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The Effect of Real Exchange Rate Uncertainty on Exports: Empirical Evidence

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Unless very specific assumptions are made, theory alone cannot determine the sign of the relation between real exchange rate uncertainty and exports. On the one hand, convexity of the profit function with respect to prices implies that an increase in price uncertainty raises the expected returns in the export sector. On the other, potential asymmetries in the cost of adjusting factors of production (for example, investment irreversibility) and risk aversion tend to make the uncertainty-exports relation negative. This article examines these issues using a simple risk-aversion model. Export equations allowing for uncertainty are then estimated for six developing countries. Contrary to the ambiguity of the theory, the empirical relation is strongly negative. Our estimates indicate that a 5 percent increase in the annual standard deviation of the real exchange rate can reduce exports by 2 to 30 percent in the short run. These effects are substantially magnified in the long run.

More than half a decade after the onset of the debt crisis, many countries are still struggling to achieve a current account situation that is compatible with reduced external financing and a moderate but sustainable rate of output growth. Given the sudden decrease in the availability of external funds, most of the initial adjustments have involved drastic reductions in imports and investment but only marginal increases in exports.

A key element of a successful medium-term strategy of adjustment and growth is to move resources into the export sector. If the economy is close to full employment, the necessary reallocation of resources will require restrictive aggregate demand policies and a sustained real effective depreciation to make net exports more profitable (Fischer 1986, Khan 1987, Killick and others 1984).

Unfortunately, this type of policy usually entails sharp short-run recessions. Export incentives which have smaller costs are needed. The main purpose of this article is to show that real exchange rate uncertainty is one of these

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variables. Reducing exchange rate uncertainty may decrease the real devaluation required to improve the current account balance while avoiding a recession.

Theory alone cannot determine the sign of the relation between real exchange rate uncertainty and exports. Here we construct a simple two-period model that highlights some of the issues involved in this ambiguity. Given the simplicity of the model, there is a one-to-one relation between the effects of real exchange rate uncertainty on exports and its effects on investment.

Hartman (1972) and Abel (1983) showed that if one assumes perfect competition, convex and symmetric costs of adjusting capital, and risk neutrality, investment increases with price (real exchange rate) uncertainty. Their argument is a straight application of Jensen's inequality (that is, the fact that $E[f(x)] \geq f(Ex)$ if $f''(x) > 0$), since the profit function, as well as the marginal profitability of capital, are convex functions of prices (see Varian 1978). Thus, an increase in uncertainty raises the expected marginal profitability of capital, increasing investment (exports).

Hartman and Abel's result is robust to important departures from its basic assumptions as long as competition remains perfect (or close to it) and economic agents are risk-neutral (Caballero 1989). If the assumption of perfect competition is relaxed, however, asymmetric adjustment costs (for example, irreversibility of investment) can—although it need not—reverse the sign of the investment-uncertainty relation (Pindyck 1986, Bertola 1988, Caballero and Corbo 1988), even under risk neutrality. Alternatively, if firms are risk averse similar results will emerge. This is the assumption made in this article to motivate the uncertainty-exports relation and to provide the basis for the empirical section.

Several empirical studies have shown a negative relation between real exchange rate volatility and exports (see Behrman 1976 on Chile, Diaz-Alejandro 1976 on Colombia, Coes 1979 on Brazil, and Paredes 1986 on Peru). Given that the main motivation of this article is empirical, the model outlined below presents only the minimum elements needed to test the implications of real exchange rate variability on the level of exports.

One should be careful about the possible confusion between the direct role of uncertainty on exports and the effect of uncertainty through the expected value of the real exchange rate. The latter is referred to as the credibility problem. In making decisions in response to policy-induced improvements in export incentives, agents attempt to predict whether, for how long, and under what conditions such favorable policies will continue. Lack of credibility in the sustainability of the policy change will result in a smaller response of exports to real exchange rate changes, since higher uncertainty is perceived as a *lower* expected real exchange rate for any given currently high real exchange rate. In this article we are concerned with the effects of changes in uncertainty *given* the expected real exchange rate.

I. The Model

We assume a representative firm in the export sector facing the following demand curve:

$$X^d(t) = A_1(t)\left[\frac{P_x(t)}{P_w(t)}\right]^{-\eta}$$

and production function:

$$X(t) = A_2(t)N_t^{\alpha}K_t^{1-\alpha}$$

where X^d and X represent exports demanded and produced, P_x and P_w are the export price and world price indexes, η is the (absolute value) price-elasticity of demand, N and K are labor and capital used in production, α is the labor share of output, and $A_1(t)$ and $A_2(t)$ are arbitrary functions of time.

The real exchange rate, $V(t)$, and the real wage, $W(t)$, are defined as the nominal exchange rate and wages are deflated by the consumer price index (CPI). Both are exogenous to the firm.

From the above production function we can now define the maximized operating profits, $\pi(K,t)$, as follows:

$$\pi(K,t) \equiv \max_{N(t)} V(t)P_w(t)A_1(t)^{\frac{1}{\eta}}X(t)^{\mu}W(t)N(t)$$

where $\mu = (\eta - 1)/\eta$ represents an inverse index of monopoly power.

To make notation less burdensome, let us assume that only the real exchange rate is uncertain: there is no major increase in the complexity of solving the problem when there are multiple sources of uncertainty. We summarize all the remaining state variables as a deterministic function of time, $B(t)$, so:

$$\pi[K(t),\, t] = B(t)K(t)^{\theta_1}V(t)^{\theta_2} \tag{1}$$

where

$$\theta_1 = \frac{\mu(1-\alpha)}{1-\alpha\mu} < 1 \text{ and } \theta_2 = \frac{1}{1-\alpha\mu} > 1$$

It is also useful to write exports as a function of prices and the capital stock

$$X(t) = D(t)[\mu P(t)]^{\frac{\alpha}{1-\alpha}}K(t) \tag{2}$$

where $D(t)$ is just a function of time, and $P = P_x V(t)$.

Equation 2 shows that the variance of $V(t)$ may affect exports [given a period's real exchange rate, $V(t)$] only if there is some type of capital rigidity. If there is no capital rigidity, it is possible to optimize equation 1 with respect to $K(t)$, so that exports are only a function of time and the realization of the real exchange rate, $V(t)$.

To establish the real exchange rate process let us assume that the logarithm of $V(t)$ is an independently and identically distributed normal variable, with mean $-\sigma^2/2$ and variance σ^2. (Adding serial correlation is a trivial extension). The only purpose of the correction $\sigma^2/2$ is to allow the separation between the mean and the variance of log normally distributed variables. Thus an increase in σ keeps the expected value of $V(t)$ constant.

Assume that a firm has to purchase its capital one period before it is actually used. From equation (1) it can be seen that the marginal operating profit function, $\pi[K(t), t]$ is convex in the real exchange rate ($\theta_2 > 1$). Therefore, when capital is predetermined and agents are risk-neutral, investment is an increasing function of uncertainty, as shown by the models of Hartman and Abel. As a result, exports will be higher for every real exchange rate realization.

This apparent paradox can be explained by the fact that the firm loses less (with respect to the frictionless allocation) when the change in the real exchange rate is unfavorable than it does when it is favorable. When the change in the real exchange rate is unfavorable, the firm wishes to reduce production and has too much capital compared with its optimal value. Conversely, when change in the real exchange rate is favorable, the firm has a capital stock that is too low. Given the convexity of the profit function, the potential profits forgone due to an insufficient capital stock given a rise in the real exchange rate is higher than losses due to an underutilized capital stock when the real exchange rate falls, so a profit-maximizing firm would invest more, and thereby raise exports in the face of uncertainty.

Alternative Assumptions

The positive relation established in the Hartman and Abel models can be reversed by assuming that firms can adjust their capital stock in the second period (at a higher cost), or that firms are risk averse. In the first case, firms pay a premium in order to reduce the investment lag (for a full explication of this approach, see Caballero and Corbo 1988). This is similar to the model of irreversible investment. When capital is irreversible but not predetermined, it is always possible to invest in the second period, limiting the losses of having invested too little in the previous period to possible higher costs of "rushing" capital installation in the second period. At the same time, the losses of being caught with too much capital are the same as in the case in which capital is predetermined. Given this ability to adjust (at a higher cost), for a broad set of parameters uncertainty reduces rather than increases exports.

A second alternative that produces a negative effect of uncertainty on exports is the introduction of risk aversion. If the concavity of the utility function is large enough to offset the convexity of the profit function with respect to prices, as real exchange rate uncertainty rises, investment, and therefore exports, will be reduced. We modify our previous model to illustrate this mechanism.

Assume that the preferences of the firm's owners can be characterized by an intertemporally separable utility function with a constant coefficient of relative risk aversion, γ:

$$U(C(t)) = \frac{C(t)^{1-\gamma}}{1-\gamma}$$

where $C(t)$ denotes average consumption at time t.

The standard consumption-capital asset pricing model implies that if markets are complete, in equilibrium the price of a unit of capital, P^k—here assumed to be one—must equal the present value of its return, discounted by the marginal rate of substitution between today's and tomorrow's consumption:

$$(3) \qquad P_k = 1 = \beta E_t\left\{\left[\frac{C(t+1)}{C(t)}\right]^{-\gamma} \frac{\pi[K(t+1), t+1]}{K(t+1)}\right\}$$

where β is the subjective discount factor. Because this is a two-period model with a delivery lag, there is only one productive period, and due to the delivery lag, $K(t + 1)$ is known at time t.

Solving the stock of capital from equation 3 yields:

$$(4) \qquad K(t+1) = \left\{\beta B(t+1) E_t\left[\left[\frac{C(t+1)}{C(t)}\right]^{-\gamma} V(t+1)^{\theta_2}\right]\right\}^{\frac{1}{1-\theta_1}}$$

We further assume that the pair $(\ln[C(t + 1)/C(t)], \ln V(t + 1))$ is joint-normally distributed with mean $(\gamma/2, -\sigma^2/2)$, variance $(1,\sigma^2)$, and covariance $\rho\sigma$. The assumptions on the mean and variance of the rate of growth of consumption are only made for notational convenience and do not affect the main results. Given this assumption, there is an explicit relation between the capital stock and real exchange rate uncertainty:

$$(5) \qquad K(t+1) = \left\{\beta B(t+1) e^{[\theta_2(\theta_1-1)\sigma^2/2 - \gamma\theta_2\rho\sigma]}\right\}^{\frac{1}{1-\theta_1}}$$

Equation 5 shows that as long as there is a positive correlation between the rate of consumption growth and changes in the real exchange rate ($\rho > 0$), there is always a coefficient of relative risk aversion, γ, large enough to produce a negative relation between investment, and therefore exports, and real exchange rate uncertainty. It is also clear, nonetheless, that the effect of uncertainty is in general ambiguous, confirming the issue discussed. The empirical evidence shown in section II, however, is unambiguous, and strongly suggests that the exports-uncertainty relation is negative for the case of developing countries reviewed here.

II. Application of the Model

The Data

The World Bank's export price and volume data which we use here are described in depth in Moran and Park (1986; for advantages, limitations, and biases, see pp. 9–11). We used Paasche indexes of unit values as proxies for individual commodity prices. Volumes of exports are export values divided by their respective price indexes.

Quarterly data on consumer price indexes, nominal exchange rates, and the proxy for world demand (industrial-country real gross domestic product) were obtained from the *International Financial Statistics* of the International Monetary Fund (IMF). The real exchange rate is defined as the export price times the nominal exchange rate divided by the CPI.

To develop a measure of uncertainty, we first calculated quarterly standard deviation estimates of the real exchange rate. Each quarter's standard deviation was estimated using the real exchange rate of the current and previous three quarters. The annual uncertainty level was measured by averaging the standard deviation of the real exchange rate of the four quarters of each year. This is equivalent to the standard deviation estimates of a generalized autoregressive conditionally heteroskedastic model on the exchange rate equation with a very long and restricted moving average structure.[1]

Adaptation and Assumptions

Section I highlighted the relation between exports and the average level and variance of the real exchange rate. This section tests these relations on the time-series data of six developing countries: Chile, Colombia, Peru, Philippines, Thailand, and Turkey.[2]

The theoretical model was only designed to motivate the relation between exports and uncertainty. In order to preserve the main implications of the theoretical model and at the same time provide a simple (and feasible) export equation, we have made the following assumptions:

- The relation between the logarithm of capital and the standard deviation of the real exchange rate is approximately linear.
- Adjustment is slow and firms learn through experience, so lagged exports are included in the right-hand side to account for these delayed responses (see Caballero and Corbo 1987).
- The (inverse) index of monopoly power, μ, is linearly related to world demand.

The third assumption resulted from our earlier finding using the same six countries, that tying the relative price effect to the effect of world demand seemed to marginally improve the behavior of export equations (Caballero and Corbo 1986). This restriction is introduced by assuming that the inverse of the mark-up, μ, is linearly related to the level of world demand. The restriction is thoroughly tested, and the results without the world demand variable are also reported. All the fundamental results of the paper are shown to be robust to

1. Notice that if a generalized autoregressive conditionally heteroskedastic model in which the mean is dependent on the higher moments is applied directly to the export equation, there may not be a good approximation of uncertainty if the econometrician has less information than the firms themselves.

2. The Korean data showed a negative relation between uncertainty and the level of exports, but there were clear symptoms of strong specification error. We excluded these results to avoid distracting the reader with too many second-order arguments.

the omission of world demand from the export equations. Furthermore, the model with flexible mark-up seems to perform marginally better than the fixed mark-up model, supporting our specification. To be strictly rigorous, the flexible mark-up model should also affect the coefficient of the uncertainty term, but we have omitted this highly nonlinear complexity.

We can now proceed to expand the exports equation (2), as follows:

$$(6) \quad x(t) = c_0 + c_1[p(t) + wd(t)] + c_2\sigma(t) + c_3x(t-1) + c_4a(t) + \epsilon(t)$$

where x, p, and wd denote the logarithm of exports, the real exchange rate, and world demand, respectively, and $a(t)$ is just a function of time.

Before entering into the econometric issues involved in the estimation of equation 6, it is worth presenting the results of a simple regression in which no dynamic components are present (table 1). Although the parameters shown are almost surely inconsistent because important dynamic elements are omitted, they provide informal evidence in favor of our central hypothesis. With the exception of Colombia and Peru, uncertainty seems to have strong depressive effects on export levels.

Results and Implications

Estimates of equation 6 were made using ordinary least squares (OLS) and instrumental variables (IV) and are shown in table 2. The instruments chosen for the IV estimates are: a constant, the log of industrial countries' CPI divided by the developing countries' domestic CPI, the log of world demand, the standard deviation of the log of the real exchange rate, the log of lagged exports,

Table 1. *Static Approach Estimates of the Export Equation*

	Coefficient of:			
Country	*[p(t) + wd(t)]*	*σ(t)*	*Durbin-Watson statistic*	R^2
Chile	1.86 (0.29)	−5.45 (2.14)	1.14	0.64
Colombia	1.78 (0.40)	−0.84 (2.90)	1.18	0.64
Peru	3.97 (2.25)	−1.24 (5.57)	0.27	0.07
Philippines	3.27 (0.72)	−7.89 (3.43)	0.48	0.57
Thailand	4.00 (0.45)	−10.90 (3.24)	0.68	0.76
Turkey	3.91 (0.30)	−5.67 (2.22)	1.87	0.93

Note: $p(t)$ = logarithm of the real exchange rate; $wd(t)$ = logarithm of world demand; $\sigma(t)$ = standard deviation of the logarithm of the real exchange rate. The model was estimated using instrumental variables. Robust standard deviations are in parentheses (White 1980). A constant was also included in the regressions.

Source: Calculations based on World Bank data (see Moran and Park 1986).

Table 2. *Export Equations: The Flexible Mark-up Model*

Country and procedure	Short-run effect; coefficient of: $p(t) + wd(t)$	Short-run effect; coefficient of: $\sigma(t)$	Short-run effect; coefficient of: $x(t-1)$	Long-run effect; coefficient of: $p(t) + wd(t)$	Long-run effect; coefficient of: $\sigma(t)$	DW	R^2	LM	HM
Chile									
OLS	0.49	−0.83	0.77	2.15	−3.62	1.92	0.91		
	(0.20)	(1.50)	(0.09)	(0.54)	(5.95)				
IV	0.92	−1.99	0.61	2.38	−5.17	1.75	0.89	1.04	4.30
	(0.31)	(1.80)	(0.12)	(0.47)	(3.55)				
Colombia									
OLS	0.24	−0.87	0.74	0.92	−3.33	1.84	0.89		
	(0.23)	(1.21)	(0.11)	(0.63)	(4.49)				
IV	0.85	−0.45	0.50	1.70	−0.91	1.69	0.85	1.81	3.08
	(0.37)	(1.97)	(0.14)	(0.50)	(3.97)				
Peru									
OLS	0.62	−1.13	0.76	2.57	−4.70	1.47	0.99		
	(0.25)	(0.46)	(0.07)	(1.35)	(2.22)				
IV	0.87	−1.08	0.78	3.91	−4.86	1.74	0.99	3.24	1.22
	(0.28)	(0.52)	(0.08)	(1.88)	(2.89)				

Philippines									
OLS	0.52	−0.12	0.91	5.61	−1.35	2.07	0.98		
	(0.27)	(1.66)	(0.04)	(1.56)	(18.10)				
IV	0.79	−0.30	0.86	5.60	−2.16	1.86	0.97	2.63	3.25
	(0.35)	(1.35)	(0.04)	(1.30)	(9.74)				
Thailand									
OLS	0.29	−4.11	0.90	2.84	−40.20	2.27	0.98		
	(0.20)	(0.89)	(0.05)	(1.38)	(18.61)				
IV	1.20	−5.93	0.70	3.99	−19.70	1.56	0.96	0.09	3.81
	(0.51)	(1.22)	(0.11)	(0.58)	(5.50)				
Turkey									
OLS	3.71	−5.83	—	—	—	1.76	0.93	0.00	6.32
	(0.29)	(2.25)							
IV	3.92	−5.66	—	—	—	1.88	0.93		
	(0.29)	(2.22)							

— Not available.

Note: $p(t)$ = log of the real exchange rate; $wd(t)$ = log of world demand; $\sigma(t)$ = standard deviation of the log of the real exchange rate; $x(t-1)$ = lagged exports; DW = Durbin-Watson statistic; LM = Lagrange multiplier test statistic; HM = Hausman specification test statistic; OLS = ordinary least squares estimation procedure; IV = instrumental variable estimation procedure.

Standard deviations are in parentheses. A constant, included in the regressions, is not shown here. In the case of Peru, we also included log t as a regressor. The estimated value of its coefficient with the respective standard errors in parentheses were 2.19 (0.67) for the OLS estimation and 1.95 (0.76) for the IV estimation.

Source: Calculations based on World Bank data (see Moran and Park 1986).

and the log of time. In both approaches, time and lagged exports are excluded when they are not significant.

It is apparent from this table that most of the models using the assumption of a predetermined real exchange rate (and therefore using OLS procedures) are likely to underestimate seriously the price-elasticity of exports. Once simultaneity is corrected for, the price-elasticity estimates increase for every country in our sample, and in several countries the price-elasticity estimates more than double. The evidence of OLS specification error is also supported by Hausman's specification test (Hausman 1978), presented in the last column of table 2. This statistic is above the critical level for five out of our six countries (the exception is Peru) at the 10 percent significance level. Moreover, the Lagrange multiplier test shows that the overidentifying restrictions imposed in the IV procedure cannot be rejected at any reasonable significance level (with the exception of Peru, where they are rejected at the 10 percent significance level).

When the IV procedure is used, the estimates of the effects of real exchange rate uncertainty on exports are always negative, although this can be validated statistically for only half the countries studied (Peru, Thailand, and Turkey). However, the probability that the coefficient of uncertainty comes out negative in *every* country as a result of sampling error is considerably less than the significance level of each individual test. Thus there is clear evidence of the depressing effect of real exchange rate uncertainty on export levels.

Not only does real exchange rate uncertainty depress exports, but it does so by a substantial amount, even in the short run. For example, according to our estimates, an increase of 5 percentage points in real exchange rate variability in the Chilean economy leads to a total decline in exports of about 10 percent. And this is a very conservative example: the swings in Chilean exchange rate regimes suggests changes that far exceed the 5 percent change used in this example. The example is far more dramatic in the case of Thailand and Turkey, where similar increases in variability would lead to a 30 percent decline in exports.

With the exception of Turkey, where adjustment seems to be very fast, all the effects previously described are magnified in the long run. A 5 percent increase in real exchange rate uncertainty would lead to a long-run decline in exports of 25 percent in the case of Chile, and it would eliminate Thailand's exports.

Constant Elasticity of Demand

At the outset of this section, we argued that allowing for the index of monopoly power, μ, to depend linearly on world demand provided an improvement in the fit of export equations in some countries. In this subsection we show that this is indeed the case. More important, however, we show that none of our main qualitative results relies on this assumption.

The estimates in table 3 provide clear evidence of this. Both the downward bias of OLS estimates and the depressing effects of real exchange rate uncertainty are fully carried over to the perfectly competitive case in which firms are price-takers. If anything, the results are even more supportive of our hypothesis under this constant mark-up assumption.

Table 4 shows that there is marginal evidence in support of the imperfect markets assumption and in favor of the flexible μ specification. The table presents the results of a *J* test modified to account for the correlation between regressors and disturbances (see MacMinnon, White, and Davidson 1982). The results of this test are inconclusive for Chile, Peru, the Philippines, and Thailand but favor the flexible demand-elasticity specification in the cases of Colombia and Turkey. This result is backed, in the case of Colombia, by a high χ^2 statistic for the Lagrange multiplier test in the constant demand-elasticity specification.

Unanticipated Exchange Rate Changes

The derivations in the theoretical section suggested that the expected and the realized exchange rate would affect exports separately and thus that both should enter the right-hand side of the exports equation (6). The expected rate should enter through its effect on the capital stock, and the actual through its effect on hiring and firing of flexible factors. In the empirical section we disregarded this difference and proceeded using just the realized exchange rate, $p(t)$. If all the information on which production decisions depend were available in the first period, the unanticipated change would not be relevant. Similarly, in our model if all the instruments belonged to the information set at $(t-1)$, the estimation procedure would be unable to distinguish between the realized and expected exchange rates. However, some of the instruments used correspond to contemporaneous variables and hence do not belong to the information set at $(t-1)$, when capital stock decisions are made. In spite of this, the results of the Lagrange multiplier test presented in table 2 suggest that this did not imply substantial biases. In other words, the covariance between the "news" component of the real exchange rate and instruments seems to be very small.

That this is so is also suggested by the approach shown in table 5. For each country, the first row reproduces the IV results of table 2. The second row shows an alternative set of IV estimates in which the contemporaneous instruments are lagged one period. The results are very similar, so that it seems safe to conclude that our results are robust despite the fact that we do not account separately for the anticipated and unanticipated elements of exchange rate variation.

III. Conclusion

In this article we briefly reviewed the theoretical ambiguity of the relation between uncertainty and exports. We then proceeded to assess the sign of this relation from a set of empirical estimations of export functions. In contrast with the theoretical ambiguity, the empirical results of the estimation of the

Table 3. *Export Equations: The Constant Mark-up Model*

Country and procedure	Short-run effect; coefficient of: $p(t)$	$\sigma(t)$	$x(t-1)$	Long-run effect; coefficient of: $p(t)$	$\sigma(t)$	DW	R^2	LM	HM
Chile									
OLS	0.52	−0.48	0.87	3.97	−3.66	1.99	0.91		
	(0.23)	(1.48)	(0.07)	(1.76)	(10.80)				
IV	1.16	−1.72	0.77	4.94	−7.36	1.78	0.87	0.02	4.40
	(0.41)	(1.80)	(0.09)	(1.75)	(6.17)				
Colombia									
OLS	0.07	−1.03	0.82	0.42	−5.76	1.78	0.88		
	(0.25)	(1.14)	(0.08)	(1.29)	(5.78)				
IV	0.52	−0.98	0.75	2.08	−3.97	1.83	0.86	5.10	1.94
	(0.34)	(1.61)	(0.09)	(1.21)	(6.11)				
Peru									
OLS	0.64	−1.11	0.75	2.58	−4.46	1.46	0.99		
	(0.27)	(0.46)	(0.07)	(1.37)	(2.09)				
IV	0.99	−1.03	0.77	4.31	−4.50	1.82	0.99	3.07	1.43
	(0.31)	(0.54)	(0.08)	(2.01)	(2.80)				

Philippines									
OLS	0.56	0.19	0.99	36.10	12.50	2.07	0.97		
	(0.33)	(1.88)	(45.40)	(55.60)	(124.50)				
IV	1.17	−0.62	0.97	35.30	−18.70	1.76	0.96	1.76	3.58
	(0.56)	(1.43)	(0.03)	(37.90)	(43.50)				
Thailand									
OLS	0.17	−3.94	0.95	3.57	−83.7	2.28	0.98		
	(0.24)	(1.06)	(0.04)	(5.34)	(72.20)				
IV	1.64	−7.54	0.88	13.90	−63.60	1.53	0.95	0.14	3.10
	(0.86)	(2.19)	(0.06)	(6.07)	(26.10)				
Turkey									
OLS	6.05	−9.22	—	—	—	1.56	0.81		
	(0.73)	(3.53)							
IV	7.27	−9.29	—	—	—	1.93	0.78	2.34	6.56
	(0.91)	(3.45)							

— Not available.

Note: $p(t)$ = log of the real exchange rate; $\sigma(t)$ = standard deviation of the log of the real exchange rate; $x(t-1)$ = lagged exports; DW = Durbin-Watson statistic; LM = Lagrange multiplier test statistic; HM = Hausman specification test statistic; OLS = ordinary least squares estimation procedure; IV = instrumental variable estimation procedure.

Standard deviations are in parentheses. A constant, included in the regressions, is not shown here. In the case of Peru, we also included log t as a regressor. The estimated value of its coefficient with the respective standard errors in parentheses were 2.94 (0.69) for the OLS estimation and 3.03 (0.84) for the IV estimation.

Source: Calculations based on World Bank data (see Moran and Park 1986).

Table 4. *Evidence for the Perfectly Competitive and Imperfectly Competitive Cases*
(t-statistics)

Country	*Constant mark-up*	*Variable mark-up*	*Conclusion*
Chile	−0.13	1.19	Inconclusive
Colombia	2.00	−1.29	Flexible mark-up
Peru	0.04	−0.04	Inconclusive
Philippines	−0.40	0.43	Inconclusive
Thailand	0.55	−0.37	Inconclusive
Turkey	2.41	0.04	Flexible mark-up

Source: Calculations based on World Bank data (see Moran and Park 1986).

model based on data from Chile, Colombia, Peru, the Philippines, Thailand, and Turkey are unambiguous in showing a clear and strong negative effect of real exchange rate uncertainty on exports. In the empirical investigation, simultaneity was carefully treated and shown to have a substantial effect on the estimates of the price-elasticity of exports. Also, the results were shown to be robust to a variety of possible sources of specification error. The point estimates obtained indicated that increases as small as 5 percentage points in the annual standard deviation of the real exchange rates can lead to a short-run decline in exports of 2.5 (Colombia) to 30 percent (Thailand and Turkey). These effects are substantially magnified in the long run.

The results suggest an option for governments that want to increase export supply without exacerbating the recessionary effect of sectoral reallocation due to exchange rate depreciation and restrictions on aggregate demand. This study indicates that such reallocation policies could be accompanied by a program to stabilize the real exchange rate and to make its policy-induced changes more predictable. The consequent decline in real exchange rate uncertainty would be expected to increase exports and thus minimize the length and costs of the adjustment period.

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Table 5. *Export Equations: Contemporaneous versus Lagged Information Instruments*

Country and Instruments	*Coefficient of:* $p(t) + wd(t)$	$\sigma(t)$	$x(t-1)$
Chile			
Contemporaneous	0.92	−1.99	0.61
	(0.31)	(1.80)	(0.12)
Lagged	0.94	−2.24	0.60
	(0.28)	(1.76)	(0.11)
Colombia			
Contemporaneous	0.85	−0.45	0.50
	(0.37)	(1.97)	(0.14)
Lagged	0.86	−0.58	0.49
	(0.37)	(1.91)	(0.14)
Peru			
Contemporaneous	0.87	−1.08	0.78
	(0.28)	(0.52)	(0.08)
Lagged	0.85	−1.19	0.78
	(0.26)	(0.50)	(0.08)
Philippines			
Contemporaneous	0.79	−0.30	0.86
	(0.35)	(1.35)	(0.04)
Lagged	0.69	−0.64	0.87
	(0.30)	(1.47)	(0.05)
Thailand			
Contemporaneous	1.20	−5.93	0.70
	(0.51)	(1.22)	(0.11)
Lagged	1.18	−5.60	0.70
	(0.48)	(1.13)	(0.10)
Turkey			
Contemporaneous	3.92	−5.66	—
	(0.29)	(2.22)	
Lagged	3.93	−5.11	—
	(0.30)	(2.71)	

— Not available.

Note: $p(t)$ = log of the real exchange rate in period (t); $wd(t)$ = log of world demand; $\sigma(t)$ = standard deviation of the log of the real exchange rate; $x(t-1)$ = lagged exports. In the case of Peru, we also included log t as a regressor. The estimated coefficient was 1.95 (with a standard error of 0.76) for both cases. Standard deviations are in parentheses.

Source: Calculations based on World Bank data (see Moran and Park 1986).

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