The volatility of international trade flows and exchange rate uncertainty*

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Abstract

Empirical evidence obtained from data covering Eurozone countries, other industrialized countries, and newly industrialized countries (NICs) over 1980–2006 shows that exchange rate uncertainty has a consistent positive and significant effect on the volatility of bilateral trade flows. A one standard deviation increase in exchange rate uncertainty leads to an eight per cent increase in trade volatility. These effects differ markedly for trade flows between industrialized countries and NICs, and are not mitigated by the presence of the Eurozone. Contrary to earlier findings, our results also suggest that exchange rate uncertainty does not affect the volume of trade flows of either industrialized countries or NICs.

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1 Introduction

It is generally presumed that an increase in uncertainty will lead to adverse effects on the optimal behavior of economic agents. However, it is not unusual that research yields contradictory results. To that end, one problem that has puzzled many international economists over the last 40 years involves the effects of exchange rate volatility on international trade. Since the breakdown of the Bretton Woods system of fixed exchange rates, the vast literature that has been generated finds that the impact of exchange rate volatility on international trade flows is mixed. Analytical results which predict positive, negative or no effect of exchange rate volatility on the volume of international trade are based on varied underlying assumptions and only hold in certain cases.[1]

Empirical results from studies of this relationship are similarly inconclusive. They are generally sensitive to the choices of sample period, model specification, form of proxies for exchange rate volatility, and countries considered (developed versus developing). However, when we turn to more recent empirical literature, a certain pattern seems to emerge. It appears that exchange rate volatility has a weak but positive effect on industrialized nations' trade flows while it has a negative and pronounced effect on newly industrialized countries' trade flows. For instance, Baum et al. (2004), relying on a nonlinear specification rather than linear alternatives, show that the effect of exchange rate uncertainty on trade flows is positive yet complex.

[1]The research of Clark (1973), Baron (1976) yields negative effects of exchange rate uncertainty on trade, while Cushman (1983) finds similar effects from uncertain changes in the real exchange rate. Cushman (1988) finds negative effects for US bilateral trade flows. Others, including Franke (1991), Sercu , Vanhulle (1992) have shown that exchange rate volatility may have a positive or ambiguous impact on the volume of international trade flows depending on aggregate exposure to currency risk (Viaene , de Vries (1992)) and the types of shocks to which the firms are exposed (Barkoulas et al. (2002)). We should also note models that study the impact of exchange rate uncertainty on trade and its welfare costs within a general equilibrium framework such as Obstfeld , Rogoff (2003), and Bacchetta , van Wincoop (2000).

They also consider the role of income volatility on trade flows among several industrialized countries but its effects are not clear. A subsequent analysis by Grier , Smallwood (2007) reports a significant role for exchange rate uncertainty for developing countries' exports as well as a strong role for income uncertainty in most countries. Their results for developing countries provide support to earlier studies including Arize et al. (2000), Sauer , Bohara (2001) who report negative effects of exchange rate uncertainty on trade flows for countries of similar levels of industrialization.[2] A more recent study by Baum , Caglayan (in press, hereafter BC) also documents that exchange rate uncertainty has a small positive impact on trade flows among industrialized countries. Finally, results obtained from sectoral data generally confirm these findings.[3]

In this paper, using data from both industrialized and newly industrializing countries, we present a broad empirical investigation of the effects of exchange rate volatility on trade flows. Our work is motivated by the propositions of Barkoulas et al. (2002) who claim that one must study both the effects of exchange rate uncertainty on the second moments of trade flows in addition to the (often indeterminate) first moments to fully understand the effects of exchange rate volatility on international trade flows. It is generally expected that exchange rate uncertainty would depress the volume of international trade as the riskiness of trading activity increases. Contradictory empirical findings could be due to the presence of real options inherent in operating in an uncertain environment. Although empirical research shows that

[2] We should also note that researchers implementing gravity models (see Frankel , Wei (1993), Dell'Ariccia (1999), Rose (2000), and Tenreyro (2003) among others) have generally found a negative relationship between exchange rate variability and trade. However, Clark et al. (2004) indicate that 'this negative relationship, however, is not robust to a more general specification of the equation linking bilateral trade to its determinants that embodies the recent theoretical advances in a gravity model' (p. 2).

[3] See, for instance, Caglayan , Di (2008) who study the linkages between exchange rate uncertainty and bilateral U.S. sectoral exports to her top 13 trading partners, some of which are industrialized and others are newly industrializing.

the average impact of exchange rate uncertainty on trade flows is small, the inherent volatility of imports and exports is very high: generally higher than that of GDP. There is no doubt that this high volatility in trade flows has profound effects on exporters' and importers' decision-making processes. In particular, these firms' optimal capital structures, production, capital investment and hiring decisions will be affected because of the uncertainty related to their trade-sensitive activities. According to Barkoulas et al. (2002), trade flow variability can significantly impact the state of the overall level of economic activity resulting in 'financial sector illiquidity, reductions in real output, and/or heightened inflationary pressures' (p. 491).

We aim to provide broad evidence on the effects of exchange rate volatility on the first and second moments of trade flows. To pursue this goal we expand on the scope of BC who report positive and significant effects of exchange rate volatility on trade volatility for the period 1980–1998, ending with the advent of the Euro. The current study extends BC in three ways. First, we study a larger set of countries, with differing degrees of industrialization. This allows us to explore the potential impact of the level of development on the relationship. Second, we scrutinize more recent data for 1980–2006. In the current study, we consider monthly bilateral trade flows for 22 countries: the US, the UK, Austria, Denmark, France, Germany, Italy, Netherlands, Norway, Sweden, Switzerland, Canada, Japan, Finland, Portugal, Spain, Turkey, South Africa, Brazil, Mexico, Peru, and South Korea.[4]

As a third distinction, our empirical methodology in this study substantially improves upon BC's approach as we generate uncertainty proxies for trade and exchange rate volatility using a bivariate GARCH-in-mean (GARCH-M) model rather than the simpler version utilized in the earlier study. This approach allows us to capture both

[4] We do not investigate intra-Eurozone trade flows before 1999 as these trade flows were covered in the earlier study. Issues related to the introduction of the Euro on intra-Eurozone trade are covered in Baldwin (2006).

volatility series more accurately by allowing trade flow volatility to be affected by exchange rate volatility as well as by its lags. Furthermore, we allow trade flows to be affected by exchange rate volatility in the mean equation. Overall, we investigate more than 250 bilateral models for each of the two relationships and discuss our findings across industrialized (*Ind*), newly industrialized (*NIC*), and Eurozone (*Ezone*) countries.

Our analysis reveals two sets of findings. The first set shows that the relationship between exchange rate volatility and the volume of bilateral trade flows is weak. We find that only 38 out of 254 models tested yield statistically significant steady-state effects of exchange rate volatility on the volume of trade flows. The overall effect is on average weak with almost equal numbers of positive and negative significant effects. Carefully considering results for the three groups of trade flows, we conclude that exchange rate volatility does not appear to have a significant impact on either industrialized or *NIC* trade flows.

Our second set of findings provides wide support for the proposition that exchange rate volatility has a significant effect on the *volatility* of trade flows. Considering the full dataset, we find that a one standard deviation increase in exchange rate volatility leads to an 8.16% increase in trade flow volatility. The weakest support to the second hypothesis comes from trade among the newly industrialized economies where only 30% of the models attribute a significant role to exchange rate volatility. This may reflect the presence of managed exchange rate regimes, exchange controls and other market imperfections. In contrast, 52% of the models of trade among industrialized countries produce significant findings. Surprisingly, we find that the single currency of the Eurozone does not lead to a lower volatility of trade flows for exports crossing the Eurozone's borders. 56% of the models representing exports into the Eurozone and 39% of the models representing Eurozone exports exhibit a significant role for exchange rate volatility.

The rest of the paper is constructed as follows. Section 2 discusses earlier research on trade flow variability to motivate the empirical analysis that follows. Section 3 discusses the data set and our empirical model. Section 4 documents our empirical findings while Section 5 concludes and draws implications for future research.

2 Motivation

After the breakdown of the Bretton Woods agreement in 1973, a substantial amount of attention has been devoted to understanding the effects of exchange rate movements on international trade flows. While some developments, including the liberalization of capital flows and the increase in cross-border financial transactions, might have amplified the magnitude and the impact of exchange rate movements, other changes such as widespread use of financial hedging instruments may have reduced firms' vulnerability to exchange rate fluctuations.[5] Nevertheless, it is not obvious how international trade should be affected by changes in exchange rate volatility. The most general argument is that exchange rate uncertainty could depress the volume of international trade by increasing the riskiness of trading activity. The existing empirical evidence is mixed. It appears that trade flows of industrialized countries are not significantly (or at best positively) affected by exchange rate volatility while those of developing countries are negatively affected.

Despite the fact that the volatility of trade flows relative to that of aggregate output is often a factor of two or three times GDP volatility and as volatile as fixed investment spending, to our knowledge, only BC empirically investigate the effects of exchange rate volatility on trade flow volatility along with the first moment

[5] Wei (1999) finds no empirical support for the hypothesis that the availability of hedging instruments reduces the impact of exchange rate volatility on trade.

effects. Their empirical analysis is mainly based on the analytical results of Barkoulas et al. (2002) who propose that to understand the behavior of exporters one should investigate the volatility of trade flows in addition to the level of trade flows. Those authors specify a simple partial equilibrium model to address the linkages between exchange rate volatility arising from three different sources, and the first and the second moments of trade flows. They show that the association between exchange rate uncertainty and trade flows is indeterminate. More interestingly, their model yields unambiguous associations between exchange rate variability and trade flow variability. Depending on the type of shock that affects exchange rates, its impact on trade volatility can be signed.

Notably, the literature on international business cycles yields two other studies, Zimmermann (1999) and Engel , Wang (2007), which incorporate the volatility of trade flows in their analysis. Although their premise is different from ours, as we emphasize that uncertainty in trade flows arising from exchange rate uncertainty will have serious effects on the macroeconomy, it is valuable to discuss these studies. Zimmermann (1999) provides an international real business cycle model to explain the behavior of components of GDP while attempting to rationalize trade flow variability. He points out that trade flows are as volatile as investment and that prior research has not addressed this issue. Using a three-country business cycle model he suggests that while there might be other variables driving the volatility in trade flows, shocks to exchange rates may be important in explaining the observed volatility in trade flows. In his model, exchange rate affects trade flows due to the fact that it takes time for imports to be delivered and the exchange rate relevant for invoicing is determined at delivery; he assumes no hedging. His simulation exercises yield reasonable results relative to observed data, yet the results are sensitive to the choice of parameters.

In their recent contribution to the international real business cycle literature, Engel , Wang (2007) point out that imports and exports tend to be much more volatile than GDP. Noting that trade in durable goods on average accounts for about 70% of imports and exports for OECD countries, they focus on the role of trade in durable consumption goods to explain the volatility of trade flows. In their view, despite the stylized fact that high volatility of exchange rates is a feature of the data, it is unlikely that high volatility in international trade flows arises from volatility of exchange rates. Their simulation results provide support for their model capturing the volatility of trade flows along with several other features of the data such as the fact that imports and exports are both procyclical, and positively correlated with each other.

Our empirical investigation below employs reduced form models to understand how movements in real exchange rates affect the level and the volatility of exports.[6] Although the two hypotheses require investigation of shocks originating from different sources as Barkoulas et al. (2002) claim, we do not attempt to decompose exchange volatility with respect to the types of shocks. In that sense our proxy for exchange rate uncertainty can be interpreted as a composite index for real exchange rate uncertainty. Prior to modeling the two relationships, we scrutinize the time series characteristics of the variables and consider the possibility that they may be cointegrated. Having found no empirical evidence supporting cointegration, we use a model in first differences.

In the next section, we provide information about our data, the mechanism that generates measures of exchange rate and trade volatility and the model that we implement to test the two hypotheses that exchange rate volatility may impact both the first and second moments of trade flows.

[6]Results are presented for exports only for imports of one country are exports of another country.

3 Data

We use monthly data on bilateral aggregate real exports, in each direction, to carry out our empirical investigation. Our dataset spans the period between January 1980 and December 2006 and includes 22 industrialized or newly industrialized countries (*NIC*s): the US, UK, Austria, Denmark, France, Germany, Italy, Netherlands, Norway, Sweden, Switzerland, Canada, Japan, Finland, Portugal, Spain, Turkey, South Africa, Brazil, Mexico, Peru, and South Korea. The latter six countries (Turkey–South Korea) are classified as *NIC*s. Inclusion of both industrialized and newly industrializing countries in the dataset is important given the earlier results that exchange rate volatility may have quite different effects on economies at different levels of development.

These data are constructed from bilateral export series available in the IMF's *Directions of Trade Statistics* (DOTS) and export price deflators, consumer price indices and monthly spot foreign exchange rates from the IMF's *International Financial Statistics* (IFS).[7] The export data are expressed in current US dollars; they are converted to local currency units (LCU) using the spot exchange rate vis-à-vis the US dollar, and deflated by the country's export price deflator to generate real exports. The real exchange rate is computed from the spot exchange rate and the local and US consumer price indices, and is expressed in logarithmic form. We also adjusted the log(real exchange rate) series using seasonal dummies, for the series entering the computation of the real exchange rate are not seasonally adjusted.

Our regression models include measures of foreign GDP extracted from *International Financial Statistics* as a control variable. Because GDP figures are available on a quarterly basis only, we generate a proxy for monthly foreign GDP to match the

[7] Our analysis starts in 1980 with the availability of consistent trade flow data from DOTS.

monthly frequency of export data.[8] To generate a proxy for monthly economic activity, we apply the proportional Denton benchmarking technique (Bloem et al. (2001)) to the quarterly real GDP series. The proportional Denton benchmarking technique uses the higher-frequency movements of an associated variable—in our case monthly industrial production—as an interpolator within the quarter, while enforcing the constraint that the sum of monthly GDP flows equals the observed quarterly total.

In Table 1 we provide summary statistics on the variability of trade flows and how it compares to GDP volatility. These figures provide evidence that the volatilities of both real exports and real imports vary widely across the countries studied. The foot of the table provides values averaged for three groups of countries: newly industrialized countries (NICs), Eurozone countries and non-Eurozone industrialized countries. It is evident that GDP volatility for NICs is higher than for industrialized countries. This holds true for real export volatility scaled by GDP volatility as well. Interestingly, NICs as a group have the lowest real import volatility relative to GDP volatility. Both real exports and real imports of Eurozone to non-Eurozone countries have higher relative volatilities than do non-Eurozone industrialized countries.

3.1 Generating proxies for the volatility of trade volumes and real exchange rates

In order to proceed with our investigation of the effects of real exchange rate uncertainty on the volume and volatility of trade flows, we must produce proxies that capture the volatilities of the exchange rate and trade flow series. We implement a bivariate GARCH-in-mean (GARCH-M) system for the real exchange rate and the

[8]It might be possible to use monthly industrial production itself to generate such a proxy, however given we are using bilateral trade data we chose not to use industrial production in that context, as such a measure provides a limited measure of overall economic activity.

volume of trade flow data to estimate these volatility measures.[9] This strategy allows us to estimate internally consistent conditional variances of both series which we use as proxies for exchange rate and trade flow volatility.[10,11]

Prior to estimation of the GARCH-M system, we scrutinize the time series properties of the data to determine the appropriate characterization of the order of integration of each series. We subject these series to a rigorous analysis of their order of integration, and find that 426 of 462 series can be characterized as unit root ($I(1)$) processes. Next, we explore whether those exchange rate and trade flow series that exhibit $I(1)$ characteristics exhibit a long-run cointegrating relationship. The data do not provide any support for the existence of cointegration between the variables of interest. Given this information, we include only the first differences of log trade volume, log real exchange rate and log foreign country GDP in our bivariate system as in BC. Using a similar notation to theirs, we denote the first differences of the log real export series, log real exchange rate and log real GDP by s_t, x_t and y_t, respectively. Our bivariate GARCH-M model for bilateral trade volumes and real exchange rates takes the following form:

$$s_t = \theta_0 + \theta_1 s_{t-1} + \theta_2 x_{t-1} + \theta_3 \sigma^2_{s_t} + \eta_t + \theta_4 \eta_{t-1}, \quad (1)$$

$$x_t = \vartheta + \vartheta_1 s_{t-1} + \vartheta_2 x_{t-1} + \vartheta_3 y_{t-1} + \vartheta_4 \sigma^2_{s_t} + \omega_t + \vartheta_5 \omega_{t-1}, \quad (2)$$

$$\mathbf{H}_t = \mathbf{C}'\mathbf{C} + \mathbf{A}' u_{t-1} u'_{t-1} \mathbf{A} + \mathbf{B}' \mathbf{H_{t-1}} \mathbf{B}. \quad (3)$$

In equation (1), s_t is defined as a function of its own lag, lagged trade volume and its own conditional variance ($\sigma^2_{s_t}$) as well as a first-order moving average innovation.

[9]It may be possible to use a moving standard deviation of the series to compute such a proxy. However, this approach induces substantial serial correlation in the constructed series. Given that trade and exchange rates are interrelated, we can capture the interrelationships between the two volatilities by employing a bivariate system.

[10]BC utilize a standard bivariate GARCH system which does not allow the mean equations to be affected by exchange rate volatility.

[11]Grier et al. (2004) and Grier , Perry (2000) use a similar approach to jointly model the effects of uncertainty inflation and output growth in a bivariate GARCH-in-mean framework.

Likewise, x_t in equation (2) is modeled as a function of its own lag, the lagged real exchange rate, lagged output (y_{t-1}) and the conditional variance of the exchange rate $(\sigma^2_{s_t})$, with a first-order moving average innovation. The vector of innovations is defined as $u_t = [\eta_t, \omega_t]'$. The diagonal elements of $\mathbf{H}_t$ are the conditional variances of Δ log real exchange rate, $\sigma^2_{s_t}$ and Δ log trade volume, $\sigma^2_{x_t}$ respectively.

Following Karolyi (1995), the matrix $\mathbf{C}$ is parameterized as lower triangular while matrices $\mathbf{A}$ and $\mathbf{B}$ are 2×2 matrices, so that there are eleven estimated parameters in equation (3). We assume that the errors are jointly conditionally normal with zero means and conditional variances given by an ARMA(1,1) structure as expressed in equation (3). The structure of Equation (3) allows the conditional variance of the exchange rate to have an effect on that of trade flows and *vice versa*. Overall, this system of equations provides a well-specified minimal framework for estimation of the mean and the volatility of trade flow series in equations (4) and (5) below, where we take into account more complex dynamic relationships between the variables. The system is estimated using the multivariate GARCH-M-BEKK model, implemented in RATS 7.10.

3.2 Modeling the dynamics of the mean and the variance of trade flows

In this section, we describe the two reduced form models that we use to investigate the impact of exchange rate volatility on the first and second moments of trade flows. For both sets of relationships we must introduce several lags of the independent variables to capture the delayed impact of these variables on the dependent variable. Earlier research has shown that changes in income and exchange rate volatility may have delayed effects on trade flows, involving up to six periods' lags at a monthly frequency. We must also take into account the dynamics of the dependent variable which may arise due to the time lags associated with agents' decisions to purchase and the completion of that transaction. These two issues require an estimated model which is computationally tractable and yet sufficiently flexible to capture the dynamic pattern that exists between the variables. We employ a simple distributed lag structure which has been successfully implemented in similar contexts.

To investigate the impact of exchange rate volatility on trade flows, we employ the following distributed lag model:

$$x_t = \alpha + \varphi \sum_{j=1}^{6} \delta^j x_{t-j} + \beta_1 \sum_{j=1}^{6} \delta^j \sigma^2_{s_{t-j}} + \beta_2 \sum_{j=1}^{6} \delta^j y_{t-j} + \beta_3 \sum_{j=1}^{6} \delta^j s_{t-j} + \xi_t \tag{4}$$

where the model contains the first difference of log real GDP (denoted y_t) of the importing country as a control variable in our basic equation.[12] The lag parameter δ is set to a specific value to ensure dynamic stability in that relationship while we estimate a single coefficient associated with each of the variables expressed in distributed lag form: φ, β_1, β_2 and β_3, respectively.[13]

To study the second hypothesis that exchange rate volatility may have a significant impact on the variability of trade flows, we employ a similar model

$$\sigma^2_{x_t} = \alpha + \lambda \sum_{j=1}^{6} \delta^j \sigma^2_{x_{t-j}} + \phi_1 \sum_{j=1}^{6} \delta^j \sigma^2_{s_{t-j}} + \phi_2 \sum_{j=1}^{6} \delta^j y_{t-j} + \phi_3 \sum_{j=1}^{6} \delta^j s_{t-j} + \zeta_t \tag{5}$$

where on the left hand side we have trade flow variability, $\sigma^2_{x_t}$. In this model, we are interested in the sign and the significance of the coefficients of exchange rate

[12] We use the modified log-periodogram regression test of Phillips (2007) to examine if y_t exhibits $I(1)$ vs. $I(0)$ properties. These test statistics provide clear evidence that the log-differenced y_t series is stationary. Given the uniformity of the unit root test results, which are available from the authors, we do not tabulate them.

[13] We tried different values for δ in the range of (0.3, 0.5). These results, which are available from the authors upon request, are similar to those we report here for $\delta = 0.4$. We also experimented with lag length, and found that six lags were sufficient to capture the series' dynamics. Given the monthly frequency of the data and the large number of coefficients on highly correlated regressors to be estimated, we did not find that an unconstrained distributed lag approach produced usable nor dynamically stable estimates.

volatility, $\sigma^2_{s_{t-j}}$.[14] As control variables, we introduce the differences of the log real exchange rate (s_t) and log real GDP (y_t) of the importing country into this basic relationship. Similar to the model of equation (4), we choose the lag parameter δ to ensure dynamic stability.

4 Empirical results

4.1 Timeseries properties of the data

Prior to carrying out the bivariate GARCH system of equations, we test each series for a unit root using the modified log-periodogram regression test of Phillips (2007) as implemented in Baum , Wiggins (2001). We find that the overwhelming proportion of the log exchange rate, log trade flow and log real GDP series exhibit $I(1)$ characteristics. The analysis is then conducted using those series that are clearly classified as $I(1)$ while we drop the remaining series. In total, this approach causes us to discard 36 potential country-pairs out of a possible 462 cases.[15] We also test if there is a long run relationship between the exchange rate, trade flow and foreign income series using an Engle–Granger regression on levels of the series.[16] We fail to establish a cointegrating relationship between any of the pairs involved. Hence, we use the first differences of the series in our bivariate GARCH-M system as discussed in the previous section.

[14]Note that $\sigma^2_{s_t}$ is a generated regressor, as is the dependent variable in this equation, which is an augmented autoregression. The presence of two generated regressors, each produced in the nonlinear context of a multivariate GARCH specification, may have consequences for the conventional estimates of coefficients' standard errors. To our knowledge the econometric literature has not addressed this problem.

[15]We do not estimate all 462 possible country-pair models, as described below, as some country-pairs relate to intra-Eurozone trade, which is excluded from the analysis.

[16]Unlike Johansen's maximum likelihood-based method, the Engle–Granger methodology is appropriately robust to deviations from normality of the underlying series.

4.2 Generating proxies for conditional variance

We have employed the bivariate GARCH-M model described above to estimate $\mathbf{H_t}$, the conditional covariance matrix of log real trade flows and log real exchange rates, for each point in time.[17] Although the conditional covariance between GARCH-M errors is not currently employed in our analysis, it is important to note that this measure of contemporaneous correlation is generally nonzero, signifying that estimation of Equations (1–3) as a system is the preferred approach to modeling the two conditional variances.

We present two summary statistics, mean and interquartile range (IQR), for the three elements of the conditional covariance matrix. These are shown in Tables 2–4 for three exporting countries: the US, France and Brazil. Similar statistics for the other 19 exporting countries are available on request. In each of these tables, it is evident that the conditional variances of trade flows—in terms of either mean or interquartile range across the sample—differ quite widely across partner countries. The conditional variances of real exchange rates for the US are similar for most countries with the exception of Canada (perhaps reflecting the close economic relationship between those NAFTA partners) with those related to NICs being an order of magnitude larger (except for South Korea). The mean conditional covariance for the US is negative and positive in almost half of the cases.

For French real exports (Table 3) to countries outside the Eurozone, the lowest volatility of trade flows is related to US and UK imports. In terms of exchange rate volatility, the lowest values occur for Denmark (whose currency may closely track the Euro) and Switzerland, a partner in bilateral relations with the European Union. For Brazil (Table 4), one of the NICs in our sample, both volatility measures are an

[17] Detailed estimation results from the bivariate GARCH-M models are available on request from the authors.

order of magnitude higher than that of the US or France, with the mean conditional covariance uniformly positive.

4.3 Estimation results

In this section we first discuss our estimation results on the effects of exchange rate volatility on trade flows: *the first moment effects*. Next, we focus on the impact of exchange rate volatility on trade flow volatility: *the second moment effects*. For each hypothesis, we provide the effects of exchange rate volatility for the full data followed by several categories of bilateral flows defined in Section 4.3.2 below. The analyses of these categories of bilateral flows are important as they can shed light on various questions including whether exchange rate volatility affects the first and the second moments of trade flows differently across trading country-pairs with different levels of industrialization. Throughout the presentation of our results, we concentrate on the sign and the significance of point and interval estimates of β_1 and ϕ_1, obtained from equations (4) and (5), along with their corresponding steady state values, to explain the effects of exchange rate volatility on the mean and variance of trade flows, respectively. We compute the steady state values $\hat{\beta}_1^{SS} = \left(\hat{\beta}_1 \sum_{j=1}^{6} \delta^j\right) / \left(1 - \hat{\psi} \sum_{j=1}^{6} \delta^j\right)$ and $\hat{\phi}_1^{SS} = \left(\hat{\phi}_1 \sum_{j=1}^{6} \delta^j\right) / \left(1 - \hat{\lambda} \sum_{j=1}^{6} \delta^j\right)$.

We initially construct a table, which is available upon request, containing regression results for all countries in our dataset. Using this table, we document and discuss in Tables 5–8 the impact of exchange rate volatility on trade, $\hat{\beta}_1$, the steady state impact of exchange rate volatility on trade flows, $\hat{\beta}_1^{SS}$, the impact of exchange rate volatility on trade volatility, $\hat{\phi}_1$ and the steady state impact of exchange rate volatility on trade volatility, $\hat{\phi}_1^{SS}$, as well as their corresponding p-values for each country-pair. In our discussions, we provide special attention to results for several different categories which are constructed as described below.

4.3.1 First-moment effects of exchange rate volatility on trade flows

Table 5 presents summary information on the first moment effects of exchange rate volatility on trade flows for the full sample. The first column gives the exporting country in the order above. The second column shows all admissible relationships that we investigate for each country. This figure can at most be 21 for each country, but it is less than that for two reasons: (i) countries in the Eurozone (*Ezone*) are not counted as trading with other *Ezone* members even before 1999,[18] and (ii) our model of a particular bivariate GARCH-M system relationship did not converge for a number of the country-pairs.[19] So, for instance, we investigate as many as 20 models for Japan and 18 for the US, but as few as five models for Portugal and South Africa.

In columns three and four, we display the median value of $\hat{\beta}_1$ when that coefficient is significant, followed by the number of occurrences that the impact of exchange rate uncertainty on trade flows is distinguishable from zero at the five per cent level. The fifth and sixth columns present equivalent information for the steady-state coefficient, $\hat{\beta}_1^{SS}$.[20] The standard deviation of the conditional variance, $\bar{\tau}_s$, is calculated from the timeseries of each bilateral relationship. The figure given in column seven is averaged over those trading partners for which we have estimated significant steady-state effects. Finally, the last column gives the median impact of exchange rate uncertainty on trade flows in percentage terms, computed as $100 \times \left(\hat{\beta}_1^{SS} \cdot \tau_s\right)$ for each bilateral relationship possessing a significant steady-state effect. The impact measure expresses the median impact (over that subset of trading partners) of a one standard deviation increase in exchange rate volatility on the transformed log level

[18]See Baum et al. (2004) or Baum , Caglayan (in press) who investigate the 1980–1999 period in their analysis for most of the *Ezone* countries.

[19]Of the 368 GARCH-M systems estimated, 114 failed to converge, and are excluded from further analysis.

[20]Although the numbers of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients are equal, that is not necessarily the case.

of trade flows.

Inspection of column three of Table 5 shows that exchange rate volatility does not generally affect international trade. We find that only in 38 out of 254 possible models (fewer than 15% of the cases) does exchange rate uncertainty has a significant impact on trade flows. In light of earlier research, this result is not surprising as several researchers have concluded that there is little or no systematic relationship between the two variables. However, vis-à-vis earlier research, these results more conclusively refute the claim that exchange rate uncertainty affects trade flows. At the 5% level of significance, there are 21 positive estimates and 17 negative estimates of the steady-state value, $\hat{\beta}_1^{SS}$.[21] The greatest number of significant effects is registered by Brazil (8) and Peru (7). Given this information, it may seem that exchange rate volatility mainly affects the trade flows of NICs. However, when we inspect the remaining NICs, we find no significant impact of exchange rate volatility on trade flows for Turkey, South Africa and Korea, and for Mexico the relationship is significant in only one instance. Overall these findings provide very strong empirical support to the idea that exchange rate uncertainty does not have an impact on trade flows of industrialized countries. Furthermore, the evidence convincingly refutes the claim that exchange rate uncertainty affects NICs' trade flows.

4.3.2 First-moment effects for categories of bilateral trade flows

To understand these results, we consider categories of bilateral trade flows involving three groups of countries: Eurozone economies (*Ezone*), industrialized economies (*Ind*) and newly industrialized economies (*NIC*). We construct seven categories of bilateral flows based on the results that generated Table 5: (i) Non-*Ezone* to Non-*Ezone*, (ii) *Ezone* to Non-*Ezone*, (iii) Non-*Ezone* to *Ezone*, (iv) *Ind* to *Ind*, (v)

[21] At the 10% level of significance, there are 51 significant steady-state estimates: 25 positive and 26 negative.

Ind to *NIC*, (vi) *NIC* to *Ind*, and (vii) *NIC* to *NIC* countries. These groups are not mutually exclusive, but defined in order to provide a sharper view of how the relationships may vary across different categories of country-pairs' trade. A summary view of the findings for these categories is provided in Table 6, which gives the percentage of models yielding significant findings and the median significant $\hat{\beta}_1$ for each category. Detailed results by category are presented in Tables 9–15 in the Appendix, respectively.

The percentage of models that support a significant role for exchange rate volatility on trade flows from *Ind* to *NIC* is only 10% while that from *NIC* to *Ind* countries is 33%. However, this latter figure is somewhat artificial as it is driven mainly by results from Brazil and Peru, whose exports seem to be highly affected by exchange rate volatility when they trade with industrialized partners. These significant effects may be due to excessive volatility during the currency crises of the mid- to late-1990s, which severely affected most Latin American countries. When we consider trade flows to and from *NIC*s it is clear that exchange rate uncertainty does not impede *NIC*s' trade. For trade flows among industrialized countries, only 11% of the models estimated signal an important role for exchange rate volatility. For all categories except *NIC*–*Ind* and *NIC*-*NIC* trade flows are weakly and positively affected by exchange rate volatility. In the case of trade from *NIC*s to *Ind*, trade flows are negatively affected by exchange rate volatility. For the case of trade from *NIC*s to *NIC*s, we find no effect of exchange rate volatility on trade flows. These results provide firm evidence that exchange rate volatility does not affect international trade, including trade among NICs and bilateral trade between industrialized countries and NICs.

4.3.3 Second-moment effects of exchange rate volatility on trade flows

Table 7 presents a summary of our findings for the hypothesis that exchange rate volatility affects trade flow volatility. The table is constructed similar to Table 5 with one exception. Because the dependent variable in Equation (5) is the variability of trade flows and not its logarithm, we present a percentage impact measure in the last column. In other words, we compute the impact of a one standard deviation increase in exchange rate uncertainty on trade flow volatility for each bilateral relationship as $100 \times \left(\hat{\phi}_1^{SS} \cdot \tau_s / \bar{\sigma}^2_{x_t}\right)$ where $\bar{\sigma}^2_{x_t}$ is the volatility of trade flows for that country-pair. The impact measure displayed is the median of those values across trading partners, and is expressed as a percentage of the exporting country's mean volatility of trade flows.

Inspecting Table 7, we see that exchange rate volatility has an economically meaningful impact on the *volatility* of trade flows. We find that in 119 out of the 254 models tested (47%) there is support for a statistically significant steady-state effect of exchange rate volatility on trade volatility at the 5% level. We obtain a positive and significant relationship in 104 models and a negative and significant relationship in only 15 models.[22] We detect the greatest number of significant effects for the US (12), UK, Denmark and Japan (10 each). For several countries, the table registers only two instances where the effects of exchange rate volatility is significant. However, given that the number of bilateral relationships for each country differs, the number of significant models masks which countries' trade flows are most affected. To overcome this problem, we compute the percentage of cases where the effect is significant for each exporting country. Denmark records the highest fraction of significant cases (83% of the estimated models) followed by the US (67%). The countries whose trade

[22] At the ten per cent level of significance, we find 137 significant coefficients: 112 positive, 25 negative.

volatilities are least affected by exchange rate volatility are Finland and Austria. For these countries, trade flow volatility is affected by exchange rate volatility in 20% and 37% of the estimated models, respectively. For all other countries the significant models exceed 40% of possible cases.

Beyond the presence or absence of a statistically significant relationship, it is important to consider the magnitude of the effects: the average impact of exchange rate volatility on trade volatility. For exporting countries, the impact of a one standard deviation increase in exchange rate volatility on trade volatility has a median estimated value of 8.16%, with first and third quartiles of 2.69% and 19.38%, respectively. A significant negative median value of any meaningful size is only found for Mexico. These findings have strong implications for the behavior of exporters: increases in the volatility of trade flows will have marked effects on the value of their real options to export. Although those option values are only one of the countervailing forces on the volume of trade, the effects we detect are sizable enough to play a role in the expansion or contraction of trade which would affect managers' decisions regarding fixed investment, hiring or financing.

4.3.4 Second-moment effects for categories of bilateral trade flows

We now provide an overview of the effects of exchange rate volatility on trade flow volatility to scrutinize our findings for each category of country-pairs defined above. In particular, Table 8 gives the percentage of models yielding significant findings and the median significant $\hat{\phi}_1$ for each category. Detailed results by category are presented in Tables 16–22 in the Appendix, respectively. Significant effects of exchange rate volatility on trade flow volatility are captured for 30–56% of the models, with trade volatility among industrial countries showing important effects for 52% of the models. These two sets of findings show that exchange rate volatility affects trade flow volatility similarly across *Ezone* and other industrialized countries. The single currency does not lead to a lower volatility of trade flows. Given these results, it is evident that trade volatility is a problem for *Ezone* countries as it is for the remaining industrialized countries in our data set.

The weakest support is obtained from newly industrialized countries' bilateral trade. For this group, we have only 10 models to scrutinize and only three models yield significant results at the 5% significance level. This is understandable given that NICs trade mainly with industrialized countries and they have little trade with each other. When we look at the same effect for trade from *NIC* to *Ind* countries, the significance of exchange rate uncertainty on trade flow volatility increases to 47% of the models. For trade flows from *Ind* countries to *NIC*, the same ratio is 38%. These percentages are the two lowest figures across all groups, perhaps reflecting the managed exchange rate regimes and exchange controls prevalent in many NICs.

When we compare size effects of exchange rate volatility across groups, we find the largest impacts are related to flows among NICs (36%), between NICs and industrialized countries (13%) as well as non-Eurozone exports to Eurozone partners (11%). The smallest effects on trade flow volatility are registered in the larger group of non-Eurozone countries' internal flows (3.6%). Nevertheless, these effects are sizable enough to impact the decision making processes that lead to capital investment, hiring and financing practices of exporters and importers.

Given that the overwhelming majority of the models provide a positive relationship between exchange rate uncertainty and the volatility of trade flows, it is useful to consider how our findings reconcile with predictions of Barkoulas et al. (2002). In light of our results, it seems that either (i) the preponderance of shocks to the exchange rate process are associated with shocks to the fundamentals or (ii) fundamental shocks are larger or have a greater impact on the real exchange rate than other shocks.[23] However, we must also note that possibly due to the use of an aggre-

[23] Given the policy implications of knowing the type of shocks that lead to trade volatility, we think

gate proxy for exchange rate volatility, different types of shocks to the exchange rate may cancel one another and leave us with a lower number of models where the relationship is significant. This may be a likely scenario for developing countries where shocks from different sources affect the exchange rate, explaining the comparatively low success of our models.

We would like to note that despite the simplicity of our models, we were able to detect significant effects of exchange rate volatility on trade flow volatility. What is surprising is that this relationship seems to hold across almost all countries in our data. Being a Eurozone member country does not seem to mitigate such effects. Our results suggest that the potential effects of *volatility* of trade flows on the macroeconomy are important, and decision makers must seriously consider their wider effects on economic agents.

5 Conclusions

In this paper, we investigate two sets of hypotheses that fluctuations in exchange rates affect the level and volatility of trade flows employing a broad set of bilateral trade data. Our dataset contains information on bilateral trade flows for both industrialized and newly industrialized countries over the period 1980–2006. We document that the preponderance of bilateral trade volume and real exchange rate series can be characterized as unit root processes which do not enter into a cointegrating relationship. We then generate internally consistent measures of trade and exchange rate volatility employing a bivariate GARCH-M methodology. Using these proxies, we investigate the impact of exchange rate volatility on the mean and the variance of trade flows. Our first set of results is conclusive: exchange rate volatility does

that it would be useful to verify our conjecture. However, given the complexity of our methodology, we leave this issue for future investigation.

not have a significant impact on either industrialized or newly industrialized country trade flows, as a very small fraction of the estimated models present significant relationships. Hence, we believe that it is not productive to devote more effort to further investigate this hypothesis.

Our second set of findings supports and broadens the results in Baum , Caglayan (in press). We show that bilateral trade volatility is higher than GDP volatility for both developed and newly industrialized countries, a stylized fact documented by Engel , Wang (2007) and Zimmermann (1999) for aggregate trade flows. We compute significant effects of exchange rate volatility on trade flow volatility, where these effects vary with respect to the state of industrialization of the exporting and importing countries. Overall, we find that the median impact of a one standard deviation increase in exchange rate volatility on trade volatility is an economically meaningful 8.16% increase. In addition, we investigate if the relationship holds in seven subset groups of bilateral country-pairs, and find significant differences in its prevalence and strength among them. Notably, despite the claims of the proponents of the Eurozone, it seems that the single currency has failed to play a significant role in reducing trade volatility for flows crossing the Eurozone's borders, as those flows show significant (if modest) effects in both directions.

We would like to close our discussion noting that it might be more useful to shift our attention to understanding the volatility of trade flows rather than their levels. Despite the simplicity of our models, we detect significant effects of exchange rate volatility on trade flow volatility. What is surprising is that this effect seems to exist across almost all countries in our data; being a Eurozone member country does not seem to mitigate such effects. Given the size of the effects of exchange rate volatility on trade flow volatility, we can readily conjecture that the behavior of exporters and importers would be significantly affected. We believe that further analysis of the volatility of trade flows, and its real effects on economic activity, would seem to be warranted: particularly in light of the marked differences across categories of countries with different levels of industrialization.

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Table 1: GDP, Export and Import Volatility

	GDPvol	RXvol	RX/GDP	RMvol	RM/GDP
US	0.2417	0.4383	1.8138	1.2947	5.3573
UK	0.2017	0.4600	2.2806	1.2835	6.3628
AT	0.1910	0.5150	2.6961	1.6314	8.5412
DK	0.1629	0.5524	3.3913	1.3229	8.1219
FR	0.1591	0.4426	2.7819	1.7402	10.9374
DE	0.2019	0.4884	2.4198	1.6170	8.0111
IT	0.1506	0.6272	4.1648	1.6185	10.7476
NL	0.2160	0.5871	2.7187	1.6639	7.7047
NO	0.2392	0.6982	2.9193	1.3909	5.8153
SE	0.1837	0.6345	3.4546	1.2261	6.6756
CH	0.0645	0.4108	6.3725	1.3695	21.2450
CA	0.2431	0.5279	2.1711	1.5056	6.1925
JP	0.1703	0.3773	2.2159	1.3332	7.8300
FI	0.1970	0.9256	4.6994	1.8639	9.4635
PT	0.2193	0.6652	3.0329	1.7749	8.0921
ES	0.2465	0.8998	3.6509	2.1593	8.7611
TR	0.2889	1.1174	3.8685	1.7376	6.0156
ZA	0.0995	0.4443	4.4632	1.5379	15.4509
BR	0.3488	9.6938	27.7920	1.1364	3.2580
MX	0.2014	0.9556	4.7450	1.7608	8.7436
PE	0.1990	6.5228	32.7735	1.1769	5.9133
KO	0.5303	1.2478	2.3530	1.6497	3.1109
NIC	0.2780	3.3303	12.6659	1.4999	7.0820
Eurozone	0.1977	0.6439	3.2706	1.7586	9.0323
OtherInd	0.1884	0.5124	3.0774	1.3408	8.4500

Notes: GDPvol, RXvol and RMvol represent the timeseries volatilities of real GDP, real exports and real imports, respectively, for each exporting country. Real export and real import volatilities are averaged over the other trading partners. The ratio RX/GDP (RM/GDP) is the ratio of averaged real export (import) volatility to GDP volatility. NIC, Eurozone and OtherInd rows are averages of those countries' values. NICs include TR–KO, Eurozone includes AT, FR, DE, IT, NL, FI, PT, ES and OtherInd includes all other countries.

Table 2: Conditional variance and covariance estimates for US exports

Impt.	$\bar{\sigma}^2_{x_t}$	$\bar{\sigma}^2_{s_t}$	$\overline{covar}$	IQR $\sigma^2_{x_t}$	IQR $\sigma^2_{s_t}$	IQR *covar*
UK	0.0136	0.0006	-0.0001	0.0029	0.0002	0.0004
AT	0.0891	0.0007	-0.0004	0.0497	0.0002	0.0014
FR	0.0149	0.0007	0.0004	0.0017	0.0001	0.0004
DE	0.0092	0.0007	0.0003	0.0014	0.0001	0.0004
IT	0.0182	0.0006	0.0000	0.0016	0.0001	0.0002
NL	0.0160	0.0009	0.0001	0.0072	0.0004	0.0008
NO	0.0641	0.0005	0.0002	0.0338	0.0001	0.0007
CH	0.0709	0.0005	0.0004	0.0501	0.0001	0.0009
CA	0.0094	0.0002	0.0001	0.0000	0.0000	0.0000
JP	0.0076	0.0009	0.0002	0.0030	0.0002	0.0004
PT	0.1003	0.0006	0.0003	0.0314	0.0000	0.0006
ES	0.0318	0.0006	-0.0004	0.0051	0.0002	0.0006
TR	0.0440	0.0057	-0.0030	0.0070	0.0032	0.0014
ZA	0.0377	0.0025	-0.0011	0.0145	0.0011	0.0008
BR	0.0244	0.0034	0.0000	0.0175	0.0016	0.0009
MX	0.0095	0.0031	-0.0003	0.0050	0.0009	0.0002
PE	0.0390	0.1749	-0.0067	0.0045	0.0320	0.0010
KO	0.0155	0.0007	0.0001	0.0041	0.0002	0.0004

Notes: *Impt.* denotes the importing country. $\bar{\sigma}^2_{x_t}$ is the mean conditional variance of the log real export series. $\bar{\sigma}^2_{s_t}$ is the mean conditional variance of the log real exchange rate series. $\overline{covar}$ is the mean covariance. The latter three columns contain the interquartile range (IQR) of the same three series.

Table 3: Conditional variance and covariance estimates for FR exports

Impt.	$\bar{\sigma}^2_{x_t}$	$\bar{\sigma}^2_{s_t}$	$\overline{covar}$	IQR $\sigma^2_{x_t}$	IQR $\sigma^2_{s_t}$	IQR *covar*
US	0.0161	0.0008	0.0004	0.0082	0.0004	0.0004
UK	0.0160	0.0004	-0.0000	0.0000	0.0003	0.0000
DK	0.0172	0.0001	0.0001	0.0016	0.0000	0.0001
SE	0.0280	0.0003	0.0000	0.0007	0.0002	0.0002
CH	0.0222	0.0002	-0.0000	0.0117	0.0001	0.0003
CA	0.0756	0.0006	-0.0002	0.0083	0.0001	0.0003
JP	0.0269	0.0007	-0.0000	0.0111	0.0004	0.0004
ZA	0.0570	0.0009	0.0002	0.0286	0.0006	0.0007
BR	0.0670	0.0018	-0.0004	0.0340	0.0009	0.0014
MX	0.1073	0.0029	0.0007	0.0784	0.0009	0.0025
PE	0.2657	0.2919	0.1286	0.1324	0.0372	0.0070

Notes: see notes to Table 2.

Table 4: Conditional variance and covariance estimates for BR exports

Impt.	$\bar{\sigma}^2_{x_t}$	$\bar{\sigma}^2_{s_t}$	$\overline{covar}$	IQR $\sigma^2_{x_t}$	IQR $\sigma^2_{s_t}$	IQR *covar*
US	0.0820	0.0065	0.0065	0.0058	0.0039	0.0043
UK	0.0907	0.0040	0.0027	0.0181	0.0023	0.0032
AT	0.3024	0.0041	0.0003	0.1444	0.0020	0.0060
FR	0.1253	0.0043	0.0009	0.0048	0.0026	0.0018
DE	0.0942	0.0040	0.0030	0.0109	0.0020	0.0026
IT	0.1086	0.0043	0.0041	0.0156	0.0033	0.0051
NL	0.1406	0.0043	0.0056	0.0192	0.0020	0.0044
NO	0.3052	0.0034	0.0030	0.0031	0.0020	0.0018
CA	0.1183	0.0050	0.0016	0.0075	0.0037	0.0021
JP	0.0984	0.0038	0.0014	0.0095	0.0022	0.0024
FI	0.2352	0.0040	0.0036	0.0235	0.0023	0.0033
ZA	0.2279	0.0056	0.0021	0.2758	0.0040	0.0049
MX	0.1145	0.0072	0.0032	0.0480	0.0049	0.0032

Notes: see notes to Table 2.

Table 5: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig*$\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
US	18	410.951	2	170.710	2	0.198	0.303
UK	16	-75.076	2	-30.014	2	3.176	-1.378
AT	8	-	0	-	0	-	-
DK	12	-	0	-	0	-	-
FR	11	-722.824	1	-290.378	1	0.065	-1.889
DE	14	33.455	4	15.997	4	1.901	0.159
IT	14	83.615	2	31.068	2	0.957	0.472
NL	11	-	0	-	0	-	-
NO	12	-	0	-	0	-	-
SE	11	514.146	4	231.114	4	0.128	2.108
CH	13	-151.849	2	-55.807	2	53.106	-4.672
CA	15	1662.196	1	625.386	1	0.018	1.121
JP	20	-	0	-	0	-	-
FI	10	-188.041	2	-62.366	2	0.308	-2.387
PT	5	-142.693	1	-54.603	1	0.355	-1.940
ES	9	561.892	1	197.315	1	0.317	6.264
TR	8	-	0	-	0	-	-
ZA	5	-	0	-	0	-	-
BR	13	14.811	8	7.325	8	6.946	3.723
MX	9	16.846	1	6.298	1	3.573	2.250
PE	7	-0.168	7	-0.072	7	2814.587	-6.104
KO	13	-	0	-	0	-	-

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 6: Summary results for β_1 over country-pair categories

	# Models	*# Signif.*	*Pct. Sig.*	*med sig* $\hat{\beta}_1$
NonEuro–NonEuro	91	13	14.3	0.8811
Euro–NonEuro	82	11	13.4	-0.8647
NonEuro–Euro	81	14	17.3	1.1211
Ind–Ind	149	17	11.4	0.3026
Ind–NIC	50	5	10.0	-0.9486
NIC–Ind	45	15	33.3	2.2504
NIC–NIC	10	1	10.0	-5.5478

Notes: The categories of country-pairs are not mutually exclusive. $\#\ Models$ gives the total number of models estimated for that category, while $\#\ Signif.$ indicates how many estimates of β_1 were significant at the 5% level. *Pct. Sig.* is the fraction of models with significant coefficients. *med sig* $\hat{\beta}_1$ is the median value of the estimated coefficient over those models in which it was significantly different from zero.

Table 7: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
US	18	3.179	12	5.252	10	1.284	5.918
UK	16	3.489	10	4.246	10	0.893	6.152
AT	8	3.645	3	3.770	3	1.318	18.774
DK	12	25.102	10	22.857	10	8.353	102.096
FR	11	11.142	2	0.459	3	925.594	51.960
DE	14	2.525	9	2.045	9	1.649	8.785
IT	14	0.292	7	0.469	8	2.981	2.688
NL	11	3.422	5	6.227	5	4.510	24.004
NO	12	36.238	5	56.222	4	1.022	17.276
SE	11	2.225	6	1.241	5	0.349	0.683
CH	13	12.749	5	14.700	5	0.875	16.229
CA	15	3.852	7	4.738	6	0.669	1.074
JP	20	1.763	10	2.773	10	1.139	4.301
FI	10	5.986	2	1.397	3	0.345	1.845
PT	5	13.254	2	17.635	2	0.269	5.713
ES	9	32.309	2	47.169	2	0.280	10.893
TR	8	-0.015	4	-0.022	4	7.569	-0.032
ZA	5	13.903	2	10.206	2	1.981	19.382
BR	13	2.477	6	1.620	6	6.179	7.542
MX	9	-2.279	5	-7.983	3	5.236	-5.178
PE	7	2.999	3	1.700	4	4513.083	273.740
KO	13	137.641	4	44.957	5	11.459	147.542

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 8: Summary results for ϕ_1 over country-pair categories

	# Models	*# Signif.*	*Pct. Sig.*	*med sig* $\hat{\phi}_1$
NonEuro–NonEuro	91	44	48.4	3.6603
Euro–NonEuro	82	32	39.0	9.8389
NonEuro–Euro	81	45	55.6	11.2146
Ind–Ind	149	78	52.3	5.8999
Ind–NIC	50	19	38.0	8.7847
NIC–Ind	45	21	46.7	13.6422
NIC–NIC	10	3	30.0	36.0725

Notes: The categories of country-pairs are not mutually exclusive. $\#\ Models$ gives the total number of models estimated for that category, while $\#\ Signif.$ indicates how many estimates of ϕ_1 were significant at the 5% level. *Pct. Sig.* is the fraction of models with significant coefficients. *med sig* $\hat{\phi}_1$ is the median value of the estimated coefficient over those models in which it was significantly different from zero.

Appendix Tables

Table 9: *NonEuro–NonEuro* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
US	11	-84.132	1	-34.030	1	0.348	-1.185
UK	8	-7.317	1	-2.731	1	6.166	-1.684
DK	6	-	0	-	0	-	-
NO	5	-	0	-	0	-	-
SE	5	1143.698	2	441.077	2	0.102	2.400
CH	6	0.208	1	0.084	1	105.297	0.881
CA	8	-	0	-	0	-	-
JP	12	-	0	-	0	-	-
TR	4	-	0	-	0	-	-
ZA	3	-	0	-	0	-	-
BR	7	18.001	4	8.779	4	7.591	3.577
MX	5	16.846	1	6.298	1	3.573	2.250
PE	3	-0.042	3	-0.020	3	5983.717	-5.548
KO	8	-	0	-	0	-	-

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 10: *Euro–NonEuro* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
AT	8	-	0	-	0	-	-
FR	11	-722.824	1	-290.378	1	0.065	-1.889
DE	14	33.455	4	15.997	4	1.901	0.159
IT	14	83.615	2	31.068	2	0.957	0.472
NL	11	-	0	-	0	-	-
FI	10	-188.041	2	-62.366	2	0.308	-2.387
PT	5	-142.693	1	-54.603	1	0.355	-1.940
ES	9	561.892	1	197.315	1	0.317	6.264

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 11: *NonEuro–Euro* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
US	7	906.033	1	375.449	1	0.048	1.790
UK	8	-142.834	1	-57.297	1	0.187	-1.071
DK	6	-	0	-	0	-	-
NO	7	-	0	-	0	-	-
SE	6	514.146	2	231.114	2	0.154	2.108
CH	7	-303.907	1	-111.698	1	0.915	-10.225
CA	7	1662.196	1	625.386	1	0.018	1.121
JP	8	-	0	-	0	-	-
TR	4	-	0	-	0	-	-
ZA	2	-	0	-	0	-	-
BR	6	12.801	4	6.292	4	6.301	3.881
MX	4	-	0	-	0	-	-
PE	4	-0.377	4	-0.156	4	437.740	-6.852
KO	5	-	0	-	0	-	-

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 12: *Ind–Ind* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
US	12	410.951	2	170.710	2	0.198	0.303
UK	12	-142.834	1	-57.297	1	0.187	-1.071
AT	6	-	0	-	0	-	-
DK	10	-	0	-	0	-	-
FR	7	-722.824	1	-290.378	1	0.065	-1.889
DE	8	193.443	2	77.466	2	0.227	1.219
IT	8	177.230	1	66.081	1	0.243	1.604
NL	8	-	0	-	0	-	-
NO	11	-	0	-	0	-	-
SE	10	514.146	4	231.114	4	0.128	2.108
CH	11	-303.907	1	-111.698	1	0.915	-10.225
CA	12	1662.196	1	625.386	1	0.018	1.121
JP	15	-	0	-	0	-	-
FI	8	-188.041	2	-62.366	2	0.308	-2.387
PT	5	-142.693	1	-54.603	1	0.355	-1.940
ES	6	561.892	1	197.315	1	0.317	6.264

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 13: *Ind–NIC* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
US	6	-	0	-	0	-	-
UK	4	-7.317	1	-2.731	1	6.166	-1.684
AT	2	-	0	-	0	-	-
DK	2	-	0	-	0	-	-
FR	4	-	0	-	0	-	-
DE	6	-11.035	2	-4.190	2	3.574	-1.238
IT	6	-10.000	1	-3.946	1	1.671	-0.659
NL	3	-	0	-	0	-	-
NO	1	-	0	-	0	-	-
SE	1	-	0	-	0	-	-
CH	2	0.208	1	0.084	1	105.297	0.881
CA	3	-	0	-	0	-	-
JP	5	-	0	-	0	-	-
FI	2	-	0	-	0	-	-
ES	3	-	0	-	0	-	-

Notes: #*Impt.* is the number of bilateral relationships analyzed. The #*Sig.* values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 14: *NIC–Ind* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
TR	8	-	0	-	0	-	-
ZA	3	-	0	-	0	-	-
BR	11	14.811	8	7.325	8	6.946	3.723
MX	8	16.846	1	6.298	1	3.573	2.250
PE	6	-0.258	6	-0.101	6	3154.911	-6.852
KO	9	-	0	-	0	-	-

Notes: #*Impt.* is the number of bilateral relationships analyzed. The #*Sig.* values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 15: *NIC–NIC* Exports: Coefficient estimates and impact of β_1

Expt.	*# Impt.*	*med sig* $\hat{\beta}_1$	*# Sig.*	*med sig* $\hat{\beta}_1^{SS}$	*# Sig.*	$\bar{\tau}_s$	*% Impact*
ZA	2	-	0	-	0	-	-
BR	2	-	0	-	0	-	-
MX	1	-	0	-	0	-	-
PE	1	-0.168	1	-0.072	1	772.645	-5.548
KO	4	-	0	-	0	-	-

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\beta}_1$ and $\hat{\beta}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 16: *NonEuro–NonEuro* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
US	11	0.026	5	1.721	4	2.873	5.863
UK	8	0.504	4	0.967	4	1.858	2.591
DK	6	22.627	4	21.929	4	8.304	98.122
NO	5	4.464	1	8.005	1	3.407	6.344
SE	5	0.197	3	-8.313	2	0.606	-6.355
CH	6	4.874	2	2.871	2	0.775	2.720
CA	8	-0.719	4	-2.461	3	1.253	-4.621
JP	12	1.650	6	1.867	6	1.684	2.747
TR	4	-0.753	1	-1.277	1	15.137	-23.290
ZA	3	3.264	1	1.622	1	1.879	5.024
BR	7	1.730	4	1.092	4	5.965	4.574
MX	5	-4.001	3	-16.766	2	3.984	-23.209
PE	3	1.686	2	0.758	2	8589.253	440.540
KO	8	137.641	4	44.957	5	11.459	147.542

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 17: *Euro–NonEuro* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
AT	8	3.645	3	3.770	3	1.318	18.774
FR	11	11.142	2	0.459	3	925.594	51.960
DE	14	2.525	9	2.045	9	1.649	8.785
IT	14	0.292	7	0.469	8	2.981	2.688
NL	11	3.422	5	6.227	5	4.510	24.004
FI	10	5.986	2	1.397	3	0.345	1.845
PT	5	13.254	2	17.635	2	0.269	5.713
ES	9	32.309	2	47.169	2	0.280	10.893

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 18: *NonEuro–Euro* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
US	7	9.895	7	10.866	6	0.225	5.918
UK	8	6.606	6	5.133	6	0.249	10.515
DK	6	26.057	6	23.201	6	8.386	103.036
NO	7	64.069	4	79.289	3	0.227	28.209
SE	6	2.295	3	2.194	3	0.178	1.110
CH	7	46.304	3	18.739	3	0.943	17.071
CA	7	42.025	3	36.129	3	0.086	2.778
JP	8	7.012	4	9.569	4	0.321	9.424
TR	4	0.059	3	0.115	3	5.046	1.176
ZA	2	24.542	1	18.790	1	2.083	33.739
BR	6	3.560	2	2.113	2	6.606	11.215
MX	4	9.129	2	34.861	1	7.739	46.034
PE	4	3.807	1	10.437	2	436.912	165.464
KO	5		0		0		

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 19: *Ind–Ind* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
US	12	3.726	11	6.207	9	0.225	6.087
UK	12	3.793	9	4.497	9	0.307	5.040
AT	6	38.792	2	19.663	2	0.240	9.387
DK	10	25.102	10	22.857	10	8.353	102.096
FR	7	0.245	1	0.459	1	0.379	0.647
DE	8	9.203	6	6.313	6	0.181	8.173
IT	8	1.079	3	2.532	4	0.353	1.620
NL	8	18.858	4	15.808	4	3.954	13.279
NO	11	64.069	4	79.289	3	0.227	28.209
SE	10	2.295	5	1.717	4	0.269	0.896
CH	11	12.749	5	14.700	5	0.875	16.229
CA	12	23.212	4	21.704	4	0.107	2.270
JP	15	4.453	8	4.521	8	0.359	4.301
FI	8	5.986	2	1.397	3	0.345	1.845
PT	5	13.254	2	17.635	2	0.269	5.713
ES	6	32.309	2	47.169	2	0.280	10.893

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 20: *Ind–NIC* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
US	6	0.026	1	0.044	1	10.820	5.076
UK	4	0.983	1	1.891	1	6.166	9.071
AT	2	3.645	1	3.770	1	3.472	20.308
DK	2		0		0		
FR	4	22.040	1	6.778	2	1388.201	114.939
DE	6	0.361	3	0.353	3	4.586	8.785
IT	6	0.231	4	0.369	4	5.609	4.455
NL	3	3.422	1	6.227	1	6.735	24.004
NO	1	4.464	1	8.005	1	3.407	6.344
SE	1	-11.697	1	-16.972	1	0.672	-13.335
CH	2		0		0		
CA	3	-1.437	3	-7.107	2	1.793	-5.819
JP	5	1.164	2	1.443	2	4.257	36.640
FI	2		0		0		
ES	3		0		0		

Notes: *#Impt.* is the number of bilateral relationships analyzed. The *#Sig.* values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 21: *NIC–Ind* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
TR	8	-0.015	4	-0.022	4	7.569	-0.032
ZA	3	13.903	2	10.206	2	1.981	19.382
BR	11	1.786	5	1.295	5	6.561	7.903
MX	8	-2.279	5	-7.983	3	5.236	-5.178
PE	6	2.999	3	1.700	4	4513.083	273.740
KO	9	437.644	2	273.387	3	11.432	638.180

Notes: *#Impt.* is the number of bilateral relationships analyzed. The *#Sig.* values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.

Table 22: *NIC–NIC* Exports: Coefficient estimates and impact of ϕ_1

Expt.	*# Impt.*	*med sig* $\hat{\phi}_1$	*# Sig.*	*med sig* $\hat{\phi}_1^{SS}$	*# Sig.*	$\hat{\tau}_s$	*% Impact*
ZA	2		0		0		
BR	2	3.757	1	3.832	1	4.270	7.181
MX	1		0		0		
PE	1		0		0		
KO	4	10.616	2	19.629	2	11.498	64.964

Notes: $\#Impt.$ is the number of bilateral relationships analyzed. The $\#Sig.$ values refer to the number of significant $\hat{\phi}_1$ and $\hat{\phi}_1^{SS}$ coefficients estimated, respectively. $\bar{\tau}_s$ is reported as 1000 $\times$ the average standard deviation of σ_{s_t}.