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Real exchange rate effects of monetary disturbances under different degrees of exchange rate flexibility: An empirical analysis

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Abstract

We examine the effects of monetary policy on the real exchange rate. In a cross-country analysis, we find that the variability of money shocks and the degree of informativeness of the exchange rate are important determinants of the magnitude of the real exchange rate effects of domestic money shocks. Our results are consistent with previous cross-country evidence on the output effects of money shocks but also highlight the role of the exchange rate regime.

Key words: Exchange rate; Monetary policy

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1. Introduction

In equilibrium rational expectations models, incomplete information enables monetary policy to have real effects. These models imply that the

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real effects of monetary disturbances vary across regimes and decrease in magnitude as the variance of the disturbances increases (see, for example, Lucas, 1972; Barro, 1976; Weiss, 1980, 1982; King 1982). Lucas (1973), Kormendi and Meguire (1984), Fry and Lilien (1986), and Montiel and Zaidi (1987) find support for this hypothesis in cross-country analyses of the real output effects of nominal shocks. Kretzmer (1989) finds evidence supporting a closely related proposition concerning the variance of real disturbances in a cross-industry analysis. These results support the view that the reduced-form output effects of nominal shocks depend on the parameters of the process governing the nominal shocks.

Previous cross-country analyses measure monetary policy regime differences only in terms of the variance of domestic monetary disturbances. However, exchange rate regimes as well as domestic monetary regimes may matter across countries. In a small open economy model, Kimbrough (1983, 1984) shows that the magnitude of output responses to disturbances depends on the exchange rate regime since the information content of the exchange rate varies across regimes. In particular, he finds that monetary disturbances have real effects under flexible exchange rates through their influence on the information content of the nominal exchange rate. In his model the nominal exchange rate does not convey information about current disturbances under fixed exchange rates, and monetary shocks then have no real effects.

Glick and Wihlborg (1990) derive explicit formulations for the real effects of unanticipated disturbances in a two-country model. They show that, under reasonable assumptions, the effect on the real exchange rate of an unanticipated monetary disturbance is larger under a flexible exchange rate regime than under a fixed regime, while the effect on the real exchange rate of an unanticipated real disturbance is smaller under flexible rates. They also derive testable implications of differences across exchange rate regimes for the strength of the relationship between the real exchange rate effects of unanticipated monetary disturbances and the relative variance of money disturbances. Their two-country framework enables analysis of the role of foreign as well as domestic disturbances.

The purpose of this paper is to test hypotheses regarding the real exchange rate effects of monetary disturbances across nominal exchange rate regimes as well as across domestic monetary regimes. We are particularly concerned with establishing empirically that the real exchange rate effects of monetary disturbances depend negatively on the variance of monetary shocks and positively on the degree of exchange rate flexibility. While there exist companion hypotheses concerning the effects of real shocks on real exchange rate adjustment, in this paper we do not test these hypotheses because of the relative lack of high frequency empirical proxies for real disturbances.

In addition to explicitly considering the degree of exchange rate flexibility, the tests performed differ from previous work in that they focus on the adjustment of real exchange rates rather than output, and they define monetary disturbances as unanticipated changes in the domestic component of the monetary base, rather than the total monetary base (or M1). The emphasis on real exchange rate effects, rather than on output effects, stems from two factors. First, the behavior of the real exchange rate is of major policy concern due to its impact on the composition as well as the level of real output. Second, monetary influences may be more apparent in the case of a real price variable, such as the real exchange rate, than in the case of output, since the latter can be subject to adjustment lags. Our definition of monetary disturbances is motivated by the observation that, unless exchange rates are perfectly flexible or changes in foreign exchange reserves are offset by sterilization policy, changes in the total monetary base (or M1) are endogenous in open economies; their use in generating measures of exogenous disturbances is therefore inappropriate, particularly when international capital mobility is high.

Despite the differences between our work and that of others, our tests should be viewed as open economy analogues of the cross-regime tests of the output effects of nominal disturbances, as done, for example, by Lucas (1973) and Kormendi and Meguire (1984). Anticipating our results, we do indeed find evidence of a strong negative relation between the magnitude of the real exchange rate effects of money shocks and the variance of money shocks. We also find evidence that this relation depends on the exchange rate regime. Thus, our results can be interpreted as lending support to the class of incomplete-information rational expectations models.

The organization of the paper is as follows. In Section 2 we state the hypotheses to be tested and summarize the theoretical arguments behind them. In Section 3 we discuss measures of the degree of exchange rate flexibility. We argue that conventional definitions do not capture the informational distinctions between different exchange rate regimes. Section 4 presents the data and results of our empirical tests based on an estimation procedure similar to that of Kormendi and Meguire (1984). We conclude with a summary in Section 5.

2. Theory and hypotheses

Before stating the specific hypotheses to be tested, we present an intuitive framework within which to understand how the nature of the exchange rate regime influences the information content of the market signals and hence the real effects of disturbances.[1]

Consider a two-country, two-good world with unobservable domestic and

[1] Specific structural models of open economies generating the hypotheses can be found in Flood and Hodrick (1985) and Glick and Wihlborg (1990).

foreign monetary as well as real disturbances. Assume that information and expectations are identical across agents. Assume also perfect capital mobility, implying that the domestic real interest rate is equated to the foreign real interest rate plus the expected rate of change in the real exchange rate. In this framework, agents' perceptions of each period's disturbances depend on signals obtained by observing current goods and asset market prices and conditions. The nature of the signals provided by conditions in asset markets, in particular, depends importantly on the exchange rate regime.

With purely flexible exchange rates, money market equilibrium requires that money supply in each country equals money demand. Each money market provides a signal about local money conditions which depends purely on local money supply and velocity disturbances as well as on real disturbances to the market for the locally-produced good through their effect on the transactions demand for money. To the extent that these real disturbances are independent across countries, the domestic and foreign money markets are informationally independent. In this case, confusion about local domestic shocks is insulated from further confusion due to foreign shocks, and local monetary disturbances are confused only with local real disturbances.

With a perfectly fixed exchange rate regime, the two local money markets become one with only the locally-created component of the money supply subject to policy control in each country. World money market equilibrium then requires that total world money supply equals total money demand arising from the two countries. When the nominal exchange rate is fixed, it conveys no information, and one signal is lost. The single signal provided by the world money market under a fixed rate regime is a composite of the local money market conditions in both countries. As a result, agents are unable to distinguish the impact of domestic from foreign disturbances on local money markets conditions, unless associated changes in foreign exchange reserves can be directly observed or are revealed to agents.

Under both exchange rate regimes the confusion between local money and real disturbances allows current monetary shocks as well as real shocks to have real effects.[2] The effect of lagged disturbances depends on the lag with which agents acquire knowledge of the magnitudes of past disturbances, the degree of persistence of disturbances, and the length of

[2] In typical open economy 'island' models (e.g. Kimbrough, 1983; Flood and Hodrick, 1985; von Hagen, 1990) the incentive to resolve this confusion arises from an asymmetric information structure in which individual agents in different markets are differentially informed. In Glick and Wihlborg (1990) all agents have homogeneous information and the incentive to resolve confusion arises from the assumption that disturbances are serially correlated and equilibrium depends on forward-looking prices – the real interest rate and the expected real exchange rate. This gives agents an incentive to infer the serially-correlated component of future disturbances under both fixed and flexible exchange rates.

structural adjustment lags due to inventory changes and capital accumulation.[3]

The effect of changes in the relative variance of monetary shocks follows from the effect on the confusion of agents. The larger the relative variance of domestic monetary shocks, the smaller is the confusion about domestic monetary shocks, and the smaller the absolute response of the real exchange rate to domestic monetary shocks. As in closed-economy incomplete information models, a negative relation between the relative variance of domestic monetary shocks and the real effects of money shocks is implied.[4]

Because of the differences in the information content of market prices across exchange rate regimes, the real effects of monetary disturbances vary across exchange rate regimes as well. In particular, because domestic and foreign money markets are more informationally independent under flexible exchange rates, it is only the relative variances of the disturbances *within* each country that determine the magnitude of adjustment to domestic monetary shocks under flexible rates. Under fixed rates, because world money market conditions are a composite of local money market conditions in the two countries, monetary shocks can be confused with foreign shocks as well as with domestic real shocks. Consequently, the response to domestic monetary shocks differs from that under flexible exchange rates. Analogous implications can be obtained for the effects of real disturbances.

2.1. Hypotheses

We are now in a position to state our hypotheses concerning real exchange rate adjustment under fixed and flexible exchange rates. To present the testable hypotheses within a cross-country framework we specify the following equation for the (absolute) magnitude of the real exchange rate response to domestic money shocks, B_{mj}:

$$|B_{mj}| = a_0 + a_1\sigma^2_{mj} + a_2 D_j + a_3\sigma^2_{yj} + a_4 D_j \sigma^2_{mj} , \quad (1)$$

[3] If, for example, agents learn the magnitude of all disturbances with a one-period lag, all disturbances follow AR1 processes, and there are no other structural adjustment dynamics, then the information confusion problem is fully resolved after one period. In this case, only current monetary shocks have real effects; lagged monetary shocks have no real effect (see Glick and Wihlborg, 1990).

[4] Kormendi and Meguire (1984) and others have analyzed the real output effects of monetary shocks in limited information, rational expectations models. Glick and Wihlborg (1990), using a model with a Lucas-type supply function in which output supply depends positively on the local real interest rate, show how the real exchange rate effects of monetary shocks translate directly into real interest rate and output effects. Specifically, a monetary shock inducing an expected real depreciation of the domestic currency also causes a rise in the domestic real interest rate (to maintain interest parity), and hence leads to a (temporary) rise in aggregate supply.

where σ^2_{mj} denotes the variance of unanticipated domestic money disturbances, σ^2_{yj} denotes the variance of unanticipated domestic real disturbances, D_j is a dummy shift variable with a value of 1 for a flexible exchange rate regime and a value of 0 for a fixed exchange rate regime, and j indexes these variables by country.[5] The main hypotheses and their rationales are stated below.

(i) $a_1 < 0$. An increase in the variance of the domestic money supply causes a *decrease* in the adjustment to domestic monetary disturbances. Intuitively, the larger is the relative variance of domestic money shocks, the less is the confusion about these shocks, and the smaller is the response of the real exchange rate.

(ii) $a_2 > 0$. The magnitude of the real effect of monetary disturbances depends on the exchange rate regime and is *stronger* under flexible rates.[6] As discussed above, in a rational expectations, incomplete information framework, the exchange rate regime affects the information inferred about underlying disturbances and hence the magnitude of the real response to these disturbances. Under a fixed exchange rate regime, when world money market conditions are a composite of local money market conditions in the two countries, the ability to discern between domestic and foreign shocks is less than under a flexible exchange rate regime where confusion about local domestic shocks is insulated from further confusion about foreign shocks. Assuming that domestic and foreign real shocks have opposing effects on the real exchange rate, i.e. a domestic real shock causes a real exchange rate appreciation while a (similar) foreign real shock causes a real depreciation, the response of the real exchange rate to a given monetary shock is *less* under fixed rates. The reason is that there is greater confusion with a fixed rate regime about whether a shock is domestic or foreign in nature and therefore about whether the exchange rate will appreciate or depreciate.[7]

(iii) $a_3 > 0$. The adjustment to domestic monetary disturbances is *positively* related to the variance of domestic real disturbances under both fixed and

[5] Variances of foreign shocks are not incorporated in (1) because there is no cross-sectional variance in these shocks in our empirical tests below. The United States is the foreign country for all countries indexed by j.

[6] See Glick and Wihlborg (1990, Proposition 2). This property can be interpreted as the information model analogy to the greater effectiveness of monetary policy under a flexible exchange rate regime in a Mundell–Fleming model. The result is not affected if domestic and foreign real disturbances are correlated.

[7] This hypothesis is derived under the strong rational expectations assumption that equilibrium movements in the real exchange rate arise solely from fundamentals. If there is substantial variability in the exchange rate unrelated to fundamentals, then it can be shown that the informativeness of the exchange rate decreases. Such 'noise' in the exchange rate reduces the difference in informativeness between fixed and flexible exchange rate regimes.

flexible rates. The intuitive explanation for this hypothesis is that confusion about monetary disturbances increases with the variance of real disturbances.

(iv) $a_4 < 0$. The effect of an increase in the variance of the domestic money supply on adjustment to domestic monetary disturbances is *stronger* (i.e. more negative) under flexible rates than under fixed rates. Intuitively, this hypothesis follows from hypotheses (i) and (ii). By hypothesis (i), as the relative variance of domestic monetary shocks rises, all shocks are perceived more as monetary in nature, and the real effects of domestic monetary shocks decline. By hypothesis (ii), the real exchange rate effects of monetary shocks are relatively large under a flexible exchange rate regime since domestic money market conditions are insulated to a greater degree from confusion about foreign disturbances. Since shocks are less 'diversified' under flexible exchange rates, a shock from a particular source represents a larger share of the total shocks about which there is confusion. Consequently, a given increase in the variance of domestic money shocks leads to a proportionately larger decline in the sensitivity of the real exchange rate to money shocks.[8]

It should also be pointed out that rigid price models, such as the Mundell–Fleming framework, also imply that monetary shocks have real effects. Moreover, they also yield the hypothesis that the real effects of monetary shocks are greater under flexible exchange rates than under fixed rates because of the extra 'kick' of exchange rate adjustment through net exports. However, rigid price models generally do not yield hypotheses concerning the variances of shocks.

3. Degree of exchange rate flexibility

In the previous section the nominal exchange rate was assumed to be either perfectly flexible ($D = 1$) or perfectly fixed ($D = 0$). In reality, such extremes seldom exist, raising the question of how best to measure the degree of exchange rate flexibility. The IMF uses the labels 'independently floating', 'other managed floating', 'adjusted according to a set of indicators', 'cooperative exchange arrangement', 'currency pegged to dollar', etc. to classify countries by exchange rate regime. However, in a limited-information, rational expectations framework such as ours, it is not obvious

[8] Glick and Wihlborg (1990) show that the effect of an increase in σ^2_{mj} on the sensitivity to real disturbances may be stronger or weaker under flexible exchange rates than under fixed exchange rate.

that the IMF's definitions appropriately distinguish among fixed and flexible exchange rate regimes for empirical purposes.[9]

Consider a country in which exchange rate changes are driven by hyperinflationary monetary policy. In this case, exchange rate changes may primarily reflect inflation anticipated by private agents, and the exchange rate in any given period may not convey information about contemporaneous disturbances. In terms of our model, this country's exchange rate regime can be regarded as perfectly fixed in terms of its information content.

Conversely, consider a country in which the nominal exchange rate is officially pegged and characterized by relatively low variance. Exchange rate changes may convey information if intervention points are sufficiently far apart that the exchange rate is allowed to adjust to disturbances, and unanticipated exchange rate changes are not large enough to trigger intervention. Similarly, if the central bank always intervenes in order to reduce unanticipated exchange rate changes by a given proportion, or by some rule known by agents and involving only observable variables, then exchange rate changes are still informative. In this case, the exchange rate regime is informationally equivalent to a perfectly flexible regime, since the degree of intervention can be inferred by agents from observation of the exchange rate and whatever variables may enter into the monetary authorities' policy rule. As a result, foreign money market conditions can be distinguished from domestic money market conditions.

This discussion suggests that to test our hypotheses it is desirable to measure the degree of exchange rate flexibility by the degree to which exchange rate changes potentially convey information about contemporaneous disturbances rather than simply by the variance of nominal exchange rate changes. It accords with the point made in Section 2 that under the assumptions that contemporaneous foreign exchange reserve changes are not revealed to agents by monetary authorities and cannot otherwise be inferred by agents from observable variables, (unsterilized) foreign exchange market intervention potentially creates noise in the exchange rate as a signal about disturbances (see Flood and Hodrick, 1985; and Kimbrough, 1984).

Accordingly, we define an alternative measure of the degree of exchange rate flexibility, reflecting informativeness, as

$$DF_j = \sigma_{sj}^2 / \left(\sigma_{sj}^2 + \sigma_{rj}^2\right),$$

where σ_{sj}^2 denotes the variance of unanticipated changes in the nominal exchange rate, and σ_{rj}^2 denotes the variance of unperceived changes in

[9] See Glick et al. (1990) for further discussion of different approaches to measuring exchange rate flexibility.

foreign exchange reserves in domestic currency terms, measured as a fraction of the monetary base for country j, where by 'unperceived' changes we mean unanticipated current reserves changes not revealed by observation of current exchange rate changes. Under the assumption that foreign exchange market intervention that changes the monetary base by 1 percent prevents a 1 percent exchange rate movement, the variance of incipient exchange rate changes prevented by exchange market intervention is captured by σ_{rj}^2.

This measure of exchange rate flexibility implies that the exchange rate is perfectly flexible ($DF = 1$) and informative if all changes in foreign exchange reserves can either be anticipated or inferred from available information $(\sigma_r^2 = 0)$. The exchange rate is perfectly fixed ($DF = 0$) and uninformative if all exchange rate changes are anticipated $(\sigma_s^2 = 0)$ or if the variance of unperceived changes in foreign exchange reserves is very large relative to the variance of unanticipated exchange rate changes $(\sigma_r^2 \to \infty)$. Values of DF between 0 and 1 denote intermediate degrees of exchange rate flexibility and informativeness.

The hypotheses of Section 2 can be tested using this measure of exchange rate flexibility. The coefficient D in (1) which previously was interpreted as having one value under fixed exchange rates (0) and another value under flexible exchange rates (1) will take on a continuum of values depending on the *degree* of exchange rate flexibility.

Columns (1) and (2) of Table 1 present estimates of two versions of the degree of flexibility measure, DF, for 31 countries. One version employs bilateral nominal exchange rate changes against the dollar; the second version employs multilateral (i.e. effective trade-weighted) nominal exchange rate changes. These estimates are used subsequently to test the hypotheses of Section 2.[10]

The calculations were performed by first running regressions for monthly percent changes in the nominal exchange rate (expressed as domestic currency units/U.S. dollar for the bilateral estimates and as domestic currency units/foreign currency basket for the multilateral estimates) and for monthly changes in foreign exchange reserves (in domestic currency terms) as a fraction of the lagged monetary base, for each country. The independent variables in the exchange rate regression were percent changes in the lagged exchange rate, lagged growth in domestic and U.S. industrial production, lagged growth in the monetary base created domestically and in

[10] All data were obtained from the IMF *International Financial Statistics*. The data range, with some exceptions, generally extended from August 1973 through September 1991. The actual data range for each country is presented in column (1) of Table 2. The sample of countries was dictated by considerations discussed in Section 4. See the Data appendix for further details.

Table 1
Exchange rate flexibility measures, 1973–1991

Country	*DF* bilateral[a] (1)	*DF* multilateral[b] (2)	IMF Classification[b] (3)	*DN*[c] (4)	*DB*[d] (5)
Japan	0.91	0.78	FI	x	x
Argentina	0.84	–	FM/P$(3/91)	–	x
Spain	0.81	0.52	FM/FC(6/89)	–	x
Israel	0.78	–	PC/FM(10/77)/PC(8/86)	–	x
Bolivia	0.78	0.82	P$/FM(8/85)	–	–
Italy	0.77	0.27	FC	x	x
Germany	0.77	0.15	FC	x	x
Greece	0.64	0.34	FM	–	x
France	0.61	0.09	FC	x	x
Mexico	0.60	–	FM/FI(8/82)	–	x
Switzerland	0.58	0.16	FI	x	x
Austria	0.54	0.02	PC	–	–
Chile	0.51	0.50	P$/PI(12/83)	–	–
Belgium	0.50	0.04	PC	x	x
Ecuador	0.40	0.45	P$/FM(8/86)	–	–
Netherlands	0.39	0.02	PC	x	x
South Africa	0.36	0.32	P$/FI(1/79)	x	x
Portugal	0.35	0.04	PI	–	–
Australia	0.34	0.18	FM/FI(10/83)	–	x
Venezuela	0.33	0.23	P$/FI(3/89)	–	–
Sweden	0.21	0.04	PC	–	–
Iceland	0.17	0.09	FM/PC(8/84)	–	x
Canada	0.16	0.08	FI	x	x
Singapore	0.14	–	PC	–	–
Ireland	0.13	0.01	P£/FC(3/79)	x	x
Korea	0.11	–	P$FM(2/80)	–	x
Finland	0.08	0.01	PC	–	–
Norway	0.07	0.01	PC	–	–
Malaysia	0.05	0.02	PC	–	–
Denmark	0.05	0.00	FC	x	x
Colombia	0.01	0.18	PI	–	–

[a] $DF = \sigma_s^2/(\sigma_s^2 + \sigma_r^2)$, where σ_s^2 = variance of unanticipated percent monthly change in nominal exchange rate, σ_r^2 = variance of unperceived monthly changes in foreign exchange reserves as a fraction of the monetary base. *DF* bilateral and *DF* multilateral constructed using nominal bilateral $ and multilateral exchange rates, respectively. See text for details of construction.

[b] source of IMF classifications: IMF *International Financial Statistics*; FI = independent float; FM = managed float; FC = cooperative float; P$ = peg to $; P£ = peg to £; PC = peg to composite; PI = indicator adjustment. Dates in parentheses denote timing of classification changes.

[c] *DN*: independent float (FI) + cooperative float countries (FC).

[d] *DB*: narrow definition (DN) + managed float countries (FM).

the United States,[11] and lagged changes in foreign exchange reserves measured as a fraction of the (lagged) monetary base.[12] Six monthly lags of all variables were used. The specification of the equation for the change in foreign exchange reserves as a fraction of the monetary base was identical except that percent changes in the current nominal exchange rate were included in order to capture any systematic relation between contemporaneous and observable exchange rate changes and foreign exchange market intervention.[13]

The variance of the residuals of each exchange rate equation was used to measure the variance of unanticipated exchange rate changes (σ_s^2), while the variance of the residuals of each reserve equation was used to measure the variance of unperceived changes in foreign exchange reserves as a fraction of the monetary base (σ_r^2). *DF* was then calculated as the variance of the residuals of the first equation relative to the sum of the variances of the residuals of the two equations.

In Table 1 we order the 31 countries in the sample in terms of their degree of exchange rate flexibility by our bilateral *DF* measure, reported in column (1). The results based on the multilateral measure are in column (2). Higher values of *DF* correspond to increasing degrees of exchange rate flexibility. For comparison we also present the IMF's exchange rate regime classifications for these countries.

Table 1 shows that the IMF's classification is an imperfect predictor of the degree of flexibility by our definition. For example, Bolivia, classified by the IMF as pegged to a currency basket for most of the sample period, was found to have a high degree of exchange rate flexibility by our measure. Canada and Australia, classified by the IMF as independently floating, have relatively low *DF* values.

The last two columns of Table 1 categorize the countries in the sample into fixed and flexible exchange rate groupings based on the IMF's classifications. We employ these groupings as a benchmark against which our alternative measure of exchange rate flexibility can be compared. The

[11] The domestically-created monetary base was defined as the monetary base minus foreign exchange reserve holding (excluding gold), measured in domestic currency terms.

[12] Interest-rate parity considerations suggest that the beginning-of-month one-month interest rate differential between the U.S. dollar and the local currency should be useful for predicting monthly exchange rate changes. However, such data are not available over the time period used for many of the countries in our sample. Our specification can be viewed as proxying for the omitted interest rate differential.

[13] Bilson and Frenkel (1979), Edwards (1983a), and others have estimated partial adjustment international reserve demand equations, with a focus on measuring the speed of adjustment. They point out that if the errors in these equations are serially correlated, then OLS estimates are inconsistent, and propose the use of pooled time-series cross-section estimates to overcome this potential problem. In general, the residuals in our foreign exchange reserve equations are not serially correlated.

'narrow' definition of exchange rate flexibility (DN) according to the IMF's classification includes independently floating currencies plus all cooperatively floating countries; and the 'broad' definition (DB) includes managed-float currencies as well. Countries that were reclassified at some time over the period were categorized according to the regime in effect for the bulk of the sample period.

4. Tests of the hypotheses

This section tests the hypotheses presented in Section 2. The estimation procedure that we follow has three stages, and is similar to the procedures of Kormendi and Meguire (1984) and Kretzmer (1989). The first stage is the estimation of unanticipated money growth and unanticipated industrial production growth for each of the 31 countries in our sample plus the United States. This stage provides us with estimates of money shocks and real shocks. The second stage employs the unanticipated variables from the first stage in a reduced-form equation for the real (bilateral or multilateral) exchange rate for each country. We include unanticipated domestic money, unanticipated domestic industrial production, unanticipated U.S. money, unanticipated U.S. industrial production, and lags of these variables in both the bilateral and multilateral cases. It is from the second stage that we derive estimates of B_{mj}, the effect of domestic money shocks on the real exchange rate, for use in the third stage regressions. The third stage regressions directly test the hypotheses of Section 2 by using estimates of domestic monetary variance and domestic real variance proxied by the sample variances of the shocks from the first stage, together with our measures of exchange rate flexibility, to explain the variation in $|B_{mj}|$ across countries.[14] Each stage will be discussed in turn after a brief discussion of the data.

[14] Efficiency could be gained if the three stages were combined by taking account of the appropriate cross-equation restrictions into one estimation procedure. Mishkin (1982) and Pagan (1984) discuss this with respect to combining the first and second stages in tests of the real output effects of unanticipated money. Fry and Lilien (1986) combine the second and third stages in an analysis of differential output growth rates by estimating a pooled cross-section, time-series regression.

However, as long as the errors in the first-stage prediction are serially uncorrelated and there is no measurement error, the estimates of the coefficients of exchange rate adjustment derived in the second stage of our procedure are consistent, though inefficient. Since we are unconcerned with hypothesis testing in the second stage, our results are not affected by this problem. Moreover, we found that substituting the cross-equation restriction implied by the third stage into the second stage and performing pooled cross-section time series estimation using the contemporaneous exchange rate effect gave results very similar to those obtained by the procedure described in the text. It is problematic to implement the joint second and third stage estimation using the peak effects since the peak effects vary across countries.

The data are taken from the IMF *International Financial Statistics* and are described fully in the Data appendix. The sample of 31 countries, plus the United States, includes all countries for which the following monthly series are available for a sufficiently long sample period: (i) the domestically-created money supply, defined as the monetary base minus foreign exchange reserve holdings (excluding gold); (ii) the real exchange rate, defined as the end-of-period foreign exchange value of the dollar times the U.S. wholesale price index divided by the domestic wholesale price index; and (iii) industrial production. The countries selected are listed alphabetically in Table 2. Nominal and real multilateral exchange rate measures were available for all but five of these countries. Due to nonstationarity in the levels, all variables were transformed into growth rates when used in estimation.[15]

4.1. Stage 1: Money and industrial production equations

The first stage in our procedure is the decomposition of domestically-created money supply growth and industrial production growth into their anticipated and unanticipated components. We employ one decomposition uniformly for all countries, regressing money supply growth and the growth rate of industrial production each on six lags of domestic money supply growth, six lags of growth in domestic industrial production, six lags of U.S. money supply growth, six lags of growth in U.S. industrial production, an intercept term, monthly seasonal dummies, and a time trend.[16] Table 2 gives the sample period used for each country. The residuals of each country's money and industrial production equations serve as measures of unanticipated money and unanticipated industrial production for use in the second-stage regressions. The sample variances of unanticipated money and unanticipated industrial production for each country (denoted by σ^2_{mj} and σ^2_{yj} for country j) are used in the third-stage regressions and are reported in columns (2) and (3) of Table 2.

The estimates of σ^2_{mj} reported in Table 2, column (2), have been multiplied by 1000 for expositional purposes. They range from a minimum of 0.17 for the United States to a maximum of 330.64 for Argentina. Four of

[15] Domestically-created money supply growth was defined as monetary base changes minus foreign exchange reserve holding changes (excluding gold) measured in domestic currency terms, all as a proportion of the lagged monetary base level.

[16] There are certainly other variables that might be considered for inclusion in these prediction equations. Edwards (1983b) and Montiel and Zaidi (1987), for example, emphasize that the role of fiscal deficits in the growth of the money supply is quite important for many developing countries. However, such data are not available on a monthly basis. Given the number of variables already included in the money and industrial production equations as well as the desire to follow a procedure that can be uniformly implemented for each country in the sample, we feel that our specification is acceptable.

Table 2
Stage 1: Money and industrial production equations

Country	Data range of dependent variable (1)	σ_m^2 [a] (2)	σ_y^2 [b] (3)	Q-m.s.l. for residuals from money equation[c] (4)	Q-m.s.l. for residuals from industrial production equation[d] (5)
Argentina	73:08–87:06	330.64	0.97	1.00	0.22
Australia	73:08–91:09	1.95	4.24	0.81	0.61
Austria	73:08–91:09	1.00	2.86	0.95	0.59
Belgium	73:08–91:09	1.26	5.42	0.30	0.00
Bolivia	73:08–91:09	23.46	26.98	1.00	0.93
Canada	73:08–91:09	1.51	1.85	0.65	0.01
Chile	79:03–91:09	4.51	10.14	0.18	0.72
Colombia	73:08–86:09	1.65	13.04	0.34	0.99
Denmark	74:08–91.09	27.75	7.58	0.93	0.06
Ecuador	75:08–87:02	4.32	59.74	0.80	0.21
Finland	73:08–91:09	6.58	5.94	0.46	0.17
France	73:08–91:09	2.19	2.67	0.70	0.27
Germany	73:08–91:09	1.50	2.47	0.21	0.00
Greece	73:08–91:09	1.28	9.11	0.74	0.90
Iceland	73:08–91:09	7.27	1523.20	0.92	0.00
Ireland	76:08–91:09	7.41	8.09	0.75	0.56
Israel	73:08–91:09	13.94	12.62	0.12	0.97
Italy	73:08–91:09	0.47	4.34	0.64	0.38
Japan	73:08–91:09	0.44	1.19	0.86	0.12
Korea	73:08–91:09	4.62	5.77	0.73	0.14
Malaysia	73:08–91:05	4.13	16.50	0.38	0.09
Mexico	73:08–91:09	8.21	7.40	0.47	0.28
Netherlands	73:08–91:09	2.86	5.12	0.47	0.88
Norway	73:08–91:09	12.43	19.01	0.39	0.15
Portugal	76:08–91:09	4.82	6.05	0.85	0.88
Singapore	74:08–91:09	3.76	0.19	0.79	0.00
South Africa	73:08–91:09	3.56	4.54	0.38	0.32
Spain	73:08–91:09	2.00	3.46	0.62	0.24
Sweden	73:08–91:09	5.16	6.44	0.25	0.91
Switzerland	73:08–91:09	0.85	0.38	0.93	0.01
United States	73:08–91:09	0.17	0.57	0.24	0.82
Venezuela	73:08–91:09	11.15	22.51	0.99	0.08

[a] Sample variance of residuals from money equation. Numbers have been multiplied by 1000.

[b] Sample variance of residuals from industrial production equation. Numbers have been multiplied by 10,000.

[c] Marginal significance level for Q test that all residual autocorrelations for each money equation are zero.

[d] Marginal significance level for Q test that all residual autocorrelations for each industrial production equation are zero.

the estimates of monetary variance are less than one, and 21 of the estimates, almost two-thirds of the 32 country estimates, are less than five. The skewed distribution of the σ^2_{mj} estimates across the countries is amenable to a log transformation. A log transformation is appropriate for the estimates of σ^2_{yj} as well, which range from a minimum of 0.19 for Singapore to a maximum of 1523.20 for Iceland. In the remainder of the analysis we drop Argentina because of its extreme value of σ^2_m and Iceland because of its extreme value of σ^2_y.

A necessary (but not sufficient) condition for a variable to be unanticipated is that it should be unpredictable from its own past values. Table 2, column (4), reports the marginal significance levels by country for the test (based on the Q-statistic for the residuals) that all residual autocorrelations from the money equation are zero. The null hypothesis that all residual autocorrelations are zero cannot be rejected at the 0.05 level of significance for any of the 32 countries, including the United States. Only two countries (Chile and Israel) have a marginal significance level less than or equal to 0.20. Our money equations do a good job of whitening the residuals.

Table 2, column (5), reports the evidence on the residual autocorrelations of each country's industrial production equation. The null hypothesis that all residual autocorrelations are zero can be rejected at the 0.05 level of significance for six of the 32 countries in the sample. Nine countries have marginal significance levels less than or equal to 0.10, and 13 countries have marginal significance levels less than or equal to 0.20. These statistics suggest that additional modeling of industrial production may be helpful. Since we employ the domestically-created base as our measure of money, excluding foreign reserve holdings, our monetary measure base should be less sensitive to exchange rate changes than alternatives such as the monetary base or M1. Thus our measures of monetary variance should be largely independent of our measure of exchange rate flexibility. This is confirmed by the relatively low cross-country correlations between σ^2_m and the bilateral and multilateral exchange rate flexibility measures of -0.15 and 0.09, respectively.

4.2. Stage 2: Real exchange rate equation

The second stage in our procedure involves estimating the sensitivity of the real exchange rate to monetary (and real) disturbances, using the measures of unanticipated money and unanticipated industrial production from the first stage. Specifically, we regress the monthly rate of change of the bilateral or multilateral real exchange rate for each country on contemporaneous and six monthly lags of unanticipated domestic money growth, unanticipated U.S. money growth, unanticipated growth in domestic industrial production, and unanticipated growth in U.S. industrial production.

We also include an intercept term, time trend, and monthly dummies as regressors.[17]

This specification may be interpreted as a differenced version of a reduced-form expression for the (log of the) equilibrium real exchange rate as a function of a general set of monetary and real disturbances of both domestic and foreign origin. We model the real shock effects by including contemporaneous and six monthly lags of unanticipated domestic and U.S. industrial production. The number of lags seemed appropriate after examination of cross-correlations of the rate of change of the real exchange rate and unanticipated domestic and U.S. industrial production for a variety of countries.

As noted in Section 2, the duration of the real effects of monetary shocks in an incomplete information framework depends on the lag with which agents acquire knowledge of the magnitudes of past disturbances, the degree of persistence of disturbances, and the length of structural adjustment lags due to inventory changes and capital accumulation. In order to allow for the possibility that real exchange rate effects may take several months to achieve, six lags of unanticipated (domestic and U.S.) money are included in the real exchange rate equations.[18]

Table 3 presents estimates of the magnitude of the effect on bilateral and multilateral real exchange rates of unanticipated domestic money. Following (1), the effect of unanticipated domestic money on the real exchange rate is denoted B_{mj}. Both 'impact' and 'peak' effect measures of B_{mj} are reported for each country. The impact effect refers to the coefficient on contemporaneous unanticipated domestic money in the real exchange rate equation. The peak effect is defined as the maximum algebraic sum of the coefficients on contemporaneous *and* lagged unanticipated domestic money. The peak effect measure of the response of the real exchange rate to unanticipated money allows for differences in the timing of the response across countries.

[17] Identification of the effects of unanticipated domestic money in the second stage require that the residuals from the first and second stage regressions are uncorrelated. This rules out the possibility that domestic money depends on contemporaneous real exchange rate changes or shocks through a policy feedback mechanism.

Barro (1978), Mishkin (1982), and others test another implication of rational expectations macro models, namely that only unanticipated monetary policy has real effects, while anticipated policy changes do not. As discussed in Barro (1978, 1981), such tests require additional identifying restrictions, such as the presence of regressors in the first-stage equations that do not directly affect the dependent variable of the second stage. Whether or not such additional restrictions are imposed does not affect the identification of unanticipated policy effects nor the feasibility of subsequent cross-regime tests.

[18] In the long run however, the real exchange rate should be neutral with respect to unanticipated money shocks (see footnote 21). We also estimated these equations with the lagged real exchange rate change as an explanatory variable; this change in specification had little effect on the results.

Table 3
Stage 2: Bilateral and multilateral real exchange rate effects of domestic unanticipated money (B_{mj})

Country	Bilateral exchange rate			Multilateral exchange rate		
	Impact effect (1)	Peak effect (2)	*N*	Impact effect (3)	Peak effect (4)	*N*
Australia	−0.16 (3.04)**	−0.16 (3.04)**	0	−0.17 (4.26)**	−0.29 (3.63)**	3
Austria	−0.43 (5.89)**	−0.43 (5.89)**	0	−0.03 (1.39)	−0.04 (1.15)	1
Belgium	−0.13 (2.05)**	−0.13 (2.05)**	0	−0.04 (2.29)**	−0.10 (2.31)**	4
Bolivia	0.17 (1.53)	0.17 (1.53)	0	0.10 (0.63)	−0.19 (0.47)	6
Canada	−0.16 (7.24)**	−0.16 (7.24)**	0	−0.09 (4.27)**	−0.22 (3.82)**	6
Chile	−0.01 (0.17)	0.07 (0.64)	6	−0.05 (1.22)	−0.20 (1.80)*	6
Colombia	−0.00 (0.01)	−0.19 (3.17)**	5	0.06 (1.17)	0.25 (1.57)	6
Denmark	−0.03 (2.36)**	−0.03 (1.32)	2	−0.01 (1.18)	−0.03 (2.11)**	6
Ecuador	−0.07 (0.93)	0.08 (0.49)	4	−0.11 (1.12)	0.24 (0.91)	5
Finland	−0.03 (1.10)	−0.08 (1.37)	5	0.02 (2.30)**	0.06 (2.52)**	5
France	−0.06 (1.31)	−0.15 (1.26)	5	−0.01 (0.62)	−0.07 (2.44)**	2
Germany	−0.22 (3.62)**	−0.34 (2.50)**	4	−0.06 (3.22)**	−0.16 (3.65)**	5
Greece	−0.06 (1.02)	−0.22 (1.43)	5	−0.00 (0.01)	0.08 (0.85)	3
Ireland	−0.08 (2.84)**	−0.15 (2.07)**	5	0.00 (0.46)	0.04 (1.79)*	6
Israel	−0.08 (2.14)**	−0.22 (2.10)**	6	– –	– –	–
Italy	−0.26 (2.61)**	−0.33 (2.38)**	1	−0.04 (0.92)	−0.07 (0.70)	4
Japan	−0.37 (3.33)**	−0.37 (3.33)**	0	−0.37 (4.33)**	−0.63 (2.82)**	5
Korea	0.01 (0.50)	−0.06 (1.06)	5	– –	– –	–
Malaysia	−0.06 (3.32)**	−0.06 (1.44)	5	0.01 (0.72)	0.09 (2.22)**	5
Mexico	−0.00 (0.03)	−0.14 (0.86)	4	– –	– –	–
Netherlands	−0.19 (4.27)**	−0.21 (1.91)*	5	−0.03 (2.45)**	0.05 (1.88)*	6
Norway	−0.07 (3.97)**	−0.08 (1.79)*	6	−0.00 (0.22)	0.01 (0.81)	4
Portugal	−0.06 (1.69)	−0.13 (1.95)*	2	0.01 (0.57)	0.05 (1.32)	4
Singapore	−0.19 (9.47)**	−0.23 (5.86)**	3	– –	– –	–
South Africa	−0.12 (2.47)**	−0.31 (2.39)**	6	−0.10 (1.62)	−0.40 (3.32)**	3
Spain	−0.00 (0.06)	−0.12 (0.90)	5	−0.01 (0.18)	−0.18 (2.00)**	6
Sweden	0.00 (0.08)	−0.16 (2.03)**	6	0.01 (0.58)	−0.05 (1.19)	6
Switzerland	−0.14 (1.57)	−0.32 (1.56)	4	−0.06 (1.66)	−0.13 (1.53)	4
Venezuela	0.09 (1.64)	0.12 (1.50)	1	0.02 (0.50)	−0.09 (0.71)	6

Notes:

Impact effect denotes coefficient on contemporaneous domestic money. Peak effect denotes maximum (algebraic) sum of coefficients on contemporaneous and up to six lags of domestic money. *N* denotes number of months (0–6) at which peak effect is achieved.

t-statistics for null that the effect equals zero are in parentheses after coefficient estimates; *denotes significance at the 0.10 level; **denotes significance at 0.05 level.

The coefficients for the impact and peak effects of money shocks on bilateral exchange rates are shown in Table 3, columns (1) and (2), along with *t*-statistics in parentheses for the test that the effect equals zero.[19] The

[19] Since the procedure followed selects the maximum sum from among several partial sums, it likely does not have the standard Student-*t* distribution. The *t*-statistics reported in Table 3 thus should be viewed as being illustrative only. We are grateful to a referee for pointing this out to us.

impact and peak effect estimates of B_{mj} based on multilateral exchange rates are shown in columns (3) and (4) of Table 3.

Columns (1) and (2) of Table 3 indicate that in 25 of 29 countries (Argentina and Iceland are excluded) a domestic money shock leads to a contemporaneous real domestic-currency depreciation against the dollar ($B_m < 0$) for both the impact and peak effect measures of the real exchange rate response.[20] The impact effect of unanticipated domestic money on the bilateral real exchange rate is significant at the 0.05 level for 15 countries. The peak effect is significant at the 0.05 level for 13 countries and at the 0.10 level for 16 countries.[21]

The estimates of the contemporaneous and peak real exchange rate effect of unanticipated domestic money using multilateral real exchange rates are reported in columns (3) and (4) of Table 3. The peak real exchange rate effect is significant at the 0.05 level for 11 countries and at the 0.10 level for 14 countries, out of a total of 25.[22] The contemporaneous effect is significant at the 0.05 level only in seven countries. The peak effect measures of the real exchange rate response to money shocks are used as dependent variables in the cross-country and cross-regime tests that follow.

4.3. Stage 3: Cross-country and cross-regime tests

The third step in our estimation procedure is the explanation of country real exchange rate responses to domestic money shocks, $|B_{mj}|$, using country estimates of unanticipated monetary and real variances, σ^2_{mj} and σ^2_{yj}, respectively, and measures of exchange rate flexibility. We carry out this

[20] With the impact effect measures, the real exchange rate appreciates for four countries – Bolivia, Korea, Sweden, and Venezuela. With the peak measure, appreciation occurs for Bolivia, Chile, Equador, and Venezuela. These effects are significant at the 0.20 level only for Bolivia and Venezuela.

Glick and Wihlborg (1990) show that B_m theoretically can be either negative or positive. In their framework, agents underestimate the magnitude of the domestic monetary disturbance and partially interpret the corresponding excess supply of money as arising from a decrease in the demand for domestic goods. Since observed price changes in the goods market signal the composite effects of domestic real demand and cost shocks, the underestimate of the demand shock from the money market implies an overestimate of the cost shock in the goods market. The net effect of these misperceptions on the real exchange rate depends on the relative persistence of real demand and cost shocks.

[21] We tested for long-run money neutrality by allowing 12 lags of money shocks in the second-stage regressions with bilateral exchange rates. The sum of the coefficients on the contemporaneous and 12 lag terms was insignificantly different from zero at the 0.10 level in all but three countries (Colombia, Singapore, South Africa).

[22] As an alternative measure of money we used the total monetary base, including foreign exchange reserves. If capital mobility is not perfect, a country may target this monetary aggregate rather than the domestic component. In the second-stage regressions, we found fewer significant coefficients for exchange rate adjustment to monetary base shocks.

step by estimating (1) as a linear regression. The estimates of real exchange rate response are derived from the second stage of our procedure, reported in Table 3. All results reported below use the absolute value of the maximum (algebraic) sum of the coefficients on contemporaneous and up to six lags of unanticipated domestic money from the real exchange rate equations, as the dependent variable for the third stage. The estimates of monetary and real variances for each country in our sample are derived from the first stage, reported in Table 2. The measures of exchange rate flexibility—*DF*, *DN*, and *DB*—are discussed in Section 3 and reported in Table 1. As suggested earlier, we apply a log transformation to the dependent variable, $|B_{mj}|$, and to the measures of monetary and real variance, σ^2_{mj} and σ^2_{yj}.[23] The results of the third stage are reported in Tables 4 and 5.[24]

We first test the hypotheses (i) that the magnitude of the real effects of monetary disturbances, $|B_{mj}|$, depends negatively upon the variance of domestic money shocks ($a_1 < 0$), and (ii) the real exchange rate response to domestic money is stronger the higher the degree of exchange rate flexibility ($a_2 > 0$).

The results provide corroboration for the results of Kormendi and Meguire (1984). We find strong evidence in column (1) of Table 4, where no distinction is made between exchange rate regimes, of a negative relation between the magnitude of the real exchange rate effects of domestic money and the variance of domestic money.[25] (Kormendi and Meguire conducted

[23] Since our peak effect estimates of B_m are all less than one in magnitude, the log transformation implies that the dependent variable and hence the intercepts in our third-stage regressions will be negative.

[24] The third-stage regressions using bilateral exchange rates that are reported include 29 of the 31 countries in Table 3. Argentina, with its extreme estimate for σ^2_m, and Iceland, with its extreme estimate for σ^2_y are excluded. The regressions using multilateral exchange rates also exclude Israel, Korea, Mexico and Singapore, for which no such data were available.

[25] As Kormendi and Meguire (1984) point out, measurement error in our estimates of σ^2_m and B_m may create problems for the interpretation of the subsequent cross-regime tests. In fact, their analysis implies that if the first-stage measurement error is relatively constant across countries and the cross-country variation in the estimates of σ^2_m is primarily attributable to variations in the true variances, then the results of the third stage are biased *against* finding a negative relation between the real exchange rate response to money shocks and unanticipated monetary variance, as we hypothesize.

On the other hand, it should be pointed out that if most of the cross-country variation in the estimated variance of unanticipated money is attributable to variations in the degree of measurement error across countries, there may be bias towards finding a negative relation. We do not view such a cross-country measurement bias as a likely explanation for our results. Kormendi and Meguire have shown on the basis of Monte Carlo simulations that in order for this bias to become severe enough to have a significant effect on cross-country inferences, the observed cross-country variation in the estimated variance of unanticipated money would have to be almost entirely attributable to cross-country variation in measurement error.

Table 4
Stage 3: Cross-country and cross-regime results–explaining the real bilateral exchange rate effects of domestic money supply shocks ($|B_{mj}|$)

Explanatory variables	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−1.44	−1.89	−1.50	−1.48	−1.33	−1.54
	(10.03)**	(8.74)**	(8.04)**	(7.08)**	(8.08)**	(5.14)**
σ_m^2	−0.36	−0.28	−0.35	−0.36	−0.30	−0.51
	(4.16)**	(3.28)**	(3.63)**	(3.78)**	(2.91)**	(3.14)**
DF		0.84				0.27
		(2.56)**				(0.56)
DN			0.10			
			(0.47)			
DB				0.05		
				(0.27)		
σ_y^2					−0.11	
					(1.24)	
$DF\sigma_m^2$						0.42
						(1.64)
$\bar{R}^2$	0.37	0.48	0.35	0.35	0.38	0.51
D.W.	1.86	2.14	1.92	1.91	2.04	2.07
Q-m.s.l.	0.10	0.47	0.11	0.11	0.16	0.64

Notes:

The dependent variable is the peak sum of coefficients on contemporaneous shocks and up to six lags. Dependent variable and variances of money shocks (σ_m^2) and of output shocks (σ_y^2) are in logs. *DF*, *DN*, and *DB* denote the continuous, narrow IMF-based definition, and broad IMF-based definition measures of exchange rate flexibility, respectively, from Table 1.

t-statistics in parentheses; *denotes significance at the 0.10 percent level; **denotes significance at the 0.05 percent level. *Q*-m.s.l. denotes marginal significance level for *Q* test that residual autocorrelations are zero.

their tests using output data.) This negative relation holds when our various exchange rate flexibility measures are added, as reported in columns (2), (3), and (4).

Table 4 also indicates that our continuous measure of exchange rate flexibility (*DF*) has a significant impact on the real exchange rate effects of domestic money. With this measure, the results support the hypothesis that $a_2 > 0$, i.e. the real exchange rate effects of domestic money shocks are directly related to the degree of exchange rate flexibility. Exchange rate flexibility as measured by the narrow (*DN*) or broad (*DB*) IMF classifications have little effect.

As noted in Section 2, the result $a_2 > 0$ is also consistent with Mundell–

Table 5
Stage 3: Cross-country and cross-regime results–explaining the real multilateral exchange rate effects of domestic money supply shocks ($|B_{mj}|$)

Explanatory variables	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−1.89 (7.79)**	−2.48 (11.33)**	−1.83 (5.37)**	−1.96 (4.92)**	−1.92 (5.61)**	−2.41 (8.75)**
σ_m^2	−0.35 (2.22)**	−0.29 (2.50)**	−0.37 (2.13)**	−0.33 (1.83)*	−0.37 (1.71)*	−0.34 (2.01)*
DF		2.45 (4.58)**				2.25 (3.02)**
DN			−0.10 (0.28)			
DB				0.09 (0.23)		
σ_y^2					0.03 (0.13)	
$DF\sigma_m^2$						0.14 (0.38)
$\bar{R}^2$	0.14	0.54	0.10	0.10	0.10	9.52
D.W.	1.87	2.19	1.88	1.86	1.86	2.23
Q-m.s.l.	0.72	0.50	0.74	0.74	0.73	0.37

Notes:

The dependent variable is the peak sum of coefficients on contemporaneous shocks and up to six lags. Dependent variable and variances of money shocks (σ_m^2) and of output shocks (σ_y^2) are in logs. *DF*, *DN*, and *DB* denote the continuous, narrow IMF-based definition, and broad IMF-based definition measures of exchange rate flexibility, respectively, from Table 1.

t-statistics in parentheses; *denotes significance at the 0.10 percent level; **denotes significance at the 0.05 percent level. *Q*-m.s.l. denotes marginal significance level for *Q* test that residual autocorrelations are zero.

Fleming type models where the real effects of monetary shocks are attributable to price rigidities. However, the strong significance of our information-based measure of exchange rate flexibility, *DF*, in combination with the insignificant results for the alternative measures, *DN* and *DF*, lends credence to our explanation for the stronger effects of monetary shocks under flexible exchange rates.

Table 4 also shows evidence on the remaining hypotheses that (iii) the variance of domestic real shocks and the magnitude of the real exchange rate response to domestic money shocks are positively related ($a_3 > 0$), and (iv) the effect of an increase in the variance of domestic money shocks on the magnitude of the real exchange rate response to domestic money is stronger the more flexible is the exchange rate regime ($a_4 < 0$). As reported in column (5), the variance of domestic real shocks has a negative, though insignificant, impact on the magnitude of the real exchange rate effects of domestic money, regardless of the exchange rate regime. This result may stem from the inappropriateness of using industrial production to measure 'real' shocks.

The negative relation between the real effects of domestic money and its variance, predicted to be stronger under flexible rates than under fixed rates, is found to be somewhat weaker in the flexible case; i.e. our estimate of a_4 reported in column (6) is positive, though only bordering on significance at the 0.10 level. However, the significance of this result is weakened by the apparent multicollinearity between DF and the interaction term $DF\sigma^2_m$. The comparison of columns (2) and (6) indicates a substantial decline in the magnitude and significance of the coefficient on DF following the addition of $DF\sigma^2_m$ to the regression. Moreover, in an unreported regression with $DF\sigma^2_m$ and σ^2_m, but not DF, the interaction term is positive and strongly significant. Thus the high correlation of DF and $DF\sigma^2_m$ (the simple correlation between the two variables is 0.68) implies that the positive coefficient on the latter merely reflects the positive effect of DF.

Table 5 reports tests of our third-stage hypotheses using multilateral real exchange rates; the results are broadly similar to those based on bilateral real exchange rates discussed above. The magnitude of the real exchange effects of domestic money depends negatively on the variance of domestic money and positively on the continuous measure of exchange rate flexibility. The effects of the IMF classifications of exchange rate flexibility, the variance of domestic real shocks, and the interaction term $DF\sigma^2_m$ are all insignificant.

5. Conclusions

In a rational expectations model of the real exchange rate effects of monetary disturbances, both the domestic monetary regime (the variance of domestic money shocks and real shocks) and the exchange rate regime (exchange rate flexibility or informativeness) should affect the magnitude of the real exchange rate effects. Our results indicate that the exchange rate's informativeness is important for determining the magnitudes of the real exchange rate effects of domestic money. The real exchange rate response to domestic money varies positively with our continuous information-based measure of exchange rate flexibility. This measure has superior explanatory power than conventional measures.

Our study also corroborates the results of Kormendi and Meguire (1984). The strong negative relation between the output effects of money shocks and the variance of money shocks also emerges in our study of the real exchange rate effects of domestic money shocks. We find strong evidence of a negative relation between the magnitude of the real exchange rate effects of domestic money shocks and the variance of unanticipated domestic money. Although the finding of stronger real effects of monetary shocks under flexible exchange rates is consistent with models with price rigidities, these alternative models typically do not yield hypotheses concerning the variances of the underlying shocks.

Overall, our results indicate that the degree of exchange rate flexibility is an important determinant of the real exchange rate effects of money shocks. The results lend empirical support to the class of rational expectations models tested, highlight the role of information, and underline the value of cross-regime testing.

Data appendix

Monthly data for all countries were obtained from IMF *International Financial Statistics*. The bilateral exchange rate was measured by the end-of-month foreign currency per dollar rate (line ae). The domestically-created money supply was defined as reserve money (line 14) minus total reserves, excluding gold (line 11.d), the latter converted to national currency units using the end-of-month exchange rate. The bilateral real exchange rate was defined as the end-of-month exchange rate times the U.S. wholesale price index divided by the domestic wholesale price index (line 63). Nominal and real multilateral exchange rate indices were obtained from lines n ec and r ec, respectively. Industrial production was obtained from line 66. The sample ranges for the variables for each country are given in Table 2 of the text, augmented by seven initial observations to allow for lagged variables and first differencing of the data. Exceptions to these procedures are noted below.

Exchange rate data for Mexico were obtained by splicing together two end-of-month series. The principal rate (line we) representing the middle rate between average buying and selling rates reported by main commercial banks, was used through 1982.11. (For the period after 1982.11, the principal rate series refers to the controlled exchange market rate.) The secondary rate (line xe), representing the free market rate introduced in 1982.12, was used for the period 1982.12–1991.11, when the dual exchange rate market was abolished.

Reserve money was unavailable for Ecuador for 1977.1. We interpolated this figure from the average of the previous and following months. The reserve money series for Colombia had missing values for 1983.1–1983.2, 1983.4–1983.5, 1985.7–1985.8, 1985.11, 1986.1, 1986.7 and were estimated ex ante from an ARIMA(1,1,0) process fit over the period 1973.8–1982.12. Reserve money for Bolivia was only available quarterly for the period 1988.Q1–1988.Q4. The missing monthly observations for this period were interpolated using a distribution procedure provided by the econometric software package RATS. This procedure assumes a simple time-series process (specified to be a random walk) for the observable quarterly data.

Owing to the unavailability of monthly wholesale price index numbers for Bolivia, Iceland, and Malaysia, the consumer price index (line 64) was used instead. Consumer prices for Iceland were only available quarterly for the period prior to 1984.1, and were distributed as above for the missing monthly observations. In the case of France and Portugal, wholesale price data were only available through 1985.12 and 1986.12, respectively; again consumer prices were used for the entire sample period.

Industrial production data for Australia and Portugal were obtained from the OECD *Main Economic Indicators*. For Argentina, Singapore, and Switzerland, monthly industrial production data were interpolated from quarterly data using the distribution procedure described above. For Bolivia, monthly values were missing for the 1986.10–1989.10, and quarterly values were missing for 1988.Q2–1988.Q4. We estimated an ARIMA(1,1,0) process for the quarterly values over the period 1973.Q3–1988.Q1 to forecast the missing quarters. These quarterly figures were then distributed as above for the missing monthly observations.

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