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FUTURES MARKET VOLATILITY, EXCHANGE RATE UNCERTAINTY AND CEREALS EXPORTS: EMPIRICAL EVIDENCE FROM FRANCE

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Futures market volatility, exchange rate uncertainty and cereals exports: Empirical evidence from France

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Abstract

This paper investigates the impact of both exchange rate and futures price volatility on bilateral cereals exports from France. Using the Poisson pseudo-maximum likelihood (PPML) estimator developed by Santos Silva and Tenreyro (2006) to deal with the problem of zero trade flows when estimating a gravity equation, we show that exchange rate uncertainty has a strong negative impact on French cereals trade. Surprisingly, we find also that higher futures price volatility is associated with increased French cereals exports. Since the PPML method allows for commodity specific estimation of this relationship, we demonstrate that these results are rather commodity-specific and not uniform across individual cereals commodities. For example, we find that realized futures price volatility has a significant and positive impact on French exports of four commodities: barley, durum wheat, maize and oats. We suggest that the storage behaviour of grains elevators and physical traders can explain this seemingly counter-intuitive result. In contrast to currencies, basis variability, i.e. the instability surrounding the spread between commercial spot prices and futures prices, can matter more than price instability, and can lead market participants to reduce their stocks, i.e. to sell, when the level of this instability is high.

Keywords: Exports, exchange rate, futures prices, volatility, Poisson pseudo-maximum likelihood (PPML)

JEL Classification: C23, F14, F31, G13, Q13

1. Introduction

Whereas the impact of exchange rate volatility on the export performance of a given economy or an industry has been widely studied in the literature, not much has been done on the consequences of futures prices volatility. Indeed, most theoretical papers on the subject show that exchange rate volatility may fuel uncertainty about the exporter's expected earnings once denominated in its own currency and, consequently, may decrease the incentive to sell abroad if hedging instruments (forward contracts and

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financial derivatives) are not available. These analyses usually rely on the implicit hypothesis that producing countries can determine the domestic prices of their agricultural products, but have to contend with exchange rate volatility and freight (river barge and ocean freight) price instability (Haigh and Bryant, 2001). However, they largely underestimate the influence of financial commodity markets or so-called "*paper markets*", on physical agricultural markets for at least three well-known reasons.

First, it is generally acknowledged that futures prices can be used by producers and buyers as a proxy for expected cash prices. Based on the Efficient market hypothesis (EMH) framework, this question remains debated. Many studies have tested this hypothesis, but there is no definitive conclusion on the causal relationships between spot prices and futures prices, especially in the case of agricultural products. For example, Kenyon *et al.* (1993) show that contracts with a December maturity for maize (corn) and a November date for soy can be taken as the benchmark for anticipating prices for future harvests, for the period 1952 to 1968, but not after 1973. Using co-integration and causality tests, Ali and Gupta (2011) provide evidence of the links which, over the long term, unite futures prices and spot prices in the Indian National Derivatives Commodity Exchange (NDCEX) market. They show that the prices of futures contracts do not cause, in the Granger sense, spot prices for a number of farm commodities, including castor seed, chickpeas and sugar. However, an even more recent study by Joseph *et al.* (2014), based on the frequency domain causality test and focusing on eight commodities negotiated on the NDCEX and on the Multi Commodities Exchange (MCX), confirms the existence of a long-term link between these two types of price, and demonstrates that a causality link between futures prices and spot prices does indeed exist, especially for farm products such as soya and chickpeas. In other words, in some cases although it cannot be generalized, futures prices can be used to predict spot prices with varying degrees of accuracy. Therefore, from a more operational point of view, and putting aside the controversy over whether they are biased or unbiased predictors, futures prices can provide an indication about the prices that may prevail in the future, given all the currently available information. Therefore, as stated by Black (1976), "*looking at futures prices for various transaction months, participants in this market can decide on the best times to plant, harvest, buy for storage, sell from storage, or process the commodity*".

Second, futures contracts can be used not only to collect information about expected spot prices but also to manage commodity price risk, usually for periods of less than one year. Depending upon the hedger's cash market situation, he will either buy or sell futures. Any producer (end-user) could hedge his long (short) initial position by a short (long) hedge, i.e. by selling (buying) futures contracts that would allow him to fix the price of the commodity he will sell (buy) in the future. A physical trader can use the futures markets in a similar way to protect his margin. This operator buys (sells) without knowing when and at what price he will sell. He may use futures contracts in such a way that any loss (profit), physical or cash market, resulting from a disadvantageous movement of prices would be offset by a profit (loss) on the paper market. A major consequence of this is that almost all commercial agricultural contracts use futures prices as the reference price in order to ensure a correlation between the physical position and the futures position. This is not difficult since futures prices are publicly available on commodity financial exchanges, for example, the Chicago Mercantile Exchange (CME) or the Minneapolis Grain Exchange (MGE). Hence, any cash prices for commodities, such as wheat, corn or soya beans, can be split into an financial price and a price differential which accounts for discrepancies between the future contract, standardized by nature, and the corresponding commercial contract. If this differential is stable, then the correlation between these two prices will tend to be high, and the hedging strategy, all things being equal, will be successful. If, at maturity, the physical price is higher than the futures price, it is in the buyer's interest to take delivery of the commodity using the futures contracts, and in the seller's interest to clear his position, thereby assuring convergence between cash and and futures prices.

Finally, it is important also to consider that futures contracts are not only useful to hedge risk but also to improve inventory management (Tomek and Gray, 1970). Inventories on commodity markets should balance any disequilibrium between a seasonal supply and a continued demand, both inelastic in the short-run, because these inventories create the need for traders to find financial hedging solutions (Peck, 1985), and ensure the con-substantial relationship between physical prices and commodity futures. In the absence of forward commercial deals, carrying stock is based on a fundamental uncertainty, which, all things being equal, imposes an increase on profit margins in order to remunerate the risk. Conversely, it is usually assumed that the use of futures (or forward-type) contracts allows for more competitive prices since the risk premium partly disappears. According to the well-known storage theory (Working, 1948; Brennan, 1958; Fama and French, 1987), the spread between futures prices and cash prices, i.e. the basis, shows whether the market is experiencing shortage or not. If the market is in surplus, then futures prices will be higher than cash prices. In this so-called "contango situation", the return on storage activity for the stock holder is independent of the price level when the sale occurs, provided that there is a convergence with the basis at maturity date. This return then will be equal to the magnitude of the contango minus the storage costs. If there is a commodity shortage, the market will turn to a "backwardation" situation in which any storage activity would lead to an immediate financial loss. In other words, it is wise to store when the market is in contango, and to sell otherwise. Thus, we can say that futures prices influence production and, also, storage activity and export strategies, since most commodities are traded internationally. The contribution of this paper, therefore, is that it takes account of these variables in an analysis of the determinants of French bilateral export flows of five commodities: barley, maize, oats, rice and durum wheat. Indeed, in 2013, France was the second largest exporter of cereals (including rice) behind the United States (U.S.) with an export market share of 9,6 % and the largest European producer with 61 million tonnes produced in 2011.

The rest of the paper is organized as follows. Section 2 reviews the theoretical and empirical literature on the relationship between price and exchange rate uncertainty, and trade. Section 3 presents the data and econometric specification of the estimated models. Section 4 summarizes the results of our gravity equation estimations. Section 5 provides some concluding remarks.

2. Literature review

Numerous studies investigate the relationship between the uncertainty arising from exchange rate volatility, and export flows. Since De Grauwe (1988), the ambivalence of this link has been highlighted: an increase in export prices leads, *ceteris paribus*, to an increase in the rate of profit, risk aversion and production adjustment costs due to exchange rate variability which, in turn, can reduce the total volume of exports of a given country. While early work highlights the existence of a negative relationship (Thursby and Thursby, 1987), more recent studies tend to temper this conclusion. Tenreyro (2007) develops a Poisson pseudo-maximum likelihood (PPML) technique in order to take account of the problems in previous papers generated by heteroskedasticity and zero-trade observations and concludes, for a broad sample of countries from 1970 to 1997, that there is no statistical link between increased currency volatility and reduced international trade. One of the explanations proposed by Tenreyro (2007) for this result, is that derivatives contracts (swaps, options and futures) are now widely used to hedge foreign exchange risk arising from exchange rate volatility, but do not really influence export strategies.

Can this statement be generalized? While early studies attempted to describe the nature of this link for an economy as a whole, more recent analyses focus on particular industries, and demonstrate the specificity of the agricultural sector. For instance, May (2010) investigates the determinants of Thai exports of five agricultural products (corn, rice, rubber, sugar and tapioca), using various explanatory variables including short-run real exchange rate volatility. His analysis reveals a direct link between increased volatility (whatever the measure of volatility adopted - Moving Average (MA) of the standard deviation, residual of AutoRegressive Moving Average (ARMA), AutoRegressive Integrated Moving Average (ARIMA), or General Autoregressive Conditional Heteroskedasticity (GARCH) models of the daily or monthly bath/US dollar real exchange rate), and a reduction in export volume. The author also tests the hypothesis that it is production rather than the firm's export decision that is influenced by exchange rate volatility, but finds little evidence that producers choose to produce less in times of high exchange rate volatility.

A study by Cho *et al.* (2002) confirms that export volumes of agricultural products from the G-10 countries, are more sensitive to the uncertainty resulting from erratic currency movements than exports from other sectors. Kandilov (2008) extends this analysis by comparing exchange rate sensitivity of agricultural trade in the G-10 to that in two other groups of countries - emerging and developing countries. Kandilov (2008) uses a gravity model to test the determinants of bilateral trade in the period 1975-1997 and demonstrates that the link between export volume and real exchange rate variability is weak and not statistically significant if the whole economy is considered. However, he finds that this relationship is more pronounced in the case of agricultural products, although there are important differences between countries. In the context of bilateral trade, the elasticity of the volume of agricultural exports to variability in the exchange rate is much higher for the G-10 than for emerging countries. Kandilov (2008) suggests three explanations for this counter-intuitive result: failure to take account of the non-linearity properties of exchange rate uncertainty; the choice of currency billing, which can significantly change the forex risk faced by each partner country (Goldberg and Tille, 2005); and the existence in the G-10 countries of export subsidies which appear to be statistically dependent on the variability of the exchange rate. However, he shows that if these factors are taken into account, the sensitivity of agricultural exports from developing countries to changes in the exchange rate, is higher than in the case of industrialized countries. Using a similar econometric methodology, Karemera *et al.* (2011) focus on international trade in fruits and vegetables among the OECD countries. They show that in 1996-2002, short and long run currency variability has a positive effect on OECD countries' exports of certain commodities. More specifically, they highlight that although the link between exchange rate volatility and aggregate volume of agricultural exports is statistically proven, it is not uniform, and varies considerably from one commodity to another.

Zhang *et al.* (2010) is the only paper that provides a more comprehensive analysis of the different types of risks an exporter faces. They assume that exchange rate volatility is not the only uncertainty faced by exporters and that multiple volatilities (exchange rate, but also commodity prices and ocean freight costs) need to be considered to explain trade flows. The study includes these variables to explain Brazilian and US soya bean exports between January 1996 and January 2006. The authors provide evidence that exchange rate volatility is statistically significant for explaining Brazilian and US exports, while volatility in soya bean and heating oil prices seems to have no influence on US soya bean exports. They explain this result by suggesting that the availability of commodity derivatives allows exporters to hedge their price risk and render export flows insensitive to volatility. This statement seems reasonable although it ignores the fact that fixed-price hedging strategies traditionally used by producers and users do not protect them against basis risk, i.e. the difference between export spot prices and futures prices. In other words, the protection offered by commodity derivatives instruments, especially futures contracts, is often imperfect. Therefore, findings on the impact of exchange rate volatility cannot be merely extrapolated to commodity price risk, and it is not an unreasonable hypothesis, as mentioned earlier, to consider that the variability of commodity prices may have an impact on export flows, even when hedging instruments are available.

3. Empirical model and data

3.1. The gravity model of trade

Since the pioneering work of Anderson (1979), the gravity equation of trade "has gone from an embarrassing poverty of its theoretical foundations to an embarrassment of riches" (Frankel, 1997, pp. 53). It is generally acknowledged that the gravity equation can be derived from very different models of trade. For example, Anderson's (1979) model assumes that goods are differentiated by country of origin, as in Armington (1969), and that consumers have preferences which are defined over all differentiated products. Bergstrand (1985, 1989) derives the gravity equation directly from a model of trade with monopolistic competition, and demand for variety. Deardorff (2001) indicates that the gravity equation of trade can emerge from a simple Hecksher-Ohlin model. Eaton and Kortum (2002) use a Ricardian type model to derive their gravity equation while Helpman *et al.* (2008) and Chaney (2008) refer to the Melitz (2003) model of firm heterogeneity.

In its general formulation, the model predicts that the volume of trade between two countries is proportional to their gross domestic product (GDP) and inversely proportional to their mutual transaction and transportation costs, as in:

$$T_{ij} = e^{\alpha_0} Y_i^{\alpha_1} Y_j^{\alpha_2} D_{ij}^{\alpha_3} \tag{1}$$

Where Y_i and Y_j represent, respectively, country i's and country j's GDP, D_{ij} is the bilateral distance between country i and country j, which is a proxy for transaction and transportation costs, and α_0, α_1, α_2 and α_3 are the parameters to be estimated.

However, the contribution of the prominent research developed by Anderson and van Wincoop (2003) on the micro-foundations of the gravity equation, highlights the importance of controlling in the model for relative trade costs. Adopting a constant elasticity of substitution demand function, and assuming the Armington's (1969) hypothesis of product differentiation, Anderson and van Wincoop show that trade flows between two countries are determined by a trade barrier relative to the average trade barrier of each country with all its partners. They call this the "multilateral resistance". They indicate also that the empirical gravity literature fails to include any form of multilateral resistance in the gravity equation which results in biased estimates. This omission is described by Baldwin and Taglioni (2006) as the "gold medal mistake". In a short sample period, Baldwin and Taglioni (2006) indicate that this mistake can be resolved by including importer and exporter fixed effects in the gravity equation. However, in the case of a long sample period, we could expect the multilateral resistance to change over time. As a consequence, country fixed effects are not appropriate to evaluate multilateral resistance. In this case, the introduction of importer time-varying fixed effects takes account of the fact that multilateral resistance evolves through time.

The gravity equation of trade has been widely used to investigate the relationship between exchange rate volatility and agricultural trade (Cho *et al.*, 2002; Kandilov, 2008; Karemera *et al.*, 2011; Sheldon *et al.*, 2013). In these studies, the traditional gravity model of trade (equation (1)) is augmented with other factors that may create trade resistance, such as exchange rate volatility, and includes trade costs. We assume that volatility of international futures prices of commodities will have an impact on bilateral trade, and include it in our gravity model of trade. As a consequence, our model yields the following equation (Tenreyro, 2007):

$$X_{Fjkt} = e^{\alpha_0} Y_{Ft}^{\alpha_1} Y_{jt}^{\alpha_2} D_{Fj}^{\alpha_3} e^{(\alpha_4 cont_{Fj} + \alpha_5 lang_{Fj} + \alpha_6 col_{Fj} + \alpha_7 RTA_{Fjt} + \alpha_8 XV_{Fjt} + \alpha_9 PV_{kt})} \epsilon_{ijt} \quad (2)$$

where X_{ijk} is exports of product k from France to country j in t, Y_{it} is French GDP in t, Y_{jt} is country j's GDP in t, XV_{Fjt} is the exchange rate volatility of the Euro against the currency of country j in t, PV_{kt} is the futures price volatility of commodity k in t, D_{Fj} is the bilateral distance between France and country j, and $cont_{Fj}$, $lang_{Fj}$, col_{Fj}, RTA_{Fjt} are dummy variables capturing respectively whether France and country j share a common border, share a common language, were ever in a colonial relationship, and are members or not in a regional trade agreement, ϵ_{ijt} is an error term assumed to be statistically independent of the regressors, and α's are the parameters to be estimated.

Standard practice in the empirical literature on exchange rate volatility and trade consists of log-linearizing equation (2) as follows:

$$\begin{aligned} Ln(X_{Fjkt}) = \alpha_0 + \alpha_1 ln(Y_{Ft}) + \alpha_2 ln(Y_{jt}) + \alpha_3 D_{Fj} + \alpha_4 contig_{Fj} + \alpha_5 lang_{Fj} \\ + \alpha_6 col_{Fj} + \alpha_7 RTA_{Fjt} + \alpha_8 XV_{Fjt} + \alpha_9 PV_{kt} + ln(\epsilon_{ij}) \end{aligned} \quad (3)$$

Note that if importer specific effects θ_j are added to the model to account for multilateral resistance, all time-invariant variables are perfectly collinear with these fixed effects and are removed from the estimated equation. Also, if time-varying country fixed effects are added to take account of the changing nature of the multilateral resistance term, as suggested by Baldwin and Taglioni (2006), all importer time-varying characteristics, such as GDP or exchange rate uncertainty, are perfectly collinear with these effects and are removed from the model.

3.2. Estimation issues

As discussed in the previous section, the most frequent approach in the empirical literature on exchange rate volatility and trade is to estimate a log-linearized model (equation (3)) using the Ordinary Least Squares (OLS) estimator (De Grauwe and Skudelny, 2000; Cho *et al.*, 2002; Kandilov, 2008; Karemera *et al.*, 2011; Sheldon *et al.*, 2013). This estimation procedure entails two serious problems (Santos Silva and Tenreyro, 2006; Tenreyro, 2007). First, the error term in equations (2) and (3) is generally heteroskedastic which means the OLS estimator of the log-linearized model can suffer from serious bias

(Santos Silva and Tenreyro, 2006). Jensen's inequality implies that the expected value of the logarithm of a random variable does not equal the logarithm of its expected value, namely $E(ln(x)) \neq ln(E(x))$. As a consequence, in the presence of heteroskedasticity, the parameters generated by the OLS estimator and interpreted as elasticities can be very misleading (Santos Silva and Tenreyro, 2006). Second, all zero-value observations are dropped from the estimation, creating selection bias. This applies especially when working with disaggregated data such as agricultural products (Haq *et al.*, 2013).

In this paper we investigate the impact of futures price volatility on cereals exports using data at the *6-digit* level of disaggregation in the Harmonized System (HS), where zero trade flows are frequent. Indeed, our dataset contains more than 60 % zero-value observations for French bilateral cereals exports. Thus, dropping zero trade flows would result in selection bias which could lead to inaccurate and biased interpretations of the impact of exchange rate and price volatility on French cereals exports. As Tenreyro (2007) points out, zero trade flows need to be included in the sample when investigating the relationship between exchange rate uncertainty and trade.

Various empirical methods have been implemented to overcome this problem. The first attempt to deal with zero value observations was by Eaton and Tamura (1994). They modelled $ln(\alpha + X)$ and estimate the parameter α rather than setting it arbitrarily (e.g. to 1) using a Tobit approach. The main limitation of this approach is that the estimated parameter α lacks a compelling structural interpretation (Head and Mayer, 2014).

Helpman *et al.* (2008) develop an empirical process based on a two-step Heckman model. The first stage relies on a probit model which estimates the probability that a firm in country i imports a positive amount from country j. The second stage involves estimation of the gravity equation on the positive value observations including selection correction. One limitation of this approach is related to the choice of exclusion variable, which has to be correlated only with firm's propensity to export abroad and not to the current level of exports. Since both equations include country fixed effects, this variable needs to be dyadic in nature (Head and Mayer, 2014). Also, Santos Silva and Tenreyro (2013) emphasize that parameter estimation based on the Helpman *et al.* (2008) model is possible only under the assumption that all the random components of the model are homoskedastic.

In the present paper, as in Sun and Reed (2010), we prefer to use the Poisson pseudo-maximum likelihood (PPML) estimator proposed by Santos Silva and Tenreyro (2006) to estimate the gravity model of trade. In their seminal paper, they provide evidence that this estimator out-performs OLS in the presence of heteroskedasticity. Further, Head and Mayer (2014) show that the PPML remains consistent in the case of over-dispersion in the data. They argue also that for empirical estimation of gravity equations of trade use of negative binomial estimators that rely on a Poisson distribution and estimate the amount of over-dispersion should be avoided. Boulhol and Bosquet (2013) pointout that the Negative Binomial Poisson Maximum Likelihood (NBPML) estimator is not appropriate for application to a continuous dependent variable because the estimates depend heavily on an arbitrary choice of measurement for the dependent variable. Finally, Santos Silva and Tenreyro (2011) using Monte Carlo simulations, confirm that even in the presence of a high frequency of zeros, the PPML outperform alternatives such as OLS or Tobit. These results show that datasets that contain a large number of zeros as in our case, do not undermine the performance of the PPML estimator.

3.3. Measuring exchange rate and futures price volatility

There is no consensus in the empirical literature on how to evaluate exchange rate uncertainty (Clark *et al.*, 2004), with the result that a variety of methods is implemented in the literature. However, the choice of the exchange rate volatility measure can affect the empirical results. We choose two different measures of exchange rate uncertainty in order to test the robustness of our results. In both cases, we use the real rather than the nominal exchange rate, although McKenzie (1999) shows that both methods produce very similar results.

The first measure of exchange rate uncertainty that we compute evaluates the standard deviation of the first difference of the logarithm of the monthly exchange rate between France and its trading partner, as in Dell' Ariccia (1999) and Tenreyro (2007):

$$XV^{S}_{Fjt} = Std.\ dev.[ln(e_{Fjt,m}) - ln(e_{Fjt,m-1})] \tag{4}$$

where $e_{Fjt,m}$ is the real exchange rate between country j and France in month $m = 1, 2, ..., 12$, of year $t = t-5, ..., t-1$. This measure, based on the standard deviation of the bilateral exchange rate, captures short-run volatility (Koray and Lastrapes, 1989; Chowdhury, 1993). We construct this measure for the period 2000 to 2011 using monthly average real exchange rates for the five years previous to year t.

We also implement a measure of long-run exchange rate volatility as a robustness check. Although firms can cover themselves against short-run uncertainty using a hedging strategy, this is more difficult over the long-run. As suggested by McKenzie (1999), firms may be exposed to higher and possible unhedgeable exchange rate risks in the long-run. Thus, in line with the measure proposed by Peree and Steinherr (1989) and applied by Cho *et al.* (2002), Karemera *et al.* (2011) and Sheldon *et al.* (2013), we implement the long-run volatility of the exchange rate between France and its trading partners as:

$$XV^{L}_{Fjt} = \frac{max\ e^{t}_{t-z} - min\ e^{t}_{t-z}}{min\ e^{t}_{t-z}} + \left[1 + \frac{|e_t - e^{p}_{t}|}{e^{p}_{t}}\right] \tag{5}$$

where e_t is the real exchange rate in year t, $max\ e^t_{t-z}$ and $min\ e^t_{t-z}$ denote the maximum and minimum values of the real exchange rate over a time interval of size z up to time t, and e^p_t is the equilibrium exchange rate. The first term in equation (5) captures learned experience, and the second term reflects a correction factor derived from current exchange rate misalignment from its equilibrium value. However, the evaluation of the equilibrium exchange rate remains unresolved in forecasting models. As a consequence, in previous empirical studies, the equilibrium exchange rate is measured as the average of the real exchange rate over the previous years (Peree and Steinherr, 1989; Cho *et al.*, 2002; Karemera *et al.*, 2011; and Sheldon *et al.*, 2013). Following the analysis by Peree and Steinherr (1989), we set the value of z to 5[3].

In contrast to previous studies on agricultural products (Cho *et al.*, 2002; Kandilov, 2008; Zhang *et al.*, 2010; Karemera *et al.*, 2011; Sheldon *et al.*, 2013), we assume also that commodity price volatility has a significant impact on bilateral cereals trade. Indeed, firms are exposed to both exchange rate and commodity price uncertainty on the cereals market. As for exchange rate volatility, the choice of the uncertainty measure can affect the results. Therefore, in order to test the robustness of our estimations, we implement two different measures of commodity prices volatility.

The first measure refers to the standard deviation of the logarithm of the daily commodity futures price as in:

$$PV^U_{kt} = Std.\ dev.[ln(P_{kt,d}) - ln(P_{kt,d-1})] \tag{6}$$

where $P_{kt,d}$ is the futures price of commodity k on day $d = 1, 2, ..., 360$ of year $t = t-1, t-2$. To compute this volatility measure we use the futures price volatility for five commodities: durum wheat, barley, oats, maize and rice. We use daily price data and refer to the two years previous to t to construct this measure of volatility for the period 2000 to 2011.

The previous measure reflects the unconditional realized volatility. To capture price uncertainty *ex ante* and estimate the conditional futures price volatility, we implement a second measure based on a General Autoregressive Conditional Heteroskedasticity (GARCH) model. It is well known that futures prices are characterized by a heavy-tailed probability distribution, which can be dealt with using a GARCH model. This method has been widely used in the empirical literature to model exchange rate uncertainty (Baillie and Bollerslev, 1989; Clark *et al.*, 2004; Kandilov, 2008; May, 2010) or commodity price volatility e.g. crude oil prices (Agnolucci, 2009). This family of models allows volatility clustering and model persistence and serial correlation to be described by the in the volatility dynamics. In our study, we estimate a GARCH (1,1) process using daily data for each of our five commodities. For a given year t, we estimate five

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[3] Previous studies e.g. Karemera *et al.* (2011) and Sheldon *et al.* (2013), provide evidence that the results are robust to the choice of parameter z.

versions of the following GARCH (1,1) model (one for each of the five commodities):

$$ln(P_{kb,d}) = \mu + \phi_1 ln(P_{kb,d-1}) + \epsilon_{kt,d} \tag{7}$$

where $\epsilon_{kj,d} \sim N(0, h_{t,d})$ and the conditional variance is:

$$h_{t,d} = \omega + \beta\epsilon^2_{kbd} + \alpha h_{b,d-1} \tag{8}$$

where P_{ktd} is the futures price of commodity k in da$d = 1, 2, ..., 360$, in year $b = t-1, t-2$. Since we study five different commodities, we estimate 60 (12 years*5 commodities) different GARCH (1,1) models. We use the last estimated conditional standard deviation to proxy for conditional volatility, PV^C_{kt} at the beginning of the next period. For instance, the conditional volatility for 2000 is the estimated conditional standard deviation for the last day of 1999 in the GARCH (1,1) process using data for 1^{st} January 1998 to 31^{st} December 1999.

3.4. Data

The panel dataset used in this analysis covers the period 2000 to 2011, for a sample of 59 of France's trading partners[4]. The variable to be explained is the bilateral exports from France to these 59 countries in five commodities: durum wheat, barley, oats, maize and rice. Hence, our sample consists of 3,540 observations of bilateral exports from France. Information on bilateral exports at the *6 digit* level of the Harmonized System (HS) expressed in current dollars is from the UNcomtrade database.

GDP data expressed in constant US dollar are taken from the World Bank's *World Development Indicators* (WDI). Bilateral nominal exchange rates are taken from the International Monetary Fund *International Financial Statistics* (IFS). This variables is expressed in real terms using the Consumer Price Indexes (CPI) for France and its trading partners, taken from the World Bank WDI. Bilateral distance is computed using the distance in kilometres between France and its trading partners' capital cities. This variable is taken from the CEPII's GeoDist database. Dummy variables capturing common border, common language, and whether France and its trading partners were ever in a colonial relationship are from the CEPII database. Information on regional trade agreements is from the World Trade Organization (WTO)[5].

Information on daily futures prices of durum wheat, barley, oats, maize and rice comes from the *Datastream* database, which offers continuous series for these derivatives instruments listed on the Chicago board of trade (wheat, oats, maize and rice) and the International continental exchange Canada (barley). These prices are defined, for each commodity, as the daily average of the settlement prices of all futures contract traded

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[4] The complete list is contained in Appendix Table A.1.

[5] The list of free trade agreements considered in the analysis is displayed in Appendix Table A.1.

at that time. This method, which relies on the definition of synthetic futures prices time-series, allows us to avoid unrealistic jumps when time-series of futures prices from different contracts are created using a simple roll-over procedure. Summary statistics are presented in Appendix Table A.2.

4. Empirical results

4.1. Baseline results

Table 1 presents the regression results and the test statistics for the OLS and PPML specifications. Columns 2 to 5 report the OLS estimates using the logarithm of trade as the dependent variable. Columns 6 to 9 show the results of the estimations of equation (2), using the PPML method proposed by Santos Silva and Tenreyro (2006), for the whole sample. In all the estimations, we control for product heterogeneity using fixed effects.

A first look at this table shows that OLS and PPML estimates give very similar results for GDP. We find that French bilateral cereals exports depend heavily on French and trading partners' GDP. The elasticity of foreign income ranges from 0.43 to 0.55 according to the PPML estimator, and is around 0.57 according to the OLS estimates. This confirms that higher foreign income stimulates export demand for French cereals. The results indicate also that, as expected, bilateral distance affects French cereals trade, while sharing a common border (known as the border effect) strongly increases French export volumes. We observe that the OLS estimator seems to overestimate all the estimated coefficients especially the variable capturing the border effect.

Several of the coefficients estimated using the PPML method differ significantly from those generated by OLS. Santos Silva and Tenreyro (2006) and Westerlund and Whilhelmsson (2011) attribute these differences to the problem of heteroskedasticity when using the OLS estimator which biases results. This is the case especially for the two measures of exchange rate volatility. Indeed, the results using the OLS estimator indicate that short-run exchange rate volatility has a significant positive impact on French cereals exports. This finding is reversed for the PPML estimator. In that case, we find that both short-run and long-run exchange rate volatility measures have a negative impact on French cereals exports, which confirms previous studies on agricultural products (Cho *et al.*, 2002; Kandilov, 2008: Karemera *et al.*, 2011; Sheldon *et al.*, 2013). However, only the long-run measure is statistically significant at the 5 % level.

Table 1: Baseline results with product fixed-effects

	OLS				PPML			
Y_{Ft}	4.003**	3.812**	4.528**	4.230**	6.493***	6.307***	6.611***	6.261***
	(1.624)	(1.710)	(1.623)	(1.767)	(1.279)	(1.029)	(1.290)	(1.014)
Y_{jt}	0.573***	0.562***	0.572***	0.527***	0.439***	0.556***	0.439***	0.556***
	(0.123)	(0.128)	(0.123)	(0.128)	(0.111)	(0.113)	(0.110)	(0.113)
RTA_{Fjt}	0.727*	0.559	0.732*	0.557	0.985	0.783	0.989	0.783
	(0.428)	(0.434)	(0.428)	(0.434)	(0.819)	(0.831)	(0.819)	(0.831)
XV^S_{Fjt}	2.649***		2.630***		-2.653		-2.595	
	(0.983)		(0.986)		(8.462)	(8.243)		
XV^L_{Fjt}		-0.472		-0.518		-3.530**		-3.534**
		(1.037)		(1.026)		(1.715)		(1.700)
PV^U_{kt}	22.391	18.797			12.722***	3.998		
	(14.387)	(15.043)			(4.791)	(4.544)		
PV^C_{kt}			10.900	9.095			8.376**	3.470
			(9.195)	(9.332)			(4.334)	(3.798)
$contig_{Fj}$	2.081***	2.074***	2.086***	2.072***	1.184***	0.653	1.185***	0.652
	(0.403)	(0.448)	(0.403)	(0.448)	(0.449)	(0.427)	(0.447)	(0.425)
col_{Fj}	1.792**	1.801**	1.780**	1.781**	1.798	1.081	1.799	1.080
	(0.781)	(0.849)	(0.782)	(0.848)	(1.157)	(1.081)	(1.155)	(1.081)
$lang_{Fj}$	-1.014	-1.080	-1.009	-1.067	-0.266	0.406	-0.265	0.407
	(0.613)	(0.664)	(0.613)	(0.663)	(0.727)	(0.624)	(0.727)	(0.624)
D_{Fj}	-0.657**	-0.635**	-0.651**	-0.627**	-0.697**	-0.502*	-0.697**	-0.502*
	(0.275)	(0.287)	(0.274)	(0.286)	(0.338)	(0.283)	(0.338)	(0.283)
$Intercept$	-115.150**	-108.805**	-129.862***	-120.517**	-201.100***	-195.993***	-204.397***	-194.674***
	(45.683)	(48.318)	(45.809)	(49.805)	(35.290)	(28.026)	(35.770)	(28.876)
Product FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1455	1455	1455	1455	3540	3540	3540	3540
R-squared	0.436	0.434	0.436	0.434				
Pseudo LL					-80.169	-79.436	-80.176	-79.435

Clustered standard errors in parentheses.

The single (*), double (**), or triple (***) asterisk denote significance at the 10 %, 5 %, and 1 % levels, respectively.

Returning to the main focus of the paper, we find that, relying on the baseline equation including product fixed-effects, futures price volatility tends to have a strong positive effect on French cereals exports. We find that the realized futures price volatility is significant at the 5 % level, and that the conditional futures price volatility is significant at the 1 % level. To get a better sense of the actual effect of the futures price volatility on French cereals exports, we can implement the impact of increasing both realized and conditional volatility from zero up to its mean value. In this case, when the realized price volatility is reduced from its mean value of 0.018, it increases French cereals exports to around 24 % (12.72*0.018). If we consider conditional volatility (PV_{kt}^{C}), the impact is smaller and about 15 %. Therefore, French cereals exporters are interested in both exchange rate volatility and commodity futures price volatility, although the impact remains lower in terms of intensity. The two types of volatility exhibit opposite signs. Previous studies find that exchange rate volatility has a strong negative impact on exports; however, we find that the two measures of futures price volatility that we implemented, have a significant and positive impact on French exports. The explanation for this positive sign is as follows. In the short-run, price elasticities of supply and demand for agricultural products are low, which means that only storage capabilities can mitigate price volatility. It is important to consider that inventories help producers to reduce the costs of changing production in response to fluctuations in demand. As a consequence, producers determine their production along with their expected inventory holdings (Pindyck, 2001). Accordingly, when inventory levels are low, reflecting a shortage of supply and, therefore, high spot prices, price volatility will tend to be higher, since quantity adjustment is largely constrained on the market. This is confirmed by Symeonidis *et al.* (2012) who empirically investigated the theory of storage using a dataset of physical inventories of 21 different commodities over the period 1993-2011. They found that low inventory levels are associated not only with a backwardation market structure but also with high price volatility for the majority of commodities considered. In fact, the price that the inventory holder has to pay will be equal to the marginal convenience yield, which has three components: the physical cost that holding a given commodity entails; the cost of capital (i.e. the interest forgone by paying for a commodity at time t_0); and the expected decrease in the commodity price which can be calculated precisely using futures prices. Hence, any increase in futures price volatility will bring uncertainty to the value of the convenience yield, i.e. the opportunity cost of holding an inventory. As a consequence, producers or third parties (elevators) will sell their inventories when futures price volatility is high, and will increase their exports.

Nevertheless, although the two measures of price volatility are positive in every estimation, they are significant at the 5 % level in only two out of four regressions using the PPML estimator. To test the robustness of our results, in Table 2 we estimate the baseline equation including both product and country fixed-effects. This controls for unobserved heterogeneity and resolves the “gold medal mistake” (Baldwin and Taglioni, 2006). In this case, all time invariant variables, such as bilateral distance, are removed from the equation.

Table 2: Results with product and country fixed effects

	OLS				PPML			
Y_{Ft}	6.224***	5.987***	6.464***	6.315***	6.531***	6.440***	6.558***	6.494***
	(1.876)	(1.827)	(1.900)	(1.859)	(0.950)	(0.958)	(0.906)	(0.926)
Y_{jt}	0.201	0.742	0.334	0.892	0.142	0.438	0.191	0.474
	(0.871)	(0.842)	(0.855)	(0.829)	(0.640)	(0.577)	(0.663)	(0.603)
RTA_{Fjt}	0.149	0.333	0.161	0.340	1.206***	1.294***	1.211***	1.293***
	(0.316)	(0.331)	(0.319)	(0.333)	(0.229)	(0.232)	(0.231)	(0.235)
XV^S_{Fjt}	1.422		1.384		0.022		0.030	
	(1.152)		(1.161)		(0.584)		(0.602)	
XV^L_{Fjt}		2.247***		2.143***		0.892**		0.832*
		(0.599)		(0.592)		(0.448)		(0.444)
PV^U_{kt}	20.235	25.301*			12.956***	13.584***		
	(14.008)	(13.999)			(4.825)	(4.942)		
PV^C_{kt}			8.336	9.504			8.648**	8.747**
			(8.888)	(8.768)			(4.242)	(4.247)
$Intercept$	-174.068***	-186.641***	-184.680***	-200.045***	-203.193***	-218.392***	-214.789***	-220.602***
	(39.442)	(38.340)	(39.843)	(38.866)	(22.699)	(23.224)	(20.805)	(21.790)
Importer FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Product FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1455	1455	1455	1455	3540	3540	3540	3540
R-squared	0.611	0.614	0.610	0.613				
Pseudo LL					-71.499	-71.481	-71.506	-71.490

Clustered standard errors in parentheses.

The single (*), double (**), or triple (***) asterisk denote significance at the 10 %, 5 %, and 1 % levels, respectively.

This estimation confirms our previous results and shows that French and foreign income strongly increase French cereals exports. This is especially true for the PPML specifications. We find also that, unlike the OLS, the PPML generates a significant and positive coefficient of the variable capturing regional trade agreements. This support the idea that the development of free trade agreements and free trade areas strongly increases bilateral trade.

The positive impact of both realized and conditional futures price volatility on French cereals exports is strongly supported by the PPML estimation with country fixed effects. Indeed, the coefficient associated with price volatility is significant at the 5 % level in every estimation, whatever the measure implemented. Moreover, the estimated elasticities are very similar to those previously computed using the PPML estimator. These results provide evidence of an increase in PV_{kt}^{U} from zero to its mean value and an increase of French cereals exports ranging from 24.6 % to 25.8 %. Exports should increase by around 15.6 % to 15.7 % as a result of an increase in PV_{kt}^{C} from zero to its mean value (0.018). However, the predictions for exchange rate volatility are not supported by the country fixed-effects estimation.

Although the results for the whole sample clearly confirm the positive relationship between futures price volatility and French cereals exports, they are different at the disaggregated level. To test the robustness of our results, we need to estimate commodity-specific gravity equations of trade (Karemera *et al.*, 2011; Sheldon *et al.*, 2013).

4.2. Commodity-specific results

Table 3 reports the estimated parameters from the non-linear form of the model specification presented in equations (2), using the PPML estimator, for each of the five commodities in the sample.

First, the results for all the different specifications and all the different commodities confirm the key role of foreign income in explaining French exports. The results indicate also that regional trade agreements enhance French exports of durum wheat, oats and rice.

Table 3: Commodity-specific results using the fixed-effect PPML estimator

	Durum Wheat (100110)				Barley (100300)				Rice (100610)	
Y_{Ft}	4.861**	4.757**	4.670**	4.519**	7.692***	7.450***	7.172***	6.984***	-7.619	0.821
	(2.384)	(2.409)	(2.129)	(2.153)	(1.634)	(1.640)	(1.682)	(1.690)	(5.919)	(3.564)
Y_{jt}	4.456***	4.188***	4.255***	3.955***	-0.719*	-0.093	-0.782*	-0.162	5.013*	4.522*
	(1.601)	(1.461)	(1.215)	(1.055)	(0.406)	(0.380)	(0.421)	(0.394)	(2.832)	(2.563)
RTA_{Fjt}	1.065**	0.878	1.070**	0.879	0.430	0.557**	0.441	0.564**	2.568***	5.134
	(0.500)	(0.600)	(0.480)	(0.584)	(0.272)	(0.275)	(0.273)	(0.276)	(0.866)	(3.331)
XV^S_{Fjt}	-0.388		-0.372		0.780		0.734		-55.815***	
	(0.241)		(0.245)		(1.401)		(1.396)		(17.097)	
XV^L_{Fjt}		-1.920		-1.978		1.670***		1.626***		11.269
		(1.281)		(1.359)		(0.440)		(0.425)		(10.224)
PV^U_{kt}	-0.631	-0.110			12.513**	13.795**			-22.313	19.326
	(14.478)	(12.272)			(5.549)	(5.836)			(16.879)	(18.824)
PV^C_{kt}			8.050	9.751			12.854***	13.203***		
			(10.891)	(8.823)			(4.669)	(4.734)		
Intercept	-284.255***	-271.286***	-273.258***	-258.021***	-222.072***	-231.521***	-205.895***	-216.630***	48.248	-194.097***
	(63.416)	(61.552)	(54.683)	(52.523)	(38.614)	(39.411)	(39.478)	(40.517)	(121.846)	(67.534)
Importer FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	708	708
Observations	708	708	708	708	708	708	708	708	708	708
Pseudo LL	-14.217	-14.207	-14.215	-14.206	-31.482	-31.437	-31.464	-31.421	-0.778	-0.777

	Oats (100400)				Maize (100510)				Rice (100610)	
Y_{Ft}	4.620**	4.520**	3.902*	3.858**	2.423**	2.421**	3.733***	3.736***	-3.948	1.707
	(2.297)	(1.868)	(2.060)	(1.826)	(1.232)	(1.170)	(1.212)	(1.189)	(2.574)	(1.762)
Y_{jt}	1.667	1.725	1.599	1.629	1.788*	1.853**	2.164**	2.029**	4.681**	4.167*
	(1.579)	(1.401)	(1.500)	(1.364)	(0.994)	(0.882)	(0.911)	(0.817)	(2.370)	(2.318)
RTA_{Fjt}	-0.264	-0.223	-0.324	-0.235	0.567***	0.478***	0.574***	0.431**	2.506***	4.821*
	(0.500)	(0.404)	(0.443)	(0.397)	(0.176)	(0.174)	(0.202)	(0.176)	(0.840)	(2.882)
XV^S_{Fjt}	9.278		5.612		-3.085		-2.274		-39.747*	
	(16.703)		(14.587)		(2.754)		(2.536)		(21.022)	
XV^L_{Fjt}		4.275**		4.012**		-1.179**		-1.626***		10.221
		(1.820)		(1.828)		(0.496)		(0.529)		(9.036)
PV^U_{kt}	35.974***	39.749***			34.990***	30.811***				
	(7.207)	(5.480)			(5.503)	(5.752)				
PV^C_{kt}			20.090***	23.788***			5.208	2.209	10.239	11.443
			(6.184)	(4.116)			(6.048)	(6.163)	(12.773)	(11.502)
Intercept	-198.561***	-202.573***	-175.563***	-180.398***	-134.184***	-134.318***	-180.480***	-175.064***	-46.155	-206.236***
	(43.736)	(30.206)	(38.642)	(28.799)	(17.548)	(17.365)	(20.479)	(19.831)	(58.209)	(53.286)
Importer FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	708	708	708	708	708	708	708	708	708	708
Pseudo LL	-0.886	-0.885	-0.886	-0.885	-17.692	-17.688	-17.716	-17.707	-0.777	-0.776

Clustered standard errors in parentheses.

The single (*), double (**), or triple (***) asterisk denote significance at the 10 %, 5 %, and 1 % levels, respectively.

Our findings show also that long-run exchange rate volatility has a negative and significant effect on French exports of maize. Results concerning short-run exchange rate volatility are significant and negative only for rice. This findings illustrates that both short and long-run exchange rate volatility can affect agricultural trade, which is in line with previous studies (Cho *et al.*, 2002; Kandilov, 2008).

The results for price volatility support the idea that the positive effect is rather commodity-specific and not uniform across individual cereals commodities. Indeed, we find that realized futures price volatility has a significant and positive impact on French exports of only three commodities: barley, oats and maize. The impact is especially strong for French exports of maize. An increase from zero to its mean value (0.017) of unconditional futures price volatility should lead to an increase of 52 % of French maize exports. The results for conditional futures price volatility are also diverse and depend on the commodity under scrutiny. Indeed, the coefficient associated with this variable is significant only for French exports of barley and oats. The different outcomes for durum wheat and rice can be explained as follows. Use of the rice futures market for hedging purpose is limited. Government intervention policies in the rice market are common in order to protect domestic prices from international prices volatility and, therefore, negate the structural need for a futures contract. Futures price volatility then will have a limited impact on exports. With respect to the result for durum wheat, one potential explanation is that French exporters may rely more on the London International Financial Futures and options Exchange (LIFFE) wheat contract to hedge their price risk, than the CBOT contract, which is less related to their commercial needs, but is accounted for in this analysis. This conclusion is confirmed by the results in Table 4. If we consider the LIFFE wheat contract rather than CBOT contract as the reference price for French exports of wheat, we find that the realized unconditional price volatility has a strong positive impact on French exports of wheat.

5. Conclusion

The main purpose of this paper was to investigate the relationship between exchange rate volatility, futures price volatility and French exports of five cereals: durum wheat, barley, oats, maize and rice. To address this question, we ran a product-level analysis using data on French exports, in relation to 5 commodities and 59 trading partners during the 2000-2011 period. Like Santos Silva and Tenreyro (2006), we argue that the standard empirical procedures to estimate gravity equations are inappropriate. Indeed, estimation of gravity trade models using OLS, leads to biased results due to the problem of heteroskedasticity and failure to take account of zero-value observations (Westerlund and Whilhelmsson, 2011). To address these issues, we used the solution proposed by Santos Silva and Tenreyro (2006) and implemented a PPML method to estimate our gravity equations.

Our main results confirm the conclusions in previous studies and highlight that the two measures of exchange rate uncertainty are significant and negative. Thus, exchange

Table 4: Results for durum wheat using the LIFFE as reference price

	PPML			
Y_{Ft}	4.644**	4.482**	4.613**	4.410*
	(2.058)	(1.992)	(2.340)	(2.457)
Y_{jt}	3.573***	3.186***	4.488***	4.248***
	(0.771)	(0.693)	(1.213)	(1.187)
RTA_{Fjt}	1.086**	0.874	1.028*	0.814
	(0.432)	(0.550)	(0.561)	(0.707)
XV^S_{Fjt}	-0.332		-0.408	
	(0.217)		(0.259)	
XV^L_{Fjt}		-2.229		-2.117
		(1.719)		(1.623)
PV^U_{kt}	48.702***	52.580***		
	(12.313)	(16.959)		
PV^C_{kt}			15.369	19.762
			(25.712)	(31.161)
$Intercept$	-253.641***	-234.319***	-278.307***	-263.151***
	(59.542)	(58.160)	(57.810)	(58.421)
Importer FE	Yes	Yes	Yes	Yes
Observations	708	708	708	708
Pseudo LL	-14.195	-14.183	-14.214	-14.203

Clustered standard errors in parentheses.

The single (*), double (**), or triple (***) asterisk denote significance at the 10 %, 5 %, and 1 % levels, respectively.

rate volatility strongly affects French cereals exports. We found also that the two measures of price volatility that we implemented have a significant and positive impact on French exports. Since we use futures prices, these two measures reflect the price volatility anticipated by French producers, who try to manage to their stocks over time. Any increase in the volatility of futures prices will introduce uncertainty into the opportunity cost of holding inventories and will lead producers, elevators or traders to sell their stocks. Our investigation of the commodity-specific results shows that this holds for French exports of barley, durum wheat, oats and maize, but not rice. These different results can be explained by the differences, in terms of liquidity, that can be observed between the futures markets in this study, and also the different pricing strategies used by exporters. This is one of the limitations of the present paper. The liquidity of a futures contract is an essential criterion for its serving as a reference for commercial contracts and this needs to be taken into account in further developments. We also suggest that from a theoretical point of view, market structure volatility (i.e. variation in the contango or backwardation levels, but also the probability to move from contango to backwardation, or vice versa), should be considered alongside futures prices volatility, to explain the link between commodity futures prices and exports.

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Appendix

Table A.1. List of partners and countries and regional trade agreements (RTA)

Country	RTA	Country	RTA	Country	RTA
Albania	Yes (2006)	Greece	Yes	Poland	Yes (2004)
Algeria	Yes (2005)	Hong Kong	No	Portugal	Yes
Australia	No	Hungary	Yes (2004)	Romania	Yes (2007)
Austria	Yes	Ireland	Yes	Russia	No
Belgium	Yes	Israel	Yes	Saudi Arabia	No
Bulgaria	Yes (2007)	Italy	Yes	Senegal	No
Burkina Faso	No	Japan	No	Slovakia	Yes (2004)
Cameroon	Yes (2009)	Korea	Yes (2011)	Slovenia	Yes (2004)
China	No	Latvia	Yes (2004)	South Africa	Yes
Congo	No	Lithuania	Yes (2004)	Spain	Yes
Cote d'Ivoire	Yes (2009)	Luxembourg	Yes	Sweden	Yes
Cyprus	Yes (2004)	Mali	No	Switzerland	Yes
Czech Republic	Yes (2004)	Malta	Yes (2004)	Togo	No
Denmark	Yes	Mauritania	No	Tunisia	Yes
Egypt	Yes (2004)	Mauritius	No	Turkey	Yes
Estonia	Yes (2004)	Mexico	Yes	Ukraine	No
Finland	Yes	Morocco	Yes	United Kingdom	Yes
Gabon	No	Netherlands	Yes	United States	No
Germany	Yes	Nigeria	No	Yemen	No
Ghana	No	Norway	Yes		

Date of the RTA's implementation in brackets (only if after 2000)

Table A.2. Summary statistics

Variable	Obs.	Mean	Std. Dev.	Min	Max
Y_{Ft}	3540	28.387	.0444	28.311	28.442
Y_{jt}	3540	25.34	2.020	21.278	30.213
RTA_{Fjt}	3540	0.539	0.498	0	1
XV^S_{Fjt}	3540	0.032	0.097	0.001	1.088
XV^L_{Fjt}	3540	1.196	0.188	1.005	2.521
PV^U_{kt}	3540	0.019	0.005	0.008	0.031
PV^C_{kt}	3540	0.018	0.007	0.004	0.051
$contig_{Fj}$	3540	0.101	0.302	0	1
col_{Fj}	3540	0.220	0.414	0	1
$lang_{Fj}$	3540	0.254	0.435	0	1
D_{Fj}	3540	7.736	0.946	5.570	9.737

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