ESCUELA DE NEGOCIOS Y ECONOMÍA

PONTIFICIA UNIVERSIDAD CATÓLICA DE VALPARAÍSO

Working Papers

2018-01

The Exchange Rate and Export Variety: A cross-country analysis with long panel estimators

Daniel Goya

Pontificia Universidad Católica de Valparaíso

ene.pucv.cl/wp

The Exchange Rate and Export Variety: A cross-country analysis with long panel estimators

Daniel Goya*

Monday 16th December, 2019

Abstract

I study how the number of categories exported by countries is related to the level and the volatility of the exchange rate. I find that export variety is positively related to a weaker exchange rate and negatively related to its volatility. These relationships seem stronger for goods with higher technological intensity.

Using data for a long panel of countries, I investigate these relationships using a methodology that allows for heterogeneous coefficients across countries and discuss two sources of bias that are often overlooked.

Keywords: Export diversification, export variety, exchange rate, exchange rate volatility, pooled mean group.

JEL codes: F14, F40, O30.

1 Introduction

Export diversification, in terms of the diversity of the products exported,[1] is usually stated to be a desirable policy objective, aligning with arguments related to the stability of growth and

*School of Business and Economics, Pontificia Universidad Católica de Valparaíso. Avenida Brasil 2830, Valparaíso, Chile, daniel.goya@pucv.cl, +56 32 2273585, ORCID 0000-0003-0461-845X. I would like to thank Gabriel Palma, Giancarlo Corsetti, Melvyn Weeks, Sean Holly, M. Hashem Pesaran, Giammario Impullitti, Donald Robertson, Jorge Katz, Richard Smith, Meredith Crowley, Beata Javorcik, Markus Eberhardt and Tadashi Ito for helpful comments and suggestions. I also thank participants at the 16th European Trade Study Group Conference, the Econometrics Workshops in Cambridge and the Intelis Seminars at Universidad de Chile. **Funding**: I thank Iniciativa Científica Milenio [Nucleo Milenio NS100017 'Intelis Centre'] and Comisión Nacional de Investigación Científica y Tecnológica [Becas Chile 79090016] for financial support during different stages of this work. All remaining errors are my own.

[1]This paper is concerned with the *variety* aspect of diversification (i.e., the number of product categories exported) rather than the *concentration* dimension (i.e., whether one or a few sectors represent the lion's share of exports).

the Schumpeterian idea of the continuous creation of new products and the growth they may induce. There is growing empirical evidence suggesting that weak and stable real exchange rates (RER) as well as higher export variety are both associated with output growth.[2] The central issue that this paper explores is whether the exchange rate is related to export variety, which could be one of the channels through which the exchange rate has an impact on growth.

The level and the volatility of the exchange rate could have an impact on firm-level decisions about exporting new products or abandoning existing ones. These decisions could, in turn, have an impact on *total export variety*, which is defined as the number of categories that a country exports.

If the exchange rate is related to export variety, is the relationship the same for all types of products? Many growth models (e.g., those described in Lucas, 1988 and Young, 1991) and empirical evidence (e.g., Johnson et al., 2007; Dell et al., 2008; and Hausmann et al., 2007) have suggested that long-term productivity growth potential seems to vary across different types of goods. Subsequently, a second issue is whether the eventual impact of the exchange rate on export variety is the same (or not) for goods with different degrees of sophistication or technological intensity. Third, I seek to determine whether the relationships of interest are the same for countries with different income levels.

The results indicate that a positive relationship exists between depreciation and export variety and that a negative relationship exists between exchange rate volatility and export variety. These relationships seem to be stronger for products with higher technological intensity. In terms of the income level of the countries, no clear conclusions can be drawn.

This paper contributes in two ways: one is the relationship that is estimated, and the other is the econometric methodology used, which allows us to look at two sources of bias that are often overlooked.

The relationship between the exchange rate and export variety has been explored at the firm level (e.g., Tang and Zhang, 2012 and Iacovone and Javorcik, 2008) and by focusing on bilateral variety (the number of products exported to each destination market). The issue studied here is whether the level and the volatility of the exchange rate are determinants of the *total* number

[2] Regarding the exchange rate, Eichengreen (2008) reviews the literature and concludes that both RER level and volatility matter for growth (see Rodrik (2009), Schnabl (2007) and Eichengreen and Leblang (2003), among others). Diversification, in terms of the variety of products exported, has been associated with growth by Funke and Ruhwedel (2001a, 2005) and Saviotti and Frenken (2008). In addition, Addison (2003) and Feenstra and Kee (2008) find that a relationship exists between export variety and productivity growth. The direction of causality between diversification and income, however, is unclear, as argued in the review by Cadot et al. (2012).

of varieties exported by a country, which could be more important from a productivity growth perspective than the number of varieties exported to each destination market.

The second way this paper contributes is the econometric methodology that is used. All previous related papers use estimators that assume the coefficients across units (firms or countries) are homogeneous. As shown by Robertson and Symons (1992) and Pesaran and Smith (1995), this assumption might introduce bias into the estimates. This paper uses estimators that allow for heterogeneity of the coefficients across countries, and two potential sources of bias that are often overlooked are discussed. Although these sources of bias do not seem to drive the results, there is evidence of bias due to the cross-sectional dependence of the residuals and of the heterogeneous slopes (especially for the exchange rate level), which suggests that the estimates of similar models might be biased, underestimating the effect of the exchange rate level.

The next section reviews the theoretical arguments and the existing empirical evidence on the issues at hand. Section 3 describes the data and the variables, and Section 4 describes the econometric approach. Section 5 presents the results and discusses potential econometric issues and the robustness tests. Section 6 summarises the paper's findings and concludes.

2 Literature Review

2.1 Exchange rate level and export variety

In a Melitz (2003) like framework, with sunk entry costs and heterogeneous productivity levels, currency depreciation can be considered to have the same effect as a reduction in trade costs. That is, a weaker currency would reduce the productivity cutoff for exporting, resulting in an increase in the number of exporting firms (which is equivalent to varieties in monopolistic competition models). This increase would occur in models with constant demand elasticity, such as Melitz (2003) and Chaney (2008), and in models with heterogeneous pricing to market, such as Berman et al. (2012) and Chatterjee et al. (2013). Models of multi-product firms, such asBernard et al. (2011) and Chatterjee et al. (2013), also predict that firms export a broader range of products under a weaker currency. At the margin, as depreciation induces firms to start exporting and to add varieties to their export baskets, the number of varieties exported at the country level may increase.

Iacovone and Javorcik (2008) find that devaluations precede export 'discoveries' across Mexican firms. Freund and Pierola (2008) study surges in manufacturing exports in developing countries and find that these surges are preceded by strong real devaluations and a reduction in exchange rate volatility. Freund and Pierola find that depreciation increases entry into new products and new markets and that these new flows account for 25% of the growth during the surges. Tang and Zhang (2012) find that an exchange rate appreciation has a negative impact on the firm-level extensive margin, which is measured as each product-destination pair served by a firm. Tang and Zhang (2012) and Freund and Pierola (2008) consider entries into new products and new markets. In contrast, the focus here is only on product variety.

The paper most closely related to this study is the one by Colacelli (2010), who studies the responses of exports to bilateral RER fluctuations. She decomposes trade into extensive and intensive margins, following Feenstra (1994) and Hummels and Klenow (2005),[3] and finds that '*the extensive margin of trade has a significant role in overall yearly export responses to real exchange rate fluctuations*', especially among less substitutable exports. In this paper, I provide new evidence reinforcing her finding that the level of the exchange rate has a differentiated impact on the export variety of different types of goods.

2.2 Exchange rate volatility and export variety

Baldwin and Taglioni (2004) develop a Melitz-style partial equilibrium model, i.e., with sunk exporting costs and heterogeneous marginal costs, which means that only firms above a certain productivity cutoff will export. In their model, firms are risk averse, and exchange rate volatility reduces the firm's utility from profits. Given this setup, a reduction in volatility lowers the productivity cutoff, and more firms start exporting.

Lin (2012) develops a general equilibrium monetary model similar to that in Bacchetta and Van Wincoop (2000), with heterogeneous productivity and fixed entry costs, which depend on the exchange rate. The nominal exchange rate is a function of the money supplies of the two countries in the model, and uncertainty comes from randomly distributed disturbances in the money supplies (which can have different degrees of correlation between the two countries). As in Baldwin and Taglioni (2004), volatility affects the productivity cutoff for exporting, reducing the number of exporters.

[3] This extensive margin measure, which will be discussed below, adjusts for the importance of the products.

Similar to the discussion above regarding the level of the exchange rate, the idea is that the entries or exits of varieties at the firm level could, at the margin, have an effect on aggregate, country-level export variety.

Evidence that exchange rate volatility has a negative relationship with measures of export variety is found by Lin (2012), Berthou and Fontagné (2008), Álvarez et al. (2009) and Héricourt and Poncet (2013). All of these studies, however, look at firm-level measures or bilateral measures of export variety or define variety in a different way than this paper, which looks at the *total number of different product categories exported by each country*, as its motivation is the potential relationship between variety at the country level and long-run productivity growth.

3 Data and Variables

3.1 Measuring export variety

Disaggregated export data is needed to build a measure of export variety. The main source of data is the World Trade Flows dataset compiled by Feenstra et al. (2005), which contains four-digit SITC revision 2[4] data for over 100 countries for the period between 1962 and 2000. The length of the panel is the key factor for choosing this database, because the estimation methods used require the use of long panels.[5]

An alternative data source was also used to check for robustness: the BACI database (Gaulier and Zignago, 2010), which is based on data obtained from Comtrade. This database has a higher level of disaggregation, although is covers a shorter time span, which is critical for the estimation methods that will be used.[6]

The interest here is in *total* export variety (i.e., the total number of product categories exported, not the number of product categories exported to each country). The simplest measure of export variety is the number of categories exported, which at a high level of disaggregation, can be interpreted as types of products. This measure will be used throughout the paper.

Another possibility is using a measure that considers the importance of the products. Hum-

[4]SITC stands for 'Standard International Trade Classification'. At the four digit level, the revision 2 classification comprises 778 product categories.

[5]In practice, at most 59 countries (listed in each regression table) and 29 time periods are used in any single regression, due to the data requirements of the estimation method.

[6]There is another issue with more disaggregated data. It can be argued that disaggregated data allows for a better measurement of the number of categories exported; however, due to the updates in the classification systems (HS codes are updated roughly every five years), there is a higher risk of the misclassification of exports, which would make the measure of export variety more noisy.

mels and Klenow (2005), following Feenstra (1994), decompose exports into intensive and extensive margins, with the latter representing a weighted count of the exported varieties. More weight is given to the varieties that represent a larger share of the exports of a reference group. The appeal of this measure arises from its weighting and its sound theoretical basis (it is derived from a CES utility function). The downside (for this study) is that this measure does not exclusively capture changes in the diversity of categories that are exported.

To investigate whether the impact of the exchange rate differs for categories with varying degrees of technological intensity, it is necessary to find ways to classify the product categories. After classifying the product categories, variety measures for each group can be constructed. The implied productivity measure proposed by Hausmann et al. (2007), which they call 'prody', is used as the baseline. Prody is defined as *'a weighted average of the per capita GDPs of countries exporting a given product'* (Hausmann et al., 2007). Prody is calculated for the four-digit SITC categories; then, the number of product categories exported with values of 'prody' higher and lower than its median are counted separately. Alternative measures are used for robustness checks.

3.2 Exchange rate measures

The data for the exchange rate measures was obtained from the IMF's International Financial Statistics (IFS). The baseline measures are simple: for the exchange rate level, a yearly real effective exchange rate index (based on a price deflator) is used. An effective rate is used because the interest is in total variety rather than bilateral variety. A higher value is associated with a more competitive currency.

To build yearly volatility measures, the monthly version of the variable described above was used. The most commonly used volatility measure is preferred: the standard deviation of the log differences of monthly rates.

The results were confirmed using other measures, including nominal rates for the level and the volatility, as well as measures based on black market rates obtained from Reinhart and Rogoff (2004).

Table 1 shows the main variables used.

Table 1: Summary of the main variables.

	Concept	**Measured used**
Dependent variable	Export variety	Number of different four-digit SITC categories exported.
Exchange rate	Exchange rate level	Real effective exchange rate index (REER). Higher is weaker.
(Independent variables)	Exchange rate volatility	Standard deviation of the log differences of monthly REER.

3.3 Additional controls

Based on the findings of previous theoretical and empirical works, four variables are used as controls: GDP per capita,[7] population, openness to trade (imports plus exports over GDP), and public education expenditures (current plus capital, as a share of GDP), which is the education measure available for more country-year pairs. These variables were obtained from the World Bank's World Development Indicators (WDI).

Other controls were discarded because the available time series were not long and continuous enough for Pooled Mean Group (PMG) estimation.

4 Econometric Approach

The previous empirical studies most closely related to this paper are based on fixed effects estimation (e.g., Colacelli, 2010 and Freund and Pierola, 2008). This estimation method, as well as other commonly used ones (instrumental variables, dynamic GMM, etc.), make a strong assumption that is usually overlooked: they assume homogeneity of the coefficients across groups (i.e., that the relationships between the variables are the same for all the countries in the sample). This assumption allow these estimators to pool data over groups and increase efficiency. However, Robertson and Symons (1992) and Pesaran and Smith (1995) show that when the regressors are autocorrelated, the methods traditionally used with short dynamic panels, such as fixed effects, instrumental variables, and GMM estimators *'can produce inconsistent, and potentially very misleading estimates of the average values of the parameters in dynamic panel data models unless the slope coefficients are in fact identical'* (Pesaran et al., 1999).

The differences in market and institutional conditions across countries make it reasonable to think that the way that export variety adjusts to changes in the level or the volatility of the exchange rate can differ across countries, especially in the short run. To account for these

[7] Although there is evidence that a nonlinear relationship exists between income per capita and export variety (Klinger and Lederman, 2006; Cadot et al., 2011), to preserve degrees of freedom, the regressions include only one term for GDP per capita. Adding a second term to capture nonlinearity did not appear to affect the conclusions.

differences and assess the possible bias unaccounted for in previous work, this paper focuses on estimators designed for large T and large N datasets, which allow for heterogeneity in the coefficients for different groups.

Below I describe the estimation methods that are used in this paper. First, assume that the following long-run relationship holds:

$$Variety_{i,t} = \theta_{0,i} + \theta_{1,\mathrm{i}} GDPpc_{i,t} + \theta_{2,i} RER_{i,t} + \theta_{3,i} X^3_{i,t} + ... + \theta_{R,i} X^R_{i,t} + u_{i,t} \tag{1}$$

$i = 1, 2, ..., N, \quad t = 1, 2, ..., T,$

where $Variety_{i,t}$ is a measure of export variety in country i in year t,[8] $GDPpc$ stands for GDP per capita, RER is the real exchange rate level, the X^r represent the other controls and u_{it} represents the unobserved determinants of export variety.

There are N countries, T time periods and R explanatory variables. All the regressions reported will include both one measure of the exchange rate level and one of its volatility, but only the former and one additional control will be included here for simplicity of exposition.

The vector θ_i contains the point estimates of each coefficient for country i. It is important to emphasise that the coefficients are allowed to be country-specific.

The variety measure is highly persistent,[9] suggesting that a dynamic model should be used. The starting point is the following ARDL specification:[10]

$$Variety_{i,t} = \mu_i + \rho_i Variety_{i,t-1} + \delta_{1,0,i} GDPpc_{i,t} + \delta_{1,1,i} GDPpc_{i,t-1} + \delta_{2,0,i} RER_{i,t} + \delta_{2,1,i} RER_{i,t-1} + \varepsilon_{i,t} \tag{2}$$

with its corresponding error correction form:

$$\Delta Variety_{i,t} = \phi_i (Variety_{i,t-1} - \theta_{0,i} - \theta_{1,i} GDPpc_{i,t-1} - \theta_{2,i} RER_{i,t-1}) - \delta_{1,1,i} \Delta GDPpc_{i,t} - \delta_{2,1,i} \Delta RER_{i,t} + \varepsilon_{i,t} \tag{3}$$

where

$\theta_{0,i} = \frac{\mu_i}{1-\rho_i}$, $\theta_{1,i} = \frac{\delta_{1,0,i}+\delta_{1,1,i}}{1-\rho_i}$, $\theta_{2,i} = \frac{\delta_{2,0,i}+\delta_{2,1,i}}{1-\rho_i}$, $\phi_i = -(1-\rho_i)$,

$\theta_i = (\theta_{0,i} \quad \theta_{1,i} \quad \theta_{2,i})$ and $\theta = E(\theta_i)$.

[8]As defined in the previous section. It could be argued that count data models should be used due to the nature of the dependent variable, but with relatively large numbers such as those occurring here, these models are less appealing. The well-known issue of zeroes in trade regressions is not a problem here.

[9]The first-order autocorrelation of the baseline variety measure is 0.9875.

[10]To simplify the exposition, an ARDL model with one lag for each variable is used here.

The structure of the unobserved term $\varepsilon_{i,t}$ will be discussed in more detail in Section 5.4.2.

Using yearly data to estimate multi-country relationships, there are two extreme opposite ways to proceed: one is to assume that the slopes and intercepts are homogeneous and to pool over groups (pooled OLS). The other is to allow for full heterogeneity by estimating the relationship separately for each country without imposing cross-country restrictions on the parameters. These estimates can then be averaged over groups to obtain consistent estimates of the mean short-run and long-run parameters: this is Pesaran and Smith's (1995) Mean Group (MG) estimator.

There are several alternatives between these two extremes. The Dynamic Fixed Effects (DFE) estimator imposes slope homogeneity but allows for heterogeneity in the intercepts.[11] The Pooled Mean Group (PMG) estimator developed by Pesaran, Shin, and Smith (1999) allows for heterogeneity in the intercepts, short-run adjustment parameters (δ_i in equation 3), and error variances, but it imposes homogeneity on the long-run parameters (θ_i in equation 3 becomes θ).

The main assumptions required for consistent PMG estimation are: a) the ARDL model in equation (2) is stable (ensuring the existence of a long-run relationship between the dependent variable and the independent variables) ; b) the long-run coefficients are the same across every group ($\theta_i = \theta \ \forall i$); and c) the disturbances $\varepsilon_{i,t}$ are independently distributed across i and t and independent of the regressors.[12]

Assumption a) can be informally tested by checking that the error correction model adjustment speed coefficients ϕ_i are significantly negative but above -1. A formal test can be conducted, following Pesaran, Shin and Smith (2001). PMG estimation requires the existence of a long-run relationship, but consistent estimation is possible *regardless of the order of integration of the regressors* (Pesaran et al., 1999). Nevertheless, the variables and the regression residuals were checked for stationarity.

Assumption b) points to the usual trade-off between consistency and efficiency. Efficiency increases as stronger homogeneity restrictions are imposed but at the expense of a loss in robustness. In other words, the estimators with stronger cross-country restrictions will be more efficient, but if the assumptions behind these restrictions are not valid, then they will produce

[11] This is simply fixed effects estimation of the ARDL model, reporting the implied long-run parameters from the error correction form.

[12] There are other more technical assumptions, such as the true parameter being an interior solution, positive variance of the unobserved, etc.

inconsistent estimates.

In this context, the Mean Group (MG) estimator (Pesaran and Smith, 1995) is useful. This estimator will provide consistent estimates of the mean of the long-run parameters, even if they are heterogeneous across countries, but these estimates will be inefficient if the long-run slopes are in fact homogeneous. Therefore, this estimator can be used as the basis for a Hausman (1978)-style test for the assumption of long-run slope homogeneity, which is needed by the PMG estimator.

In economic terms, the PMG estimator assumes that the relationship of interest is the same *in the long run* across all countries, but that *short-run adjustment dynamics* can differ. Export variety can react differently to changes in the exchange rate level or volatility in the short-run due to differences in financial development, labour market flexibility, availability of educated labour, etc.

Assumption c) has several parts. Regarding regressor exogeneity and independence across time, Pesaran and Shin (1998) have shown that sufficient augmentation of the lag order of the ARDL model can, in principle, address these issues, and thus, standard inference on the long-run parameters is valid. Moreover, endogeneity is more of an issue for the short-run parameters (Pesaran et al., 1999), which are not of central interest here. Independence of $\varepsilon_{i,t}$ across groups, or cross-sectional independence, is a more complicated issue—to such an extent that most empirical studies assume it away.[13] This issue will be discussed in detail in Section 5.4.2. The next section presents the results, using the estimation methods explained above.

5 Results

This section presents three sets of results. First, it shows the relationship between the exchange rate and export variety, measured as the (log) count of four-digit SITC (SITC4) categories exported. Then, it presents the results showing whether this relationship is heterogeneous across different types of products. Finally, the results for low and middle income countries are compared to those of high income countries. A discussion about econometric concerns and robustness tests follows the results.

Unreported unit root tests indicate that the demeaned variables used in the following regres-

[13] Some exceptions include Eberhardt and Teal (2010); Holly et al. (2010) and Cavalcanti et al. (2011).

sions are stationary.[14]

The tables below present the Mean Group (MG), Pooled Mean Group (PMG) and Dynamic Fixed Effects (DFE) estimates of two different specifications The models differ in terms of their lag structures and the included regressors. Model 1 includes only the first lag of the dependent variable (log of the number of exported varieties) and of each independent variable (the logs of the real exchange rate, real exchange rate volatility, GDP per capita, trade openness, population and public education expenditures). Model 2 augments the lag structure of the ARDL model and includes the first two lags of the dependent variable and of each independent variable. However, to preserve degrees of freedom, only GDP per capita is included as an additional regressor.

Country dummies are always included, and data is always cross-sectionally demeaned (which is equivalent to including time dummies). Only the implied long-run coefficients are reported. The level of the exchange rate is defined so that a higher value is associated with a weaker (more competitive) currency.

5.1 Exchange rate and export variety

The results in Table 2 show that variety is positively related to depreciation and negatively related to exchange rate volatility. The coefficient for the level of the exchange rate is significantly positive for all models except the MG version with fewer lags. The coefficient for exchange rate volatility is significantly negative under both the PMG and DFE specifications. The estimated real exchange rate level elasticities of export variety are in the approximate range of 0.17 to 0.53; the exchange rate volatility elasticities lie in a more narrow approximate range, between -0.07 and -0.12.

In most cases, GDP per capita is significantly positive. Trade openness is always negative, as expected if openness induces specialisation. Population and education expenditures do not appear to be significant determinants of export variety.

For all regressions in Table 2, the speed of adjustment is significantly negative and smaller than one in absolute value, as required for a long-run relationship to exist. Residual autocorrelation is evaluated for each country's equation and reported (as the number of groups for which the correlation is significant) only for the MG and PMG estimates, which have country-specific equations and thus can be compared more directly. Only MG model 1 presents some problems

[14]Including augmented Dickey-Fuller type (Im et al., 2003) and Fisher type (Choi, 2001) tests.

here, suggesting that including only one lag might not be enough.

All specifications show difficulties in terms of the assumption of cross-sectional independence of the residuals, as indicated by Pesaran's (2004) CD test.[15] This assumption will be discussed in detail in Section 5.4.2. In terms of residual stationarity, all specifications pass the test.[16]

The Hausman test does not reject the null hypothesis in any of the specifications; i.e., the differences between the coefficients do not appear to be systematic for both the PMG and DFE estimates vis-à-vis the MG estimates. This result means that in principle, we can rely on the assumption that the coefficients are homogeneous and prefer the more efficient DFE estimates. However, the coefficients for GDP and especially for the exchange rate level are much smaller for DFE estimation than for PMG estimation, which suggests that bias may have been caused by mistakenly imposing homogeneity on the short-run coefficients.[17] The point estimates for Model 2 are the most similar across the different models.

5.2 Heterogeneity across product types

The possibility suggested by previous empirical studies, namely, that the exchange rate could have a heterogeneous impact on export variety across different sectors (e.g. Colacelli, 2010), is especially important for the potential growth effects of export variety. To explore this issue, the 'prody' measure of implied productivity proposed by Hausmann et al. (2007) is used.[18] Using the median prody as the cutoff, the number of exported categories with a high or low 'prody' value were counted separately and are used as the dependent variables.

Table 3 shows the results for the number of low 'prody' (Panel A) and high 'prody' (Panel B) varieties, respectively, for Models 1 and 2, as before. All the PMG and DFE estimates are significantly positive for the exchange rate level and significantly negative for its volatility, and most of the MG estimates are also significant for Model 2. The estimated elasticities are much larger in magnitude for the high-prody exports: the estimates for high prody are between 1.3 and 2.5 times as large as those for low prody for the level of the exchange rate and between

[15] The null hypothesis is that there is only weak cross-sectional dependence, and the test statistic is distributed as a standard normal under the null.

[16] Pesaran's (2007) augmented Dickey-Fuller (ADF) test, which includes cross-sectional averages to be robust to cross sectional dependence, was used. The null hypothesis is that all series are non-stationary. And given that the demeaned regressors are stationary, there is no risk of cointegration *between the regressors*.

[17] Section 5.4.3 discusses this in detail.

[18] Prody is defined as *'a weighted average of the per capita GDPs of countries exporting a given product'* (Hausmann et al., 2007).

Table 2: Log of the number of *total* SITC4 categories exported

	Model 1: 1 lag in short run eq.			Model 2: 2 lags in short run eq.		
	MG	PMG	DFE	MG	PMG	DFE
RER level	2.023	0.459***	0.170***	0.369***	0.527***	0.216***
	(0.200)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)
Exchange rate volatility	0.0159	-0.0711***	-0.0946***	-0.0144	-0.124***	-0.0936***
	(0.781)	(0.000)	(0.000)	(0.768)	(0.000)	(0.000)
GDP per capita	2.458	0.655***	0.453***	0.596	0.579***	0.448***
	(0.104)	(0.000)	(0.000)	(0.275)	(0.000)	(0.000)
Trade openness	-0.854*	-0.209***	-0.0107			
	(0.074)	(0.001)	(0.897)			
Population	4.274	-0.112	0.228*			
	(0.380)	(0.316)	(0.077)			
Education expenditure	0.911	-0.0310	0.0217			
	(0.381)	(0.585)	(0.829)			
Adjustment speed	-.875***	-.405***	-.355***	-.996***	-.639***	-.482***
N	1200	1200	1200	1084	1084	1084
Pesaran CD	3.571	7.451	9.702	8.431	6.032	6.112
CD p value	0.000	0.000	0.000	0.000	0.000	0.000
ADF p (1 lag)	0.000	0.000	0.000	0.000	0.000	0.000
Groups with serial correlation	16	1		6	3	
Hausman test		0.999	1.000		0.811	1.000

Notes. Mean Group, Pooled Mean Group, and Dynamic Fixed Effects estimation using Feenstra et al.'s (2005) World Trade Flows dataset. Only the long run coefficients are reported. Time and country dummies are included (implicitly) in all regressions. All variables in logs. Models 1 and 2 have one and two lags respectively for every variable in the short run equation. 'ADF p' reports the p-value for Pesaran's (2007) panel unit root test. The null is that all series are non-stationary. At most one lag seemed to be necessary for these tests. 'Pesaran CD' and 'CD p value' are the test statistic and p-value for Pesaran's (2004) cross-sectional dependence test for the residual (see 5.4.2). Residual autocorrelation was evaluated equation by equation for MG and PMG estimates, the number of countries where it was significant at the 5% level is reported. 'Hausman test' reports the p-value for the Hausman test comparing the MG to PMG or to DFE estimates (rejection means the efficient estimator is inconsistent).

Countries included in the regressions: Australia, Austria, Bahrain, Belgium, Canada, Cyprus, Denmark, Finland, France, Greece, Hungary, Israel, Japan, Malta, Netherlands, New Zealand, Norway, Portugal, Saudi Arabia, Singapore, Spain, Sweden, Switzerland, Trinidad and Tobago, United Kingdom, United States, Algeria, Belize, Bolivia, Brazil, Cameroon, Central African Republic, Chile, China, DR Congo, Costa Rica, Côte d'Ivoire, Dominican Republic, Ecuador, Fiji, Gambia, Ghana, Guyana, Iran, Malawi, Malaysia, Mexico, Morocco, Nicaragua, Pakistan, Paraguay, Philippines, Sierra Leone, Togo, Tunisia, Uruguay, Venezuela, Zambia.

p-values in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

2.6 and 3.6 times as large for volatility.[19] This result supports the idea that the impact of the exchange rate is not the same for all types of products, as found by Colacelli (2010).

All specifications passed the Hausman and ADF tests. There is some evidence of groups (countries) with serial correlation, but not an important number of them for the PMG estimates. As in Table 2, the CD test provides evidence of strong cross-sectional dependence. For all specifications, the CD test statistic is much higher for the high prody exports. The MG results for model 2 should be interpreted carefully. Their adjustment speeds are not consistent with the existence of a long-run relationship; however, the coefficient estimates are roughly in line with those of the other estimators.

As in the previous subsection, Models 1 and 2 lead to the same conclusions.

One possible explanation for the finding that a heterogeneous relationship exists between exchange rate volatility and export variety for different types of products is the idea proposed by Rauch (1999) that homogenous goods, which can be traded in organised exchanges, are not affected by uncertainty in the same way as differentiated goods.

The results are also consistent with the idea of 'costly discovery' that was proposed by Hausmann and Rodrik (2003); there are information externalities that reduce experimentation in new varieties (i.e., the experimenter must pay the discovery costs; subsequently, potential new entrants have access to this information for free). If we assume that varieties with higher technological intensity are more difficult to imitate, then the impact of this information externality would be reduced. Therefore, when there is a marginal change in profitability due to a change in the exchange rate, we can expect that the categories that were marginally unprofitable before and that can easily be imitated will *not* be developed, while those that are difficult to imitate might be developed.

[19] It is not straightforward to test whether these differences are significant.

Table 3: Log of the number of *low* and *high 'prody'* SITC4 categories exported

	Model 1: 1 lag in short run eq.			Model 2: 2 lags in short run eq.		
	MG	PMG	DFE	MG	PMG	DFE
			Panel A — Low prody			
RER level	0.296**	0.286***	0.141***	0.351**	0.359***	0.198***
	(0.036)	(0.000)	(0.000)	(0.026)	(0.000)	(0.000)
Exchange rate volatility	-0.00360	-0.0386***	-0.0592***	-0.0462*	-0.0497***	-0.0581***
	(0.905)	(0.001)	(0.007)	(0.064)	(0.000)	(0.007)
GDP per capita	0.484	0.438***	0.342***	-0.147	0.502***	0.335***
	(0.134)	(0.000)	(0.000)	(0.599)	(0.000)	(0.000)
Trade openness	-0.252	-0.0771**	0.0115			
	(0.242)	(0.045)	(0.879)			
Population	-0.559	-0.0781	0.159			
	(0.502)	(0.301)	(0.307)			
Education expenditure	-0.131	-0.0453	-0.0366			
	(0.490)	(0.158)	(0.630)			
Adjustment speed	-.989***	-.488***	-.361***	-1.05***	-.601***	-.405***
N	1095	1095	1095	989	989	989
Pesaran CD	2.455	4.827	9.764	6.548	4.439	6.443
CD p value	0.007	0.000	0.000	0.000	0.000	0.000
ADF p (1 lag)	0.000	0.000	0.000	0.000	0.000	0.000
Groups with serial correlation	14	3		7	0	
Hausman test		0.999	1.000		0.840	1.000
			Panel B — High prody			
RER level	0.462	0.660***	0.187***	0.233	0.901***	0.275***
	(0.333)	(0.000)	(0.004)	(0.279)	(0.000)	(0.003)
Exchange rate volatility	-0.125	-0.112***	-0.158***	-0.120*	-0.181***	-0.157***
	(0.502)	(0.000)	(0.000)	(0.058)	(0.000)	(0.000)
GDP per capita	2.006**	0.883***	0.630***	0.539	0.754***	0.602***
	(0.048)	(0.000)	(0.000)	(0.191)	(0.000)	(0.000)
Trade openness	-0.759	-0.396***	0.0263			
	(0.192)	(0.000)	(0.846)			
Population	-0.115	-0.710***	0.170			
	(0.955)	(0.000)	(0.336)			
Education expenditure	0.441	-0.0243	-0.00935			
	(0.309)	(0.764)	(0.944)			
Adjustment speed	-.913***	-.457***	-.460***	-1.01***	-.598***	-.610***
N	1095	1095	1095	989	989	989
Pesaran CD	5.989	13.321	15.034	7.877	7.807	11.038
CD p value	0.000	0.000	0.000	0.000	0.000	0.000
ADF p (1 lag)	0.000	0.000	0.000	0.000	0.000	0.000
Groups with serial correlation	13	1		9	8	
Hausman test		0.999	1.000		0.750	1.000

Notes. Mean Group, Pooled Mean Group, and Dynamic Fixed Effects estimation using Feenstra et al.'s (2005) World Trade Flows dataset. Products classified on those below and above the median *prody* value. Only the long run coefficients are reported. Time and country dummies are included (implicitly) in all regressions. All variables in logs. Models 1 and 2 have one and two lags respectively for every variable in the short run equation. 'ADF p' reports the p-value for Pesaran's (2007) panel unit root test. The null is that all series are non-stationary. At most one lag seemed to be necessary for these tests. 'Pesaran CD' and 'CD p value' are the test statistic and p-value for Pesaran's (2004) cross-sectional dependence test for the residual (see 5.4.2). Residual autocorrelation was evaluated equation by equation for MG and PMG estimates, the number of countries where it was significant at the 5% level is reported. 'Hausman test' reports the p-value for the Hausman test comparing the MG to PMG or to DFE estimates (rejection means the efficient estimator is inconsistent).

Countries included in the regressions: Australia, Bahrain, Canada, Cyprus, Denmark, Finland, Greece, Hungary, Israel, Japan, Malta, Netherlands, New Zealand, Norway, Portugal, Saudi Arabia, Singapore, Spain, Sweden, Switzerland, Trinidad and Tobago, United Kingdom, United States, Algeria, Belize, Bolivia, Brazil, Cameroon, Central African Republic, Chile, China, Costa Rica, Côte d'Ivoire, Dominican Republic, Ecuador, Fiji, Gambia, Ghana, Guyana, Iran, Malawi, Malaysia, Mexico, Morocco, Nicaragua, Pakistan, Paraguay, Philippines, Sierra Leone, Togo, Uruguay, Venezuela, Zambia.

p-values in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

5.3 Developed and developing countries

Cadot et al. (2011) show that as countries grow, first, they increase export diversification and then reconcentrate their exports, and this occurs mostly through changes in the extensive margin. Here, I explore whether the relationship between the exchange rate and export variety is different for countries at different stages of the development process. Table 4 presents the results by splitting the sample between low and middle income countries and high income countries.[20]

The most important finding here is that there is no evidence of cross-sectional dependence of the residuals in the regressions for low and middle income countries. This result is important because it confirms that the bias due to cross-sectional dependence (discussed below) is not driving the findings, including the fact that the relationship between the exchange rate and export variety is stronger for 'high prody' goods.

When comparing the two groups of countries, the coefficients for total variety are larger for high income countries, but when distinguishing between low and high prody exports, the relative magnitudes are not clear, as they change across estimators. Moreover, it is risky to compare the two groups, due to the possibility of cross-sectional dependence bias in only one of them.

5.4 Econometric concerns

5.4.1 Endogeneity

It is possible to think of endogeneity due to a 'Dutch Disease' type of effect. Finding and exporting oil (for example) has a strong impact on a country's currency. This could be biasing the estimates downwards for the exchange rate level for low-prody exports, if most of the commodities that could cause a Dutch Disease are in this group. However, the type of export discovery that can have an impact on the exchange rate is a rare event. Most changes in export variety are due to small new exports or to abandoning (at the country level) products that are no longer profitable. Nevertheless, to eliminate this risk, it is possible to isolate the cases where the changes in variety are associated with non-marginal changes in export volumes. All countries that in any single year had entries or exits that represented over 5% of their exports were dropped to eliminate possible reverse causality running from variety to the exchange rate. The results are the same as before, including the differences between the low and high-prody exports.

[20] For succinctness, only the specification with longer lags is displayed.

Table 4: Log of the number of SITC4 categories exported – *Low-middle vs high income countries*

	Total variety			**Low prody** variety			**High prody** variety		
	MG	PMG	DFE	MG	PMG	DFE	MG	PMG	DFE
Panel A — Low and middle income countries									
RER level	0.0974	0.226***	0.194***	0.316	0.214***	0.193***	-0.391	0.493***	0.202***
	(0.377)	(0.000)	(0.000)	(0.119)	(0.000)	(0.000)	(0.356)	(0.000)	(0.004)
Exchange rate volatility	-0.0246	-0.0541***	-0.0638	-0.228	-0.0405***	-0.0287	0.0879	-0.0911***	-0.0880
	(0.600)	(0.000)	(0.106)	(0.272)	(0.000)	(0.450)	(0.628)	(0.000)	(0.118)
GDP per capita	1.571*	0.490***	0.285	-9.819	-0.279***	0.158	2.087**	1.032***	0.0831
	(0.069)	(0.000)	(0.283)	(0.341)	(0.000)	(0.499)	(0.024)	(0.000)	(0.878)
Adjustment speed	-1.16***	-.691***	-.374***	-1.11***	-.543***	-.328***	-1.15***	-.678***	-.542***
N	564	564	564	525	525	525	525	525	525
Pesaran CD	-0.252	0.742	-1.454	-0.875	-0.300	-1.705	2.106	1.637	0.469
CD p value	0.600	0.229	0.927	0.809	0.618	0.956	0.018	0.051	0.319
ADF p (1 lag)	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Groups with serial correlation	3	1		4	0		6	0	
Hausman test		0.973	>.999		0.983	>.999		0.805	>.999
Panel B — High income countries									
RER level	2.495	0.656***	0.522**	0.495*	0.107***	0.473**	1.093*	0.202***	0.440*
	(0.160)	(0.000)	(0.017)	(0.073)	(0.000)	(0.030)	(0.080)	(0.000)	(0.074)
Exchange rate volatility	0.213	-0.0966***	-0.0814***	0.0977*	-0.0232***	-0.0633**	0.0283	-0.0694***	-0.116**
	(0.228)	(0.000)	(0.001)	(0.086)	(0.000)	(0.012)	(0.632)	(0.000)	(0.014)
GDP per capita	1.613	0.144*	0.341	0.423	0.117***	0.178	1.391	-0.238***	0.429**
	(0.325)	(0.055)	(0.155)	(0.491)	(0.000)	(0.456)	(0.156)	(0.000)	(0.014)
Adjustment speed	-.661***	-.345***	-.336***	-.818***	-.532***	-.339***	-.659***	-.397***	-.310***
N	520	520	520	464	464	464	464	464	464
Pesaran CD	5.602	10.574	7.465	3.479	5.709	3.776	6.622	11.095	8.603
CD p value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
ADF p (1 lag)	0.000	0.000	0.002	0.000	0.000	0.000	0.000	0.000	0.000
Groups with serial correlation	2	0		2	0		1	0	
Hausman test		0.446	>.999		0.851	>.999		0.355	>.999

Notes. Mean Group, Pooled Mean Group, and Dynamic Fixed Effects estimation using Feenstra et al.'s (2005) World Trade Flows dataset. Only the long run coefficients are reported. Time and country dummies are included (implicitly) in all regressions. All variables in logs. Two lags of every variable included in the ARDL model. 'ADF p' reports the p-value for Pesaran's (2007) panel unit root test for the residuals. The null is that all series are non-stationary. At most one lag seemed to be necessary for these tests. 'Pesaran CD' and 'CD p value' are the test statistic and p-value for Pesaran's (2004) cross-sectional dependence test for the residual (see 5.4.2). Residual autocorrelation was evaluated equation by equation for MG and PMG estimates. The number of countries where it was significant at a 5% is reported. 'Hausman test' is the p-value for the Hausman test comparing the MG to PMG or DFE estimates (rejection means the efficient estimator is inconsistent).

Countries included (low-middle income): Algeria, Belize, Bolivia, Brazil, Cameroon, Central African Republic, Chile, China, DR Congo, Costa Rica, Côte d'Ivoire, Dominican Republic, Ecuador, Fiji, Gambia, Ghana, Guyana, Iran, Malawi, Malaysia, Mexico, Morocco, Nicaragua, Pakistan, Paraguay, Philippines, Sierra Leone, Togo, Tunisia, Uruguay, Venezuela, Zambia.

Countries included (high income): Australia, Austria, Bahrain, Belgium, Canada, Cyprus, Denmark, Finland, France, Greece, Hungary, Israel, Japan, Malta, Netherlands, New Zealand, Norway, Portugal, Saudi Arabia, Singapore, Spain, Sweden, Switzerland, Trinidad and Tobago, United Kingdom, United States.

p-values in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Another concern is the possibility of omitted variable bias. The results are robust to adding a trade liberalisation dummy, which is almost always insignificant. If a measure of monetary policy is included, the sample size drops dramatically, but the results still hold for model 1.

Considering the importance of *expectations* about the level and the volatility of the exchange rate, a specification including the leads of these variables in the short-run equation was also checked. None of the results discussed before change under this specification.

The MG, PMG and DFE estimators are all affected by the well-known bias of dynamic fixed effects models (see Nickell, 1981). As this bias is of order $1/T$, it should not be a first order concern here.

As a general check against different potential sources of endogeneity, Table A1 in Appendix A presents Difference and System GMM estimates (see Arellano and Bond, 1991; Arellano and Bover, 1995; Blundell and Bond, 1998) for both low and high-prody export varieties.[21,22] The data is averaged over four years, instead of the more common five years, to increase the sample size. Sargan, Hansen and second-order residual AR tests are passed, and the coefficients for the lagged dependent variables for the GMM estimators lie between the upper and lower bounds provided by the OLS and the two-way fixed effects estimates, respectively, as expected. The coefficients are consistent with the results presented before.

The results discussed here cannot be taken as evidence of a causal relationship, but they do show that the correlations between the exchange rate and the export variety measures are robust to some of the most evident concerns.

5.4.2 Cross-sectional dependence

If the assumption that the residuals are independent across countries does not hold, that could cause in the best case a loss of efficiency, and in the worst inconsistent estimates (Coakley et al., 2006). This can actually be a problem not only for PMG estimation but also for a wide variety of estimators.[23]

Often, empirical papers using the PMG estimator mention the issue in passing, stating that by including time dummies, they *expect* that cross-sectional independence will be achieved. The

[21] Dynamic GMM estimators (e.g. Arellano and Bond, 1991; Blundell and Bond, 1998), in addition to the parameter homogeneity assumption, have problems in terms of the exogeneity and the strength of the internal instruments they exploit as well as with their specification tests (see Roodman, 2009 and Bowsher, 2002).

[22] The well-known instrumental variable approach used by Tenreyro (2007) can only be used for bilateral exchange rates and thus is not applicable here.

[23] For reviews of this in a stationary setting, see Sarafidis and Wansbeek (2012) and Breitung and Pesaran (2008) for nonstationary panels.

problem is that this only works when the unobserved factors have the same impact on all groups (i.e., when they are effectively period fixed effects).[24]

In economic terms, cross-sectional correlation could be the result of spillover effects (e.g., the diffusion of new products across countries) or common macroeconomic shocks that affect countries in a heterogeneous way (Eberhardt et al., 2013).

There are two questions then: is there evidence of cross-sectional dependence in the residuals? If there is, could it be biasing the coefficients and driving the results?

The first question is evaluated using Pesaran's (2004) CD test. The null hypothesis of the test, in contrast with Lagrange multiplier type tests (see Breusch and Pagan, 1980), is not cross-sectional independence but rather weak dependence, as defined by Chudik et al. (2011). As argued by pes, this is a more appropriate test for panels where the cross section dimension is large relative to the time dimension, where strong dependence rather than weak dependence might cause inference problems and complete independence might be unnecessarily restrictive. This test might have low power when time dummies are included, but we do not observe this here.[25]

Pesaran's CD test rejects the null of only weak dependence of the residuals for all specifications in Tables 2 and 3. The test statistics are much higher for high-prody exports, suggesting there could be some difference in the way that the variety of different types of products is correlated across countries, for example, in the form of stronger spillovers across high-prody goods.

The next question is whether the cross-sectional dependence that remains after time-demeaning could be driving the results. There are two main reasons to think that the answer is no: first, the unreported results without demeaning show an interesting pattern; the CD test statistics

[24] We can impose the following structure on the unobserved $\varepsilon_{i,t}$ in equation (2):
$\varepsilon_{i,t} = \gamma_i'\mathbf{f}_t + \nu_{i,t}$,
where $\nu_{i,t}$ is white noise, $\mathbf{f}_t$ is a column vector with k unobserved period-specific shocks (or 'factors') and γ_i represents the k group-specific (here country-specific) factor loadings indicating how group i is affected by each of the k different factors. If $\gamma_i = \gamma \ \forall \ i$, then $\gamma_i'\mathbf{f}_t$ is a period-specific constant, and it can be absorbed by the fixed effects. However, if the factor loadings are heterogeneous ($\gamma_i \neq \gamma_j$), then the time dummies are not able to remove the contemporaneous correlation of the errors across countries. In addition, when the common factors $\mathbf{f}_t$ are present in the unobserved and in the regressors, there is an endogeneity problem and standard estimates will be inconsistent. This problem will occur by construction in dynamic models if the common factors are serially correlated; if $\mathbf{f}_t = \lambda_t\mathbf{f}_{t-1} + \xi_{i,t}$, then $Diversif_{i,t-1}$ is correlated with the unobserved $\varepsilon_{i,t}$ through $\mathbf{f_{t-1}}$. ($\mathbf{\lambda_t}$ is a square matrix defining factor persistence, which is diagonal if the factors are independent).

[25] De Hoyos and Sarafidis (2006), Sarafidis and Wansbeek (2012) and Chudik and Pesaran (2013) argue that the CD test will have low power and might not be consistent after cross-sectional demeaning. However, there are reasons to believe that the CD test is working here: the test is rejecting the null hypothesis after demeaning, and the opposite should happen if the test lacked power. Regardless of this power issue, the test remains the standard in the literature.

are much higher but only for total variety and the low-prody variety. High-prody exports have very low CD statistics, to the point that in some cases, the null is not rejected. The direction and relative magnitudes of the point estimates are the same as before. This result indicates that even if there was a bias, the main results are not altered when the degree of dependence increases.

Second and most important, the results clearly hold for the sub-sample of low and middle income countries, where there is no evidence of cross-sectional dependence, as can be seen in Table 4.[26]

5.4.3 Heterogeneity bias

If the true coefficients are not homogeneous across countries, then the pooled estimators (DFE and to a lesser extent, PMG) are at risk of the heterogeneity bias described by Robertson and Symons (1992). The presence of this bias could be observed in the adjustment speeds, which would be underestimated under heterogeneity bias. Consistent with this, as homogeneity restrictions are imposed going from MG to PMG and from PMG to DFE, the adjustment speed drops in Tables 2, 3 and 4.[27]

The results for model 2 in Table 3 (panels A and B) suggest that the findings for volatility, including the fact that the coefficient is larger in absolute value for high-prody exports, cannot be driven by the heterogeneity bias, as they hold for the robust MG estimator. The results for the level also hold for the MG estimator for total variety (Table 2, model 2) and for low prody variety (Table 3).

As discussed before, there is evidence of cross-sectional dependence in the residuals, which may induce a further bias in an unknown direction. Thus, the most important table for analysing the effect of the bias due to imposing parameter heterogeneity is Table 4, where cross-sectional dependence bias does not seem to be a concern for low and middle income countries (panel A). Here, the estimates for the exchange rate level, when significant, are always smaller under estimators that impose homogeneity, suggesting that previous studies (those estimating long-run effects at the country level with dynamic models) might be understating the effect of the level

[26]For some specifications, the null is rejected but only at the 5% and 10% levels, while it is rejected at the 0.1% level for all other specifications.

[27]The only exception is with the DFE estimates for high-prody exports (Table 3 Panel B), which also have relatively high CD test statistics, suggesting that there could be bias due to cross-sectional dependence in the residuals at the same time.

of the exchange rate on export variety.[28] The coefficients for volatility are much more stable across models, indicating that the way that variety reacts to exchange rate volatility may be more homogeneous across countries.

5.5 Robustness checks

This section will briefly discuss some of the additional tests that were conducted to confirm that the main results are robust to changes in the sample, the dataset and the definition of the variables. For succinctness, the results are not reported.

Sample and dataset

The results are robust to dropping all countries that have entries or exits of products representing over 5% of total exports in any given year and to dropping everything prior to 1984, when the original data source changes.[29] Alternatively, the variety measure can be redefined to consider only products with exports over USD 100,000, to avoid possible inconsistencies across countries or periods.[30] The results also hold in this case.

Countries that have experienced several episodes of exchange rate crises could be affecting or even driving the results. Dropping the observations for the lower and upper five centiles of the real exchange rate level or of its volatility, the results hold.

Checking the results with another database is a good way to eliminate issues with the classification system. It is also interesting to check whether the results hold if variety is defined at a higher level of disaggregation. The results were verified using the BACI database (Gaulier and Zignago, 2010), which classifies products at the six digit level using the Harmonised System. Due to the shorter time span available, PMG estimation was not feasible, but the results hold for the exchange rate level with the DFE estimator.

Variable definition

When using Feenstra's extensive margin measure (see Section 3.1) instead of a rough count of exported varieties, the results still hold.

When using the level and the volatility of black market nominal exchange rates (from Re-

[28] In all other specifications, there is also a marked decrease in the coefficient for the level of the exchange rate when going from PMG to DFE estimation but not always between MG and PMG.

[29] This also drops the few pre-1973 observations, when exchange rate behaviour might have been different.

[30] Feenstra et al.'s (2005) dataset includes exports with volumes smaller than this only for the years before 1984 and for some countries after that year.

inhart and Rogoff, 2004), the results hold for the level but are ambiguous for volatility. When the nominal effective exchange rate from the IFS is used, most of the specifications studied confirmed the original results. If an exchange rate volatility measure based on this nominal exchange rate is used, all results continue to hold.

Finally, all results hold if PPP GDP per capita (from the WDI) is used as a control instead of GDP per capita in constant dollars.

Heterogeneity across product types

Two alternative product classifications were considered. Separate count variables were built for the variety of primary goods and for the variety of manufactures (using the Eurostat classification) and for the variety of homogeneous goods and differentiated goods, following Rauch (1999). The coefficients on the exchange rate variables are larger for manufactures than for primaries, and for differentiated relative to homogenous goods, confirming the findings in Table 3.

6 Summary and conclusions

This paper explores the relationship between the exchange rate and export variety at the country level. The empirical results show that a competitive and stable exchange rate is associated with a higher number of exported varieties. This relationship appears to be stronger for products with higher technological intensity or sophistication—the kind of products that are usually associated with technological spillovers and dynamic growth effects. No clear conclusions were drawn regarding whether this relationship differs for developed and developing countries. The results are robust to using different samples, datasets, lag structures, variable definitions and estimation methods (including fixed effects, dynamic GMM, Mean Group and Pooled Mean Group estimators).

The results are consistent with the previous findings reported by Freund and Pierola (2008), Álvarez et al. (2009) and Colacelli (2010). The difference is that here, it is found that *both the level and the volatility* of the exchange rate are related to export variety, and to *total* (country-level) export variety rather than bilateral export variety (for country pairs). The finding that the exchange rate has a stronger impact on the variety of more sophisticated or technologically intensive goods is consistent with the findings of Colacelli (2010), who found that the exchange rate level has a stronger impact on the (bilateral) extensive margin of less substitutable products. This finding could be especially important if not all sectors have the same potential to contribute to output growth (see Young, 1991; Lucas, 1988; Hausmann et al., 2007).

Although they do not seem to drive the results, there is evidence of two types of biases that are often neglected in empirical works: one is due to imposing parameter homogeneity, and the other is caused by strong cross-sectional dependence in the residuals. The latter seems to be different across income levels and across different types of goods, suggesting that there might be unobserved economic factors that differ across those categories, for example, spillovers that are stronger for high-prody goods than for low-prody goods or spillovers that exist across developed countries but not across developing countries.

The bias generated by imposing homogeneity when the slopes are actually heterogeneous seems to be more of an issue for the coefficients for the exchange rate level than for those for volatility.

Although the discussions above show that the results do not seem to be driven by these biases, both issues should be considered more seriously in empirical studies when long enough panels are available.

The main policy implication that can be derived from this evidence is that appreciation that is not in line with fundamentals, and excessive exchange rate volatility should be avoided. Trying to aim for a particular exchange rate level is likely to do more harm than good. Exchange rate volatility can be more effectively targeted by policy and with fewer negative effects, and this could, in turn, prevent large negative level shocks.

There is evidence of an inverted U shaped pattern for the extensive margin of exports as countries develop (Cadot et al., 2011). For advanced countries with an already diversified export structure, the impact of the exchange rate on export variety may not be crucial. However, for developing countries with concentrated export baskets and little export variety, an appreciated and volatile currency could reinforce this situation and harm their long-run growth prospects.

References

(????).

Addison, D. M. (2003). Productivity growth and product variety : gains from imitation and education. Policy Research Working Paper Series 3023, The World Bank.

Álvarez, R., Doyle, M., and López, R. A. (2009). Exchange rate volatility and export margins. *Documentos de Trabajo (Banco Central de Chile)*, (539), 1.

Arellano, M., and Bond, S. (1991). Some tests of specification for panel data: Monte carlo evidence and an application to employment equations. *The Review of Economic Studies*, *58*(2), 277–297.

Arellano, M., and Bover, O. (1995). Another look at the instrumental variable estimation of error-components models. *Journal of econometrics*, *68*(1), 29–51.

Bacchetta, P., and Van Wincoop, E. (2000). Does exchange-rate stability increase trade and welfare? *American Economic Review*, *90*(5), 1093–1109.

Baldwin, R., and Taglioni, D. (2004). Positive oca criteria: Microfoundations for the rose effect. *Mimeo, Graduate Institute of International Studies, Geneva.*

Berman, N., Martin, P., and Mayer, T. (2012). How do different exporters react to exchange rate changes? *The Quarterly Journal of Economics*, *127*(1), 437–492.

Bernard, A. B., Redding, S. J., and Schott, P. K. (2011). Multiproduct firms and trade liberalization. *The Quarterly Journal of Economics*, *126*(3), 1271–1318.

Berthou, A., and Fontagné, L. (2008). The euro and the intensive and extensive margins of trade: Evidence from french firm level data. *Document de travail CEPII*, *6*.

Blackburne III, E. F., and Frank, M. W. (2007). Estimation of nonstationary heterogeneous panels. *Stata Journal*, *7*(2), 197–208.

Blundell, R., and Bond, S. (1998). Initial conditions and moment restrictions in dynamic panel data models. *Journal of econometrics*, *87*(1), 115–143.

Bowsher, C. G. (2002). On testing overidentifying restrictions in dynamic panel data models. *Economics letters*, *77*(2), 211–220.

Breitung, J., and Pesaran, M. H. (2008). Unit roots and cointegration in panels. *The Econometrics of Panel Data*, (pp. 279–322).

Breusch, T. S., and Pagan, A. R. (1980). The lagrange multiplier test and its applications to model specification in econometrics. *The Review of Economic Studies*, (pp. 239–253).

Cadot, O., Carrère, C., and Strauss-Kahn, V. (2011). Export diversification: What's behind the hump? *Review of Economics and Statistics*, *93*(2), 590–605.

Cadot, O., Carrere, C., and Strauss-Kahn, V. (2012). Trade diversification, income, and growth: What do we know? *Journal of Economic Surveys*.

Cavalcanti, T., Mohaddes, K., and Raissi, M. (2011). Commodity price volatility and the sources of growth. Tech. rep., Cambridge Working Papers in Economics.

Chaney, T. (2008). Distorted gravity: The intensive and extensive margins of international trade. *The American Economic Review*, *98*(4), 1707–1721.

Chatterjee, A., Dix-Carneiro, R., and Vichyanond, J. (2013). Multi-product firms and exchange rate fluctuations. *American Economic Journal: Economic Policy*, *5*(2), 77–110.

Choi, I. (2001). Unit root tests for panel data. *Journal of international money and Finance*, *20*(2), 249–272.

Chudik, A., and Pesaran, M. H. (2013). Large panel data models with cross-sectional dependence: a survey. Tech. rep., CESifo Working Paper.

Chudik, A., Pesaran, M. H., and Tosetti, E. (2011). Weak and strong cross-section dependence and estimation of large panels. *The Econometrics Journal*, *14*(1), C45–C90.

Coakley, J., Fuertes, A.-M., and Smith, R. (2006). Unobserved heterogeneity in panel time series models. *Computational Statistics & Data Analysis*, *50*(9), 2361–2380.

Colacelli, M. (2010). Intensive and extensive margins of exports and real exchange rates. Tech. rep., mimeo.

De Hoyos, R. E., and Sarafidis, V. (2006). Testing for cross-sectional dependence in panel-data models. *Stata Journal*, *6*(4), 482.

Dell, M., Jones, B. F., and Olken, B. A. (2008). Climate change and economic growth: Evidence from the last half century.

Eberhardt, M. (2011). Xtcd: Stata module to investigate variable/residual cross-section dependence. *Statistical Software Components*.

Eberhardt, M. (2012). Estimating panel time-series models with heterogeneous slopes. *Stata Journal*, *12*(1), 61–71.

Eberhardt, M., Helmers, C., and Strauss, H. (2013). Do spillovers matter when estimating private returns to r&d? *Review of Economics and Statistics*, *95*(2), 436–448.

Eberhardt, M., and Teal, F. (2010). Aggregation versus heterogeneity in cross-country growth empirics.

Eichengreen, B. (2008). The real exchange rate and economic growth. *Commission on Growth and Development, Working Paper*, (4).

Eichengreen, B., and Leblang, D. (2003). Exchange rates and cohesion: Historical perspectives and political-economy considerations. *JCMS: Journal of Common Market Studies*, *41*(5), 797–822.

Feenstra, R. (1994). New product varieties and the measurement of international prices. *The American Economic Review*, (pp. 157–177).

Feenstra, R., and Kee, H. L. (2008). Export variety and country productivity: Estimating the monopolistic competition model with endogenous productivity. *Journal of International Economics*, *74*(2), 500–518.

Feenstra, R. C., Lipsey, R. E., Deng, H., Ma, A. C., and Mo, H. (2005). World trade flows: 1962-2000.

Freund, C., and Pierola, M. (2008). Export surges: The power of a competitive currency. *World Bank Policy Research working paper*, *4750*.

Funke, M., and Ruhwedel, R. (2001a). Product variety and economic growth: empirical evidence for the oecd countries. *IMF Staff papers*, (pp. 225–242).

Funke, M., and Ruhwedel, R. (2005). Export variety and economic growth in east european transition economies. *Economics of transition*, *13*(1), 25–50.

Gaulier, G., and Zignago, S. (2010). Baci: International trade database at the product-level (the 1994-2007 version). (2010-23).

Hausman, J. A. (1978). Specification tests in econometrics. *Econometrica: Journal of the Econometric Society*, (pp. 1251–1271).

Hausmann, R., Hwang, J., and Rodrik, D. (2007). What you export matters. *Journal of Economic Growth*, *12*(1), 1–25.

Hausmann, R., and Rodrik, D. (2003). Economic development as self-discovery. *Journal of development Economics*, *72*(2), 603–633.

Héricourt, J., and Poncet, S. (2013). Exchange rate volatility, financial constraints and trade: empirical evidence from chinese firms.

Holly, S., Pesaran, M. H., and Yamagata, T. (2010). A spatio-temporal model of house prices in the usa. *Journal of Econometrics*, *158*(1), 160–173.

Hummels, D., and Klenow, P. J. (2005). The variety and quality of a nation's exports. *American Economic Review*, *95*(3), 704–723.

Iacovone, L., and Javorcik, B. S. (2008). Multi-product exporters: Diversification and micro-level dynamics. *World Bank Policy Research Working Paper Series, Vol*.

Im, K. S., Pesaran, M. H., and Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of econometrics*, *115*(1), 53–74.

Jenkins, S. P. (2010). Ineqdeco: Stata module to calculate inequality indices with decomposition by subgroup. *Statistical Software Components*.

Johnson, S., Ostry, J., and Subramanian, A. (2007). The prospects for sustained growth in africa: Benchmarking the constraints.

Klinger, B., and Lederman, D. (2006). Innovation and export portfolios. *World Bank Policy Research Working Paper No. 3983*.

Lewandowski, P. (2007). Pescadf: Stata module to perform pesaran's cadf panel unit root test in presence of cross section dependence. *Statistical Software Components*.

Lin, C.-Y. (2012). Exchange Rate Uncertainty and Trade. *The B.E. Journal of Macroeconomics*, *12* (1), 1–37.

Lucas, R. E. (1988). On the mechanics of economic development. *Journal of monetary economics*, *22* (1), 3–42.

Melitz, M. (2003). The impact of trade on aggregate industry productivity and intra-industry reallocations. *Econometrica*, *71* (6), 1695–1725.

Nickell, S. (1981). Biases in dynamic models with fixed effects. *Econometrica: Journal of the Econometric Society*, (pp. 1417–1426).

Persyn, D., and Westerlund, J. (2008). Error-correction-based cointegration tests for panel data. *Stata Journal*, *8* (2), 232.

Pesaran, M. H. (2004). General diagnostic tests for cross section dependence in panels.

Pesaran, M. H. (2007). A simple panel unit root test in the presence of cross-section dependence. *Journal of Applied Econometrics*, *22* (2), 265–312.

Pesaran, M. H., and Shin, Y. (1998). An autoregressive distributed-lag modelling approach to cointegration analysis. *Econometric Society Monographs*, *31*, 371–413.

Pesaran, M. H., Shin, Y., and Smith, R. J. (2001). Bounds testing approaches to the analysis of level relationships. *Journal of applied econometrics*, *16* (3), 289–326.

Pesaran, M. H., Shin, Y., and Smith, R. P. (1999). Pooled mean group estimation of dynamic heterogeneous panels. *Journal of the American Statistical Association*, *94* (446), 621–634.

Pesaran, M. H., and Smith, R. (1995). Estimating long-run relationships from dynamic heterogeneous panels. *Journal of econometrics*, *68* (1), 79–113.

Rauch, J. E. (1999). Networks versus markets in international trade. *Journal of International Economics*, *48* (1), 7–35.

Reinhart, C. M., and Rogoff, K. S. (2004). The modern history of exchange rate arrangements: a reinterpretation. *the Quarterly Journal of economics*, *119* (1), 1–48.

Robertson, D., and Symons, J. (1992). Some strange properties of panel data estimators. *Journal of Applied Econometrics*, *7*(2), 175–189.

Rodrik, D. (2009). The real exchange rate and economic growth. *Brookings Papers on Economic Activity*, *2008*(2), 365–412.

Roodman, D. (2006). How to do xtabond2: An introduction to difference and system gmm in stata. *Center for Global Development working paper*, (103).

Roodman, D. (2009). A note on the theme of too many instruments. *Oxford Bulletin of Economics and Statistics*, *71*(1), 135–158.

Sarafidis, V., and Wansbeek, T. (2012). Cross-sectional dependence in panel data analysis. *Econometric Reviews*, *31*(5), 483–531.

Saviotti, P. P., and Frenken, K. (2008). Export variety and the economic performance of countries. *Journal of Evolutionary Economics*, *18*(2), 201–218.

Schnabl, G. (2007). Exchange rate volatility and growth in small open economies at the emu periphery. Working Paper Series 773, European Central Bank.

Tang, H., and Zhang, Y. (2012). Exchange rates and the margins of trade: Evidence from chinese exporters. *CESifo Economic Studies*, *58*(4), 671–702.

Tenreyro, S. (2007). On the trade impact of nominal exchange rate volatility. *Journal of Development Economics*, *82*(2), 485–508.

Young, A. (1991). Learning by doing and the dynamic effects of international trade. *The Quarterly Journal of Economics*, *106*(2), 369.

A Dynamic GMM estimates

Table A1: Log of the number of SITC44-digit SITC categories exported

	Low prody exports				*High prody* exports			
	OLS	FE	Difference GMM	System GMM	OLS	FE	Difference GMM	System GMM
Lagged variety	0.899***	0.409***	0.518**	0.783***	0.822***	0.273***	0.819**	0.675***
	(0.000)	(0.000)	(0.027)	(0.000)	(0.000)	(0.002)	(0.026)	(0.000)
RER level	0.221***	0.128*	0.108	0.360**	0.237**	0.0318	0.223	0.472**
	(0.000)	(0.093)	(0.499)	(0.034)	(0.020)	(0.824)	(0.372)	(0.021)
Exchange rate volatility	-0.957***	-0.995***	-1.367***	-1.672***	-1.503***	-1.111***	-2.417**	-3.223***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.033)	(0.009)
N	253	253	205	253	253	253	205	253
AR(1) p			0.070	0.001			0.148	0.037
AR(2) p			0.175	0.235			0.528	0.443
Number of instruments			18	24			18	24
Overidentifying restrictions			5	10			5	10
Sargan p			0.089	0.545			0.886	0.745
Hansen p			0.927	0.827			0.946	0.496
Diff-in-Hansen for levels p			.	0.466			.	0.332
Hansen for assumed exogenous p			.	0.939			.	0.603

Notes. OLS, fixed-effects and GMM estimation for the log of the number of 4-digit SITC categories exported, controlling for GDP per capita, population, trade openness, lagged variety, RER level and exchange rate volatility. Only the last three coefficients are reported. All equations include time dummies and FE also includes country dummies. All regressors are in logs. Country-year data from 1962 to 2000, averaged over four-year periods. Robust standard errors clustered at the country level reported for all estimators. 'AR(1) p' and 'AR(2) p' : p-values for the Arellano-Bond (1999) tests for first and second order error autocorrelation in the differenced equation. 'Sargan p' and 'Hansen p' : p-values for the Sargan and Hansen tests of overidentifying restrictions. 'Diff-in-Hansen for levels p' : p-value for the Difference-in-Hansen test of exogeneity of the instruments for the levels equation. 'Hansen for assumed exogenous p' : p-value for the Hansen test for the instruments assumed exogenous for the Difference-in-Hansen test. One step GMM used because the sample is too small to estimate the optimal weight matrix appropriately. In GMM estimators, population assumed exogenous, the lagged dependent variable assumed predetermined (instrumented with its first and second lags in the differenced equation) and all other regressors assumed endogenous (instrumented with their second and third lags in the differenced equation). Instrument matrix collapsed (see Roodman, 2006).

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. p-values in parentheses.