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REAL SHOCKS AND REAL EXCHANGE RATES IN REALLY LONG-TERM DATA

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Abstract

There is little consensus concerning the sources of fluctuations in real exchange rates. In this paper I assess the nature of the shocks that drive the real U.S. dollar-U.K. pound exchange rate, analyzing 130 years of data. I first show that wars, which are examples of a (transitory) real shock, are significant. I next use an alternative empirical approach, in which I identify various types of real and nominal shocks. I find that output shocks and monetary shocks account for approximately the same percentage of the variance of the real exchange rate over short horizons. The monetary shock is decomposed into monetary base and money multiplier shocks and the output shock is decomposed into supply and demand shocks. Essentially all of the effect of the combined output shock is due to the demand shock. The effect of the monetary shock is accounted for by both money multiplier shocks and monetary base shocks in roughly equal amounts. Thus, the paper suggests that the contributions of real and monetary shocks are roughly equal overall, while shedding light on the nature of those shocks.

Real Shocks and Real Exchange Rates in Really Long-Term Data

John H. Rogers[1]

In this paper I provide empirical evidence on the nature of the shocks that drive the real exchange rate. I analyze approximately 130 years of data from the United States and United Kingdom. My motivation for focusing on the long-term pound/dollar exchange rate is that a long span of data is desirable on both statistical and economic grounds. For these countries a long, uninterrupted span of data on macroeconomic variables is available.

My strategy and findings are as follows. First, using a cointegration framework, I present evidence that transitory events involving real shocks - wars in particular - have a significant effect in moving nominal exchange rates and relative price levels away from the long run path implied by PPP. Second, using a vector autoregression framework, I identify various types of real and nominal shocks. Estimates of the forecast error variance of the real exchange rate from several structural VARs indicate that aggregate output shocks and monetary shocks, the latter of which are restricted to having zero effect in the long run, account for approximately the same percentage of the variance of the real exchange rate over short horizons. Fiscal shocks are also important. The monetary shock is decomposed into monetary base and money multiplier shocks and the output shock is decomposed into supply (productivity) and demand (preference) shocks. Essentially all of the effect on the real exchange rate from the combined output shock is due to the demand shock, while the effect of the monetary shock is accounted for by both money multiplier shocks and monetary base shocks in roughly equal amounts.

My results are interpreted in light of the competing claims from the "equilibrium" and "sticky-price" views of real exchange rate determination. Finally, I compare my results to those of recent studies. The encompassing analysis suggests that the choice of model is quite important for making inferences about the sources of real exchange rate movements.[2]

1. Related Literature

An early strand of the literature seeking to explain the factors underlying real exchange rate movements linked statistical evidence on real exchange rates, nominal exchange rates, prices, and other macroeconomic variables with theoretical models of the exchange rate. Adler and Lehman (1983) and Stockman (1983) present evidence on real exchange rate changes for over 40 countries during both fixed and flexible exchange rate periods. Mussa (1986) examines relative prices and nominal exchange rates for 16 industrialized countries. Two empirical regularities are documented by these studies. First, real exchange rates are more volatile under flexible exchange rates than fixed rates. Second, there is a high correlation between real exchange rate and nominal exchange rate changes during both fixed and flexible exchange rate regimes. Although there is a large consensus on what the stylized facts are, there is much less agreement on the issue of what types of models are supported by these facts.[3]

More recent studies have carried these ideas further. Campbell and Clarida (1987) and Meese and Rogoff (1988) both reject the hypothesis that there is a significant statistical link between real exchange rates and real interest rate differentials, and relate their findings to sticky-price theories of

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[2]Other studies using data only on the real exchange rate do not assess the sources of real exchange rate fluctuations from the perspective of a general-equilibrium macro model [see Grilli and Kaminsky (1991), Diebold, Husted, and Rush (1991), and Lothian and Taylor (1994); Kaminsky and Klein (1994) is a recent exception]. Hence my encompassing analysis relates to Clarida and Gali (1994).

[3]The task of discriminating between models is complicated in part by a "third" stylized fact: the real exchange rate is apparently the only macroeconomic variable whose behavior -- in terms of its own stochastic properties or its covariance with other aggregates -- changes significantly across exchange rate regimes. See Baxter and Stockman (1989) and Flood and Rose (1993). I do not address the neutrality of macroeconomic aggregates with respect to the exchange rate regime in this paper, but rather focus on how shocks to macroeconomic variables affect the real exchange rate.

the exchange rate. Edison and Pauls (1993) fail to find evidence of cointegration between real exchange rates and real interest rate differentials. Baxter (1994) finds evidence that real interest rate differentials are related only to the temporary component of the the real exchange rate, but that most movements in the real exchange rate are due to changes in the permanent component. Engel (1993) shows that the variance of the common-currency price of the same good across borders is larger than the variance of the relative price of different goods within a country, and relates this to the debate over sticky prices. In a similar vein, Engel and Rogers (1994) examine goods prices in U.S. and Canadian cities, in order to assess the relative importance of transportation costs, non-tradeables, and sticky prices in explaining deviations from the law of one price.[4]

An alternative approach, which is more related to this paper, is to assess the nature of various shocks to the real exchange rate and measure their relative importance. Huizinga (1987) examines monthly data on the real exchange rates of the major industrial countries from 1974-86, and finds that on average the variance of the permanent component accounts for 58% of the variance of the actual change in the real exchange rate. Baxter (1994) also analyzes the temporary and permanent components of real exchange rates, using a multivariate Beveridge-Nelson (1981) decomposition, and finds that most movements are due to changes in the permanent component.

Lastrapes (1992) and Clarida and Gali (1994) follow the methodology of Blanchard and Quah (1989), as I do in the final part of this paper, to estimate the relative contribution of various shocks to explaining the variance of real exchange rates between the United States and Canada, Germany, Japan, the United Kingdom, and Italy. Clarida and Gali (1994) estimate the system (Δy, Δr, and π),

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[4]Several papers, including Frankel (1986), Baillie and Selover (1987), Edison (1987), Mark (1990), Kim (1990), and Rogers and Jenkins (1994), have tested for unit roots in real exchange rates, or cointegration between nominal exchange rates and relative prices. The conclusion that emerges from these studies, and Diebold, Husted, and Rush (1991), who apply a fractional difference estimator to several real exchange rates during the gold standard, is that real exchange rates exhibit very slow mean reversion or "long-memory". There is some consensus in the literature that the half-life of a shock that moves the exchange rate away from parity is about 3 or 4 years. See Stockman (1988), Frankel (1990), and Froot and Rogoff (1994) for further discussion about how such statistical evidence may (or may not) be relevant for the debate about equilibrium versus sticky-price models.

where y is the home-foreign output growth differential, r is the real exchange rate and π is the inflation differential, with corresponding shocks labelled aggregate supply, aggregate demand, and nominal shocks. They find that nominal shocks account for between 34% and 47% of the variance of real exchange rate changes (over all horizons) for Germany and Japan, but no more than 2.8% for the United Kingdom and Canada. Essentially all of the rest is accounted for by aggregate demand shocks, as aggregate supply shocks are negligible at all horizons for all countries. Estimating a smaller-scale VAR, Lastrapes finds that the contribution of the real shock ranges from 63% to 94%, depending on the exchange rate.[5]

2. The Data

I use annual observations from both the United Kingdom and United States on the following:[6]

S = log of the nominal exchange rate in pounds per dollar, 1859-1992
P1 = log of the wholesale price index, U.K. relative to U.S., 1791-1992
P2 = log of the GNP deflator, U.K. relative to U.S., 1889-1992
r1 = log of the real exchange rate, (S - P1), 1859-1992
r2 = log of the real exchange rate, (S - P2), 1889-1992
y = real GNP; mill. of 1985 pounds (U.K. 1830-1992), and bill. of 1987 dollars (U.S., 1889-1992)
g = real government expenditure (from the national income accounts); millions of 1985 pounds (U.K., 1830-1992), and billions of 1987 dollars (U.S., 1889-1992)
M = nominal M2 money stock in millions of pounds (U.K., 1871-1992), and billions of dollars (U.S., 1867-1992)
H = nominal monetary base in millions of pounds (U.K., 1871-1992), and billions of dollars (U.S., 1867-1992)

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[5]Kaminsky and Klein (1994) estimate a VAR with contemporaneous restrictions to investigate the effects of aggregate supply and demand shocks, shocks to money demand, shocks to expected devaluation, and three different types of fiscal shocks on the real pound-dollar exchange rate during the gold standard. At long horizons, the contribution to the variance of the real exchange rate is approximately 35% for the supply shock, 25% for the shock to expected devaluation, and about 13% each for the money demand and budget deficit shocks. Eichenbaum and Evans (1994) also estimate VARs with contemporaneous zero restrictions, in order to investigate the effects of shocks to U.S. monetary policy on the bilateral exchange rates with Japan, Germany, Italy, the U.K., and France in the post-Bretton Woods era. Although they find that exogenous shocks to U.S. monetary policy contributed to the variability of exchange rates, their variance decompositions indicate that other shocks also explain a lot of these movements.

[6]The sources for the data on prices, output, and government spending are Mitchell (1988) for the United Kingdom and Kendrick (1961) for the United States. The exchange rate series is also taken from Mitchell (1988), while the money supply and monetary base data are taken from Friedman and Schwartz (1982). Updates are from Britain's Central Statistical Office (1992), and various issues of the Survey of Current Business for the United States. Additional detail is available from the author.

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P_N = price of non-traded goods, as proxied by the government expenditure deflator; 1985 = 1.00
for the U.K. (1830-1992), and 1987 = 1.00 for the U.S. (1889-1992)

m = log of the real money stock, M/P_N.

Notice that I use two different measures of the real exchange rate, r1 and r2, which are based on the two different price indexes, wholesale prices (P1) and the GNP deflator (P2). An increase in the exchange rate (real or nominal) is a depreciation of the pound and an appreciation of the dollar.

Figure 1 contains plots of the real and nominal exchange rate series. The stylized facts discussed above seem to be verified. The real exchange rate appears to be more volatile during the periods of flexible exchange rates, and there is clearly a high correlation between real exchange rate and nominal exchange rate changes during both fixed and flexible exchange rate regimes.

In addition, the data suggests that there is mean reversion in the real exchange rate over the full sample, consistent with the recent view in the literature on PPP. However, this inference is not free of uncertainty, even in this long span of data.[7] In the remainder of the paper, I explicitly take into account this uncertainty by using two different empirical methodologies. Notice that either:

(i) r is stationary; or (ii) r contains a unit root, and is stationary in first-differences.

A presumption that the "truth" is (i) suggests undertaking the cointegration analysis of section 3. There I use Stock and Watson's (1993) dynamic OLS procedure to perform inference on the estimated cointegrating relationship between S and P, subject to a number of different configurations for the

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[7] I ran Augmented Dickey-Fuller (ADF) τ_τ and τ_μ tests for a unit root in the exchange rate and price ratio series. The tests indicate that the nominal exchange rate, S, and relative price ratios, P1 and P2, are each I(1) processes. However, the unit root null hypothesis for the real exchange rates r1 and r2 can be rejected. For example, the τ_τ test statistics for S, P1, and r1 are -2.30, -1.71, and -3.49, respectively (the 5% critical value is -3.43). These results are somewhat borderline, however. First, consider that the 95% confidence interval for the largest root in r1 or r2, estimated using Stock's (1991) procedure, contains the value 1.0. Second, Johansen's (1991) trace test for cointegration fails to reject the null hypothesis of zero cointegration vectors in systems containing the nominal exchange rate and prices (either separately or as the ratio P1 or P2). For the system containing S and each country's wholesale price index, for example, the test statistic for the null hypothesis of zero cointegration vectors is 25.8, which is less than the 10% critical value of 26.8. However, the Johansen test results are somewhat at odds with Engle and Granger's (1987) residual-based tests. For example, an ADF test for a unit root in the residual from a regression of S on P1 produces a t-statistic of -3.51, which implies that these two series are cointegrated. The ambiguity of these tests motivates the "two-track" analysis of the real exchange rate I follow in sections 3 and 4.

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short-run dynamics. These different configurations include allowing the short-run dynamics to differ during war and non-war years. This analysis sheds light on the issue of whether or not transitory events have a significant effect in moving nominal exchange rates and relative price levels away from the long run path implied by PPP, and if they do, what is the nature of those events. On the other hand, one might assign a higher probability that (ii) is true. Working under this alternative prior, I estimate different structural vector autoregression models in section 4.[8] The combined evidence gives us a good picture of the sources of real exchange rate movements.

3. Cointegration Analysis

As noted above, rejecting the null hypothesis that the real exchange rate is a unit root (or that S and P are not cointegrated) suggests undertaking an analysis of the cointegrating relationship between S and P. In this section I use Stock and Watson's (1993) dynamic OLS procedure to estimate the cointegrating relationship, and perform hypothesis testing on the vectors. I use this framework as the first of my two ways of assesing the nature of the shocks that drive the real dollar-pound exchange rate.

Before discussing results, I describe the Stock-Watson method. Consider an n-dimensional vector of I(1) variables, y_t, whose first-differences are stationary stochastic processes with zero mean. If there are p cointegrating vectors among the n elements of y_t, the system can be written:

$$\Delta y_{1t} = u_{1t} \quad , \quad y_{2t} = \theta y_{1t} + u_{2t} \tag{1}$$

where y_{1t} is (n-p)-by-1, y_{2t} is p-by-1, θ is p-by-(n-p), and u_{1t} and u_{2t} are zero mean stationary processes. Stock and Watson show that with $u_t = (u_{1t}, u_{2t})'$ stationary and Gaussian, the above system has the following triangular representation:

$$\Delta y_{1t} = C_{11}(L)\epsilon_{1t} \quad , \quad y_{2t} = \theta y_{1t} + d(L)\Delta y_{1t} + C_{22}(L)\epsilon_{2t} \tag{2}$$

with ϵ_{1t}, ϵ_{2t} independent, and d(L) a matrix of two-sided lag polynomials.

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[8]I also assess the robustness of the VAR estimates to the assumption that the real exchange rate is trend-stationary instead of a unit root.

Stock and Watson demonstrate that under fairly general conditions the maximum likelihood estimate of θ is asymptotically equivalent to OLS. Stock and Watson call this the dynamic OLS (DOLS) estimator. Moreover, provided the moving average nature of the error term in the equation for y_{2t} is accounted for, and the standard error is adjusted for serial correlation, conventional inference can be carried out for hypothesis tests on the cointegrating vectors.

To examine the long-run relationship between the nominal exchange rate and relative price ratio, I estimate the following equations:

$$S_t = \alpha_0 + \alpha_1 P1_t + d_1(L)\Delta P1_t + u_t \tag{3a}$$

$$S_t = \beta_0 + \beta_1 P2_t + d_2(L)\Delta P2_t + v_t \tag{3b}$$

where, as noted above, the terms involving $d(L)$ imply that both leads and lags of the variable are included in the regression. As a further check, I also run the tests with the dependent and independent variables switched. Using White's (1980) correction for heteroscedasticity, I then test the relevant hypotheses on the individual coefficients ($\alpha_1 = 1$ and $\beta_1 = 1$).

Results are reported in table 1. Examining the rows where no dummy variables are included in the regression (indicated by a "No Break" in column 4), we reject the null hypothesis of a unit coefficient in all cases. In the 1859-1992 sample period, for example, the estimated cointegrating vector for (S, P1) is (1, -1.10), with a robust standard error of .044.

One possible explanation for the deviations of the nominal exchange rate and relative price ratio from the long-run path implied by PPP centers on the effects that transitory events may have on these variables.[9] That is, it may be that the long-run relationship between S and P2 is as PPP implies, but that large, transitory changes in prices or the exchange rate during, e.g., major war years, have had persistent effects.

One way to test for such effects is to allow the constant and short-run dynamics to be different in years of a major war. Hence, I estimate for both P1 and P2 regressions of the form:

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[9]Froot and Rogoff (1994, section 2.4) discuss several reasons, in addition to that I present here, why one might not find a (1, -1) cointegrating vector.

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$$S_t = \alpha_0 + \alpha_0'DUM + \alpha_1 P_t + d_1(L)\Delta P_t + d_1'(L)(DUM*\Delta P_t) + u_t \qquad (4a)$$

$$P_t = \beta_0 + \beta_0'DUM + \beta_1 S_t + d_2(L)\Delta S_t + d_2'(L)(DUM*\Delta S_t) + v_t \qquad (4b)$$

where DUM is a dummy variable for (major) "war" versus "non-war" years. War years for the U.S. are 1861-1865, 1917-1919, and 1941-1946, and for the U.K. 1899-1902, 1914-1919, and 1939-1946.

In principle, there are at least two reasons why wars may have the postulated effect. First, wars disrupt the flow of goods across borders, and so interfere with the arbitrage needed for the law of one price to hold. This will take place even when the U.S. and U.K. are fighting on the same side of the war. Second, wars necessitate large temporary fiscal expansions and destroy part of a nation's stock of physical and human capital, either of which could result in large and persistent changes in the real exchange rate. Both are examples of a real shock.

In estimating equations (4), I consider two war/non-war dummies: (i) "War/Non-War A", which is unity for any year in which either the U.S. and/or U.K. was at war and zero during non-war years, and (ii) "War/Non-War B", which is equal to the ratio of military personnel to the general population in major war years (the ratio for the U.K. less that of the U.S.) and zero for non-war years. Thus, War/Non-War A treats all wars as alike, thereby capturing the "trade disruption" channel, while War/Non-War B treats each war (and each war year) as different, and hence captures the "fiscal expansion" channel.[10]

The results are also displayed in table 1. The null hypothesis of a unit coefficient in the cointegrating vector is no longer rejected in seven of the eight cases in which the war/non-war dummy variables are included. For example, without allowing the short-run dynamic behavior of

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[10]Either dummy variable could be capturing the "destruction of human and physical capital" channel. This evidence is reduced-form in nature. A structural model is estimated in section 4. Thus, in this section I am agnostic about the precise channels through which events, like wars or changes in exchange rate regimes, cause deviations from PPP. Typically, both countries were fighting a war at the same time. Changes in the exchange rate regime occured essentially at the same time for both countries. However, it is impossible for both countries to simultaneously experience, say, a real depreciation due to the disruption of trade. Hence, one ought to think about the channels outlined above as depending on the effect that wars have on one economy relative to the other. For instance, it is obvious from figure 2a that the wars had a much larger expansionary effect on government size in Britain than the United States.

relative prices to be different between war and non-war years, the estimated long-run relationship between S and P2 is 1.20, which is significantly greater than unity. Once the effect of the wars is taken into account with the dummy War/Non-War A, the estimated long-run coefficient becomes insignificantly different from unity at 1.08. Finally, notice from table 1 that the same result is obtained when the DOLS regression is run with P as the dependent variable.

The remaining results in table 1 are estimates of equations (4a) and (4b) with the dummy variable defined to be unity in years of floating exchange rates (1914, 1919-25, 1931-40, 1949, and 1973-present) and zero the rest of the time. These results are similar to the estimates of the equations without dummy variables: in each case, the null of a (1, -1) cointegrating vector is rejected. Thus, allowing for different behavior of the short-run dynamics during fixed versus flexible exchange rate regimes has no effect on the long-run relationship between nominal exchange rates and relative prices.

The purpose of this section is to identify shocks or events -- transitory ones in this case -- that could be the source of deviations of the nominal exchange rate and relative price ratio from the path implied by PPP. The results of this section are consistent with evidence expressed elsewhere in the literature in support of PPP as a long-run concept. It is shown that, by allowing the short-run dynamics to be different during wars, which are examples of a real shock, estimates of the long-run relationship between the nominal exchange rate and relative price ratio are significantly closer to the implied value of unity. However, there is no apparent effect on the long-run relationship when I allow the short-run dynamics to differ across periods of fixed and floating exchange rate regimes.

The choice of empirical methodology used in this section was dictated by the prior that S and P are cointegrated. In the next section I use a VAR framework to identify various types of real and nominal shocks and assess their relative contribution to real exchange rate variability.

4. Identification and Derivation of the Structural Vector Autoregression Model

In this section I estimate a small-scale macro model which includes the real exchange rate and fiscal and monetary variables. I begin by briefly discussing the econometric procedure, then give a

justification for interpreting the structural shocks. This highlights the long-run restrictions used to estimate these shocks, as in Blanchard and Quah (1989). Because most macroeconomic debates are about short-run phenomena, it is generally less controversial to use long-run rather than short-run restrictions. It is also true though, that the use of long-term data opens up the possibility of various types of regime changes. Hence, it is necessary to discuss how the identifying restrictions are robust to different assumptions concerning the exchange rate regime and capital mobility.

A. The Estimated Model

Consider a vector of stationary variables X and a vector of structural shocks ϵ. The structural model can be compactly written,

$$X_t = C(L)\epsilon_t \tag{5}$$

where C is a non-singular matrix of coefficients, and L denotes the lag operator. A reduced form of the structural system that can be estimated is given by,

$$\Gamma(L)\Delta X_t = \Phi X_{t-1} + \epsilon_t^*. \tag{6}$$

Assuming that the long-run moving average coefficient matrix, C(1), is lower-triangular and that the elements of ϵ are mutually uncorrelated, one can follow the procedure developed by Ahmed, Ickes, Wang, and Yoo (1993) to retrieve the structural coefficients from the reduced form.

Let the vector of stationary variables $X = \{\Delta(g/y), \Delta(y), \Delta(r), \Delta(mm), \Delta(h)\}$, and the vector of shocks $\epsilon = \{\epsilon^g, \epsilon^S, \epsilon^D, \epsilon^{mm}, \epsilon^h\}$.[11] The notation is as used in section 2: g denotes the real value of government spending, y is the log of real output, r is log of the real exchange rate, mm is the log of the money multiplier (the ratio of nominal M2 to the monetary base), and h is the log of the real monetary base. The structural shocks are, in order, permanent disturbances to government size, output, the real exchange rate, the money multiplier, and real monetary base. These are interpreted as fundamental shocks based on the theoretical model laid out in the appendix. The way in which the

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[11]In this section I work primarily with the prior that the real exchange rate contains a unit root, but also consider the case in which (g/y) and r are stationary.

transformed variables are assumed to be related in the long run (i.e., ignoring the lagged dynamic terms in the VAR) can be written:

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\begin{bmatrix} \Delta(g/y) \\ \Delta(y) \\ \Delta(r) \\ \Delta(mm) \\ \Delta(h) \end{bmatrix} = \begin{bmatrix} \bar{g} \\ \bar{y} \\ \bar{r} \\ \bar{m} \\ \bar{h} \end{bmatrix} + \begin{bmatrix} c_{11} & 0 & 0 & 0 & 0 \\ c_{21} & c_{22} & 0 & 0 & 0 \\ c_{31} & c_{32} & c_{33} & 0 & 0 \\ c_{41} & c_{42} & c_{43} & c_{44} & 0 \\ c_{51} & c_{52} & c_{53} & c_{54} & c_{55} \end{bmatrix} \begin{bmatrix} \epsilon^g \\ \epsilon^S \\ \epsilon^D \\ \epsilon^{mm} \\ \epsilon^h \end{bmatrix} \tag{7}
$$

where $\bar{g}$, $\bar{y}$, $\bar{r}$, $\bar{m}$, and $\bar{h}$ are constant and independent of the structural shocks. In the estimation, I use the value of the U.K. variable less that of the U.S. for (g/y), y, mm, and h. These variables are depicted in Figure 2, both in levels and first-differences.[12]

A general justification for the lower triangularity of the model is as follows. First, the assumption that the long-run share of government spending in total output is exogenous implies the zero-restrictions in the top row. Second, assuming the classical dichotomy between real and monetary variables holds in the long run implies that for rows 1, 2, and 3, the elements in columns 4 and 5 are zero. Third, the assumption of a long-run unit elasticity of the money supply with respect to the monetary base justifies the restriction that $c_{45}=0$.[13] The final zero-restriction, in row two and

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[12]The plots suggest that my prior of each series being I(1) is not implausible. Formal tests for a unit root do not suggest clearly that this prior can be rejected. For example, using a lag length of 5, the ADF τ_τ test statistic for the variables (g/y), y, mm, and h are -3.33, -1.19, -2.54 and -2.56, respectively. These are all less than the 5% critical value of -3.43. Note that for the government size differential, which is bounded by -1 and 1, this result is borderline. Indeed, when an optimal lag length is chosen using the AIC, the ADF τ_τ test statistic for the four variables are -3.85, -1.20, -2.40 and -2.05, implying a rejection in the case of (g/y). Also, the test result for the government size variable is sensitive to the choice of deflator. Furthermore, consider Rogoff (1992), who examines the ratio of government consumption to GNP for Japan relative to the U.S. and Germany relative to the U.S. from 1975-90. He finds a unit root in the former but not the latter. I retain my unit root prior for most of the VAR analysis, but also estimate VARs with government size assumed to be stationary in levels as a check for robustness.

[13]This assumption does not imply that the money multiplier is constant in the long run. On the contrary, through ϵ^{mm} the money multiplier may change for any of several reasons: (i) portfolio demand shocks (exogenous changes in the currency-deposits ratio), (ii) changes in the reserves-

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column three of the C(1) matrix, implies that the preference shock (for traded goods relative to non-traded goods) has no effect on output in the long run.

Hence, the assumptions that (i) C(1) is lower triangular, and (ii) the shocks are orthogonal, enables me to identify the fundamental disturbances. The expected signs of the elements of the C(1) matrix, and the corresponding interpretation of the shocks, are discussed in the two-good, two-sector theoretical model which serves as a reference point for interpreting the empirical work. In order to save space, the model is placed in the appendix. Note that the method of identification imposes no restrictions on the short-run movements of the variables, which are instead determined by the data. Furthermore, identifying restrictions such as long-run monetary neutrality and exogenous government size seem to be plausible for different exchange rate regimes and degrees of capital mobility.

Before turning to the estimation results, it is worthwhile discussing some of the critiques of this empirical approach. First, as pointed out by Lippi and Reichlin (1993), we must assume that the MA representation is fundamental since otherwise the structural VAR methodology would be improper. Blanchard and Quah (1993) emphasize that this assumption is required in standard macroeconometric work and is made implicitly in most time-series empirical studies. In other words, ruling out nonfundamental MA representations is common in practice.

Second, Faust and Leeper (1994) critically address the validity of imposing the long-run restrictions and the implicit aggregation of shocks in the Blanchard-Quah methodolgy. Both concerns motivate our robustness checks. Because we are imposing infinite order (long-run) restrictions on a finite order VAR, the estimation results are correct only if the estimated reduced-form VAR is the correct representation. In this sense, imposing long-run restrictions is not necessarily more robust to possible misspecification than imposing short-run restrictions as in Bernanke (1986). Also, as pointed

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deposits ratio brought about by new bank legislation or by changes in reserve requirements, or (iii) real aggregate demand changes that have no long-run effect on government size, output, or the relative price of non-tradeables, but get transmitted to broad money through the dependence of the money supply on interest rates. Ahmed (1994) makes the same identifying assumption in a closed-economy context, and provides an expanded discussion of these shocks.

out by Blanchard and Quah (1989) and elaborated upon by Faust and Leeper (1994), if one identified structural shock indeed consists of two independent shocks, then the methodology is valid only if the underlying macroeconomic variables respond to the two shocks in the same directions.

In practice one can consider only a limited set of variables identifying a limited set of shocks. Because many theoretical models suggest that fiscal and output shocks lead to opposite effects on the real exchange rate (see Froot and Rogoff (1991)), it is crucial to separate the two. Of course, one may argue that the "output" shock may consist of both demand and supply shocks and the "monetary" shock may contain both money demand and money supply shocks (or shocks to inside money as well as to outside money). These disaggregated shocks will affect the real exchange rate differently. As a consequence, in the model I (i) identify a "preference" or "demand" shock separately from the output "supply" shock, based on the assumption that output is supply-determined in the long run, and (ii) decompose the monetary shock into monetary base and money multiplier shocks, under the assumption that shocks to the monetary base have a one-to-one effect on the money supply in the long-run. Finally, I relate the results to Clarida and Gali (1994), estimating their model on my data and my model on their data. The estimated models arguably exhaust the most important forces -- though not necessarily all forces -- underlying real exchange rate movements.

B. Results

Table 2 presents the results from estimating the VAR model (7).[14] Notice first, estimates of the elements of the C(1) matrix in table 2B. Of particular importance, both c_{31} and c_{32} are negative, implying that fiscal shocks and output supply (productivity) shocks lead to a long-run real appreciation. This effect would be found in several theoretical models. In terms of the two-sector, illustrative theoretical model used in this paper, the results imply that fiscal shocks are felt primarily in the non-traded goods sector (since a positive fiscal shock raises the relative price of non-tradeables), while productivity shocks have a relatively larger effect in the traded goods sector.

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[14]Johansen's (1991) trace test indicates the absence of cointegration in our five-variable systems (see table 2A), and thus implies that estimating a VAR in first-differences is appropriate.

According to the estimate of c_{33}, a postive demand shock, ϵ^D, leads to a long-run real depreciation. This is consistent with interpreting ϵ^D as a shift in preferences toward traded goods. Finally, note that c_{55}, the long run effect of ϵ^h on the real monetary base, is negative. This suggests interpreting ϵ^h as reflecting postive shocks to the growth rate of the monetary base.[15]

The impulse responses of r1 and Δ(r1) are displayed in figures 3a and 3b (available on request are results for the system containing r2; these are similar). Dashed lines represent plus and minus one standard deviation bands around the point estimates. There is a persistent response of the real exchange rate to a shock to either government size or productivity. Either shock leads to a real appreciation in the long run, a result that was seen in the estimates of c_{31} and c_{32}. The (traded goods) preference shock leads to a persistent real depreciation, as expected. The response of the real exchange rate to either the money multiplier or the monetary base shock is short-lived. A positive money multiplier shock leads to a real depreciation in the short run, which is consistent with the interpretation given to ϵ^{mm} below. Finally, following a shock to the monetary base the real exchange rate drops on impact, then immediately rises (see figure 3b). Relative to its initial level, the real exchange rate is depreciated significantly over horizons three through ten, after which there is a gradual return to its (unchanged) long-run level.

The variance decomposition (VDC) results are displayed in table 2C. Consider the estimates for Δ(r1) in the top panel. These indicate that the demand (preference) shock, the money multiplier shock, the monetary base shock, and -- to a lesser extent and at long horizons -- the fiscal shock are all significant. The monetary shocks combined account for approximately 45 percent of the variance of real exchange rate changes. The effect is split into money multiplier shocks and monetary base shocks in roughly equal amounts. Preference shocks account for over 35 percent, while the fiscal

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[15]Although the long-run estimates are of the correct sign, many are insignificant, especially those associated with the fiscal shock. This may be because the permanent component of the fiscal variable is small. That most of the changes in government size are due to temporary events like wars (whose effects were estimated in section 3) is suggested by the plots of the variables in Figure 2 and by the results of the unit roots tests on (g/y). Estimates of the model with stationary (g/y), undertaken below, provide evidence on the importance of this.

shock accounts for over 10 percent. The productivity shock is insignificant at all horizons, however. The VDC of $\Delta(r2)$ is very similar, the only noticeable difference being that in the model with r2 the contribution of ϵ^h (ϵ^{mm}) is almost 10 points higher (lower) than in the model with r1. The lower panel of table 2C contains the VDCs of the levels of the real exchange rate. Notice that the contribution of the monetary shocks goes to zero in the long run, which is of course an identifying assumption. At shorter horizons, the relative contributions of each shock are as described for $\Delta(r_j)$.

In light of these results, it is important to be very specific in interpreting the shocks. First, consider the types of disturbances in the real world that might correspond to ϵ^D. Recall that this is any disturbance that may have a long-run effect on the relative price of non-tradeables but not government size or output. For this long-term data set, important examples are wars and changes in trade policy. Large negative realizations of ϵ^D presumably resulted from the disruption of trade during the war years in the sample. In this sense, wars represent "involuntary" shifts in preferences away from traded goods. In addition, the histories of both countries, especially the United States, are filled with important changes in trade policy.[16]

Second, consider the monetary shocks. Exogenous disturbances to the real monetary base, ϵ^h, correspond most closely to monetary policy shocks. Possible interpretations of money multiplier

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[16]Notice that although wars have dramatic transitory effects on relative government size, they do not seem to have had much of a long-run effect according to figure 2. As for changes in trade policy, consider an overview of the U.S. legislation on external tariffs during the sample period, as described by Robertson (1973). In 1890, the McKinley Act raised the average U.S. tariff rate significantly, to 50%, which was slightly above the level during the Civil War. In 1897, tariffs were raised to almost 60%, except for raw materials and semi-finished products. In 1913, the Underwood-Simmons bill lowered tariff rates and increased the the list of items that were exempt from tariffs. This move toward freer trade left the average tariff rate at 25%. In 1921, an emergency tariff act was passed, and one year later, the Fordney-McCumber Act raised the average tariff rate to 33%. The notorious Hawley-Smoot Act, passed in 1930, raised the average tariff rate to 40%. Between 1934 and the end of WWII, the President was given expanded powers to negotiate tariffs. Hence, significant changes (mostly increases) in U.S. tariffs ocurred every 10 years or so in the first half of my sample. Beginning with the creation of the GATT in 1947, an era of freer trade was ushered in. A notable exception to this was a 10% surcharge on dutiable imports levied by Nixon in 1971. Tariffs are important in Kaminsky and Klein's (1994) analysis of the real exchange rate during the gold standard. Marquez (1994) discusses the importance of U.S. tariffs for estimating trade elasticities.

shocks, ϵ^{mm}, were discussed above (see footnote 13). The VDCs suggest that about half of the effect of monetary shocks on the real exchange rate is due to monetary policy shocks.[17]

VAR Estimates with Stationary Government Size and Real Exchange Rate

Acknowledging the uncertainty associated with the stochastic properties of some of the series, I examine specifications with government size and the real exchange rate assumed to be stationary (the real exchange rate is trend-stationary).[18] Dropping the unit root specification implies that both fiscal shocks and preference shocks are now modeled as temporary disturbances. The results, reported in the appendix to this paper, are very similar to those in table 2. All but 1 of the 30 coefficients in the estimated $C(1)$ matrices are different in sign from those in table 2B. The VDCs of the (detrended) real exchange rate indicate that monetary shocks contribute approximately 50 to 75 percent over short horizons. This is 10 to 35 percent larger than in the model with (g/y) and r modeled as unit roots. The fiscal and demand shocks, ϵ^g and ϵ^D, account for most of the rest. This is especially true of the long run, when the contribution of the monetary shocks goes to zero. The combined effect of these two shocks is slightly less than it was in the model with $\Delta(g/y)$ and $\Delta(r)$: aproximately 85% in the long run in table 2 and around 70% in the current model. Once again, the contribution of ϵ^S is negligible.[19]

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[17]There is a large debate in the literature about identifying monetary policy shocks (see Gordon and Leeper (1994) for a recent discussion). This work focuses mostly on closed-economy modeling issues, but it does suggest that the monetary base shock I identify is not necessarily the most appropriate proxy for policy shocks. This criticism is less importance in the present context, however, because my primary goal is to assess the relative contribution of real versus nominal shocks to the real exchange rate, as opposed to measuring, e.g., a liquidity effect on interest rates.

[18]An alternative would be to estimate a vector error correction model (VECM) with log(S) and log(P) added individually, and the cointegrating relationship between them imposed. The VECM methodology is developed in King, Plosser, Stock, and Watson (1991), whose application has been extended by Ahmed and Rogers (1994).

[19]Three additional checks for robustness are also worth noting. First, I estimated the model with alternative deflators for the real monetary base: the GNP deflator and the wholesale price index (the government spending deflator (P_N) is appropriate according to the theoretical model). The results are largely insensitive to this alteration; each estimate is within one standard deviation of the corresponding estimate reported in table 2C. Second, in order to examine the potential problem associated with aggregation of shocks, I considered estimates of a set of three-by-three systems, in

Implications for Existing Literature

Do the VDC results provide any useful information beyond the stylized facts discussed in the introduction? As I argue next, it is desirable to go beyond establishing stylized facts in order to shed light on the sources of movements in real exchange rates. The results above represent a basis for discriminating between different views in the debate.

First, it has traditionally been assumed that the high correlation between nominal and real exchange rate changes is evidence in support of the Dornbusch (1976) model, which emphasizes monetary shocks and sticky prices. However, Stockman (1987, 1988) points out that there are always shocks to technology, tastes, and trade and fiscal policies that would move the real exchange rate under fixed or flexible exchange rate regimes. That these shocks also tend to show up in the nominal exchange rate instead of relative price levels is because monetary authorities try to stabilize the price level. Thus, the sticky-price model is not necessarily the only consistent explanation. My finding that nearly one-half of the forecast error variance of the real exchange rate can be attributed to monetary shocks is thus quite supportive of the Dornbusch model, but also suggests that there are important sources of real exchange rate variation in addition to monetary shocks.

Second, both equilibrium and sticky-price models are claimed to be consistent with the apparent non-neutrality of the real exchange rate across exchange rate regimes. The sticky-price view asserts that fluctuations in the real exchange rate are greater under flexible rates because nominal exchange rates are more volatile, while price levels are sluggish under either regime. On the other hand, as Stockman (1988) points out, although the real shocks discussed above affect the nominal exchange rate under flexible exchange rate regimes, these disturbances affect the central bank's

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which the two output shocks and the two monetary shocks are combined. The results, which were reported in a previous draft of this paper, are broadly consistent with those of the more informative five-by-five system. Third, and also reported in the previous draft, I estimated models with the real interest rate (differential) replacing government size. The sign of each of the elements of C(1) was the same as in the model with government size as the first variable, and the VDC results indicated that real and monetary shocks contribute approximately equally to the VDC of the real exchange rate. These results are available on request.

international reserves under fixed rates. When countries suffer a loss of international reserves due to a shock which gives rise to a real depreciation, for example, they are more likely to impose trade or capital controls -- the type of policy change considered to be a part of ϵ^D above -- in order to stem the flow of reserve loss which would otherwise force a devaluation. Optimizing agents forsee that the government will undertake policies to stabilize the real exchange rate (stem reserve losses), and as a consequence, the expectation that the future real exchange rate will be stabilized acts to stabilize the current rate.[20] Through this mechanism, real exchange rate fluctuations are smaller in response to a given real disturbance under a system of fixed exchange rates than under flexible rates.

The VAR models I estimate have the potential to discriminate between these different views of real exchange rate changes. After identifying shocks to fiscal policy, shocks to technology and tastes, and monetary shocks, I find evidence to support elements of both the "sticky-price view" of real exchange rate determination, which emphasizes monetary shocks, and the "equilibrium view", with its emphasis on real shocks. The estimates are also informative concerning the nature of the real shocks influencing the real exchange rate: these shocks are due more to fiscal and trade polcies than technology changes (which, for the U.K. relative to the U.S., perhaps were not very large over much of the sample), while the effect of monetary shocks is split into money multiplier shocks and monetary base shocks in roughly equal amounts.

C. Encompassing

In a recent paper, Clarida and Gali (1994) also use the Blanchard-Quah VAR methodolgy to examine the sources of movements in real exchange rates. Their sample consists of Canada, Germany, Japan, and the United Kingdom for the post-Bretton Woods era.

Clarida and Gali's VAR system contains the stationary variables (Δy, Δr, and π), where y is the U.S.-foreign country output growth differential, r is the real exchange rate, and π is the inflation

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[20]This stabilizing effect of expected future prices on current prices is standard in rational expectations models. It is analogous to the main result in Krugman's (1991) seminal paper on target zones: the expectation of future central bank intervention to stabilize the nominal exchange rate at the edge of the band makes the exchange rate more stable (than under a pure float) within the band.

differential. The corresponding shocks are labelled aggregate supply, aggregate demand, and nominal shocks. The identifying assumptions thus include that, in the long run, permanent disturbances to relative prices have no effect on the real exchange rate or the output differential, and that permanent disturbances to the real exchange rate have no effect on the output differential.

As indicated in Table 3, where I reproduce the authors' results, Clarida and Gali find that most of the variance of the real exchange rate is accounted for by aggregate demand shocks, although nominal (monetary) shocks are important for Germany and Japan. Aggregate supply shocks are insignificant at all horizons for all countries. Collecting data for the same countries and the same time period as Clarida and Gali, I attempt to replicate their results.[21] In part A of Table 3 I display the VDC results for $\Delta(r)$. The similarity with Clarida and Gali's results is evident from the table. Consider Japan. At the 20-quarter horizon, the contribution of each shock, respectively, is (3.60, 61.2, 35.2) in Clarida and Gali and (6.31, 59.2, 34.5) in my replication.

In part B of Table 3, I estimate Clarida and Gali's model using my long-term U.S.-U.K. data set. Once again, I focus only on the VDCs of $\Delta(r)$ in the interest of space. The results suggest that the aggregate demand shock (i.e., the permanent disturbance to the real exchange rate) is the dominant influence. Furthermore, the monetary shock accounts for approximately 15-25 percent of the forecast error variance, an amount that is in-between what Clarida and Gali report for the post Bretton Woods era (essentially zero for the U.K.) and what I find in my five-by-five model using the long-term data (around 40-45 percent).

Finally, in part C of table 3 I estimate my five-by-five model on the Clarida-Gali data set. The results for Japan and the U.K. are very similar to each other, and very similar to my findings for

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[21]Clarida and Gali obtain their price data from the IFS tape, the output data from the OECD Main Economic Indicators, and the exchange rate data from the New York Fed's database. I obtain these same variables from the FAME database of the Federal Reserve Board, in order to make things consistent with my source for the monetary base and money supply data used below. The different data sources accounts for the slight difference in our results.

the U.K. in the long-term data set.[22] Specifically, monetary shocks combined account for approximately one-half of the forecast error variance of $\Delta(r)$, and the effect is split between shocks to the money multiplier and shocks to the monetary base in roughly equal amounts. Second, the preference shock is the most influential of the remaining real shocks, contributing over 35 percent to the VDC of $\Delta(r)$ over short horizons and 25 percent in the long-run.

The estimates are also informative concerning the differences in results for the U.K. in table 3A and table 3B. Comparing table 3A (Clarida-Gali model on Clarida-Gali data) to 3B (Clarida-Gali model on my long-term data) suggests that the effect of using the longer data is to increase the contribution of the monetary shock at the expense of the aggregate demand shock, by approximately 20 to 25 percent. The contribution of the supply shock is not found to have increased noticeably in the long-term data. Now compare table 3A (Clarida-Gali model, Clarida-Gali data) to table 3C (my model, Clarida-Gali data), where the expanded model (i) allows certain AD disturbances -- permanent disturbances to government size -- to affect output in the long run, and (ii) contains a decomposition of the monetary disturbance. The effect of estimating the expanded model is to increase the contribution of monetary shocks from essentially zero to nearly one-half. Hence, at least for Clarida and Gali's estimates for the U.K., the effect of multiple aggregation of shocks discussed by Faust and Leeper (1994) is important. Finally, to reiterate what is noted in the preceding paragraph, a comparison of table 3C (my model, Clarida-Gali data) and table 2C (my model, my data) indicates that the choice of model is much more important than the choice of sample period for making inferences about the sources of real exchange rate movements.

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[22]The Canadian results are similar to those in the three-by-three model. It is difficult to make comparisons for Germany, because of the dramatic changes in the government spending and money supply data as a result of re-unification. Also note that Johansen's trace test indicates some evidence of 1 cointegration vector in the German and U.K. systems, but none for Canada or Japan. The test statistic for the null hypothesis of zero cointegration vectors (versus the alternative of 1) for Germany, the U.K., Canada, and Japan are 74.1, 72.5, 54.3, and 62.8, respectively, compared to the 10%, 5%, and 1% critical values of 64.8, 68.5, and 71.8. Hence, the German results are probably not worth discussing, and some pre-caution should be used when interpreting the U.K. results.

5. Conclusions

A long-standing issue in international finance is whether shocks to the real exchange rate are primarily real or monetary in nature. Previous authors have shown that both equilibrium and sticky-price models of exchange rate determination are consistent with the stylized facts concerning real exchange rates. Hence, I argue that, in order to assess the claims of the competing models, it is desirable to empirically estimate the relative contributions of real and monetary shocks on real exchange rate movements. My main finding is that output shocks and monetary shocks, the latter of which are restricted to having zero effect in the long run, account for approximately the same percentage of the variance of the real exchange rate over short horizons. Essentially all of the effect of the combined output shock is due to shocks to preferences, or "demand shocks", rather tahn technology shocks, which perhaps were not very large for the U.K. relative to the U.S. over much of the sample. The effect of the combined monetary shock is accounted for by both money multiplier shocks and monetary base shocks in roughly equal amounts. Thus, the paper suggests that the contributions of real and monetary shocks are roughly equal overall, while shedding light on the nature of those shocks.

APPENDIX: Interpreting the Shocks in the VAR Models: An Illustrative Long-Run Model

In order to illustrate how the residuals may be interpreted as fundamental structural shocks, I develop the following general-equilibrium optimizing model. The point of departure is Rogoff (1992) [see also section 3 of Froot and Rogoff (1994)], who explores the predictions of theoretical models concerning the types of shocks that will affect the real exchange rate. He considers some very stylized benchmark models, then makes the case that his empirical examples lie in between. First, in the classic Balassa-Samuelson model with perfect competition, perfect mobility of factors across sectors and internationally, and a small country setting, only productivity shocks will affect the real exchange rate. However, demand factors do have an impact in Rogoff's "fixed factor, open capital markets" model. More generally, there is role for demand factors in the short run (and possibly the long run) if the country is "large"; if there is some capital immobility internationally; if capital and labor are not instantaneously mobile across sectors domestically; and if there are increasing returns to scale (demand factors affect the real exchange rate in the long run if there is a fixed factor, such as land, which can be transferred across sectors, thus implying increasing returns to the mobile factors).

In the model that follows, I add two things to Rogoff's analysis: taste shocks, in the form of a disturbance to traded goods consumption, and money, via a cash-in-advance contraint. It should be noted that, like Rogoff, I specify a small country model. The U.K. and U.S. are not small economies. An alternative way of thinking about how the model produces the zero restrictions in (7) is to assume that equations analogous to the ones presented below hold for a "foreign" country. In the empirical estimation I take values of the U.K. variable less those of the U.S., of course.

I have in mind a model of a small, open economy which produces both traded and non-traded goods, according to the production technology,

$$y_{Tt} = A_{Tt} L_{Tt}^{\omega_T} K_{Tt}^{1-\omega_T} \tag{A1a}$$

$$y_{Nt} = A_{Nt} L_{Nt}^{\omega_N} K_{Nt}^{1-\omega_N} \tag{A1b}$$

where y_{it} is output in sector i (i=T,N), L_{it} and K_{it} represent the labor and capital input into y_{it} and A_{it} is a stochastic productivity shock in sector i. I will abstract from considering growth in, or depreciation of, the inputs. Assume that L and K are immobile across sectors, although, as noted above, this is not strictly necessary for there to be a role for factors other than productivity.

There is a representative agent with time-separable utility over consumption,

$$V_t = E_t \sum_{s=0}^{\infty} \beta^{s-t} \left[\frac{(c_{Ns}^{\alpha} \theta_s c_{Ts}^{1-\alpha})^{1-\gamma}}{1-\gamma} \right] \tag{A2}$$

where E is the expectations operator, β is the subjective discount rate, c_{it} is consumption of good i at time t, and θ_t is a shift parameter that captures changes in preferences toward the traded good.

Assume that domestic agents hold domestic money M and foreign money M*. With P_N (P_T) denoting the price of the non-traded (traded) good, and $Q = P_N/P_T$, we can define real money balances $m = M/P_N$ and $m^* = M^*/P_T$, so that domestic real balances in terms of tradeables is given by $\tilde{m} = Qm = M/P_T$. In this case, we add to the model the individual's budget constraint,

$$\tilde{m}_{t+1} + m_{t+1}^* = z_t - c_{Tt} - Q_t c_{Nt} - \pi_t \tilde{m}_t - \pi_t^* m_t^* - g_{Tt} - Q_t g_{Nt} \tag{A3}$$

where z is total income from domestic production (in terms of tradeables), $\pi = \Delta P_N/P_N$, and $\pi^* = \Delta P_T/P_T$. I assume that government consumption of traded and non-traded goods, g_T and g_N respectively, is financed by lump-sum taxation, and that the representative agent integrates the government budget constraint into her own.

It is also assumed that the representative agent needs to hold various amounts of domestic and foreign real balances in order to purchase consumption goods,

For the sake of tractibility, I assume that the money demand function f( ) takes the form,

$$c_{Tt} + Q_t\, c_{Nt} \le f(\tilde{m}_t, m_t^*) \qquad\qquad\text{(A4)}$$

$$f(\tilde{m}_t, m_t^*) = \rho \cdot \tilde{m}_t^\phi (m_t^*)^{1-\phi} \qquad\qquad\text{(A5)}$$

where ϕ $(1-\phi)$ is the share of total money balances held in the form of domestic (foreign) money, and ρ denotes the fraction devoted to "cash goods", where $0 < \rho \le 1$.

The agent's problem is to maximize (A2) subject to (A3)-(A5). The Hamiltonian is written,

$$H = E_t \sum_{s=0}^{\infty} \beta^{s-t} \left[\frac{(c_{Ns}^\alpha \theta_s c_{Ts}^{1-\alpha})^{1-\gamma}}{1-\gamma}\right] + \lambda_2\, [\rho \tilde{m}_s^\phi (m_s^*)^{1-\phi} - c_{Ts} - Q_s c_{Ns}]$$

$$+ \lambda_1\, [z_s - \tilde{m}_{s+1} - m_{s+1}^* - c_{Ts} - Q_s c_{Ns} - \pi_s \tilde{m}_s - \pi_s^* m_s^* - g_{Ts} - Q_s g_{Ns}]$$

Using the first-order conditions for both traded and non-traded goods consumption gives us,

$$Q_t = \theta_t\left(\frac{\alpha}{1-\alpha}\right)\left(\frac{c_{Tt}}{c_{Nt}}\right) \qquad\qquad\text{(A6)}$$

while the first-order conditions for domestic and foreign money balances imply,

$$\tilde{m}_t = \left(\frac{\phi}{1-\phi}\right) m_t^* \left(\frac{\pi_t^*}{\pi_t}\right) \qquad\qquad\text{(A7)}$$

To derive an expression for domestic real money balances, rewrite the cash-in-advance contraint as,

$$\frac{c_{Tt}}{Q_t} + c_{Nt} = \rho m_t \left[\frac{m_t^*}{\tilde{m}_t}\right]^{1-\phi} \qquad\qquad\text{(A8)}$$

where it will be recalled that $m_t = \tilde{m}_t/Q_t$. Also, note that (A6) can be rewritten $c_T/Q = \theta^{-1}c_N[(1-\alpha)/\alpha]$, so that this and (A7) substituted into (A8) becomes, upon rearranging,

$$m_t = \theta_t^{-1}\left(\frac{1}{\alpha}\right)c_{Nt} \cdot \rho^{-1}\left[\left(\frac{1-\phi}{\phi}\right)\left(\frac{\pi_t}{\pi_t^*}\right)\right]^{\frac{1}{1-\phi}} \qquad\qquad\text{(A9)}$$

Assuming that equilibrium in the non-traded goods market holds,

$$y_{Nt} = c_{Nt} + g_{Nt} \qquad \text{(A10)}$$

and that government consumption of non-traded goods is financed by lump-sum taxes,

$$g_{Nt} = \tau_t \, y_{Nt} \qquad \text{(A11)}$$

then (A9) can be rewritten,

$$m_t = \theta_t^{-1}(\frac{1}{\alpha})\rho^{-1}[(\frac{1-\phi}{\phi})(\frac{\pi_t}{\pi_t^*})]^{\frac{1}{1-\phi}}(1-\tau_t)y_{Nt} \qquad \text{(A12)}$$

In log-differences, this becomes,

$$\Delta\ln(m_{t+1}) = (\frac{1}{\phi})\Delta\ln(\mu_{t+1}) - \Delta\theta_{t+1} - \Delta\tau_{t+1} + \Delta\ln(A_{N,t+1}) \qquad \text{(A13)}$$

where I have used the steady-state result that $\pi_t = \mu_t \equiv \Delta M_t/M_t$, normalized π^* to unity, and used equations (A11) and (A1b).

Now rewrite (A6), using (A10) and (A11), as

$$Q_t = \theta_t(\frac{\alpha}{1-\alpha})[\frac{c_{Tt}}{(1-\tau_t)y_{Nt}}] \qquad \text{(A14)}$$

or in log-differences,

$$\Delta\ln(Q_{t+1}) = \Delta\ln(c_{T,t+1}) + \Delta\theta_{t+1} + \Delta\tau_{t+1} - \Delta\ln(A_{N,t+1}) \qquad \text{(A15)}$$

Note that the first-order condition for traded goods consumption implies,

$$(\frac{c_{Nt}}{c_{Tt}})^\alpha \cdot (c_{Nt}^\alpha c_{Tt}^{1-\alpha})^{-\gamma} = E_t \, \beta[(\frac{c_{N,t+1}}{c_{T,t+1}})^\alpha \cdot (c_{N,t+1}^\alpha c_{T,t+1}^{1-\alpha})^{-\gamma}] \qquad \text{(A16)}$$

Taking logs of both sides of (A16) and rearranging allows us to approximate the equation as,

$$E_t\Delta\ln(c_{T,t+1}) = [\frac{\alpha(1-\gamma)}{\alpha(1-\gamma)+\gamma}] E_t\Delta\ln c_{N,t+1} \qquad \text{(A17)}$$

Assuming for simplicity that the expected change in consumption equals the actual change, then (A15) can be rewritten, using (A17) along with (A1b), (A10), and (A11), as:

$$\Delta\ln(Q_{t+1}) = \Delta\theta_{t+1} + \Delta\tau_{t+1}[\frac{\gamma}{\alpha(1-\gamma)+\gamma}] - \Delta\ln(A_{N,t+1})[\frac{\gamma}{\alpha(1-\gamma)+\gamma}] \quad\quad (A18)$$

Now consider the relations $\Delta\theta_t = \bar{r} + \epsilon^D$, $\Delta\tau_t = \bar{g} + \epsilon^g$, $\Delta\ln(A_{Nt}) = \bar{y} + \epsilon^S$, and

$\Delta\ln(\mu_t) = \bar{M} + \epsilon^m$. In order to derive equation (7), it is necessary to use these and to discuss two

more aspects of the model. First, since equilibrium output is supply-determined, the preference shock

ϵ^D, will not affect output in the long run. This implies the ordering of variables two and three. Also,

allowing for non-optimizing government activities, the fiscal shock affects output, and implies the

ordering of variables one and two [i.e., not imposing the over-identifying restriction $C_{21}=0$].

Second, the final two equations in (7) are derived using equation (A13), which indicates that

real money balances, m, are a function of the fiscal, productivity, preference, and money growth

shocks. Writing m = mm + h, where mm is the log money multiplier and h is the log real monetary

base, the ordering of variables four (Δmm) and five (Δh) reflects the assumption that shocks to the

monetary base have a one-to-one effect on the money supply in the long-run.

Finally, two caveats are in order. First, replacing $\ln(Q) = p_N-p_T$ with r (= $s+p^*-p$) implies a

switch in the signs of c_{21} and c_{22} in going from (A18) to (7). Writing the home price index as

$p = ap_N+(1-a)p_T$, and assuming an analogous expression for foreign prices, allows us to write

$r = (s+p_T^*-p_T) - a(p_N-p_T) + a^*(p_N^*-p_T^*)$, and implies a negative relationship between Q and r. Also note

that the real exchange rate equation in (7) would be obtained if we considered an equation analogous

to (A18) holds for a "foreign" country. If relative PPP holds for traded goods, so that ($s+p_T^*-p_T$) is

constant, this alternative two-country specification would imply an exact correspondence between Q

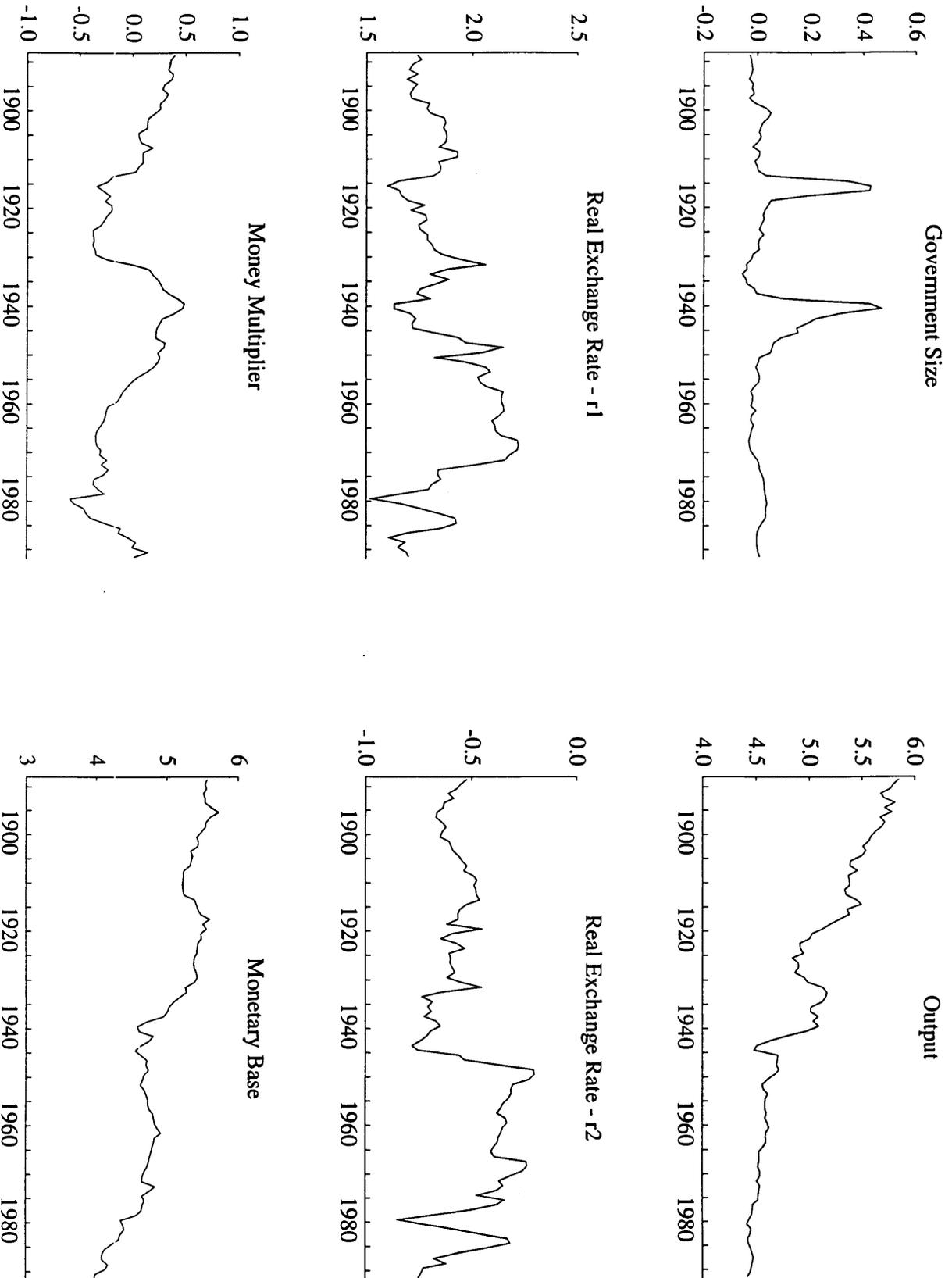(or "Q - Q^*"), which is used in the model, and r, which is used in the empirical work.

Second, the signs of c_{21} and c_{22}, the long-run effect of the fiscal shock and the supply shock

on the real exchange rate, respectively, are unclear in theory. Equation (A6) indicates that in general

the signs depend on whether the shocks have a larger effect on the consumption of traded or non-

traded goods. This becomes somewhat hidden in the derivation of (7), because I have used (A16) to

solve for traded goods consumption in terms of non-traded goods consumption, and hence have

written the model in terms of the productivity shock in the non-traded goods sector and government consumption of non-traded goods (which is probably the more accurate choice, since most government spending is on non-tradeables). That fiscal shocks have an ambiguous effect on the real exchange rate is a feature of several models.
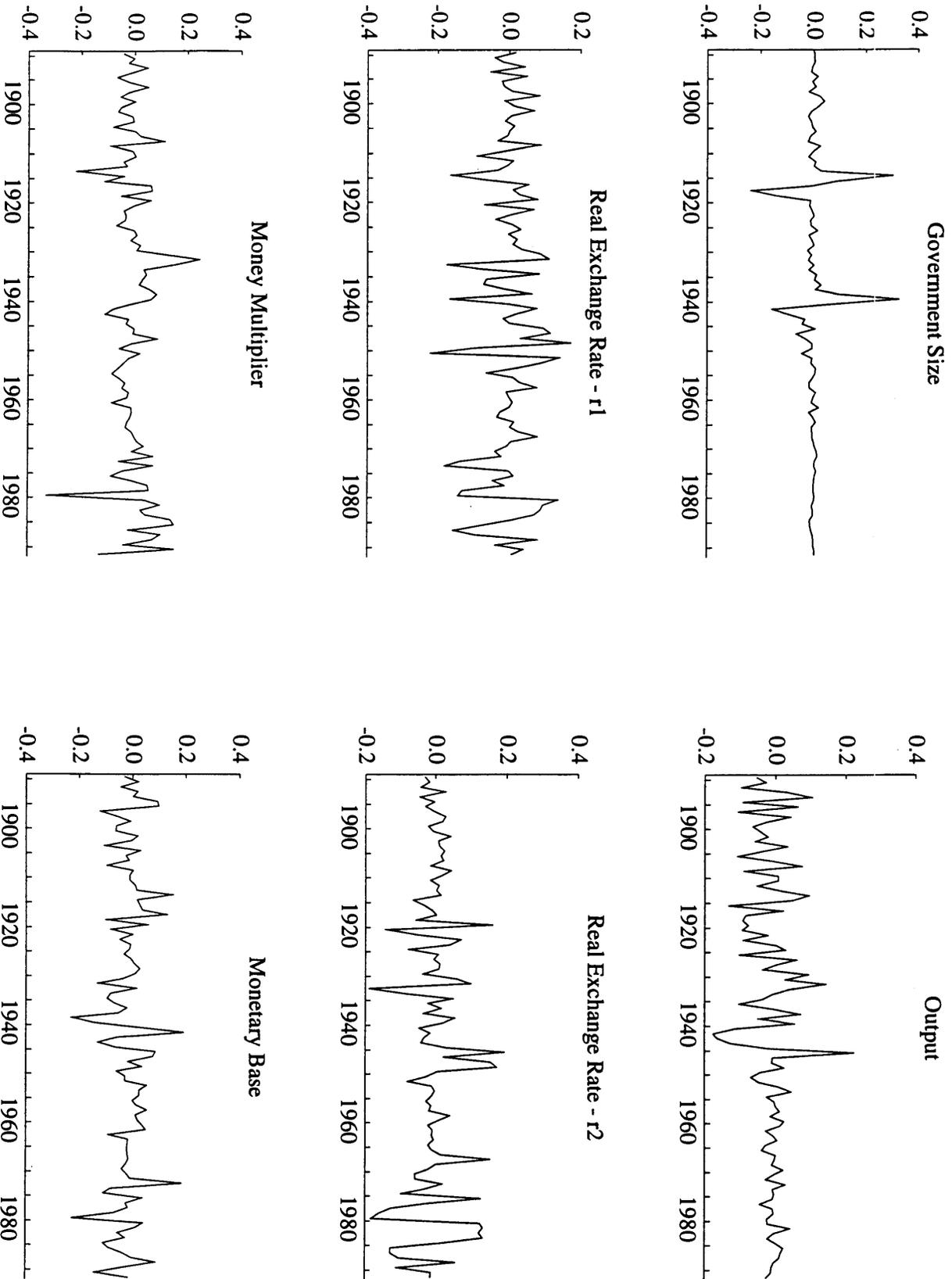
# Figure 1: Real and Nominal Exchange Rates

Nominal Rate and Wholesale Price-Based Real Rate

R1 ←

S ⇢

Nominal Rate and GDP Deflator-Based Real Rate

R2 ←

S ⇢
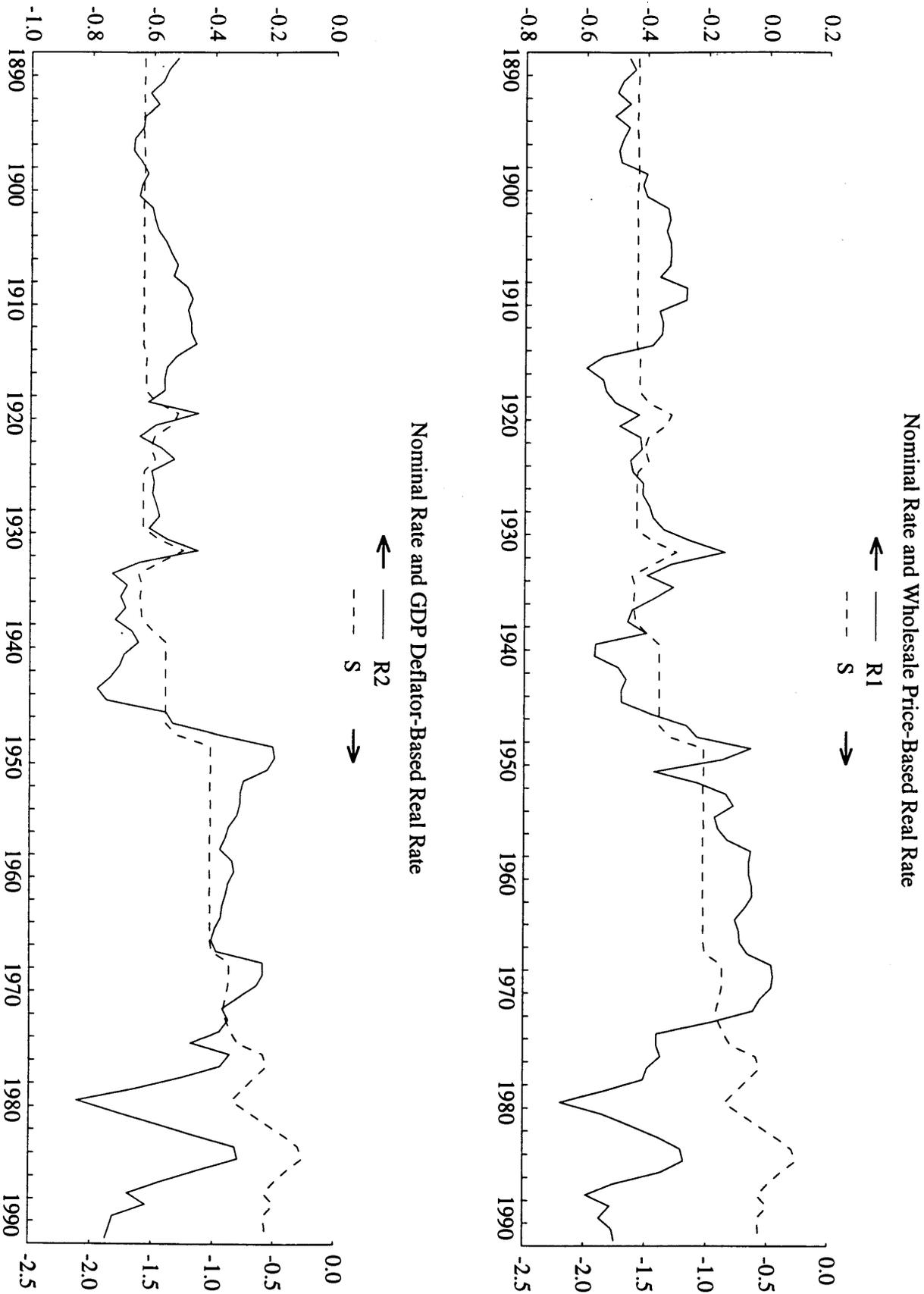
Figure 2a: Variables in the VAR - Levels

Government Size

Real Exchange Rate - r1

Money Multiplier

Output

Real Exchange Rate - r2

Monetary Base

Figure 2b: Variables in the VAR - First-Differences

Government Size

Output

Real Exchange Rate - r1

Real Exchange Rate - r2

Money Multiplier

Monetary Base

Figure 3a; Impulse Responses of r1

Fiscal Shock
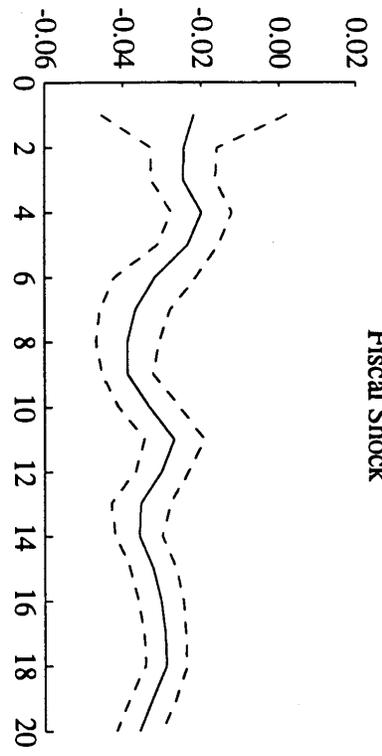
Money Multiplier Shock
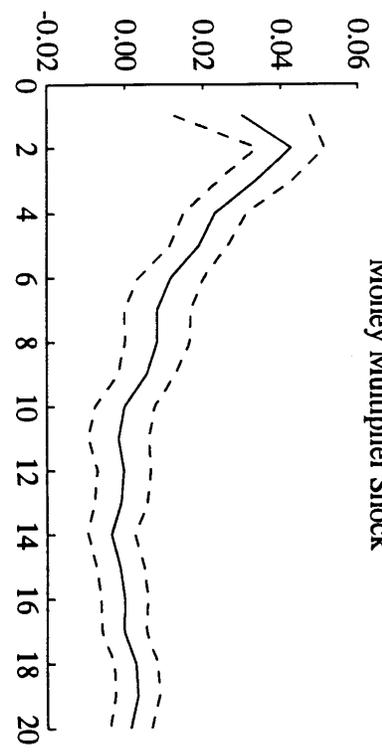
Supply (Productivity) Shock
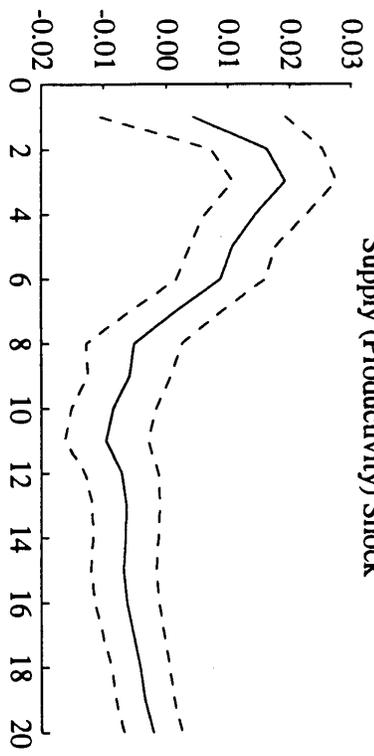
Monetary Base Shock
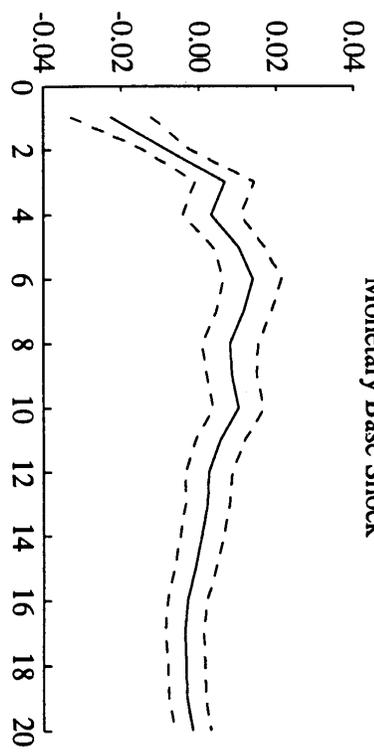
Demand (Preferences) Shock
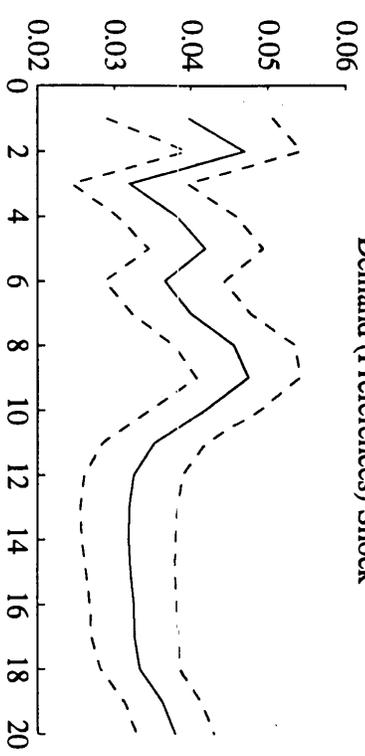
Figure 3b; Impulse Responses of Dr1

Fiscal Shock

Money Multiplier Shock

Supply (Productivity) Shock

Monetary Base Shock

Demand (Preferences) Shock
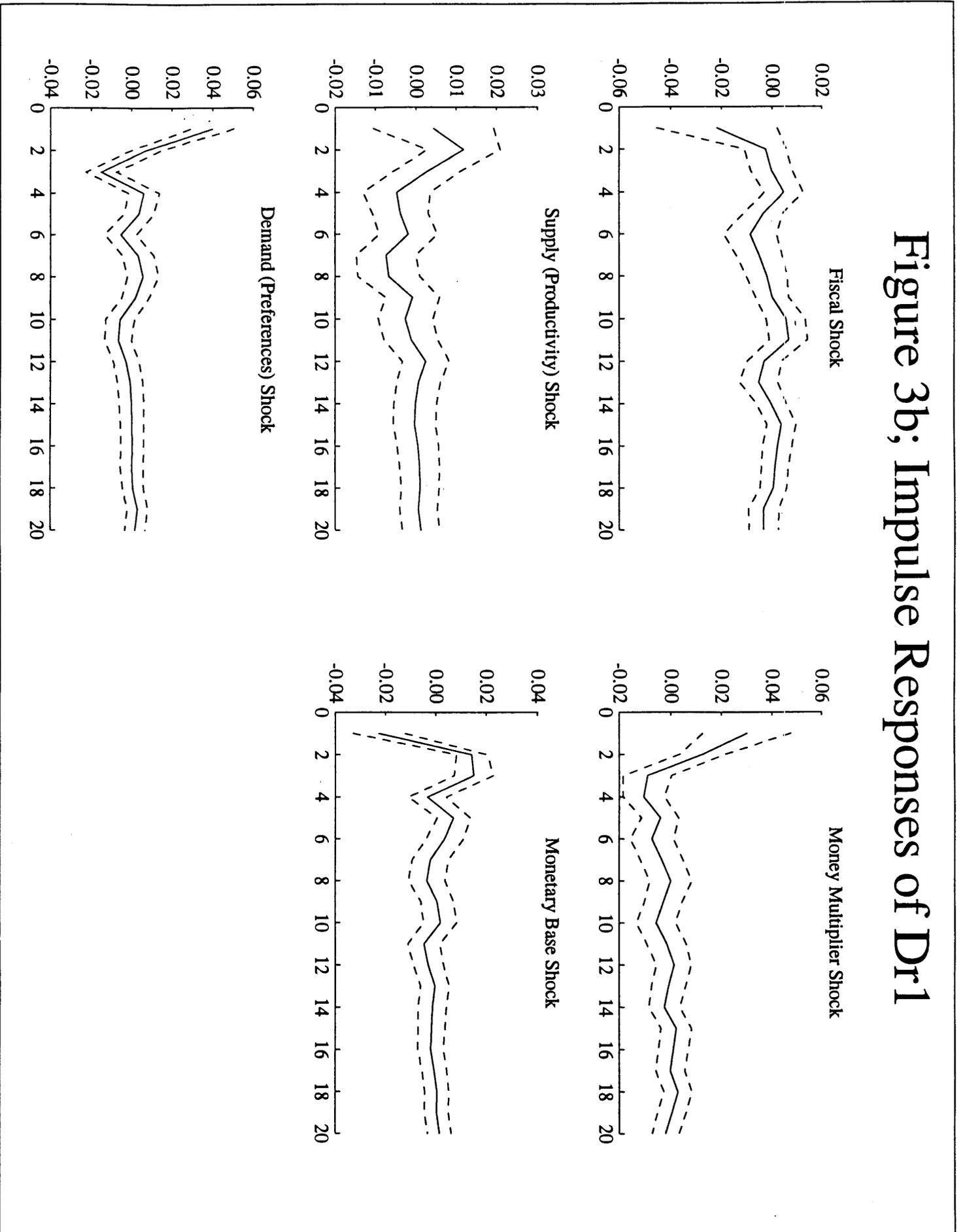
Table 1:
Stock and Watson's Dynamic OLS Estimates of the Cointegrating Relationship Between S and P

Regression	Estimated Coeff. (standard error)	Sample Period	Type of Dummy used for Breaks ?
S on P1	1.10 (.044)*	1859-1992	No Break
S on P1	1.06 (.043)	1859-1992	War/Non-War A
S on P1	.955 (.059)	1859-1992	War/Non-War B
S on P1	1.22 (.052)*	1859-1992	Fixed / Float
S on P2	1.20 (.061)*	1889-1992	No Break
S on P2	1.08 (.074)	1889-1992	War/Non-War A
S on P2	1.09 (.076)	1889-1992	War/Non-War B
S on P2	1.24 (.052)*	1889-1992	Fixed / Float
P1 on S	.843 (.051)*	1859-1992	No Break
P1 on S	.882 (.047)*	1859-1992	War/Non-War A
P1 on S	.906 (.062)	1859-1992	War/Non-War B
P1 on S	.642 (.041)*	1859-1992	Fixed / Float
P2 on S	.804 (.052)*	1889-1992	No Break
P2 on S	.926 (.050)	1889-1992	War/Non-War A
P2 on S	.908 (.052)	1889-1992	War/Non-War B
P2 on S	.755 (.035)*	1889-1992	Fixed / Float

Notes: The cointegrating vectors above are estimated using Stock and Watson's dynamic OLS procedure. The lead/lag length of 5 on the dynamics was determined by a likelihood ratio test. In parentheses are standard errors, which have been corrected for heteroscedasticity using White's (1980) technique. A * denotes that the estimate is significantly different from the value of 1.00 predicted by PPP. The final column indicates the type of dummy variable employed in conjunction with the short-run dynamic terms. The War/Non-War A dummy variable is zero during non-war years and unity when either the U.S. and/or U.K. was at war. The War/Non-War B dummy variable is zero during non-war years and is equal to the ratio of military personnel to the general population (the U.K. ratio less that of the U.S.) during war years. The war years are 1899-1902, 1914-19, and 1939-46 (U.K.) and 1861-1865, 1917-19 and 1941-46 (U.S.). The Fixed/Float dummy is unity for the years 1914, 1919-25, 1931-40, 1949, and 1973-present.

A. Johansen's Trace Test for Cointegration

System	p=0	p ≤ 1	p ≤ 2	p ≤ 3	p ≤ 4	Lags
(g/y), y, r1, mm, h	40.9	21.0	10.0	3.28	.0001	7
(g/y), y, r2, mm, h	47.3	24.2	12.7	3.42	0.15	7

Notes: The null is that there are no more than p cointegrating vectors in the system (n-p distinct unit roots). Statistics are calculated using the Cheung and Lai (1993) correction. Critical values for (n-5) equal to 1,2,3,4,5 are 3.76, 15.5, 29.7, 47.2, 68.5 (95%) and 2.69, 13.3, 26.8, 44.0, 64.8 (90%). These are taken from Table 1 of Osterwald-Lenum (1992). For each VAR, I use a procedure based on a sequence of likelihood ratio tests to obtain the appropriate lag length. The procedure starts with 12 lags, tests if the last lag is significant, then reduces the lag length by one and continues.

B. Coefficients of the Long-Run Moving Average Matrix

Coeff./System	Δ(g/y), Δy, Δr1, Δmm, Δh	Δ(g/y), Δy, Δr2, Δmm, Δh
C_{11}	0.06	0.001
C_{21}	0.03	0.001
C_{22}	0.49	0.32
C_{31}	-0.07	-0.001
C_{32}	-0.07	-0.007
C_{33}	0.55**	0.44**
C_{41}	0.11	0.003
C_{42}	0.27	0.01
C_{43}	-0.10	-0.32
C_{44}	1.14*	1.00#
C_{51}	-0.11	-0.002
C_{52}	-0.19	-0.07
C_{53}	0.19	0.30
C_{54}	-0.75*	-0.58#
C_{55}	-0.54**	-0.54**

Notes: The estimates represent the long-run effects of (i) the fiscal shock on government size (C_{11}), output, (C_{21}), the real exchange rate (C_{31}), the money multiplier (C_{41}), and real monetary base (C_{51}); (ii) the output supply shock on output (C_{22}), the real exchange rate (C_{32}), the money multiplier (C_{42}). and real monetary base (C_{52}); (iii) the demand shock on the real exchange rate (C_{33}), the money multiplier (C_{43}), and real monetary base (C_{53}); (iv) the money multiplier shock on the money

multiplier (C_{44}), and real monetary base (C_{54}); and (v) the monetary policy shock on the real monetary base (C_{55}). A (**, *, #) indicates significance at the 1%, 5%, and 10% level, respectively.

C. VDCs of the Real Exchange Rate

The VDC of $\Delta(r_j)$

Shock/System:	$\Delta(g/y)$, Δy, $\Delta r1$, Δmm, Δh	$\Delta(g/y)$, Δy, $\Delta r2$, Δmm, Δh
ϵ^g	13.6, 11.8, 10.5, 13.6	12.6, 15.7, 14.7, 15.4
	(16.5)(14.4)(11.4)(8.46)	(12.5)(11.1)(8.89)(6.92)
ϵ^S	0.55, 3.99, 4.22, 5.80	0.12, 2.57, 5.86, 7.72
	(8.86)(8.44)(7.21)(5.91)	(5.39)(5.28)(5.16)(4.51)
ϵ^D	45.3, 40.3, 38.9, 36.7	40.4, 39.2, 37.4, 35.5
	(19.7)(17.2)(14.3)(10.9)	(18.2)(16.1)(14.1)(10.6)
ϵ^{mm}	26.0, 26.3, 26.0, 24.8	18.7, 16.4, 14.5, 15.0
	(18.8)(16.3)(13.6)(9.80)	(16.4)(14.1)(11.9)(8.36)
ϵ^h	14.7, 17.6, 20.4, 19.1	28.2, 26.0, 27.4, 26.4
	(12.3)(10.1)(9.67)(7.39)	(14.9)(12.0)(11.6)(8.92)

The VDC of (r_j)

Shock/System:	$\Delta(g/y)$, Δy, $\Delta r1$, Δmm, Δh	$\Delta(g/y)$, Δy, $\Delta r2$, Δmm, Δh
ϵ^g	13.6, 12.6, 15.2, 34.5	12.6, 19.1, 22.9, 46.9
	(16.5)(16.6)(16.6)(19.4)	(12.5)(13.4)(13.7)(17.8)
ϵ^S	0.55, 3.37, 5.73, 2.81	0.12, 0.77, 5.46, 2.63
	(8.86)(10.4)(10.9)(8.21)	(5.39)(5.67)(7.93)(5.83)
ϵ^D	45.3, 44.8, 47.0, 51.1	40.4, 45.9, 43.7, 38.3
	(19.7)(19.8)(18.9)(17.8)	(18.2)(18.4)(17.5)(15.2)
ϵ^{mm}	26.0, 32.3, 27.8, 9.15	18.7, 17.1, 17.6, 6.19
	(18.8)(19.8)(17.5)(6.83)	(16.4)(16.1)(15.1)(5.57)
ϵ^h	14.7, 6.92, 4.34, 2.51	28.2, 17.2, 10.4, 6.06
	(12.3)(7.83)(3.95)(2.51)	(14.9)(11.6)(7.40)(4.38)

Notes: In the table above are the variance decompostion (VDC) results. These indicate the percentage of the forecast error variance of $\Delta(r_j)$ or (r_j) that is due to each shock at the 1-year, 2-year, 5-year, and 20-year horizons. By construction, the contributions of ϵ^{mm} and ϵ^h to the variance of (r) go to zero asymptotically. Lag length is determined as described in table 2A above. In parenthesis are standard errors, calculated using Monte Carlo simulations with 1,000 draws.

Table 3: Encompassing (VDCs of Real Exchange Rate Changes)

A. My Replication of Clarida and Gali's Results

Country	Supply Shock	Demand Shock	Money Shock
Canada (CG)	2.20 / 4.50 / 4.30	97.0 / 94.3 / 93.2	0.70 / 1.20 / 2.50
Canada	8.48 / 9.82 / 10.5	87.0 / 85.0 / 84.0	4.49 / 5.18 / 5.55
Germany (CG)	0.70 / 4.60 / 10.4	51.8 / 50.3 / 48.5	47.4 / 45.0 / 41.1
Germany	8.56 / 16.0 / 15.2	47.1 / 43.1 / 43.9	44.4 / 40.9 / 40.9
Japan (CG)	0.90 / 3.40 / 3.60	63.1 / 62.7 / 61.2	35.9 / 33.9 / 35.2
Japan	2.71 / 4.92 / 6.31	67.2 / 66.0 / 59.2	30.1 / 29.1 / 34.5
U.K. (CG)	0.50 / 5.50 / 6.10	97.3 / 91.9 / 91.1	2.20 / 2.50 / 2.80
U.K.	0.10 / 5.46 / 7.09	99.9 / 89.5 / 86.6	0.03 / 5.06 / 6.32

Notes: Entries indicate the percentage of the forecast error variance of $\Delta(r)$ that is due to each shock at the 1-quarter, 4-quarter, and 20-quarter horizons. For each country, the first row of results (with (CG) following the country name) indicates that the estimates are taken from Table 4 of Clarida and Gali (1994). The second row of results is my replication of their results.

B. Estimating Clarida and Gali's Model on My Data

System	Supply Shock	Demand Shock	Money Shock
$\Delta y, \Delta r1, \pi$	9.10, 11.9, 13.3	79.3, 74.8, 69.7	11.6, 13.3, 16.9
$\Delta y, \Delta r2, \pi$	2.30, 5.56, 7.65	80.0, 68.7, 65.4	17.6, 25.7, 26.9

Notes: Entries indicate the percentage of the forecast error variance of $\Delta(r)$ that is due to each shock at the 1-year, 5-year (20-quarters), and 20-year horizons.

C. Estimating My Five-By-Five Model on Clarida and Gali's Data

Shock/ Country	Canada	Germany[a]	Japan	U.K.[b]
ϵ^g	20.8, 28.7, 28.4	9.18, 17.9, 23.6	1.99, 7.67, 9.75	1.91, 14.4, 13.9
ϵ^s	4.51, 4.58, 6.06	45.4, 36.5, 32.6	5.56, 10.4, 15.2	0.10, 5.99, 5.97
ϵ^D	72.1, 62.3, 58.1	41.9, 38.6, 35.6	42.0, 34.7, 25.2	36.3, 30.4, 28.3
ϵ^{mm}	0.61, 0.73, 1.45	3.43, 5.99, 5.48	22.0, 23.2, 27.3	36.5, 28.9, 28.3
ϵ^h	1.96, 3.64, 6.00	0.07, 1.05, 2.60	28.5, 24.0, 22.6	25.1, 20.3, 23.6

Notes: Entries indicate the percentage of the forecast error variance of $\Delta(r)$ that is due to each shock at the 1-quarter, 4-quarter, and 20-quarter horizons. (a) Data for Germany end in 1990:4, due to

definitional problems associated with the fiscal data as a result of German reunification. (b) M4 is used because of two large definitional changes in U.K. M2 during the sample.

APPENDIX TABLES:

Results from the Benchmark Model with Stationary (g/y) and Trend-Stationary r

Coefficients of the Long-Run Moving Average Matrix

Coeff./System	(g/y), Δy, r1, Δmm, Δh	(g/y), Δy, r2, Δmm, Δh
C_{11}	1.68	1.82
C_{21}	-0.21	-0.20
C_{22}	0.85	0.83
C_{31}	-6.04	-5.19
C_{32}	-0.96	-0.69
C_{33}	5.23	2.80
C_{41}	0.90	1.10
C_{42}	1.08	1.20
C_{43}	-0.98	-0.67
C_{44}	0.72	0.42
C_{51}	-0.48	-0.54
C_{52}	-0.36	-0.55
C_{53}	0.69	0.45
C_{54}	-0.46	-0.24
C_{55}	-0.64	-0.54

Notes: The estimates represent the long-run effects of the shocks, as in table 2B.

The VDC of (r_j)

Shock/System:	(g/y), Δy, r1, Δmm, Δh	(g/y), Δy, r2, Δmm, Δh
ϵ^g	32.7, 38.7, 49.0, 48.7	16.4, 32.4, 46.3, 41.5
ϵ^S	3.19, 7.04, 10.8, 6.67	0.02, 0.07, 5.04, 5.75
ϵ^D	7.23, 3.85, 4.35, 22.0	8.30, 8.20, 6.65, 27.1
ϵ^{mm}	35.0, 37.4, 27.4, 16.3	34.9, 30.0, 21.2, 11.6
ϵ^h	21.8, 13.0, 8.40, 6.32	40.3, 29.3, 20.8, 14.0

Notes: The table indicates the percentage of the forecast error variance of (r_j) that is due to each shock at the 1-year, and 20-year horizons, as in table 2C.

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