HARALD HAU

Real Exchange Rate Volatility and Economic Openness: Theory and Evidence

This paper relates the volatility of the trade-weighted effective real exchange rate to the degree of trade openness of an economy. The theoretical part presents an intertemporal monetary model of a small open economy with nominal rigidities. Both monetary and aggregate supply shocks are shown to produce (ceteris paribus) smaller real exchange rate movements if the country is more open to foreign trade. Empirical evidence on a cross section of forty-eight countries confirms this relationship: Differences in trade openness explain a large part of the cross-country variation in the volatility of the effective real exchange rate.

ELIMINATING REAL EXCHANGE RATE VOLATILITY represents the foremost economic objective of Europe's common currency. The benefits of this monetary union have often been evaluated against the alternative scenario of free floating exchange rates in an fairly integrated European product market. But how are trade integration and openness related to real exchange rate volatility? Does increasing economic integration provide more stable real exchange rates even in the absence of a monetary union? While these questions are of great policy interest, surprisingly little is known about the fundamental determinants of real exchange rate volatility. Supporters of European Monetary Union, for example, have pointed to large real exchange rate volatility of the DM/dollar rate or yen/dollar rate as the alternative to a common European currency.[1] However, both these exchange rates are intercontinental exchange rates and therefore correspond to a relatively low degree of underlying trade integration. We argue on theoretical and empirical grounds that economic openness and real exchange rate volatility are inversely related. A high de-

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1. See, for example, Wyplosz (1997). Padoa-Schioppa (1995) affirms that trade integration is no remedy to the incompatibility of fixed exchange rates, free capital movement and independent national monetary policies often termed the "inconsistent trinity." But even if this is correct, the "consistent trinity" of flexible exchange rates, free capital movement, and independent monetary policies may be a much more attractive alternative if there exists a structural link between high trade integration and low volatility.

HARALD HAU *is an associate professor in finance at INSEAD (France) and research affiliate at the Center for Economics Policy Research (CEPR), London. E-mail:* harald.hau@free.fr

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gree of trade integration tends to provide more stable real exchange rates, while intercontinental exchange rates like the DM/dollar or the euro/dollar rates with low degrees of underlying integration should be extremely volatile.

Our insights in this paper build on recent work on the international transmission mechanism under nominal rigidities. A robust theoretical prediction emphasized by Hau (2000a, 2000b) and Obstfeld and Rogoff (2000) is that more-open economies should exhibit less volatile real exchange rates. The intuition for this effect is straightforward. More imported goods provide a channel for a quick adjustment of the domestic aggregate price level. This in turn reduces any short-run effect of a money supply or real shock on the real household balances and therefore reduces the scope of such a shock to develope real effects on either domestic consumption or the real exchange rate. Relatively closed economies with little aggregate price flexibility due to a lower import share are deprived of this aggregate price level flexibility transmitted through the exchange rate and therefore produce (ceteris paribus) more pronounced effects on consumption and the real exchange rate. Incomplete exchange rate pass-through in the short run does not change the nature of the argument, but may just imply that the structural link between real exchange rate volatility and openness is more difficult to detect over short measurement periods. Measuring real exchange rate volatility at a sufficiently low frequency allows us to find clear evidence that more-open economies have indeed substantially lower real exchange rate volatility.[2] We consider this evidence the main contribution of the paper.

It is by now well established that the law of one price (LOP) does not hold tightly even for fairly disaggregated commodity categories. More recently the role of distance and border effects for the effective segmentation of markets has been explored in detail (Engel and Rogers 1996, 2001). Notwithstanding such evidence for incomplete pass-through, Obstfeld and Rogoff (2000) provide evidence that the conventional pass-through assumption is empirically more relevant as the medium and long-run macroeconomic benchmark. Our statistical approach tries to minimize the role for the short-run LOP violations by measuring the exchange rate volatility for three-year changes. We also highlight that our volatility measures are based on effective real exchange rates. Previous work was frequently based on bilateral U.S. dollar exchange rate which can be of little relevance for structural macroeconomic linkages if the country in question does not have the United States as its predominant trading partner and vice versa.[3] By contrast the effective exchange rate is the statistical equivalent to the exchange rate in models that summarize foreign countries as a single hypothetical foreign trading partner.

2. Our distinction between a traded and a nontraded goods is based on the representation of this good in the foreign country's consumption basket as opposed to a technological definition of transportability. Large closed economies might have as many transportable goods in their consumption basket as small open economies, but tend to consume fewer foreign imports. Openness in this paper is suitably captured by the average of the import and export share of GDP.

3. An example is Engel (1999). Using bilateral real dollar exchange rates in a volatility decompostion, he finds a role of relative nontradeable prices only for Canada, but not for various European countries. This is not surprising since only Canada and the United States are each other's predominant trading partners. Martin (1998) also uses bilateral exchange rates for OECD countries and examines the role of country size for real exchange rate volatility.

The statistical work is based on the effective real exchange rate reported in the International Financial Statistics (IFS). For the period 1980 to 1998, we retain a sample of forty-eight countries and find a significant negative relationship between openness and real exchange rate volatility. But since the theoretical linkage between openness and real exchange rate volatility depends also on the magnitude of the monetary and real shocks of each country, it seems reasonable to concentrate the analysis on the more homogeneous OECD subsample. It seems more plausible that OECD countries experience relative shocks of similar magnitude. And the results are indeed more pronounced for the OECD subsample. Trade openness explains a large part of the cross-country variation in the long-run volatility of the effective real exchange rate. We also find that the openness-volatility linkage is robust to various control variables, for example, per capita GDP, central bank independence, and exchange rate regime choice.

In addition, we explore three alternative explanations for the openness-volatility linkage. First, we examine the reverse causality from real exchange to openness. Previous research has tried to identify exchange rate risk as a trade impediment. Following Romer (1993), we use the land area of a country as a suitable exogenous instrumental variable for openness. A reestimation of the volatility-openness regression does not change the point estimates and provides evidence against the risk-trade channel. Second, the presented evidence might be simply based on the failure of the law of one price (LOP) and unrelated to nontradeables. If LOP deviations are related to distance (Engel and Rogers 1996, 2001) and openness is related to country size and distance, we might pick up the failure of the LOP for tradeables. While this explanation is plausible for high-frequency measures of the real exchange rate, we argue that the LOP deviations are a less plausible explanation for measurement intervals of three years. Also the openness-volatility linkage becomes less pronounced as we measure volatility at higher frequency, for which LOP deviations become a more reasonable assumption. Third, the openness-volatility linkage might simply result from a closer co-movement of domestic and foreign monetary shocks for more-open economies. We reject this explanation because effective relative price level volatility measured at the same low frequencies is uncorrelated with openness.

The remainder of the paper is organized as follows. Section 1 establishes the theoretical result for the simple case of a small open economy. We consider both a monetary and a labor supply shock. Simplicity and transparency in the presentation is given preference over generality of the model framework. Section 2 explains the data and presents the basic empirical results. Alternative explanations for these results are examined and discussed in section 3. Section 4 concludes.

1. THE MODEL

For the sake of simplicity, we use a model from the Obstfeld and Rogoff textbook (1996, p. 689). A small open economy is endowed with a competitively priced export good and imports a competitively priced import good. The nontraded sector makes monopolistic production decisions and sets domestic product prices. This elementary setup can be generalized to allow for monopolistic competition in both the traded and nontraded sector in a two-country framework. Further extensions by Hau (2000a, 2000b) and Obstfeld and Rogoff (2000) add labor markets and labor market rigidities. However, the basic result for the role of nontradeables as a determinant of real exchange rate volatility can be obtained in the most elementary setup and is robust to these various extensions.

A representative home household j is endowed with a constant quantity $\bar{y}_T$ of an export good in each period and also participates in the nontraded sector as the monopolistic producer of a nontraded good z on the unit interval of nontraded goods, $z \in [0,1]$. The intertemporal utility function of the household is given by

$$U_t^j = \sum_{s=t}^{\infty} \beta^{s-t} \left[\gamma \log C_{T,s}^j + (1-\gamma) \log C_{N,s}^j + \chi \log \frac{M^j}{P} - \frac{\kappa}{2} y_{N,s}(j)^2 \right], \quad (1)$$

where C_T denotes the consumption of an imported good with price P_T^{im} and C_N the composite nontraded goods consumption defined by

$$C_N = \left[\int_0^1 c_N(z)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta-1}{\theta}} . \quad (2)$$

Households associated utility benefits with holding real money balances M/P and disutility with the production of the nontraded good z at quantity $y_N(j)$. We can define a price index P as the minimal costs of buying one unit of the composite consumption $C_T^\gamma C_N^{1-\gamma}$. This price index follows as

$$P = \frac{P_T^\gamma P_N^{1-\gamma}}{\gamma^\gamma (1-\gamma)^{1-\gamma}} , \quad (3)$$

where P_N represents the nontraded goods price index defined as

$$P_N = \left[\int_0^1 p(z)^{1-\theta} dz \right]^{\frac{1}{1-\theta}} . \quad (4)$$

Domestic import and export prices are identical ($P_T = \mathrm{P}_T^{im} = \mathrm{P}_T^{ex}$) and linked to a constant world price P_T^* by the exchange rate E, such that

$$P_T = EP_T^* . \quad (5)$$

We assume the existence of an international bond market with real bonds denoted in terms of tradeables. Let the world interest rate be given by $1 + r = 1/\beta$. The intertemporal budget constraint for the representative household is given by

$$P_{T,t}B^j_{t+1} + M^j_t = P_{T,t-1}(1+r)B^j_t + M^j_{t-1} + p_{N,t}(z)y_{N,t}(j) + p_{T,t}\bar{y}_T \\ - P_{T,t}C^j_{T,t} - P_{N,t}C^j_{N,t} - \tau_t \,, \tag{6}$$

where B_t represents the bond portfolio and τ_t denotes household taxes. We simplify the analysis by assuming a balanced government budget without government consumption. Seignorage income is distributed to the households, hence $-\tau_t = M_t - M_{t-1}$.

1.1 First-Order Conditions

The monopolistic competition under the constant price elasticity θ implies a demand function for nontradeables given by

$$y^d_N(z) = \left[\frac{p(z)}{P_N}\right]^{-\theta} C^A_N \,. \tag{7}$$

Equation (7) is linear in the aggregate demand in nontradeables, $C^A_N = \int_0^1 C^j_N dj$.

Next we maximize the utility function (1) under the budget constraint (6) with respect to the household's choice variables $C^j_{T,t}$, $C^j_{N,t}$, M^j_t, and $p_N(j)$. The four resulting first-order conditions are

$$C^j_{T,t+1} = C^j_{T,t} \,; \tag{8}$$

$$\frac{\gamma}{P_{T,t}C^j_{T,t}} = \frac{\chi}{M^j_t} + \beta\frac{\gamma}{P_{T,t+1}C^j_{T,t+1}} \,; \tag{9}$$

$$\frac{\gamma}{P_{T,t}C^j_{T,t}} = \frac{1-\gamma}{P_{N,t}C^j_{N,t}} \,; \tag{10}$$

$$\frac{\gamma}{P_{T,t}C^j_{T,t}} p_{N,t}(j) = \frac{\theta}{\theta-1}\kappa y_{N,t}(j) \,. \tag{11}$$

Equation (8) represents the common Euler condition for optimal intertemporal consumption smoothing for traded goods. It is obtained for a bond market discount rate equal to the utility discount rate, $\beta(1 + r) = 1$. The utility maximizing trade-off between consumption spending in period t (providing marginal utility $\gamma/P_{T,t}C^j_{T,t}$) and a combination of one-period money holding and consumption spending in period $t + 1$ is characterized by equation (9). Equation (10) states that the marginal utility of traded and nontraded consumption must be equal at any given time. Finally, the condition for optimal monopolistic price setting is expressed in equation (11). The left-hand side states the marginal consumption utility of a unit of nontraded good with price $p_{N,t}(j)$ and the right-hand-side term $\kappa y_{N,t}(j)$ denotes the marginal disutility of production for an additional unit. Monopolistic price setting incorporates a price mark-up of $\theta/(\theta - 1)$ relative to the competitive equilibrium where the marginal consumption utility equals marginal production cost.

Substituting equation (8) into equation (9) allows us to express the demand for real balances as

$$\frac{M_t}{P_t} = \frac{\chi}{\gamma} \frac{C_{T,t} \frac{P_{T,t}}{P_t}}{\left(1 - \beta \frac{P_{T,t}}{P_{T,t+1}}\right)} \tag{12}$$

The demand for real balances is therefore determined by the consumption of tradeable, $C_{T,t}$, but also depends on the tradeable price change, $P_{T,t}/P_{T,t+1}$, and on the price ratio $P_{T,t}/P_t$.

1.2 Steady-State Equilibrium

The next step consists in describing the steady-state solution under flexible prices. It provides the reference values (denoted by overbars) around which we can linearize the model to capture the model dynamics under nominal rigidities.

We assume that the domestic households initially hold a zero net stock of foreign bonds. Given the consumption-smoothing desire expressed in equation (8) and a constant net endowment of the export good, it follows directly that domestic households consume tradeables at the endowment level, $\bar{y}_T = C_{T,t} = \bar{C}_T$. Similarly, market clearing for nontradeables under symmetric household production implies $C^A_{N,t} = y_{N,t}(z) = C_{N,t}$ for all z. Together with equations (10) and (11), we obtain the nontradeable steady-state production and consumption as

$$\bar{y}_N = \bar{C}_N = \left[\frac{(\theta - 1)(1 - \gamma)}{\kappa\theta}\right]^{\frac{1}{2}}. \tag{13}$$

Constant tradeable prices in steady state determine the aggregate price level as

$$\bar{P} = \frac{\gamma}{\chi} \frac{(1 - \beta)}{\bar{C}_T} \bar{M}, \tag{14}$$

and the steady-state nominal exchange rate as

$$\bar{E} = \frac{\gamma}{\chi} \frac{(1 - \beta)}{\bar{C}_T} \frac{\bar{M}}{P*}, \tag{15}$$

1.3 Unanticipated Permanent Monetary Shocks

We can now consider the effect of an unanticipated permanent monetary shock that occurs in period 1. Let a new (sans serif) typeface mark the percentage change of any variable. We distinguish the temporary percentage change, $\mathsf{X} = (X_1 - \bar{X}_0)/\bar{X}_0$, from the permanent percentage change of the steady state, $\bar{\mathsf{X}} = (\bar{X} - \bar{X}_0)/\bar{X}_0$.

Nominal rigidities for nontraded goods are important features of our model, therefore $\mathsf{P}_N = 0$.[4] By contrast, tradeable prices react to the monetary shock. In our setup, they are tied to the constant world price of tradeables P_T^* according to $P_T = P_T^* E$ or $\mathsf{P}_T = \mathsf{E}$ expressed in percentage changes. More generally, one can show that optimal monopolistic price-setting behavior of the foreign exporters indeed consists of full exchange rate pass-through given constant price elasticities of the demand. The definition of the price index allows us to state percentage price level changes as $\mathsf{P} = \gamma\mathsf{P}_T = \gamma\mathsf{E}$. It is straightforward that for a more-open economy (larger γ) any given (percentage) exchange rate change E has a more pronounced impact on the overall domestic price level. Log-linearization of the money demand in equation (12) implies

$$\mathsf{M} - \mathsf{P} = \mathsf{P}_T - \mathsf{P} + \frac{\beta}{1-\beta}(\mathsf{P}_T - \bar{\mathsf{P}}_T) \,. \tag{16}$$

Let $\mathsf{M}^S = \bar{\mathsf{M}}^S$ denote the permanent home money supply expansion. For $\mathsf{P}_T = \bar{\mathsf{P}}_T$, the monetary equilibrium (16) implies long-run neutrality of money with $\bar{\mathsf{P}}_T = \bar{\mathsf{M}}^S$. Similarly, the short-run percentage tradeable price change and also the exchange rate change are proportional to the home money supply increase,

$$\mathsf{P}_T = \mathsf{M}^S = \mathsf{E} \,. \tag{17}$$

We recall that consumption smoothing implies a constant consumption of tradeables, hence $\mathsf{C}_T = 0$. But the nominally rigid nontradeable prices imply that the real price of nontradeables decreases and their demand increases. Log-linearizing equation (10) shows that this consumption expansion in nontradeables is proportional to the tradeable price increase,

$$\mathsf{C}_N = \mathsf{P}_T \,. \tag{18}$$

The consumption expansion is, of course, transitory. The nontradeable prices increase after one period and bring nontradeable consumption back to the original steady-state level.

We summarize our findings as follows: Money supply shocks have only a transitory effect on the real price of nontradeables, while the real price of tradeables is constant. The separability of the utility function in tradeable and nontradeable consumption implies that the real price change for nontradeables is without conse-

4. Nominally, rigidities for the nontradeables should be expected if the factor or labor market prices have sticky prices or wages. Optimal price setting for the nontradeable producers then implies under constant price elasticities that the constant price mark-up should be maintained.

quence for the tradeable good consumption. This latter feature of the model is certainly extreme and will not be obtained under a more general utility specification.[5] But we highlight two more robust features of our setup. First, exchange rate changes imply a proportional change in the import prices. We obtained this result by assuming a competitive market for tradeables. But the competitive market structure is not a necessary condition as highlighted by Hau (2000a, 2000b) and Obstfeld and Rogoff (2000). Optimal monopolistic price setting of the foreign exporters also implies full exchange rate pass-through if the demand elasticities are constant. More-open economies therefore have generally a more responsive aggregate price level. Second, this more flexible aggregate price level reduces the effect of the domestic money supply shock on the real household balances and therefore its scope for short-term consumption and real exchange rate changes. The higher aggregate price level flexibility implies lower real exchange rate volatility for a more-open economy.

1.4 Real Exchange Rate Volatility and Openness

We can formally derive the relationship between the real exchange rate and the openness of the economy. Using equation (10), we can define openness as the import (or identical) export share and express it as

$$Openness = \frac{P_T C_T}{P_N C_N + P_T C_T} = \gamma \,. \tag{19}$$

The percentage real exchange rate change follows as

$$\mathsf{E} - \mathsf{P} = \mathsf{P}_T - \mathsf{P} = (1 - \gamma)\mathsf{P}_T = (1 - Openness)\mathsf{M}^S \,. \tag{20}$$

The real exchange rate effect decreases in the openness of the economy. More-open economies should ceteris paribus have smaller changes in their real exchange rate.

The previous section explored the exchange rate effect for a one-time permanent money supply expansion in period 1. The model can be generalized to a dynamic setting with repeated monetary shocks. Percentage changes in the money supply may, for example, be randomly drawn from common distribution of mean zero and variance σ_M^2. We choose as a volatility measure the standard deviation of the real exchange rate and we obtain for expectations $\mathcal{E}$ the following openness-volatility linkage:

$$Vol = [\mathcal{E}(\mathsf{E} - \mathsf{P})^2]^{\frac{1}{2}} = (1 - Openness)\sigma_M \,. \tag{21}$$

Countries obviously differ in both their degree of openness and the volatility of their money supply shocks. However, if domestic supply shocks are largely exogenous to the openness of the economy and of similar magnitude, we expect to find a negative

5. Also, a more general utility framework will produce a current account effect with a (modest) permanent impact on consumption, output, and the real exchange rate.

correlation between openness and real exchange rate volatility in a cross-sectional comparison of real exchange rate volatility.

1.5 Unanticipated Supply Shocks

Real shocks can be shown to generate the same negative relationship between real exchange rate volatility and openness. We illustrate this for the parameter κ, which governs the marginal disutility of a household's nontradeable production. Let $\boldsymbol{\kappa}$ denote a permanent percentage increase in marginal disutility in period 1. Log-linearizing equation (11) implies

$$-\mathsf{P}_T - \mathsf{C}_T = \boldsymbol{\kappa} + \mathsf{y}_N \,. \tag{22}$$

The constant endowment of tradeables, $\bar{y}_T$, constant net foreign assets, and the consumption-smoothing motive imply again that the tradeable consumption is constant irrespective of the labor supply shock ($\mathsf{C}_T = 0$). Equation (10) can be log-linearized to

$$\mathsf{C}_N = \mathsf{P}_T \,, \tag{23}$$

given nominally rigid nontradeable prices ($\mathsf{P}_N = 0$). Market clearing for nontradeables ($\mathsf{C}_N = \mathsf{y}_N$) then determines the price change for tradeables as

$$\mathsf{P}_T = \frac{1}{2}\boldsymbol{\kappa} \,. \tag{24}$$

Tradeable prices are tied to the world price level P_T^*; hence, again $\mathsf{E} = \mathsf{P}_T$. The real exchange rate change follows as

$$\mathsf{E} - \mathsf{P} = \mathsf{P}_T - \mathsf{P} = (1 - \gamma)\mathsf{P}_T = (1 - \mathit{Openness})\,\frac{1}{2}\boldsymbol{\kappa} \,. \tag{25}$$

The unanticipated labor supply shock generates the same negative relationship between openness and real exchange rate volatility as the monetary shocks. Therefore, the following empirical section does not require a separate identification of monetary and real shocks.

2. EVIDENCE ON THE VOLATILITY-OPENNESS LINKAGE

This empirical section examines the cross-sectional pattern that links real exchange rate volatility to economic openness. First, we need to define a suitable measure of real exchange rate volatility. Previous studies have often focused on the bilateral dollar exchange rates. However, any specific bilateral real exchange rate might not be representative of the trade flows of some countries. We therefore focus on the effective real exchange rates that aggregate bilateral real exchange rates in one trade-weighted real exchange rate. A second choice concerns the frequency at which we measure real exchange rate changes. Since monetary models tend to perform better at low frequencies for which LOP deviations are presumably less pronounced, we measure real exchange rate changes over three-year periods.[6]

In accordance with the theoretical consideration in the previous section, we measure real exchange rate volatility as the standard deviation for the percentage changes of the effective real exchange rate *RE* over intervals of thirty-six months. Thus, for monthly data of country *i*, we define the standard deviation of the percentage effective real exchange rate change as

$$Vol_i = \left[\frac{1}{T} \sum_t \left(\frac{RE_{t+36,i} - RE_{t,i}}{RE_{t,i}} \right)^2 \right]^{\frac{1}{2}} . \tag{26}$$

The theoretical model suggests (ceteris paribus) a negative relationship between volatility and openness, where openness is measured by the import share of GDP. A simple cross-sectional regression can be formulated as

$$Vol_i = \alpha_0 + \alpha_1 Openness_i + \alpha_2 Z_i + u_i . \tag{27}$$

The variable Z_i represents different control variables. The model developed in the previous section implies for the coefficients that $\alpha_0 > 0$ and $\alpha_1 < 0$.

2.1 Data

The analysis is based on IMF data from the International Financial Statistics (IFS). For the measures of real exchange rate volatility we use effective real exchange rates (IFS code: xxx..RECZF...) for the period January 1980 to December 1998.[7] These trade-weighted effective real exchange rates are computed from consumption-based price indices and therefore correspond to our theoretical measure. We exclude all countries for which we do not have data going back to 1980. This eliminates all countries with central planning in the 1980s and also many developing countries. As a second selection criteria we require that a country's real exchange rate history does not have any extreme values. An extreme value is a change in the real exchange rate of more than 100 percent from one month to the next. We consider such extreme events related to political events outside the scope of an economic theory of real exchange rate determination and discard these countries. A total of forty-eight countries have complete data and pass the extreme value filter. The sample includes twenty-three OECD countries.[8]

6. Mark (1995) and Mark and Choi (1997) report, for example, that monetary models outperform the random walk hypothesis in out-of-sample tests only for low frequency movements above twelve months.

7. We can motivate this sample period. Many effective real exchange rate series start in January 1980. The December 1998 date marks the introduction of the euro.

8. The only OECD country without complete effective real exchange rate data is Turkey.

The openness of each country is calculated based on the arithmetic mean of the import and export share [IFS code: xxx90C.CZF (Exports), xxx98C.CZF (Imports), xxx99B.CZF (GDP)] and averaged over the nineteen-year sample period. Additional control variables include the per capita income in 1980 measured in U.S. dollars, the average number of revolutions and coups as a measure of political stability (Barro 1991), and measures of central bank independence (Cukierman 1992). Finally, I add a set of dummy variables for those countries which have exchange rate commitments in 1995 according to the International Financial Statistics (IMF 1996).[9] An additional dummy is introduced for the core members of the European Monetary System, which closely followed German monetary policy, namely, Austria, Belgium, Luxemburg and the Netherlands.

Figure 1 presents a scatterplot of volatility against the degree of openness for the full sample. Three countries with extremely high openness are Bahrain, Luxemburg,

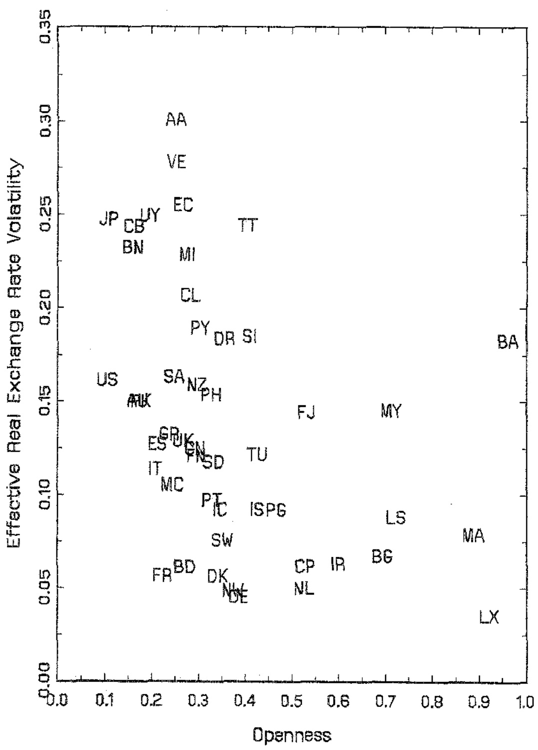

FIG. 1. Volatility of the Real Effective Exchange Rate for the Full Sample of Forty-eight Countries over the Period 1980 to 1998 as a Function of Their Openness Measured by the Average of the Import and Export Share. For the abbreviations of the OECD countries see Figure 2. The non-OECD countries are Algeria (AA), Bahrain (BA), Burundi (BN), Chile (CL), Colombia (CB), Cyprus (CP), Dominican Republic (DR), Ecuador (EC), Fiji (FJ), Israel (IS), Lesotho (LS), Malawi (MI), Malaysia (MY), Malta (MA), Morocco (MC), Pakistan (PK), Papua New Ginea (PG), Paraguay (PY), Philippines (PH), Saudi Arabia (SI), South Africa (SA), Trinidad and Tobago (TT), Tunisia (TU), Uruguay (UY), and Venezuela (VE).

9. OECD countries with exchange rate commitments are Austria, Belgium, Denmark, France, Germany, Ireland, Luxemburg, Netherlands, Portugal, and Spain.

and Malta. Figure 2 shows a separate scatterplot for the twenty-three OECD countries. Average volatility for the OECD countries is 0.099 compared to 0.137 in the full sample. Developing countries are on average characterized by much higher real exchange rate volatility. The following analysis reports results for the full sample and the OECD subsample separately.

2.2 *Basic Results*

Table 1 shows the regression results for the full sample of forty-eight countries. Column 1 reports the specification without any controls. The correlation between volatility and economic openness is statistically significant on a 1 percent level with an adjusted $\bar{R}^2$ of 0.150. Inclusion of the variable "log per capita GDP"in column 2 considerably improves the fit of the regression. Controlling for per capita GDP increases the adjusted $\bar{R}^2$ to 0.269, while the negative coefficient for openness remains significant at the 1 percent level. Per capita GDP represents an important control for a data set that combines countries of various development levels. Developing coun-

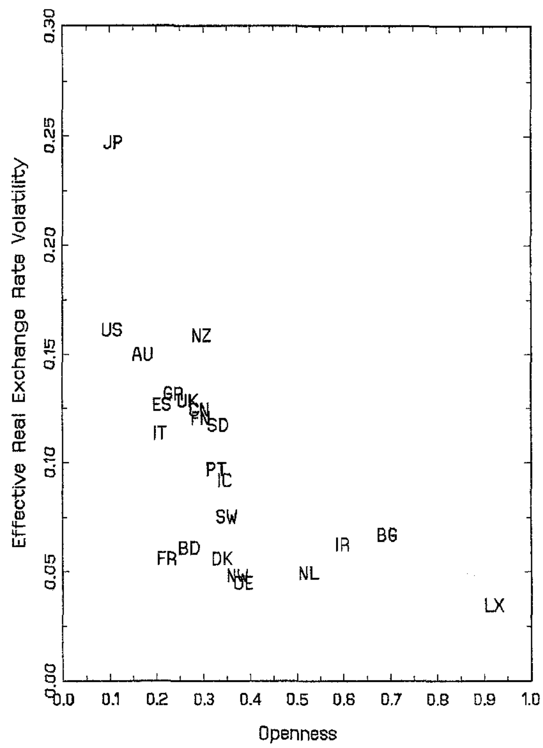

FIG. 2. Volatility of the Real Effective Exchange Rate for Twenty-three OECD Countries over the Period 1980 to 1998 as a Function of Openness Measured by the Average of Import and Export Share. The countries are Australia (AU), Austria (OE), Belgium (BG), Canada (CN), Denmark (DK), Finland (FN), France (FR), Germany (BD), Greece (GR), Iceland (IC), Ireland (IR), Italy (IT), Japan (JP), Luxemburg (LX), Netherlands (NL), New Zealand (NZ), Norway (NW), Portugal (PT), Spain (ES), Sweden (SD), Switzerland (SW), United Kingdom (UK), and United States (US).

TABLE 1

EFFECTIVE REAL EXCHANGE RATE VOLATILITY IN THE FULL SAMPLE

	1	2	3	4
Constant	0.189*** (0.019)	0.229*** (0.023)	0.241*** (0.028)	0.235*** (0.023)
Openness	−0.140*** (0.046)	−0.133*** (0.044)	−0.170*** (0.043)	−0.110** (0.046)
Log per capita GDP	–	−0.073*** (0.025)	−0.085*** (0.027)	−0.076*** (0.025)
Revolutions	–	–	0.038 (0.068)	–
Oil export dummy	–	–	0.141 (0.044)***	–
FX commitment	–	–	–	−0.025 (0.019)
$\bar{R}^2$	0.150	0.269	0.402	0.282
Sample size	48	48	48	48

NOTES: The effective real exchange rate volatility for the period January 1980 to December 1998 is measured as the standard deviation of three-year trade-weighted exchange rate changes for the full sample of forty-eight countries and regressed on a set of independent variables comprising the economic openness (average of import and export shares), the log per capita GDP in 1980 (in thousands of U.S. dollars), a political stability measure for the number of revolutions and coups (Barro 1991), a dummy variable for oil-exporting countries (Romer 1993) and a dummy variable for countries with exchange rate commitments in 1995 according to the International Financial Statistics (1996). We state the standard errors in parentheses and indicate significance on a 10 percent (*), 5 percent (**), and 1 percent level (***).

ties are plausibly exposed to larger real and monetary shocks than developed countries. Explanations range from their higher dependence on volatile world commodity prices to larger monetary shocks as a consequence of political instability. Column 3 includes the number of revolutions and coups in each country as measures of political stability and a dummy for oil-exporting countries. Oil exports are related to a more volatile real exchange rate, which is not surprising given the high volatility of the world oil price itself. But the point estimate for the openness effect remains significant on a 1 percent level even after inclusion of these control variables. In column 4 we test if the openness-volatility linkage is robust to a control for the exchange rate regimes. A dummy variable is used to distinguish all countries for which the IMF lists exchange rate commitments in 1995. Again the openness variable is significant, though on a weaker 5 percent significance level. Overall, the full sample supports the conjectured linkage between real exchange rate volatility and openness. But other factors such as development level or exposure to world commodity prices appear to be additional important volatility determinants. The full sample (plotted in Figure 1) constitutes a country group that presumably faces monetary and real shocks of very uneven magnitude. This sample heterogeneity in terms of the underlying shocks can only obscure the volatility-openness linkage in the sample cross section.

We therefore repeat these regressions for the subsample of twenty-three OECD countries. The OECD countries are more likely to be exposed to monetary and real shocks of similar magnitude. Table 2 shows the regression results. The basic regression without controls reported in column 1 now shows a surprisingly high adjusted $\bar{R}^2$ of 0.402. A high percentage of cross-sectional volatility is therefore explained by economic openness. Inclusion of the variable "log per capita GDP" in column 2 does

TABLE 2

EFFECTIVE REAL EXCHANGE RATE VOLATILITY IN THE OECD SAMPLE

	1	2	3	4	5
Constant	0.159***	0.176***	0.217***	0.192***	0.163***
	(0.017)	(0.038)	(0.041)	(0.106)	(0.041)
Openness	−0.176***	−0.175***	−0.178***	−0.113**	−0.145**
	(0.044)	(0.045)	(0.042)	(0.047)	(0.058)
Log per capita GDP	–	−0.020	−0.021	−0.039	−0.012
		(0.041)	(0.038)	(0.037)	(0.042)
CB independence	–	–	−0.105*	–	–
			(0.051)		
FX commitment	–	–	–	−0.045**	–
				(0.017)	
Core EMS	–	–	–	–	−0.022
					(0.026)
$\bar{R}^2$	0.371	0.402	0.380	0.466	0.518
Sample size	23	23	23	23	23

NOTES: The effective real exchange rate volatility for the period January 1980 to December 1998 is measured as the standard deviation of three-year trade-weighted exchange rate changes for the OECD subsample and regressed on a set of independent variables comprising the economic openness (average of import and export share), the log per capita GDP in 1980 (in thousands of U.S. dollars), a measure of central bank independence (Cukierman 1992) and a dummy variable for countries with exchange rate commitments in 1995 according to the International Financial Statistics (1996). The dummy variable "core EMS" marks the four countries Austria, Belgium, Luxemburg, and the Netherlands which closely followed German monetary policy. We state the standard errors in parentheses and indicate significance on a 10 percent (*), 5 percent (**) and 1 percent level (***).

not improve the fit, unlike for the full sample. Columns 3 and 4 consider an index of central bank independence and a dummy for exchange rate commitment as additional controls. Both measures are negatively correlated with exchange rate volatility. The exchange rate commitment dummy shows a statistical significance level above 2 percent.[10] But the point estimate for openness remains significant on a 3 percent level. A more restrictive measure of exchange rate commitment is captured by the variable "core EMS," which marks Austria, Belgium, Luxemburg, and the Netherlands. The negative point estimate for the openness effect is again robust to this alternative control. We conclude that the cross-sectional data and particularly the OECD subsample identifies openness as an important determinant of real exchange rate volatility. This evidence is therefore in accordance with the theoretical prediction of the structural monetary model presented in section 1.

3. ALTERNATIVE EXPLANATIONS

3.1 Endogenous Openness

The theoretical framework and the cross-sectional evidence in the previous section treat openness as an exogenous variable. However, exchange rate risk may have a feed-back effect on trade integration itself. This risk-trade channel is particularly plausible for low-frequency movements of the real exchange rate for which hedging

10. It is plausible that exchange rate commitments are made primarily by more-open economies.

strategies are either unavailable or very costly. A large empirical literature is devoted to this issue with contradictory results.[11]

The cross-sectional data allow us to address the issue of trade endogeneity with an instrumental variable (IV) approach. We simply reestimate the basic regression treating openness as endogenous and using a country's land area as an instrumental variable. The land area of a country is exogenous and not determined by exchange rate risk, but it is also a good proxy variable for openness with a correlations of −0.733 and −0.759 for the full sample and the OECD sample, respectively. If trade is endogenously depressed by exchange rate risk, then the IV point estimates should neglect this reverse causality and show a less pronounced negative relationship between volatility and openness than the OLS point estimates.

Table 3 shows the instrumental variable estimates for the full sample and the OECD sample. Column 1 reports the estimates for the full sample corresponding to the OLS specification of column 2 in Table 1. The IV point estimate for the openness coefficient is slightly lower (more negative) than the OLS counterpart. A similar result is obtained if we include exchange rate commitments in the regression specification. This is evidence against a reverse causality from exchange rate risk to lower levels of trade. We repeat the same regressions for the OECD sample. The IV point estimates for the openness coefficient are again negative and significant on a 1 percent level. Comparing column 3 in Table 3 to column 2 in Table 2, we find again a lower (more negative) IV coefficient than the corresponding OLS estimates. We therefore conclude that the openness-volatility linkage is not induced by a trade-risk channel.

TABLE 3

EFFECTIVE REAL EXCHANGE RATE VOLATILITY USING INSTRUMENTAL VARIABLES

Sample	1 Full	2 Full	3 OECD	4 OECD
Constant	0.233*** (0.026)	0.237*** (0.025)	0.186*** (0.038)	0.196*** (0.032)
Openness	−0.144** (0.056)	−0.118* (0.063)	−0.209*** (0.055)	−0.136** (0.061)
Log per capita GDP	−0.072*** (0.024)	−0.076*** (0.024)	−0.019 (0.038)	−0.036 (0.027)
FX commitment	–	−0.023 (0.019)	–	−0.040** (0.018)
$\bar{R}^2$	0.228	0.254	0.334	0.491
Sample size	48	48	23	23

NOTES: Land area of a country (in logs) is used as instrumental variable to estimate the relation between the effective real exchange rate volatility (for the period January 1980 to December 1998) and the economic openness (average of import and export share). Additional control variables are the log per capital GDP in 1980 (in thousands of U.S. dollars) and a dummy variable for countries with exchange rate commitments in 1995 according to the International Financial Statistics (1996). We state the standard errors in parentheses and indicate significance on a 10 percent (*), 5 percent (**) and 1 percent level (***)

11. Studies by Hooper and Kohlhagen (1978), Gotur (1985), and Asseery and Peel (1991), among others, do not find support for the trade-depressing effect of volatility. On the other hand, Cushman (1983, 1986, 1988), Akhtar and Hilton (1984), Kenen and Rodrik (1986), Thursby and Thursby (1987), De Grauwe (1988), Perée and Steinherr (1989), Koray and Lastrapes (1989), and Arize (1995, 1996) claim the opposite result.

3.2 Deviations from the Law of One Price

Deviations from the law of one price (LOP) are well documented in the economic literature. Furthermore, they appear related to geographic distance (Engel and Rogers 1996, 2001). But since economic openness strongly correlates with land area, country size, and distance, we might capture distance-related LOP deviations in our regressions rather than a fundamental macroeconomic volatility-openness linkage.

The role of LOP deviations in our regressions is presumably minimized by the choice of a three-year measurement frequency. Obstfeld and Rogoff (2000) examine the empirical case for incomplete exchange rate pass-through based on terms-of-trade data for the OECD countries. They find that the correlation of the nominal exchange rate E and a terms of trade index P^{im}/EP^{im*} composed of home currency prices of foreign imports (P^{im}) and foreign currency prices of domestically produced imports of the foreign country (P^{im*}) tends to be positive at a quarterly frequency contrary to what the incomplete pass-through scenario predicts. Over three-year periods, we expect LOP deviations to matter even less.

This leads to the interesting question of whether the volatility-openness linkage is more pronounced for high-frequency measures of real exchange rate volatility. Table 4 shows the standard regression for the OECD sample with volatility measured at frequencies of one month, three months, twelve months, and thirty-six months. Openness is found to be highly significant for all four measurement intervals. But the t-value on the openness coefficient decreases as we consider higher-frequency measures of real exchange rate volatility. A greater prominence of LOP deviations in high-frequency data in conjunction with distance-related LOP deviations should give a more pronounced short-run linkage. We therefore conclude that distance-related LOP deviations are not a very plausible explanation for the observed volatility-openness linkage.

3.3 Openness and the Intensity of Shocks

A third alternative explanation for the volatility-openness linkage might be that more-open economies do not experience relative monetary and real shocks of the same intensity as do closed economies. This corresponds to the hypothesis that the volatility σ_m of the monetary supply shock in equation (21) is itself a decreasing

TABLE 4

Effective Real Exchange Rate Volatility at Higher Frequencies

	1	2	3	4
Constant	0.0198***	0.0406***	0.0881***	0.1591***
	(0.0020)	(0.0041)	(0.0085)	(0.0172)
Openness	−0.0179***	−0.0380***	−0.0836***	−0.1761***
	(0.0053)	(0.0106)	(0.0219)	(0.0443)
$\bar{R}^2$	0.324	0.353	0.382	0.402
Sample size	23	23	23	23

Notes: The effective real exchange rate volatility is measured for the period January 1980 to December 1998 as the standard deviation of trade-weighted exchange rate changes for the OECD subsample at the following frequencies: one month (1), three months (2), twelve months (3), and thirty-six months (4). The independent variables are economic openness (average of import and export share) and the log of the per capita GDP in 1980 (in thousands of U.S. dollars). We state the standard errors in parentheses and indicate significance on a 10 percent (*), 5 percent (**), and 1 percent level (***).

function in openness. More-open economies might, for example, adopt institutions and policies that reduce the intensity of relative shocks. Examples of such policies include imitation of German monetary policy by the rather open economies of Austria, Belgium, and Netherlands. We marked these countries as core EMS countries. If such policy endogeneity is important for the entire OECD sample, we can potentially explain the observed relationship without recourse to the exchange rate magnification effect of nontradeables.

The literature distinguishes sterilized policies without relative price impact from nonsterilized policies, which change the relative inflation rate. Sterilized interventions have often been judged as largely ineffective.[12] Empirically, their impact seems to be small and difficult to detect (Edison 1993). Real exchange rate effects over a period of three years appears even more questionable. Our analysis therefore focuses on nonsterilized policies that stabilize the relative price level. If more-open economies engage in nonsterilized real exchange rate smoothing, we should expect their inflation rate to follow more closely those of their main trading partners. Inflation convergence can be measured directly by looking at the standard deviation of the effective relative price level over three-year periods. We calculate the relative price level P^*/P for twenty OECD countries based on a trade-weighted index P^* of the price levels of their major trading partners.[13] The volatility is measured (like for the real exchange rate) as the standard deviation of relative price changes over thirty-six-month intervals. If nonsterilized policies smooth real exchange rates for more-open economies, we should see a negative correlation between effective relative price volatility and openness.

Figure 3 plots the standard deviation of the relative price change for the twenty OECD countries. The correlation between this volatility measure and openness is negative, but insignificant as shown in Table 5, column 1. Controlling for "log per capita GDP," central bank independence and exchange rate regime choice does not indicate a statistically significant correlation either. Closer inspection of Figure 3 reveals that three core EMS countries, Austria, Belgium, and the Netherlands, show indeed low relative price volatility. Endogenous monetary policy aimed at tracking the German inflation rate is the plausible explanation for these three countries. We therefore use the core EMS dummy variable to mark these three countries and obtain the expected negative dummy coefficient (column 5). The correlation between openness and relative price volatility in the remaining sample is now positive rather than negative.[14] We conclude that a negative cross-sectional relationship between the magnitude of relative price shocks and economic openness is not supported by the data. The volatility-openness linkage for the real exchange rate is therefore not explained by the hypothesis that more-open countries simply have smaller relative shocks.

12. According to Obstfeld and Rogoff (1995, p. 76) "sterilized intervention operations are largely smoke and mirrors. Because they do not change the relative money supplies, sterilized interventions can have only modest effects, if any, on interest and exchange rates."

13. We used a matrix of trade weights available from the Bank of England for twenty OECD countries.

14. A positive coefficient is indeed predicted by the theory presented in section 1. Note that for an openness measure γ and a percentage money supply change M^S, the domestic price level change is $P = \gamma P_T = \gamma M^S$. More-open economies should therefore have more volatile relative price levels.

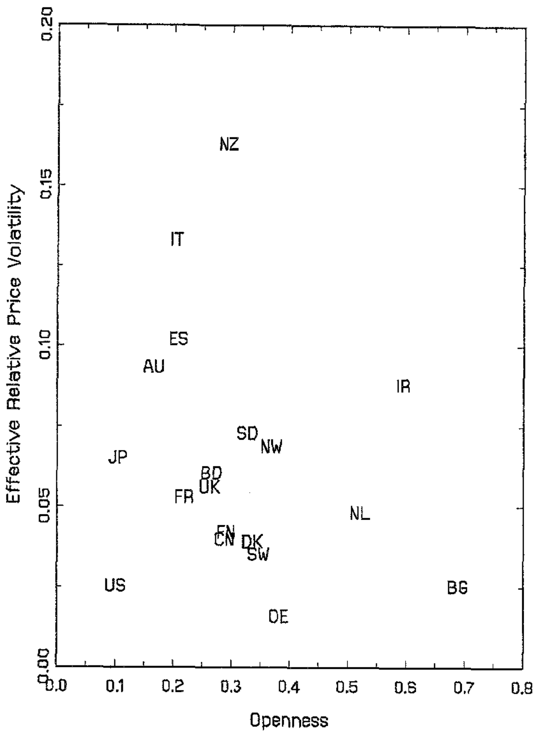

FIG. 3. Volatility of the Effective Relative Price Level for Twenty OECD Countries over the Period 1980 to 1998 as a Function of Openness. The abbreviations are explained in Figure 2.

TABLE 5

EFFECTIVE RELATIVE PRICE VOLATILITY

	(1)	(2)	(3)	(4)	(5)
Constant	0.091** (0.038)	0.346*** (0.057)	0.353*** (0.061)	0.068*** (0.013)	0.060*** (0.014)
Openness	−0.047 (0.109)	−0.117 (0.072)	−0.115 (0.074)	−0.002 (0.019)	0.002 (0.019)
Log per capita GDP	–	−0.027*** (0.005)	−0.027*** (0.006)	−0.005*** (0.001)	−0.004*** (0.001)
CB independence	–	–	−0.026 (0.069)	–	–
FX commitment	–	–	–	−0.005 (0.005)	–
Core EMS	–	–	–	–	−0.008 (0.007)
$\bar{R}^2$	−0.045	0.553	0.529	0.402	0.419
Sample size	20	20	20	20	20

NOTES: The volatility of the effective relative price levels are calculated based on trade weights between OECD countries and regressed on variables comprising the economic openness, the log per capita GDP in 1980, a measure of central bank independence (Cukierman, 1993), and a dummy variable for countries with exchange rate commitments in 1995 according to the International Financial Statistics (1996). The dummy variable "core EMS" marks the four countries Austria, Belgium, Luxemburg, and the Netherlands which closely followed German monetary policy. We indicate significance on a 10 percent (*), 5 percent (**), and 1 percent level (***).

4. CONCLUSION

This paper examines the role of nontradeables in the volatility of the real exchange rate. Macroeconomic theory suggests that real exchange rate volatility is negatively related to economic openness. Nontradeables increase the degree of aggregate price rigidity unlike tradeables which facilitate the relative price level adjustment under exchange rate pass-through. This argument remains valid even if exchange rate pass-through of tradeable prices is itself incomplete in the short run. Larger real exchange rate changes are therefore needed for a more-closed economy in order to achieve the same degree of relative price adjustment that accommodates the asymmetric monetary and real shocks. By contrast, more-open economies behave more like flexible price economies with (ceteris paribus) smaller real effects.

The empirical evidence on a cross section of forty-eight countries and particularly the OECD subsample is supportive of the conjectured openness-volatility linkage. The negative relationship between volatility and openness is robust to the inclusion of various control measures. We also discard an endogenous risk-trade channel, distance-related LOP deviations, or a systematic concentration of relative shocks in closed economies as alternative explanations for the observed relationship. Our results predict large real exchange rate volatility for the future euro/dollar rate or the euro/yen rate if the level of intercontinental trade remains on its presently modest level.

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