ON THE EFFECTS OF EXCHANGE RATE UNCERTAINTY ON DOMESTIC PRODUCTION: A DYNAMIC PANEL APPROACH.

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1 Introduction

The Asian currency crisis fueled anew the debate over the use of devaluations as an important component of economic adjustment and stabilization programs. It is widely accepted that devaluations improve a country's balance of payments position, and accumulate more foreign exchange reserves, however, the issue of how they affect aggregate output still remains subject to disagreement.

The orthodox approach that has dominated the textbooks contends that devaluation enhances competitiveness, increases exports and bends demand toward domestically produced goods, thus expanding the production of tradables. However, frequent output declines in the aftermath of devaluations hinted that the benign relative price adjustment caused by devaluations could bring about a recession. Most of the early models in the literature focused on the effects of devaluations on the demand side. Studies by Diaz-Alejandro (1963), Krugman and Taylor (1978), Barbone and Rivera-Batiz (1987) are some of the seminal papers of the contractionary devaluation literature. Later studies deal with supply side factors that make devaluations contractionary: Bruno (1979), Gylfason and Schmid (1983), van Wijnbergen (1986), Agenor (1991), Gylfason and Radetzki (1991), and Taye (1999).

The most important reasons for a devaluation to trigger an **aggregate demand** contraction include: (1) A redistribution of income towards those with high marginal propensity to save [Diaz-Alejandro (1963); Cooper (1971); and Krugman and Taylor (1978)]; (2) A fall in investment [Branson (1986); Buffie (1986b); and van Wijnbergen (1986)]; (3) An increased debt service burden [Cooper (1971); Gylfason and Risager (1984); and van Wijnbergen (1986)]; (4) A reduction in real wealth [Bruno (1979); Gylfason and Schmid (1983); Hanson (1983); and Gylfason and Radetzki (1991)]; (5) A low government marginal propensity to spend out of tax revenue [Krugman and Taylor (1978)]; (6) Real income declines under an initial trade deficit [Cooper (1971); and Krugman and Taylor (1978)]; (7) Increased interest rates [Bruno (1979); and van Wijnbergen (1986)]; (8) Increased foreign profits [Barbone and Rivera-Batiz (1987)].

On the other hand, **aggregate supply** may suffer after a devaluation because of these three factors: (1) More expensive imported production inputs [Bruno (1979); Gylfason and Schmid (1983); Hanson (1983); Gylfason and Risager (1984); Islam (1984); Gylfason and Radetzki (1985); Branson (1986); Solimano (1986); and Wijnbergen (1986)]; (2) Wage indexation programs [Hanson (1983); Gylfason and Risager (1984), Islam (1984); Gylfason and Radetzki (1985); Branson (1986); Edwards (1986b). Solimano (1986); and Wijnbergen (1986)]; (3) Costlier working capital [Bruno (1979); and Wijnbergen (1986)].

The existing literature is dominated by macro-simulation studies. Econometric studies represent a smaller part. To mention some of the most cited empirical articles, Sheehey (1986), Morley (1992), Rogers and Wang (1995), Santaella and

Vela (1996), Copelman and Wemer (1996), Kamin and Rogers (1997), Rodriguez and Diaz (1995), Hoffmaister and Vegh (1996) find the contractionary devaluation hypothesis plausible. Other empirical studies by Edwards (1986, 1989b), Kamin and Klau (1998), Nunnenkamp and Schweickert (1990), Agenor (1991), and Bahmani-Oskooee and Rhee (1997) obtain less straightforward results. The conclusions of some of these latter studies are susceptible to the choice of period, type of devaluation, short- and long-run effects, econometric techniques, and country selections.[1]

Since we utilize a panel data approach in our estimation, we pay particular attention to existing panel data studies. Edwards (1986) is one of the most important early econometric studies. He uses least squares dummy variable model of real output, based on data for 12 developing countries for the period 1965-80, and finds that devaluations appear to be neutral in the long run. The first year sees an output reduction, but the second offsets it completely. Khan (1988), including policy variables and terms of trade disturbances in a panel framework covering 67 countries over the period 1973-1986, finds the exchange rate variable insignificant. Another study by Edwards (1989b) estimates a reduced form equation with panel data on 12 developing countries for the period 1965-84. Again, Edwards finds devaluations contractionary in the short-run, but they remain neutral for the long run only in two out of seven regressions. The coefficients of the lagged exchange rate turn out to be insignificant in all seven

[1] For a recent review of the literature see Bahmani-Oskooee and Miteza (2003).

regressions. Kamin and Klau (1998) examine a dataset of pooled data from 27 countries while including external shocks and considering regional grouping. They do not find support for the contractionary devaluation hypothesis in the long run. Moreover, they conclude that there is no evidence that makes contractionary devaluations only a developing country phenomenon. Nunnenkamp and Schweickert (1990) do pooled time-series cross-section analysis on 48 countries over the period 1982-1987. They conclude that there is no evidence to expect devaluations to hurt growth rates. They find that exporters of manufactures are more susceptible to contractionary devaluations in the short run, while expansionary devaluations are more likely in heavy exporters of agricultural products.

Agenor (1991) considers a pooled sample of 23 developing countries, and analyses the deviation of actual from expected changes in real exchange rates, foreign income, the money supply, and government spending. He concludes that unexpected real exchange rate depreciation is expansionary, but anticipated real depreciations are irreversibly contractionary.

Existing panel studies suffer from some shortfalls. First, regressions which include the real exchange rate among other variables (Edwards, 1986) create an endogeneity problem inasmuch as the real exchange rate is a function of the price of home goods. Second, the use of current and lagged terms for the exchange rate in the literature (Edwards, 1986; Agenor, 1991) in order to capture the immediate effects and the medium-run effects of devaluation on output can be problematic. Since these terms tend to have a high correlation, the size and sign of their coefficients becomes quite unreliable. Third, a negative correlation may arise between exchange rate changes and real output from a selectivity bias if devaluations are undertaken in periods of low growth. To allow for this, some authors bring in the lagged value of output on the right hand side of the regression. However, this can backfire if the estimation technique does not allow for the necessary dynamics. Finally, some reservations about period selection in some panel studies. The purpose of devaluations has changed over time, from opening up the economy to foreign trade, to overcoming balance of payments problems caused largely by the oil crisis, to reacting to an international debt crisis. Hence controlling for numerous environment shifts becomes technically complex when selecting a relatively long time span for the data set.

This study examines the effect of devaluation on aggregate output. It makes a new contribution to the literature in two ways: First, we attempt to bifurcate the effects of devaluation from those of exchange rate volatility. We expand the traditional formulations of this problem by including an additional factor – namely, *exchange rate volatility*. Exchange rate volatility has been extensively analyzed mainly in the context of its effect on trade volumes. High exchange rate volatility may cause the coefficient on the exchange rate to undeservedly capture the growth-reducing or expanding effect of exchange rate volatility. Isolating the impact of exchange rate volatility on output is potentially an informative exercise that could serve policy analysis.

Second, this paper improves upon several technical approaches of the existing empirical literature by introducing dynamics into the system through the adoption of a dynamic panel model. Furthermore, the use of a generalized method of moments (GMM) estimator allows for endogeneity in the regressors.

2 MODEL AND VARIABLE SELECTION

We use an output model by Edwards (1989b) in the following reduced form equation:

$$\log y = \text{constant} + a_1 \log E + a_2 \log M + a_3 \log G + u$$

where y is real GDP, E is the nominal exchange rate, M is a measure of the money stock, and G represents government demand for nontradables. Therefore a_1 measures the elasticity of real GDP with respect to the nominal exchange rate, and is our main object of estimation; a_2 represents the elasticity of real GDP with respect to changes in the money stock, and is expected to have a positive sign; lastly, a_3 captures the effect of fiscal policy on output and is also expected to carry a positive sign.

As mentioned above, the ultimate objective of our empirical estimation is the computation of the exchange rate coefficients. Since the exchange rate has been defined as units of foreign currency per unit of domestic currency, a negative coefficient for the exchange rate would imply that devaluations are expansionary. The inclusion of lags, however, may make it possible to distinguish between the short run and the long run effects of the exchange rate if any.

Recognizing the significance of the variables included in the above equations, as well as the merits of the literature that make use of them in one form or another, we expand on these formulations to include *exchange rate volatility*. While this variable has been extensively analyzed in the context of its effect on trade volumes, it has not been a part of the model specifications in the literature on contractionary devaluations.

We believe that high exchange rate volatility may have an impact on output because it adds to the uncertainty of the economic environment in the very same way inflation volatility does. If exchange rate volatility does play a role in trade volumes and output determination, then in some of the empirical studies the coefficient on the exchange rate may have in fact undeservedly captured the effect of exchange rate volatility.

We therefore include a measure of exchange rate volatility in the reduced form equation as follows:

$$\log y = \text{constant} + a_1 \log E + a_2 \log M + a_3 \log G + a_4 V + u \qquad (2.1)$$

The measure of exchange rate volatility (V) is generated by computing the standard deviation of the monthly percentage changes in the nominal exchange rate over a period of 12 months.

In the process of estimating this enhanced reduced form equation we hope to bring together the literature on the impact of exchange rate volatility on trade with the contractionary devaluation literature.

3 METHODOLOGY AND ESTIMATION PROCEDURE

We have chosen a panel framework to estimate the effects of devaluation on output mainly because it can control for heterogeneity in individual behavior, offers more variation, less collinearity among regressors, and more efficient estimators. Moreover panel models intrinsically present less measurement error problems as well as a mitigated omitted variable bias.

A number of pooled time series cross-section studies in the literature of contractionary devaluations have applied simple pooled OLS techniques that might well suffer from ignoring the different intercepts as well as a biased slope.

In estimating the reduced form equation (2.1), the explanatory variables are nominal effective exchange rates (NEER), broad money (M), real government spending, and the 12-month exchange rate volatility (V).

The negative correlation between changes in the exchange rate and real output in the year of the devaluation may result from a selectivity bias if devaluations are undertaken in periods of low growth. To avoid this, some studies introduce the lagged value of output (the dependent variable) on the right hand side of the regression. While a valid strategy, this may prove damaging if the estimation technique does not take into account that this represents a dynamic model that has to be estimated as such.

Consider the following model:

$$\log y_{it} = \beta_{1i} + \sum_{s=0}^{K} \beta_{2s} \log e_{i,t-s} + \sum_{s=0}^{K} \beta_{3s} \Delta \log M_{i,t-s} + \sum_{s=0}^{K} \beta_{4-s} \Delta \log G_{i,t-s} + \beta_5 y_{i,t-1} + \varepsilon_{it}$$

Although the panel framework is quite versatile in the estimation of dynamic models like the above, use of the ordinary fixed or random effects estimators is problematic. The lagged dependent variable included on the right hand side of the above equation will be correlated with the error term, even under the assumption of no autocorrelation in the ε_{it}. The application of fixed effects on a dynamic model that includes lagged values of the dependent variable among its regressors is bound to produce biased estimates if the panel time dimension is small (Nickell, 1981). According to Judson and Owen (1996), "… macroeconomists should not dismiss the least squares dummy variable bias as insignificant. Even with a time dimension as large as 30, we find that the bias may be equal to as much as 20% of the true value of the coefficient of interest."

While the choice of a fixed effects over a random effects model in macroeconomics is well justified by the fact that omitted variables, captured by the individual effect, are likely to be correlated with the explanatory variables, still it often is inadequate. In point of fact, macroeconomists often face the necessity of including lagged values of the dependent variable as regressors in the hope to better comprehend the dynamics of adjustment. However, since the lagged dependent variable is correlated with the error term the least squares dummy variable estimator (fixed effects) will be biased and inconsistent even when the error terms are not serially correlated. This becomes more of a concern for a short period of time (T) (Nickell (1981)). The random effects GLS estimator for a dynamic panel suffers no less from the same regressor-error term correlation problem. Anderson and Hsiao (1981) resort to first differencing the model so as to cancel away the individual effect in order to apply an instrumental variable estimation method. Valid instruments would constitute that subset of the terms that are correlated with the regressors but orthogonal to the error term. As indicated by Ahn and Schmidt (1993), since this instrumental variable technique does not allow for the differenced structure of the disturbances, nor it does exploit all the available moment conditions, it yields consistent but potentially inefficient parameter estimates. Furthermore the choice of differences is inferior to the choice of levels as instruments, according to Arellano (1989), as it yields large parameter variances (Arellano and Bond, 1991).

The general form of the dynamic panel model that we are concerned with here is that of a single equation with individual effects as in Arellano and Bond (1998).

$$y_{it} = \sum_{k=1}^{p} \alpha_k y_{i(t-k)} + \beta'(L) x_{it} + \lambda_t + \eta_i + \upsilon_{it}$$
$$(t = q+1, \ldots, T_i; i = 1, \ldots, N)$$

The above equation includes time and individual specific effects, denoted respectively by λ_t and η_i. The number of observations for the i^{th} individual is T_i, and the maximum lag length is q. In order for the model to be identified, a number of restrictions have to be made about the regressors and the error term υ_{it}.

If the error term was initially autoregressive, the assumption is that the model is transformed in such a way that the slope coefficients α's and β's satisfy some set of common factor restrictions, which implies that the model allows only for serially uncorrelated (moving average) errors. The υ_{it} errors are assumed to be independently distributed across individuals with zero mean and heteroskedasticity across individuals and time. The explanatory variables may be endogenous, predetermined or strictly exogenous with respect to υ_{it}. Every individual will have (Ti-q) equations, which may also be written as:

$$y_i = W_i\delta + \iota_i\eta_i + \upsilon_i$$

where the aforementioned coefficients α_K, β's and λ's are conveniently included in the parameter δ, while W_i is a matrix that includes the lagged dependent variables, the explanatory variables and the time dummies. The individual effects are premultiplied with a $(T_i - q) \times 1$ vector of ones ι_i. The general form of the linear GMM estimator that we employ in this study is:

$$\hat{\delta}_{GMM} = \left[\left(\sum_i W_i^{*'} Z_i\right) A_N \left(\sum_i Z_i' W_i^*\right)\right]^{-1} \left(\sum_i W_i^{*'} Z_i\right) A_N \left(\sum_i Z_i' y_i^*\right)$$

where

$$A_N = \left(\frac{1}{N}\sum_i Z_i' H_i Z_i\right)^{-1}$$

where Z_i is a matrix of instrumental variables and W_i^* and y_i^* denote a first difference transformation of W_i and y_i. The purpose of this transformation is to do away with the individual specific effects η_i, which are likely to be correlated with the instruments. The generalized method of moments (GMM) produces efficiency gains because it exploits more moment restrictions. More precisely, the set of instruments consists of all available lagged values of the dependent variables as well as lagged values of the exogenous regressors.

Similar estimators are examined in detail in studies by Arellano (1988), Arellano and Bond (1991), Arellano and Bover (1995) and Blundell and Bond (1998).

The GMM procedure is carried out in two steps. At first, the procedure yields the so-called one-step estimates, by using a known matrix for H_i. In a second stage, a two-step estimator is computed by exploiting the one-step residuals $\hat{\upsilon}_i^*$ to form $H_i = \hat{\upsilon}_i^* \hat{\upsilon}_i^{*\prime}$. The resulting estimator is more efficient when υ_{it} are heteroskedastic (White (1982)). In a GMM estimation framework the number of instruments available expands with every time period, though selecting an adequate instrument matrix Z is of paramount importance and hinges on the assumptions we make about the regressors. First, if the explanatory variables can safely be assumed to be strictly exogenous, and thus uncorrelated with future, present and past errors, they are all available to be used as instruments at every time period. The matrix of instruments Z_i for one individual would have the largest. Second, when the explanatory variables are predetermined by nature, they will be uncorrelated with the current and future errors. As a result, less instruments are available for every time period. Lastly, if the explanatory variables are endogenous then they will be contemporaneously correlated with the errors. Consequently the number of instruments for each time period declines by one term.

Since higher order lags of the dependent variable are used to instrument the lagged dependent variable a no serial correlation assumption in the disturbances υ_{it} is needed to ensure the consistency of the estimators. When there is no serial correlation in υ_{it}, the first differenced residuals $\Delta\upsilon_{it}$ will display significant negative first order serial correlation but no second order serial correlation. We therefore test the null of no correlation separately for first and second order autocorrelation.

Accordingly, a one-degree of freedom test (m_2) suggested by Arellano and Bond (1991), which tests for the absence of first-order and second-order serial correlation in the first-differenced residuals, is employed. Under the null of no autocorrelation, the standardized average residual autocovariances, upon which this test is built, are asymptotically N (0,1) variables.

The validity of the moment conditions that GMM estimation makes use of is typically tested via the Sargan test of overidentifying restrictions (Blundell, Bond and Windmeijer (2000)). The test statistic that pertains to the GMM estimator of the first-differenced model is constructed in Arellano and Bond (1991) as follows:

$$s = \hat{\upsilon}' Z \left(\sum_{i=1}^{N} Z'_i \hat{\upsilon}_i \hat{\upsilon}'_i Z_i \right)^{-1} Z' \hat{\upsilon}$$

where $\hat{\upsilon} = y - X\hat{\delta}$, and $\hat{\delta}$ is the two-step estimator for a certain matrix of instruments Z (not necessarily containing the optimal set of instruments). Under the null that the moment conditions are valid, the Sargan test statistic is asymptotically distributed as chi-square with p-k degrees of freedom (equal to the number of overidentifying restrictions), where p is the number of columns in Z and k is the number of columns in the matrix of regressors.

4 RESULTS

The GMM estimation procedure is executed in Gauss version 3.2.38 using the DPD98 program by Arellano and Bond (1998). Because the number of instruments increases with the time periods we have chosen to use only the most recent number of observations as instruments. When more instruments are used, we face memory limitation problems. The earliest instrument we use is t-7 in some cases.

The model reported in column (1) of table 4.1 assumes the explanatory variables are endogenous and correlated with the country fixed effects, while no serial correlation is assumed in the structure of errors, υ_{it}. This GMM model includes only current-dated first-differenced explanatory variables and instruments them with lagged levels of each explanatory variable from t-2 up to t-7. The following represents its specification:

$$\Delta Y_{it}=\alpha_i+\gamma_1 \Delta NEER_{i,t}+\gamma_2 \Delta M_{i,t}+\gamma_3 \Delta G_{i,t}+\gamma_4 \Delta V_{i,t}+e_{i,t} \tag{4.1}$$

This GMM model yields significant coefficients for the one-step and two-step estimators, which are not very different in size. As expected, however, the two-step estimator, which uses an optimal weighting matrix, is more efficient. The two-step estimator shows lower standard errors, by about 70%, for all slope coefficients. The positive coefficient on the nominal effective exchange rate signifies that a 10% devaluation leads to a 1% contraction of real output. On the other hand, exchange rate volatility appears to have a small expansionary effect on output rejecting the traditional view that volatility is contractionary. Money and fiscal measures carry expansionary signs. The m2 test for second-order serial correlation fails to reject the null of no serial correlation, and the Sargan test confirms the validity of the instruments. The Wald test rejects the null of jointly insignificant coefficients and a second Wald test for the joint significance of the nominal effective exchange rate and exchange rate volatility rejects the null as well.

Although none of these tests indicates model misspecification, we recognize that the above model excludes any dynamics of adjustment. Therefore, we estimate a first difference model with a structure similar to Edwards' (1989) that includes one lag for the exchange rate and another for money. Column (2), shows the results of this estimation, specified as

$$\Delta Y_{it}=\alpha_i+\gamma_1\Delta NEER_{i,t}+\gamma_2\Delta NEER_{i,t-1}+\gamma_3\Delta M_{i,t}+\gamma_4\Delta M_{i,t-1}+\gamma_5\Delta G_{i,t}+\gamma_6\Delta V_{i,t}+e_{i,t} \quad (4.2)$$

The coefficients of this model are quite similar with those of the previous one, providing further evidence that devaluations are contractionary.

Table 4.1 Basic GMM Models in First Differences with Endogenous Regressors

	(1)		(2)	
	1-step	2-step	1-step	2-step
$\Delta y_{i,t-1}$				
$\Delta NEER_{it}$	0.1184	0.1002*	0.0429	0.0231
	(0.0513)	(0.0158)	(0.0414)	(0.0149)
$\Delta NEER_{it-1}$			0.0757	0.0647*
			(0.0677)	(0.0139)
ΔM_{it}	0.1202	0.0954*	0.0748	0.0114
	(0.0551)	(0.0185)	(0.0648)	(0.0277)
ΔM_{it-1}			0.0429	0.0622*
			(0.0448)	(0.0143)
ΔG_{it}	0.1228	0.1177*	0.1284	0.1101*
	(0.0567)	(0.0094)	(0.0587)	(0.0131)
ΔV_{it}	0.0011	0.0010	0.0000	0.0002
	(0.0013)	(0.0002)	(0.0015)	(0.0003)
m_2	0.126	0.139	0.083	0.061
z_1	0.000	0.000	0.000	0.000
z_2	0.000	0.000	0.001	0.000
Sargan		30.506		23.982
Deg. of freedom		212		210
Wald: $NEER_t$ and $NEER_{t-1}$			0.054	0.000
Instruments	$x_{t-2} \ldots x_{t-7}$		$x_{t-2} \ldots x_{t-7}$	

Notes:

1) The test statistics and standard errors (reported in parenthesis) are robust to general time-series and cross-section heterosckedasticity

2) All equations include time dummies as regressors and instruments

3) m_2 is a test for second order serial correlation in the residuals. Under the null of no serial correlation, it is asymptotically distributed as N(0,1). P-value shown.

4) z1(k) is a Wald test of joint significance of the listed coefficients. Under the null of no relationship, it is asymptotically distributed as χ^2 (k). P-value shown.

5) z2(k) is a Wald test of joint significance of the time dummies. P-value shown

6) The Sargan statistic is a test of the overidentifying restrictions. Under the null, it is asymptotically distributed as χ^2 (k)

7) The Wald statistic tests the joint significance of NEERt and NEERt-1. P-value shown

8) In the list of Instruments, x represents NEER, M, G and V.

9) Two-step estimates significant at the 5% and 10% levels are marked with a (*) and (**) respectively.

In contrast with Edwards' findings that devaluations produce contractionary effects during the first year but turn expansionary in the following year, we find devaluations to be contractionary in both years, although NEERit is not significant. Moreover, a Wald test for the joint significance of the current-dated and lagged nominal effective exchange rate rejects the null hypothesis that they are jointly insignificant, which implies that devaluations do have a long-run impact on output. However these findings are somewhat questionable since at the 10% significance level, the m_2 test for second-order serial correlation rejects the null of no serial correlation, indicating possible dynamic misspecification. The exclusion of exchange rate volatility in an attempt to obtain a more loyal replication of Edwards's (1989) model made no significant difference.

Table 4.2 reports a broader set of GMM models to ascertain the robustness of the estimates. In column 3 of Table 4.2, we include a lag of all the changes in explanatory variables and a lag of the change in the dependent variable on the right hand side. This allows for an autoregressive structure of output as follows:

$$Y_{it}=\alpha_i+\gamma_0\Delta Y_{it-1}+\gamma_1\Delta NEER_{i,t}+\gamma_2\Delta NEER_{i,t-1}+\gamma_3\Delta M_{i,t}+\gamma_4\Delta M_{i,t-1}+\gamma_5\Delta G_{i,t}+\gamma_6\Delta G_{i\ i,t-1}+\gamma_7\Delta V_{i,t}+e_{i,t} \quad (4.3)$$

At the same time, we enrich the instrument set with $y_{i,t-2}$. These changes, however, do not result in a well-determined model (most coefficients are insignificant). Overall, there is a substantial downward bias in the two-step estimates as compared to the one-step estimates, which are typically more reliable for inference purposes (Blundell and Bond (1998)). The lagged first difference of the dependent variable is highly significant and its inclusion reduces the size of most other coefficients including the nominal effective exchange rate. NEER is no longer significant for either year.

Table 4.2. Broader GMM Models in Differences Allowing for Endogenous Regressors and Measurement Error

	(3)		(4)		(5)	
	1-step	2-step	1-step	2-step	1-step	2-step
$\Delta y_{i,t-1}$	0.7807	0.6935*	0.8148	0.6681*	0.8133	0.6525*
	(0.1092)	(0.0922)	(0.0850)	(0.1078)	(0.0950)	(0.1039)
$\Delta NEER_{it}$	0.0140	0.0253	0.0271	0.0463*	0.0273	0.0370*
	(0.0229)	(0.0173)	(0.0195)	(0.0160)	(0.0246)	(0.0134)
$\Delta NEER_{it-1}$	0.0114	-0.0008	-0.0106	-0.0113	-0.0005	-0.0209*
	(0.0214)	(0.0116)	(0.0168)	(0.0116)	(0.0200)	(0.0090)
ΔM_{it}	0.0427	0.0153	0.0147	-0.0064	0.0421	-0.0232
	(0.0380)	(0.0229)	(0.0254)	(0.0166)	(0.0351)	(0.0210)
ΔM_{it-1}	0.0027	0.0160	0.0198	0.0478*	0.0084	0.0397*
	(0.0222)	(0.0138)	(0.0228)	(0.0186)	(0.0202)	(0.0147)
ΔG_{it}	0.0413	0.0414*	0.0820	0.0752*	0.0596	0.0405*
	(0.0567)	(0.0096)	(0.0355)	(0.0123)	(0.0429)	(0.0122)
ΔG_{it-1}	-0.0287	-0.0233	-0.0500	-0.0473*	-0.0603	-0.0577*
	(0.0340)	(0.0136)	(0.0247)	(0.0122)	(0.0232)	(0.0113)
ΔV_{it}	0.0013	0.0007	0.0028	0.0018*	0.0021	0.0014*
	(0.0018)	(0.0004)	(0.0017)	(0.0005)	(0.0016)	(0.0004)
m_2	0.333	0.392	0.316	0.381	0.318	0.377
z_1	0.000	0.001	0.000	0.000	0.000	0.000
z_2	0.000	0.000	0.002	0.000	0.004	0.000
Sargan		24.705		23.514		20.134
Deg. of freedom		217		253		181
Wald: $NEER_t$ and $NEER_{t-1}$	0.433	0.207	0.293	0.013	0.354	0.013
Instruments	$x_{t-2} \ldots x_{t-7}$ y_{t-2}		$x_{t-1} \ldots x_{t-7}$ y_{t-2}		$x_{t-3} \ldots x_{t-7}$ y_{t-2}	

Notes: As for table 4.1

In column (4) we entertain the possibility that our regressors might be predetermined, which requires that we increase the number of instruments available by including $x_{i,t-1}$ (one more instrument for every explanatory variable). This assumption results in a larger initial impact of nominal effective exchange rates as well as exchange rate volatility on output. The signs of the two-step estimates remain as in model (3). The coefficients on lagged nominal effective exchange rates and current-dated money continue to be insignificant. According to the Sargan and the m_2 tests, the instruments used are valid. One source of skepticism about models in columns (3) and (4) is the insignificant coefficient for current money.

The inclusion of x_{it-1} and x_{it-2} in the instrument set is inappropriate if measurement error is present. For that reason, the instrument set of the model reported in column (5) eliminates both of these lags for every explanatory variable. Theoretically, the incidence of measurement error causes the coefficient estimates to be biased towards zero. Measurement error appears to be of little concern because we do not observe an increase in the slope coefficients in model (5) when compared to model (4). Surprisingly, the standard errors of the slope coefficients in model (5) do not appear to show a substantial loss in efficiency vis-à-vis model (4), although model (5) utilizes two less instruments (for every explanatory variable).

Replacing the instrument y_{it-2} with y_{it-3}, to account for the possibility of measurement error in the dependent variable, makes little difference in the coefficient estimates compared to model (5).

Table 4.3 continues the results of a broader set of GMM models. In model (6), while accounting for the presence of measurement error (i.e. we exclude x_{it}, x_{it-1}

and x_{it-2}), future instruments are included so as to examine the assumption of exogenous regressors. Otherwise, we continue to keep the same specification as in equation (4.3). The inclusion of future instruments does not produce a well-determined model. The model assumptions, however, are implausible in that both the current and lagged money variables turn out insignificant. At the same time, the coefficients on the current nominal effective exchange rate are insignificant and remain so even jointly according to the Wald test of joint significance. Nevertheless the model passes the tests for serial correlation and validity of instruments. Indeed, under the new instrument set, the slope coefficients would increase substantially, compared to model 5, if υ_{it} were serially correlated (Blundell et al., 1992). With the exception of the lagged dependent variable, no such increase is observed.

Table 4.3. Broader GMM Models in Differences Allowing for Exogenous Regressors and Measurement Error

	(6)		(7)		(8)	
	1-step	2-step	1-step	2-step	1-step	2-step
$\Delta y_{i,t-1}$	0.8290	0.7703*	0.8540	0.8070*	0.8501	0.7348*
	(0.0871)	(0.0673)	(0.0700)	(0.1351)	(0.0773)	(0.0993)
$\Delta NEER_{it}$	0.0196	0.0099	0.0214	0.0216	0.0233	0.0322*
	(0.0193)	(0.0225)	(0.0153)	(0.0176)	(0.0194)	(0.0157)
$\Delta NEER_{it-1}$	-0.0244	-0.0246*	-0.0252	-0.0144	-0.0146	-0.0237*
	(0.0192)	(0.0109)	(0.0256)	(0.0177)	(0.0205)	(0.0091)
ΔM_{it}	0.0284	-0.0053	-0.0413	-0.0441	0.0236	-0.0162
	(0.0429)	(0.0277)	(0.0428)	(0.0328)	(0.0278)	(0.0217)
ΔM_{it-1}	-0.0173	-0.0084	0.0528	0.0583*	0.0042	0.0317*
	(0.0432)	(0.0277)	(0.0373)	(0.0272)	(0.0226)	(0.0177)
ΔG_{it}	0.1034	0.0800*	0.1201	0.0857*	0.0974	0.0827*
	(0.0397)	(0.0163)	(0.0532)	(0.0213)	(0.0431)	(0.0127)
ΔG_{it-1}	-0.0325	-0.0202	-0.0623	-0.0335	-0.0539	-0.0485*
	(0.0401)	(0.0168)	(0.0233)	(0.0250)	(0.0237)	(0.0144)
ΔV_{it}	0.0028	0.0019*	0.0035	0.0020	0.0025	0.0018*
	(0.0017)	(0.0005)	(0.0021)	(0.0006)	(0.0017)	(0.0004)
m_2	0.320	0.354	0.263	0.302	0.305	0.372
z_1	0.000	0.000	0.000	0.000	0.000	0.000
z_2	0.001	0.000	0.004	0.000	0.002	0.000
Sargan		23.741		24.792		24.253
Deg. of freedom		213		321		220
Wald: $NEER_t$ and $NEER_{t-1}$	0.405	0.079	0.372	0.296	0.453	0.029
Instruments	$x_{t+4} \dots x_{t+1}$ $x_{t-3} \dots x_{t-5}$ y_{t-2}		$x_{t+4} \dots x_{t-5}$ y_{t-2}		$G_{t+5} \dots G_{t-5}$ *Other than* G: $x_{t-3} \dots x_{t-7}$ y_{t-2}	

Notes: As in table 4.1

The assumption of strict exogeneity of our explanatory variables with respect to υ_{it} is incorporated in model (7) by including past, present and future levels of the values of x_{it} in the instrument set. The same assumption of strict exogeneity is also a prerequisite for the within groups and the random effects GLS estimators.

We present and examine these latter models in the next section. The observed increase in the significant slope coefficients of model (7) compared to model (5) may likely be a result of the simultaneity bias induced by the inclusion of future instruments, offsetting any existing downward bias due to measurement error from the inclusion of x_{it}, x_{it-1} and x_{it-2} in the instrument set. All the same, this model remains poorly determined with insignificant coefficients for nominal effective exchange rates and current money.

Recognizing the fact that some variables can be exogenous while others are endogenous or predetermined, we consider assuming government spending as exogenous on the grounds that it is determined by "outside forces" such as the political process. Therefore we estimate the "hybrid" model presented in column (8). Because government spending is considered strictly exogenous, we use its past, current and future levels as instruments while maintaining the configuration of model (5) for the rest of the instruments. The model is relatively well determined, showing no signs of serially correlated disturbances, and valid instruments. The impact of devaluation on output is consistent with models (3) or (4) in terms of signs – causing a contraction in the first year and reversing most of it in the second. Its long run effect though is a significant contractionary effect as determined by the Wald test of joint significance for both nominal effective exchange rate terms. Money and Government spending have net expansionary impacts on output. The exchange rate volatility variable is positive and significant. The sign and size of the exchange rate volatility variable is consistent across the various models.

Of all the specifications we have presented we tend to favor model (8). The rationale for choosing model (8) over the rest of the models rests on several points. First, this model allows for the presence of measurement error by excluding the instruments lagged one and two times. Second, despite having fewer instruments than model (4), model (8) displays no loss in efficiency; in general, its standard errors are lower. Third, as several of the other GMM models, it passes the tests for serial correlation in the error term as well as the Sargan test for the validity of instruments. The z_1 and z_2 Wald statistics indicate that all coefficients are jointly significant.

The main implication of model (8) is that a devaluation of 10 percent causes output to decline by 0.32 percent in the first year. However, in the second year, devaluation turns expansionary and offsets some of the initial contraction in output. While the time pattern in output reactions is in line with Edwards' (1989) conclusions, the long run effects are not. The contractionary effect in the first year is larger than the positive impact in the second year, resulting in a contractionary long run effect. This result is in contrast with Edwards' conclusion that devaluations are neutral in the long run. The Wald test of joint significance of the two nominal effective exchange rate terms for model (8) verifies that they are jointly significant. The other variable of special interest is exchange rate volatility, which has an expansionary impact on output as it does consistently in all our models. This result contradicts the more intuitive theory that treats volatility in exchange rates as undesirable and inhibitive not only for trade flows but also for economic activity in general. As can be inferred from the literature review pertaining to the exchange rate volatility, this result is in line with the more recent literature that deems increased exchange rate volatility as expansionary for trade flows.

5 CONCLUSION

The objective of this work is to add to the existing empirical literature on the effect of devaluation on aggregate output. This paper expands the traditional formulations of this problem by including *exchange rate volatility* – under the assumption that high exchange rate volatility is a source of additional uncertainty in the economic environment and it may affect output just as it affects trade volumes. Our objective is to be able to make a distinction between the effects of devaluation per se and the effects of higher exchange rate variability on output.

In addition, we introduce dynamics into the system through the adoption of a dynamic panel model that uses a generalized method of moments (GMM) estimator. Such an estimation technique allows for endogenous regressors.

We believe that using data from the most recent decade will render our conclusions more useful for policy analysis as well as avoiding the regime shifts that come with longer time series. Hence, we use data on a panel of 42 countries, of which 18 are OECD countries, for the period 1988-1997. Variation in real output is expressed as a function of variation in nominal effective exchange rates, broad money, real government spending, and the 12-month exchange rate volatility.

We specify different configurations of GMM dynamic panel models in first differences that pertain to various assumptions about the endogeneity or exogeneity of the regressors. A model where government spending is assumed exogenous while the rest of the variables are taken as endogenous is our preferred specification. The impact of devaluation on output is, for the most part, consistent in all our GMM models – causing a contraction in the first year and reversing most of it in the second. Its medium-run effect then is a significant contractionary effect.

The exchange rate volatility variable is consistently positive across our models and significant in many instances. Contrary to initial intuition, our empirical findings suggest that a more volatile exchange rate has a minor positive impact on output. This result is more in line with the recent part of the literature on the impact of exchange rate volatility on trade volumes. Several arguments can be made in support of this finding. First, this result may reflect the fact that output grows faster where exchange rate regimes are more flexible. Second, while increased volatility in exchange rates, reduces economic activity in the related domains (substitution effect), more resources might be devoted to that activity (income effect) to compensate for the reduced expected total utility of the activity (Cote, 1994). Third, exchange rate volatility is a means of staying in line with purchasing power parity. Reducing such volatility may give rise to divergence from PPP and thus could increase profit uncertainty. Finally, exchange rate volatility can be seen as a risk, but also as an opportunity to make profits (Gros, 1987).

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APENDIX

All data were extracted from the International Financial Statistics (IFS) of the International Monetary Fund, in CD-ROM format. IFS Is the International Monetary Fund's principal statistical publication.

The dataset contains information for 42 countries and spans the period from 1988 to 1997 with yearly observations. The following countries were included based on data availability.

OECD Countries: Australia, Austria, Belgium, Denmark, Finland, France, Greece, Iceland, Italy, Netherlands, New Zealand, Norway, Portugal, Sweden, Switzerland, United Kingdom, United States.

Non-Oecd Countries: Belize, Burundi, Cameroon, Chile, Colombia, Costa Rica, Ivory Coast, Cyprus, Dominican Republic, Philippines, Saudi Arabia, Sierra Leone, Singapore, South Africa, St. Kitts And Nevis, St. Vincent & Grens., Togo, Trinidad And Tobago, Tunisia, Uganda, Uruguay, Venezuela, Zambia.

Real GDP, which is denoted by Y, has been extracted from subject code 99 of the section on National Accounts and Population and refers to constant price GDP. Real GDP is expressed in index form with base year 1995.

Nominal Effective Exchange Rates (NEER) were obtained from the section of Exchange Rates and Exchange Rate Arrangements. NEER is an index that equals the ratio of an index of the period average exchange rate of a given currency to a weighted geometric average of exchange rates for the currencies of a group of its trading countries. The base year for the index is 1990.

Money (M) represents a broad measure of money comparable to what is commonly referred to as the M2 monetary aggregate. This variable was computed as the total of subject codes 34 and 35, which stand for Money and Quasi-Money respectively.

Real government spending (G) was obtained from the section of Government Finance, more specifically from subject code 82 corresponding to Expenditure. In a further step, government expenditure was deflated using a domestic price index.

Finally, Exchange Rate Volatility was generated by computing the standard deviation of the monthly percentage changes in the nominal exchange rate over a period of 12 months. The monthly nominal exchange rate is an end-of-period units of national currency per U.S. dollar. Market determined nominal exchange rates were selected where available. For the rest of the panel, nominal exchange rates constitute the official exchange rate, or the principal rate as the case may be.