITIONAL VARIANCE OF THE REER OF THE CXE COUNTRIES OVER 1994-2014
(monthly data)



Source: Developed by authors based on IMF data.

the Euro currency. Finally, note that the conditional variance of REER is much higher for South American countries than European and Oceanian countries.

5. Empirical strategy

In the macroeconomic literature, there are basically two ways of considering the interdependence of relationships between variables. One option is to build a general equilibrium model, where there are specified optimizer agents, preferences, technologies and constraints. These models are extremely useful because they provide answers to economic policy issues and allow a clear understanding of welfare issues. However, by construction, these models impose certain constraints that are not always compatible with the statistical properties of the data. In this context, the policy prescriptions that can be derived are strongly related to the related assumptions (Canova and Ciccarelli, 2013). An alternative approach is to construct vector autoregressive models (VAR). All variables in a VAR system are typically treated as endogenous, although identification restrictions based on theoretical models or on statistical procedures may be imposed to disentangle the impact of exogenous shocks to the system (Sims, 1980).

In this paper we additionally develop the method, by performing a dynamic empirical analysis of simultaneous equations using the Panel-VAR (P-VAR) approach (as done by Love and Zicchino, 2006).[10] P-VAR analysis combines traditional VAR methodology, considering the whole set of system variables as endogenous and interdependent, with a panel data technique, which allows to control for individual and temporal heterogeneity and to estimate causality of relationships between endogenous variables (Canova and Ciccarelli, 2013). P-VAR methodology, first, allows us to specify the model with little theoretical information about the relationships among the variables. Second, it is also useful to deal with the endogeneity problem, given that all variables are potentially endogenous. Finally, the P-VAR model allows us to make more complete use of the information available in the data since it exploits the time-series and cross-sectional dimensions of our database (Grossmann *et al.*, 2014).

The original P-VAR model can be specified as a model of k endogenous variables with an order of lags p, as follows:

$$Y_{it} = Y_{it-1}A_1 + Y_{it-2}A_2 + \ldots + Y_{it-p}A_p + X_{it}B + u_i + d_t + e_{it} \tag{4}$$

where $i = 1, \ldots, N$ represents the country and t is the time over 1994/01-2014/12. Y_{it} is the 1 x k vector of endogenous variables, X_{it} is the 1 x m vector of exogenous variables, d_t is a 1 x N temporal dummy that captures the specific shocks that affect all countries in period t, while u_i represents the country-effects variable that captures unobservable individual heterogeneity, and e_{it} are idiosyncratic errors, both of dimensions 1 x k. The k x k matrices $A_1, A_2, \ldots A_p$ m x k matrix B are the parameters to be estimated. Finally, it is assumed that $E(e_{it}) = 0$, $E(e_{it}, e_{it}) = \Sigma$ and $E(e_{it}, e_{iS}) = 0 \; \forall \; t > s$.

The Y_{it} vector of endogenous variables is comprised of: total goods exports, REER volatility and commodity non-fuel price index. The exogenous variables are global demand for goods and the commodity fuel price index. Finally, d_t is a temporal exogenous shock that reflects the impact of the international financial crisis of 2008/2009 that takes the value of 1 from August 2008 to December of 2014, and 0 otherwise.

Following prior trade literature, we specify the total goods export equation as:[11]

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10 Love and Zicchino (2006) make the STATA pvar code available for the use of researchers; the most recent version of this pvar code is in Abrigo and Love (2016).

11 See also Chowdhury (1993), Arize (1997), Arize and Malindretos (1998), Arize *et al.* (2008), as well as Bayar (2018) for an excellent survey.

$$X_{it} = \alpha_1 X_{it-p} + \alpha_2 P_{it-p} + \alpha_3 Vol_{it-p} + \beta_1 M^*_{it} + \beta_2 P^*_{it} + u_i + d_t + e_{it} \quad (5)$$

with p lags, where $i = 1, \ldots, 27$ represents the country and t is the time between 1994/01 and 2014/12. The endogenous variables are total goods exports (X), the non-fuel commodity price index (P) and the different measures of REERV (Vol). The exogenous variables of the model are world goods imports (M^*) and the fuel commodity price index (P^*). In this case, u_i represents the country effects that capture unobservable individual heterogeneity, the dummy variable d_t captures the international financial crisis of 2008/2009,[12] and e_{it} contains the idiosyncratic errors. Finally, the coefficients α_1, α_2, α_3, β_1 and β_2 are the parameters to be estimated.

Specifically, we estimate a dynamic P-VAR model with country effects to preserve the orthogonality between the regressors (lags of the dependent variables). Also, and following Love and Zicchino (2006), Love and Turk (2014) and Grossmann *et al.* (2014), in order avoid biases in coefficients, we use the Helmer transformation to remove the forward mean, *i.e.*, the mean of all the future observations available for each country-year. This transformation preserves the orthogonality between transformed variables and lagged regressors, making it possible to use lagged regressors as instruments and estimate coefficients by Generalized Method of Moments (GMM) (Arellano and Bover, 1995).[13] Additionally, once P-VAR models have been estimated, we perform simulation exercises using impulse response functions (IRF's). Finally, it is important to point out that the P-VAR methodology also allows us to include supposedly exogenous variables in our model.

6. Empirical results

In this section, first, we present the results of the unit root test. Second, we report the estimations for the panels of commodity-exporting countries

[12] Situ (2015) and Vilela and MacDonald (2016) take into account the effects of the 2008/2009 international financial crisis on exports. To capture this effect, the first article subdivides the analysis period and the second article introduces an intervention to the model. Here, we follow the second one by introducing a dummy variable that take the value 1 from August 2008 to December of 2014 and 0 otherwise (in a model specified in levels). However, when we estimate our empirical model following the methodology of Situ (2015), the estimations results doesn't change. These last results are not reported due to space problems, but are available upon request.

[13] The GMM estimation deals with potential endogeneity issues (Arellano and Bond, 1991; Arellano and Bover, 1995; Blundell and Bond, 1998). More specifically, both are general estimators designed for situations with: 1) a linear functional relationship; 2) one left-hand-side variable that is dynamic, depending on its own past realizations; 3) independent variables that are not strictly exogenous, meaning they are correlated with past and possibly current realizations of the error; 4) fixed individual effects; and 5) heteroskedasticity and autocorrelation within individuals but not across them.

(CXE) and manufactures-exporting countries (MXE). Third, we show the post-estimation outcomes. Finally, an extension of the empirical analysis is conducted; specifically, we report the estimates splitting the panel of countries by development level following the classification of the International Monetary Fund (see Nielsen, 2011).

6.1 Unit root test

In order to estimate the P-VAR model, the integration order of the series (stationarity) was analyzed. Following Grossmann *et al.* (2014), a first-generation panel unit root tests is used. We have information from a strongly balanced macro-panel for MXE ($n1$ = 15) and for CXE ($n2$ = 12) over the 1994/01-2014/12 period (t = 252). The t dimension is sufficiently large, and larger than both the $n1$ and $n2$ dimensions. Therefore, the Levin, Lin and Chu (2002) (LLC) unit root test is used (Levin *et al.*, 2002).[14]

Table 1 presents the main results of the unit root test for the MXE and CXE panels of countries for the entire 1994/01-2014/12 period. It is found that the export series is integrated of first order, I (1). The REER volatility series estimated as the standard deviation moving averages with 4, 8, 12 and 24 periods (*V4*, *V8*, *V12* and *V24*, respectively) are stationary, *i.e.* I (0).

Table 2 presents the results of the unit root test for those series which are common to all countries in the panel $\{M^*, P, P^*\}$. The results of the Augmented Dickey-Fuller test show that this set of level variables have units roots and are stationary in the first difference, *i.e.* I(1).

In the case of the REER conditional volatility (V), the MXE and CXE panels are no longer balanced, since the data does not contain the basis of observation for all 27 countries and all months throughout the 1994/01-2014/12 period. The t dimension is sufficiently large and greater than the $n1$ and $n2$ dimensions, so the Fisher-type unit root test is used (see Choi, 2001).[15] Table 3 presents the results of the unit root test for REER conditional volatility (V), and we reject the null hypothesis that all panel series contain unit roots for the MXE and CXE panels.

It should be noted that by construction the GARCH processes are stationary, I (0), and therefore at least one panel series is stationary; also, the IGARCH processes are first order integrated, I(1). Thus, we consider V to be a first order integrated process for both panels.

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[14] There is a wide variety of unit root tests for panel data. The tests present different assumptions for implementation (whether the panel is balanced or not; whether the panel number ratio, n, divided by the size of the temporal dimension, t, tends to infinity; whether n or t is fixed) (Maddala and Wu, 1999). Moreover, see Hurlin (2010) for a discussion about use of first- and second-generation panel unit root tests.

[15] Choi (2001) describes four ways of combining the p-value: when n is finite the inverted chi-squared test, the inverted normal test and the inverted logit test, and when n tends to infinity, suggests to use a modified inverted chi-squared test.

TABLE 1
LLC UNIT ROOT TEST RESULTS: CXE AND MXE PANELS

Variable	Level		First Difference	
	Adjusted statistic t*	Integration order	Adjusted statistic t*	Integration order
Panel: MXE				
X	2.544 [0.995]	I(1)	–31.737 [0.000]	I(0)
V4	–8.347 [0.000]	I(0)		
V8	–8.146 [0.000]	I(0)		
V12	–9.208 [0.000]	I(0)		
V24	–5.499 [0.000]	I(0)		
Panel: CXE				
X	1.260 [0.896]	I(1)	–33.546 [0.000]	I(0)
V4	–10.171 [0.000]	I(0)		
V8	–8.069 [0.000]	I(0)		
V12	–10.264 [0.000]	I(0)		
V24	–6.748 [0.000]	I(0)		

Note: LLC refers to Levin-Lin-Chu unit root test. Null hypothesis: panels contain the integrated series. Level of significance of the test is 95%. In […] p-value. Number of panels A = 12 and number of panels B = 15. The number of delays was selected by the Akaike criterion, max. delays = 10. The variables were considered as logarithm. Cross-sectional dependence was eliminated (as per Levin *et al* 2002). Sample: 1994/01-2014/12.
Source: Developed by authors.

TABLE 2
ADF UNIT ROOT TEST: UNIVARIATE ANALYSIS

Variable	Level		First Difference	
	Statistical value	Integration order	Statistical value	Integration order
M*	1.835 (15 lags)	I(1)	–4.507 (14 lags)	I(0)
P	0.553 (14 lags)	I(1)	–3.847 (13 lags)	I(0)
P*	–0.941 (13 lags)	I(1)	–4.728 (12 lags)	I(0)

Note: Augmented Dickey-Fuller (ADF). Null hypothesis: there is a unit root. The number of delays was determined according to the Akaike criterion. The ADF model was specified without a constant; it was non-significant/insignificant in all cases. The variables were considered as logarithm. Level of significance: 10% (*), 5% (**) and 1% (***).
Source: Developed by authors.

TABLE 3
FISHER-TYPE UNIT ROOT TEST: CONDITIONAL VOLATILITY

Test		Statistic	
		MXE panel	CXE panel
Inverse chi-squared	P	178,141***	319,687***
Inverse normal	Z	–10,483***	–15,412***
Inverse logit	L*	–12,735***	–25,640***
Modified inv. chi-squared	Pm	19,125***	42,679***

Note: Fisher-type unit root test based on augmented Dickey-Fuller tests. Null hypothesis: All panels contain unit roots; alternative hypothesis: at least one panel is stationary. Specification with constant, no trend and removed cross-sectional shear mean. Level of significance: 10% (*), 5% (**) and 1% (***).
Source: Developed by authors based on IMF data.

Consequently, and based on the above unit root tests results, we include the stationary I (0) variables in levels and the non-stationary I (1) variables in first differences in equation 5 (see, for example, Love and Turk, 2014 and Gevorkyan, 2019, for a similar analysis).

6.2. Estimation results

P-VAR estimation was carried out for five different specifications of the REERV. Specifically, models 1 to 5 differ only in the way in which the measure of volatility was built. From the first to the fourth estimated equations, the REERV was calculated using the moving standard deviation for 4, 8, 12 and 24 periods, respectively. The fifth specification used the measure of conditional volatility.[16]

6.1.1. *Manufactures-exporting countries*

Table 4 presents the estimation results of Equation 7 for the MXE panel of countries using the alternative measures of REERV. The main findings of models 1 to 5 can be summarized as follows. On the one hand, regarding endogenous variables, firstly, the export variable lag is positive and significant at 1%. In other words, past changes in exports are relevant in explaining the contemporary exports. Secondly, the non-fuel commodity price index is negative and significant at 1%, except for model 5 where it is significant at 5%. These results are consistent with the fact that non-fuel commodities price

[16] Since we use a P-VAR model, *i.e.* a reduced and unrestricted simultaneous equations model, all endogenous variables affecting the model should be represented. However, for simplicity, we only report the equation that has exports as a variable to be explained; the rest of estimations of the different equations are available upon request.

index represent a loss of term of trade for these MXE countries. Thirdly, the volatility variable measured as the moving standard deviation of the REER is negative and significant at 1% (models 1-4) and is insignificant in the case of the conditional volatility specification (model 5). This result is associated with risk averse traders; therefore, episodes of high (low) exchange rate volatility are followed by a reduction (increase) of export flows. Likewise, the negative effect can be explained by the greater capacity to adjust production in response to exchange rate variability which is partly determined by the type of goods they export. Among the empirical literature that supports this negative result we can mention Chowdhury (1993) for the G-7 countries over the 1973-1990 period, Arize (1997) for seven industrial economies over 1973-1992, Verheyen (2012) for the bilateral trade from 11 countries of the European Monetary Union to US from 1995 to 2010, and Situ (2015) for developed countries during 1994-2014.

Regarding exogenous variables, first, the fuel commodity price index has negative and significant (at 1%) coefficients from models 1 to 5. This is due

TABLE 4

ESTIMATION RESULTS: MXE COUNTRIES

Equation: X	*V4* (1)	*V8* (2)	*V12* (3)	*V24* (4)	*V* (5)
L1.X	0.959***	0.922***	0.568***	0.579***	0.971***
	(0.008)	(0.014)	(0.036)	(0.035)	(0.006)
L1.P	–0.057***	–0.076***	–0.087***	–0.089***	–0.041**
	(0.019)	(0.020)	(0.018)	(0.019)	(0.018)
L1.Volatility	–1.332***	–1.554***	–1.985***	–2.098***	–0.010
	(0.310)	(0.405)	(0.380)	(0.658)	(0.008)
*M**	0.183***	0.257***	0.665***	0.669***	0.141***
	(0.027)	(0.034)	(0.041)	(0.041)	(0.024)
*P**	–0.060***	–0.057***	–0.078***	–0.075***	–0.055***
	(0.008)	(0.008)	(0.007)	(0.008)	(0.008)
d_t	–0.104***	–0.094***	–0.059***	–0.066***	–0.112***
	(0.010)	(0.012)	(0.012)	(0.011)	(0.010)
No. of obs.	3615	3495	3390	3210	3693
No. of countries	15	15	15	15	15
Avg. no. of T	241.000	233.000	226.000	214.000	246.200

Note: We considered the first difference of the variables' logarithms. Level of significance: 10% (*), 5% (**) and 1% (***). Equations 1, 2, 3 and 4 use the volatility of the REER calculated through the 4-, 8-, 12- and 24-period standard deviation moving averages, respectively. Equation 5 uses the measure of conditional volatility.

Source: Developed by authors.

to the fact that the MXE countries are mainly net importers of fuels; therefore, an increase in that index raises production and transportation costs, negatively affecting exports. Second, the global demand conditions have a positive impact on variation in exports, at a 1% significance level. Finally, the Great Recession variable had a clear negative effect on changes in exports, at a 1% significance level. This result is in line with prior theoretical and empirical trade literature. The impacts of the international financial crisis in 2008/2009 occurred in advanced economies, which in this sample of countries coincide mainly with the MXE countries panel. One possible explanation for the negative effect on exports involves the role of bank financing in trade.[17] According to Shelburne (2010), if an import transaction (the other side of the export transaction) is guaranteed by the banks' financing, there is a lower risk for the exporter to obtain the payment, whereas in the international financial crisis context, bank lending became more expensive, and export activity was reduced as a result of increased risk of and reduced access of importers to bank financing. However, even though bank financing has contributed as one of the mechanisms through which crises could affect exports, this is not the only one. In this sense, the OECD (2010) describes important additional channels through which crisis have affected exports. Firstly, crisis affects international trade indirectly through reduced consumption and therefore through the decline in demand for goods. With a declining demand for foreign goods, fewer imports are purchased and fewer exports are sold.[18] Secondly, the OECD (2010) argues that the way international trade reacts to financial crisis depends on the economic development level of the exporting country. Developing countries can be more dependent on trade exports relative to their GDP than developed economies. A trade slump therefore can have an amplified affect for developing countries. Available data indicates that trade in some regions –Asia, Middle East and Northern Africa and South America– was more severely impacted by changes in short-term trade finance than other regions (Europe and North America).[19] This may be due to the fact that some countries in these regions were considered higher risk, or their level of risk was re-evaluated after the onset of the crisis and thus due to increasing trade finance prices it became unaffordable for those countries. On the other hand the lack of integration with the international financial system could have been a blessing in disguise in protecting developing and emerging countries against negative chain reactions and providing those countries with a regional advantage and a gain in a competitive edge that would lead to a lesser decline

17 According to the IMF (2009a, 2009b) several banks reported sharp increases in the cost of trade finance-70% of the surveyed banks reported that the price for trade finance services has increased.

18 For more detailed analysis of this point see, for example, Eaton *et al*. (2016) and Cheung and Guichard (2009).

19 See Didier *et al*. (2012)

in trade and faster recovery. Finally, also is important to note that some studies have detailed additional mechanisms through which crises could affect exports. For example, Berman *et al*. (2012) analyzes the effect of the financial crisis on international trade covering the whole post-war era on a global scale and using a gravity-based approach. The fall in trade caused by financial crises is magnified by the time-to-ship goods between the origin and the destination country. In this sense, these authors strongly suggest that financial crises affect trade not only through demand but also through financial frictions that are specific to international trade.

6.1.2. *Commodity-exporting countries*

Table 5 shows the results of the P-VAR estimations for the different measures of the REERV for the CXE countries panel (models 1 to 5). On the one hand, the lag of the endogenous variables, such as exports, is significant at 1%. Moreover, the non-fuel commodity price index is significant at 5% in models 3 and 4. The positive sign on the non-fuel commodity price index means that the increase in prices encourages producers to increase exports. The REERV impact is not significant allowing us to disregard this variable as relevant in the model to explain the export variations (except in model 4). In other words, these results are consistent with prior evidence of a not insignificant effect. More specifically, this finding is consistent with the theoretical works of Clark (1973) and Ethier (1973), whose models suggest a negative or insignificant effect. In addition, Grier and Smallwood (2007) argue that it is possible that such an effect of exchange rate uncertainty on exports may be because export contracts are possible to adjust only in the long term. Finally, results are consist with the conclusion of Vilela and MacDonald (2016), who argue that there is no negative and significant effect of exchange rate volatility on exports for emerging and developing countries when oil export countries are excluded; our CXE country sample does not include them.

The exogenous variable, the fuel commodity price index is positive and significant. In other words, an increase in it leads to a rise in the energy commodities exports, and consequently, in total exports. This is because the share of fuel commodities in the exports of many of these economies is high, so rather than being a cost, it is an opportunity to increase their exports earnings. Moreover, the global demand conditions variable positively impacts exports and is the major determinant of them –a similar finding is reported in Bahmani-Oskooee and Gelan (2018) for African countries. Finally, we found the Great Recession to have negatively affected the exports of both MXE and CME countries, but it had only an insignificant impact on exports for the CXE countries. This is in line with the fact that agricultural and processed foods exports (relevant for CXE) experienced a smaller decline than manufactures exports (relevant for MXE) during the 2008 crisis.

TABLE 5
ESTIMATION RESULTS: CXE COUNTRIES

Equation: X	*V4* (1)	*V8* (2)	*V12* (3)	*V24* (4)	*V* (5)
L1.X	0.892*** (0.016)	0.812*** (0.029)	0.566*** (0.021)	0.564*** (0.021)	0.918*** (0.012)
L1.P	–0.006 0.029	–0.010 (0.032)	0.080*** (0.029)	0.071** (0.030)	0.006 (0.028)
L1.Volatility	0.038 (0.088)	–0.002 (0.087)	–0.123 (0.089)	–0.271** (0.131)	0.010 (0.007)
*M**	0.107** (0.045)	0.209*** (0.061)	0.302*** (0.044)	0.319*** (0.046)	0.066* (0.039)
*P**	0.023* (0.014)	0.029** (0.014)	0.054*** (0.013)	0.051*** (0.014)	0.020 (0.014)
d_t	–0.064 (0.043)	–0.056 (0.041)	–0.026 (0.039)	–0.024 (0.039)	–0.066 (0.043)
No. of obs.	2885	2793	2712	2568	2945
No. of countries	12	12	12	12	12
Avg. no. of T	240.417	232.750	226.000	214.000	245.417

Note: We considered the first difference of the variables' logarithms. Level of significance: 10% (*), 5% (**) and 1% (***). Equations 1, 2, 3 and 4 use the volatility of the REER calculated through the 4-, 8-, 12- and 24-period standard deviation moving averages, respectively. Equation 5 uses the measure of conditional volatility.
Source: Developed by authors.

6.3. Post-estimation tests

In this subsection, first, we report the Granger causality test, and then the IRF's of the endogenous and exogenous series for the MXE and CXE panels of countries.[20]

6.3.1. *Granger test*

The presence of correlation between two variables does not always imply causality (where changes in one of them determine the changes in the values of the other). In order to observe if causality exists between variables, we carried out a Granger causality test (Granger, 1969). Rejecting the null hypothesis implies that past changes in one variable affect, or precedes the changes of the other variable. Table 6 shows the results of the Granger causality test for the MXE and CXE panels; they are reported for REERV and exports.

[20] Variance decomposition results are not reported for brevity, but are available upon request.

TABLE 6
GRANGER CAUSALITY TEST (WALD)

H_0: Excluded variable does not Granger-cause equation variable H_1: Excluded variable Granger-cause equation variable			
Equation	*Excluded*	*Panel MXE*	*Panel CXE*
X	*V4*	18.526***	0.180
V4	*X*	16.424***	0.270
X	*V8*	14.753***	0.001
V8	*X*	7.517	0.082
X	*V12*	27.299***	1.917
V12	*X*	0.894	0.310
X	*V24*	10.175***	4.249**
V24	*X*	9.975***	0.111
X	*V*	1.500	2.112
V	*X*	1.707	3.239*

Notes: Rejection of the null hypothesis: 10% (*), 5% (**) and 1% (***) of significance (prob.>chi2). Sample: 1994/01-2014/12. The variables were considered as logarithm. Results are reported for exports and the different measures of volatility. V4, V8, V12, V24, and V refer to the 4-, 8-, 12- and 24-period standard deviation moving averages and the conditional volatility, respectively.
Source: Developed by authors.

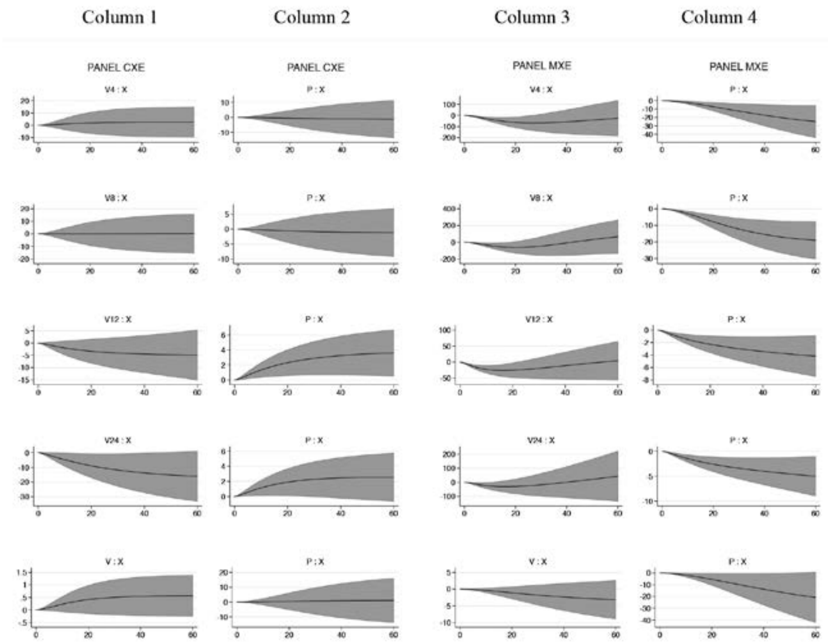
Here, we find a unidirectional significant relationship wherein the 8- and 12-period moving standard deviations Granger-cause exports for the MXE countries panel. While bidirectional Granger causality is found for the 4- and 24-period moving standard deviations, causality in these relationships is not conclusive. As far as CXE countries are concerned, the 24-period moving standard deviation causes exports in the Granger sense.

6.3.2. *Impulse-response functions*

Here, we discuss the simulation of the accumulated IRF's. The focus of the analysis is to quantify macroeconomic shocks one at a time to see how they affect exports, with particular interest in the impact of an exchange rate volatility shock. In the IRF's graphs, the export response is represented by an orthogonal impulse or shock, one standard deviation in magnitude, to the non-fuel commodity price index and the REERV measures. The exports response is considered for a period of 60 months (5 years). We assume the following recursive order to construct the IRF:

$$P \rightarrow V \rightarrow X$$

FIGURE 4
ACCUMULATED IMPULSE RESPONSE FUNCTION: ENDOGENOUS VARIABLES



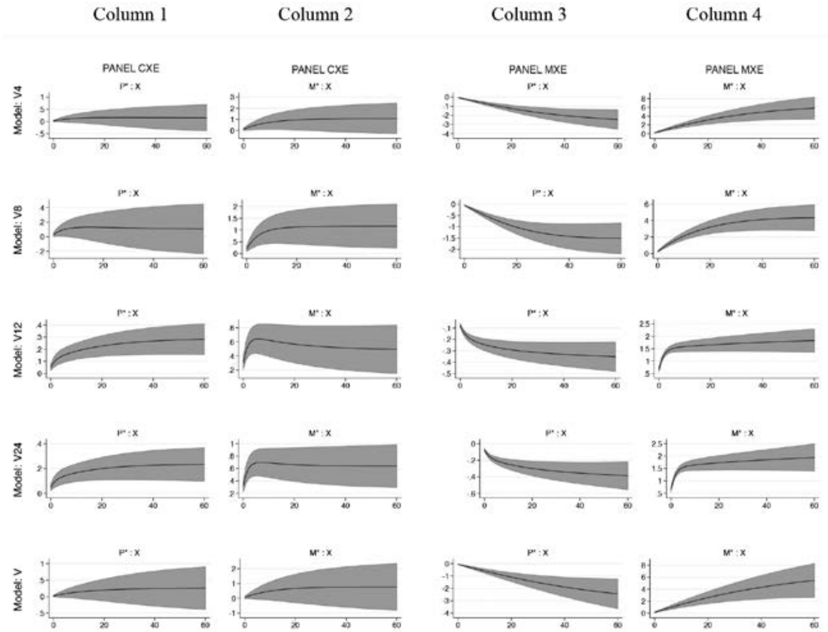
Note: The impulse is the endogenous variable and the response variable is exports. The band containing the cumulative IRF corresponds to the 95% confidence.
Source: Developed by authors.

The economic intuition of this Cholesky order can be expressed as follows: firstly, the non-fuel commodities price index is the most important variable for the MXE and CXE panels, based on its effect on the terms of trade and thus on the decision of the countries to export.[21] Secondly, due to the effect of uncertainty on exports, the exchange rate volatility cannot be accurately predicted. Given that exports are presumed to respond at the same time as the rest of the variables in the system, it is in last position in Cholesky's order.

Figure 4 illustrates the accumulated IRF's of the endogenous variables pertaining to the non-fuel commodity price index and REERV (by row) for the MXE panel (columns 1-2), and the CXE panel (columns 3-4). Meanwhile, an REERV shock generates an export response in the short- and medium-term for the MXE panel, but this is not significant for the CXE panel.

[21] See Gevorkyan (2019), for more detailed explanation of the Cholesky order considered.

FIGURE 5
ACCUMULATED IMPULSE RESPONSE FUNCTION: EXOGENOUS VARIABLES



Note: The impulse is the exogenous variable and the response variable is exports. The band containing the cumulative IRF corresponds to the 95% confidence.
Source: Developed by authors.

In addition, the P-VAR methodology allows an IRF to simulate a shock (here, a twofold increase) to the exogenous variable and its effects on the endogenous variable of interest. The results are illustrated in Figure 5: REERV specifications are depicted in each row, and the accumulated IRF's of the exogenous variables associated with the fuel commodity price index is shown in columns 1-2 for the MXE panel and those associated with global demand in columns 3-4 (CXE panel).

A shock to global demand generates a positive short- and medium-term exports response for both panels. Specifically, a one-standard deviation unit shock to global demand results in about 0.8% increase in exports for the CXE in twenty periods (months), and a one standard deviation shock to global demand results in about 3% increase in export for MXE in twenty periods (months), where the shocks seem to stabilize. Both short- and medium-term negative export responses are generated by an impulse of the exogenous variable (the fuel commodity price index) for the MXE panel, and positive or insignificant export responses for the CXE panel. Particularly, a visual inspection of IRF's allow us to observe that a one standard deviation shock to fuel commodity price index causes a significant decrease in exports for MXE countries for twenty periods after which the effect dissipates. The decrease peaks is in period twelve. And a one standard deviation shock to fuel commodity price index cause significant increase in exports for CXE countries for twenty periods after which the effect dissipates. The increase peak is in period ten.

6.4. Extension: Advanced economies vs. developing and emerging economies

In this subsection we propose an additional empirical analysis. Now, instead of focusing on the export-related characteristics of our sample of countries, we split the sample by development level of countries, *i.e.* advanced economies and developing and emerging economies. Therefore, Australia and New Zealand are excluded from the CXE sample and included in the MXE sample.

General speaking, the impact of REERV on exports does not change when excluding Australia and New Zealand from the CXE sample (see Table 7), significance and sign do not change relative to the reference model (see Table 5). When Australia and New Zealand are grouped together with the European countries, significance and sign do not change relative to the reference model (see Table 4). Thus, these results suggest that the level of development does affect the relationship between REERV and exports for this sample of countries; significant for advances economies and insignificant for developing and emerging economies.

7. Conclusions

This paper focused on the relationship between exchange rate uncertainty and exports for a novel panel of 27 countries over 1994/01–2014/12 using the P-VAR empirical methodology. This issue was tackled by building a high frequency dataset and employing a novel empirical methodology (P-VAR). Also, differently from prior empirical analysis that focuses on the level of development of economies (see, for example, Sauer and Bohara, 2001 and Grier and Smallwood, 2007); we provide novel insight into the relationship between exchange rate volatility and exports by considering the production characteristics of the countries, *i.e.* manufactures-exporting economies (MXE) and commodity-exporting economies (CXE).

Our main empirical findings suggest the following conclusions. First, REERV is important for modeling the exports of MXE countries, but is not relevant in the case of CXE countries. The economic interpretation of the results obtained could be based on the response in the "average" exporting country with respect to exchange rate risk. While the negative effect of REERV on exports in the MXE sample appears to be associated with countries that display risk-averse behaviors or have some contract flexibility to adjust their exports in the short term, the lack of the effect of REERV on exports in the CXE sample seems to

TABLE 7

ESTIMATION RESULTS: ADVANCED ECONOMIES VS. DEVELOPING AND EMERGING ECONOMIES

Equation: X	*V4* (1)	*V8* (2)	*V12* (3)	*V24* (4)	*V* (5)
Advanced economies					
L1.X	0.959*** (0.008)	0.919*** (0.014)	0.582*** (0.032)	0.586*** (0.032)	0.970*** (0.006)
L1.P	–0.048*** (0.018)	–0.061*** (0.018)	–0.062*** (0.017)	–0.065*** (0.018)	–0.032* (0.017)
L1.Volatility	–1.130*** (0.232)	–1.299*** (0.299)	–1.532*** (0.294)	–2.232*** (0.498)	–0.008 (0.008)
*M**	0.166*** (0.025)	0.234*** (0.032)	0.598*** (0.036)	0.606*** (0.036)	0.125*** (0.022)
*P**	–0.052*** (0.008)	–0.497*** (0.007)	–0.064*** (0.007)	–0.061*** (0.007)	–0.047*** (0.008)
d_t	–0.086*** (0.018)	–0.078*** (0.018)	–0.047*** (0.018)	–0.046*** (0.018)	–0.094*** (0.017)
No. of obs.	4097	3961	3842	3638	4183
No. of countries	17	17	17	17	17
Avg. no. of T	241.000	233.000	226.000	214.000	246.059
Developing and emerging economies					
L1.X	0.891*** (0.018)	0.811*** (0.032)	0.560*** (0.022)	0.559*** (0.023)	0.918*** (0.013)
L1.P	–0.010 (0.033)	–0.022 (0.037)	0.068** (0.034)	0.056 (0.035)	0.001 (0.032)
L1.Volatility	0.054 (0.091)	0.019 (0.089)	–0.112 (0.090)	–0.222* (0.134)	0.008 (0.007)
*M**	0.106** (0.052)	0.223*** (0.071)	0.336*** (0.052)	0.351*** (0.054)	0.066 (0.045)
*P**	0.024 (0.016)	0.031* (0.016)	0.057*** (0.016)	0.054*** (0.016)	0.020 (0.016)
d_t	–0.086* (0.048)	–0.078* (0.047)	–0.044 (0.045)	–0.043 (0.045)	–0.088* (0.049)
No. of obs.	2403	2327	2260	2140	2455
No. of countries	10	10	10	10	10
Avg. no. of T	240.300	232.700	226.000	214.000	245.500

Note: We considered the first difference of the variables' logarithms. Level of significance: 10% (*), 5% (**) and 1% (***). Equations 1, 2, 3 and 4 use the volatility of the REER calculated through the 4-, 8-, 12- and 24-period standard deviation moving averages, respectively. Equation 5 uses the measure of conditional volatility. Advanced economies: European countries, Australia and New Zealand. Developing and emerging economies: South American countries (see Table A.1).

Source: Developed by authors.

be associated with countries that have contract rigidities, which enables them to adjust exports in the short term. Second, this paper also reports evidence of the relationship between exports and other explanatory macroeconomic variables. Furthermore, world demand conditions are one of the most important factors explaining variations in exports. In contrast with Vilela and MacDonald (2016), who argue for an increase in exports after the financial crisis period, our finding reveals that the Great Recession reduced exports of MXE countries.

Our results provide important insights in relation to macroeconomic policy. Note that REERV is not a policy variable directly controlled by policymakers. If policymakers ignore the unpredictability of exchange rate movements, however, export markets may underlie the uncertainty of outcomes. Thus, this empirical analysis leads us to suggest to minimize exchange-rate volatility and its persistence, by mitigating nominal exchange rate fluctuations, in order to reduce the risks associated with export activity, and consequently, to stabilize the external trade position. Finally, it is important to note that using the same policies would likely have divergent effects on the two panels of countries, particularly given that MXE countries have higher market integration and more advanced production than CXE countries.

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APPENDIX

TABLE A.1
COUNTRY LIST

Panel: CXE		Panel: MXE	
South América	*Oceanía*	*Europe*	
Argentina Bolivia Brazil Chile Colombia Ecuador Paraguay Peru Uruguay Venezuela	Australia Nueva Zealand	Austria Belgium Denmark Finland France Germany Greece Ireland Italy Luxemburg	the Netherlands Portugal Spain Sweden United Kingdom

TABLE A.2
DEFINITIONS AND VARIABLE SOURCES

Variable	*Description*	*Source*
Exports (*X*)	Total exports of goods in millions of constant dollars (Base January 1994 = 100) (exports in millions of current FOB dollars, deflated by the US CPI).	IMF; Luxemburg 1994/01-1996/12 and Greece 1994/09 and 1994/10, source Eurostat.
World Demand (*M**)	World imports of goods in millions of constant dollars (Base January 1994 = 100) (imports in millions of current CIF dollars, deflated by the US CPI).	IMF
CPI	United States Consumer Price Index (US CPI) (Base January 1994 = 100).	Department of Labor Bureau of Labor Statistic U.S.
Real Effective Exchange Rate (*REER*)	The index considers the weighted average of the bilateral real exchange rates with the main trading partners (using as weighting the share of trade in the economies) (Base January 1994 = 100).	IMF; Perú (ECLAC); Argentina (CEI).
Volatility (*V4*)	Volatility of the real effective exchange rate average, 4 periods.	IMF; Perú (ECLAC); Argentina (CEI).
Volatility (*V8*)	Volatility of the real effective exchange rate average, 8 periods.	IMF; Perú (ECLAC); Argentina (CEI).
Volatility (*V12*)	Volatility of the real effective exchange rate average, 12 periods.	IMF; Perú (ECLAC); Argentina (CEI).
Volatility (*V24*)	Volatility of the real effective exchange rate average, 24 periods.	IMF; Perú (ECLAC); Argentina (CEI).
Volatility (*V*)	Standard deviation of the conditional variance.	IMF; Perú (ECLAC); Argentina (CEI).
P	Index of non-fuel commodities prices (Base January 1994 = 100).	IMF
*P**	Index of fuel commodities prices (energy) (Base January 1994 = 100).	IMF

Source: Developed by authors.

TABLE A.3
SUMMARY STATISTICS: CXE AND MXE PANELS

Variable		*Averages*	*Standard deviation*	*Minimum*	*Maximum*	*Observations*
Panel: CXE						
X	Overall	7.050	1.293	4.162	9.734	*N = 3024*
	Between		1.234	5.227	8.803	n = 12
	Within		0.525	5.842	8.435	*T = 252*
V4	Overall	0.022	0.023	0.002	0.247	*N = 2976*
	Between		0.007	0.011	0.038	*n = 12*
	Within		0.022	–0.013	0.248	*T = 248*
V8	Overall	0.023	0.022	0.003	0.179	*N = 2928*
	Between		0.008	0.012	0.042	*n = 12*
	Within		0.021	–0.013	0.180	*T = 244*
V12	Overall	0.024	0.021	0.003	0.147	*N = 2880*
	Between		0.009	0.012	0.044	*n = 12*
	Within		0.019	–0.014	0.147	*T = 240*
V24	Overall	0.026	0.020	0.004	0.108	*N = 2736*
	Between		0.010	0.012	0.048	*n = 12*
	Within		0.017	–0.015	0.105	*T = 228*
V	Overall	4.753	0.518	3.569	6.622	*N = 3007*
	Between		0.429	3.845	5.433	*n = 12*
	Within		0.315	4.056	6.830	*T–bar = 250.583*
Panel: MXE						
X	Overall	8.943	1.229	5.638	11.463	*N = 3780*
	Between		1.237	6.537	10.904	*n = 15*
	Within		0.284	7.624	9.706	*T = 252*
V4	Overall	0.008	0.005	0.001	0.059	*N = 3720*
	Between		0.003	0.005	0.013	*n = 15*
	Within		0.005	–0.003	0.054	*T = 248*
V8	Overall	0.008	0.005	0.001	0.046	*N = 3660*
	Between		0.003	0.005	0.014	*n = 15*
	Within		0.004	–0.001	0.040	*T = 244*
V12	Overall	0.008	0.005	0.002	0.039	*N = 3600*
	Between		0.003	0.005	0.014	*n = 15*
	Within		0.004	0.000	0.033	*T = 240*
V24	Overall	0.008	0.004	0.002	0.031	*N = 3420*
	Between		0.003	0.005	0.014	*n = 15*
	Within		0.003	0.001	0.025	*T = 228*
V	Overall	4.659	0.432	3.651	6.193	*N = 3759*
	Between		0.376	4.141	5.578	*n = 15*
	Within		0.234	3.555	5.592	T-bar = 250.6

Note: All variables are expressed as logarithm. Period: 1994 to 2014.
Source: Developed by authors based on IMF data.

TABLE A.4
SUMMARY STATISTICS

Variable		*Averages*	*Standard deviation*	*Minimum*	*Maximum*	*Observations*
P	Overall	4.855	0.308	4.396	5.473	*N = 6804*
	Between		0.000	4.855	4.855	*n = 27*
	Within		0.308	4.396	5.473	*T = 252*
*P**	Overall	5.582	0.697	4.331	6.756	*N = 6804*
	Between		0.000	5.582	5.582	*n = 27*
	Within		0.697	4.331	6.756	*T = 252*
*M**	Overall	13.308	0.359	12.599	13.848	*N = 6804*
	Between		0.000	13.308	13.308	*n = 27*
	Within		0.359	12.599	13.848	*T = 252*

Note: All variables are expressed as logarithm and are the same for each country. Period: 1994 to 2014.
Source: Developed by authors based on IMF data.

TABLE A.5
EQUATIONS OF THE CONDITIONAL VARIANCE

Country	Specification	Coefficients		
		C	$RESID^2_{t-1}$	$GARCH_{t-1}$
Argentina	GARCH(1,1)	3.78E-05** (1.78E-05)	0.4647*** (0.1280)	0.4508*** (0.1147)
Bolivia	GARCH(1,1)	3.86E-05* (2.16E-05)	0.1817*** (0.0656)	0.5758*** (0.1766)
Brazil	GARCH(1,1)	0.0001** (3.98E-05)	0.3189*** (0.0793)	0.6142*** (0.0855)
Chile	GARCH(1,1)	2.74E-05* (1.57E-05)	0.0568* (0.0327)	0.8764*** (0.0611)
Colombia	GARCH(1,1)	0.0002*** (5.54E-05)	0.1895*** (0.0450)	0.5009*** (0.0770)
Ecuador	GARCH(1,1)	3.31E-05** (1.47E-05)	0.4190*** (0.1089)	0.5598*** (0.0974)
Paraguay	GARCH(1,1)	4.46E-05*** (1.61E-05)	0.1517*** (0.0405)	0.7846*** (0.0617)
Uruguay	GARCH(1,1)	0.0002*** (2.52E-05)	0.4784*** (0.1243)	0.2000** (0.0842)
Venezuela	GARCH(1,1)	0.0004*** (3.76E-05)	0.4454*** (0.1180)	0.3355*** (0.0597)
Australia	IGARCH(1,1)	—–	0.0632*** (0.0227)	0.9368*** (0.0227)
New Zealand	IGARCH(1,1)	—–	0.0587** (0.0233)	0.9413*** (0.0233)
Germany	IGARCH(1,1)	—–	0.0785*** (0.0284)	0.9215*** (0.0284)

Table A.5 (Cont.)

Country	Specification	Coefficients		
		C	$RESID^2_{t-1}$	$GARCH_{t-1}$
Austria	IGARCH(1,1)	—	0.0433*** (0.0146)	0.9567*** (0.0146)
Belgium	IGARCH(1,1)	—	0.0727*** (0.0259)	0.9273*** (0.0259)
Denmark	IGARCH(1,1)	—	0.0540*** (0.0142)	0.9460*** (0.0142)
Spain	IGARCH(1,1)	—	0.0553*** (0.0191)	0.9447*** (0.0191)
Finland	IGARCH(1,1)	—	0.0454*** (0.0145)	0.9546*** (0.0145)
France	IGARCH(1,1)	—	0.1007*** (0.0249)	0.8993*** (0.0249)
Greece	GARCH(1,1)	1.27E-05* (7.36E-06)	0.2075*** (0.0710)	0.5661*** (0.1721)
Ireland	GARCH(1,1)	8.89E-06** (4.41E-06)	0.1835*** (0.0547)	0.7514*** (0.0772)
Italy	GARCH(1,1)	1.85E-06** (2.1181)	0.0935*** (0.0296)	0.8703*** (0.0340)
Luxemburg	IGARCH(1,1)	—	0.0853*** (0.0206)	0.9147*** (0.0206)
the Netherlands	IGARCH(1,1)	—	0.1023*** (0.0227)	0.8977*** (0.0227)
Portugal	IGARCH(1,1)	—	0.0689*** (0.0129)	0.9311*** (0.0129)
United Kingdom	IGARCH(1,1)	—	0.0958*** (0.0166)	0.9042*** (0.0166)
Sweden	IGARCH(1,1)	—	0.0802*** (0.0243)	0.9198*** (0.0243)

Note: Model parameters were estimated by Maximum likelihood (ML) - Normal distribution. Level of significance to: 10% (*), 5% (**) and 1% (***).
Source: Developed by authors based on IMF data.

TABLE A.6
CONDITIONAL VARIANCE EQUATION FOR PERU

Variable	Coefficient	Std. Error	Prob.
C(3)	–2.4353	1.3962	0.0811*
C(4)	0.2947	0.1330	0.0267**
C(5)	–0.1659	0.0885	0.0610*
C(6)	0.7531	0.1532	0.0000***

Note: LOG(GARCH) = C(3)+C(4)*ABS(RESID(-1)/@SQRT(GARCH(-1)))+C(5)*RESID(-1)/@SQRT(GARCH(-1))+C(6)*LOG(GARCH(-1)).
Source: Developed by authors based on IMF data.

TABLE A.7
EIGENVALUE STABILITY CONDITION

Model	Eigenvalue		Modulus
	Real	Imaginary	
Panel: MXE			
V4	0.9625	–0.0120	0.9625
	0.9625	0.0120	0.9625
	0.8394		0.8394
V8	0.9545	–0.0290	0.9550
	0.9545	0.0290	0.9550
	0.8846		0.8846
V12	0.9842		0.9842
	0.9009		0.9009
	0.5595		0.5595
V24	0.9952		0.9952
	0.9316		0.9316
	0.5686		0.5685
V	0.9750		0.9750
	0.9572		0.9572
	0.9174		0.9174
Panel: CXE			
V4	0.9608		0.9608
	0.8921		0.8921
	0.8553		0.8553
V8	0.9547		0.9547
	0.9372		0.9372
	0.8109		0.8109
V12	0.9542		0.9542
	0.9400		0.9400
	0.5701		0.5701
V24	0.9568	–0.0122	–0.9568
	0.9568	0.0122	0.9568
	0.5687		0.5687
V	0.9579		0.9579
	0.9203		0.9203
	0.7940		0.7940

Source: Developed by authors.

A sensitivity analysis on the impact of regional trade agreements in bilateral trade flows*

Un análisis de sensibilidad sobre el impacto de los acuerdos comerciales regionales en los flujos bilaterales de comercio

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Abstract

We estimate the effect of RTAs on bilateral exports by means of a gravity model analyzing its sensitivity to different specifications and methods. RTAs generate a sizable positive effect. However, shifting to country-pair and time-varying fixed effects systematically reduces coefficients. Nevertheless, the RTA effect is consistent across methods and specifications.
The RTA effect attributable to particular trade agreements displays high variability. While most RTAs increase trade, others present non-significant or negative results. We apply robustness checks to individual RTA estimates by presenting PPML time-invariant fixed effects and next to these, country-pair and time-varying fixed effects estimates. Thus, 38.2% of RTAs are positive and significant in both specifications. RTAs trade creation effects tend to prevail over trade diversion effects.

Key words: *International Trade, Trade Liberalization, Regional Trade Agreements RTA, Gravity Model, Economic Integration.*

JEL Classification: *F13, F14, F15, F53, F55.*

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Resumen

En este artículo estimamos, mediante un modelo de gravedad, el efecto de los Acuerdos Comerciales Regionales (ACR) en las exportaciones bilaterales, y realizamos un cuidadoso análisis de sensibilidad, considerando diferentes métodos y especificaciones. Los ACR presentan, por lo general, un efecto positivo considerable. Este impacto se reduce substancialmente al incluir efectos fijos país variables en el tiempo y efectos fijos individuales. No obstante, el efecto de los ACR es consistente a través de los métodos y especificaciones aplicados. Cuando el impacto de los ACR es calculado para cada acuerdo en particular, los coeficientes presentan una alta variabilidad. La mayoría de ACR presenta un impacto positivo. Otros presentan resultados no significativos o negativos. Para una mayor robustez de los resultados, los impactos de los ACR particulares fueron estimados con efectos fijos invariables en tiempo, y también con efectos fijos variables en el tiempo bajo el método de PPML. Así, el 38,2% de los ACR son positivos y significativos en ambas especificaciones. A su vez, los efectos de creación de comercio tienden a prevalecer sobre los efectos de desviación del comercio.

Palabras clave: *Comercio Internacional, liberalización económica, Acuerdos Comerciales Regionales ACR, modelo de gravedad, integración económica.*

Clasificación JEL: *F13, F14, F15, F53, F55.*

1. Introduction

After its creation in 1995, the World Trade Organization has been able to convince most of the countries to abide by the rules of multilateral trade. Nevertheless, its rounds of negotiations have come to a deadlock, partly explained by the difficulty of making agreements among too many countries of a heterogeneous nature. In the midst of this, Regional Trade Agreements RTAs, appeared as a more effective way to close trade deals. They presented exponential growth from the 80s to the first decade of the current century, to slow down in recent years. The question, then, arises about its effectiveness.

Trump's administration has dispensed with more than 70 years of liberal tradition in the United States, by dumping the Trans-Pacific Partnership (TPP) and the Transatlantic Trade and Investment Partnership (TTIP), the two most ambitions RTAs ever negotiated, while initiating a frontal trade war with China, and imposing duties on aluminium and steel worldwide. In the same direction, the United Kingdom has officially divorced from the most admired and profound RTA, the European Union. In a time where liberal ideas are under strain, answering the question of whether RTAs really increase trade is even more important. Despite substantial progress to compute RTA estimates, the debate about the effectiveness of RTAs remains open. The main objective of this paper is then to help answer the question: To what extent are RTAs able to create trade?

To do it, we employ the widely accepted approach of the gravity model. We build on the works of Baier and Bergstrand (2007), Martínez-Zarzoso *et al.* (2009); Kohl (2014) and Baier *et al.* (2019).

The main contribution of this paper is the presentation of a comprehensive sensitivity analysis based on a battery of relevant regression methods and specifications applied to the gravity equation on an updated database, providing the possibility of easy visualization and comparison, see Table 1. This information will also be valuable for future meta-analysis studies about the effect of RTAs on trade. We subsequently explore the RTA effect on bilateral trade for an ample sample of particular RTAs. Thus, coefficients for 123 particular RTAs comparing PPML time-invariant fixed effects (TIFE) and time-varying fixed effects (TVFE) estimates are presented in Table 3, which enables us to carry out robustness checks on their effectiveness. As far as we know, we are the first to present this comparison and analysis at disaggregated level for a large number of RTAs over a long period, see Table 4. Finally, we present results on trade creation and trade diversion for 25 relevant RTAs in Table 5.

Results from our most relevant specifications and methods point to a positive and significant effect of RTAs between 4.7% and 51.3% on bilateral exports. Considering particular RTAs, their impact is predominantly positive and significant. Trade creation effects in most of the cases offset trade diversion effects.

Gravity model estimations define what should be the normal pattern of trade, and then enable us to seek deviations from it, originated, for example, in the implementation of institutional arrangements. Given the counterfactual it offers, and its widespread use, the gravity model is tenable for calculating outcomes such as the expected gains from the entry into force of an RTA, or other institutional changes.

One important advantage of gravity models according to Bussière (2009) is that their results stem not only from a measure of multilateral trade integration (a country against all its trading partners), but also of bilateral trade integration (a country and each of its trading partners).

Our interest in finding the effects of RTAs in bilateral trade flows, hinges on the belief that higher international competition leads to greater productivity and higher cross-border exchanges increase wellbeing. We do not intend to disentangle this effect, although from Sachs *et al.* (1995), Wacziarg and Welch (2008) we have evidence that international trade promotes economic growth and then wellbeing. Similarly, Halpern *et al.* (2015) have established a positive relationship between firm import input access and productivity in the Hungarian economy, and Bas and Ledezma (2010) provided evidence of trade barriers reduction and with-in plant productivity increases in Chile.

In 1980, the GATT counted up to 83 signatories. In 2020 the number practically doubles, reaching 164 countries, now under the label of the WTO. Hayakawa and Kimura (2015) found that free trade agreements (FTAs) successfully reduce tariff rates and non-tariff barriers (NTBs). Nevertheless, mixed effects were reported by Afesorgbor (2017), Caporale *et al.* (2012), Didia *et al.* (2015), Kahouli & Maktouf (2015), Martin-Mayoral *et al.* (2016) who studied the impact of RTAs on exports by trade blocs in different regions as the Americas, Africa or Europe. Their results maintain alive a long-standing debate on the optimal mechanism for liberalizing international trade, confronting the multilateral negotiation approach to RTAs.

It is expected that membership to multilateral trade institutions would bear a strong positive effect on trade. Strong evidence for a positive WTO membership effect was found by Rose (2005), Subramanian and Wei (2007) and Kim (2011). Nevertheless, Eicher and Henn (2011) found evidence of an attenuated WTO membership impact after preferential trade agreements had entered into force. In view of the historical importance of this institution, this paper controls for country membership status in the WTO.

In parallel, the number of physical RTAs in force has steadily grown from 1980 to 2019. There were only 15 RTAs in 1980. They rose to 51 in 1995, 137 in 2005 to reach the number of 303 in 2019. Despite a slowdown in the number of new RTA negotiations worldwide, more RTAs are expected see the light in the years to come.

Special attention has been paid to Baier and Bergstrand (2007) who, using panel data on five years intervals, found that the average treatment effect of an RTA implies an increase of bilateral exports around 100% in 10 years. Another important contribution came from Magee (2008) who let the RTA dummy take leads and lags, thus finding significant anticipatory and slow motion impacts. Thus, in the long-run, an RTA increases trade on average by 89%. Regarding dynamics, Martínez *et al.*, (2009), remark that bilateral exports are persistent and find significant effects for the lagged bilateral export flows, as well as for RTA coefficients at the disaggregated level.

RTA estimates have recently been reviewed downwards, a result that we confirm in this paper. This erosion effect was detected by De Sousa (2012) who focused on the effect of currency unions, and later by Kohl (2014) applying the Baier and Bergstrand's technique where he found that RTAs increased trade by at most 50%. Proving the reasons behind this behaviour goes beyond the scope of this paper. Yet, some hypothesis point out to the appearance of diminishing returns as more and more countries engage in RTAs; a rise in transaction costs coming from the multiplication of non-tariff measures such as rules of origin and local content requirements, or even a relaxation in the enforceability of existing RTAs due to a political movement of resistance to trade liberalization.

Following this introduction, the paper is organized as follows. Section 2 discusses methodological issues. Section 3 presents the data. Section 4 sets up the econometric specifications to be estimated. Section 5 presents and analyses results and section 6 concludes the paper.

2. Gravity model and methodology

Important advances in the micro-foundation of the gravity model are attributed to Anderson (1979) and Anderson and Van Wincoop (2003). They set up a model in which consumers maximize a homothetic Cobb-Douglas utility function that is identical in all countries; goods are differentiated by their country of origin, iceberg costs are assumed and only a fraction of the goods arrives at destination.

The mathematical approach developed by them puts multilateral resistance in the spotlight of the analysis. Their model takes us to estimate:

$$x_{ij} = \frac{y_i y_j}{y^w}\left(\frac{t_{ij}}{P_i P_j}\right)^{1-\sigma} \qquad (1)$$

Where x_{ij} represents exports from country $_i$ to country $_j$; y^w, y_i and y_j represent world, country i and country j's GDPs, respectively; t_{ij} is a trade cost factor between i and j, consisting of geographical, political and institutional barriers. The parameter σ represents the elasticity of substitution between all goods and P_i and P_j are the multilateral resistance terms, which give us a measure of the relative openness of the economies. The Gravity equation is compatible with several underlying theories. A detailed discussion about the gravity model micro-foundation is available in Head and Mayer (2014).

Augmented gravity models control for confounders, which, if omitted, would bias the estimate of our parameter of interest on RTAs. Hence, we control for border contiguity and other cultural or institutional variables such as the use of a common language, Melitz and Toubal (2014), and colonial links, Head *et al*. (2010).

Anderson and van Wincoop (2003) pointed out the difficulties in estimating unbiased coefficients through cross-sections as well as the threat of omitted variable bias derived from multilateral resistance. Thus, Panel data models enabling for fixed effects specifications provided a solution to the fact that P_i and P_j, the so-called multilateral resistance terms in equation (1) are unobservable and the procedure to estimate them implies a non-linear routine. De Benedictis and Taglioni (2011) examine the sensitivity of OLS estimates to variations in fixed effects. These procedures control for endogeneity from unobservable heterogeneity and then for omitted variable bias derived from multilateral resistance. The authors consider the introduction of time-varying fixed effects for importing and exporting countries a robust solution.

Apart from the multilateral resistance difficulty, the possibility of endogeneity between bilateral trade and institutional trade liberalization variables is also prominent. Trefler (1993) pointed out that a country's decision to sign a regional trade agreement could not be completely exogenous. In the same way, Ghosh and Yamarik (2004) based on extreme bounds analysis showed that the RTAs coefficient computed with cross-sectional data could be biased in the presence of endogeneity and Baldwin and Jaimovich (2012) found that free trade agreements could be contagious (domino effect). When endogeneity is present, traditional estimation methods could result in inconsistent estimates. Instrumental variable methods can deal with endogeneity, allowing for stronger causal claims.

In that vein, Baier and Bergstrand (2007, 92) stated "standard cross-section techniques using instrumental variables and control functions do not provide stable estimates of RTA average treatment effect in the presence of endogeneity, and tests of over-identifying restrictions generally fail". They suggested that panel data methodologies must be implemented to estimate the RTA coefficient.

A panel approach will then be preferred over cross-section because it accounts better for country observed and unobserved time-varying or time-invariant heterogeneity. It provides the possibility of controlling for relevant relationships over time, avoiding the risk of choosing an unrepresentative year Antonucci and Manzocchi (2006). Panels also improve the efficiency of the estimates, Cheng Hsiao (2003). The panel structure would deal relatively well with the endogeneity problem considering that the reasons linked to RTAs not being exogenous should most probably be related to time-invariant heterogeneity (huge pre-existing trade flows, or contiguity).

Not all RTAs are equal. Dür *et al.* (2014) created deep integration indicators proving that differences on the depth of the agreements produces weaker effects for shallow agreements. Considering the heterogeneity of economic integration agreements, Egger and Nigai (2015) concluded that shifting to deeper trade agreement increases welfare, this effect being particularly high for some countries. Kohl *et al.* (2016); Ahcar and Siroën (2017) confirmed the effects of deep integration on trade, where deeper agreements result in larger gains. Baier *et al.* (2018) also found that certain integration settings produce greater impacts on the intensive margin than on the extensive margin.

Seeking better predictions of the effect of new economic integration agreements, Baier *et al.* (2018) and Baier *et al.* (2019) went beyond the importance of accounting for RTA heterogeneity. They found asymmetries in the RTA effect linked to the direction of trade. They also proved that country-pair heterogeneity is relevant as any given integration agreement can produce different effects on trade. For example, partners engaged in pre-existing economic integration agreements and distant pairs of countries obtain weaker gains out of further integration.

Considering that the main objective of this paper is to compare the average effect of an RTA through different specifications and methods, we want to make the caveat that not all RTA are equally designed. Hence, we do not expect to interpret these coefficients as precise predictions of the effect of any new RTA agreement, as literature acknowledges that information on RTA heterogeneity is required for accurate forecasting purposes. Nevertheless, to mitigate these shortcomings, we estimate the RTA effect for particular couples and blocks, where we can observe a long range of variability on the effect of RTA, possibly caused by this heterogeneity.

3. Data

To deal with the challenges mentioned above and to successfully estimate our variables of interest, this research set up an exhaustive data set to run a gravity model. It consists of bilateral trade flows for 153 countries from 1980 to 2018 that add up to 715.626 individual bilateral trade flows and an extensive set of control variables.

Bilateral Exports are taken in current dollars at fob values from the International Monetary Fund (IMF) Direction of Trade Statistics Database DOTS (2020). The current GDP in dollars, population in number of inhabitants and urban participation in percentages are provided by the World Development Indicators (WDI) database of the World Bank (2020). The surface in square meters as well as island and landlocked status were constructed by the author based on data from the World Factbook of the Central Intelligence Agency of the United States of America CIA (2020). Weighted distance in Km, common land border and colonial links stem from the CEPII (2013): Head *et al*. (2010) Gravity dataset.

The dummy variable for Regional Trade Agreements was constructed by the author based on the Regional Trade Agreements Information System (RTA-IS) of the World Trade Organization WTO (2020), and from de Sousa RTA data set for De Sousa (2012). Generalized System of Preferences GSP is built by the author based on the Database on Preferential Trade Arrangements of the World Trade Organization WTO (2020). The author based on the World Trade Organization WTO information (2020) constructed GATT membership and OECD membership based on information from the Organisation for Economic Co-operation and Development (OECD) (2020).

4. Econometric specifications

The equation to estimate with OLS, with time-fixed effects and exporter and importer time-invariant fixed effects is presented in (2) below:

$$lnX_{ijt} = \beta_0 + \beta_1 RTA_{ijt} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{jt}} + \varphi_g \boldsymbol{Z_{ijt}} + \alpha_t + \alpha_i + \alpha_j + \varepsilon_{ijt} \qquad (2)$$

Where, the dependent variable lnX_{ijt} represents the natural logarithm of current dollar fob export values from country i to country j; β_1 is the RTA coefficient, our parameter of interest; β_0 is a constant term, α_t represents the time-fixed effects, α_i represents time-invariant exporter fixed effects, α_j are the importer time-invariant fixed effects and ε_{ijt} is an idiosyncratic error term.

Likewise, $\boldsymbol{S}_{it}$ and $\boldsymbol{M_{jt}}$ are vectors of time-varying monadic controls for exporters and importers respectively composed of h variables: $lnGDP_{it}$, $lnpop_{it}$, $urpart_{it}$, $OECD_{it}$ and $GATT_{it}$, $gspprovider_{it}$, $gspben_{it}$ as well as, $lnGDP_{jt}$, $lnpop_{jt}$, $urpar_{jt}$, $OECD_{jt}$ and $GATT_{jt}$.

Here, ψ and ϕ are vectors of coefficients to be estimated concerning the above control variables, and the subscript h indicates variables.

We define $lnGDP_{it}$ and $lnGDP_{jt}$ as the natural logarithms for current dollar GDPs from countries *i* and *j*; $lnpop_{it}$, $lnpop_{jt}$ are natural logarithms for the population in number of inhabitants of countries *i* and *j*; $urpart_{it}$ and $urpart_{jt}$ stand for the percentage of urban population in country *i* and *j* respectively; this could be seen as a measure of the degree of development of countries, as more developed countries tend to be relatively more urbanized.

Other non-dyadic variables attempt to control for institutional traits related to commerce; these are $gatt_{it}$ and $gatt_{jt}$ that take on 1 if countries *i/j* belong to the GATT/WTO respectively. We use variable $gspben_{it}$ that takes on 1 if country *i* is receiving the generalized system of preferences or any other unilateral preference scheme from country *j*, otherwise 0; $gspprovider_{it}$ takes on 1 if country *i* is granting the generalized system of preferences or any other unilateral preference scheme to country *j*; $oecd_{it}$ and $oecd_{jt}$ take on 1 if the countries *i/j* belong to the Organization of Economic Cooperation and Development OECD.

When no country fixed effects are introduced, controlling for time-invariant monadic variables such as the total surface of a country, the fact of being an island or being landlocked, helps to improve results. Then, vectors $\boldsymbol{S_{it}}$ and $\boldsymbol{M_{it}}$ are augmented with variables $lnarea_{it}$, isl_{it} and $landlocked_{it}$; and $lnarea_{jt}$, isl_{jt} and $landlocked_{jt}$ respectively. Here, $lnarea_{it}$ and $lnarea_{jt}$ are the natural logarithms for the surface in square km of country i and j; *Isl* takes on 1 if country i/j is an island, otherwise 0; and *landlocked* takes on 1 if country i/j is deprived of a direct access to the sea, otherwise 0.

Finally, $\boldsymbol{Z_{ijt}}$ is a vector of dyadic variables that helps to minimize possible bias, composed of g variables: $contg_{ijt}$, $comlang_{ijt}$, $col45_{ijt}$ and $lndist_{ijt}$ and φ is a vector of coefficients to be estimated concerning these dyadic variables; the subscript *g* is to indicate variables, where $lndist_{ijt}$ is the natural logarithm for the weighted distance between countries *i* and *j*; $contig_{ijt}$ takes on 1 if there is a common land frontier between *i* and *j*, otherwise 0; $comlang_{ijt}$ takes on 1 if at least 9% of the pair population share the same language, otherwise 0; $col45_{ijt}$ takes on 1 if both countries were under a colonial relationship before 1945, otherwise 0; and finally our variable of interest rta_{ijt} takes on 1 if both countries share a free trade agreement, otherwise 0.

The equation to be estimated with random effects or with country-pair fixed effects is presented in (3) below. Here we follow (4) assumption.

$$lnX_{ijt} = \beta_0 + \beta_1 RTA_{ijt} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{it}} + \alpha_t + \alpha_{ij} + \mathcal{E}_{ijt} \tag{3}$$

$$Cov(EV_{ijtg}\ , \alpha_{ij}) = 0\,, \qquad t = 1,2,\dots,T\ ;\ ij = 1,2\dots,N\ ;\ g = 1,2\dots,k. \tag{4}$$

Where EV stands for explanatory variables, (*ij*) represents the entities, *t* represents years, and *g* is to enumerate the explanatory variables.

α_{ij} represents country-pair fixed effect. For the traditional fixed effect model (within transformation) (4) assumption is modified to allow for a differential intercept for each country pair *ij*, then, a correlation between at least some of the explanatory variables and the country-pair fixed effects is permitted. See (5). This method does not allow controlling for time-invariant exporter and importer fixed effects at the same time, as the pair-fixed effects are collinear with country fixed effects. Thus, all time-invariant variables are dropped by the within transformation, Greene (2011).

$$Cov(EV_{ijtg}\ ,\ \alpha_{ij}) \neq 0\ ,\ t = 1,2,\ldots,T\ ;\ ij = 1,2\ldots,N\ ;\ g = 1,2\ldots,k. \tag{5}$$

Increasing acceptance to estimate gravity models is acknowledged to the Poisson Pseudo Maximum Likelihood estimator. This technique has been defended by Santos Silva and Tenreyro (2006; 2011) and Fally (2015) as the more reliable method to estimate the gravity equation because it deals with heteroscedasticity problems better than traditional OLS methods. Furthermore, in their work of 2011, they presented further evidence that the PPML estimator generates consistent estimates, even in the presence of a large number of zero values in the data set, a recurrent difficulty in gravity models.

(6) presents the PPML specification when we introduce year fixed effects and exporter and importer time-invariant fixed effects:

$$X_{ijt} = exp(\beta_0 + \beta_1 RTA_{ijt} + \varphi_g \boldsymbol{Z_{ijt}} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{jt}} + \alpha_t + \alpha_i + \alpha_j)\, u_{ijt} \tag{6}$$

Here, X_{ijt} represents the value of the fob merchandise exports from country *i* to country j in current dollars and $u_{ijt} = \exp\left((1-\sigma)\,\varepsilon_{ijt}\right)$. We chose this specification to evaluate trade diversion for a set of interesting RTAs. Thus we introduce a vector of $\mathbf{RTA}_{\boldsymbol{it}}$ trade diversion dummies next to their associated vector of $\mathbf{RTA}_{\boldsymbol{ijt}}$. The subscript *k* stands for the number of RTA dummies included. (6) can now be read as:

$$X_{ijt} = exp(\beta_0 + \beta_k \boldsymbol{RTA_{ijt}} + Y_k \boldsymbol{RTA_{it}} + \varphi_g \boldsymbol{Z_{ijt}} + \psi_h \boldsymbol{S_{it}} + \phi_h \boldsymbol{M_{jt}} + \alpha_t + \alpha_i + \alpha_j)\, u_{ijt} \tag{7}$$

Below in (8) we relax the assumption of the maintenance of unchanging gaps among different intercepts, or stable tendencies, through time. The inclusion of time-varying country fixed effects in the PPML specification leads us to estimate.

$$X_{ijt} = \exp\,(\beta_0 + \beta_1 RTA_{ijt} + \varphi_g \boldsymbol{Z_{ijt}} + \alpha_{it} + \alpha_{jt})u_{ijt} \tag{8}$$

Where α_{it} stands for time varying exporter fixed effects and α_{jt} are the importer time-varying fixed effects. In (9) we include country-pair fixed effects in a specification that not only control for time varying unobserved heterogeneity at the country level but also for time invariant unobserved heterogeneity at the individual level. This is the literature preferred specification.

$$X_{ijt} = \exp\,(\beta_0 + \beta_1 RTA_{ijt} + \varphi_g \boldsymbol{Z_{ijt}} + \alpha_{it} + \alpha_{jt} + \alpha_{ij})u_{ijt} \tag{9}$$

Improvements in Stata procedures documented by Correia *et al*. (2019) and Larch *et al*. (2019) have made possible the estimation of models with larger number of fixed effects with the PPML estimator, such as those presented in equations (8) and (9).

5. Results

In accordance with Baldwin and Taglioni (2006) this paper includes specifications that control for the passing of time using time-fixed effects. This approach allows us to work properly with GDP dollars, avoiding the so-called bronze medal mistake, which occurs when deflating these time series to obtain their real values. Non-averaged bilateral trade data to avoid the silver medal mistake is also used. The inclusion of time-invariant country fixed effects permits the partial offsetting of the endogeneity problem caused by omitted variables, in what is known as the gold medal mistake.

Under the OLS and PPML method, this paper also controls for time-varying country fixed effects for importers and exporters. This procedure would furnish a robust estimate of RTAs that controls for multilateral resistant and other omitted variables that change with the passing of time. The summary of the results will be presented in Table 1 to make comparison easier.

5.1. Traditional methods of estimation

5.1.1. *Pooled ols specifications results*

In its first row, Table 1 presents results based on the pooled OLS specifications. An analysis of the RTA coefficients shows that the model with no fixed effects in column 1 estimates a rise of 39.0%, ($e^{0.329}$ -1) in bilateral exports affected by RTAs relative to flows not influenced by them. It underestimates the impact of RTAs on international bilateral trade with respect to other OLS models that control for fixed effects, excepting for model 8 which simultanusly control for TVFE and country-pair fixed effects. This specficiation indicates a rise of 28.5% in bilateral trade.

When only time-fixed effects are controlled for, model 2 on the pooled OLS specification, the RTA coefficient overreacts, see column 2 of Table 1, producing a rise of 104.6% in bilateral exports affected by a RTA with respect to bilateral trade not affected by RTAs. This is the highest global RTA estimate computed in this paper.

5.1.2. *Random effects and country-pair fixed effects results*

A random effects model assumes that unobserved individual effects are uncorrelated with the explanatory variables Wooldridge (2012). The random effects model moves RTA estimates downward with respect to pooled OLS, yet a positive and significant effect is persistent.

The RTA random effects estimates in model 6 (see column 6 and random effects row) imply a slightly less important increase in trade than the model 2, controlling for time-fixed effects but omitting time-invariant fixed effects. Thus, when time-fixed effects and exporter and importer time-invariant fixed effects are introduced together as in model 6, we obtain an increase of about 31.3% in bilateral exports. To distinguish which model performs better between OLS and random effects we applied Breusch and Pagan (1980) test that checks if random effects are present. Based on their Lagrangian multiplier test for random effects, the OLS pooled model is outperformed by the random effects model.

When we relax the assumption that country-pair individuals' effects are uncorrelated with covariates, we obtain a fixed effect model. This model creates fixed effects for each bilateral export flow that remains invariant through time. Thus, observed and unobserved time-invariant heterogeneity at the country-pair level is kept at bay.

The fixed effects model in column 3 of the Fixed Effects row estimates that bilateral exports sharing a RTA increase by 19.1%, relative to flows without RTAs. Introducing time fixed effects to this model results in an increase of 18.9% in bilateral trade.

Results from the Hausman's specification test, establish that the fixed effect model fits better than the random effects. Particularities at individual level are then correlated with the explanatory variables. Nevertheless, the fixed effect regression at the individual country-pair level generates estimates that could be underestimating the RTA effect on bilateral trade, particularly when time fixed effects are accounted for.

5.2. Current methods of estimation

5.2.1. *PPML specification results*

The PPLM seems to be the more reliable method to estimate the gravity model. Martínez-Zarzoso (2013) validates this method through a series of Monte Carlo experiments. Fally (2015), Montenegro *et al*. (2011) and Martín-Montaner *et al*. (2014) gives additional support to PPML estimations over another techniques.

In the specification without time-fixed effects or country fixed effects, Column 1 in Table 1, the PPML estimate of RTA presents a rise of 10.5%. The introduction of time-fixed effects and country time-invariant fixed effects corrects PPML estimates upwards.

Likewise, PPML estimations shift upward to the introduction of time-varying fixed effects for exporter and importer countries, see column 7. The coefficient is lower in the TIFE and time-fixed specification, column 6, than in column 7, by 0.074 points, equivalent to 10.8 percentage points, which is a relevant difference that deserves attention. The TIFE and time-fixed effect model estimated by PPML produces an increase of 40.5% in bilateral exports affected by a RTA, compared with bilateral export flows that do not profit from any RTA; the comparable result using TVFE estimated by PPML is 51.3%.

We observe a sharp reduction of the RTA estimate when using PPML controlling simultaneously for TVFE and country-pair fixed effects, see column 8. This behaviour is also present under the OLS method, but PPML accentuates the decline. Sharing an RTA will only increase trade by around 4.7% in this specification, which is the preferred approach in literature.

5.2.2. *The Baier and Bergstrand method*

The Baier and Bergstrand technique consists of controlling for multilateral resistance and RTA endogeneity by the means of introducing country-pair fixed effects and time varying fixed-effects on a panel of non-successive years that we call periods. In accordance with Baier and Bergstrand we estimated our model keeping 10 periods, so we retain information for intervals of four years. Baier and Bergstrand (2007) kept data for intervals of five years[1]. In model 8 of the Baier and Bergstrand specification where country-pair and time-varying country fixed effects are computed together, the introduction of a RTA will increase bilateral exports by around 30.0% for OLS and 3.9% for PPML. RTA estimates under the Baier and Bergstrand method, where country fixed effects are accounted for; present higher values compared with those including country-pair fixed effects specifications. See Table 1.

The Baier and Bergstrand method simplifies the analysis of the dynamics of RTA through time. Variable rta_{ijt-1} and rta_{ijt-2} will capture the impact of RTAs on bilateral exports four and eight years before their entry into force or phase-in effects. It also enables the evaluation of anticipatory effects. Thus, rta_{ijt+1} describes the effects of the announcement and pre-entry into force of RTAs. Table 2 presents results for OLS and PPML regressions based on the time varying fixed effects and country-pair fixed effects specification.

Using the Baier and Bergstrand method with OLS, the first lag of the RTAs is positive and significant, but this does not hold for PPML specifications. Introducing a second lag in the specification drops the significance of the first lag. Using OLS, the RTAs effect experience a cumulative increase of 30.5% in bilateral exports during the first four year of entry into force, and approximately a 32.3% cumulative effect during the eight first years. The four years prior to its entry into force, known as the anticipatory effect of RTAs, is non-significant, which suggest strict exogeneity of RTAs, mitigating doubts about reversal causality where an increase in trade could cause RTA appearance. Estimates using PPML show an absence of cumulative and anticipatory effects. See columns (5-8) in Table 2.

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1 The results we show are estimated with data for years 1980, 1984, 1988, 1992, 1996, 2000, 2004, 2008, 2012 and 2016.

5.3. Instrumental variable methods and dynamics

5.3.1. *The Hausman and Taylor instrumental variable estimator*

To deal with RTA endogeneity problem of the type suggested by Baldwin and Jaimovich (2012) where a new free trade agreement between A and B increases the probability that C will sign a RTA with A or B or due to pre-existing overtrading patterns raised by Baier and Bergstrand (2007), we resort to the Hausman and Taylor estimator. These authors have proposed an instrumental variables estimator that uses only the information within the model by taking deviations from group means that can be used as instrumental variables. (Greene, 2011). The correct use of instrumental variable methods requires data on a sufficient number of instruments that are both exogenous and relevant. Swamy *et al.* (2015) argue that such instruments, weak or strong, are often impossible to find.

The Hausman-Taylor estimator assumes that some of the explanatory variables are correlated with the individual-level random effect α_{ij}, while none of the explanatory variables are correlated with the idiosyncratic error ε_{ijt}. In addition to standard assumptions, we also assume that RTA was the only endogenous variable in the model. Through Monte Carlo simulations, this estimator has proved to be robust for endogenous time-varying variables in large sample and perfect knowledge gravity model frameworks. (Mitze, 2010).

H-T estimates of RTA indicate an increase in bilateral trade between 18.2% and 30.2%. We should take these estimates with prudence as the Hausman test applied indicate that we should prefer the fixed effects estimator over the Hausman and Taylor estimator. Nevertheless, being the Hausman and Taylor estimator an instrumental variable estimator we should gain some confidence on making causal claims about the RTA effect.

Other instrumental variable techniques were also considered. The instrumental variable fixed effects and random effects estimators were computed using as instruments for RTA its lags on t-3 and t-4, under the assumption that these variables only influence bilateral exports by the influence they exert on the variable of interest RTA. Results suggest an upward bias for RTA in the IV-Dynamics using the third and fourth lags of the RTA, compared to H-T estimates.

5.3.2. *GMM regression and the Arellano and Bond estimator*

The main purpose of the Arellano and Bond (1991) method is to consistently estimate the dependent variable lags. This technique also allows setting other explanatory variables as endogenous by using GMM-type instruments to compute the causal effects of endogenous covariates. We tried to take advantage of this possibility and intended to use it to correct a possible endogeneity bias on the RTA estimates, nevertheless the Sargan test and the Hansen test failed, suggesting that lagged RTA instruments and lagged bilateral exports instruments were not valid because of overidentification.

TABLE 1
RTA ESTIMATES SUMMARY

Fixed Effects Controls	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(9)
Time invariant Exporter Fixed Effects	NO	NO	NO	NO	YES	YES	NO	NO
Time invariant Importer Fixed Effects	NO	NO	NO	NO	YES	YES	NO	NO
Time Varying Exporter Fixed Effects	NO	NO	NO	NO	NO	NO	YES	YES
Time Varying Importer Fixed Effects	NO	NO	NO	NO	NO	NO	YES	YES
Country-pair Fixed Effects	NO	NO	YES	YES	NO	NO	NO	YES
Time Fixed Effects	NO	YES	NO	YES	NO	YES	NO	NO
Econometric Method								
Pooled OLS	0.329***	0.716***			0.591***	0.631***	0.679***	0.251***
Random Effects	0.164***	0.296***			0.210***	0.272***		
Fixed Effects (within)			0.175***	0.173***				
PPML	0.100***	0.257***	0.067***	0.080***	0.336***	0.340***	0.414***	0.046***
Baier and Bergstrand (OLS)	0.148***	0.362***	0.170***	0.225***	0.607***	0.644***	0.706***	0.262***
Baier and Bergstrand (PPML)	0.086**	0.251***	0.041*	0.066***	0.330***	0.339***	0.410***	0.038**
IV: Hausman and Taylor	0.167***	0.264**			0.183***	0.246***		
IV-Dynamics: RTA lags 3 and 4	0.268***	0.473***	0.304***	0.402***	0.356***	0.451***		
IV-Dynamics: Arellano-Bond (GMM)	-0.247**	0.493***						

Note: *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Robust standard errors.
Source: Elaborated by the authors.

TABLE 2
DYNAMICS BAIER AND BERGSTRAND REGRESSION ON 153 COUNTRIES FOR 10 PERIODS FROM 1980 TO 2016. OLS AND PPML ESTIMATORS

	(1) *lnxijt*	(2) *lnxijt*	(3) *lnxijt*	(4) *lnxijt*	(5) *Xijt*	(6) *Xijt*	(7) *Xijt*	(8) *Xijt*
rtaijt	0.251***	0.061***	0.059***	0.079***	0.046**	0.023	0.020	0.022
rtaijt-1		0.205***	-0.019	0.009		0.019	–0.002	0.002
rtaijt-2			0.240***	0.236***			–0.021	0.021
rtaijt+1				–0.021				–0.009
Constant	15.323***	15.331***	15.340***	15.352***	4.837***	4.850***	4.865***	4.838***
Observations	126,441	115,169	103,085	84,271	103,583	91,237	79,087	63,654
R-squared	0.878	0.885	0.892	0.902	0.957	0.958	0.960	0.959

Note: *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Robust standard errors. All estimations included country-pair and time varying exporter and time varying importer fixed effects. Columns 1-4 for OLS. Columns 5-8 for PPML.
Source: Elaborated by the authors.

Among the techniques and methods applied in this paper, Arellano Bond specifications are the only ones to produce a significant and negative effect for the RTA dummy, although it disappears after dummy year inclusion. We present the Arellano-Bond RTA results in Table 1, assuming the RTA dummy as an exogenous variable and using the dependent variable lags as instruments to correct for the endogeneity of the first lag of the bilateral exports.

The results reported in Table 1 for Arellano-Bond RTA coefficients come from a regression that uses the first lag of the dependent variable as a regressor to account for the dynamics of the model, and all available dependent variable lags as instruments for this. In column 1 we show the result for the equation without year fixed effects, and in column 2 we control with time dummies. For robustness, we verified that the RTA coefficient is systematically negative in the regressions with no year effects as we reduce the number of lagged instruments, and overidentification problems also persist.

These results could be interpreted as evidence of the static approach strength over the dynamic approach, and also as an invitation for further research on dynamics of the dependent variable and RTAs. Some interesting works on dynamics for international trade gravity models can be found in Caporale *et al.* (2012), Didia *et al.* (2015) and Kahouli & Maktouf, (2015).

5.4. RTA estimates summary

Table 1 summarizes RTA estimates results, taking into account the econometric method and the fixed effect mix introduced. The methods used include pooled regression, random effects, fixed effects (within), PPML, Baier and Bergstrand for OLS and for PPML, Hausman and Taylor, Instrumental variables and dynamics models. Thus, in static models, the RTA coefficient is always positive and significant. Depending on the method and specification employed, statics models coefficients can vary from an estimate 0.038 in the Baier and Bergstrand-PPML method with TVFE and country-pair fixed effects, to levels as high as 0.716 that comes from the OLS pooled regression using only time fixed effects.

5.5. RTAs effects at the disaggregated level

The effect of particular RTAs computed by means of dummy variables for each scheme such as EU, NAFTA or MERCOSUR has been reviewed in Magee (2008), Eicher and Henn (2011) and Kohl (2014), Baier and Bergstrand (2019) among others.

Most of the preceding studies on the effects of particular RTAs are estimated by OLS techniques. This paper offers 123 RTA estimates based on PPML over a database across 153 countries and observations from 1980 to 2018, see Table 3. We apply a robustness check to our individual RTAs estimates by presenting PPML time-invariant fixed effects and next to them time-varying fixed effects estimates. On the time-invariant country fixed effects specification, we control for RTA membership other than the RTA of interest, distance between countries

TABLE 3
PPML ESTIMATES FOR A GROUP OF 123 REGIONAL TRADE AGREEMENTS FROM A 153 COUNTRIES 1980-2018 DATA SET

Agreement	Year	(1) TIFE	(2) TVFE	Agreement	Year	(1) TIFE	(2) TVFE
		RTA coef.	RTA coef.			RTA coef.	RTA coef.
ASEAN free trade area	1992	0.165***	–0.211***	EC–Jordan	2002	–0.279***	–0.237***
ASEAN-Australia	2010	0.352***	0.028	EC–Lebanon	2003	0.366***	–0.249***
ASEAN-India	2010	0.242***	0.009	EC–Mexico	2000	–0.161***	–0.204***
ASEAN-Korea	2010	0.799***	0.156***	EC–Moldova	2015	0.322**	0.229***
ASEAN-New Zealand	2010	0.402***	0.348***	EC–Morocco	2000	0.763***	0.149***
Australia-Japan	2015	0.652***	0.218***	ECOWAS	1993	1.041***	0.697***
Australia-Korea	2015	0.758***	0.111	EC–Peru	2012	0.224***	–0.044
Australia-New Zealand	1983	1.269***	0.486***	EC–South Africa	2000	0.485***	–0.446***
Australia-Singapore	2003	0.274***	0.135**	EC–Syria	1977	0.515***	–1.139***
Australia-Thailand	2005	0.777*	0.483***	EC–Tunisia	1998	0.913***	–0.003
CAN (Andean Community)	1988	0.928***	0.956***	EC–Turkey	1996	0.462***	0.106***
Canada Colombia	2012	–0.524***	–0.104	Ecuador–EC	2017	–0.006	0.139**
Canada-EC	2018	–0.236**	0.163***	EFTA–Israel	1993	0.354***	–0.599***
Canada-EFTA	2009	0.226**	–0.261***	EFTA–Korea	2006	0.437***	–0.009
Canada-Jordan	2013	–0.805***	0.563***	EFTA–Peru	2012	1.574***	–0.016
Canada-Peru	2009	0.857***	0.487***	GCC	2003	–0.812***	0.032
CEFTA	2007	0.389***	0.181***	Group of Three	1995	0.504***	0.895***
Chile Colombia1	1994	0.905***	–0.025	India–Japan	2011	–0.685***	–0.140***
Chile Colombia2	2009	0.924***	0.106	India–Malaysia	2011	0.436***	–0.029
Chile-Australia	2009	–0.852***	0.524***	India–Singapore	2005	0.300***	–0.053
Chile-China	2006	1.494***	0.606***	India–Sri Lanka	2001	1.251***	0.532***
Chile-EC	2003	0.201***	–0.247***	Japan–ASEAN	2008	0.581***	0.108***
Chile-India	2008	0.669***	–0.021	Japan–Indonesia	2008	0.648***	–0.111***
Chile-Japan	2007	0.867***	0.274***	Japan–Malaysia	2006	0.661***	0.121***
Chile-Korea	2004	1.579***	0.246***	Japan–Mexico	2005	–0.099	0.034
Chile-Malaysia	2012	–0.919***	0.035	Japan–Mongolia	2016	–0.355	0.736***
Chile-Peru1	1999	0.819***	0.250***	Japan–Peru	2012	0.318**	0.056
Chile-Peru2	2009	0.657***	–0.292***	Japan–Philippines	2008	0.507***	0.294***
Chile-Thailand	2015	0.120	0.333***	Japan–Singapore	2002	0.250***	0.084**
Chile-Turkey	2011	–0.297***	0.209**	Japan–Switzerland	2009	0.654***	0.100
Chile-Vietnam	2014	0.719***	0.431***	Japan–Thailand	2007	0.885***	0.237***
China-ASEAN	2005	–0.160***	0.146***	Japan–Vietnam	2009	0.636***	–0.050
China-Costa Rica	2012	0.075	0.229	Korea Republic–Canada	2015	–0.033	–0.039
China-New Zealand	2008	0.188*	0.506***	Korea Republic–India	2010	0.058	–0.023
China-Pakistan	2007	–0.096	0.181***	Korea Republic–New Zealand	2016	0.156***	0.030
China-Peru	2010	1.351***	0.339***	Korea Rep.–Singapore	2006	0.649***	0.300***
China-Singapore	2009	–0.256***	–0.025	Korea Republic–Turkey	2013	0.424***	0.301***
CIS	1994	1.561***	0.020	Korea–Peru	2012	1.292***	0.348***
COL (CAN) MERCOSUR	2005	–0.006	0.247***	Mauricio–Turkey	2013	0.865***	0.833***
Colombia Northern Triangle	2009	0.514***	0.489***	MERCOSUR	1991	1.139***	0.592***
Colombia-Costa Rica	2017	–0.257	–0.220*	MERCOSUR–India	2009	0.325***	0.206**
Colombia-EC	2013	0.185**	–0.0317	Mercosur–Peru	2006	0.024	–0.105**
Colombia-EFTA	2011	0.345**	–0.396***	NAFTA	1994	0.871***	0.439***
Colombia-Korea	2017	0.369***	0.036	PAFTA	1998	–0.699***	0.547***
COMESA	1994	1.273***	1.067***	SAFTA	2006	0.284**	–0.041
EAEC	1997	1.155***	0.417***	Southern African Develop. Comm.	2000	1.992***	0.229***
EC Enlargement (10)	1981	0.239***	0.018	Turkey–EFTA	1992	0.031	–0.268***
EC Enlargement (12)	1986	0.259***	0.106***	Ukraine–Belarus	2006	1.658***	0.519***
EC Enlargement (15)	1995	0.309***	0.083***	Ukraine–Kazakhstan	1998	1.864***	0.188
EC Enlargement (25)	2004	0.277***	0.086***	Ukraine–Turkmenistan	1995	3.266***	0.366
EC Enlargement (27)	2007	0.496***	0.143***	US–Australia	2005	–0.781***	–0.237***
EC Enlargement (28)	2013	0.534***	0.076***	US–Bahrain	2006	–0.163	0.089
EC-Albania	2006	0.976***	–0.031	US–CAFTA–DR	2006	0.477***	0.132***
EC-Algeria	2005	0.262***	–0.227***	US–Chile	2004	–0.228**	0.324***
EC-Cameroon	2009	0.433***	–0.464***	US–Colombia	2012	0.153**	0.026
EC-Caricom	2008	–0.623***	–0.122***	US–Israel	1985	1.093***	0.253***
EC-Côte d'Ivoire	2009	0.275***	–0.283***	US–Jordan	2001	0.339***	0.778***
EC-Croatia	2002	0.803***	–0.130***	US–Morocco	2006	–0.716***	0.367***
EC-EFTA	1973	0.220***	–0.082***	US–Oman	2009	–0.703***	0.302***
EC-Egypt	2004	0.185***	–0.342***	US–Peru	2009	–0.147	0.096**
EC-Ghana	2017	–0.416	–0.472***	US–Singapore	2004	–0.021	–0.273***
EC-Israel	2000	0.042	–0.283***				

Source: Elaborated by the author. *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors.

Note: Columns (1) are estimated with time-invariant fixed effects and time fixed effects. Columns (2) include time-varying fixed effects and country-pair fixed effects.

i and *j*, common land frontier between *i* and *j*, if the country-pair shares the same language, and if both countries were under a colonial relationship before 1945. Profiting from recent PPML computing power improvements, Correia *et al* (2019), we respectively estimate PPML country-pair and time-varying fixed effects for individual RTAs. This is one of the major contributions of this paper.

As can be seen in Table 4, column 1, most of the RTA estimates, 94 out of the 123, equivalent to 76.4% of the sample, show a positive sign. Column 2, gives a less optimistic view, presenting 80 positive estimates. In addition, 62 RTA estimates bear out positive signs in both specifications, and 47, equivalent to 38.2% are positive and significant in both PPML specifications. These results, point to a larger proportion of trade agreements that are successful in promoting trade than in Kohl (2014), who reported that only 44 out of 166 RTAs, equivalent to 26.5% of their sample presented a positive and significant effect. Another interesting comparison is Baier *et al.* (2019) who used a sample of 65 RTAs and found positive statically significant effects for the majority of the agreements, 54%. Nevertheless, their results are not strictly comparable to ours, as they account for the effect of lagged RTAs.

The median RTA on this sample increases trade by 42.2%, ($e^{0.352} - 1$). Despite the dispersion, around 75.6% of RTA's estimates fall within one standard deviation of this median effect and 93.5% within two standard deviations.

Some straightforward outliers are the Chile-Malaysia, Gulf Council Countries GCC, PAFTA and EC-Caricom agreements, which seem to be highly counterproductive to trade creation, while the largest positive effects are posted by Ukraine-Turkmenistan, SADC, the Chile-Korea, EFTA-Peru and the Ukraine-Kazakhstan agreement. The latest impressive results of these cases concern former Soviet Union countries and could be attributed to some kind of transition effect or measurement error that could bias their estimates upward. Chile-Malaysia and GCC, Ukraine-Turkmenistan, EFTA-Peru and Ukraine-Kazakhstan become non-significant under the TVFE and country-pair specification. As we can observe in Table 4, around 23% of the RTA lose significance when shifting from TIFE to TVFE. The number of agreements significant in both specifications is 77, equivalent to 63%. One intriguing result is that only 50.4% of RTAs are positive and significant under the country-pair and TVFE specification. That number is substantially higher under the TIFE specification reaching 71.5% of RTAs.

Considering results in both specifications, United States agreements present mixed results, showing trade creation with Israel, Jordan and Colombia while the agreement with Bahrein is non-significant. Counterproductive effects appear with Australia. Similarly, European Union agreements outside its zone tend to produce mixed results. Particularly successful seem to be the agreements with Albania, Turkey and Moldova. Agreements with CARICOM, Jordan and Mexico present significant negative effects in both specifications. On the other side of the Pacific Ocean, 64% of the RTAs signed by Japan show a positive sign, and only its agreement with India produces a negative impact. China's RTAs tend to promote trade. Robust results are present in its agreements with Chile and Peru.

TABLE 4
DESCRIPTIVE STATISTICS FOR 123 RTA COEFFICIENTS ESTIMATED BY PPML ON A GRAVITY MODEL

	Specifications	
	TIFE	TVFE
Total number of RTAs	123	123
RTA coefficients (average)	0.407	0.119
Number of RTAs presenting a positive effect	94	80
% of RTAs presenting a positive effect	76.4%	65.0%
Average of positive RTA coefficients	0.653	0.29
Number of RTAs presenting a negative effect	29	43
% of RTAs presenting a negative effect	23.6%	35.0%
Average of negative RTA coefficients	0.391	0.2
Number of (+ or -) significant RTA coefficients (below the 0.10 level of significance)	106	89
% of (+ or -) significant RTA coefficients	86.2%	72.4%
Average coefficient for significant RTAs	0.484	0.155
Number of positive (+) and significant RTA coefficients	88	62
% of positive and significant RTA coefficients	71.5%	50.4%
Average coefficient for positive and significant RTAs	0.694	0.352
Number of negative (-) and significant RTA coefficients	18	27
% of negative and significant RTA coefficients	14.6%	21.9%
Average coefficient for negative and significant RTAs	–0.541	–0.298
	Total	%
Number of RTAs losing significance by shifting from TIFE to TVFE	29	23.5%
Number of non-significant RTAs gaining significance by shifting from TIFE to TVFE	12	9.8%
Number of significant (+ or -) RTAs on both specifications (TIFE and TVFE)	77	62.6%
Number of non-significant (+ or -) RTAs on both specifications (TIFE and TVFE)	5	4.1%
Number of positive (+) and significant RTAs on both specifications (TIFE and TVFE)	47	38.2%
Number of negative (-) and significant RTAs on both specifications (TIFE and TVFE)	5	4.1%

Source: Elaborated by the authors. Note: (TIFE) stands for time-invariant fixed effects. (TVFE) stands for time-varying fixed effects.

A final caveat: RTA coefficients at the disaggregated level should be read with caution. The scope and depth of the agreements change considerably from one RTA to the other. In theory it could be expected that deeper agreements produce higher increases in cross-border flows than those which are shallow. Equally important is the enforceability of these arrangements, especially in the case of politically unstable developing countries.

5.6. RTAs trade creation or trade diversion

Following Ghosh and Yamarik, (2004) and Eicher, Henn and Papageorgiou (2012) we use two sets of dummy variables to pick up RTA trade creation and trade diversion effects. Trade diversion occurs if a trade block creates trade in detriment of more productive third countries excluded from the agreement. The first, RTA_{ijt}, in (7) implies that both trading partners are members of the same RTA, the second, $DivRTA_{it}$ indicates that one country, whereas exporter or importer is a member of the RTA we are estimating.

Ghosh and Yamarik, (2004) define $DivRTA_{it}$ as a vector of variables which measures current membership of either country i or j in a RTA and thus, captures the external effects of the RTA on trade with countries outside the zone. The coefficient Υ_k for $DivRTA_{it}$ is interpreted as a measure of lower or higher than normal trade between nations in the trading bloc, and a country outside the bloc relative to a random pair of countries.

Hence, a negative sign for Υ_k indicates less trade with non-members and is interpreted as evidence of trade diversion.

In this section, we select a group of 25 interesting RTAs to evaluate whether trade diversion is actually mitigating the impact of RTAs on trade. As in Magee (2008) our estimates point to trade creation effects for ASEAN, MERCOSUR and NAFTA. The following analysis will be based on results for the TIFE specification because some trade diversion effects could not be estimated under the TVFE specification, due possibly to collinearity problems as too many fixed effects were dropped to perform estimations.

Thus, a third of the agreements trade creation effects are mitigated by trade diversion effects. In 6 cases the intra-block trade creation effect is sufficiently strong to resist trade diversion as in Australia-Korea, Colombia-Northern Triangle, the Group of 3, ECOWAS, EC-Turkey and NAFTA. For half of the sample of analysed RTAs, the extra-block effect reinforces the intra-block trade creation effects. Conversely, the trade diversion effect outstrips the trade creation intra-block effect in ASEAN-Japan and adds to intra-block negative effects in Canada-Colombia, Chile-EC, EC-Israel and Peru-United States. See Table 5.

6. Conclusion

This paper examines the effect of regional trade agreements RTAs on international bilateral trade flows. Based on the gravity model, we perform a sensitivity analysis to the effect of the RTA dummy, applying a wide range of econometric methods and model specifications. Our database consists of an unbalanced panel for 153 countries, including observations from 1980 to 2018. Particular attention is given to Poisson Pseudo Maximum Likelihood

TABLE 5

TRADE CREATION AND TRADE DIVERSION: PPML ESTIMATES FOR A GROUP RTAS FROM A 153 COUNTRIES 1980-2018

Agreement	1 Intra-block effect *RTAijt*	2 Extra-block effect *DivRTAit*	3 Net effect	4 Intra-block effect *RTAijt*	5 Extra-block effect *DivRTAit*	6 Net effect
Australia - Korea Republic	0.436***	–0.029	0.407	0.114	0.028	0.14
Andean Community	0.809***	0.019	0,82	0.709***	.	
Association of Southeast Asian Nations ASEAN	0.200***	0.137***	0,36	–0.192***	.	
ASEAN - Japan	0.091**	–0.346***	–0,26	0.262***	0.105***	0.37
Canada - Colombia	–0.463***	–0.154***	–0,60	–0.122	.	
Canada - European Union	0.591***	0.092**	0.683	0.277***	0.038	0.32
Caribbean Community CARICOM	3.342***	0.108	3,45	0.828*	.	
Chile - Colombia	0.902***	0.203***	1,17	–0.195*	.	
Chile - European Union	–0.319***	–0.158***	–0,44	–0.373***	–0.051***	–0.42
Chile - Vietnam	0.534***	0.179***	0.355	0.031	–0.245***	–0.21
Common Market for Eastern and Southern Africa	1.234***	–0.019	1,21	–0.245	–0.671***	–0.92
Colombia - EFTA	0.153	0.097***	0,24	0.084	0.275***	0.36
Colombia-(Guatemala-El Salvador-Honduras)	0.433***	–0.008	0.813	0.437***	0.022	0.46
Colombia-Mexico-Venezuela (The Group of 3)	0.655*	–0.013	0,41	0.560***	.	
Colombia - United Estates	0.294***	0.05	0,32	–0.031	.	
Economic Community of West African States	0.963***	–0.093**	0,88	0.662***	.	
European Free Trade Association EFTA	0.665***	0.196***	0,86	–0.009	.	
European Union EU(27)	0.591***	0.084***	0,67	0.414***	0.055	0.47
European Union - Israel	–0.127	–0.055	–0,20	–1.288***	–0.511***	–1.80
European Union - South Africa	0.786***	0.140***	0,93	0.611**	0.516***	1.13
European Union - Tunisia	0.939***	0.012	0,94	0.051	0.026	0.08
European Union - Turkey	0.490***	–0.002	0,51	0.166***	0.002	0.17
Southern Common Market MERCOSUR	1.311***	0.044	1,36	0.556***	.	
North American Free Trade Agreement NAFTA	0.463***	–0.220***	0,25	0.405***	.	
Peru - United States	–0.328***	–0.106***	–0,43	0.068	.	
Observations	674133	674133		457250	457250	
R-squared	0.899	0.899		0.958	0.958	

*** $p<0.01$, ** $p<0.05$, * $p<0.1$. Robust standard errors. Source: Elaborated by the author.
Note: Columns (1) and (2) include time-invariant country fixed effects and time fixed effects. Columns (4) and (5) stand for time-varying fixed effects and country-pair fixed effects

(PPML), which is the method that, when applied with fixed effects, best seems to contend with heteroscedasticity problems and bias from a high proportion of trade flows registered as zero.

A strong positive impact for RTA is consistently found on most specifications. Once multilateral resistance and other unobserved variable bias are controlled by the introduction of time-varying country fixed effects in a PPML regression, we find that RTAs increase bilateral trade flows by 51.3%, with respect to those trade flows with no agreements. When country-pair fixed effects are added to the previous specification, the RTA effect is reduced to 4.7%, still economically significant, as it confirms that efforts to close international trade deals are fruitful.

RTA cumulative effects are found using OLS, but its significance disappears with the PPML method. Instrumental variable methods were also tested. Using the third and the fourth lags of RTAs as instruments for RTA, as well as, employing the Hausman and Taylor estimator that introduces instrumental variables to deal with endogeneity, gives sizable and significant results.

We found considerable variations in the estimates of RTAs at the disaggregated level. While most of these successfully increase trade, others seem to destroy it, or are non-significant. When only time-invariant fixed effects and time fixed effects were included, 71.5% of RTA were positive and significant; this number slid to 50.4% for the time-varying and country-pair fixed effects specification. Robustness checks for individual RTAs based on the comparison of the PPML time-invariant fixed effects specification and the time-varying and country-pair fixed effects specification show that 38.2% of RTAs are positive and significant in both specifications.

The wide range of individual RTA estimates [-1.139; 3.266] could be explained by the fact that RTAs are heterogeneous in scope and depth. Another hypothesis points to lack of enforceability, meaning that a number of RTAs are not completely implemented in practice and remain only as a written statement, a line of research worth exploring.

Trade diversion effects were computed for a sample of RTAs. At large, trade creation effects tend to be stronger than trade diversion effects or even be reinforced by an open trade block expansion effect. Nevertheless, the potentiality of RTA to improve well-being must not be given for granted, as in certain cases trade diversion is found to outstrip trade creation effects.

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APPENDIX

List of countries included in the gravity model database

Albania, Algeria, Angola, Argentina, Australia, Austria, Azerbaijan, Bahrain, Bangladesh, Barbados, Belarus, Belgium, Belize, Benin, Bermuda, Bolivia, Brazil, Brunei, Bulgaria, Burkina Faso, Burundi, Cambodia, Cameroon, Canada, Cape Verde, Central, African Republic, Chad, Chile, China, Colombia, Congo Democratic, Congo Republic, Costa Rica, Ivory Coast, Croatia, Cuba, Cyprus, Czech Republic, Denmark, Djibouti, Dominican Republic, Ecuador, Egypt, El Salvador, Equatorial, Guinea, Estonia, Ethiopia, Fiji, Finland, France, Gabon, Gambia, Georgia, Germany, Ghana, Greece, Grenada, Guatemala, Guinea, Guinea-Bissau, Guyana, Haiti, Honduras, Hong Kong, Hungary, Iceland, India, Indonesia, Iran, Iraq, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kazakhstan, Kenya, South Korea, Kuwait, Kyrgyzstan, Latvia, Lebanon, Liberia, Libya, Lithuania, Luxembourg, Madagascar, Malawi, Malaysia, Mali, Malta, Mauritania, Mauritius, Mexico, Moldova, Mongolia, Morocco, Mozambique, Nepal, Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Norway, Oman, Pakistan, Panama, Papua New Guinea, Paraguay, Peru, Philippines, Poland, Portugal, Qatar, Romania, Russia, Rwanda, Samoa, Saudi Arabia, Senegal, Sierra Leone, Singapore, Slovakia, Slovenia, South Africa, Spain, Sri Lanka, Sweden, Switzerland, Syria, Tajikistan, Tanzania, Thailand, Togo, Tonga, Trinidad and Tobago, Tunisia, Turkey, Turkmenistan, Uganda, Ukraine, United Arab Emirates, United Kingdom, United States, Uruguay, Uzbekistan, Venezuela, Vietnam, Yemen, Zambia, Zimbabwe.

Spillover effects of economic complexity on the per capita GDP growth rates of Mexican states, 1993-2013*

Efectos derrame de la complejidad económica en las tasas de crecimiento del PIB per cápita de los estados Mexicanos, 1993-2013

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Abstract

The opening up of the Mexican economy completely transformed the growth dynamics of the per capita Gross Domestic Product (GDP) of the country's various states, with a clear tendency towards growth being concentrated in specific regions. In this study, we quantify the indirect or spillover effect of economic complexity on growth based on the following two facts: i) economic complexity is an important factor in explaining GDP growth rates, and ii) there is a clear regional pattern in the states' economic complexity, i.e., the economic complexity variable shows a positive spatial autocorrelation. Our results provide two insights: first, that the estimated positive spillover effect of complexity on growth is not negligible, particularly for states in the north of the country, whose own economic complexity is as important as that of their neighbors. In contrast, the spillover effect in southern states is negative. Being located next to states with low levels of economic complexity has a significant negative externality that almost overrides the positive effect of a state's own level of complexity. Our findings lead us to conclude that spillover effects may

* The views and conclusions contained in this article are those of the authors and do not necessarily reflect the point of view of Banco de México. The comments and remarks by three anonymous referees are greatly acknowledged. They have significantly enhanced this research work.

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have played a more important role in explaining the diverse pattern of growth between northern and southern Mexico than previously thought.

Key words: *Economic complexity, spillover effects, spatial econometrics.*

JEL Classification: *O10, O14, O47.*

Resumen

La apertura de la economía mexicana transformó por completo la dinámica de crecimiento del Producto Interno Bruto (PIB) per cápita de los diversos estados del país, con una clara tendencia a concentrar el crecimiento en regiones específicas. En este estudio, cuantificamos el efecto derrame o indirecto de la complejidad económica sobre el crecimiento con base en los siguientes dos hechos: i) la complejidad económica es un factor importante para explicar las tasas de crecimiento del PIB, y ii) hay un patrón regional claro en la complejidad económica de los estados, i.e., la variable complejidad económica muestra una autocorrelación espacial positiva. Nuestros resultados muestran: i) que el efecto derrame o indirecto estimado de la complejidad en el crecimiento es positivo y no insignificante, particularmente para los estados del norte del país, cuya propia complejidad económica es tan importante como la de sus vecinos. Por el contrario, el efecto indirecto en los estados del sur es negativo. Estar ubicado al lado de estados con bajos niveles de complejidad económica tiene una externalidad negativa significativa que casi anula el efecto positivo del propio nivel de complejidad de un estado. Nuestros hallazgos nos llevan a concluir que los efectos indirectos pueden haber jugado un papel más importante para explicar el patrón diverso de crecimiento entre el norte y el sur de México de lo que se pensaba anteriormente.

Palabras clave: *Complejidad económica, efectos contagio, econometría espacial.*

Clasificación JEL: *O10, O14, O47.*

1. Introduction

The structure of the Mexican economy has undergone a significant transformation since the economic liberalization period, with the impact on the various states being remarkably heterogeneous.[1] This fact has inspired a growing

[1] The opening-up period is generally considered to have begun in the mid-1980s with Mexico's joining the General Agreement on Tariffs and Trade (GATT). However, most studies analyze the effects after its signing of the North America Free Trade Agreement

literature, which attempts to document the changes brought about by the economic reforms, including: i) studies that analyze the changes in the localization of specific industries or the specialization of specific states as a result of the trade reforms; ii) studies that seek to determine the key factors in explaining the diverse economic growth performance of Mexican states in this period, and; iii) studies that document the increase in the concentration of economic activity, primarily manufacturing, just as traditional trade models, new trade theories or new economic geography models predict.[2] Therefore, studies that endeavor to establish the causes of growth during this period should take into account the agglomeration of economic activity, since spatiality represents an important component of the regional growth process in Mexico following the reforms.

Hidalgo and Hausmann (2009) (henceforth HH) propose a measure of the amount of *productive knowledge* that economies have, which they call *economic complexity*. Traditional approaches to performing this task seek to gauge the latter by taking into account all of the productive elements (inputs) that economies possess, e.g., abundance of resources, human and physical capital, infrastructure (communications, transportation, etc.), technology, quality of institutions, to mention just a few. In contrast, HH's method looks at the products that are already being produced by economies or the economic activities they already undertake.[3] They show that their measure of productive knowledge can account for the per capita GDP differences among countries and, furthermore, that it can be used to predict their future growth rates.[4] They do this by estimating growth regressions, using the growth rate as the dependent variable explained by the economic complexity. They argue that economic complexity alone is much more predictive than other development indicators combined, such as, aggregate measures of human capital, various measures of physical capital, and measures of social capital and of the health of their institutions (institutional quality, measures of enforcement of the rule of law, etc.).

(NAFTA) in 1994, as this is regarded as a more influential event and, more importantly, there is little or no reliable data for the 1980s or before.

2 Theoretical models predict that trade will lead to greater concentration and only differ in regard to the explanation of the factors that cause it. For example, Ricardo's model predicts that trade will lead to regional specialization and to a higher level of industrial localization due to the productivity differences among economies, whereas in the Heckscher-Ohlin model, economies specialize in economic activities that are intensive in those factors of production in which they are relatively abundant. Models from the literature known as "new economic geography" explain that trade costs, increasing returns to scale, input-output linkages (among companies in the same or different industrial sectors), and so on can lead to increased agglomeration of economic activity [see Krugman (1991), Krugman and Venables (1995, 1996), among others].

3 We believe this makes perfect sense, since it is easier to measure the goods being produced by an economy than the inputs needed to produce them, e.g., the quality of institutions, etc.

4 Hartmann *et al.* (2017) state that: "These measures of economic complexity have received wide attention because they are highly predictive of future economic growth."

Chávez *et al.* (2017) apply the ideas of HH to the Mexican case and use information on the productive structure of each of the country's 32 states to calculate a measure of economic complexity, then show that this variable goes a long way to explaining the different growth patterns of Mexican states during the period 1998-2013. However, they do not consider the spatial dimension of economic complexity. As theoretical models predict and various empirical studies illustrate, increased trade tends to lead to a concentration of economic activity; consequently, the authors may underestimate economic complexity as a predictor of growth rates, since they ignore the spillover effects. In this study, we find empirical evidence to affirm that the growth rates of the states during this period depend not only on their own economic complexity measure but also on that of their neighboring states.

The present study follows on from the work of Chávez *et al.* (2017) and expands upon it in various ways: i) we extend the sample period by adding data from the 1993 economic census to the analysis of the 1998, 2003, 2008, and 2013 census data that they employed, thus covering more of the post-liberalization period; ii) we provide evidence to affirm that the automotive industry also helps to explain the different growth rates of the states during this period, with states specializing in the economic activities associated with the latter experiencing growth rates that were above the national average, and, more importantly: iii) we confirm that economic complexity is an important factor in explaining the observed growth rates of Mexican states in the post-liberalization period. Indeed, the level of economic complexity has a direct impact, with more complex states growing faster than their less complex counterparts; furthermore, we document an indirect or spillover effect that can be generated in different ways (the existence of technology dissemination, agglomeration effects, economies of scale, network effects, etc.). States whose neighbors have more complex economies tend to grow faster than those with less complex neighbors. Moreover, this spillover effect is not homogeneous among the states: northern states (the most complex) have a positive influence on their neighbors' growth rates, whereas southern states (the least complex) have a negative impact. Compared to the direct effect, the magnitude of the estimated indirect effect is not negligible.

Panel data studies looking to measure the relationship between growth and its determinants –using growth regressions à la Barro– find it very straightforward to investigate if those determinants have both direct and indirect (spillover) effects on economic growth. As defined by Halleck-Vega and Elhorst (2017), a direct effect measures the marginal impact of a change in one explanatory variable in a particular cross-sectional unit on the dependent variable of that unit itself. Meanwhile, an indirect (or spillover) effect is defined as the marginal impact of a change in the explanatory variable in a particular unit i on the dependent variable values in another unit j ($\neq i$). Spatial econometrics literature includes a range of models to estimate different types of interaction effects among units: i) endogenous interaction effects among the dependent variables, (ii) exogenous interaction effects among the explanatory variables, and (iii) interaction effects among the error terms. The General Nesting Spatial (GNS) model is the most general specification, containing all three of the types of interactions previously mentioned, while the Spatial Autoregressive Combined (SAC), Spatial Durbin (SDM), and Spatial Durbin Error (SDEM) models contain only two. The Spatial Autoregressive (SAR), Spatial Lag of X (SLX), and Spatial Error (SEM) models contain only one of the three interactions.[5]

Our study employs the simplest model, the SLX, to estimate the spillover effect of economic complexity on growth. As SLX only considers exogenous interaction among the explanatory variables, it can be estimated using OLS. Therefore, our growth regressions incorporate economic complexity as independent variable in two distinct ways: i) as the specific economic complexity of each state to estimate the direct effect of that particular variable, and; ii) as the average economic complexity of the neighbors of each state to estimate the indirect effect of complexity, *i.e.*, to estimate the effect that the complexity of a state's neighbors has on its own growth.[6]

The remainder of the article is organized as follows. In Section 2, we present a brief review of studies that document the main changes in the Mexican economy in the post-liberalization period. In Section 3, we present the data to be used in the empirical analysis and explain the method for calculating the measure of economic complexity that we will use to explain the states' growth rates. Appendices 1 and 2 show the computed values of the complexity variable for all the economic censuses considered, along with the evidence for the need to include a spatial dimension when attempting to explain per capita GDP growth rates based on complexity. In Section 4, we present and discuss the main results. Section 5 presents the final remarks.

2. Related Studies

The change in Mexico's development strategy –from import substitution to economic liberalization and trade promotion– resulted in a significant change in the growth performance of its individual states. Esquivel (1999) finds evidence in favor of the per capita output convergence hypothesis for Mexican states during the period 1940-1995, *i.e.*, that poor states tended to grow faster than rich states during this period. In general, rich states tend to be located in the north of the country, with the notable exception of Mexico City, while poor states tend to be located in the south.[7] This would imply that the gap between rich and poor states decreased during this period. In line with these findings, Chiquiar (2005) uses

5 Excellent references for spatial econometrics include Elhorst (2013), LeSage and Pace (2009), LeSage (2014), Halleck-Vega and Elhorst (2015), and Elhorst and Halleck-Vega (2017)

6 To estimate the spatially lagged level of complexity, we employ the simplest contiguity matrix: the queen matrix.

7 An analysis of subperiods reveals a clear pattern in the rates at which states converge. The convergence rate from 1940 to 1960 is higher than that for the period 1960-1980, while

a similar methodology to that of Esquivel (growth regressions), though finds that the trade reforms led to a divergent pattern in the per capita output levels of states during the period 1985-2001. Other studies also affirm that the gap between rich and poor states has been widening since the mid-1980s.[8]

What can explain these changes in the states' growth rates? Hanson (1998) describes how there was an important reallocation of manufacturing industry within the country after the enactment of NAFTA, from the country's center (Mexico City and Mexico State) to states in the north, mainly those sharing a border with the U.S. (Baja California, Chihuahua, Coahuila, Nuevo León, Tamaulipas, and Sonora).[9] He argues that this reallocation of industry sought, in part, to reduce transportation costs to what would become the most important market after the signing of the agreement: the U.S. Mosqueda *et al.* (2017) state that the sectors that contributed most to the increase in manufacturing concentration in the first ten years of NAFTA were: transportation equipment, chemicals, food products, and primary metal industries. In 1993, these four manufacturing subsectors accounted for 32% of the concentration of all manufacturing production; ten years later, the figure was 52%. Chiquiar (2005) reports that states more favorably endowed in terms of human and physical capital and better levels of transport and communications infrastructure (*i.e.*, states in the north) have grown faster since the signing of NAFTA. Rodríguez-Oreggia (2005) also finds that human capital plays a decisive role in explaining the difference in growth rates, as well as evidence to affirm that public investment causes greater growth. Jordaan and Rodríguez-Oreggia (2012) argue that Foreign Direct Investment (FDI) and agglomeration have acted as important drivers of state growth since the trade reforms. Moreover, they affirm that there is a spatial dimension to the structural change in the Mexican economy, since many economic activities have agglomerated in the states that share a border with the U.S., fostered by FDI, which also tends to localize in certain economic activities. Cabral and Varella-Mollick (2012) document that trade, FDI, and international migration contributed significantly to the growth of the output per capita of Mexican states during the period 1993-2006. The role of migration in explaining growth rates is more important for states located on the northern border, in the center, and in the northern-central region. Cabral, Varella-Mollick, and Saucedo (2016) study the effect of violence on the evolution of the productivity (GDP per worker) of

both are greater than that for 1940-1995 and 1960-1995. For the period 1980-1995, the rate is estimated to be statistically not different from zero.

8 See Aguayo-Téllez (2006), Gómez-Zaldívar and Ventosa-Santaulària (2010, 2012), Rodríguez-Oreggia (2005), and Rodríguez-Pose and Sánchez-Reaza (2002), among others.

9 Mosqueda *et al.* (2017) affirm that during the first ten years of NAFTA: the contribution of Mexico City and Mexico State to domestic manufacturing value added decreased from 37.3 to 18.3 percent; that of the six states along the northern border rose from 23.8 percent to 33.4 percent, and that of Aguascalientes, Durango, Guanajuato, Querétaro, San Luis Potosí, and Zacatecas (states in the North-Center of the country) rose from 8.7 percent to 14.8.

Mexican states during the period 2003-2013 and find that crime has negative and statistically significant effects on labor productivity, particularly across those categories of crime prosecuted by local authorities.

Using municipal-level data, Garduño (2014) shows that output per worker grew faster in regions located closer to the U.S.-Mexico border and slower in regions located further away from it. According to him, the trade agreement increased inequality and the localization of economic activity. Finally, Chávez *et al.* (2016) find empirical evidence of a positive relationship between the average GDP growth rate of Mexican states and a measure of efficiency of the judicial system in the states; in particular –for the period 2006-2013–, the time it takes to solve commercial disputes brought before local courts.[10] They explain that their goal was to find evidence of a positive correlation between the rule of law and economic growth; however, constructing a rule of law measure for Mexican states –a multidimensional concept that should be constructed from indicators of property rights, the efficiency and independence of the judicial system, crime rates, efforts to combat corruption, political stability, and so on– is a difficult task, since there is not enough data available.

More recently, Chávez *et al.* (2017) show evidence to affirm that *economic complexity* (or *productive knowledge*) is an important factor in explaining the disparities in the growth rates of Mexican states in the period 1998-2013.[11] They conclude that the states that have reaped most benefit from the trade reforms are those with a more complex structure, *i.e.*, those specializing in more economic activities (are more diverse) or in economic activities that are more complex or sophisticated (are less ubiquitous). As in HH, they find evidence that using one variable, economic complexity, to explain state growth rates is at least as good as the traditional approach, where a numerous of variables are necessary to explain these rates.

Several variables have been found to be relevant in explaining the states' growth rates after trade liberalization (including human and physical capital, various measures of infrastructure and agglomeration, FDI, and the efficiency of the judicial system, among others); however, economic complexity seems to provide the most parsimonious explanation. Nevertheless, a flaw of Chávez *et al.* (2017) is their failure to take into account the spatial dimension of economic complexity.

As classical models of trade, new trade theories, and new economic geography models predict,[12] and previous studies applied to Mexico have documented,

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10 The data on the ease of enforcing contracts come from the World Bank's *Doing Business* reports.

11 Economic complexity as a predictor of economic growth is illustrated empirically at the international level by HH and at the subnational level for Mexico by Chávez *et al.* (2017).

12 These models expect more integration or trade to lead to an increase in economic concentration, either in the form of industrial localization or in the level of specialization of the states. Diverse studies have evaluated the predictions of these models by examining the changes in the patterns of localization and specialization and found evidence in favor

the concentration of production has increased since the signing of NAFTA; therefore, economic complexity must be useful in explaining the growth rate of any given state and that of its neighboring states. To show this, we use spatial growth panel regressions.

3. Data and Methodology for Calculating the Economic Complexity Index (ECI) and its Spatial Lag

In this section, we describe the variables used in the spatial growth panel regressions that we will calculate to show the connection between growth rates and ECI. This includes an explanation of the methodology used to compute the two main independent variables: ECI and ECI spatial lag.

The dependent variable, average state per capita GDP growth rate, is computed using data from the Economic Information Bank of Mexico's National Institute of Statistics and Geography (INEGI) and the National Population Council (CONAPO).

The main independent variable, the ECI, is computed using data on the number of people employed (PE) in each state and each economic activity from INEGI's economic censuses.[13] We employ the Method of Reflections (MR) proposed by HH to calculate the ECI for each state. The ECI measures the productive knowledge embedded in each state economy or the sophistication of its productive structure. It is calculated by combining information on the diversity of each state (*i.e.*, the number of economic activities in which each state specializes) and the ubiquity of economic activities (*i.e.*, the number of states that specialize in each economic activity). Intuitively, more complex economies are, in general, diverse and specialize in less ubiquitous economic activities.

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of this hypothesis. The studies that analyze specific countries focus principally on the E.U. [see, Amiti (1999), Storper *et al.* (2001), Ezcurra *et al.* (2006), and Krenz and Rübel (2010), among others]. At the regional level, they primarily discuss the experience of developed economies, for example, the U.S. [see Kim (1995), Kim (1999), and Mulligan and Schmidt (2005), among others]; France (Maurel and Sédillot, 1999), and Spain (Paluzie *et al.*, 2001), to mention just a few.

13 The economic census years are 1993, 1998, 2003, 2008, and 2013. The 1993 census classifies economic activities according to the Mexican Classification of Activities and Products (CMAP) system. From 1998 onwards, the censuses use the North American Industry Classification System (NAICS). The 1994 data were adapted to make them consistent with the NAICS system. We use the data at the six-digit level of aggregation and the total number of economic activities are, 620, 797, 866, 882, and 883, respectively. GZ only considers the last four censuses, *i.e.*, 1998, 2003, 2008, and 2013.

First, using the definition of Location Quotient (LQ) commonly employed in regional science literature,[14] we construct a binary matrix, $M_{s,a}$, for each year for which we have data:[15]

$$lq_{s,a} = \frac{\frac{p_{s,a}}{\sum_{a=1}^{n} p_{s,a}}}{\frac{\sum_{s=1}^{32} p_{s,a}}{\sum_{s=1,a=1}^{s=32,a=n} p_{s,a}}} \tag{1}$$

where $p_{s,a}$ is the number of people employed by state s in economic activity a; $\sum_{a=1}^{n} p_{s,a}$ is the total number of people employed by state s; $\sum_{s=1}^{32} p_{s,a}$ is the total number of people employed in economic activity a throughout the country; $\sum_{s=1}^{32} \sum_{a=1}^{n} p_{s,a}$ is the total number of people employed in the entire country.

The matrix, $M_{s,a}$, is defined as follows:

$$m_{s,a} = \begin{cases} 1 & \text{if} \quad lq_{s,a} \geq lq^{*} = 1 \\ 0 & \text{in any other case} \end{cases}$$

Intuitively, state ***s*** is considered to be specialized in economic activity ***a*** if the percentage of PE in that activity with respect to the total PE in state ***s*** is greater than or equal to the analogous percentage nationwide.

Secondly, from matrix $M_{s,a}$ we define the two dimensions needed to calculate the ECI, which describe the economic structure of states and economic activities:

Diversity of states $$\kappa_{s,0} = \sum_{a=1}^{n} m_{s,a} \tag{2}$$

Ubiquity of economic activities $$\kappa_{a,0} = \sum_{s=1}^{32} m_{s,a} \tag{3}$$

The diversity vector is obtained by summing each of the rows of matrix $M_{s,a}$; each entry of this vector indicates the number of economic activities in which a given state is specialized. Diversity is the first approximation of a state's ECI; this measure is refined later with the information that provides the ubiquity. The ubiquity vector is obtained by summing each of the columns of matrix $M_{s,a}$; each entry of this vector indicates the number of states that specialize in each economic activity. The iterative process that combines these two dimensions is:

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[14] Analogous to the definition of Revealed Comparative Advantage employed by HH.

[15] The dimensions of the matrix M are 32*n; the number of rows (32) is the number of states in Mexico and the number of columns (n) represents the number of economic activities to be considered.

$$\kappa_{s,N} = \frac{1}{\kappa_{s,0}} \sum_{a=1}^{n} m_{s,a} \cdot \kappa_{a,N-1} \qquad (4)$$

$$\kappa_{a,N} = \frac{1}{\kappa_{a,0}} \sum_{s=1}^{32} m_{s,a} \cdot \kappa_{s,N-1} \qquad (5)$$

where N is the number of iterations, which continue until the process reaches a fixpoint that occurs when the relative ranking of the $\boldsymbol{k}_{s,N}$ remains unchanged for three consecutive iterations.[16] We will refer to the complexity variable, $\boldsymbol{CX}_{s,t}$.

To quantify the spillover effect of economic complexity, we compute the spatial lag of the ECI variable. To do this, we use a row-standardized "queen contiguity" spatial weight matrix, $\boldsymbol{W}$.[17] Therefore, the variable ($\boldsymbol{W} \cdot \boldsymbol{CX}_{s,t}$) represents the average complexity of the neighbors of each state.[18]

To complement the complexity measure, we first consider a control variable that captures the growth derived from natural resource endowment, since this source of growth cannot be explained by the ECI. In some Mexican states, the exploitation of natural resources (petroleum) accounts for an important portion of their GDP. This variable is constructed with data from INEGI.

Similarly, the automotive industry has always made a very important contribution to the country's output, and its impact has increased since the signing of NAFTA. By 2016, the industry represented 3 percent of overall GDP and 18 percent of manufacturing GDP, and accounted for almost 900,000 direct jobs. Motor vehicle production has increased so much that Mexico is now the seventh largest automobile producer in the world. However, the localization of this industry is limited to just a few of the country's states. Only certain states manufacture motor vehicles, whereas all 32 manufacture motor vehicle parts, though there are huge disparities among them.[19] Whilst Chihuahua has 122,704 persons employed in motor vehicle parts manufacturing, Coahuila has 115,758, Nuevo León has 49,939, and Tamaulipas 56,507; there are eleven states (Baja

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[16] Appendix 1 shows the estimation of the ECI of each state in each census year. These results will be used in the empirical application.

[17] A queen-contiguity spatial weight matrix considers a state to be the neighbor of another if they share a common border. Each entry of this matrix takes the value of one if states share a border and zero otherwise.

[18] In Appendix 2, we offer empirical evidence of the nature of the spatial autocorrelation of the economic complexity variable (ECI). Moran's I test statistics and scatterplots do not support the null hypothesis that states are randomly distributed; instead, the results suggest that there is a positive and statistically significant autocorrelation. States with high ECI values are surrounded by high ECI states (these tend to be located in the north of the country), while states with low ECIs are surrounded by low ECI states (which tend to be located in the south) in each census year.

[19] This industry comprises three different industrial groups: motor vehicle manufacturing, motor vehicle parts manufacturing, and motor vehicle body and trailer manufacturing, representing around 52%, 46%, and 2% of the total value added of the industry, respectively.

California Sur, Campeche, Chiapas, Guerrero, Hidalgo, Michoacán, Nayarit, Quintana Roo, Tabasco, Veracruz, and Yucatán), mainly in the south of the country, that have fewer than 1,000 persons employed in this activity. Since state growth rates during this period may also be explained, in part, by the performance of this industry, we believe it necessary to control for it.

4. Results

To illustrate the spatial effects of economic complexity on growth, we use three different panel estimation methods: Pooled, Random effects, and Panel Corrected Standard Errors (PCSE). The Hausman test suggests that using random effects is more appropriate than fixed effects.[20] The Breusch-Pagan (BP) test based on the Lagrange Multiplier (LM) for random effects suggests that Pooled OLS estimation is preferred over random effects.[21] In addition, we report the PCSE estimations, as this method provides a more efficient estimation, according to Beck and Katz (1995). As can be seen in Table 1, the estimations obtained by Pooled OLS and PCSE are identical, the only difference being the estimated standard errors.[22]

We begin by showing that economic complexity is related to future economic growth or that a state's future growth rates are correlated with its initial level of complexity, exactly as Chavez *et al.* (2017) did, the only difference being that our estimations include an additional five-year period.

We do this by estimating a panel growth regression model [Equation (6)] that has as a dependent variable the average annual growth rate of per capita GDP, $\gamma_{s,t}$. As independent variables, we have the logarithm of initial per capita GDP, $\log(y_0)$;[23] a dummy variable, *Oil*, which identifies the oil mining states;[24] a dummy variable, *Aut*, which identifies states specializing in the automotive

20 The random effects models appears to be more appropriate than the fixed because: i) the Hausman test indicated it was, as reported in Table 1, and more importantly; ii) as Barro (2015) mentions, "…*with country fixed effects, it is challenging to estimate statistically significant coefficients on X variables that do not have a lot of independent variation over time within economies*," as is the case with our independent variable, economic complexity.

21 This test is also reported in Table 1.

22 We applied three different tests of cross-sectional correlation: the Frees, the Friedman, and the Pesaran (see De Hoyos and Sarafidis, 2006). In none of these were we able to reject the null hypothesis of cross-sectional independence.

23 This variable is always included in growth regressions because of the convergence hypothesis, which implies that, *ceteris paribus*, poor economies tend to grow faster than rich ones.

24 This variable is included to complement the economic complexity variable, given that the measure of complexity (ECI) cannot explain the income that comes from the exploitation of natural resources. It takes the value of 1 for states where oil mining represents more than 5 percent of the state's GDP (Campeche, Tabasco, Tamaulipas, Chiapas, and Veracruz), and 0 in all other cases.

industry;[25] and the states' economic complexity in the initial year of the period, ($CX_{s,t}$). In addition, we include time-fixed effects dummies, (p_i), one for each five-year period analyzed, which captures the common factors that affect all states in each period.[26]

$$\gamma_{s,t} = \delta + \sum_{i=1}^{3} \alpha_i \cdot p_i + \beta_0 \cdot \log(y_{s,t_0}) + \beta_1 \cdot Oil + \beta_2 \cdot Aut + \beta_3 \cdot CX_{s,t} + \varepsilon_{s,t} \qquad (6)$$

where s identifies the states, s = 1,2,…32; t identifies the periods, t = 1993-1998, 1998-2003, 2003-2008, and 2008-2013.

Columns (1), (4), and (7) in Table (1) show the results of estimating Equation (6), which are comparable to those presented in Chávez *et al.* (2017). All parameters have the expected sign. The results confirm the positive correlation between growth rates and economic complexity, with more complex states growing faster. The estimated parameter associated with this variable is always statistically significant at the one per cent level and slightly greater in value to that estimated in Chávez *et al.* (2017). The parameters associated with the dummy variable that identifies states that specialize in the automotive industry show that, in general, these had higher growth rates than the rest of the states in the country, and they are also highly significant. Similar to the estimations presented in Chávez *et al.* (2017), the parameter associated with the variable that identifies the oil-mining states is always estimated to be statistically insignificant.

Once we have shown that future growth is related to the initial level of economic complexity of a state, our aim is then to show how that future growth is also correlated to the initial economic complexity of its neighboring states. To do so, we need to include a term that incorporates the spatial effects of the ECI variable into the previous model.

Equation (7) includes a term to calculate the spillover effect of the ECI variable, ($W \cdot CX_{s,t}$). W represents the row- standardized queen contiguity matrix to compute the spatially lagged economic complexity, *i.e.* the average economic complexity of the states' neighbors. This specification is known in spatial econometrics literature as the spatial externality model (SLX). It includes the spatial lag $CX_{s,t}$ as independent variable (LeSage and Pace, 2009); it is the simplest specification for measuring spillover effects (Halleck-Vega and Elhorst, 2015), yet the most appropriate based on the spatial distribution of the ECI variable, as

25 Takes the value of 1 for states with car assembly plants (Aguascalientes, Coahuila, Guanajuato, Morelos, Puebla, San Luis Potosí, and Sonora) and 0 in all other cases.

26 This model was estimated using few variables, just as HH (2009) presented it, their argument being that if complexity and all the other variables normally included in growth regressions to capture the different capacities of economies (*i.e.*, human capital, various measures of physical capital, institutional quality measures, measures of enforcement of the rule of law, etc.) are controlled for, this last group of variables proves to be redundant.

shown in Appendix 2.[27] We expect the estimate of parameter β_4 to be positive and significant, *i.e.*, we expect that states whose neighbors have a high ECI will tend to grow faster than those whose neighbors have, on average, a low ECI. This would imply that growth depends not only on a state's own ECI but also on the ECI of its neighbors, due to the spillover effects.

$$\gamma_{s,t} = \delta + \sum_{i=1}^{3} \alpha_i \cdot p_i + \beta_0 \cdot \log(y_{s,t_0}) + \beta_1 \cdot Oil + \beta_2 \cdot Aut + \beta_3 \cdot CX_{s,t} + \beta_4 \cdot (W \cdot CX_{s,t}) + \varepsilon_{s,t} \tag{7}$$

As can be seen in columns (2), (5), and (8), the estimated values of the parameters that Equations (6) and (7) have in common –δ, β_0, β_1, β_2, and β_3– are fairly similar. The parameter of interest, β_4 –the one associated with the spatial lag of the ECI–, is always estimated to have the expected sign, regardless of the estimation method, though is nevertheless marginally statistically insignificant in all cases. These results are quite unexpected given the strong evidence in favor of the positive spatial association of the ECI. We presume that this may be occurring because the spillover effect is not homogeneous among all states (or regions) and depends instead on the ECI level of neighboring states. As shown by the maps in Appendix 2, in general, states located in the north of the country have higher levels of economic complexity, while states located on the south have lower levels of economic complexity.

To find evidence of the heterogeneity of the spillover effect, we estimate a slightly modified Equation (7), one in which the spillover effect of highly complex states is different from the spillover effect of less complex states. Equation (8) is design to quantify the difference in the spillover effects of the ECI between states with high and low ECIs. Equation (8) is similar to (7) except for its last term, which includes a dummy variable, φ, which takes the value of 1 if the state has neighbors with a higher than average mean ECI, and 0 otherwise. Therefore, $\beta_4 + \beta_5$ estimate the spillover effect among the most complex states (as shown in Appendix 2, these tend to be located in the northern part of the country).

27 There is a plethora of alternative model specifications to study spatial spillover effects, not only the SLX model. We also considered the estimation of other models: the SAR (Spatial Autoregressive) and the SDM (Spatial Durbin) models [as LeSage (2014) pointed out, the nature of spillover effects in an SLX specification is *local*; in contrast, the SAR and SDM models allow us to study *global* spillover phenomena]. However, the autoregressive coefficients in all these other cases were not different from zero; hence, following Elhorst (2014), we discarded these models as an option for measuring spatial spillover effects, which in our case, are local in nature.

TABLE 1
ESTIMATION RESULTS[28]

Variable (Parameter)	Estimation Method								
	Pooled OLS			Random Effects			Panel Corrected Standard Errors		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
log (y_0), (β_0)	–1.044*	–1.041*	–1.239**	–1.078***	–1.078***	–1.237***	–1.044**	–1.041**	–1.239***
	(–1.92)	(–1.91)	(–2.14)	(–3.75)	(–3.77)	(–5.02)	(–2.25)	(–2.24)	(–2.58)
$CX_{s,t}$, (β_3)	0.473***	0.472***	0.605***	0.479***	0.478***	0.596***	0.473***	0.472***	0.605***
	(3.02)	(2.98)	(3.25)	(3.03)	(3.00)	(3.56)	(3.29)	(3.26)	(3.73)
$W \cdot CX_{s,t}$, (β_4)	–	0.0398	–0.552*	–	0.0275	–0.514*	–	0.0398	–0.552*
		(0.13)	(–1.66)		(0.14)	(–1.68)		(0.19)	(–1.71)
($\varphi 1 \cdot W \cdot CX_{s,t}$), ($\beta_5$)	–	–	1.082**	–	–	0.998**	–	——	1.082**
			(2.06)			(2.51)			(2.203)
Aut, (β_2)	0.630***	0.632***	0.663***	0.628**	0.630**	0.659**	0.630***	0.632***	0.663***
	(2.87)	(2.88)	(2.95)	(2.49)	(2.50)	(2.46)	(2.97)	(2.99)	(3.07)
Oil, (β_1)	–0.112	–0.099	0.078	–0.089	–0.079	0.069	–0.112	–0.099	0.078
	(–0.25)	(–0.22)	(0.19)	(–0.16)	(–0.14)	(0.18)	(–0.25)	(–0.22)	(0.19)
Constant, (δ)	6.576**	6.557**	7.366***	6.741***	6.740***	7.370***	6.576***	6.557***	7.366***
	(2.46)	(2.45)	(2.63)	(4.67)	(4.71)	(6.02)	(2.903)	(2.897)	(3.173)
Hausman test[a]	–	–	–	0.98	0.99	0.91	–	----	----
BP-LM test for random effects[b]	–	–	–	0.06	0.06	0.20	–	----	----
Observations	128	128	128	128	128	128	128	128	128
Year FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
R-squared	0.387	0.387	0.411	–	–	–	0.387	0.387	0.411

Robust t-statistics in parentheses. The symbols ***, **, and * denote statistical significance at the 1, 5, and 10 percent, respectively.
a The null hypothesis of random effects cannot be rejected. Its test statistic is asymptotically distributed as a $\chi^2_{(1)}$.
b BP is testing the null of the pooled model against the alternative of random effects. Under the null hypothesis of no random individual effects, the test statistic is asymptotically distributed as a $\chi^2_{(1)}$. Its result suggests that the null cannot be rejected.

28 Following the suggestion of one of the referees, we estimate these models, excluding the last 5-year period, this as a robustness check to evaluate the impact of the global financial crisis. The estimated values of the parameters are similar, the only difference being the estimated standard errors, which increase slightly.

$$\gamma_{s,t} = \delta + \sum_{i=1}^{3} \alpha_i \cdot p_i + \beta_0 \cdot \log(y_{s,t_0}) + \beta_1 \cdot Oil + \beta_2 \cdot Aut + \beta_3 \cdot CX_{s,t} + \beta_4 \cdot (W \cdot CX_{s,t}) + \beta_5 \cdot (\varphi \cdot W \cdot CX_{s,t}) + \varepsilon_{s,t} \tag{8}$$

The results in columns (3), (6), and (9) show that there is a positive spillover effect among states with the highest levels of ECI, which is estimated to be of a similar magnitude regardless of the estimation method: -0.552+1.082=0.530, –0.514 + 0.998 = 0.484, and –0.594 + 0.936 = 0.342. This implies that the growth rates of the most complex states (in general, those closer to the U.S.) were higher not only because of their own level of complexity, but also due to the positive impact of the higher level of complexity of their neighbors.

The same results for states whose neighbors have lower than average ECIs show the spillover effects to be negative, their magnitudes being: 0.531-1.082 = –0.551, 0.484-0.998 = -0.514, and 0.342-0.936 = –0.594. In both cases, it is important to note that the magnitude of the indirect effect is high (whether positive or negative) compared to the direct effect of ECI, β_3.

The results can be summarized as follows: future growth rates are positively related to the initial level of economic complexity of a state, *i.e.*, the higher the initial level of complexity of a state, the higher its future growth rate. Furthermore, future growth rates are also correlated with the average level of complexity of a state's neighbors, *i.e.*, complexity has a spillover effect. Nevertheless, the level of economic complexity of a state's neighbors can affect growth rates either positively or negatively. States with highly complex neighbors are affected positively, *i.e.*, their future growth rates rise, whereas states with less complex neighbors are negatively affected by being geographically close to states with low levels of development.

The existence of important externality effects suggests that regional development policies require greater coordination among the various levels of government: federal, state, and municipal. The efforts of one state to improve its economic, social or demographic conditions may not be successful if the states surrounding it do not take similar actions to reach the same goal, in which case the failure to harmonize their policies would result in a waste of valuable economic resources.

Regional development would be enhanced by policies aimed at developing specific productive capabilities. A successful policy in one region might not necessarily be the best policy for other regions, *i.e.*, there is no *universal* strategy that is perfect for every region, since each region has a different economic structure, with dissimilar strengths and weaknesses. Therefore, policies should be designed carefully so as to boost the economic activities in which regions have a relative comparative advantage, where the participation of local stakeholders in the design, implementation, and management of these strategies is essential. In the literature, policy interventions aimed at spurring regional development that take into account regional diversity and are conditional on the specific characteristics of the target region are usually referred to as bottom-up policies.

5. Final Comments

The amount of productive knowledge available in any given Mexican state measured by its economic complexity index (ECI) is strongly related to its per capita GDP growth rate. Nevertheless, a state's ECI is not only related to its own rate of growth, but also to that of its neighboring states, *i.e.*, it has a spillover effect. This indirect effect is estimated to be just as important as the direct effect and is not homogeneous among all states in the country, since northern and southern states differ markedly in terms of their productive structure.

Although previous studies have mentioned the existence of spillover effects, none found them to be as significant. We believe that the spatial dimension of the adjustments experienced by the Mexican economy occurred because northern states are alike in terms of their endowment of human capital, infrastructure (transportation, communications, industry, health, etc.), inflows of foreign direct investment, distance to the most relevant market (the U.S. is the main market for Mexican exports), and so on, and decidedly different from those in the south. This is also why northern states have proved more capable of taking advantage of the new sources of growth brought by liberalization.

We consider the southern half of the country to be a region immersed in a sequence of cause-and-effect events that mutually intensify and exacerbate one another, leading to an inexorable worsening of the economic performance of the states there relative to those in the north.

One way to break this vicious circle is to implement regional development policies to trigger short-, medium-, and long-term economic growth in the south of the country.

In an effort to increase productive opportunities in three of the most economically and socio-demographically disadvantaged regions of the country,[29] the administration of President Peña Nieto (2012–2018) proposed the implementation of a Special Economic Zones (SEZ) program, inspired by the success of China's SEZ created in the 1980s (in Shenzhen, Zhuhai, and Shantou). By promoting local and foreign direct investment through tax benefits, customs and business facilitation measures, and so on, the program sought to develop the economic activities in which these regions had a comparative advantage. The administration of President López Obrador has proposed an alternative yet similar program in its National Development Plan 2019–2014. The specific project for the country's southern regions includes different incentives: modernizing the Tehuantepec Isthmus railway; improving the ports of Coatzacoalcos in Veracruz and Salina Cruz in Oaxaca; developing road infrastructure and the airport network; constructing a gas pipeline to supply domestic businesses

[29] Puerto Chiapas in the state of Chiapas, the port of Lázaro Cárdenas–La Unión (shared by the states of Michoacán and Guerrero), and the Isthmus of Tehuantepec region that includes the ports of Salina Cruz in the state of Oaxaca and the port of Coatzacoalcos in the state of Veracruz.

and consumers in 76 municipalities in the two states; and tax incentives (*i.e.*, a reduction in valued-added tax and income tax).

The expected benefits of these types of projects for southern states, which seek to strengthen their economy, may be augmented by the recent new trade agreement between the U.S., Mexico, and Canada (USMCA). If expectations are actually met and these regions succeed in developing new competitive economic activities, the southern regions may take advantage of the new sources of growth that international trade offers, just as the north of the country did more than two and a half decades ago with NAFTA. Without a doubt, access to the greatest market in the world is a huge opportunity that could help them overcome their historical lag.

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Appendix 1

Estimated Values of the Economic Complexity Index (ECI)

Table A1 shows the estimated ECI values. The state rankings according to their complexity show very little variation; this is because economies can only accumulate productive capacities gradually over time. The results are robust if the computations are done with different levels of aggregation of economic activities (*i.e.*, 4 or 5-digits).

TABLE A1
STANDARDIZED ECONOMIC COMPLEXITY INDEX (ECI)*

Estados	1993	1998	2003	2008	2013
Nuevo León	2.04 (1)	2.09 (1)	1.94 (1)	1.84 (1)	2.05 (1)
México	1.85 (2)	1.34 (4)	1.01 (7)	0.74 (8)	0.65 (9)
Chihuahua	1.48 (3)	1.51 (3)	1.77 (2)	1.68 (2)	1.43 (5)
Coahuila	1.33 (4)	1.25 (5)	1.41 (4)	1.46 (5)	1.61 (2)
Distrito Federal	1.18 (5)	1.75 (2)	1.69 (3)	1.34 (6)	1.25 (6)
Baja California	1.15 (6)	1.24 (6)	1.35 (5)	1.48 (4)	1.53 (4)
Querétaro	1.05 (7)	1.09 (7)	1.06 (6)	1.58 (3)	1.56 (3)
Tlaxcala	0.69 (8)	0.05 (15)	–0.36 (18)	–0.55 (21)	–0.39 (17)
Tamaulipas	0.63 (9)	0.63 (10)	0.88 (8)	1.10 (7)	1.04 (7)
Jalisco	0.53 (10)	0.82 (8)	0.76 (9)	0.66 (10)	0.70 (8)
Aguascalientes	0.50 (11)	0.78 (9)	0.47 (10)	0.50 (11)	0.50 (11)
Guanajuato	0.44 (12)	0.49 (11)	0.31 (13)	0.33 (12)	0.56 (10)
Sonora	0.28 (13)	0.43 (12)	0.33 (11)	0.71 (9)	0.43 (13)
Durango	0.21 (14)	–0.09 (16)	0.31 (12)	0.02 (14)	0.10 (14)
Hidalgo	0.12 (15)	–0.35 (18)	–0.50 (20)	–0.36 (16)	–0.43 (18)
San Luis Potosí	0.12 (16)	0.13 (14)	0.15 (14)	0.25 (13)	0.44 (12)
Puebla	0.11 (17)	0.13 (13)	–0.12 (16)	–0.46 (18)	–0.36 (16)
Morelos	–0.45 (18)	–0.50 (19)	–0.67 (22)	–0.69 (23)	–0.72 (23)
Yucatán	–0.48 (19)	–0.29 (17)	0.01 (15)	–0.36 (17)	–0.46 (19)
Michoacán	–0.57 (20)	–0.74 (22)	–0.79 (26)	–0.81 (27)	–0.76 (26)
Sinaloa	–0.59 (21)	–0.70 (21)	–0.27 (17)	–0.19 (15)	–0.29 (15)
Zacatecas	–0.65 (22)	–0.89 (26)	–0.96 (27)	–0.78 (26)	–0.23 (25)
Baja California Sur	–0.83 (23)	–0.83 (23)	–0.54 (21)	–0.50 (20)	–0.64 (20)
Veracruz	–0.90 (24)	–0.87 (25)	–1.01 (28)	–0.75 (25)	–0.79 (27)
Colima	–0.93 (25)	–0.85 (24)	–0.70 (23)	–0.65 (22)	–0.65 (22)
Tabasco	–1.02 (26)	–0.91 (27)	–0.76 (25)	–0.89 (28)	–0.75 (24)
Quintana Roo	–1.03 (27)	–0.69 (20)	–0.48 (19)	–0.49 (19)	–0.64 (21)
Campeche	–1.18 (28)	–1.01 (28)	–0.76 (24)	–0.71 (24)	–0.81 (28)
Guerrero	–1.25 (29)	–1.28 (31)	–1.40 (30)	–1.59 (31)	–1.56 (32)
Oaxaca	–1.26 (30)	–1.20 (30)	–1.50 (32)	–1.60 (32)	–1.36 (31)
Nayarit	–1.27 (31)	–1.18 (29)	–1.21 (29)	–1.09 (29)	–1.21 (29)
Chiapas	–1.31 (32)	–1.35 (32)	–1.43 (31)	–1.23 (30)	–1.27 (30)

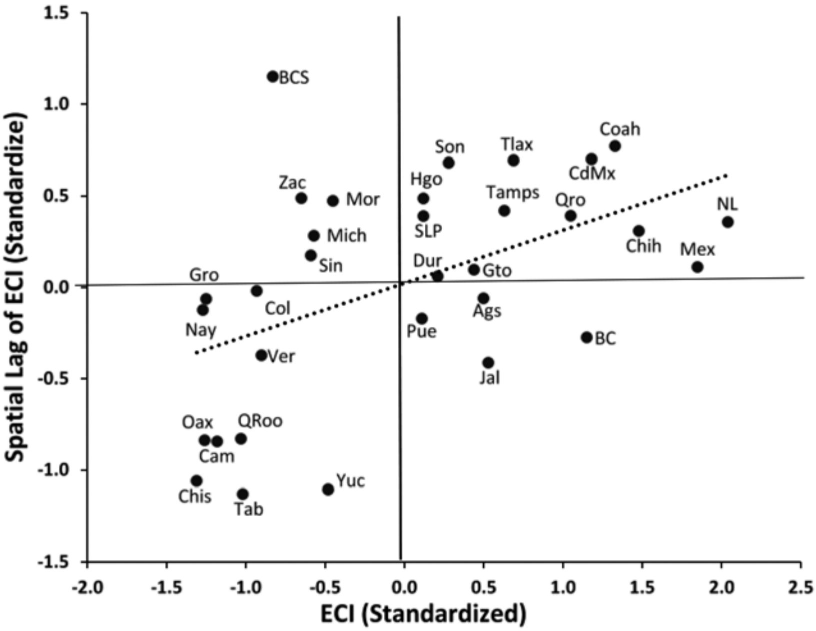
* The number in parenthesis indicates the state position in the ranking.

Appendix 2

Spatial Autocorrelation of the ECI Variable.

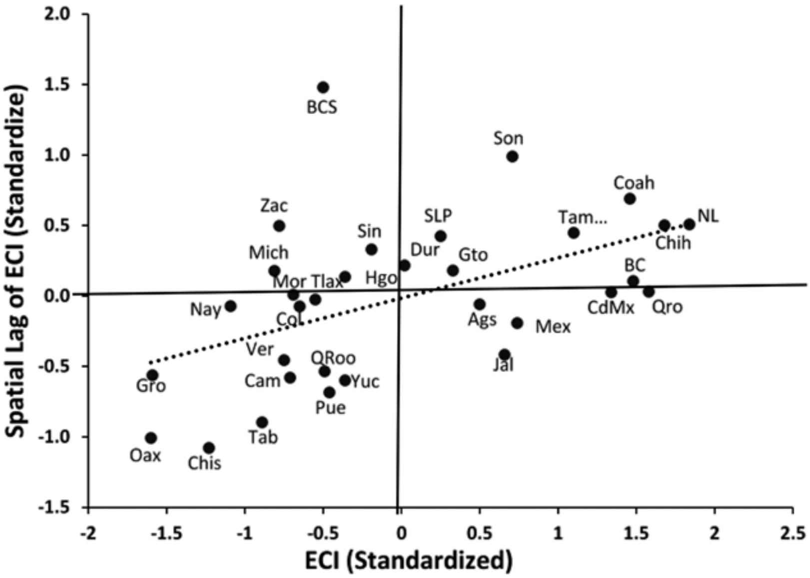
The scatterplots (and their corresponding Moran's I statistic) and maps show evidence of a very strong positive spatial dependence on the ECI variable.**

FIGURE A2.1
MORAN SCATTERPLOT OF THE ECONOMIC COMPLEXITY INDEX (1993)



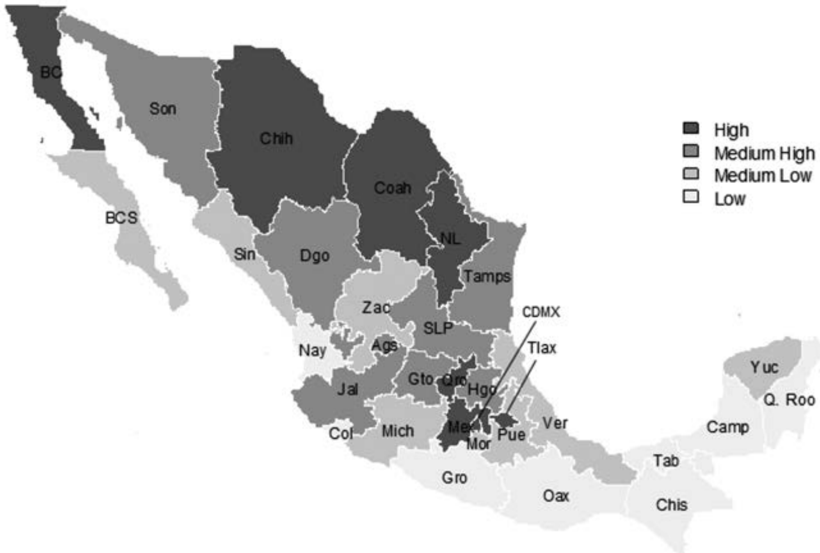
** For the sake of brevity, we show the two years for which the evidence of positive spatial dependence is more conclusive. Moran's I statistic allows us to reject the null of no spatial dependence in favor of positive spatial dependence at the 1 percent level for 1993 and 2008; at 3 percent for 2013; at 5 percent for 2003, and; at 11 percent for 1998.

FIGURE A2.2
MORAN SCATTERPLOT OF THE ECONOMIC COMPLEXITY INDEX (2008)



The maps below show the distribution of states according to their estimated ECI. There is a clear regional pattern, with more complex states being located, in general, in the northern part of the country. For 1993, we divide all the states into 4 different groups and for 2008 into 2 groups.

MAP 1
LEVEL OF ECONOMIC COMPLEXITY (ECI) OF THE STATES, 1993



MAP 2
LEVEL OF ECONOMIC COMPLEXITY (ECI) OF THE STATES, 2008

Valuing local and dual-class IPOs in the Alternative Investment Market*

Valoración de Ofertas Públicas de Venta (IPO) en el Mercado Alternativo de Inversiones

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Abstract

Initial Public Offerings are, by definition, not seasoned securities. They have not been subjected to valuation by the community of investors. It is often difficult or impossible to forecast their future cash flows because most do not have a long history of publicly disclosed financial information. Consequently, valuing IPOs in any market is more difficult than valuing seasoned equities. In this paper, we address the valuation of IPOs in the Alternative Investment Market, (hereafter the AIM.) The purpose of this study is to determine the observable factors that affect valuation in the AIM. We apply OLS, LASSO regression, and Extreme Bounds Analysis (EBA) techniques on historical accounting data to test our theory of valuation. The statistical sample consists of 2,185 IPOs issued on the AIM between 1995 and 2020. Our findings suggest that the market valuation of IPOs in the AIM is systematically related to a multiplicity of factors. These include earnings per share (EPS) in the after-market, operating cash follow per share, and the percentage of shares issued to the public. The findings of the study have a practical value for investors who are interested in buying IPOs in the AIM.

Key words: V*aluation of IPOs, AIM, Dual-class IPOs, LASSO regression, Extreme Bounds Analysis.*

JEL Code: *G12; G14; C1.*

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Resumen

Las ofertas iniciales de venta al público (IPOs), por definición, no tienen historia y no han sido valoradas por inversionistas. Por ello, es difícil o imposible realizar proyecciones de flujos futuros, al no existir información financiera pública. Este trabajo centra la valoración en el Mercado de alternativo de inversiones, determinando los observables que afectan la valoración. Aplicamos distintas técnicas econométricas a datos contables de 2,185 IPOs entre los años 1995 y 2020. Nuestros resultados sugieren que las valoraciones se relacionan con diversos factores.

Palabras clave: *Valoración, IPOs, Regresión LASSO, Análisis de límites extremos.*

Clasificación JEL: *G12; G14; C1.*

1. Introduction

All investors recognize the difficulty of valuing an IPO; the current value of these firms depends on either historical accounting information or forecasted cash flows from products and services not yet marketed. The value of an IPO in the Alternative Investment Market is even more difficult than in heavily regulated trading stock exchanges because the financial disclosure requirements for listing on the AIM are much more modest (Wahid, Mumtaz, & Mantell, 2020). Similarly, the regulatory framework of the AIM permits listing firms whether they comply or not with the relatively few rules the AIM publishes. If companies elect did not comply, they must explain why they have decided not to comply (Colombelli, 2010; Wahid, Khan, & Mumtaz, 2019). The main reason why the AIM is growing as an international stock exchange is its relatively light regulatory burden. That transactional cost advantage makes the AIM a more favorable market for cross-border or offshore listing, enabling companies to avoid the cost burden imposed by the US Sarbanes-Oxley Act (Akyol, Cooper, Meoli, & Vismara, 2014; Reutzel & Belsito, 2015).

There are several practical and theoretical reasons why the valuations of IPOs listed on the AIM are of interest. During the past two decades, only 22% of new issues were listed on the main market e.g. London Stock Exchange (hereafter LSE) whereas 78% of new issues were enlisted on the AIM (Miguel Á. Acedo-Ramírez & Francisco J. Ruiz-Cabestre, 2016; Wahid *et al.*, 2020). This shows the popularity of the AIM which is growing relative to the LSE. That growth can be expected to result in an increased incidence of mispricing. Companies selecting to launch their IPOs on the AIM are not required to disclose any specific financial credentials as a precondition for listing. That flexibility encourages newly incorporated small firms to go public which further leads to underpricing (Akyol *et al.*, 2014). The scarcity of reliable financial and accounting information makes the valuation of IPOs in the AIM more difficult that it would be in the main market. The regulatory framework of the AIM allows foreign companies to list their securities, which exacerbates the difficulties for underwriters because they must take into account the complexities of foreign exchange variability as well as parental market dynamics. Moreover, there is no minimum requirement on the AIM for the size of the listing firm or the number of shares to be held by the investing public (Wahid, Mumtaz, & Mantell, 2019).

To determine the price performance of IPOs across markets and time periods, many studies have been conducted and they documented that IPOs underprice in the short-run (Acedo-Ramírez, Díaz-Mendoza, & Ruiz-Cabestre, 2019; Hawaldar, Naveen Kumar, & Mallikarjunappa, 2018; Mumtaz, Smith, & Ahmed, 2016) and in the long-run (Ali, 2017; Fine, Gleason, & Mullen, 2017; Mumtaz, Smith, & Ahmed, 2016). The level of IPO underperformance varies across the nationality of the issuers and exchanges (Mudambi, Mudambi, Khurshed, & Goergen, 2012). Similarly, Doukas and Hoque (2016) found that firms make their own decisions and show that these two markets, *i.e.* the AIM and the main market, attract companies with different characteristics and post-listing investment and financing priorities. Acedo-Ramírez and Ruiz-Cabestre, (2016) also found the nexus between the IPO characteristics and underpricing in the AIM. They also differentiated between AIM firms that meet the main market pre-requisites and those firms that do not. Our paper goes beyond these studies by including when a firm goes public in AIM with unique size and nationality, underwriters have little information beyond traditional valuation methods employed in the technique where a supposedly comparable firm is analyzed as a surrogate for the listing firm. The statistical incidence of mispricing in the AIM has been documented by earlier studies (Abdullah, Jia'nan, & Shah, 2017; Acedo-Ramírez *et al.*, 2019; Miguel Á. Acedo-Ramírez & Francisco J. Ruiz-Cabestre, 2016; Wahid, Khan, *et al.*, 2019; Wahid *et al.*, 2020; Zheng, 2007).

The regulatory and operational dynamics of the AIM suggest numerous hypothetical explanations for the mispricing of IPOs. To identify the observable factors associated with the mispricing of IPOs in AIM, this study is focused on four research questions: (a) How can one characterize the pricing of IPOs based on the accounting information disclosed? (b)What are the financial factors that appear to be systematically related to the pricing of IPOs? (c) What are the robust predictors of IPO offer prices? and (d) Does the domicile status of the firm offering the IPO affect its offer price? This study employs the firm size, the age of the firm, market conditions, the offer size, and classification of local and cross-listing as the control variables. In this study, we use the EBA technique and LASSO regression because it reduces the ambiguity in selecting the explanatory variables and mitigates the uncertainty associated with model specification.

The rest of the paper is structured as follows. Section 2 elaborates the literary review focuses on the theoretical discourse on the valuation of IPOs. Section 3 explains the data, sample size, the econometric model, and statistical techniques to determine the robust factors affecting the valuation of IPOs. Section 4 describes descriptive statistics, the correlation matrix of criterion and outcome variables of the study, and also shows the inferential statistics including OLS and sensitivity analysis through EBA and LASSO regressions. Finally, section 5 concludes the study.

2. The theories of IPO valuation

2.1. Methods of IPO pricing

Two methods have been used in the literature for valuing the IPOs. These are the *comparable firm* approach- and another valuation method is called the regression method. The comparable firm approach is frequently used by investment bankers to value IPOs (see Kim & Ritter, 1999). The regression method is commonly employed by academics and researchers (Bartov, Mohanram, & Seethamraju, 2002). The comparable firm method has been widely used by underwriters if they can identify a firm "comparable" to the IPO. That comparable firm has designated a benchmark for determining the IPO offering price (Kim & Ritter, 1999). This method takes into account the relative value of assets of a competitive firm and then prices the shares of the IPO company based on this relative value using various financial indicators (Agnes Cheng & McNamara, 2000; Rasheed, Khalid Sohail, Din, & Ijaz, 2018).

The most popular method used in the comparable firm approach for the valuation of IPOs is the dividend discount model. That model is based on the proposition that the value of a firm's stock is equal to the discounted value of the infinite cash flow of the expected dividends per share (Rasheed, Khalid Sohail, Din, & Ijaz, 2018; (Gacus & Hinlo, 2018; Sim & Wright, 2017). The other approach firm analysis for the IPO valuation is the discounted cash flow method. That method is based on the proposition that the value of a company is based on the expected future cash flows discounted at their present values (Alhadab, Clacher, & Keasey, 2016; Shapiro *et al.*, 2019). A third valuation method employs earnings or sales concerning the market price for determining the offer price for the shares(Fernandez, 2011; Kumar, 2016). The OLS method is commonly used in academic research to determine the factors that influence IPO pricing (Beatty, Riffe, and Thompson,2000).

2.2. The magnitude of IPO mispricing

The mispricing of IPOs seems to be ubiquitous and durable. Rock (1986) found a general trend among the investors buying stocks in the secondary markets at prices exceeding the offer prices. This phenomenon was reported at11% in the US market from 1963-1965 (Reilly & Hatfield, 1969) and after that 21.14% in USA (640), 43.95% in Japan (609), 20.16% in the UK (471), 18.04% in Australia (437), 13.12% in France (171), 37.20% in Germany (132), 34.97% in Greece (124) and 32.04% in the Indian market (292) (Wahid *et al.*, 2020). The mispricing effect was also documented emerging markets where the average initial return was 462% for 101 IPOs during the 1990-1993 period in China (Tan, Dimovski, & Fang, 2015), 231% for 308 IPOs issued in the 1985-1995 period in China (Haggard, Walkup, & Xi, 2015) and 175% for 570 IPOs issued in Malaysia (Komenkul & Kiranand, 2017).

This evidence confirms that mispricing has been a pervasive phenomenon that exists almost in every market. The extent of the pervasiveness suggests that there are factors beyond the accounting information and forecasted earnings which are systematically associated with the mispricing of IPOs. In the next section, we describe the behavioral theories purporting to explain the mispricing of IPOs and the statistical evidence consistent with those theories.

2.3. Factors affecting IPO pricing

Earlier studies suggested information asymmetry as the main factor causing mispricing of IPOs by the offering firm. (see Bouzouita, Gajewski, & Gresse, 2015; McGuinness, 2016; Naifar, 2011; Wahid *et al*., 2020). An example of the asymmetry theory suggests that investors misprice the offering due to incomplete information relating to the firm's specific characteristics (Wahid, Khan, *et al*., 2019). That study employs published accounting information to determine the factors that cause the pricing of IPOs. The information includes EPS, operating cash flow per share, sales per share (Beatty *et al*., 2000), book value per share, the annual sales growth, growth of profit (Kim & Ritter, 1999), and the percentage of shares offered.

Some studies focused on the ex-ante uncertainty hypothesis as a factor responsible for mispricing (see Mantell, 2016).That theory suggests that the risks perceived by investors can be dichotomized into pre- and post-IPO uncertainty. Other studies used the firm age at the time of offering (Rathnayake, Louembé, Kassi, Sun, & Ning, 2019), and the offer size (Mumtaz *et al*., 2016) as proxies for pre-IPO uncertainty. A theory purporting to explain the price performance of IPOs in the aftermarket is related to the prestige of the underwriters (Migliorati & Vismara, 2014). This theory suggests that the luster of the underwriter's reputation is inversely associated with the magnitude of underpricing (Arora & Singh, 2019).

The signaling hypothesis suggests that high-quality large issuers intentionally underprice their IPO to signal to investors that the quality of their offer differs from the offers of low-quality firms (Badru & Ahmad-Zaluki, 2018). Market sentiment and investor sentiment can also be explained in terms of signaling theory (Colombo, Meoli, & Vismara, 2019; Obrimah, 2018). The volatility of market activity is also thought to influence the pricing of IPOs. The window-of-opportunity hypothesis develops the nexus between the timing of an issue and its mispricing. The theory suggests that in a hot market environment issuers tend to overprice their issues (Ritter, 1991). To examine the robustness of the causal factors related to IPO pricing, we used different proxies related to the above theories. These variables include: offer size and firm age as proxies of the ex-ante uncertainty hypothesis, firm size and duality of the firm listing (cross-listed IPOs) as proxies of signaling hypothesis, underwriter's prestige as a proxy for the underwriter reputation hypothesis and market condition as a proxy of the window of opportunity hypothesis.

3. Methodology

3.1. The Data and the Sample

We divided our population into two sub-samples: (a) local IPOs and (b) dual-class IPOs (cross-listed IPOs) in the AIM during the period from July 1995 to December 2019. A total of 2,226 new issues were listed on the AIM, including 1,801 locally incorporated IPOs and 425 foreign firms listed on the AIM defined as a secondary listing. The overview of these IPOs is presented in Table 1 and Figure 1.

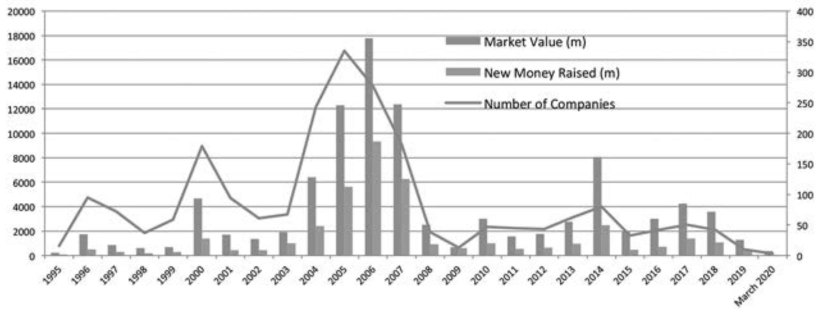
TABLE 1
THE LISTINGS OF IPOS IN THE ALTERNATIVE INVESTMENT MARKET

Year	Number of Companies	Market Value (m)	New Money Raised (m)
1995	16	208.000	69.087
1996	95	1757.000	504.257
1997	72	844.203	299.353
1998	37	602.969	185.110
1999	59	673.952	274.367
2000	179	4666.737	1395.267
2001	94	1715.668	434.913
2002	61	1338.591	433.018
2003	67	1901.531	989.820
2004	243	6385.949	2412.258
2005	335	12299.048	5632.464
2006	278	17785.840	9314.644
2007	182	12384.884	6262.350
2008	38	2508.298	917.269
2009	13	665.954	610.056
2010	47	3024.441	1012.001
2011	45	1571.542	525.095
2012	43	1779.934	642.898
2013	62	2750.771	973.588
2014	80	8064.514	2472.468
2015	33	1972.906	470.001
2016	42	3000.730	710.160
2017	50	4232.391	1379.449
2018	42	3575.811	1065.716
2019	10	1276.536	417.004
March 2020	3	370.310	48.500

Note: All monetary units are expressed in British pounds.

We selected a population of 2,185 firms. We collected the statistical and other data from the websites of the issuing firms and the London Stock Exchange (LSE).

FIGURE 1
IPOS ISSUED IN THE AIM FROM 1995 TO 2020



3.2. The Econometric Specification

To test the theories, we applied an OLS regression to evaluate the factors that influence IPO pricing (e.g. Bartov *et al*., 2002; Beatty *et al*., 2000; Kim & Ritter, 1999; Pukthuanthong-Le, 2008):

$$Price_{IPO,i} = \alpha_i + \delta_1(EPS_i) + \delta_2(OCPS_i) + \delta_3(SPS_i) + \delta_4(BVPS_i) + \delta_5(SG_i) + \delta_6(PG_i) + \delta_7(PSO_i) + \gamma_1(Firm_{Size_i}) + \gamma_2(Offer_{Size_i}) + \gamma_3(Mkt_{Cond_i}) + \gamma_4\left(Firm_{Age_i}\right) + \gamma_5\left(Undwr_{Rep_i}\right) + \gamma_6(Firm_{Class_i}) \tag{1}$$

TABLE 2
EXPLANATION OF VARIABLE

Variable	Measurement of variables
Price IPO	It is an offer price of which shares are offered to investors.
EPS	Earnings per share of a firm going public.
OCPS	This is the operating cash flow per share before the offering.
BVPS	The book value per share measured as the stockholders' offering.
SG	It is the growth of sales revenue measured by the percentage change in sales.
PG	It is the profit growth which is estimated by the percentage change in the profit.
PSO	It refers to the percentage of shares offered and calculated as the number of shares offered divided by total shares outstanding.
Firm Size	Firm size is the natural logarithm of the total assets of the issuer.
Offer size	Offer size is the total monetary value of the offering.

Table 2 (cont.)

Variable	Measurement of variables
Mktcond	It refers to the market condition and defined as a dummy variable. If the total volume of offerings in the market is higher than the average volume it is recognized as a hot market and categorizedas1, and 0 otherwise.
Firm age	Firm age at the time of offering.
Undwrep	This shows the prestige of underwriters. A dummy variable is assigned as1 if the prestige of the underwriters is high and 0 otherwise. We use total market capitalization as a measure to compute the repute of underwriters.
Firm class	This indicates the class of firm and it is a dummy variable assigned as 1 for local IPOs and 0 for cross-listed IPOs.

3.3. Statistical techniques

To test our propositions, we used robust regression in this study. The purpose of employing a robust regression method is that other techniques do not adjust for outliers. In many of those applications, outliers have been unduly influential. To overcome the problem of outliers in these techniques, researchers applied OLS with a prescription of robust regression. In the first step, we use all Z variables in the robust regression to find out the potential impact of all variables on the valuation. The basic model for choice of function ρ of the residuals is as follow:

$$\text{Huber Model} \begin{cases} \frac{x^2}{2} if\,|X| \leq c \\ c|X| - \frac{c^2}{2}\, otherwise \end{cases} \tag{2}$$

The default tuning constants for each function are taken from Holland and Welsch (1977), and are chosen so that the estimator achieves 95% asymptotic efficiency under residual normality. In the next step, we also use Median Absolute Deviation - Median Centered (MADMED) method:

$$\text{MADMED}, \hat{\sigma}^{(\delta)} = Median\left[\frac{abs\left(r_i^{(\delta-1)} - Median\left[r_i^{(\delta-1)}\right]\right)}{0.675}\right] \tag{3}$$

Maronna & Morgenthaler (1986) defines the robust R^2 statistic of robust regression as:

$$R^2 = \frac{\sum_{i=1}^{N} P_c\left(\frac{y_i-\hat{\mu}}{\hat{\sigma}\omega_i}\right) - \sum_{i=1}^{N} P_c\left(\frac{\tau_i}{\hat{\sigma}\omega_i}\right)}{\sum_{i=1}^{N} P_c\left(\frac{y_i-\hat{\mu}}{\hat{\sigma}\omega_i}\right)} \tag{4}$$

Information criteria for M-estimated equations describe the robust equivalent of the Akaike Information Criterion (AIC_R), and a corresponding robust Schwarz Information Criterion

$$(BIC_R)\colon AIC_R = 2\sum_{i=1}^{N} P_c\left(\frac{r_i(\beta)}{\hat{\sigma}\omega_i}\right) + 2k\left\{\frac{\sum_{i=1}^{n}\varphi_c\left(\frac{r_i(\beta)}{\sigma\omega_i}\right)^2}{\sum_{i=1}^{n}\varphi_c\left(\frac{r_i(\beta)}{\sigma\omega_i}\right)}\right\} \tag{5}$$

In the next step, we use two techniques that are Extreme Bounds Analysis (EBA) and Least absolute shrinkage and selection operator (LASSO) to determine the robust determinants of the price of IPOs. According to Cooley & Leroy (1981), the economic theory does not indicate which of the variables are robust and which should be kept constant while employing any statistical technique or model. To address this concern, Leamer(1983, 1985) developed the Extreme Bound Analysis (EBA) and applied by Levine & Renelt (1992). To determine the robust predictors, we construct the following regression (Moosa and Cardak, 2006):

$$Price_i = \beta_0 + \sum_{ip=1}^{n} \beta_c X_{ipi} + \mu_i \tag{6}$$

$$Price_i = \beta_0 + \sum_{ip=1}^{n} \delta_c X_{ipi} + \beta Q_i + \sum_{ip=1}^{m} \delta_c z_{ipi} + \mu_i \tag{7}$$

We estimate the coefficient of the variable of interest Q. The coefficient of that variable is an indicator of sensitivity and robustness. The methodology of robust regression requires many regressions to estimate the value of the coefficient of the independent variable. The fixed variable(s) X are included in every set of regressions. The variable of interest Q and the set of variables Z is chosen from a predetermined pool. Furthermore, to get more clarity about the specification of the model and robustness of variables, we use LASSO regression which is widely used to select both variables and measure the accuracy model. This technique was first time introduced by Santosa & Symes(1986) and used by (Tibshirani, 1996). The LASSO estimator is the OLS estimator with an L1 penalty term:

$$Price_i = \frac{1}{2m}\sum_{i=1}^{m}\left(y_i - \beta_0 + \sum_{j=1}^{p} x_i\beta_j\right)^2 + \lambda\sum_{j=1}^{p}|\beta_j| \tag{8}$$

The nature of L1 regularization penalty causes some coefficients to be shrunken to zero. Here the turning factor λ controls the strength of the penalty that is $\lambda = 0$. In this situation, coefficients are considered as simple linear regression. Likewise, when $\lambda = \infty$ then all coefficients are zero. In nutshell, $0 < \lambda < \infty$

means the coefficients between 0 and that of simple linear regression. So, when λ falls between the two extremes, we are balancing the below two ideas. The Lasso regression can perform variable selection in the linear model. Thus, as the value of λ increases, more coefficients will be set to value zero (provided fewer variables are selected) and so among the nonzero coefficients, more shrinkage is employed.

4.. Findings and analysis

4.1. The IPO market in the AIM

Table 1 depicts the history of IPO activities listed on the AIM from 1995 to 2020. Since January 2020, more than 75% of new issues were listed on the AIM and 25% were listed in the LSE. A total of 2,226 new issues have been listed on the AIM since 1995. Of those, 1,801 were locally incorporated IPOs and 425 were incorporated in foreign countries. The total market capitalization of the AIM was £97,358 million. During the period from 1995 to 2020 £39,451 million of new money was raised from IPO listings. The decade from 2001 to 2010 was unusually active for IPOs on the AIM. During that decade, 1,358 IPOs were launched, constituting more than 60% of all the IPOs listed on the AIM as mentioned in Table 3. Figure 1 demonstrates the trend of IPOs issued over the sample period.

TABLE 3
DESCRIPTIVE STATISTICS

	Local IPOs			Cross-listed IPOs		
	IPOs	Capitalization	New money raised	IPOs	Capitalization	New money raised
1995-2020	1801	69,355	28,617	425	28,003	10,834
1995-2000	419	7,726	2,437	39	1,027	290
2001-2005	665	16,478	7,167	135	7,163	2,735
2006-2010	389	22,071	12,133	169	14,299	5,983
2011-2015	203	12,169	4,039	60	3,971	1,045
2016-2020	125	10,912	2,840	22	1,543	781

Note: All monetary units expressed in British pounds This table displays the sample of 2,226 new issues. It includes 1,801 locally incorporated firm's IPOs and 425 foreign countries incorporated firms that are listed on the AIM for secondary listing or offshore listing from 1995 to 2020. Market capitalization and new money raised is quoted in millions of British pounds.

4.2. Descriptive statistics and the correlation matrix

In the Table 4, the descriptive analysis of the variables shows that the average price of IPOs in the AIM was 79.275 British pounds. Most of the IPOs listed on AIM were pre-sold by the sales process known as book building. The average firm size was 42 million pounds and the average offering size was 17 million pounds. The average EPS of the firms in the sample was 6.5%, the average operating cash flow per share was 12.99; the average operating sales per share was 10.61; and the average book value of the stock at the time of the offering was 9.88. According to Amini, Keasey, and Hudson (2012), access to market-based equity finance is easier for small firms in capital markets. The sale and profit growth are 6.46% and 8.95% relatively for young firms (less than 2 years) at the time of offering. These facts suggest that these firms are growing rapidly.

Most firms prefer to issue IPOs into what they believe to be a hot market. Most issuers prefer the offering to be managed by prestigious underwriters, if feasible. Most issuers listed on the AIM are locally incorporated small firms; the evidence relating to the effect on the offer price of the domicile of IPOs issuers shows that the majority of small IPOs are incorporated in the London-based market. According to Amini, Keasey, and Hudson (2012), access to market-based equity finance is easier for London-based firms. Additionally, the AIM is characterized by a substantial concentration of Small Medium Enterprises (SMEs), most of which are located in London. The correlation matrix (Table 4) indicates that no variable is highly correlated with any other which mitigates the difficulties associated with multi-collinearity.

4.3. Results of the basic model with all the Z variables

To disentangle the multiplicity of factors affecting the offer prices of IPOs, we apply the OLS specification represented by equation (3) above. The dependent variable is the offering price of IPOs. In the Model-I, we use all seven Z-variables. These include EPS, operating cash flow per share, sale per share, the book value of equity per share before the offering, profit growth before offering, and the percentage of shares issued. Table 5 presents the results of OLS estimates. The results of Model-I show that there is a significant and positive impact of earnings per share ($\beta = 1.758$, $p < 0.01$), operating cash flow per share ($\beta = 1.402$, $p < 0.01$), and sale per share ($\beta = 0.718$, $p < 0.01$) on the issue price of IPOs in AIM except for the percentage of share issued ($\beta = -0.771$, $p < 0.01$) which has a negative role in deciding IPOs pricing. We found insignificant effect in sales per share, book value of equity per share, profit growth in deciding IPOs pricing in AIM.

In the next step, we include control variables (*i.e.* firm size, offer size, market condition, firm age, underwriters' reputation, and dual-class IPOs). These variables are added to the specification incrementally from Model-II to Model-VII. To test the robustness of the control variables, we apply three

TABLE4
DESCRIPTIVE STATISTIC AND CORRELATION MATRIX

	Mean	Std. Dev	Price	EPS	OCPS	SPS	BVS	SG	PG	PSO	FSize	Osize	MC	FAge	URP	Class
Price	79.275	81.376	1													
EPS	6.573	2.194	.562**	1												
OCPS	12.997	3.795	.434**	.562**	1											
SPS	10.601	4.315	.401**	.530**	.562**	1										
BVS	9.884	3.245	–.027	.016	–.003	–.001	1									
SG	6.462	2.832	.006	.001	–.027	–.012	.014	1								
PG	8.985	2.586	–.024	–.025	–.025	–.002	–.009	–.009	1							
PSO	36.433	13.638	–.070**	.098**	.218**	.249**	.025	–.009	.011	1						
FSize	42.91	70.67	.523**	.436**	.342**	.281**	–.006	.007	–.045*	–.109**	1					
Osize	17.52	37.17	.486**	.410**	.330**	.300**	.004	.002	–.039	–.055*	.526**	1				
MC	0.682	0.466	.003	.006	–.011	–.006	.026	–.007	–.038	–.029	.007	–.008	1			
Firm Age	2.213	1.441	.088**	–.077**	–.233**	–.268**	.017	–.021	–.043*	–.400**	.144**	.111**	.017	1		
URP	0.501	0.500	–.014	–.024	–.011	–.027	–.011	–.015	.018	.026	–.032	–.037	.016	–.017	1	
Class	0.189	0.391	.121**	–.130**	–.345**	–.397**	.017	–.002	–.014	–.586**	.228**	.162**	.053*	.663**	–.020	1

Note: Issue price, (reported in British pounds), EPS is an average of earnings per share(reported in %), OCPS is the operating cash flow per share(reported in British pounds), BVPS is the book value per share(reported in British pounds), SG is the sale growth (reported in %) and PG is profit growth (reported in %). PSO is the percentage of share offered (reported in %), Fsize is firm size(reported in million British pounds), Osize is Offer size(reported in millions of British pounds), and MC is market condition is dummy variable. Firm age is the total age of firm URP is a dummy variable which is assigned a value of 1 if the prestige of the underwriters is high and 0 otherwise and class is categorized as a dummy variable assigned as 1 for local IPOs and 0 for cross-listed IPOs. *P<0.05; ** p<0.01 represent significance level at the 1, and 5% respectively.

TABLE 5

ORDINARY LEAST SQUARES ESTIMATES OF PARAMETERS

	Model–I	Model–II	Model–III	Model–IV	Model–V	Model–VI	Model–VII
EPS	1.758**	1.371**	1.365**	1.366**	1.353**	1.353**	1.312**
	(28.30)	(24.07)	(23.98)	(23.97)	(23.82)	(23.83)	(24.50)
Operating Cash flow/Shares	1.402**	1.149**	1.145**	1.145**	1.178**	1.177**	1.369**
	(18.90)	(17.35)	(17.31)	(17.30)	(17.77)	(17.75)	(21.55)
Sales/Shares	0.718**	0.640**	0.631**	0.631**	0.663**	0.665**	0.894**
	(12.64)	(12.75)	(12.57)	(12.56)	(13.12)	(13.15)	(18.03)
Book Value/Shares	0.029	0.038	0.036	0.037	0.032	0.033	0.014
	(0.59)	(0.87)	(0.84)	(0.85)	(0.75)	(0.77)	(0.35)
Sales Growth	0.038	0.026	0.027	0.027	0.032	0.032	0.038
	(1.15)	(0.90)	(0.93)	(0.92)	(1.09)	(1.11)	(1.38)
Profit Growth	–0.066	–0.030	–0.030	–0.030	–0.021	–0.023	–0.039
	(1.17)	(0.61)	(0.60)	(0.60)	(0.43)	(0.46)	(0.83)
% Share Offered to public	–0.771**	–0.532**	–0.534**	–0.534**	–0.482**	–0.483**	–0.094*
	(17.82)	(13.51)	(13.57)	(13.57)	(11.79)	(11.82)	(2.08)
Firm Size		0.322**	0.280**	0.280**	0.274**	0.274**	0.212**
		(24.88)	(13.65)	(13.65)	(13.42)	(13.42)	(10.83)
Offer Size			0.048**	0.048**	0.044*	0.045*	0.036*
			(2.66)	(2.66)	(2.47)	(2.50)	(2.13)
Market Condition				–0.004	–0.006	–0.007	–0.019
				(0.21)	(0.33)	(0.38)	(1.11)
Firm Age					0.106**	0.106**	–0.052*
					(4.46)	(4.46)	(2.16)
Underwriter's Reputation						0.047	0.053
						(1.56)	(1.86)
Dual Class							1.005**
							(16.62)
constant	–1.925**	–2.266**	–2.178**	–2.173**	–2.584**	–2.604**	–4.665**
	(7.70)	(10.24)	(9.75)	(9.67)	(10.68)	(10.75)	(17.97)
R^2	0.67	0.72	0.73	0.72	0.74	0.73	0.77
df	08	09	10	11	12	13	14
AIC	5265.47	4720.41 4771.61	4715.30 4772.19	4717.26 4779.84	4699.31 4767.59	4698.88 4772.84	4439.311
BIC	5310.99						4518.963

Note: This table displays the findings in a sample of 2185 IPOs. The sample consists of1773 local IPOs and 412 Cross-listed IPOs issued on the AIM from 1995 to 2021. * <0.05; ** p<0.01 represent significance level at the 1, and 5% respectively. The issue price is the dependent variable, AIC = Akaike's Information Criterion, BIC = Bayesian Information Criterion.

criteria: Akaike's Information Criterion (AIC), Bayesian Information Criterion (BIC) and the R^2. In the Model-II, we added the variable representing firm size. The test statistics are: AIC = 4720.41, BIC = 4771.61, R^2 =0.72. The value of β for that variable is significant at 99% confidence interval (p < 0.01).The explanatory power of this Model is superior to Model-I, as is signified by a lower AIC and BIC and a higher R^2. These findings indicate that firm size is systematically related to the prices of IPOs. Larger firm size leads to the probability of higher IPO pricing. In the Model-III, we included the offer size as an explanatory variable. The test statistics are: AIC = 4715.30, BIC = 4772.19 and R^2 = 0.73. The value of β for that variable is significant at 99% confidence interval ($p < 0.01$).The economic significance of the offer size is that firms have more options to generate funds in AIM because of its international scope. A reasonable explanation of this finding is that large issuers are attractive to a more diverse population of potential investors; there by generating higher prices in the after-market. In Model-IV, we found that the market condition has an insignificant effect on the pricing of IPOs. This finding constitutes evidence tending to invalidate the window of opportunity hypothesis which suggests during periods of hot market issuers tend to price their issues.

Prior literature reported the positive relationship between underpricing and firm size. (Sahoo and Rajib 2010; Diro Ejara and Ghosh 2004; Mumtaz, Smith, and Ahmed 2016).This evidence supports the ex-ante uncertainty hypothesis that the availability of historical information of firms leads to a lower probability of IPO mispricing. In general, the prestige and expertise of underwriters affect the pricing of IPOs. We found that high prestige underwriters tend to be associated with a smaller degree of mispricing. The statistical findings in Model-VI support the proposition that underwriters' reputation in our sample is not significantly related to the offer price. That finding directly contradicts the finding in Model V. We added the listing classification of IPOs in Model-VII and reported that the price of local and dual-class IPOs systematically varies concerning the nationalities of the issuer and the prestige of the underwriters in the AIM. This finding is consistent with the signaling hypothesis: High-quality large firms intentionally underprice their issue to differentiate their status in the market from the low-quality firm (Badru & Ahmad-Zaluki, 2018).

4.4. Sensitivity analysis using LASSO regression and Extreme Bounds Analysis

To test the sensitivity and robustness of the explanatory variables, this study applies the Extreme Bounds Analysis (EBA) technique. We compare the results of the EBA technique with other methods which include the Akaike's information criterion (AIC) and Bayesian information criterion (BIC).We applied a large number of regressions to predict the values of the coefficients. We include fixed variables (X) in every set of regression, a specific variable of interest, Q and the set of Z variables chosen from a predetermined pool of combinations. The sample statistics are displayed at the bottom of Table 6.

The result of the EBA indicates that EPS ($\beta = 1.205$, $p < 0.01$), operating cash flow per share ($\beta = 1.367$, $p < 0.01$), and sale per share ($\beta = 0.891$, $p < 0.01$) are the robust parameters explaining the pricing of IPOs. Firm size, offer size, firm age, and class of the firm are the fixed variables shown in table (5). Similarly, the result of LASSO indicates that EPS, operating cash flow per share, sale per share, firm size, offer size firm age, and class of the firm arise emerged as robust determinants of the value of IPOs. The optimization of this combination has been tested through lower AIC, and BIC values. Our findings indicate that EPS, operating cash flow per share, and sale per share are significantly correlated with the offer price of IPOs in the AIM.

TABLE 6
COMPARISON OF THE THE EBA AND LASSO WITH OTHER TECHNIQUES

	Pre-OLS	EBA	LASSO	Post-OLS
EPS	1.312**	1.131**	1.101**	1.205**
	(24.50)	(27.97)	(33.73)	(39.75)
Operating Cash flow/Shares	1.369**	1.490**	1.131**	1.367**
	(21.55)	(16.22)	(19.23)	(21.51)
Sale/Shares	0.894**	0.642**	0.639**	0.891**
	(18.03)	(15.12)	(11.51)	(17.96)
Book Value/Shares	0.014			
	(0.35)			
Sale Growth	0.038			
	(1.38)			
Profit Growth	–0.039			
	(0.83)			
% Share Offered	–0.094*			
	(2.08)			
Firm Size	0.212**	0.267**	0.251**	0.300**
	(10.83)	(14.89)	(10.33)	(12.71)
Offer Size	0.036*		0.031**	0.082**
	(2.13)		(3.85)	(4.02)
Market Condition	–0.019			
	(1.11)			
Firm Age	–0.052*	–0.093**	–0.101**	–0.081**
	(2.16)	(2.51)	(3.34)	(2.76)
Underwriter’s Repute	0.053			
	(1.86)			
Dual Class	1.005**	0.241**	0.176**	0.289**
	(16.62)	(3.88)	(5.54)	(4.26)
_cons	–4.665**	–0.467**	–0.498**	–0.885**
	(17.97)	(4.35)	(4.94)	(3.40)
R^2	0.64	0.60	0.61	0.61
AIC	4439.31			4038.87
BIC	4518.96			4084.38

Note: This table displays the findings in a sample of 2185 IPOs that consists of 1773 local IPOs and 412 Cross-listed IPOs issued and placed on the AIM from 1995 to 2021. * < 0.05; ** $p < 0.01$ represent significance level at the 1, and 5% respectively. The issue price is the dependent variable, AIC = Akaike’s Information Criterion, BIC = Bayesian Information Extreme Bounds Analysis (EBA) was used to predict the robust factor explaining the intrinsic value of IPOs. Total 495 combinations using n!/(k!(n-k)! formula of 7 regressors (3 level combination of variables of interest) from the Z(nx13) vector.

4.5. Firm Age and the issue prices of Initial Public Offerings

To determine the significance of control variables, we further divided our data set based on control variables such as the firm's age. We partition the sample into four sub-samples: sub-sample 1 defined as the issuers with age at the date of issue less than or equal to 1 year, sub-sample 2 defined as 2 year < firm age ≤ 3 years, Sub-sample 3 defined as 3 years < firm age ≤ 5 years and sub-sample 3 defined as 5 years < firm age. Table 7 displays the descriptive analysis relating to issue price and firm age. The data show that the issue price is positively correlated with the age of the issuer.

In Table 8, we applied the OLS separately for each sub-sample of firm age. In all four of the sub-samples, EPS, operating cash flows per share, sales per share, and percentage of shares offered came out as significant factors. In the case of the oldest firms, only EPS and operating cash flows per share emerged as robust factors for IPOs' valuation. This finding implies that strong financial history leads to lower ex-ante uncertainty in terms of the new issue.

TABLE 7

RELATION BETWEEN FIRM AGE AND ISSUE PRICE

Issuer Age (years)	Minimum	Maximum	Mean	Std. Deviation
Less than 1	1	500.000	73.506	66.844
1 ≤ Age ≤ 3	1	750.000	77.874	82.017
3 < Age ≤ 5	1	678.150	78.409	79.334
Age > 5	1	730.000	98.402	109.049

Note: This table exhibits nexus between Firm Age and Issue Price of IPOs of a selected sample of 2185 IPOs

TABLE 8

NEXUS BETWEEN FIRM AGE AND ISSUE PRICE OF IPOS

	Firm Age-I	Firm Age-II	Firm Age-III	Firm Age-IV	Overall
Earnings/ Shares	1.501**	1.762**	1.683**	1.362	1.755**
	(14.44)	(15.43)	(14.31)	(8.79)**	(28.29)
Cash flow/Shares	1.250**	1.287**	1.514**	2.540	1.402**
	(10.94)	(9.53)	(11.11)	(11.72)**	(18.93)
Sale/Shares	1.235**	0.846**	0.848**	0.183	0.713**
	(13.98)	(8.45)	(8.00)	(0.69)	(12.58)
Book Value/Shares	0.088	–0.033	0.052	–0.120	0.029
	(1.23)	(0.38)	(0.59)	(0.76)	(0.60)
Sale Growth	0.061	0.030	0.011	0.164	0.038
	(1.22)	(0.52)	(0.18)	(1.55)	(1.15)
Profit Growth	–0.075	–0.015	–0.006	–0.274	–0.066
	(0.91)	(0.15)	(0.05)	(1.55)	(1.17)

Table 8 (cont.)

	Firm Age-I	Firm Age-II	Firm Age-III	Firm Age-IV	Overall
% Share Offered	–0.181*	–0.457**	–0.755**	–0.100	-0.771**
	(2.33)	(5.37)	(9.22)	(0.57)	(17.83)
_cons	–4.681**	–3.077**	–2.637**	–3.700**	–1.913**
	(10.65)	(6.38)	(5.59)	(3.43)	(7.67)
R^2	0.79	0.70	0.68	0.59	0.67
N	621	639	661	264	2,185

Note: This table displays estimated coefficients is each of the four sub-samples. * < 0.05; ** $p < 0.01$ represent significance level at the 1, and 5% respectively.

4.6. Relationship between firm size and the issue price

The signaling hypothesis is based on the theory that large firms differentiate their status in the market from small firms by issuing IPOs with high offer prices (Badru & Ahmad-Zaluki, 2018; Wahid, Khan, *et al.*, 2019). To the extent that theory is valid, it would help to explain the statistical incidence underpricing. We tested this proposition by partitioning the sample into firm size quartiles based on the total assets of the firm. A large variation of firm size ensures that diversified IPOs are included in the sample.

Partitioning the sample into quartiles reveals a systematic relationship between firm size and the offer price. We found that as the firm sizes increase, the offer price tends to increase. This effect is displayed in Table 9. An alternative analytical method is displayed in Table 10. For each quartile formed based on the total assets of the firm, we applied OLS to find out the factors affecting the valuation of IPOs in the AIM. In small size and medium-size firms, EPS, operating cash flow per share, sales revenue per share, sale growth, and percentage of shares offered are systematically related to the pricing of IPOs in the AIM.

TABLE 9
FIRM SIZE AND ISSUE PRICE

Firm Size	Minimum	Maximum	Mean	Std. Deviation
Firm Size < 7.621 (£m)	1	478.000	28.540	43.686
7.621 (£m) ≤ Firm size ≤ 19 (£m)	1	350.000	60.663	52.503
19 (£m) < Firm size ≤ 47.170 (£m)	1	550.340	86.151	56.506
Firm Size > 47.170 (£m)	1	750.000	141.836	108.518

Note: This table displays the relationship between Firm Size and Issue Price of IPOs of a sample of 2185 IPOs. It contains 1773 local IPOs and 412 Cross-listed IPOs issued and placed on the AIM from 1995 to 2021.

TABLE 10
NEXUS BETWEEN FIRM SIZE AND ISSUE PRICE

	Firm Size I	Firm Size II	Firm Size III	Firm Size IV	Overall
Earnings/ Shares	1.628**	1.384**	1.101**	1.121**	1.755**
	(11.51)	(13.57)	(11.46)	(11.02)	(28.29)
Cash flow/Shares	1.402**	0.982**	0.868**	0.792**	1.402**
	(9.43)	(8.15)	(7.57)	(6.17)	(18.93)
Sale/Shares	1.107**	1.086**	0.534**	0.152	0.713**
	(8.01)	(11.55)	(6.40)	(1.87)	(12.58)
Book Value/Shares	0.066	0.062	0.001	–0.086	0.029
	(0.62)	(0.80)	(0.02)	(1.16)	(0.60)
Sale Growth	0.183**	–0.012	0.028	–0.033	0.038
	(2.62)	(0.22)	(0.56)	(0.67)	(1.15)
Profit Growth	–0.077	–0.083	–0.085	0.041	–0.066
	(0.64)	(0.90)	(1.02)	(0.49)	(1.17)
% Share Offered	–0.502**	–0.362**	–0.348**	–0.504**	–0.771**
	(4.82)	(5.04)	(5.24)	(7.32)	(17.83)
constant	–4.265**	–2.496**	–0.059	1.911**	–1.913**
	(7.24)	(6.05)	(0.16)	(4.79)	(7.67)
R^2	0.64	0.71	0.52	0.33	0.67
N	547	546	546	546	2,185

Note: This table displays the coefficients in each sub-sample of firm size *i.e.* Firm size-I (< = 7.621 (£m), firm size-II (> 7.621 (£m) and < = 19 (£m), firm size-III (> 19 (£m) and < = 47.170 (£m) and firm size IV (> 47.170 (£m) of overall sample of 2185 IPOs placed on the AIM during 1995 to 2021. * <0.05; ** p<0.01 represent significance level at the 1, and 5% respectively.

4.7. The Offer size and the issue prices of Initial Public Offerings

If a larger ex-ante uncertainty is associated with larger issue sizes, that would help to explain why the mispricing of large issues is generally of greater magnitude than the mispricing of smaller issues (Rathnayake *et al.*, 2019; Wahid *et al.*, 2020). We tested this proposition by partitioning the sample into quartiles defined by the size of the offer measured by gross proceeds. The lowest and the highest offer size £1 million and £750 million display large sample-variability due to the heterogeneity of the IPOs *i.e.* local and offshore listed firms. The relationship between the offer size and the issue price is shown in Table 11. The results of the OLS are displayed in Table 12. The findings suggest that there is no significant systematic relationship between the offer price and the size of the offering.

TABLE 11
OFFER SIZE AND ISSUE PRICE

Offer Size	Minimum	Maximum	Mean	Stand. Dev.
Offer Size < = 2 (£m)	1.00	285.000	28.839	38.117
Offer Size > 2 (£m) and < = 5.010 (£m)	1.00	400.000	57.692	53.984
Offer Size > 5.010 (£m) and < = 15 (£m)	5.00	730.000	96.325	69.856
Offer Size > 15 (£m)	1.10	750.000	135.917	104.969

Note: This table exhibits nexus between Offer Size and Issue Price of IPOs of a selected sample of 2185 IPOs that consists of 1773 local IPOs and 412 Cross-listed IPOs issued and placed on the AIM during 1995 to 2021.

TABLE 12
NEXUS BETWEEN OFFER SIZE AND ISSUE PRICE

	Offer SizeI	Offer Size II	Offer Size III	Offer Size IV	Overall
Earnings/ Shares	1.381**	1.448**	1.224**	1.076**	1.755**
	(10.51)	(12.29)	(13.36)	(9.86)	(28.29)
Cash flow/Shares	1.593**	1.110**	0.855**	0.623**	1.402**
	(11.39)	(8.18)	(7.70)	(4.69)	(18.93)
Sale/Shares	1.133**	1.075**	0.350**	0.269**	0.713**
	(8.46)	(10.24)	(4.26)	(3.24)	(12.58)
Book Value/Shares	0.061	–0.011	–0.053	–0.008	0.029
	(0.60)	(0.12)	(0.76)	(0.10)	(0.60)
Sale Growth	0.141*	0.042	–0.008	–0.026	0.038
	(2.03)	(0.70)	(0.17)	(0.50)	(1.15)
Profit Growth	–0.094	–0.122	0.011	0.029	–0.066
	(0.80)	(1.19)	(0.13)	(0.33)	(1.17)
% Share Offered	–0.481**	–0.699**	–0.418**	–0.455**	–0.771**
	(5.07)	(8.49)	(6.59)	(6.47)	(17.83)
_cons	–4.305**	–1.582**	0.465	1.794**	–1.913**
	(7.83)	(3.47)	(1.24)	(4.43)	(7.67)
R^2	0.65	0.68	0.51	0.30	0.67
N	555	538	562	530	2,185

Note: This table displays the coefficient of the issuer's offer size * < 0.05; ** p < 0.01 represent significance level at the 1, and 5% respectively. It contains 1773 local IPOs and 412 Cross-listed IPOs issued and placed on the AIM from 1995 to 2021.

4.8. The nationality of the IPO and the issue price

The signaling hypothesis proposes that high-quality firms intentionally set the IPO offer price high to differentiate their offering from low-quality firms (Alim & Ramakrishnan, 2017; Badru & Ahmad-Zaluki, 2018).We tested this proposition by partitioning our sample into two sub-samples: they are the sub-sample of 1,773 local IPOs and the sub-sample of 412 Cross-listed IPOs. We presume the cross-listed IPOs are high-quality firms because only those kinds of domestic firms can elect offshore listing. The firms in high-quality sub-sample are well established and have sound financial histories. Descriptive statistics in Table (13) indicate that the offer prices of cross-listed IPOs are an average of £99.640. That statistic is significantly higher than the average offer price of local IPOs.

The results of OLS are shown in Table 14, the statistical findings indicate that for local IPOs, EPS, operating cash flow per share, sales per share, and percentage of shares offered have statistically significant explanatory power. In the sub-sample of cross-listed IPOs, only EPS, and operating cash flow per share play a significant role in determining the prices of cross-listed IPOs. In summary, firm nationality or duality is powerfully influential in the determination of the offer price. The signaling hypothesis implies that cross-listed IPOs might set high offer prices to attract the attention of the local investors. Alternatively, it is also consistent with the ex-ante uncertainty hypothesis which is related to information asymmetry. The information asymmetry hypothesis proposes that the prices of cross-listed IPOs are higher than the offer prices in single-market IPOs because of strong financial track records of cross-listed IPOs in their parental market and full access to that information by underwriters. As a consequence, underwriters have more guidance and useful information for the valuation of cross-listed IPOs. This leads to more clarity and conciseness about the pricing of offshore listings.

TABLE 13

NATIONALITY OF IPOS AND ISSUE PRICE

Class	Minimum	Maximum	Mean	Std. Deviation
Local Incorporated Firm's IPOs	1	750	74.542	72.288
Cross-listed IPOs	1	730	99.640	110.212

Note: This table exhibits nexus between the nationality of IPOs and Issue Price of IPOs of a selected sample of 2185 IPOs that consists of 1773 local IPOs and 412 Cross-listed IPOs issued and placed on the AIM during 1995 to 2021.

TABLE 14
NEXUS BETWEEN NATIONALITY OF IPOS AND ISSUE PRICE

	Local IPOs	Cross-listed IPOs	Overall
Earnings/ Shares	1.559**	1.377**	1.755**
	(24.16)	(11.84)	(28.29)
Cash flow/Shares	1.347**	2.496**	1.402**
	(19.01)	(14.57)	(18.93)
Sale/Shares	1.218**	0.249	0.713**
	(22.28)	(1.21)	(12.58)
Book Value/Shares	–0.012	0.031	0.029
	(0.28)	(0.26)	(0.60)
Sale Growth	0.052	0.100	0.038
	(1.72)	(1.21)	(1.15)
Profit Growth	–0.026	–0.252	–0.066
	(0.51)	(1.79)	(1.17)
% Share Offered	–0.059	–0.189	–0.771**
	(1.19)	(1.38)	(17.83)
_cons	–5.341**	–3.772**	–1.913**
	(19.48)	(4.56)	(7.67)
R^2	0.78	0.58	0.67
N	1,773	412	2,185

Note: This table exhibits beta coefficient based on nationality selected sample of 2185 IPOs that consists of 1773 local IPOs and 412 Cross-listed IPOs issued and placed on the AIM from 1995 to 2021.

4.9. Discussion and analysis

Our findings are two-folds: First, EPS, operating cash flow per share and sales revenue per share are all significantly and positively correlated with the value of IPOs in the AIM. We also found that the percentage of shares issued is negative and significantly correlated with the price variability of IPOs. Second, the age of the firm and financial history are systematically related to the price of IPOs. Firm size and nationality are strongly correlated with the price variability of IPOs. The variables capturing underwriter's prestige, market conditions, and offer size are not significantly related to the variation in the pricing of IPOs. These findings suggest the importance of ex-ante expectations and signaling in the price behavior of IPOs listed on the AIM. The empirical evidence could not explain the role of the window of opportunity hypothesis, underwriter's reputation hypothesis, and information asymmetric hypothesis in the pricing of IPOs.

Our results further indicate that the valuation of IPOs in the AIM follows the conventional theory regarding valuation: positive earning and positive cash flows along with reasonable sales and profit growth. We found that most of the IPOs listed on the AIM are small startups, designed to exploit innovative ideas. The inception of many of these start-ups in universities contributes to the growth and survival of these firms (Amini & Keasey, 2013). Our findings are consistent with the research findings of those researchers who infer a higher probability of success of small IPOs in the AIM as compared to the IPOs of large-sized firms. The rationale behind this evidence is that new startups which are based on innovative ideas are especially likely to prosper when public shareholders consisting of local businesses are involved in generating the financial synergies. These findings suggest that prospective investors can value the IPOs based on financial performance and the position of the firm in its market before going public in the AIM.

5. Conclusions

We examined the relationship between financial indicators of performance for firms before the launch of their IPOs and the offer price of those IPOs. We used a sample of 2,185 IPOs consisting of 1,773 local IPOs and 412 Cross-listed IPOs. All the firms in the sample were issued and listed on the AIM from1995 to 2020.Our research addressed the task of identifying the set of explanatory variables that are the significant drivers of the value of the IPO prior to the offering. A secondary question we addressed is whether the value of the IPO is significantly correlated with the size of the firm, the age of the firm, market conditions, offer size, and classification of local and cross-listed firms as the control variables.

In previous studies, it is found that ex-ante uncertainty has greater penetration on the value of IPOs in the main markets, in our findings; the same patterns have been observed in AIM. Firm size, age of the firm prior to the offering, offer size, class of IPOs whether newly listed or cross-listed and the dummy variable that represents the 'hot' period for IPOs have significant contribution in the variation of the value of IPOs even having same accounting credentials prior to offering. We observe that the value of IPOs is varied for firms having different firm sizes, different offer sizes, firm age, and nationality. Similarly, accounting information specifically earning per share, sales per share, cash flow per share, and margin of share offered to the public significantly affect the value of IPOs in AIM. This depicts that IPOs' characteristics including the EPS, sale per share, cash flow per share, and ex-ante uncertainty play a vital role in defining the value of IPOs. Our findings support the view that the quality of financial statements helps reduce information asymmetries that affect IPO valuations in AIM. Higher the symmetric information, the higher the chances of defining the intrinsic valuation of IPOs. Specifically, our results point out that lessening the information gap between informed and uninformed investors leads to ease for underwriters in defining the value of IPOs.

Moreover, this study also suggests that investors working in AIM should keep the level of both local and cross-listed shareholding the same because it still has to respond to ups and downs of the home market as well as parental market dynamics which further leads to variation in the value of IPOs. Similarly, underwriters of local firms should also be aware that the competition and complexities in the primary market increases after cross-listing IPOs. Underwriters, thus, need to equip themselves with both the knowledge and the psychological preparation to deal with the complexities and frustrations associated with parental market dynamics of cross-border listed as well as AIM primary market. The findings of this study may be of interest to regulatory bodies and policymakers. The policymakers and regulatory bodies should be concerned about how they can both improve AIM regulatory framework to enhance the volume of the primary market and strengthen enforcement strategies so that both categories of IPOs would be valued fairly. In this study, we used only accounting information for the valuation of IPOs. Building on these findings, we propose that future research may be conducted to determine the value of IPOs using forecasted financial data through comparable firm methods. Secondly, a comparison between the value of IPOs quoted in AIM and the main market may also be made using accounting information and sensitivity analysis of various factors.

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Spillover effects of the US economic policy uncertainty in Latin America*

Efectos indirectos de la incertidumbre de la política económica de EE.UU.en América Latina

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Abstract

This paper is aimed at assessing the spillover effects of the US Economic Policy Uncertainty (EPU) in macroeconomic variables of major Latin American Countries (LAC): Mexico, Colombia, Brazil, and Chile. To do that, we estimate a set of two-country Structural Vector Autoregressive (SVAR) models for 1997-2019; each model includes the US and one of the LAC. We use the following variables: EPU indexes, exchange rates, consumer price indexes, industrial production (IP), and interest rates (IR) of the US and the studied LAC. The main finding is that positive shocks in the US EPU index lead to currency depreciation for all four LAC; the largest effect is for Mexico. Other statistically significant results are a brief and small positive impact on Colombia's IP and a positive impact on Mexico's IR. The remaining LAC's estimates are statistically insignificant. For this reason, we applied Rossi and Wang's (2019) robust Granger causality tests that considers structural breaks. Finally, the estimates before and after the 2008 financial crisis suggest that LAC became slightly more responsive to US EPU shocks after the crisis.

Key words: *Economic policy uncertainty, structural vector autoregressive, impulse response function, robust Granger-causality tests.*

JEL Classification: *F62, N16.*

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Resumen

Esta investigación tiene como objetivo evaluar los efectos indirectos de la incertidumbre de la política económica de EE.UU. en las variables macroeconómicas de los principales países de América Latina: México, Colombia, Brasil y Chile. Para ello, estimamos un conjunto de modelos de vectores autorregresivos estructurales de dos países para 1997-2019; cada modelo incluye a EE.UU. y un país de América Latina. Utilizamos las siguientes variables: índices la incertidumbre de la política económica, tipos de cambio, índices de precios al consumidor, producción industrial y tasas de interés de los Estados Unidos y países de América Latina estudiados. El principal hallazgo es que los choques positivos en el índice incertidumbre de la política económica de EE.UU. conducen a la depreciación de la moneda local en los cuatro países de ALC; el mayor efecto es para México. Otros resultados estadísticamente significativos son un impacto positivo breve y pequeño en la producción industrial de Colombia y un impacto positivo en la tasa de interés de México. Las estimaciones restantes para América Latina son estadísticamente no significantes. Por esta razón, aplicamos las pruebas robustas de causalidad de Granger propuestas por Rossi y Wang (2019) que consideran cambios estructurales. Por último, las estimaciones antes y después de la crisis financiera de 2008 sugieren que América Latina se tornó un poco más sensible a los choques de incertidumbre de la política económica estadounidense después de la crisis.

Palabras clave: *Incertidumbre de la política económica, vectores autorregresivo estructurales, función impulso respuesta, pruebas robustas de causalidad de Granger.*

Clasificación JEL: *F62, N16.*

1. Introduction

The influence of the US economy policy on advanced and emerging countries has been widely studied in the specialized literature. In particular, there are several papers dealing with the impact of the uncertainty of the US economic policy on economic and financial variables on several countries and regions. In this regard, Gupta, Olasehinde-Williams and Wohar (2020) assess the impact of Economic Policy Uncertainty (EPU) shocks on a panel of 50 advanced and emerging market economies These authors find that for advanced economies the exchange rate regime and financial vulnerability account for a large portion of the contraction in activity. In emerging economies the responses do not depend on the exchange rate regime, but the responses become larger when trade openness is high and weakness in the financial system. Also, Kido (2016) analyzes spillover effects of shocks the US economic policy uncertainty on real effective exchange rates of several countries duing the period 2000-2014 by using a Dynamic Conditional Correlation Generalized Autoregressive Conditional Heteroskedasticity (DCC-GARCH) model.

the other hand, Zouhair, Lanouar, and Ajmi, (2013) examine the volatility spillovers in stock markets, securitized real estate, bond markets, currency markets, and economic policy uncertainty spillovers across 7 countries. The authors' empirical findings are that spillovers are important and account for, respectively, about 72% and 50% of the dynamics of financial market stress and economic policy uncertainty, respectively, across the 7 economies examined. Also, Liow, Liao, and Huang (2018) find empirical evidence that policy uncertainty spillovers lead financial market stress spillovers in a multi-country context. In other words, changes in international economic policy uncertainty spillovers may be a predictor of changes in international financial market risk spillovers in the short run. Finally, Colombo's (2013), and Bernal, Gnabo, and Guilmin (2016) assess the impact of economic policy uncertainty on risk spillovers within the Eurozone.

Regarding the impact of the US economic policy uncertainty in the country itself, Baker, Bloom, and Davis (2016) find that the US EPU is associated with greater stock price volatility and reduced investment and employment in policy-sensitive sectors like defense, health care, finance, and infrastructure construction. Also, Bloom (2009) simulate the impact of a large EPU shock in the US and find that it generates a rapid drop, rebound, and overshoot in employment, output, and productivity growth. Hiring and investment rates fall dramatically in the 4 months after the shock because higher uncertainty increases the real-option value to waiting, so firms scale back their plans. Finally, Istiak and Serletis (2020) find that commercial bank leverage rises when geopolitical risk and macroeconomic, policy, and equity uncertainty increase. Moreover, the authors find that the leverage of broker-dealers and shadow banks declines when Chicago risk and macroeconomic, policy, financial, and equity uncertainty increase.

Concerning the convergence of uncertainties across the World, Christou, Gozgor, Gupta, Keung, and Lau (2019) analyze the convergence of a news-based measure of uncertainty across 143 countries in the period of 1996-2018. The authors use a panel data-based unit root test to the ratio of the uncertainty of individual countries relative to that of global uncertainty. These authors find empirical evidence of convergence and hence, the spillover of uncertainty across the economies of the world. Moreover, Gabauer and Gupta (2018) focuss on the transmission mechanism of country-specific and international economic uncertainty spillovers by using a Time-Varying Parameter Vector Autoregressions (TVP-VAR) approach.

For Latin American Countries (LAC), the US influence tends to be significant due to both financial and trade ties in the region. In this regard, Alam and Istiak (2020) examine the impact of US policy uncertainty on Mexico by using a Structural Vector Autoregressive (SVAR) model with linear and nonlinear tests. These authors show that an increase in the US economic policy uncertainty leads to a fall in Mexican output (industrial production), price level and interest rate. It is worth mentioning here that there is a lack of research into the possible effects on uncertainty of the US economic policy on macroeconomic aggregates in broader regions of Latin America.

This research examines the spillover effects of US economic policy uncertainty on macroeconomic aggregates in Latin American Countries by estimating a SVAR model during the period 1997-2019. Most of the empirical analysis is mainly based on Impulse Response Function (IRF) estimations from two-country SVAR[1] by using information of the EPU indexes and macroeconomic variables for the US, Mexico, Colombia, Brazil, and Chile. The IRF will be used to quantify the magnitude and persistence of a US economic policy uncertainty shock on LAC for the entire period, 1997-2019, and for specific periods before and after the 2008 financial crisis.

The rest of this paper is organized as follows: section 2 describes the nature of data and presents a summary of the descriptive statistics; section 3 depicts the SVAR specification and shows the obtained empirical results; section 4 presents the findings about Granger-causality tests from the EPU index of the US toward the currency depreciation for all four LAC; finally Section 5 concludes.

2. Nature of Data and Summary Statistics

Our dataset consists of monthly time series of EPU indexes and macroeconomic data for the US and the following LAC: Mexico, Colombia, Brazil, and Chile. The entire period of study is from January 1997 to December 2019. All series are expressed in growth rates.

EPU indexes data for all the countries were collected from the webpage of Economic Policy Uncertainty Index.[2] We also include the following macroeconomic fundamentals for all the countries: inflation (CPI), growth rate of industrial production (IP), short interest rate (IR), and the rate of depreciation of the exchange rates (ER). Macroeconomic data comes from different sources; some of them come from the Organization for Economic Cooperation and Development (OECD) webpage, other from the Federal Reserve Economic Data (FRED) database, and other more from National Central Banks of the corresponding LAC. See Appendix A for a full description of the data sources.

Figure 1 shows the EPU indexes for the US and each one of the four LAC. The left-hand side presents the EPU indexes in levels and the right-hand side presents the EPU indexes in growth rates. The right-hand graphs suggest the EPU indexes in growth rates are indeed stationary for the US and each of the four LAC. In terms of the levels, Figure 1 also shows that Mexico and the US tend to follow each other closely before the 2008 crisis, but not so much after the crisis. In fact, Mexico's index tends to be below that of the US and with

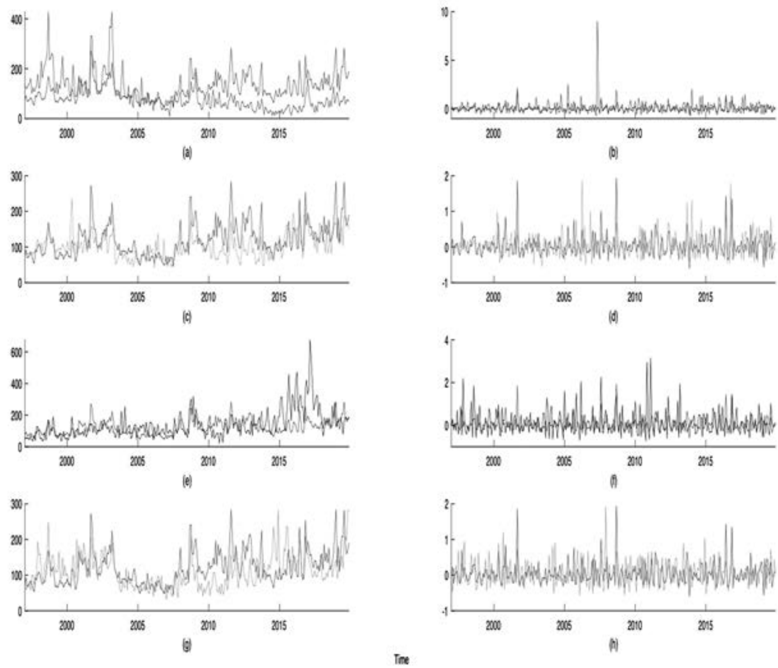
1 For measuring uncertainty see Jurado, Ludvigson, and Ng (2015) and Mumtaz and Theodoridis (2018).

2 Available in http://www.policyuncertainty.com/

lower volatility after the 2008 crisis. The rest of LAC tend to follow the US more closely.

FIGURE 1

EPU INDEXES TIME SERIES PLOT IN LEVELS (LEFT) AND GROWTH RATES (RIGHT). THE EPU INDEX FOR THE US PLOTTED AGAINST MEXICO IS IDENTIFIED WITH HIGHER PEAKS (A,B), COLOMBIA IS IDENTIFIED WITH THE LIGHT LINE (C,D), BRAZIL IS IDENTIFIED WITH HIGHER PEAKS (E,F), AND CHILE IS IDENTIFIED WITH THE LIGHT LINE (G,H).



Source: Authors' own elaboration based on the Bayesian Estimation, Analysis and Regression Toolbox (BEAR) from the European Central Bank.

Table 1 shows the descriptive statistics and the unit root test for all series under study. Results from the Residual Augmented Least Squares (RALS) test from Im, Lee, and Tieslau (2014) does not require a specific density function for the error term are reported in the last column of Table 1. For the RALS test, the null hypothesis is that there exists a unit root. The number of lags was selected using the sequential analysis proposed by Im *et al*. (2014). The null hypothesis

TABLE 1
TIME SERIES DESCRIPTIVE STATISTICS (N = 275)

Country	Variable	Mean	Median	Min	Max	Variance	SD	Skewness	Kurtosis	RALS
US	CPI	–0.041	–0.003	–11.054	6.914	0.920	0.959	–5.167	79.819	–11.239
	IP	0.001	0.001	–0.035	0.024	0.000	0.007	–1.174	8.351	–5.319
	IR	0.003	0.000	–0.454	0.800	0.014	0.117	0.997	13.045	–4.158
	EPU	0.043	–0.018	–0.601	1.935	0.103	0.321	2.207	11.930	–6.306
Mexico (MEX)	CPI	–0.005	–0.010	–0.249	0.404	0.006	0.075	0.738	6.472	–6.153
	IP	0.002	0.002	–0.053	0.049	0.000	0.012	–0.281	6.380	–7.879
	IR	–0.003	0.000	–0.225	0.484	0.004	0.062	1.488	17.532	–6.887
	ER	0.004	0.000	–0.084	0.188	0.001	0.025	1.588	12.957	–13.175
	EPU	0.119	–0.002	–0.742	9.015	0.534	0.731	6.980	81.660	–19.543
Colombia (COL)	CPI	–0.004	–0.004	–0.201	0.233	0.004	0.067	–0.097	3.507	–4.369
	IP	0.002	0.004	–0.098	0.158	0.001	0.031	0.599	7.692	–5.671
	IR	–0.005	–0.004	–0.218	0.154	0.002	0.043	–0.932	7.848	–4.859
	ER	0.005	0.001	–0.073	0.120	0.001	0.030	0.838	4.776	–12.322
	EPU	0.049	0.011	–0.548	1.869	0.116	0.340	1.716	8.631	–8.647
Brazil (BRA)	CPI	0.001	–0.003	–0.277	0.538	0.009	0.094	1.465	9.017	–4.076
	IP	0.001	0.002	–0.120	0.141	0.000	0.020	–0.511	18.475	–22.638
	IR	0.000	–0.007	–0.310	0.970	0.014	0.116	3.498	26.979	–4.265
	ER	0.006	0.004	–0.107	0.274	0.002	0.043	1.996	12.556	–4.601
	EPU	0.132	0.041	–0.736	3.155	0.349	0.591	1.728	7.700	–8.412
Chile (CHI)	CPI	0.296	–0.007	–3.160	83.725	25.621	5.062	16.369	270.336	–160.178
	IP	0.002	0.002	–0.184	0.223	0.001	0.026	0.960	31.599	–4.162
	IR	0.004	0.000	–0.503	1.880	0.031	0.175	5.251	54.317	–5.224
	ER	0.002	0.000	–0.073	0.176	0.001	0.026	1.179	9.911	–5.002
	EPU	0.060	0.003	–0.572	1.921	0.131	0.362	1.112	5.340	–5.322

Source: Authors' own elaboration

is rejected for all variables using a 99% confidence level. The results suggest the series in growth rates can be considered stationary.

3. SVAR Specification and Empirical Results

In this section, we report the empirical results to attempt to quantify the impact of an EPU shock from the US on the macroeconomic indicators of the LAC under study. Following Colombo (2013), we estimate a set of two-country structural VAR models. Each VAR model includes the US and one of the LAC from the country sample. The IRF are estimated and account for the magnitude of the impact, if any, and its dynamics.[3]

Consistent with the main objective of this paper, countries are ordered assuming that the variables in the US are "more" exogenous than the variables from LAC. Within each economy and consistent with the literature, we order the macroeconomic variables from "more" to "less" exogenous. Macroeconomic variables for each country are ordered as follows:

Mexico: $y_t = [EPU^{USA}\ CPI^{MEX}\ IP^{MEX}\ IR^{MEX}\ ER^{MEX}\ EPU^{MEX}]'$,

Colombia: $y_t = [EPU^{USA}\ CPI^{COL}\ IP^{COL}\ IR^{COL}\ ER^{COL}\ EPU^{COL}]'$,

Brazil: $y_t = [EPU^{USA}\ CPI^{BRA}\ IP^{BRA}\ IR^{BRA}\ ER^{BRA}\ EPU^{BRA}]'$,

Chile: $y_t = [EPU^{USA}\ CPI^{CHI}\ IP^{CHI}\ IR^{CHI}\ ER^{CHI}\ EPU^{CHI}]'$.

In this research IRF identification is obtained via Cholesky decomposition of the variance-covariance matrix of the reduced-form VAR estimates. SVAR models have a lag length of 1 according to Bayesian Information Criterion (BIC).

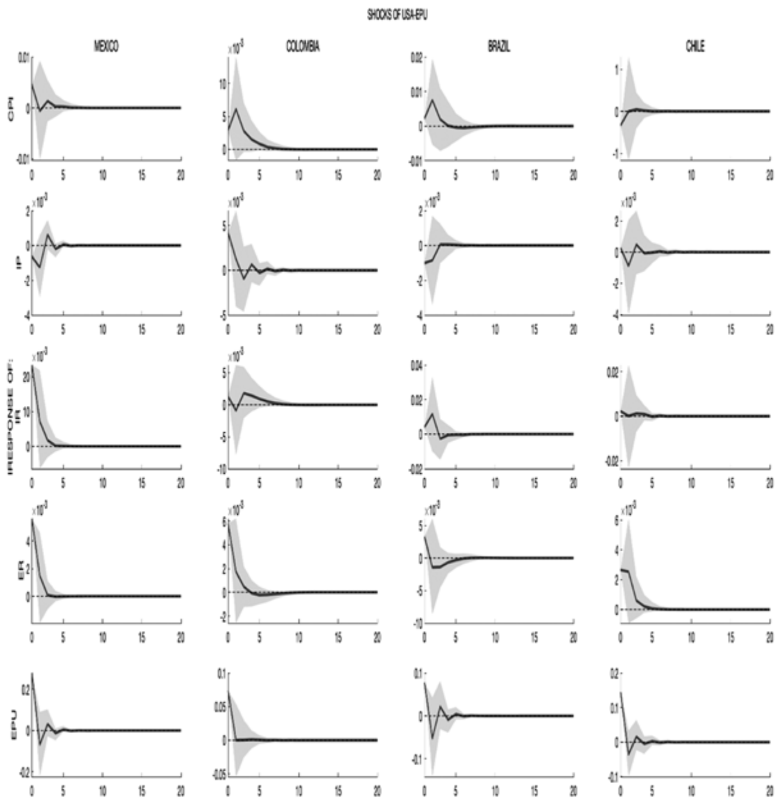
Figure 2 shows the IRF estimates from a shock of one standard deviation in the US EPU index on macroeconomic variables for each LAC obtained from a two-country SVAR model.[4] In particular, Figure 2 shows that a positive shock

3 For robustness, we applied two main specifications for VAR models: a Bayesian VAR model with stochastic volatility (BVARSV) and time-varying parameters Bayesian VAR (TV-BVAR). It has been documented extensively in the literature that stochastic volatility is important in fitting the dynamics of macro/finance variables as the ones analyzed in this research. The authors estimated the models by using Bayesian Estimation, Analysis and Regression Toolbox (BEAR) from the European Central Bank. For the structural VAR model specification, we follow Dieppe, Legrand, and Van Roye (2018). For a full description of the SVAR, BVARSV and TV-BVAR models and estimation procedures see Cogley and Sargent (2005), and Dieppe *et al.* (2018).

4 We carry out the estimation exercise using BVARSV and TV-BVAR models, but all estimates were statistically insignificant. The results of BVARSV and TV-BVAR are available upon request. We also replied the same exercise with the variables of the US as in Colombo (2013) to LAC.

to the US EPU index results in statistically significant currency depreciation for each LAC, and the magnitude of the impact is the largest for Mexico, followed by Colombia, Brazil, and Chile. This suggests the impact of an EPU index shock on LAC currencies is inversely related to the geographical distance from the US to the specific LAC.

FIGURE 2
TWO-COUNTRY SVAR MODELS
(US EPU SHOCK TO LAC)

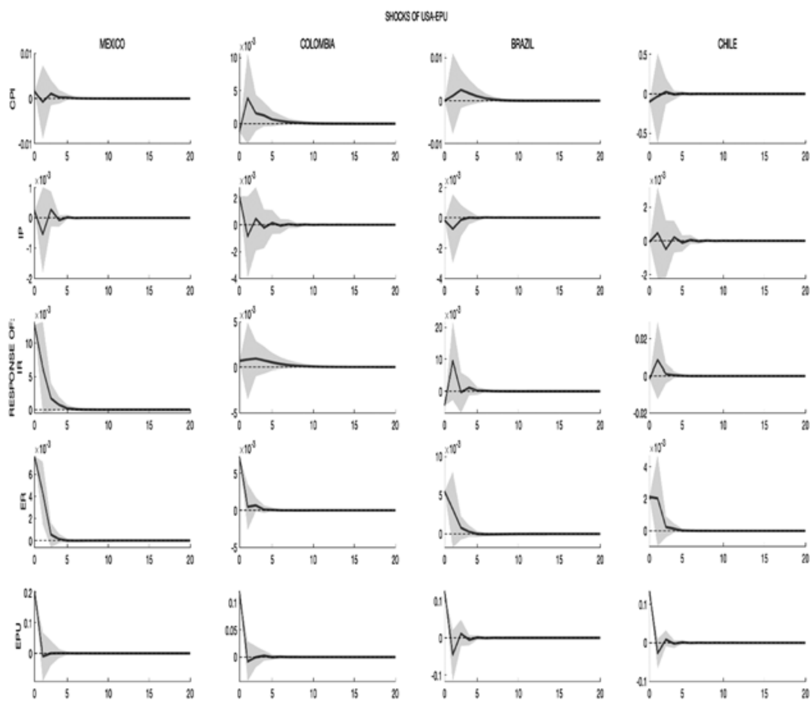


Source: Authors' own elaboration based on the sample and using the Bayesian Estimation, Analysis and Regression Toolbox (BEAR) from the European Central Bank.

In terms of the EPU index from each LAC, the impacts from a shock to the EPU index from the US are positive and statistically significant for each country, and the effects are larger for Mexico. The estimates suggest that economic policy uncertainty from the US has a direct and strong impact on economic policy uncertainty of LAC and this impact is absorbed almost entirely by currency depreciation. Other statistically significant results are: a brief and small positive impact on Colombia's IP and a positive impact on Mexico's IR. These estimates suggests that the geographical proximity to the U.S. makes Mexico and Colombia more susceptible to economic policy uncertainty shocks from the US. The IRF estimates for all other macroeconomic variables are not statistically significant.

Figure 3 and 4 present the IRF estimates from a shock of one standard deviation in the EPU index from the US on macroeconomic variables for LAC before and after the 2008 financial crisis, respectively.[5] The periods considered are for January 1997-December 2007 and for January 2008-December 2019.

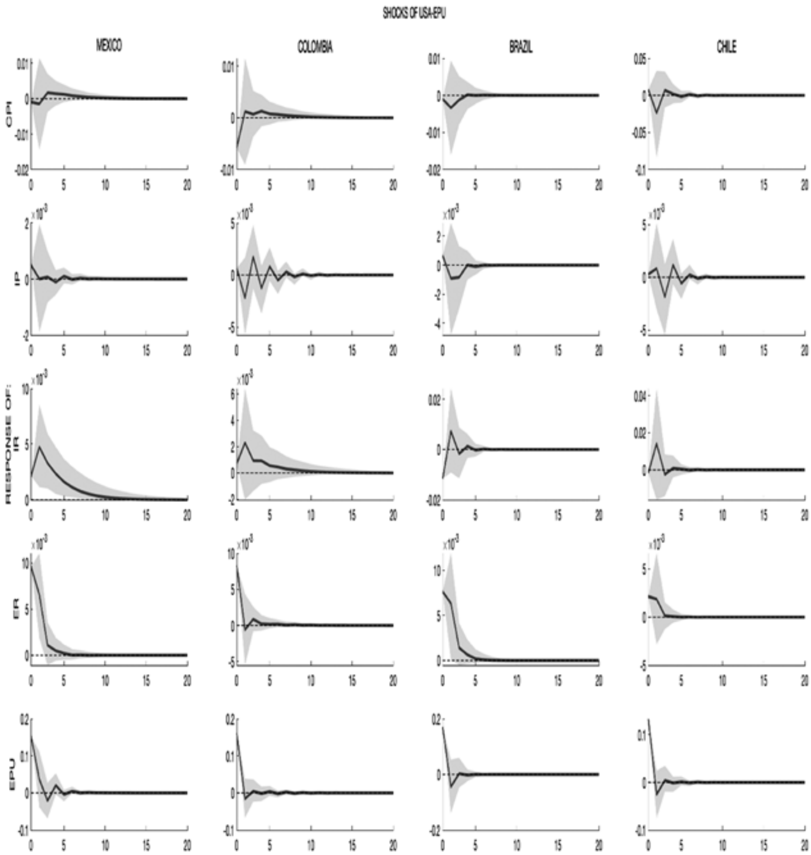
FIGURE 3
BEFORE THE 2018 CRISIS: TWO-COUNTRY SVAR MODELS
(US EPU SHOCK TO LAC)



Source: Authors' own elaboration based on the sample and using the Bayesian Estimation, Analysis and Regression Toolbox (BEAR) from the European Central Bank.

5 Following the idea of Alam and Istiak (2019) that made a split to the Mexican case.

FIGURE 4
AFTER THE 2018 CRISIS: TWO-COUNTRY SVAR MODELS
(US EPU SHOCK TO LAC)



Source: Authors' own elaboration based on the sample and using the Bayesian Estimation, Analysis and Regression Toolbox (BEAR) from the European Central Bank.

IRF estimates for the period before the 2008 financial crisis show that a shock to the EPU index from the US continues to have a statistically significant impact on the EPU indexes for each LAC, and the impact on Mexico's EPU continues to be larger than those for the other LAC. Similarly, and in terms of exchange rates, the impact of a shock to the EPU index from the US show a depreciation effect on the currencies for each LAC, and the impact is the largest for Mexico. Other positive and statistically significant impacts are for Mexico's IR and Colombia's IP.

IRF estimates for the period after the 2008 economic crisis continue to show a statistically significant impact from a US shock in the EPU index on the corresponding currency depreciation and EPU indexes for each LAC, and the largest of these impacts is observed for Mexico. For the most part, the IRF estimates for LAC show that the impact from a shock to the EPU index from the US resulted in a slightly larger impact on macroeconomic variables when compared to the estimates before the 2008 economic crisis. One potential explanation is that, given that the US was the epicenter of the 2008 economic crisis, countries started paying closer attention to the economic policy uncertainty coming from the US. In general terms, these results suggest that LAC became more responsive to EPU index shocks from the US after the 2008 economic crisis.

4. Granger Causality Tests

In order to corroborate the structural relationship between US EPU index and currency depreciation for all four LAC, Granger Causality (GC) test is applied to analyze the dynamic macroeconomic relationship between these variables.[6] Two types of GC tests are implemented, the classical test provided by Granger (2001) that assumes constant parameters and the robust GC that takes into account structural breaks. The classical GC test results may be not reliable and not significant if there are structural breaks in the period (Yang, 2015; Zeileis, Leisch, Kleiber, & Hornik, 2005). For this reason, we also applied a robust GC test that considers structural breaks due to Rossi and Wang, (2019) throughout the causal relationship among different variables.

Table 2 presents the GC tests results that correspond to the entire period. The second column corresponds to the classical test and the remaining columns correspond to different versions of the robust GC. The corresponding statistics are: ExpW (the exponential Wald test), MeanW (the mean Wald test), Nyblom (the Nyblom test), and QLR (the Quant likelihood-ratio test or SupLR test).[7] The third column (ExpW, the exponential Wald test) provides the most robust test and our preferred specification of the GC test. The results provide evidence of a Granger-Causality relationship from the EPU index of the US toward the currency depreciation for Mexico, Colombia, and Chile. The results for Brazil are not statistically significant.

Appendix B presents graphs of the Wald statistic for structural breaks in the GC for the entire period and for the periods that corresponds to before and after the 2008 financial crisis. Based on the specific dates from structural breaks, currency depreciation forecasts for all four LAC for 1 and 3 periods into the future were computed. Based on our preferred specification of the robust GC test, all estimates were statistically significant at the 5 % level for the entire period and for the period before and after the 2008 financial crisis.

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6 The usual caveat: This is not "causality" in common sense, this is predictive causality (Granger's causality or causality in the Granger sense), which tests how past values of one variable correlate with current values of another variable.

7 See Rossi (2005) for more details about these statistics.

TABLE 2
GRANGER-CAUSALITY TEST AND ROBUST GRANGER-CAUSALITY TESTS (1997M2-2019M12)

Variables	χ^2 Statistic (*p*-value)	ExpW Statistic (*p*-value)	MeanW Statistic (*p*-value)	Nyblom Statistic (*p*-value)	SupLR Statistic (*p*-value)
$EPU^{USA} \rightarrow CPI^{MEX}$	0.331	0.400	0.682	0.730	5.169
	(0.565)	(1.000)	(1.000)	(1.000)	(0.735)
$EPU^{USA} \rightarrow IP^{MEX}$	1.278	1.419	2.647	0.941	4.510
	(0.258)	(0.715)	(0.672)	(0.403)	(8.132)
$EPU^{USA} \rightarrow IR^{MEX}$	0.399	5.306	1.592	0.202	20.162
	(0.529)	(0.040)**	(0.889)	(1.000)	(0.000)**
$EPU^{USA} \rightarrow ER^{MEX}$	3.482	5.799	8.078	3.105	15.801
	(0.062)*	(0.027)**	(0.069)*	(0.031)**	(0.030)**
$EPU^{USA} \rightarrow EPU^{MEX}$	0.410	0.407	0.751	0.110	2.234
	(0.522)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow CPI^{COL}$	0.263	0.367	0.683	0.197	2.212
	(0.608)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow IP^{COL}$	0.048	2.094	3.708	0.763	6.417
	(0.827)	(0.487)	(0.466)	(0.495)	(0.578)
$EPU^{USA} \rightarrow IR^{COL}$	0.083	0.278	0.511	0.175	1.670
	(0.773)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow ER^{COL}$	2.770	4.704	7.847	0.321	13.061
	(0.096)*	(0.068)*	(0.078)*	(0.831)	(0.073)*
$EPU^{USA} \rightarrow EPU^{COL}$	2.154	1.494	2.764	1.003	5.274
	(0.142)	(0.688)	(0.647)	(0.376)	(0.717)
$EPU^{USA} \rightarrow CPI^{BRA}$	0.051	0.703)	1.257	0.139	2.778
	(0.821)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow IP^{BRA}$	0.301	0.504	0.874	0.363	2.980
	(0.583)	(1.000)	(1.000)	(0.792)	(1.000)
$EPU^{USA} \rightarrow IR^{BRA}$	1.087	3.316	2.629	0.595	13.713
	(0.297)	(0.204)	(0.676)	(0.607)	(0.060)*
$EPU^{USA} \rightarrow ER^{BRA}$	0.2358	1.162	1.769	0.391	4.616
	(0.878)	(0.808)	(0.855)	(0.768)	(0.802)
$EPU^{USA} \rightarrow EPU^{BRA}$	0.003	0.393	0.660	0.128	3.768
	(0.953)	(1.000)	(1.000)	(1.000)	(0.891)
$EPU^{USA} \rightarrow CPI^{CHI}$	0.015	0.017	0.021	0.002	0.028
	(0.901)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow IP^{CHI}$	0.217	0.696	1.312	0.426	2.474
	(0.641)	(1.000)	(1.000)	(0.737)	(1.000)
$EPU^{USA} \rightarrow IR^{CHI}$	0.017	0.231	0.417	0.082	1.847
	(0.895)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow ER^{CHI}$	0.280	13.906	12.885	2.995	34.078
	(0.596)	(0.000)**	(0.000)**	(0.034)**	(0.000)**
$EPU^{USA} \rightarrow EPU^{CHI}$	0.346	0.467	0.871	0.283	2.187
	(0.556)	(1.000)	(1.000)	(0.865)	(1.000)

Note: Second column shows the χ^2 statistic for the hypothesis null is that the VAR series x_t does not Granger cause series y_t. Third to sixth column shows ExpW statistic, MeanW, Nyblom and SupLR statistic for the hypothesis null is that the VAR series x_t does not Granger cause series y_t, assuming homoskesdatsic idiosyncratic shocks.** significant < 5% level and * significant < 10% level for all statistics.

Source: Prepared by the authors based on the sample and using the STATA Software.

Table 3 presents the GC tests results that correspond to period before the 2008 financial crisis. As before, the second column corresponds to the classical tests and the remaining columns correspond to different versions of the robust GC test. Based on our preferred specification of the robust test (third column, ExpW, the exponential Wald test), the results provide evidence only for Mexico of the Granger-Causality relationship between EPU index from the US and currency depreciation. For all other countries, the results are not statistically significant.

TABLE 3
GRANGER-CAUSALITY TEST AND ROBUST GRANGER-CAUSALITY TESTS (1997M2-2007M12)

Variables	Statistics (p-value)	ExpW Statistics (p-value)	MeanW Statistics (p-value)	Nyblom Statistics (p-value)	SupLR Statistics (p-value)
$EPU^{USA} \rightarrow CPI^{MEX}$	0.568	0.680	1.259	0.060	3.523
	(0.451)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow IP^{MEX}$	3.875	2.317	4.493	1.355	7.590
	(0.049)**	(0.422)	(0.344)	(0.249)	(0.438)
$EPU^{USA} \rightarrow IR^{MEX}$	0.011	5.447	4.824	0.549	15.859
	(0.915)	(0.036)**	(0.301)	(0.641)	(0.030)**
$EPU^{USA} \rightarrow ER^{MEX}$	0.328	13.939	14.533	2.399	35.010
	(0.567)	(0.000)**	(0.000)**	(0.069)*	(0.000)**
$EPU^{USA} \rightarrow EPU^{MEX}$	0.284	0.502	0.879	0.026	3.039
	(0.594)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow CPI^{COL}$	0.340	0.846	1.597	0.076	2.839
	(0.560)	(1.000)	(0.877)	(1.000)	(1.000)
$EPU^{USA} \rightarrow IP^{COL}$	1.487	1.221	2.324	0.881	4.631
	(0.223)	(0.788)	(0.741)	(0.433)	(0.801)
$EPU^{USA} \rightarrow IR^{COL}$	1.292	1.629	2.903	0206	6.672
	(0.256)	(0.640)	(0.617)	(1.000)	(0.547)
$EPU^{USA} \rightarrow ER^{COL}$	0.001	2.215	3.150	0.548	8.027
	(0.979)	(0.451)	(0.570)	(0.642)	(0.390)
$EPU^{USA} \rightarrow EPU^{COL}$	0.412	0.592	1.001	0.347	3.177
	(0.521)	(1.000)	(1.000)	(0.806)	(1.000)
$EPU^{USA} \rightarrow CPI^{BRA}$	0.607	0.473	0.875	0.407	2.513
	(0.436)	(1.000)	(1.000)	(0.754)	(1.000)
$EPU^{USA} \rightarrow IP^{BRA}$	0.349	3.756	3.431	0.664	12.818
	(0.555)	(0.141)	(0.516)	(0.560)	(0.081)*
$EPU^{USA} \rightarrow IR^{BRA}$	0.892	16.514	8.375	1.456	40.631
	(0.345)	(0.000)**	(0.060)*	(0.218)	(0.000)**
$EPU^{USA} \rightarrow ER^{BRA}$	0.619	1.337	2.392	0.059	4.917
	(0.431)	(0.745)	(0.726)	(1.000)	(0.763)
$EPU^{USA} \rightarrow EPU^{BRA}$	0.139	4.442	5.042	0.405	13.439
	(0.709)	(0.081)*	(0.276)	(0.756)	(0.064)*
$EPU^{USA} \rightarrow CPI^{CHI}$	0.001	0.122	0.003	0.001	0.017
	(0.982)	(1.000)	(1.000)	(1.000)	(1.000)
$EPU^{USA} \rightarrow IP^{CHI}$	1.882	1.531	2.918	1.162	6.422
	(0.170)	(0.675)	(0.614)	(0.311)	(0.577)
$EPU^{USA} \rightarrow IR^{CHI}$	0.486	0.991	1.236	0.230	6.379
	(0.486)	(0.864)	(1.000)	(1.000)	(0.583)
$EPU^{USA} \rightarrow ER^{CHI}$	0.818	0.994	1.815	0.839	4.845
	(0.366)	(0.863)	(0.847)	(0.454)	(0.773)
$EPU^{USA} \rightarrow EPU^{CHI}$	0.415	0.373	0.685	0.063	2.029
	(0.519)	(1.000)	(1.000)	(0.865)	(1.000)

Note: Second column shows the χ^2 statistic for the hypothesis null is that the VAR series x_t does not Granger cause series y_t. Third to sixth column shows ExpW statistic, MeanW, Nyblom and SupLR statistic for the hypothesis null is that the VAR series x_t does not Granger cause series y_t, assuming homoskedastic idiosyncratic shocks.** significant < 5% level and * significant < 10% level for all statistics.

Source: Prepared by the authors based on the sample and using the STATA Software.

Table 4 shows our final set of the GC tests results, which corresponds to period after the 2008 financial crisis. Based on almost all different specifications of the robust GC test, the results provide evidence of the Granger-Causality from EPU index of the US toward the currency depreciation for all four LAC. These results are in line with the perception that LAC became slightly more responsive to uncertainty shocks from the US after the 2008 financial crisis.

TABLE 4
GRANGER-CAUSALITY TEST AND ROBUST GRANGER-CAUSALITY TESTS
(2008M1-2019M12)

Variables	Statistics (*p*-value)	ExpW Statistics (*p*-value)	MeanW Statistics (*p*-value)	Nyblom Statistics (*p*-value)	SupLR Statistics (*p*-value)
$EPU^{USA} \rightarrow CPI^{MEX}$	0.0728 (0.787)	1.208 (0.792)	1.845 (0.841)	0.196 (1.000)	6.676 (0.546)
$EPU^{USA} \rightarrow IP^{MEX}$	0.083 (0.773)	0.252 (1.000)	0.400 (1.000)	0.070 (1.000)	1.766 (1.000)
$EPU^{USA} \rightarrow IR^{MEX}$	0.924 (0.336)	1.813 (0.577)	2.617 (0.678	0.345 (0.790)	7.610 (0.436)
$EPU^{USA} \rightarrow ER^{MEX}$	2.634 (0.105)	15.135 (0.000)**	15.664 (0.000)**	5.375 (0.000)**	35.649 (0.000)**
$EPU^{USA} \rightarrow EPU^{MEX}$	4.934 (0.026)*	2.992 (0.261)	5.865 (0.193)	2.561 (0.057)	7.880 (0.406)
$EPU^{USA} \rightarrow CPI^{COL}$	0.036 (0.849)	0.342 (1.000)	0.579 (1.000)	0.155 (1.000)	2.409 (1.000)
$EPU^{USA} \rightarrow IP^{COL}$	1.019 (0.313)	1.620 (0.643)	2.579 (0.686)	0.356 (0.799)	5.362 (0.706)
$EPU^{USA} \rightarrow IR^{COL}$	1.044 (0.307)	0.640 (1.000)	1.256 (1.000)	0.342 (0.811)	1.817 (1.000)
$EPU^{USA} \rightarrow ER^{COL}$	4.182 (0.041)**	6.008 (0.023)**	9.721 (0.030)**	0.893 (0.427)	16.394 (0.022)**
$EPU^{USA} \rightarrow EPU^{COL}$	1.346 (0.246)	1.818 (0.575)	3.012 (0.595)	1.215 (0.292)	6.160 (0.608)
$EPU^{USA} \rightarrow CPI^{BRA}$	0.0823 (0.774)	1.258 (0.773)	2.322 (0.742)	0.267 (0.880)	4.646 (0.799)
$EPU^{USA} \rightarrow IP^{BRA}$	0.013 (0.911)	0.938 (0.881)	1.699 (0.869)	0.338 (0.815)	4.507 (0.813)
$EPU^{USA} \rightarrow IR^{BRA}$	0.440 (0.507)	1.879 (0.555)	2.623 (0.676)	0.097 (1.000)	7.699 (0.426)
$EPU^{USA} \rightarrow ER^{BRA}$	0.886 (0.347)	7.375 (0.000)**	9.898 (0.028)**	2.998 (0.035)**	19.546 (0.000)**
$EPU^{USA} \rightarrow EPU^{BRA}$	0.001 (0.969)	0.333 (1.000)	0.571 (1.000)	0.102 (1.000)	1.780 (1.000)
$EPU^{USA} \rightarrow CPI^{CHI}$	2.528 (0.112)	1.426 (0.712)	2.854 (0.628)	0.657 (5.643)	3.735 (0.894)
$EPU^{USA} \rightarrow IP^{CHI}$	0.999 (0.752)	0.457 (1.000)	0.821 (1.000)	0.276 (0.878)	2.094 (1.000)
$EPU^{USA} \rightarrow IR^{CHI}$	0.131 (0.718)	0.332 (1.000)	0.609 (1.000)	0.238 (1.000)	1.358 (1.000)
$EPU^{USA} \rightarrow ER^{CHI}$	0.054 (0.816)	30.530 (0.000)**	30.752 (0.000)**	4.878 (0.000)**	67.766 (0.000)**
$EPU^{USA} \rightarrow EPU^{CHI}$	0.066 (0.797)	0.777 (1.000)	1.324 (1.000)	0.388 (0.770)	3.573 (1.000)

Note: Second column shows the χ^2 statistic for the hypothesis null is that the VAR series x_t does not Granger cause series y_t. Third to sixth column shows ExpW statistic, MeanW, Nyblom and SupLR statistic for the hypothesis null is that the VAR series x_t does not Granger cause series y_t, assuming homoskedastic idiosyncratic shocks.** significant < 5% level and * significant < 10% level for all statistics.

Source: Prepared by the authors based on the sample and using the STATA Software.

5. Conclusions

We have studied how a shock of economic policy uncertainty from the US might spillover to macroeconomic conditions of the LAC under study. We use EPU from the United States to attempt to account for uncertainty spillovers. We estimate a set of two-country structural VAR models. The SVAR model had a better performance to analyze the uncertainty spillovers than other models. Each model has included the US EPU index and one LAC with their corresponding macroeconomic variables from the country sample. The impulse response functions (IRF) are computed accounting for the magnitude of the impact, if any, and its shot-run dynamics.

Considering the entire period, structural VAR model estimates showed that a shock of one standard deviation in the EPU index of the US tend to lead to currency depreciation and positive impacts on EPU indexes for all four LAC, and these estimates tend to be larger for Mexico. Estimates from before and after the 2008 financial crisis suggest that LAC economies became slightly more responsive to EPU index shocks from the US after the 2008 financial crisis.

Finally, robust Granger causality tests that consider structural breaks were applied for the entire period and for the periods before and after the 2008 financial crisis. The estimates before and after the 2008 financial crisis suggest that LAC became slightly more responsive to US EPU shocks after the crisis.

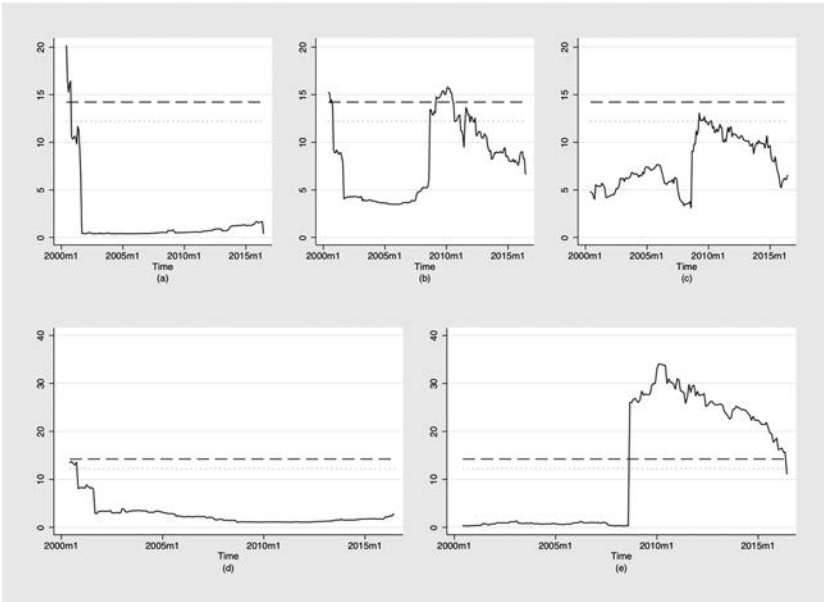
Appendix A

Variables	Source
EE.UU.	
Consumer Price Index (Inflation)	OECD Data https://data.oecd.org
Industrial Production (Industrial)	OECD Data https://data.oecd.org
Exchange Interest (Interest Rate)*	https://fred.stlouisfed.org/series/IR3TIB01USM156N
Economic Policy Uncertainty (Economy Policy, EPU)	http://www.policyuncertainty.com
Brazil	
Consumer Price Index (Inflation)	OECD Data https://data.oecd.org
Industrial Production (Industrial)	OECD Data https://data.oecd.org
Exchange Interest (Interest Rate)*	Banco Central do Brasil https://www.bcb.gov.br/
Exchange Rate	OECD Data https://data.oecd.org
Economic Policy Uncertainty (Economy Policy, EPU)	http://www.policyuncertainty.com
Chile	
Consumer Price Index (Inflation)	OECD Data https://data.oecd.org
Industrial Production (Industrial)	OECD Data https://data.oecd.org
Exchange Interest (Interest Rate)*	Banco de Chile http://www.bancochile.cl
Exchange Rate	OECD Data https://data.oecd.org
Economic Policy Uncertainty (Economy Policy, EPU)	http://www.policyuncertainty.com
Colombia	
Consumer Price Index (Inflation)	OECD Data https://data.oecd.org
Industrial Production (Industrial)	OECD Data https://data.oecd.org
Exchange Interest (Interest Rate)*	Banco de Colombia https://www.grupobancolombia.com/
Exchange Rate	OECD Data https://data.oecd.org
Economic Policy Uncertainty (Economy Policy, EPU)	http://www.policyuncertainty.com
Mexico	
Consumer Price Index (Inflation)	OECD Data https://data.oecd.org
Industrial Production (Industrial)	OECD Data https://data.oecd.org
Exchange Interest (Interest Rate)*	Banco de México http://www.banxico.org.mx
Exchange Rate	OECD Data https://data.oecd.org
Economic Policy Uncertainty (Economy Policy, EPU)	http://www.policyuncertainty.com

* 3 months.

Appendix B

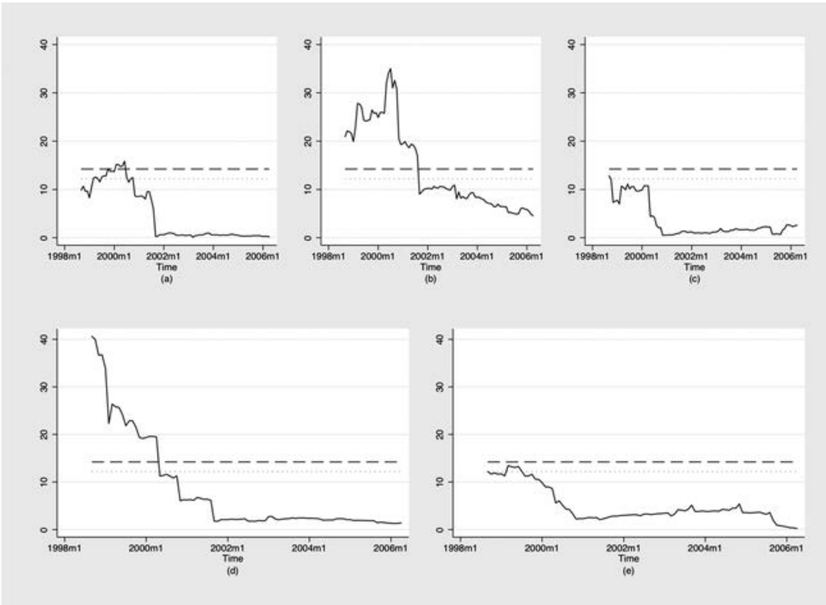
FIGURE B.1
WALD STATISTICS TESTING OF STRUCTURAL BREAKS IN GRANGER CAUSALITY FOR PERIOD 1997M2-2019M12



Note: (a) $EPU^{USA} \rightarrow IR^{MEX}$, (b) $EPU^{USA} \rightarrow ER^{MEX}$ (c) $EPU^{USA} \rightarrow ER^{COL}$, (d) $EPU^{USA} \rightarrow IR^{BRA}$, and (e) $EPU^{USA} \rightarrow ER^{CHI}$ against the alternative of break in Granger causality at time on axis. In all figures, a solid line represents the sequence of the Wald Statistic over time. The dashed line < 5% critical value, and the dotted line < 10% critical value.

Source: Prepared by the authors based on the sample and using the STATA Software.

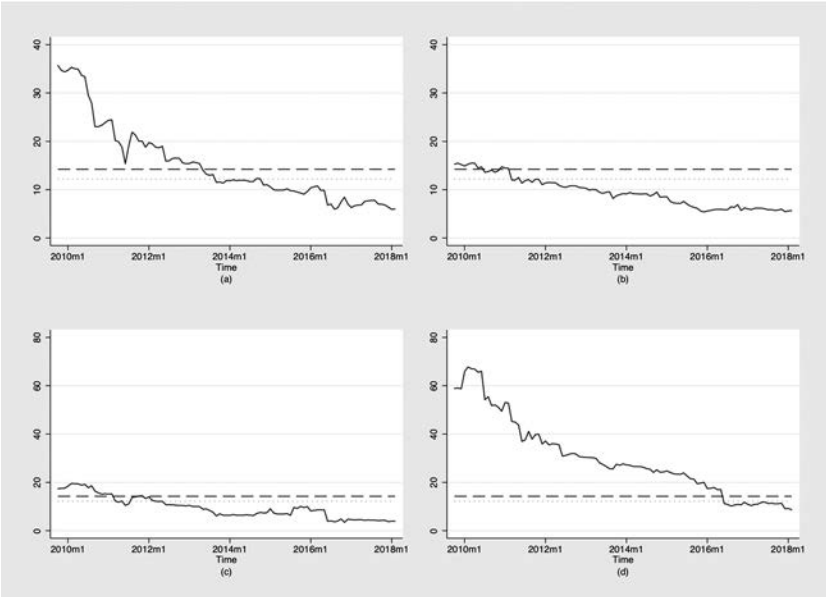
FIGURE B.2
WALD STATISTICS TESTING OF STRUCTURAL BREAKS IN GRANGER CAUSALITY FOR PERIOD 1997M2-2008M12



Note: (a) $EPU^{USA} \rightarrow IR^{MEX}$, (b) $EPU^{USA} \rightarrow ER^{MEX}$ (c) $EPU^{USA} \rightarrow ER^{COL}$, (d) $EPU^{USA} \rightarrow IR^{BRA}$. and (e) $EPU^{USA} \rightarrow ER^{CHI}$ against the alternative of break in Granger causality at time on [illegible] axis. In all figures, a solid line represents the sequence of the Wald Statistic over time. The dashed line < 5% critical value, and the dotted line < 10% critical value.

Source: Prepared by the authors based on the sample and using the STATA Software

FIGURE B.3
WALD STATISTICS TESTING OF STRUCTURAL BREAKS IN GRANGER CAUSALITY
FOR PERIOD 2008M1-2019M12



Note: (a) $EPU^{USA} \rightarrow IR^{MEX}$, (b) $EPU^{USA} \rightarrow ER^{MEX}$ (c) $EPU^{USA} \rightarrow ER^{COL}$, (d) $EPU^{USA} \rightarrow IR^{BRA}$. and (e) $EPU^{USA} \rightarrow ER^{CHI}$ against the alternative of break in Granger causality at time on X axis. In all figures, a solid line represents the sequence of the Wald Statistic over time. The dashed line < 5% critical value, and the dotted line < 10% critical value.

Source: Prepared by the authors based on the sample and using the STATA Software.

TABLE B.1

GRANGER CAUSALITY TESTS IN THE DIRECT MULTISTEP VAR-LP FORECASTING MODEL

PANEL A (1997m2-2019m12)	h=0				h=3			
Variables	ExpW Statistics (p-value)	MeanW Statistics (p-value)	Nyblom Statistics (p-value)	SupLR Statistics (p-value)	ExpW Statistics (p-value)	MeanW Statistics (p-value)	Nyblom Statistics (p-value)	SupLR Statistics (p-value)
$EPU^{USA} \rightarrow IR^{MEX}$	14.874	9.147	3.778	40.074	61.166	51.598	4.473	132.815
	(0.000)	(0.041)	(0.016)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$EPU^{USA} \rightarrow ER^{MEX}$	56.892	31.619	**1.768**	124.082	272.664	138.050	3.730	555.820
	(0.000)	(0.000)	**(1.149)**	(0.000)	(0.000)	(0.000)	(0.017)	(0.000)
$EPU^{USA} \rightarrow ER^{COL}$	14.553	7.059	**0.766**	37.913	46.490	26.184	4.413	102.447
	(0.000)	(0.110)	**(0.492)**	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$EPU^{USA} \rightarrow IR^{BRA}$	41.946	30.919	**1.643**	94.046	36.660	31.394	2.243	83.777
	(0.000)	(0.000)	**(0.178)**	(0.000)	(0.000)	(0.000)	(0.086)	(0.000)
$EPU^{USA} \rightarrow IR^{CHI}$	32.476	27.285	**1.687**	73.459	47.288	25.564	2.914	104.874
	(0.000)	(0.000)	**(0.169)**	(0.000)	(0.000)	(0.000)	(0.039)	(0.000)
PANEL B (1997m2-2007m12)	h=0				h=3			
$EPU^{USA} \rightarrow IR^{MEX}$	221.281	59.746	**1.745**	450.901	51.124	50.425	5.076	111.162
	(0.000)	(0.000)	**(0.155)**	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$EPU^{USA} \rightarrow ER^{MEX}$	24.294	14.294	**1.807**	56.802	134.009	91.971	4.755	276.990
	(0.000)	(0.000)	**(0.142)**	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$EPU^{USA} \rightarrow IP^{BRA}$	491.281	131.192	**1.078**	991.568	134.631	64.155	2.772	278.238
	(0.000)	(0.000)	**(0.000)**	(0.000)	(0.000)	(0.000)	(0.046)	(0.000)
$EPU^{USA} \rightarrow IR^{BRA}$	165.113	27.790	**1.112**	339.232	76.951	63.126	**1.464**	162.831
	(0.000)	(0.000)	**(0.331)**	(0.000)	(0.000)	(0.000)	**(0.216)**	(0.000)
$EPU^{USA} \rightarrow EPU^{BRA}$	146.483	139.878	4.492	301.972	16.549	9.830	**0.737**	41.795
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.029)	**(0.511)**	(0.000)
PANEL C (2008m1-2019m12)	h=0				h=3			
$EPU^{USA} \rightarrow IR^{MEX}$	23.615	14.585	**1.449**	55.459	268.339	182.245	4.031	545.432
	(0.000)	(0.000)	**(0.220)**	(0.000)	(0.000)	(0.000)	(0.011)	(0.000)
$EPU^{USA} \rightarrow ER^{COL}$	92.248	37.701	**1.493**	193.368	39.135	14.24	7.409	85.831
	(0.000)	(0.000)	**(0.210)**	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$EPU^{USA} \rightarrow ER^{BRA}$	74.998	72.625	**1.820**	159.092	30.847	22.997	2.222	70.572
	(0.000)	(0.000)	**(0.140)**	(0.000)	(0.000)	(0.000)	(0.088)	(0.000)
$EPU^{USA} \rightarrow ER^{CHI}$	18.963	16.440	3.614	46.321	83.473	42.691	3.789	176.117
	(0.000)	(0.000)	(0.019)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Note: ExpW, MeanW, Nyblom and SupLR are statistics for the null hypothesis that the VAR series x_t does not Granger cause series y_t assuming heteroskedastic and serially correlated idiosyncratic shocks. The bolded values are not statistically significant. The rest are significant at the < 5% level.

Source: Prepared by the authors based on the sample and using the STATA Software.

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