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George Hondroyiannis
P.A.V.B. Swamy
George S. Tavlas
Michael Ulan

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SOME FURTHER EVIDENCE ON EXCHANGE-RATE VOLATILITY AND EXPORTS

George Hondroyiannis
Bank of Greece and Harokopio University

P.A.V.B. Swamy
U.S. Bureau of Labor Statistics

George Tavlas
Bank of Greece

Michael Ulan*
U.S. Department of State

ABSTRACT

The relationship between exchange-rate volatility and aggregate export volumes for 12 industrial economies is examined using a model that includes real export earnings of oil-producing economies as a determinant of industrial-country export volumes. A supposition underlying the model is that, given their levels of economic development, oil-exporters' income elasticities of demand for industrial-country exports might differ from those of industrial countries. Five estimation techniques, including a generalized method of moments (GMM) and random coefficient (RC) estimation, are employed on panel data covering the estimation period 1977:1-2003:4 using three measures of volatility. In contrast to recent studies employing panel data, we do not find a single instance in which volatility has a negative and significant impact on trade.

Keywords: Exchange-rate volatility; Trade; Random-coefficient estimation; Generalized method of moments; Panel
JEL classification: C23; F3; F31

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*Retired

Correspondence:
George Tavlas,
Economic Research Department,
Bank of Greece, 21, E. Venizelos Ave.,
102 50 Athens, Greece
Tel. +30210-320 2370, Fax +30210-320 2432
email: gtavlas@bankofgreece.gr

1. Introduction

A large empirical literature has investigated the relationship between exchange-rate volatility and trade flows. While the earlier literature (*circa* 1980 to the mid-1990s), employing mainly time-series data and ordinary least squares (OLS) estimation, did not, by-and-large, find a negative and significant effect of exchange-rate volatility on trade volumes, recent studies, often using panel data and more-elaborate estimation techniques (fixed effects, random effects) and specifications (*e.g.*, gravity models), have uncovered some - - though by no means overwhelming - - evidence of a negative, significant relationship.[1] For example, in a study that surveyed recent work and provided new evidence, Clark, Tamirisa, and Wei (2004, p. 3) found that "some recent studies, as well as some of the evidence presented here, appear to suggest that the data support a negative relationship".[2]

This study is a further entry into an already crowded field. We aim to extend the literature in several directions, using an analytic framework proposed by Bailey, Tavlas, and Ulan (1986). Those authors investigated the relationship between exports of the seven largest industrial economies and the short-term volatility of the exchange rates of the currencies of those economies based on a specification that included real export earnings of oil-producing economies as a determinant of export volumes. The suppositions underlying this specification were as follows: (1) since the 1970s, oil-exporting economies have provided important markets for exports of industrial economies, and (2) given their levels of economic development, the income elasticities of demand of the oil producers for industrial-country exports might well differ from the income elasticities of demand for those goods in other industrial countries. Bailey, Tavlas, and Ulan found that oil-exporters' income elasticity of demand for industrial-country exports was statistically significant. In light of the large fluctuations in oil prices (and of export earnings of oil-producing economies) in recent years, the inclusion of such a variable appears to be especially appropriate in present circumstances. We apply this model to the exports of 12 industrial countries.

[1] Surveys of earlier literature include IMF (1984) and Edison and Melvin (1990). The more recent literature is surveyed by McKenzie (1999) and Clark, Tamirisa, and Wei (2004).
[2] The authors of that study were careful to point out, however, that such a negative relationship was not robust to all specifications they estimated. In his survey, McKenzie (1999) similarly found that most recent papers appeared to be having greater success in deriving such (*i.e.*, negative) relationship.

As indicated above, much of the recent literature has used panel-data estimation techniques. As our point of departure, we follow the literature in that respect. Estimation with panel data can help control for individual heterogeneity and nonstationarity, and can improve efficiency of estimators by using data with more variability and less collinearity (by combining variation across cross-sectional units with variation over time). Specifically, we follow much of the recent literature in using common-fixed-coefficient estimation, fixed-effects estimation and random-effects estimation. We extend the work of previous authors by using two additional estimation techniques, namely GMM estimation applied to dynamic panel-data specifications, and random-coefficient (RC) estimation. The GMM approach, proposed by Hansen (1982), purportedly does not require distributional assumptions, such as normality, can allow for heteroskedacity of unknown form, and can correct for the effects of misspecification errors including omission of variables (see Verbeek, 2004, pp. 148-53).[3] The RC approach deals with four major specification problems (discussed in Chang, Swamy, Hallahan and Tavlas, 2000) that almost always arise in econometric estimation.[4] The approach takes as its point of departure the premise that, although one can never be sure that a “true” model (in this case, a model of the determination of exports) exists, RC estimation, by correcting for factors that cause spurious relationships (*e.g.*, the effects of omitted variables, unknown functional forms, and measurement errors), can find the most-reasonable approximations to the “true” values of the identifiable coefficients of the “true”, but unknown, model.

The remainder of this paper consists of six sections. Section 2 briefly summarizes recent studies that use panel data estimation. Section 3 is an overview of the theoretical relationship between exchange-rate volatility and trade. Section 4 discusses the model and data. Section 5 provides a brief description of the estimation techniques. Section 6 presents the empirical results. Section 7 concludes.

[3] Having said that, the assumptions underlying the GMM approach are questioned below.

[4] For discussions of RC estimation, see Swamy and Tinsley (1980), Chang, Hallahan, and Swamy (1992), Swamy, Chang, Mehta, and Tavlas (2003), Swamy, Tavlas, and Chang (2005) and Swamy and Tavlas (2001, 2005, 2006).

2. Literature review

Previous studies employing panel data have tended to find evidence of a negative and statistically-significant relationship between exchange-rate volatility and trade. Wei (1999) estimated a panel of 63 countries over the years 1975, 1980, 1985 and 1990; a total of over 1000 country pairs was examined. Using switching regressions, the author found that, for country pairs with large potential trade, exchange-rate volatility had a negative and significant effect on bilateral trade among the countries considered. Dell'Arricia (1999) examined the effect of exchange-rate volatility on the bilateral trade of European Union members plus Switzerland over the period 1975-1994 using several definitions of volatility. In the basic OLS regression, exchange-rate volatility had a small but significant -negative impact on trade; reducing volatility to zero in 1994 would have increased trade by an amount ranging from ten to 13 per cent, depending on the measure of volatility used. Using both fixed and random effects, the impact of volatility was still negative and significant, but smaller in magnitude. The author found that elimination of exchange rate volatility would have increased trade by about 3½ per cent in 1994. Rose (2000) also obtained similar results employing a gravity model. His data set consisted of 186 countries for the five years 1970, 1975, 1980, 1985, and 1990. In his benchmark results (without random effects), he found that reducing volatility by one standard deviation would increase bilateral trade by about 13 per cent. Using random effects, he also found a small but significant negative effect; reducing volatility by one-standard deviation would increase bilateral trade by about four per cent. In general, Rose's results are consistent with those of Dell'Arricia.

Tenreyro (2004), however, cast some doubt on the robustness of Rose's results. Using annual data from 1970-1997 on a sample of 104 (developed and developing) countries, and employing a gravity model that took endogeneity into account, she found that volatility had an insignificant effect on trade. Clark, Tamirisa and Wei (2004) applied the gravity model using a battery of estimation techniques - - including fixed and random effects - - to a panel of 178 International Monetary Fund (IMF) members using every fifth year from 1975-2000. Using both country- and time-fixed effects, the authors found a negative and significant impact of exchange-rate volatility on trade; a one-standard deviation fall in exchange rate volatility, would raise trade by seven per cent. Allowing for time-varying random effects, however, a negative relationship was not evident. The authors concluded that, while there is evidence that increased exchange-rate volatility reduces the volume of trade, this finding depends on the particular estimation technique employed.

3. Analytical considerations

The argument that exchange-rate volatility reduces trade typically runs as follows.[5] In a two-country context, consider a firm located in country A that sells its product in country B (as well as in country A). For simplicity, suppose that the firm sells in a forward market in each country so that the firm knows the future price of its product at the time it incurs its costs of production. However, if there is no futures or forward market for foreign exchange, the firm has an exchange risk for the future conversion of its sales revenues from country B into the currency of country A. If the firm is risk-averse, it would be willing to incur an added cost to avoid this risk, so that the risk, if not hedged, is an implicit cost. In the presence of such a cost, this reasoning suggests that the firm's supply price at each quantity of export sales is higher than in the absence of the risk. For such firms in the aggregate, the quantity of exports supplied at a given price is smaller with this risk than without it. The same reasoning applies to firms in country B. If the risk is present for firms in both countries, the supply curve for exports from each country to the other is shifted to the left, compared with those that would exist in the absence of exchange rate risk. Trade is reduced in a way similar to that resulting from an increase in transportation costs.

Where there is a forward market for foreign exchange, a discount of the forward exchange rate in one direction, below the expected future rate, is a premium in the other direction. Thus, if expectations are similar in the two countries, such a discount cannot be a deterrent to trade in both directions. However, the brokerage cost (spread) for forward transactions is generally greater than that for spot transactions in foreign exchange, and the spread is an increasing function of the variability of the exchange rate. Hence, the risk can be hedged only at a cost; the existence of forward or futures markets for foreign exchange does not change the thrust of the above argument though it reduces its quantitative significance.

[5] For recent discussions see McKenzie (1999) and Clark, Tamirisa and Wei (2004). The discussion in the text, draws, in part, and expands on those studies as well as the studies by Bailey, Tavlas and Ulan (1987), Dellas and Zilberfarb (1993), and De Grauwe (2005).

The arguments, however, are not all on one side. Consider the following factors, which suggest that exchange-rate volatility can increase trade.

(i) Exporters may gain knowledge through trade that might help them anticipate future exchange-rate movements better than can the average participant in the foreign-exchange market. If so, the profitability of this knowledge could be used to offset the risk of exchange-rate volatility. If exporters wish to hedge longer-term investment or other transactions, rather than use the forward-exchange market, they can borrow and lend in local currency to offset their other commitments. For example, a plant in a foreign country can be financed mainly with local capital, so that investors limit their exchange risk in the basic investment.

(ii) A counter-argument of especially great weight is that one must specify the alternative to exchange-rate volatility. If the volatility is attributable to fundamental factors' influencing the exchange rate, intervention by the authorities to reduce it would be unsustainable and eventually disruptive. To achieve a reduction of apparent, observed volatility, authorities would have to intervene with exchange controls or other restrictions on trade and payments. That intervention could be more harmful to trade, and reduce it more, than would unrestrained movement of the exchange rate.

(iii) Variability of an exchange rate does not measure the effect added amounts of that foreign currency have on the overall riskiness on the firm's asset portfolio. The latter risk effect depends on the covariance of an exchange rate with the prices of the firm's other assets as well as the own variance of the exchange rate. In particular, the firm may hold a portfolio of several foreign currencies, thereby diversifying the risk. If variations in one currency's exchange rate against the home currency are negatively correlated with the variations in others, its variability reduces portfolio risk rather than increasing it when that currency is added to the portfolio. In general, variance by itself does not measure the exchange risk.

(iv) If firms can adjust factor inputs in response to movements in the exchange rate, increased variability may create opportunities to raise profits. That is, movements in exchange rates represent not only risk, but also potential reward. If a firm adjusts inputs to both high and low prices in order to take advantage of profit opportunities when prices are relatively high, its expected (or average) profits will be higher the higher is exchange-rate volatility because the firm can sell more when the price is high and less when the price is low. If risk aversion is relatively low, the positive effect of greater price volatility on expected profits may outweigh the negative impact of higher profits, and the firm will produce and export more (Clark, Tamirisa, and Wei, 2004, p. 4; De Grauwe, 2005, pp. 69-75). As pointed out by De Grauwe (2005, p. 73), exporting goods can be viewed as an option, whereby the value of the option, rises when the volatility of the underlying asset increases; when the exchange-rate becomes more favorable, the firm exercises its option to export.

In the light of the foregoing arguments, the issue of the relationship between exchange-rate volatility and trade appears to be an empirical question. In what follows, we describe the approach taken in this paper.

4. The model and the data

Following Bailey, Tavlas, and Ulan (1986), the model estimated (with one exception) is of the following form:

$$\log \mathrm{X}_{it} = \log a_1 + a_2 \log \mathrm{Y}_{it} + a_3 \log RP_{it} + a_4 \log OP_t + a_5 V_{it} + e_{it} \tag{1}$$

where X_{it} is the volume of exports of country i, Y_{it} is real GDP of industrial country trading-partner nations, RP_{it} is a measure of relative prices of exports of country i to its trading partners, OP_t represents real export earnings of oil-exporting countries, V_{it} is real exchange-rate variability, and e_{it} is a random-error term, and t indexes time. In equation (1), the coefficients are assumed to be constants. This strong assumption gets relaxed in RC estimation. Furthermore, the assumptions about the relationship between e_{it} and the explanatory variables in (1) should be based on the correct interpretation of e_{it}. Such assumptions are used in RC estimation (see Swamy and Tavlas, 2001). As discussed in the next section, RC estimation does not add an arbitrary error term to a mathematical equation to obtain an econometric model as is done in (1).

Using quarterly time-series data over the interval 1973:1-1984:3, Bailey, Tavlas, and Ulan (1986) estimated time-series regressions based on the above model for each of the G-7 economies - - Canada, France, Germany, Italy, Japan, the U.K., and the U.S. Using OLS, correcting for autocorrelation where necessary, and employing two measures of nominal exchange-rate volatility, they estimated a total of 22 regressions.[6] The authors did not find a single instance of a negative and significant coefficient on the exchange-rate volatility term in any of their regression.

This study applies the above model to 12 industrial economies - - the G-7 economies plus Ireland, the Netherlands, Norway, Spain, and Switzerland. The data frequency is quarterly and the sample period is 1977:1-2003:4. All data come from the *International Financial Statistics* (IFS).[7] In what follows, we describe these data.

The dependent variables in the estimated equations are the real exports of the countries considered. There are problems involved in devising proxies for the independent variables. Theory tells us that income in trading-partner nations should affect a country's exports. To construct an income variable for trading partners, we proceeded as follows. Real GDP data for Australia, Canada, France, Germany, Italy, Japan, the Netherlands, Spain, Switzerland, the United Kingdom, and the United States are available in the IFS for the period covered by this study. The industrial-country trading-partner income variable for Ireland and Norway was constructed by converting the real GDP data for the 12 countries listed above to U.S.-dollar terms using year-average 1998 exchange rates and summing the data for all 12 countries. For each of the other countries that is the subject of this investigation, the industrial-country trading-partner income variable is constructed by subtracting the dollar-denominated real GDP data for the country in question from the 12-country sum, *e.g.*, the industrial-country trading-partner income variable for Switzerland is the dollar-denominated 12-country aggregate GDP minus the dollar-denominated real GDP for Switzerland. These series were employed as our industrial-country-income variable, Y_{it}. In order to aggregate national GDP series, it was necessary to convert them to a common currency; we chose the US dollar. However, we wanted to ensure that our income variables were affected by only changes in real incomes in partner nations; we did not want the variables to be affected by the changing foreign-exchange value of the dollar. Thus, we converted all GDP data to dollars at a set of

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[6] Specifically, the authors used the absolute value of the quarter-to-quarter percentage change in the nominal effective exchange rate as a measure of volatility. The variable was used in both its current period form and as an eight-period, second-degree polynomial lag.

[7] The choice of both the particular countries and the sample period was constrained by data availability. Data series that were not seasonally-adjusted in the IFS were tested for seasonal adjustment using the Census X11 program (multiplicative option); those that displayed seasonality were seasonally adjusted for use here.

fixed exchange rates. We valued trading-partner income in dollars at average 1998 exchange rates.

While industrial-country trading partners purchase the bulk of the exports of the countries under study here, since 1973, oil-exporting countries have been major purchasers of the exports of these countries. Since they tend to be at a different stage of development from that of industrial-country trading partners, however, it is possible that oil-producers' income elasticities of demand for imports from industrial countries differ from the income elasticities of industrial countries. Accordingly, the oil-exporter "income" variable enters our export equations separately from the GDPs of industrial-country trading partners. The ability of oil exporters to purchase foreign goods varies with the real purchasing power of their exports rather than the countries' real outputs, and export earnings of these countries can vary with the price of oil even as their real GDPs move in the opposite direction. Thus, the oil-exporter-income variable is the sum of the dollar values of the oil-exporters' export earnings deflated by the dollar-denominated export unit value index of the industrial countries taken as a whole.[8]

Theoretically, the relative-price variables in the export equations should be the ratio of export prices in country *i* to the domestic prices of similar goods produced by its trading partners. Since that measure is not available, the relative-price variable in our export regressions is a real-exchange-rate index for each country. This variable is based on unit labour costs in manufacturing and represents the product of the index of the ratio of the relevant indicator of the country considered to a weighted geometric average of the corresponding indicators for twenty other industrial counties.

Exchange-rate volatility has been measured in the literature using either nominal or real (effective) exchange rates. As nominal and real exchange rates tend to move closely together, given the stickiness of domestic prices (especially in the short-run), the choice of measure is not likely to affect the econometric results. The decision to engage in international transactions, however, stretches over a relatively-long period of time, during which production cost and import prices in foreign-currency terms are likely to vary. This latter consideration suggests that exchange rates

[8] The following countries comprise the IMF's oil-exporters composite and are used in constructing the oil-exporter-income variable used in this study: Algeria, Indonesia, Iran, Iraq, Kuwait, Libya, Nigeria, Oman, Qatar, Saudi Arabia, the United Arab Emirates, and Venezuela.

measured in real terms are appropriate, and we have, therefore, used real rates.[9] For each country, three volatility measures are tested. Our first two measures, described below, were used by Bailey, Tavlas, and Ulan (1986) in terms of nominal values.

(1) One measure is the absolute values of the quarterly percentage change in the exporting nation's effective exchange rate:

$$A_{i,t} = \left|\left(\mathrm{E}_{i,t} - \mathrm{E}_{i,t-1}\right) / \mathrm{E}_{i,t-1}\right| \tag{2}$$

where $\mathrm{E}_{i,t}$ is the real effective exchange rate of the currency of exporting nation i. This measure of volatility is used to test for a stable and significant response of exports to a one-percentage-point change in the exchange rate.

(2) A second measure is the log of the eight-quarter moving standard deviation of the real effective exchange rate. Both this measure and the previous measure capture delayed responses of exports to exchange-rate volatility. This second measure is used to test for a stable and significant response of exports to a one-per cent change in the standard deviation.

$$S_{i,t} = \left[\frac{1}{8}\sum_{k=1}^{8}(\mathrm{E}_{i,t+k-1} - \mathrm{E}_{i,t+k-2})^2\right]^{1/2} \tag{3}$$

(3) In recent years, some authors have attempted to capture exchange-rate volatility by using the conditional second moment to proxy such volatility (see, *e.g.*, Chou, 2000, Clark, Tamirisa and Wei, 2004, Siregar and Rajan, 2004). The underlying idea is that part of the volatility can be predicted based on past values of the exchange rate. Therefore, firms engaged in trade would likely make an effort to develop such a forecast. We constructed a GARCH measure of volatility as follows:

$$\mathrm{E}_{i,t} = \alpha_0 + \alpha_1 \mathrm{E}_{i,t-1} + u_{i,t} \tag{4a}$$

$$h_{i,t} = \alpha + \beta u_{i,t-1}^2 + \gamma h_{i,t-1} \tag{4b}$$

[9] In his literature survey, McKenzie (1999, p. 85) concluded that the distinction between real and nominal rates has not impacted significantly on the results derived.

where the exchange rates are again expressed in logs and $u_{i,t}$ is a random error. The conditional variance equation in (4b) is a function of three terms: (i) the mean, α; (ii) news about volatility from the previous period, measured as the lag of the squared residual from the mean equation, $u_{i,t-1}^2$ (the ARCH term); and (iii) the last period's forecast error variance, $h_{i,t-1}$ (the GARCH term). We estimated a number of versions of GARCH models. For equation (4a), lags of up to three periods were used, depending upon whether the lags were significant. A GARCH (1,1) specification generated superior results.[10]

5. Estimation methods

This section briefly describes the five estimation techniques used.[11] We assume, realistically we believe, that RC estimation is likely to be less familiar to readers than the other approaches used. Therefore, we devote somewhat more space to describing the RC procedure.

(i) *Common fixed coefficients.* This approach applies OLS to the panel data, allowing the intercept and slopes of (1) to be the same for all the countries and time periods we considered. Under this assumption, (1) may not provide an adequate approximation to the "true" model (see Swamy and Tavlas, 2001).

(ii) *Fixed effects*. Suppose that certain unobserved country-specific variables, that are constant over time *t*, influence the dependent variable of equation (1) and are correlated with the explanatory variables in the equation. Under this assumption, a country-specific constant term is added to the right-hand side of equation (1) to allow the equation contain the country-specific variables.[12] In this connection, some authors have claimed (*e.g.*, Clark, Tamirisa, and Wei, 2004) that country-specific constant terms help control for remoteness or multilateral resistance effects. The concept of

[10] Siregar and Rajan (2004) obtained similar results in their study of Southeastern Asian economies. These authors, as well as McKenzie (1999), mention potential problems involved in ARCH-based measures of exchange rate volatility; since the exchange rate volatility generated prior to the end of the sample period incorporates knowledge about the future, as ARCH models are estimated over the entire sample period. To overcome this problem, one would need to re-estimate the ARCH model at the beginning of each quarter using information that is known to the trader at that point in time.

[11] Baltagi (2001) provides a comprehensive discussion of one-way and two-way fixed and random effects models and their use in panel-data analyses.

[12] Additionally, a time-specific constant term could be introduced.

multilateral resistance was proposed by Anderson and Van Wincoop (2003), who defined it as a function of unobservable equilibrium price indices that depend on bilateral trade barriers and income shares of the trading partners.

(iii) *Random effects*. If the unobserved country-specific variables represented by a country-specific constant term are uncorrelated with the explanatory variables of (1), then the random-effects approach specifies that the country-specific term is a country-specific random element, similar to e_{it}, except that, for each country, there is but a single draw that enters equation (1) identically in each period. With random effects, the error term has two components: the traditional error unique to each observation and an error term representing the extent to which the intercept of the ith country differs from the overall intercept. The composite error term is nonspherical, so that generalized least squares (GLS) estimation is used.[13]

(iv) *Generalized method of moments (GMM).* Equation (1) is extended to include $\log X_{i,t-1}$, $\log X_{i,t-2}$, and $\log X_{i,t-3}$ as additional explanatory variables. A GMM is used to estimate this extended equation with lagged independent variables acting as instruments; in the seven-country panel five lags of each of the independent variables were used while in the 12-country panel three lags of each of the independent variables were used. Since there are more instruments than right-hand side variables the estimated regression equations are over-identified. To assess the validity of the different specifications we compute the Sargan (1964) test for over-identifying restrictions, which amounts to a test of the exogeneity of the explanatory variables, and m_1 and m_2 tests for autocorrelation.

Each of the above four estimation techniques imposes assumptions that can be hard to fulfill. Common-fixed coefficients estimation assumes that the intercept and slopes are the same for all countries in each time period. We have already noted the serious implications stemming from the assumptions made by common-fixed-coefficients estimation. With regard to fixed-effects estimation, for consistent estimation of equation (1) using the OLS method, a necessary condition is that the country-specific variables and the included explanatory variables in equation (1) are mean independent of e_{it}. Under random effects, a necessary condition for the GLS

[13] Dell'Ariccia (1999) claims that use of both fixed and random effects can help deal with simultaneity problems. The Swamy and Arora (1972) estimators of the component variances are employed in the estimation of the random-effects equations.

estimator of the coefficient of equation (1) to be consistent is that the error components are mean independent of the explanatory variables in equation (1). GMM estimators can be inconsistent because there are obstacles in obtaining the instrumental variables needed to apply GMM, as shown by Swamy and Tavlas (2001) in their derivation of an adequate approximation to a "true" model. In the light of these factors, we turn to a procedure that can produce consistent estimators of the coefficients.

(v) *Random coefficients (RC) estimation.* In this estimation, not only the intercept but also the slopes of (1) are allowed to differ among the countries both at a point in time and through time. In this form, (1) is referred to as "the time-varying coefficients (TVC) model". Let a_{1it}^{*} = $\log a_{1it}$ + e_{it}, a_{2it}^{*}, a_{3it}^{*}, a_{4it}^{*}, and a_{5it}^{*} be the intercept and the coefficients on $\log Y_{it}$, $\log RP_{it}$, $\log OP_{t}$, V_{it}, respectively, in the TVC model. Then it follows from Swamy and Tavlas (2001) that when the "true" model exists, the TVC model is its exact representation with unique coefficients if (i) the intercept, a_{1it}^{*}, is interpreted as the sum of (a) the intercept of the "true" model, (b) the joint effect on $\log X_{it}$ of the portions of excluded variables (*i.e.*, the determinants of $\log X_{it}$ excluded from (1)) remaining after the effects of the "true" values of the explanatory variables in (1) have been removed, and (c) the measurement error in $\log X_{it}$, and (ii) a_{jit}^{*} with *j* > 1 is interpreted as the sum of (a) the *j*th coefficient of the "true" model, (b) a term capturing omitted-variables bias due to excluded variables, and (c) a measurement-error bias due to mismeasuring the *j*th explanatory variable in (1). These are the correct interpretations of the coefficients of the TVC model.

We estimate the TVC model under the following assumption: For all *i* = 1, 2, …, n (= 7 or 12) and *t* = 1, 2, …, T (= 32):

$$a_{i,t}^{*} = \bar{a}_{i}^{*} + \varepsilon_{i,t}^{*} \qquad \text{(5a)}$$

$$\varepsilon_{i,t}^{*} = \Phi_{ii}\varepsilon_{i,t-1}^{*} + v_{i,t}^{*} \qquad \text{(5b)}$$

where $a_{i,t}^{*} = (a_{1i,t}^{*}, a_{2i,t}^{*}, ..., a_{5i,t}^{*})'$, $\bar{a}_{i}^{*} = (\bar{a}_{1i}^{*}, \bar{a}_{2i}^{*}, ..., \bar{a}_{5i}^{*})'$, $\varepsilon_{i,t}^{*} = (\varepsilon_{1i,t}^{*}, \varepsilon_{2i,t}^{*}, ..., \varepsilon_{5i,t}^{*})'$, Φ_{ii} is a 5×5 matrix, $v_{i,t}^{*} = (v_{1i,t}^{*}, v_{2i,t}^{*}, ..., v_{5i,t}^{*})'$, the $\bar{a}_{i}^{*}$ are independently distributed with mean vector, $E\bar{a}_{i}^{*} = \bar{a}^{*} =$ $(\bar{a}_{1}^{*}, \bar{a}_{2}^{*}, ..., \bar{a}_{5}^{*})'$, and covariance matrix, Δ, the $v_{i,t}^{*}$ are independently distributed with mean zero and covariance matrix, Δ_{ii}, and $\varepsilon^{*}_{i,t}$ is independent of $\overline{a}_i^*$ and the explanatory variables of (1). We may have to include some observable variables with nonzero coefficients on the right-hand side of (5a) to make this independence assumption hold. We consider both zero and nonzero values of Φ_{ii}.

6. Empirical results

Two sets of regressions were estimated. First, panel estimation, using each of the five methods, was performed using data for the G-7 countries - - Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States. Second, the foregoing procedure was applied to the sample of twelve countries. In what follows, specifications corresponding to the five estimation methods - - common-fixed-coefficients, fixed-effects, random-effects, GMM, and RC - - are identified by the numbers 1 through 5, respectively. Specifications with the absolute percent change measure of volatility are identified with the subscript "a", those with the moving-standard-deviation measure with the subscript "b", and those using the GARCH measure with the subscript "c". For the RC regressions, the average of the coefficients over the entire time period and all cross sectional units are reported.

All the reported results for GMM estimation are the long-run coefficients with large sample t-ratios. Two types of diagnostic tests are performed for all the estimated equations. First, the m_1 and m_2 tests for autocorrelation are conducted for the seven and 12- country panels are reported (the values of m_1 for the seven-country panel are, 4a:0.15 4b:0.04 4c:0.09 and for the twelve-country panel are, 4a:0.55, 4b:0.60, 4c:0.67 the values of m2 for the seven-country panel are, 4a:1.66 4b:1.49 4c:1.64 and for the 12-country panel are, 4a:-1.13, 4b:-0.49, 4c:-0.83). These values do not reject the hypotheses that there is no serial correlation in the regression disturbances. Next the Sargan test (seven-country panel: 4a:27.93 4b:22.61 4c:24.15 and twelve-country panel: 4a:10.90, 4b:8.28, 4c:10.79), for all the estimated regressions, does not reject the over-identifying restrictions.

Table 1 presents regression results for the seven-country panel. The following results merit comment. First, the coefficients on the volatility variables are insignificant in each of the 15 regressions reported.[14] Second, the coefficient on the oil-exporter income variable is significant and positive in each equation, and markedly lower than the coefficient on the industrial-country income variable, indicating that oil producers and industrial countries have different income elasticities, as hypothesized. Third, comparing the coefficients of the three explanatory variables other than volatility (*i.e.*, industrial-country trading-partner income, oil-exporter income, the real exchange rate) *among* the five sets of regressions, there is a clear delineation between the results based on common fixed coefficients and the results of the other four methods. Each of the other four methods yields industrial-country income elasticities in the range of 1.6-1.7, more than twice those provided by method 1. Methods 2 through 5 give oil-exporter-income elasticities in the range of .07-.09, about one-third of the elasticities obtained by method 1. Also, method 1 yields low relative-price elasticities compared with the other four methods. Finally, unlike the other methods, method 1 produces coefficients on real exchange rate variables that are insignificant.

Table 2 reports the results of the panel based on all twelve countries. Again, there is no evidence that exchange-rate volatility reduces trade; the coefficient of the volatility term is insignificant in each of the fifteen equations reported. As was the case with the results based on the seven-country panel, the common-fixed-coefficients method gives lower elasticities for industrial-country income and higher elasticities for oil-exporter income than does each of the other four methods. Compared with the results reported in Table 1, the results of methods 2 through 5 generally show higher income (both industrial country and oil exporter) and higher relative-price elasticities. The exception is the random coefficient procedure, which yields price elasticities in Table 2 that are little different from those in Table 1.

Tables 1 and 2 also report standard error of the regression (SER) for each of the specifications. The following results are worth noting. First, fixed effects and random effects produce SERs in the range of about .10, less than half of those

[14] Nevertheless, the RC approach would not drop a variable under the condition that its coefficient is insignificant. A full treatment of RC estimation would assess whether the inclusion of a volatility measure in equation (1) could be reducing omitted-variable and measurement-error biases contained in the coefficients of the equation compared to what they would have been in the absence of a volatility measure. In other words, the fact that a volatility measure is insignificant and contributes little to the coefficient of determination does not in itself provide grounds for excluding the variable. The conditions needed to pursue this line of research are very difficult to implement in RC estimation using panel data and we leave this research for a future line of work.

produced by method 1. Second, GMM gives SERs that are about one-third of those given by fixed and random effects methods. Third, the RC method yields SERs far below those yielded by GMM.

An advantage of RC estimation is that it allows the time profiles of coefficients to be traced. In so doing, it allows us to pick up regime changes and other structural breaks quickly; by consistently estimating changing coefficients, we can attempt to estimate changes in the "true" coefficients at each point in time.[15]

To examine whether the relationship between exports and exchange-rate volatility may be changing over time, RC estimation was used to trace the coefficients on the volatility term for each of the twelve countries considered and for each of the three measures of volatility considered in this paper. The results were all qualitatively similar. Figures 1 through 3 illustrate the findings. These figures show the results based on the absolute percentage change measure of volatility. As shown in the figures, there is very little evidence that the relationship between exchange-rate volatility and export volume changed during the estimation period. In all cases, the coefficient on volatility is near zero.

7. Concluding remarks

As discussed above, some recent studies, using panel data, found evidence, though by no means overwhelming, of a significant and negative impact of exchange-rate volatility on trade. Although it is difficult to draw generalizations from this finding, two factors seem to be of importance. First, as noted by McKenzie (1999) in his literature survey, the use of a GARCH specification - - typically GARCH (1, 1) - - of volatility seems to produce effects of volatility on trade that are more-consistently negative and significant than other specifications. Second, studies employing panel data, typically using gravity models, tended to find negative and significant effects of volatility on trade, regardless of the measure of volatility employed. The gravity model relates trade between a given pair of countries to characteristics of each of the

[15] Recursive estimation also provides time profiles of coefficients. Unlike RC estimation, an underlying assumption of recursive estimation is that the coefficients are constants. Also, recursive estimation does not write-off the past. If a regime-change has occurred, it averages the old regime with the new regime with changing weights; the weight of the new regime becomes smaller as more and more observations are added. For a further discussion, see Brissimis, Hondroyiannis, Swamy and Tavlas (2003).

two countries and the characteristics of their relationship to each other, including economic mass (*i.e.*, GDP), distance, land areas, cultural similarities and historical links. Many of these characteristics are proxied through the use of dummy variables.

In our results employing panel data for sets of seven and 12 industrial countries respectively and employing three measures of exchange-rate volatility, we have found no evidence of a negative and significant impact of volatility on real exports, regardless of which of these measures of volatility was used. What accounts for the differences between our findings and those of recent studies using panel data which tend to find at least some evidence of a negative and significant impact? We suggest the following.

First, as discussed in Section 2, when more-sophisticated estimation techniques - - *e.g.*, fixed effects and random effects - - are employed, in those studies that do find negative and significant impacts of volatility, those impacts tend to be smaller than the impacts derived using less-sophisticated estimation techniques, even with use of the gravity model. Second, other studies did not examine the impact of real export earnings of oil-producing countries. We found that such a variable is significant with an elasticity of demand for industrial-country exports that is markedly different from the income elasticities of industrial countries. The omission of this variable from other panel-data studies indicates a source of specification bias. Third, the use of RC estimation helps account for omitted variables without the use of an assortment of dummy variable as in the gravity model.[16] In addition, RC estimation controls for endogeneity and also helps account for measurement errors and unknown functional forms. Finally, studies employing the gravity specification, by construction, deal with bilateral trade relations, whereas our study deals with aggregate trade.[17] The tendency of studies of bilateral trade to yield significant negative measures of the effect of exchange-rate volatility on those trade flows has long been recognized in the literature (see Bailey, Tavlas, and Ulan, 1986, p. 474).

The aggregate volume of international trade - - not the trade between any two jurisdictions - - is a measure of (or determines) the extent to which countries achieve the welfare gains that international trade can provide. Hence, it is the effect of

[16] As also pointed out by Bailey, Tavlas, and Ulan (1986, p. 474), there was a tendency in the earlier studies which did find a negative and significant effect of volatility on trade to use dummy variables.
[17] As discussed in Section 2, the study by Tenreyro (2004) is an exception.

exchange-rate volatility on a country's aggregate trade (rather than on its trade with any particular subset of trading partners) that determines whether that volatility has an adverse effect on the welfare gains a country derives from trade. While the factors that can make exchange-rate volatility to decrease the volume of international exchanges do reduce trade between some pairs of trading partners slightly, in the aggregate, the impacts of those factors are offset by those of the factors that tend to increase trade in the face of short-term exchange-rate changes. Hence, overall, exchange-rate volatility does not reduce the volume of or the gains from international trade.

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Table 1 Panel Data Estimation of Export Equations: Seven Countries						
Specification	Constant	Industrial Country Income	Oil Exporter Income	Real Exchange Rate	Exchange Rate Volatility Measure	Standard Error of Regression
Common fixed coefficients						
1a	-2.977 (-1.64)	0.758 (3.05)	0.218 (3.26)	-0.137 (-0.77)	-0.002 (-0.34)	0.287575
1b	-2.891 (-1.61)	0.75 (2.95)	0.217 (3.36)	-0.124 (-0.63)	-0.022 (-0.64)	0.287317
1c	-2.953 (-1.60)	0.757 (3.06)	0.219 (3.18)	-0.141 (-0.79)	-1.776 (-0.52)	0.287495
Fixed effects						
2a	-10.172 (-12.79)	1.662 (18.47)	0.085 (2.05)	-0.275 (-4.63)	0.001 (0.68)	0.097147
2b	-10.099 (-11.94)	1.652 (16.61)	0.084 (2.05)	-0.267 (-4.13)	-0.010 (-0.52)	0.097040
2c	-10.143 (-12.55)	1.659 (18.26)	0.085 (2.06)	-0.275 (-4.65)	-0.154 (-0.14)	0.097175
Random effects						
3a	-9.934 (-11.42)	1.625 (18.06)	0.090 (2.08)	-0.257 (-4.40)	0.001 (0.50)	0.110488
3b	-9.859 (-10.75)	1.615 (16.29)	0.089 (2.10)	-0.249 (-3.79)	-0.010 (-0.58)	0.110368
3c	-9.902 (-11.24)	1.622 (17.92)	0.090 (2.08)	-0.257 (-4.40)	-0.260 (-0.25)	0.110697
GMM						
4a	-9.565 (-6.95)	1.690 (14.23)	0.074 (3.22)	-0.502 (-3.80)	0.001 (0.25)	0.033056
4b	-9.556 (-6.97)	1.694 (13.02)	0.073 (2.94)	-0.509 (-4.11)	-0.005 (-0.16)	0.032864
4c	-9.512 (-6.36)	1.689 (13.13)	0.073 (3.17)	-0.510 (-3.72)	-0.373 (-0.05)	0.033014
Random coefficients						
5a	-9.559 (-12.33)	1.627 (16.48)	0.082 (1.98)	-0.336 (-3.36)	-0.001 (-0.35)	0.000324
5b	-9.211 (-10.00)	1.604 (14.88)	0.080 (2.11)	-0.356 (-3.19)	-0.014 (-1.05)	0.0003319
5c	-10.074 (-13.01)	1.659 (14.65)	0.084 (1.91)	-0.294 (-3.20)	-13.394 (-0.59)	0.0003222
Notes: The estimation period for all the models is 1977:1 to 2003:4. The figures in parentheses are the t-ratios.						

Table 2 Panel Data Estimation of Export Equations: Twelve Countries						
Specification	Constant	Industrial Country Income	Oil Exporter Income	Real Exchange Rate	Exchange Rate Volatility Measure	Standard Error of Regression
Common fixed coefficients						
1a	-5.564 (-1.36)	1.238 (2.87)	0.195 (2.84)	-0.480 (-1.18)	-0.011 (-1.18)	0.914058
1b	-5.212 (-1.32)	1.174 (2.91)	0.189 (3.03)	-0.397 (-1.00)	-0.110 (-1.13)	0.912097
1c	-5.515 (-1.31)	1.236 (2.84)	0.200 (2.89)	-0.493 (-1.19)	-9.040 (-0.84)	0.913945
Fixed effects						
2a	-10.779 (-8.21)	1.869 (15.08)	0.106 (3.55)	-0.553 (-2.76)	0.001 (0.20)	0.139874
2b	-10.649 (-8.57)	1.851 (15.71)	0.105 (3.61)	-0.537 (-2.52)	-0.019 (-0.67)	0.139484
2c	-10.764 (-8.05)	1.868 (14.96)	0.106 (3.54)	-0.553 (-2.76)	-0.115 (-0.12)	0.139878
Random effects						
3a	-10.778 (-9.57)	1.869 (15.14)	0.107 (3.56)	-0.552 (-2.77)	0.001 (0.20)	0.139760
3b	-10.647 (-9.89)	1.851 (15.77)	0.105 (3.63)	-0.537 (-2.53)	-0.019 (-0.68)	0.139369
3c	-10.763 (-9.37)	1.868 (15.02)	0.106 (3.55)	-0.553 (-2.77)	-0.116 (-0.12)	0.139762
GMM Estimation						
4a	-10.623 (-7.54)	1.858 (14.96)	0.089 (2.62)	-0.629 (-3.47)	0.001 (0.50)	0.051349
4b	-10.586 (-7.74)	1.856 (15.30)	0.091 (2.59)	-0.632 (-3.27)	0.002 (0.08)	0.054039
4c	-10.549 (-7.25)	1.850 (14.50)	0.091 (2.71)	-0.630 (-3.45)	-2.261 (-1.02)	0.051411
Random coefficients						
5a	-11.497 (-10.13)	1.840 (13.76)	0.101 (3.60)	-0.337 (-3.04)	-0.001 (-0.23)	0.000359
5b	-11.415 (-9.22)	1.826 (13.34)	0.099 (3.90)	-0.321 (-2.91)	-0.012 (-0.82)	0.000361
5c	-11.802 (-11.08)	1.867 (13.89)	0.103 (3.38)	-0.326 (-3.38)	-7.019 (-0.32)	0.000387
Notes: The estimation period for all the models is 1977:1 to 2003:4. The figures in parentheses are the t-ratios.						

Figure 1
Absolute Percentage Change Measure of Volatility and Time-Varying Coefficient Estimates
(volatility left scale, solid line, coefficient estimates right scale dot line)

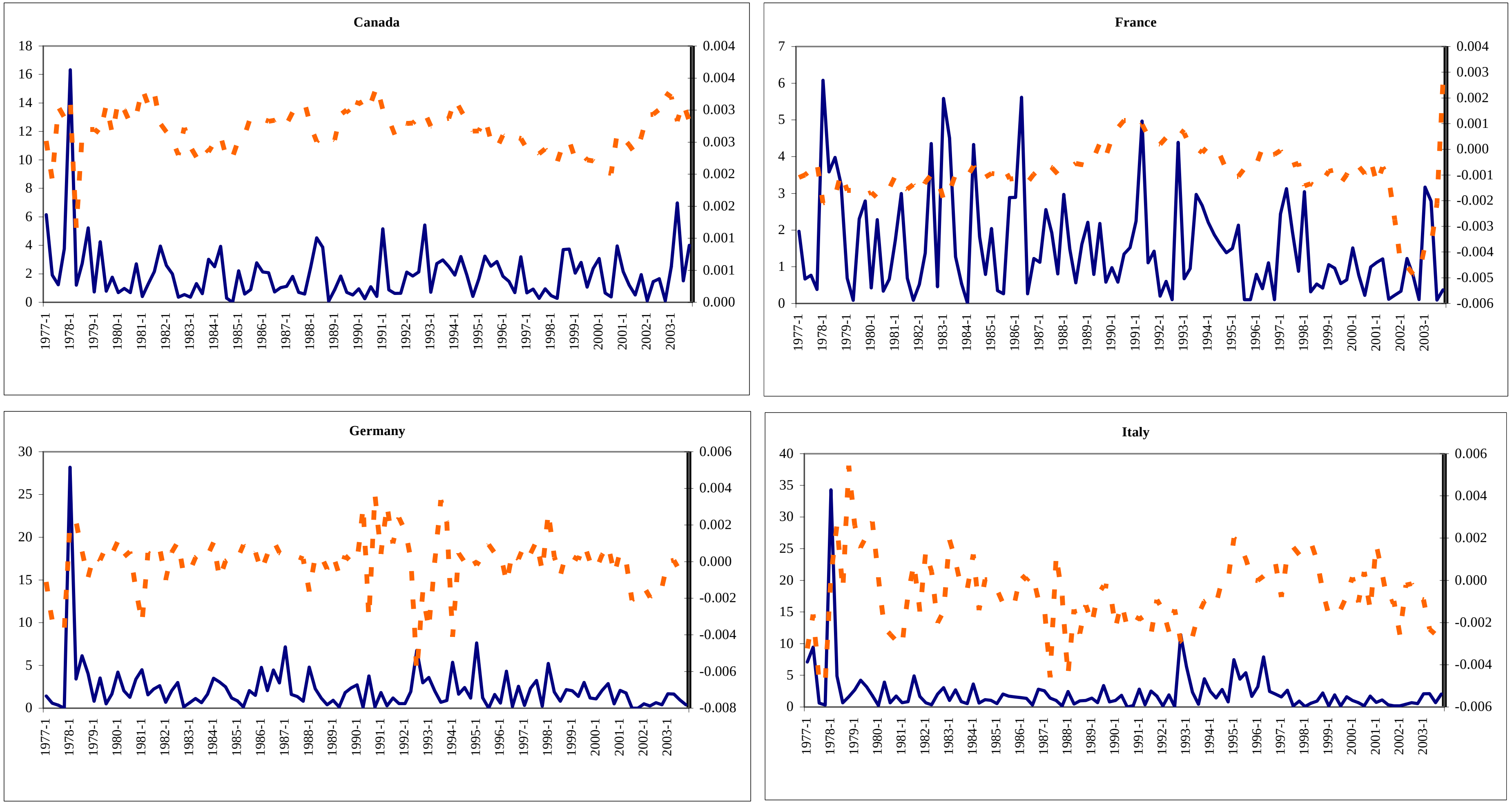


Figure 2
Absolute Percentage Change Measure of Volatility and Time-Varying Coefficient Estimates
(volatility left scale solid line, coefficient estimates right scale dot line)

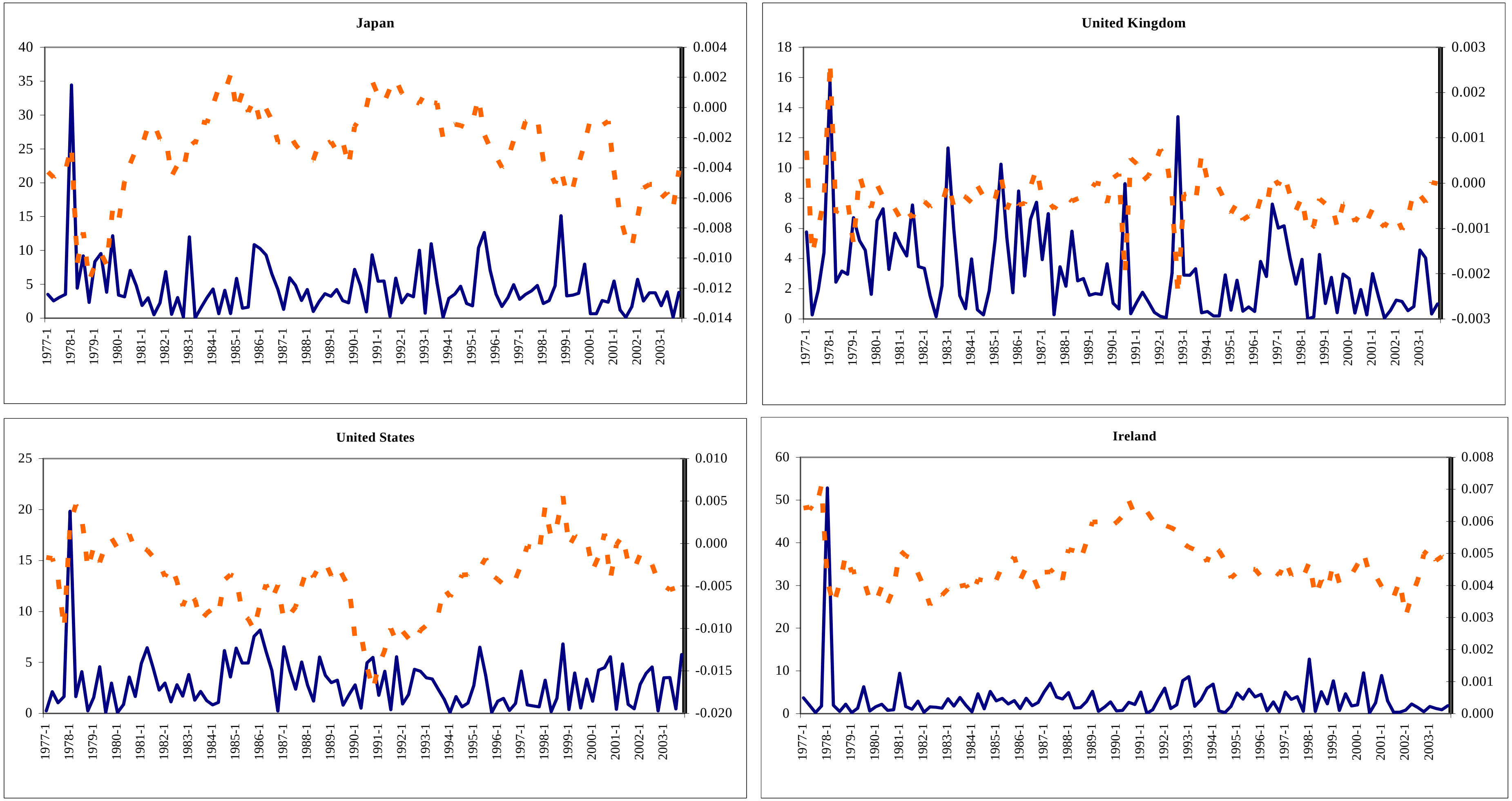
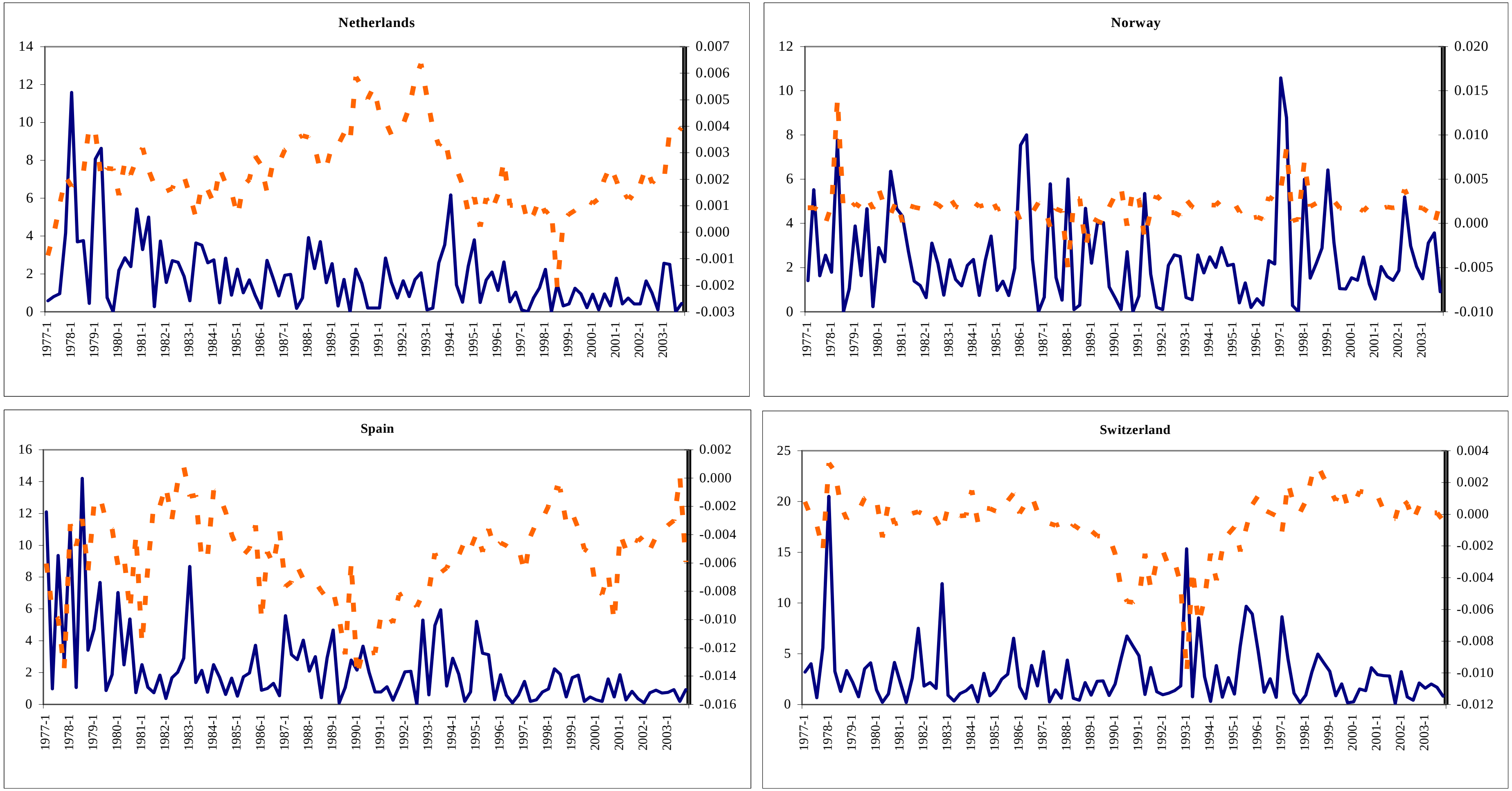


Figure 3
Absolute Percentage Change Measure of Volatility and Time-Varying Coefficient Estimates
(volatility left scale solid line, coefficient estimates right scale dot line)



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