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Exchange-rate volatility in Latin America and its impact on foreign trade

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Abstract

This paper investigates empirically the impact of real exchange-rate volatility on the export flows of eight Latin American countries over the quarterly period 1973–2004. Estimates of the cointegrating relations are obtained using different cointegration techniques. Estimates of the short-run dynamics are obtained for each country utilizing the error-correction technique. The major results show that increases in the volatility of the real effective exchange rate, approximating exchange-rate uncertainty, exert a significant negative effect upon export demand in both the short-run and the long-run in each of the eight Latin American countries. These effects may result in significant reallocation of resources by market participants.
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1. Introduction

The impact of increased exchange rate variability on foreign trade has been investigated in a large number of empirical and theoretical studies.[1] The issue is particularly important for countries that switched from a fixed to a flexible exchange rate regime due to the higher degree of variability associated with flexible exchange rates. While many Latin American countries have moved to a flexible exchange rate regime at some point in the recent past[2], it is surprising that there are very few studies that analyze the relationship between exchange rate variability and foreign trade for Latin American countries.[3]

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[1] Empirical papers on the issue include, among many others, Kenen and Rodrik (1986), Cushman (1988), Qian and Varangis (1994), Lee (1999), Doyle (2001), and Baum, Caglayan, and Ozkan (2004), while examples of theoretical contributions are Ethier (1973), Hooper and Kohlhagen (1978), De Grauwe (1988), Baldwin and Krugman (1989), Viaene and de Vries (1992), and Barkoulas et al. (2002). Surveys of the literature can be found in Côté (1994) and McKenzie (1999).

[2] The fact that some Latin American countries pegged their currency against the U.S. dollar for certain periods, such as Argentina from 1991 to 2001, does not invalidate the above statement since the real effective exchange rate used in this study continues to vary due to the fact that other Latin American countries have chosen to float their currencies against the dollar.

[3] Seabra (1995) provides estimates of the expected short-run exchange-rate uncertainty for 11 Latin American countries, but does not apply his measure to the question of trade and exchange rate variability.

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The purpose of this paper is to close this gap and provide estimates of the short- and long-run impact of exchange rate variability on export flows for eight Latin American economies.

In estimating these effects, we follow the approach introduced by Arize, Osang, and Slottje (2000) who examine the impact of exchange-rate volatility on the export flows for thirteen LDCs using both cointegration and error-correction techniques. Based on that approach, we find that the variability of the real exchange rate had a negative effect on export demand for all Latin American countries in our sample, both in the short and the long run. This result is quite surprising given that most countries in this study are middle-income economies according to World Bank classification and thus should have forward markets that would allow traders to hedge transaction exposure. But, as our results show, even fairly developed economies may not be able to completely insulate real economic flows from the fluctuations in international financial markets, and, as a result, these countries have to bear the negative consequences of such fluctuations.

Our results are on the whole consistent with the scant evidence on the relationship between exchange rate variability and export behavior of Latin American countries obtained by previous studies. Coes (1981) uses a log-level specification to examine Brazilian exports (annual data for 1965–1974) and concludes that a significant reduction in exchange-rate uncertainty in the country's economy during the crawling-peg era had a positive effect on the country's exports after the crawling peg was adopted in 1968. The study by Brada and Mendez (1988) includes 15 Latin American countries and covers the 1973–1977 period. While their conclusion is similar to ours, namely that exchange-rate uncertainty inhibits bilateral exports, they do not use a measure of exchange-rate volatility, but instead rely on a various dummy variables to account for the effects fixed versus flexible exchange rate regimes. Caballero and Corbo (1989) use a Koyck-type model and real bilateral exchange-rate volatility measure to estimate an export demand equation for six countries, among them Chile, Colombia, and Peru. They conclude that there is a strong negative effect of real exchange-rate uncertainty on the exports of all these countries.

Furthermore, the empirical results derived in this paper are also consistent with recent studies showing a significant negative (long-run) impact of exchange-rate volatility on export flows for developing countries outside of Latin America (e.g., Arize et al., 2000; Bahmani-Oskooee, 2002).[4]

The remainder of the paper is organized as follows. In Section 2, we examine the specification of our empirical model followed by a discussion of econometric methodology issues. Data sources and variable definitions are described in Section 3. In Section 4, we discuss the empirical results for the eight countries. Conclusions are drawn in the last section.

2. Model specification

A common specification of export demand in the flexible exchange-rate environment is[5]:

$$Q_t = \tau_0 + \tau_1 \cdot w_t + \tau_2 \cdot P_t + \tau_3 \cdot \sigma_t + EC_t, \tag{1}$$

where Q_t denotes the logarithm of a country's exported goods, w_t is the logarithm of a scale variable which captures world demand conditions; p_t is the logarithm of relative prices and is measured by the ratio of that country's export price in U.S. dollars to the world export price in U.S. dollars; σ_t is a measure of exchange-rate risk; and EC_t is a disturbance term. It is expected that $\tau_1 > 0$; $\tau_2 < 0$; and $\tau_3 <$ or > 0.

If foreign economic activity rises, the demand for exports will rise, so τ_1 is expected to be positive. On the other hand, if relative prices rise, the demand for exports will fall, so τ_2 is expected to be negative. Most empirical work treats exchange-rate volatility as a risk: Higher risk leads to higher cost for risk-averse traders and also to less foreign trade. This is because the exchange rate is agreed on at the time of the trade contract, but payment is not made until the future delivery actually takes place. If changes in exchange rates become unpredictable, this creates uncertainty about the profits to be made and, hence, reduces the benefits of international trade. Exchange-rate risk for developing countries is generally not hedged because forward markets are not accessible to all traders. Even if hedging in the forward markets

[4] The evidence for industrialized countries is mixed. Chowdhury (1993), Arize (1995), and Choudhry (2001) report a negative impact, while Qian and Varangis (1994) and Baum et al. (2004) find a negative effect for some countries and a positive for others. Doyle (2001) finds that in the case of Irish–U.K. trade positive effects predominate.

[5] To conserve space, no theoretical discussions on the relationship between foreign income or relative price variables and foreign trade are presented here. A treatment of this issue can be found in Arize (1990).

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were possible, there are limitations and costs. For example, the size of the contracts is generally large, the maturity is relatively short, and it is difficult to plan the magnitude and timing of all their international transactions to avail themselves of the forward markets.

There are, however, counter-arguments that suggest that τ_3 could be positive. For example, De Grauwe (1988) has stressed that the dominance of income effects over substitution effects can lead to a positive relationship between trade and exchange-rate volatility. This is because, if exporters are sufficiently risk-averse, an increase in exchange-rate volatility raises the expected marginal utility of export revenue and therefore induces them to increase exports. He suggests that the effects of exchange-rate uncertainty on exports should depend on the degree of risk aversion. A very risk-averse exporter who worries about the decline in revenue may export more when risks are higher. On the other hand, a less risk-averse individual may not be concerned with the worst possible outcome and, considering the return on export less attractive, may decide to export less when risks are higher.

Dixit (1989) has argued that sunk entry and exit costs can also influence foreign trade. Briefly, firms have to invest before they can sell their goods in another country. Some of the investments may be in research and development, relocation and distribution systems, and capital investment. The cost of these investments cannot be recouped if the firm decides to stop the export activities to a particular country. As the exchange rate becomes more volatile, firms will tend to wait longer, widening the interval in which neither exit nor entry occurs. That is, sunk exist or entry costs produce hysteresis in trade flows. As with 'risk aversion' models, it is not always clear how trade will be affected. See Arize (1997, 1995) for a more detailed discussion of this theory.

Bailey and Tavlas (1988) and Tavlas and Swamy (1997) have provided reasons why the effect of exchange-rate risk on foreign trade could be positive. The authors argue that, if traders gain knowledge through trade, enabling them to anticipate changes in exchange rates better than the average participant in the foreign exchange market, they can profit from this knowledge. That profit may offset the risk represented by movements in the exchange rate. The income earned by using such knowledge in the foreign-exchange market may offset the risk represented by movements in the exchange rate. They point out that, in any fast-changing business environment, price-affecting information is scarce and valuable, and these traders are likely to have proprietary access to some of it.

Before presentation of the empirical results, it is necessary to derive an operational measure of exchange-rate uncertainty. Various statistical measures of the variability have been suggested in the literature. The approach followed here takes into account that it is uncertainty – unpredictable component of a variable's movement – rather than variability per se, which matters most to economic agents, conditional standard deviation measures are used here. First, as Engle (1983, p. 287) notes the conditional variance is "of more relevance to economic agents planning their behavior"; therefore, exchange-rate uncertainty is proxied by the Engle (1983) model (now well known as the autoregressive conditional heteroskedasticity [ARCH] model). This model specifies the variance of a variable as a linear function of the expected squares of the lagged value of the error term from an auxiliary regression determining the mean of the variable of interest.

Assume that the conditional mean and variance of exchange rate are generated as

$$E_t = x_t\beta + u_t, \qquad u_t \sim N(0, \sigma_t^2), \qquad \sigma_t^2 = f_t\alpha \tag{2}$$

where E_t is the change in real effective exchange rate between quarter $t-1$ and t, x_t is a vector of exogenous variables in the set Ω_t of information available at t and contributing to the conditional mean $x_t\beta$ of E_t, and f_t is a vector of variables also in the information set at t and contributing to the conditional variance σ_t^2 of E_t. Expectations of the mean and variance of the change in real effective exchange rate are, in effect, assumed to be rational with respect to the information sets x_t and f_t, respectively.

Given a sample of n observations on $\{E_t, x_t, f_t\}$, estimates $\{\hat{\alpha}_t, \hat{\beta}_t\}$ of the parameter vectors can be made by maximizing the log-likelihood function for the sample, namely:

$$\ln(L) = -\frac{n}{2}\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{n} \ln(\hat{\sigma}_t) - \sum_{t=1}^{n} (\hat{u}_t/\hat{\sigma}_t^2) \tag{3}$$

The $\hat{u}_t$ are the estimated residuals $E_t - f_t\hat{\beta}$, and the σ_t^2 are the estimated variances $f_{t\hat{\alpha}}$. Again assuming that the forecaster's utility function is of an appropriate form, these time-varying variances $\hat{\sigma}_t$ can be interpreted as measures of the uncertainty surrounding the 1-quarter change in the exchange rate between $t-1$ and t.

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In order to implement this model, specific assumptions must be made concerning the elements of the vectors x_t and f_t on which the mean of exchange rate is conditioned. The function for estimating the first moment, that is, the mean of exchange rate is given as:

$$E_t = \delta_0 + \delta_1 E_{t-1} + \cdots + \delta_p E_{t-p} + \mu_t^* \tag{4}$$

where E_t represents exchange, μ_t^* is the white-noise residuals obtained from estimating Eq. (4) using ordinary least squares (OLS). It can be argued that Eq. (4) satisfies many of the requirements needed for the resulting expected exchange rate series to be rational approximations.

The first measure is represented by an ARCH (autoregressive conditional heteroskedasticity) model introduced by Engle (1983), and the pth order ARCH model of the exchange process is formulated as follows:

$$E_t = E^* + \omega_t$$

$$\omega_t^2 = h_t + e_t$$

$$h_t = E(\omega_t^2 / I_{t-1}) = \alpha_0 + \alpha_1 \omega_{t-1}^2 + \cdots + \alpha_p \cdot \omega_{t-p}^2 \tag{5}$$

where h is the conditional variance and the information set, I_{t-1} includes information available through time $t-1$, ω_t is the white-noise disturbance term and p is the lag terms in the model.

The Lagrange multiplier (LM) test for the presence of ARCH effects was conducted for lags 1, 2, 3 and 4 (four ARCH models estimated). The LM test for the ARCH is TR^2, where T is the number of observations and R^2 is the coefficient of determination for the auxiliary regression. This test statistic is distributed as $\chi^2(p)$. The ARCH variable is measured as the logarithm of the conditional standard errors. The computed LM test statistics are 10.12, 6.02, 3.30, 11.98, 5.79, 1.01, 8.02 and 25.86 for Bolivia, Colombia, Costa Rica, The Dominican Republic, Ecuador, Honduras, Peru and Venezuela. Six of these values are statistically significant at the 5% level. For Costa Rica, it is statistically significant at the 10% level, and for Honduras, the test statistic is statistically non-significant. Therefore, based on these test results, the exchange rate series could be modeled as ARCH(1) for all countries in our sample, except Honduras.

For Honduras, the exchange-rate volatility measure employed is generated by a model proposed by Antle (1983), which is denoted as the linear moment (LM) model. This model specifies the variance (and higher moments) of a variable as a linear (in the parameters) function of the regressors used in an auxiliary regression that specifies the mean of the variable of interest. Pagan, Hall, and Trivedi (1983) and Arize (1995, 1997) are but three examples of a number of studies that have used this approach. As Pagan et al. (1983) pointed out, "the major difference between the ARCH and LM methodologies lies in the type of alternative set up, with the former allowing the variance to be a function of previous forecast errors and the latter being conditional on past values of the explanatory variables." Therefore, for the purpose of implementing the linear moment model, we follow the procedure in Antle (1983) and Pagan et al. (1983) and specify the second moment of exchange rate as

$$\sigma_t^2 = c_0 + c_1 E_{t-1} + c_2 E_{t-1}^2 + \varepsilon_t' \tag{6}$$

where ε_t' is a white noise disturbance term and Eq. (4) is the conditional mean equation. The predicted values from Eq. (6) are all positive and the square root of the predicted values is denoted as the conditional standard deviation.

Finally, in order to establish whether there is a long-run equilibrium relationship among the variables in Eq. (1), we must employ the concept of cointegration. The basic idea of cointegration is that two or more nonstationary time series may be regarded as defining a long-run equilibrium relationship if a linear combination of the variables in the model is stationary (converges to an equilibrium over time). Thus, if the export demand function describes a stationary long-run relationship among the variables in Eq. (1), this can be interpreted to mean that the stochastic trend in real exports is related to the stochastic trends in the real foreign income, relative prices, and exchange-rate risk. In other words, even though deviations from the equilibrium should occur, they are mean-reverting (Arize et al., 2000). Cointegration tests in this paper are conducted using five alternative tests: Harris and Inder (1994), Shin (1994), Johansen (1992), Gregory

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and Hansen (1996) and Hansen (1992a,b,c).[6] The method of Harris and Inder (1994), Shin (1994) and Hansen (1992) employs a null hypothesis of cointegration, whereas Johansen (1992) and Gregory and Hansen (1996) estimators employ the null hypothesis of no cointegration. The use of these approaches is intended to examine the robustness of our results and increase confidence in them. It is worth noting that these tests are asymptotically equivalent to the Johansen's (1992) estimator; see for example, Stock and Watson (1993).

Also, conditional on the presence of cointegration, this study estimates and tests coefficients of the cointegrating relations using the fully modified ordinary least squares (FMLS) estimator of Phillips and Hansen (1990), the dynamic ordinary least squares (DLS) estimator of Stock and Watson (1993) and the instrumental variable estimator of Bewley (1990) and Wickens and Breusch (1988). The recent comparative Monte Carlo evidence provided by Stock and Watson (1993), Inders (1993), Banerjee, Dolado, Galbraith, and Hendry (1993), and Phillips and Loretan (1991) on the superiority of robust single equation methods (in terms of small sample properties) over several other methods guided the choice of the procedures used for obtaining long-run estimates.

3. Data and variable definitions

The eight Latin American countries examined in this study are Bolivia, Colombia, Costa Rica, The Dominican Republic, Ecuador, Honduras, Peru, and Venezuela. Brazil and Chile are left out due to the non-availability of aggregate export price indices, while Mexico is included in a previous study (Arize et al., 2000). Data were obtained from the IMF's *International Financial Statistics (IFS)*, IMF's Central Statistics Office, OECD *Main Economic Indicators* and the IMF's *Directions of Trade* (DOT) statistics and cover the 1973Q1 through 2004Q1 period.

We proxy foreign economic activity by real "world" income expressed as an index (1980 = 100) and construct a geometric average of the real income index of 17 countries: Australia, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Korea, New Zealand, the Netherlands, Sweden, Switzerland, the United Kingdom, and the United States. Following Goldstein and Khan (1978: 285), this series was calculated first as an annual series (due to lack of quarterly data on real income in a large number of countries) and then converted to a quarterly basis by using a quadratic interpolation method they recommended.

Data for individual country's export volume and unit values were taken from *IFS*, while the world export price index, PW_t, is a geometric trade-weighted average of export prices. The weights are w_{ji}, and the base period is 1980 = 100. The relative price ratio was calculated as $P_t = \ln PX_t - \ln E_t - \ln PW_t$ where PX is the export price in local currency and E is the exchange-rate index. To compute measures for exchange-rate volatility, trade-weighted effective exchange rate (*eer*) and real effective exchange rate (*reer*) were computed. They were constructed as follows (for illustrative purposes, let Bolivia be country j). The period average exchange rates are in units of domestic currency per dollar. These period averages were then expressed in index form (1980 = 1.0). The *eer* variable was calculated as: EXP $[\sum w_{ji} \ln E(i, \$, t) - \ln E(j, \$, t)]$ where EXP = exponent, ln = natural logarithm, $E(i, \$, t)$ = exchange-rate index of country i at time t and $E(j, \$, t)$ = exchange-rate index of Bolivia at time t. The real effective exchange rate was calculated as: $reer(j, t) = \mathrm{EXP}[-\ln P(j, t) + \ln E(j, \$, t) + \sum w_{ji} \ln P(i, t) - \sum w_{ji} \ln E(i, \$, t)]$ where the exchange rate terms are in units of country i (or j) currency per U.S. dollars in index form (1980 = 1.0). P is the consumer price index of country i (or j) in index form (1980 = 1.0).

4. Empirical results

4.1. Cointegration analysis

The first step in testing for cointegration in a set of variables is to test for stochastic trends in the autoregressive representation of each individual time series using conventional unit root tests. The first is the augmented Dickey and Fuller test, for which nonstationarity serves as the null hypothesis. As it is usually done in the literature, we report the

[6] The referee has pointed out that the Johansen test has several problems including the following: (i) it over-rejects the null of no cointegration, (ii) the results are not robust with respect to model specification, (iii) it tends to produce implausible point estimates of the parameters, and (iv) the cointegrating vectors have no economic interpretation. Therefore, to assess evidence of cointegration, this study uses a battery of tests to determine the presence or absence of cointegration. The Harris and Inder test used here have the advantage of distinguishing between unit root and near unit root processes (see Hargreaves, 1994).

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Table 1
Unit root tests

Country	ADF H_0: $X_t \sim I(1)$ H_a: $X_t \sim I(0)$				Lobato-Robinson H_0: $X_t \sim I(0)$ H_a: $X_t \sim I(1)$			
	Q	w	P	σ	Q	w	P	σ
Bolivia	−1.47	−2.72	−2.62	−2.57	−3.36 (0.00)	−4.77 (0.00)	−5.87 (0.00)	2.97 (0.00)
Columbia	−2.43	−2.72	−2.40	−2.77	−4.88 (0.00)	−4.82 (0.00)	−4.23 (0.00)	−3.34 (0.00)
Costa Rica	−2.79	−2.72	−1.11	−1.17	−4.99 (0.00)	−4.90 (0.00)	−2.77 (0.01)	−3.52 (0.00)
Dominican	−1.89	−2.72	−2.64	−3.33	−2.76 (0.01)	−3.20 (0.00)	−1.87 (0.06)	−3.10 (0.00)
Ecuador	−1.76	−2.72	−1.72	−3.20	−4.72 (0.00)	−4.82 (0.00)	−2.77 (0.01)	2.34 (0.02)
Honduras	−2.97	−2.72	−2.38	−2.61	−2.32 (0.02)	−3.10 (0.00)	1.78 (0.07)	−0.61 (0.54)
Peru	−1.92	−2.72	−2.45	−2.47	−4.76 (0.00)	−4.84 (0.00)	−3.23 (0.00)	−2.57 (0.01)
Venezuela	−2.59	−2.72	−2.26	−2.47	−5.35 (0.00)	−4.88 (0.00)	−3.59 (0.00)	−2.45 (0.01)

Notes: Q=real exports, w=real foreign income, P=relative price and σ=real exchange rate volatility. The augmented Dickey-Fuller (ADF) test statistics are from a model that includes a constant, trend and eight lags of the first difference of the regressand. The 5% critical value is −3.42. These test values have been adjusted using White's Heteroskedastic consistent standard errors. For the Lobato-Robinson test, the values in brackets are the p-values.

value of the ADF(k), where k is the minimum lag for white errors. According to Table 1, the ADF tests reject the null hypothesis of nonstationarity at the conventional levels, therefore we have considerable evidence that each variable is I (1) in each country. To cross-check these results, we report also the results of applying the Lobato and Robinson tests which takes stationarity as the null hypothesis. The Lobato and Robinson tests again support the conclusions of the ADF tests in regard to the integration properties of the variables.

Since a unit root has been confirmed for the data series, the question is whether there exists some long-run equilibrium relationship among real exports, foreign economic activity, relative price, and exchange-rate volatility for each country in our sample. As can be seen in Table 2, the Harris-Inder test as well as the Shin's test is conducted with

Table 2
Cointegration tests

Country	Harris-Inder			Shin's Test			Johansen		Gregory-Hansen (ADF*)			Hansen's L_c
	H_0: Cointegration H_a: No Cointegration			H_0: Cointegration H_a: No Cointegration			H_0: r=0 H_a: r=1	H_0: r=0 H_a: r=1	H_0: No Cointegration H_a: Cointegration			H_0:Cointegration H_a:No Cointegration
	3	6	8	3	6	8	λ_{max}	Trace	C	C/T	C/S	L_c
Bolivia	0.19**	0.16**	0.16**	0.045**	0.048**	0.051**	35.55	54.33	−4.08 (85Q2)	−4.10 (85Q2)	−5.02 (86Q4)*	0.54 (0.19)
Colombia	0.32**	0.21**	0.18**	0.159**	0.135**	0.123**	49.90	72.52	−4.13 (89Q4)	−5.50 (90Q3)*	−4.71 (90Q1)*	0.79 (0.06)
Costa Rica	0.10**	0.07**	0.06**	0.150**	0.109**	0.098**	37.94	53.63	−4.10 (93Q3)	−4.94 (98Q1)*	−4.80 (79Q3)*	0.43 (0.16)
Dominican	0.18**	0.13**	0.12**	0.039**	0.049**	0.076**	31.89	51.71	−4.20 (78Q3)	−3.71 (78Q3)	−4.93 (84Q2)*	0.28 (0.20)
Ecuador	0.31**	0.18**	0.15**	0.079**	0.103**	0.118**	29.07	45.14	−4.23 (93Q4)	−4.18 (86Q1)	−4.67 (93Q4)	0.49 (0.20)
Honduras	0.22**	0.15**	0.13**	0.051**	0.062**	0.072**	31.52	50.82	−3.51 (86Q2)	−3.65 (86Q2)	−4.37 (85Q2)	0.54 (0.08)
Peru	0.50	0.32**	0.27**	0.200	0.144**	0.128**	27.53	53.98	−4.73 (85Q3)*	−4.98 (85Q3)*	−5.16 (81Q4)*	0.36 (0.15)
Venezuela	0.62	0.38	0.31**	0.221	0.170	0.155**	34.21	55.02	−4.02 (78Q4)	−4.37 (80Q1)	−4.84 (89Q4)*	0.63 (0.13)

Notes: For the Harris-Inder test, the critical value is 0.32 at the 5% level. In the case of the Shin test, it is 0.159. The critical values for $p-r$=4 in the case of Johansen are 27.07for λ_{max} and 47.21 for the Trace test. Since we test for one cointegration vector; hence p, the number of variables, is 4, and r, the number of cointegrating vector is zero. The critical values are from Osterwald-Lenum (1992, Table 1.1*). The lag orders are five for The Dominican Republic, and Honduras; four for Bolivia, Colombia, Peru, and Venezuela; and two for Costa Rica, and Ecuador. The critical values for the Gregory and Hansen (1996) test are −4.34, −4.72, and −4.68 for C, C/T, and C/S, respectively. The p-values for Hansen's test are in parentheses beside the L_c test statistic.

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three, six and eight lags. The null of stationarity (or cointegration) is not rejected at all lag lengths in the case of Peru and Venezuela for these tests. However, as the authors suggest some degree of augmentation in the tests is needed for better results. As the data show in all cases at higher lags, the null hypothesis of cointegration is accepted. Also, these results are further corroborated by the Johansen test results which indicate not only presence of cointegration but also the presence of a single cointegrating relation.

Having provided evidence concerning cointegration and some relevant hypotheses, it seems prudent to examine cointegration allowing for structural change. We allow for changes in the cointegrating relations by applying the Gregory and Hansen (1996) cointegration test and implementing the augmented Dickey-Fuller (ADF*) test of no cointegration against cointegration, with a structural shift in cointegration in an unknown timing. The three models they put forward are (i) a model with level shift (mean model) (C); (ii) a model with linear trend incorporating a change in level (i.e., slope model) (C/T); and a model including both a change in level and in the coefficients of the variables in the long-term relation (C/S).

Gregory and Hansen (1996) offer suggestions as to how to use the ADF* statistic. If the null hypothesis of no co-integration is rejected by either the standard tests or the ADF* then such finding will suggest some long-run relationship among non-stationary variables. If the standard test does not support cointegration but the ADF* does reject the null hypothesis of no cointegration then structural change may be relevant. If both the standard tests and the ADF* test find evidence of cointegration, then no information concerning structural change is obtained since the ADF* has power to detect cointegration when there is no structural shift. In this case, further investigation is necessary to enable a distinction to be drawn between cointegration with stable parameters and cointegration with a structural shift. Gregory and Hansen (1996) suggest using Hansen's (1992a,b,c) parameter instability tests (e.g., Hansen's L_c) to determine whether there has been a shift in the cointegrating relationship. The results based on the ADF* statistics are set out in Table 2.

The first model (mean model) reveals that cointegration is present with a break for Peru, implying that the data support cointegration with a change in intercept. The second model (slope model) shows that cointegration is present with a break for Colombia, Costa Rica, and Peru.

Finally, the third model that takes into consideration the simultaneous presence of both a mean break and a slope break (regime shift) exhibits cointegration in seven cases; Bolivia, Colombia, Costa Rica, The Dominican Republic, Ecuador, Peru and Venezuela. In the case of Honduras, the conventional tests find evidence of cointegration, while the ADF* statistic does not, it can be concluded that the model is structurally stable. According to Hansen's L_c test results in Table 2, the null of cointegration is supported at the 10% level in all cases. All these results are consistent with cointegration and the existence of a stable long-run relationship between real exports and their determinants.

In Table 3 we show the empirical results of estimating and testing the coefficients of Eq. (1) using three alternative approaches, the fully modified ordinary least squares (FMLS) estimator of Phillips and Hansen (1990), the dynamic ordinary least squares (DLS) estimator of Stock and Watson (1993) and the instrumental variable estimator of Bewley (1990) and Wickens and Breusch (1988).

Focusing on the results obtained from the FMLS estimator, the estimated foreign economic activity (w_t) elasticity carries the expected positive sign and is significantly different from zero (at the 5% level) in all the countries in our

Table 3
Long-run elasticities and hypothesis tests

Country	Phillips-Hansen Estimator (FMLS)			Stock-Watson Estimator (DLS)			Bewley-Wickens-Breusch (BWB)		
	w	P	σ	w	P	σ	w	P	σ
Bolivia	1.09 (5.84)	−1.15 (11.18)	−0.001 (1.68)	1.05 (5.58)	−1.26 (10.95)	−0.001 (1.99)	0.51 (1.90)	−0.84 (5.35)	−0.002 (1.75)
Colombia	2.25 (16.32)	−0.57 (5.77)	−0.005 (1.86)	2.04 (20.52)	−0.68 (11.12)	−0.002 (1.45)	2.30 (9.57)	−0.62 (5.55)	−0.01 (1.49)
Costa Rica	5.84 (18.27)	−0.29 (1.21)	−0.26 (3.68)	5.22 (14.31)	0.19 (0.74)	−0.40 (4.86)	5.77 (15.38)	−0.30 (1.04)	−0.27 (3.25)
Dominican	5.23 (19.96)	−1.31 (7.58)	−0.07 (2.44)	5.05 (23.14)	−1.30 (5.90)	−0.21 (3.58)	5.06 (4.60)	−1.16 (6.54)	−0.18 (3.76)
Ecuador	1.38 (13.55)	−0.29 (3.03)	−0.001 (2.15)	1.45 (20.83)	−0.29 (4.37)	−0.07 (4.27)	1.38 (19.17)	−0.33 (4.87)	−0.001 (3.41)
Honduras	0.93 (5.22)	−0.31 (2.52)	−0.05 (2.07)	0.89 (7.69)	−0.37 (4.14)	−0.01 (4.69)	0.93 (3.81)	−0.31 (1.79)	−0.001 (1.81)
Peru	0.99 (6.28)	−0.82 (7.20)	−0.01 (3.60)	0.96 (12.67)	−0.75 (9.66)	−0.002 (7.07)	1.08 (8.97)	−0.80 (8.40)	−0.001 (3.55)
Venezuela	6.75 (12.89)	−1.38 (3.52)	−0.11 (2.89)	6.78 (27.65)	−1.50 (6.79)	−0.16 (4.95)	7.50 (8.12)	−2.29 (3.00)	−0.20 (1.78)

Notes: The numbers in parentheses report absolute t-statistics. The critical value at 10% is 1.3 and 1.67 at 5% level.

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sample. The long-run income elasticity is greater than unity in all countries except for Honduras and Peru, greater than two in four countries, and greater than three in three countries. There are several explanations for the relatively high income elasticities. First, and foremost, it must be noted that the values for the income elasticities are consistent with estimates found in other studies. As noted by Riedel (1988) most estimates of income elasticities in export demand equations, "whether for developed or developing countries, or for country aggregates or in individual countries, generally lie in the range between 2.0 and 4.0" (p. 140). Of the six studies surveyed in Marquez and McNeilly (1988, Table 1, p. 307) four report income elasticities greater than two and three report elasticities greater than three. Riedel (1988) estimates the income elasticity for Hong Kong's exports of manufactures to be greater than four.

Riedel (1988, 1989) conjectures that the high elasticities found in the literature reflect the inadequate treatment of both the supply side of exports and the normalization issue. His estimate of a simultaneous equation model with export demand normalized as a price equation yields a lower income elasticity. For a critique of Riedel's approach, see Nguyen (1989). A different explanation for high income elasticities has been given in Arize (1990). He argues that an increased penetration of world markets over the sample period can, in part, be attributed to the income elasticities of LDCs being some function of the income elasticities of the exports of the importing countries. This is plausible if exports are largely composed of semi-finished products which are used to produce final products in other countries. Finally, Adler (1970) has suggested that different income elasticities reflect the extent to which exports have been adapted to the importing country's local tastes, with higher elasticity providing evidence of greater adaptation.

The estimated price (p_t) elasticity has the expected negative sign in the eight countries studied. For Costa Rica, we obtain a negative price elasticity that is statistically insignificant.

The semi-elasticity estimates of the exchange-rate volatility (σ_t) have negative signs throughout and are statistically significant for each country. The semi-elasticity estimates range from −0.26 (Costa Rica) to −0.001 (Bolivia), implying that exchange-rate volatility exerts a significant adverse long-run effect on export volume. The results, presented in Table 3, for DLS and BWB estimators are very similar to those discussed above in the case of the FMLS estimator, despite the fact that these estimators impose different assumptions regarding endogeneity, and differ in how they correct for autocorrelation.

4.2. Error-correction model

The Granger representation theorem proves that, if a cointegrating relationship exists among a set of $I(1)$ series, then a dynamic error-correction representation of the data also exists. The methodology used to find this representation follows the "general-to-specific" paradigm (see Hendry, 1987). Initially, four lags of the first-difference of each variable in Eq. (1), a constant term and one-lagged error-correction term (EC_{t-1}) generated from the Johansen procedure were used. Then the dimensions of the parameter space were reduced to a final parsimonious specification by sequentially imposing statistically insignificant restrictions or eliminating insignificant coefficients. Given the presence of the volatility variable in the error-correction model (ECM) and the endogeneity of some of the regressors, we use the instrumental variables procedure suggested by Pagan and Ullah (1988). The list of instrumental variables consists of the constant term, the lagged EC term, and four lags in the differences of all variables included in the long-run solution. In their paper, Pagan and Ullah recommend the use of a heteroskedasticity and serial correlation consistent estimator of the covariance matrix. To ensure that the covariance is positive semi-definite, we adjust Pagan and Ullah's covariance estimator as suggested by Newey and West (1987). The results are summarized in Table 4.

Considering that each regressand in Table 4 is cast in first-difference, the empirical results suggest that the statistical fit of each model to the data is satisfactory, as indicated by the values of adjusted R^2, which range from a low of 0.33 in Ecuador to a high of 0.73 in Venezuela. Moreover, the statistical appropriateness of the equations is supported by the diagnostic tests. In particular, the stability of each estimated error-correction model is confirmed by Hansen's (1992a,b,c) parameter nonconstancy (J_t) test for stationary data. Also, each estimated model fulfills the conditions of serial noncorrelation, homoskedasticity, zero disturbance mean (i.e., no specification errors), and normality of residuals. In addition, we correct for potential endogeneity of the right-hand side variables by using an IV estimation approach.

Having provided evidence supporting the adequacy of the estimated equations, we can make the following observations regarding the obtained estimates:

First, the error-correction term's coefficient is statistically significant in each of the eight cases and is always negative, as expected. These findings support the validity of an equilibrium relationship among the variables in each

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Table 4
Regression results for error-correction models

Variables	Country							
	Bolivia	Colombia	Costa Rica	Dominican Republic	Ecuador	Honduras	Peru	Venezuela
EC_{t-1}	−0.17 (2.14)	−0.23 (3.39)	−0.12 (3.65)	−0.08 (2.86)	−0.45 (5.41)	−0.35 (3.94)	−0.08 (1.75)	−0.03 (2.06)
ΔQ_{t-1}	−0.20 (2.11)	−0.15 (1.64)						0.20 (6.19)
ΔQ_{t-2}	−0.24 (2.61)							
ΔQ_{t-3}	−0.20 (2.29)		0.16 (1.62)				0.05 (3.28)	
ΔQ_{t-4}	0.32 (3.91)	0.29 (3.86)	0.26 (3.66)				0.35 (5.38)	
ΔQ_{t-5}		−0.13 (1.58)						
Δw								
Δw_{t-1}	2.05 (1.94)				2.70 (1.99)	2.81 (2.36)		1.51 (1.74)
Δw_{t-2}			2.68 (1.66)					
Δw_{t-3}		1.27 (1.38)			3.02 (1.94)			
Δw_{t-4}								
Δw_{t-5}				1.27 (1.63)			1.68 (1.68)	
ΔP_t		−0.38 (4.30)	−1.21 (5.20)					−0.94 (53.15)
ΔP_{t-1}	0.68 (4.73)	−0.14 (1.44)	−0.35 (2.52)	−1.00 (89.41)		−0.16 (3.95)	−0.87 (34.65)	
ΔP_{t-2}					−0.29 (5.09)			
ΔP_{t-3}								
ΔP_{t-4}							0.30 (5.42)	
ΔP_{t-5}							−0.09 (3.39)	
$\Delta\sigma_t$								
$\Delta\sigma_{t-1}$	-3.00E^{-04} (1.65)		−0.02 (2.09)		-1.00E^{-04} (2.43)			−0.002 (1.79)
$\Delta\sigma_{t-2}$								
$\Delta\sigma_{t-3}$								
$\Delta\sigma_{t-4}$						-5.00E^{-05} (2.03)	-2.00E^{-04} (3.60)	
$\Delta\sigma_{t-5}$		−0.001 (3.20)		−0.002 (2.48)				
Adj. R^2	0.44	0.42	0.42	0.66	0.33	0.37	0.83	0.78
DW	2.07	2.02	2.20	1.89	1.98	2.05	2.18	1.48
Serial Corr	χ^2 [4]=5.85	χ^2 [4]=2.07	χ^2 [4]=6.70	χ^2 [4]=0.38	χ^2 [4]=0.55	χ^2 [4]=2.66	χ^2 [4]=7.48	χ^2 [4]=12.16
HET	χ^2 [1]=1.18	χ^2 [1]=0.03	χ^2 [1]=7.89	χ^2 [1]=0.002	χ^2 [1]=0.13	χ^2 [1]=12.27	χ^2 [1]=5.34	χ^2 [1]=0.09
JT (5% Critical)	1.64 (2.32)	2.30 (2.32)	1.70 (2.11)	1.41 (1.47)	1.11 (1.68)	1.41 (1.47)	1.16 (2.32)	1.32 (1.68)

Notes: Figures in parentheses are the absolute t−statistics. The critical value at 10% is 1.3 and 1.67 at 5% (1-tail). DW tests first-order residual autocorrelation. Serial Corr is an χ^2 [4] test for mth-order general autoregressive and moving-average residual autocorrelation. HET χ^2 [1] is the Koenker-Bassett test for heteroscedasticity. JT is Hansen's (1992b) parameter constancy test.

cointegrating equation. This implies that overlooking the cointegrating relationships among the variables would have introduced misspecification in the underlying dynamic structure.

Second, the change in real exports per quarter that is attributed to the disequilibrium between the actual and the long-run equilibrium levels is measured by the absolute values of the error-correction term of each equation. There is substantial inter-country variation in the adjustment speed to the last period's disequilibrium, with Ecuador having the largest value and Dominican Republic and Peru the smallest. This implies that the adjustment of export volume to changes in the regressors may take about two quarters in Ecuador to more than twelve quarters in Dominican Republic and Peru. The results point to the existence of market forces in the export market that operate to restore long-run equilibrium after a short-run disturbance.

Third, and foremost, since the sum of the estimates on current and lagged values of $\Delta\sigma_t$ is negative for all countries, we conclude that exchange-rate volatility has a negative short-run effect on foreign trade in addition to its adverse long-run effect established earlier.

Finally, the dynamics of the equation show that changes in foreign economic activity, relative price, and exchange-rate volatility have short-run effects on exports with can last for more than 58 quarters for certain variables and countries. Results regarding the mean time lag for the adjustment of exports are summarized in Table 5.

The evidence shows that, for seven of the eight countries in the sample, export volume responds faster to exchange-rate volatility changes than to relative price changes. But Table 5 also shows that for half of the countries in the sample

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Table 5
Mean time lags for adjustment of exports

Country	Mean time lags		
	Foreign income	Relative price	Exchange-rate volatility
Bolivia	4.29	11.76*	7.77*
Colombia	1.26	4.87*	4.27*
Costa Rica	17.54	7.71*	4.96*
Dominican Republic	3.38	25.00*	12.53*
Ecuador	10.48	2.87*	2.22*
Honduras	5.17	3.31*	2.86*
Peru	13.49	15.81*	7.51*
Venezuela	23.67	58.00*	26.73*

Notes: * = absolute values.

exports react faster to changes in foreign income than to changes in exchange-rate volatility. Therefore, ignoring the short- and long-run impact of exchange-rate volatility, as several previous studies on export demand have done, can produce biased results due to misspecification error.

5. Summary and conclusions

Our results concerning the effects of exchange-rate volatility on export flows suggest that there is a negative and statistically significant *long-run* relationship between export flows and exchange-rate volatility in each of the eight Latin American countries. In addition, we also find evidence for a negative *short-run* effect of exchange-rate volatility on export flows in all Latin American countries studied.

Our results have several policy implications. First, and foremost, economic policies that aim to stabilize the exchange rate (of which the establishment of a common currency area would be the most pronounced) are likely to increase the volume of trade among Latin American countries. Second, attempts to extend the North American Free Trade Agreement southward may find little support from Latin American countries, if the potential welfare gains through trade expansion are called into question through reduction in trade due to increased exchange rate variability. Finally, the intended positive effect of a trade liberalization policy may not only be doomed by a variable exchange rate but could also precipitate a balance-of-payments crisis.

It is worth noting that the approach we have used here to investigate the relationship between export flows and exchange-rate volatility for eight Latin American countries is characterized by a number of important econometric features typically not found in other empirical studies on this topic. First, the data set for each country covers the current floating exchange-rate era and thus allows us to address the stability over time of the estimated dynamic models during this period. This is essential for appropriate policy conclusions to be inferred from the estimated results. Second, by considering an error correction model, this study provides estimates of the speed of adjustment or the average time lag for adjustment of exports to changes in the explanatory variables as well as the short-run effects of exchange-rate volatility on exports. Third, each estimated model satisfies several recently developed econometric tests in the analysis of time-series data for issues such as cointegration, stationarity, specification errors, residual autocorrelation, heteroskedasticity, residual normality, and structural stability.

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