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9 Actual versus Unanticipated Changes in Aggregate Demand Variables: A Sensitivity Analysis of the Real-Income Equation

Michael R. Darby

A large, if not dominant, body of recent research in macroeconomics incorporates the Barro (1977, 1978) variant of the Lucas supply function. The analytical convenience of this approach is well known. For empirical work it has considerable attraction as well: It imposes restrictions upon how changes in money affect real income, and it may be stable despite a change in the monetary regime governing the money supply process.[1] In particular, an empirical investigator can define expected money growth by an ARIMA process, a transfer function, or other parsimonious means and then include in the real-income equation only a few money shocks (innovations)—the difference between actual and expected money growth. Thus a great saving in parameters estimated is to be achieved compared to estimating a long distributed lag on actual money growth rates as would be required to obtain effectively the same equation.[2]

This paper investigates whether the Barro restriction that only money shocks (not anticipated money growth) affect real income is supported by the data for other countries and for two other factors affecting aggregate demand: real government spending and real exports. The empirical

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1. If a change in monetary regime does not alter the predictability of the future money supply, then the coefficients on money-supply innovations (or shocks) in the Barro variant would apparently remain unchanged.

2. See Sargent (1976) on the equivalence of these two approaches in the absence of identifying information in a money-supply transfer function which is not present in the real-income equation. See also McCallum (1979) on testing for the validity of the Barro variant even in the absence of such a priori identifying information on a change in monetary regime.

results suggest that the data are not inconsistent with the Barro restrictions. However, except for the United States, it makes very little difference whether one-year distributed lags on unanticipated or actual changes in aggregate demand variables are used in the real-income regressions. Certainly the results would not suggest use of the unanticipated variables in the absence of a priori preference. While these results can be rationalized by greater measurement errors in the foreign data—no real-income regression explains much—they are sufficiently surprising to warrant further investigation and cautious application of Barro's approach.

The Mark III International Transmission Model's real-income equations (R1) and (N1) are derived in section 9.1 as a generalization of the familiar Barro real-income equation. These equations are subjected to a sensitivity analysis in research reported in section 9.2. A summary and suggestions for future research conclude the chapter.

9.1 A Generalized Barro Real-Income Equation

The real-income equation is derived by combining a Lucas supply function with a standard aggregate demand function to obtain real income as a function of lagged transitory real income and shocks in nominal money, real government spending, and real exports.[3] Other aggregate demand variables such as taxes are not included because of lack of adequate international data.[4]

The aggregate supply function is of the Lucas (1973) form:

$$\Delta \log y = a_1 - a_2 \log(y_{-1}/y^P_{-1}) + a_3 \hat{P} + \epsilon, \tag{9.1}$$

where country subscripts are omitted for simplicity, y is real income, y^P is the natural-employment or permanent value of real income, a_1 is the periodic growth rate of y^P, ϵ is a white noise disturbance, and $\hat{P}$ is the price-level shock:

$$\hat{P} = \log P - (\log P)^*, \tag{9.2}$$

where P is the price level and an asterisk denotes expectations based upon the previous period's information set.

The aggregate demand function is assumed semi-log linear:

$$\log y = b_1 + b_2 \log(M/P) + b_3 \log g + b_4 x + v, \tag{9.3}$$

3. For similar derivations, see McCallum (1978), Korteweg (1978), and Horrigan (1980).

4. No bias will result from including the effects of these variables in the error term unless their innovations are correlated with the innovations in the included variables. Were this the case, the expected values of estimated coefficients would be augmented by the product of the omitted coefficients and the regression coefficients of the omitted variables on the included variables. See Theil (1971, pp. 548–56).

where M is the nominal money supply, g is real government spending, x is exports divided by income,[5] and v is another white noise disturbance uncorrelated with ϵ. Familiar manipulations yield the semi-reduced-form real-income equation

$$\Delta \log y = a_1 - a_2 \log(y_{-1}/y^P_{-1}) + \frac{1}{1+\frac{b_2}{a_3}}\left(b_2\hat{M} + b_3\hat{g} + b_4\hat{x} + \left(\frac{b_2}{a_3}\epsilon + v\right)\right), \tag{9.4}$$

where $\hat{M}$, $\hat{g}$, and $\hat{x}$ are the differences between the actual and expected values of log M, log g, and x, respectively. It is generally argued that inventory fluctuations will lead to some lags in the adjustment of output (as opposed to final sales) so that some short distributed lags on $\hat{M}$, $\hat{g}$, and $\hat{x}$ are permitted as well as the contemporaneous terms.[6] For example, using quarterly data and assuming any inventory lags are corrected within a year,

$$\Delta \log y = a_1 - a_2 \log(y_{-1}/y^P_{-1}) + \sum_{i=0}^{3} c_{1+i}\hat{M}_{-i} + \sum_{i=0}^{3} c_{5+i}\hat{g}_{-i} + \sum_{i=0}^{3} c_{9+i}\hat{x}_{-i} + e, \tag{9.5}$$

where e is the combined residual disturbance.

This is the form of the real-income equation used in the Mark III International Transmission Model[7] and investigated in section 9.2 below. The empirical basis for including only innovations in money and not anticipated changes in money is by now well known, but it is perhaps worthwhile to comment briefly here on the corresponding basis for real government spending and exports.

The real-income equation (9.5)—assuming positive short-run effects—implies that unexpected increases in government spending or exports cause a short-run increase in real income, but this short-run increase is eliminated over time. That is, there is complete long-run *real* crowding out. I have argued elsewhere (1979, pp. 225–27) that this pattern represents a rough consensus of empirical results for the United States. As with money growth, however, alternative anticipated levels of real government spending and of real exports may imply different steady-state values of the capital-labor ratio so that their anticipated levels may belong in the

5. Recall from chapter 5 above that exports are scaled by dividing by income instead of by taking logarithms because in the Mark III International Transmission Model a balance-of-payments identity involving sometimes negative numbers is imposed. This should only cause an offsetting change in the magnitude but not the significance of the estimated export coefficients.

6. See particularly Haraf (1979).

7. The lagged value of log y is moved from the left to right side in the Mark III Model, but this has no effect on estimated coefficients and standard errors.

real-income equation even though they do not affect the natural-employment level of labor input. Further, there may be incentive effects on labor supply and efficiency effects associated with different sizes of government, but again empirical evidence is lacking to date. As specified, the real-income equation (9.5) embodies a hypothesis that the effects of anticipated M, g, and x on y via capital or otherwise are negligible.

9.2 Empirical Results

Our empirical investigation is based upon the 1955–76 quarterly data bank described in chapter 3 and the Data Appendix to this volume. Data are available for the United States, the United Kingdom, Canada, France, Germany, Italy, Japan, and the Netherlands. Two years are lost due to lagged variables which appear in the real-income equation (9.5) and in the definitions of expected values, so all estimations are for the eighty quarters from 1957I through 1976IV.

Table 9.1 reports ordinary least squares (OLS) estimates of equation (9.5), where $\hat{M}$, $\hat{g}$, and $\hat{x}$ are defined as the residuals on univariate ARIMA processes fitted according to the methods of Box and Jenkins (1976) using the programs described in Nelson (1973). As reported in chapter 6 above, these equations were also fitted by the two-stage least-squares method using principal components (2SLSPC) to take account of the endogeneity in the Mark III Model of the current money and export shocks $\hat{M}$ and $\hat{x}$. The results reported here differ little from those 2SLSPC results, and certain bugs in the TROLL system make them more useful for sensitivity analyses.[8]

Examining the results in table 9.1, we can first observe that, with the exception of the United States, the explanatory powers of the regressions are very weak: Only 10 to 20% of the residual variance around the mean growth rate of real income is explained for five countries, and for France and Japan less than 10% is explained.[9] While some of the individual t statistics would be quite significant given the maintained hypothesis that all the other variables belong in the regression, this is less true for groups of coefficients. Table 9.2 reports F statistics for the null hypothesis that all the coefficients applied to a particular shock variable are zero;[10] these are reproduced from the 2SLSPC estimates reported in chapter 6. Only the

8. The basic problem is that it is impossible to recover in TROLL the sum of squared residuals based on the fitted values of the endogenous variables. Work is under way to correct this.

9. The $F(13/66)$ value of 1.634 for France is right at the border of the critical region for rejecting the null hypothesis $a_2 = c_1 = c_2 = \ldots = c_{12} = 0$ at the 0.10 significance level, while the F value of 1.491 for Japan fails even this test. For all the other countries, this null hypothesis can be rejected at the 0.05 significance level or better. It should be noted that data reliability is a particular problem for France, Italy, and Japan.

10. That is, the F for $\hat{M}$ variables is for testing the null hypothesis $c_1 = c_2 = c_3 = c_4 = 0$.

U.S. money shock variables as a group reach significance at the 0.01 level or better. In addition, British and Canadian government spending shocks reach significance at the 0.05 level while Canadian and Italian money shocks and American and German export shocks are significant at better than the 0.10 level. I conclude that in an absolute sense the explanatory power of the generalized Barro real-income equation is weak other than for the United States.

One question is whether the use of only unanticipated changes in the aggregate demand variables is consistent with the data. Table 9.3 reports the standard errors of estimate for regressions in which actual changes are substituted for unanticipated changes for each group of aggregate demand variables.[11] The form of the regression is indicated by a combination of three U's and/or A's, where U represents unanticipated and A represents actual changes and the ordering is M, g, x. Thus a UAA specification has unanticipated changes in log M for the $c_1, \ldots, c_4$ terms and actual changes in log g and x for the $c_5, \ldots, c_{12}$ terms. The main message of table 9.3 appears to be that except for the United States it makes very little difference whether one uses actual or unanticipated changes in the real-income equation specified. If we examine the minimal-sum-of-squared-residuals regression for each country, half of the cases involve money shocks, another, partially overlapping set of four have government spending shocks, and only two have export shocks. While tests on nonnested models are difficult, it is clear from the small or no increase in the SSRs for the UUU regression form as opposed to the best alternative that a null hypothesis that UUU is the correct form is not inconsistent with the data.

Since the addition of insignificant variables may increase the standard error of estimate and reduce the (corrected) $\bar{R}^2$, the real-income regressions were also run with money shocks only as suggested by Barro.[12] The results reported in table 9.4 show that the explanatory powers of all the regressions, in fact, deteriorate slightly. The last row of the table gives the $F(4/74)$ statistic for the null hypothesis $c_1 = c_2 = c_3 = c_4 = 0$; the U.S. money shocks as a group are still significant at the 1% level and the Italian at the 10% level, but now the German and Dutch money shocks are significant at the 5% level and the Canadian money shocks not at all. If these regressions are slightly encouraging for the money shock approach, the results of replacing the money shocks with the actual changes as

11. The alternative procedure of adding additional terms for anticipated changes and testing whether they belong is not feasible in this case because the estimated ARIMA processes frequently imply extreme multicollinearity. Only variables known a priori to determine anticipated money but not to belong in the real-income equation would make this alternative approach usable.

12. That is, with coefficients $c_5, \ldots, c_{12}$ in equation (5) all set equal to 0. This would follow if b_3 and b_4 were 0 in the aggregate demand equation (3) due to short-run demand-side real crowding out.

Table 9.1 **Generalized Barro Real-Income Equation (9.5)**

$$\Delta \log y = a_1 - a_2 \log(y_{-1}/y^p_{-1}) + \sum_{i=0}^{3} c_{1+i}\hat{M}_{-i} + \sum_{i=0}^{3} c_{5+i}\hat{g}_{-i} + \sum_{i=0}^{3} c_{9+i}\hat{x}_{-i} + e$$

	US	UK	CA	FR	GE	IT	JA	NE
Coefficients								
a_1	.0079	.0056	.0109	.0125	.0108	.0114	.0204	.0100
	(.0010)	(.0016)	(.0013)	(.0020)	(.0015)	(.0015)	(.0017)	(.0015)
	8.206	3.539	8.075	6.334	7.245	7.658	11.791	6.762
$-a_2$	−.0662	−.2165	−.1262	−.0590	−.0437	−.0182	.0219	−.0880
	(.0341)	(.0826)	(.0584)	(.0660)	(.0423)	(.0435)	(.0342)	(.0527)
	−1.940	−2.621	−2.160	−.893	−1.031	−.418	.641	−1.672
c_1	.6354	−.1321	.1450	−.0180	.3476	.0972	.0322	.2522
	(.2127)	(.0923)	(.1073)	(.1829)	(.1129)	(.0966)	(.1140)	(.1037)
	2.987	−1.431	1.351	−.098	3.078	1.006	.283	2.431
c_2	.6338	.0439	.1842	.1263	.0689	.0853	.1060	.1098
	(.2145)	(.0986)	(.1014)	(.1763)	(.1105)	(.0981)	(.1138)	(.1109)
	2.955	.446	1.816	.716	.623	.870	.932	.990
c_3	−.0210	−.0281	.1128	.0923	−.0187	.2630	.1883	−.0278
	(.2254)	(.0947)	(.1020)	(.1736)	(.1092)	(.1001)	(.1131)	(.1130)
	−.093	−.296	1.106	.531	−.172	2.627	1.665	−.246
c_4	.7679	−.1249	.1670	−.0900	.0387	−.0358	.0759	.0093
	(.2247)	(.0914)	(.0980)	(.1707)	(.1099)	(.1044)	(.1118)	(.1076)
	3.417	−1.366	1.704	−.527	.353	−.342	.679	.086
c_5	−.0255	.1788	.0194	.0367	−.0361	−.0019	.0489	.0370
	(.0535)	(.0528)	(.0517)	(.0396)	(.0272)	(.0103)	(.0350)	(.0348)
	−.476	3.388	.375	.925	−1.326	−.189	1.399	1.064

c_6	.1045	.0138	−.1572	.0064	.0289	.0010	−.0114	−.0352
	(.0551)	(.0563)	(.0549)	(.0397)	(.0273)	(.0102)	(.0356)	(.0354)
	1.898	.245	−2.864	.160	1.059	.100	−.319	−.992
c_7	.0388	.1003	−.0303	.0466	−.0073	−.0013	.0480	.0139
	(.0531)	(.0555)	(.0522)	(.0392)	(.0270)	(.0100)	(.0357)	(.0353)
	.732	1.807	−.581	1.189	−.272	−.125	1.347	.395
c_8	.0742	−.0224	−.0112	.0293	.0133	.0271	−.0369	.0327
	(.0544)	(.0564)	(.0529)	(.0372)	(.0271)	(.0100)	(.0362)	(.0358)
	1.363	−.398	−.212	.788	.491	2.700	−1.020	.912
c_9	.2766	.1476	.3557	.8552	.3643	−.2336	−2.0498	.0666
	(.3815)	(.1870)	(.2225)	(.3571)	(.2225)	(.1953)	(.8332)	(.0884)
	.725	.789	1.599	2.395	1.637	−1.196	−2.460	.754
c_{10}	.3674	.4153	.1001	−.7501	−.2456	−.0368	.5237	−.0727
	(.3933)	(.1791)	(.2260)	(.3730)	(.2202)	(.1890)	(.8636)	(.0899)
	.934	2.319	.443	−2.011	−1.115	−.195	.606	−.809
c_{11}	−.0021	−.2149	.0190	.0095	−.3378	−.2348	−1.7616	.1031
	(.4131)	(.1861)	(.2264)	(.3798)	(.2279)	(.1938)	(.9134)	(.0853)
	−.005	−1.155	.084	.025	−1.483	−1.212	−1.929	1.208
c_{12}	−.9809	.0067	.4470	−.0383	−.5199	−.4591	−.7920	−.1160
	(.4119)	(.1900)	(.2298)	(.3864)	(.2258)	(.1979)	(.9213)	(.0811)
	−2.381	.035	1.945	−.099	−2.302	−2.320	0.860	−1.431
$\bar{R}^2$	.3694	.1867	.1297	.0944	.1409	.1273	.0748	.1339
S.E.E.	.0086	.0140	.0119	.0175	.0133	.0131	.0153	.0130
D-W	1.72	1.95	2.41	2.13	1.94	2.23	1.99	1.67

Note. Period: 1957I–76IV. Standard errors are in parentheses below coefficient estimates; t statistics are below the standard errors.

Table 9.2 **F Statistics for Groups of Demand Shock Variables for 2SLSPC Estimates**

	$F(4/66)$ Statistics		
Country	$\hat{M}$ Variables	$\hat{g}$ Variables	$\hat{x}$ Variables
US	7.128	1.820	2.188
UK	1.164	3.531	1.763
CA	2.315	3.191	1.858
FR	0.341	0.783	1.006
GE	1.473	0.748	2.353
IT	2.201	2.004	1.766
JA	1.152	1.141	1.660
NE	1.530	1.137	1.675

Notes. The reported F statistics are appropriate for testing the joint hypothesis that all four of the demand shock variables of the type indicated have a coefficient of zero. Such a test is conditional upon the other variables entering in the equation.

For $F(4/66)$, the 10% significance level is 2.04, the 5% significance level is 2.52, and the 1% significance level is 3.63.

reported in table 9.5 are not. Again, only for the United States is there a dramatic fall in $\bar{R}^2$ or rise in S.E.E. when actual changes are substituted for unanticipated changes. Among the other seven countries it makes little difference whether actual or unanticipated changes are used, but the $\bar{R}^2$ is higher for four countries when actual changes are used. So the money-only-matters equations tell essentially the same agnostic story as the generalized Barro real-income equations.

Errors in the independent variables are an obvious explanation for the poor explanatory power of the money and other shocks. These errors might arise from the fact that the shocks are based on constructed expectations series or from the apparent fact that the data for the other seven countries have larger measurement errors than are present in the United States data.

Table 9.3 **Standard Errors of Estimate for Alternative Real-Income Equation Specifications**

Country	Specification of Aggregate Demand Variables							
	UUU	AUU	UAA	AAA	UAU	AAU	UUA	AUA
US	0.0086	0.0098	0.0083	0.0091	0.0086	0.0098	0.0083	0.0091
UK	0.0140	0.0141	0.0133	0.0134	0.0138	0.0139	0.0134	0.0136
CA	0.0119	0.0117	0.0118	0.0116	0.0119	0.0117	0.0119	0.0116
FR	0.0175	0.0175	0.0176	0.0175	0.0174	0.0174	0.0176	0.0176
GE	0.0133	0.0133	0.0133	0.0134	0.0133	0.0133	0.0133	0.0134
IT	0.0131	0.0130	0.0132	0.0131	0.0132	0.0131	0.0131	0.0130
JA	0.0153	0.0153	0.0151	0.0150	0.0154	0.0153	0.0150	0.0149
NE	0.0130	0.0131	0.0126	0.0126	0.0128	0.0129	0.0127	0.0126

Table 9.4 **Basic Barro Real-Income Equation**

$$\Delta \log y = a_1 - a_2 \log(y_{-1}/y^P_{-1}) + \sum_{i=0}^{3} c_{1+i}\hat{M}_{-i} + e$$

	US	UK	CA	FR	GE	IT	JA	NE
Coefficients								
a_1	.0080	.0061	.0107	.0125	.0108	.0114	.0208	.0101
	(.0010)	(.0017)	(.0014)	(.0021)	(.0015)	(.0015)	(.0018)	(.0015)
	7.906	3.554	7.608	6.010	6.988	7.385	11.767	6.813
$-a_2$	−.0756	−.1946	−.1307	−.0918	−.0213	−.0151	.0120	−.1010
	(.0330)	(.0757)	(.0574)	(.0692)	(.0403)	(.0420)	(.0341)	(.0515)
	−2.286	−2.569	−2.276	−1.328	−.527	−.360	.353	−1.961
c_1	.7694	−.0629	.0131	−.1204	.3534	.1627	.1022	.3095
	(.2034)	(.0911)	(.0977)	(.1828)	(.1117)	(.0966)	(.1081)	(.0972)
	3.782	−.690	.134	−.659	3.163	1.685	.946	3.184
c_2	.6413	−.0057	.0960	.1194	.1024	.1305	.0843	.0812
	(.2026)	(.0936)	(.0961)	(.1817)	(.1107)	(.0964)	(.1060)	(.0946)
	3.165	−.061	.999	.657	.925	1.353	.795	.859
c_3	.1868	−.0321	.0768	.0862	−.0470	.2132	.2033	.0114
	(.2138)	(.0934)	(.0962)	(.1785)	(.1107)	(.0966)	(.1055)	(.0955)
	.874	−.343	.798	.483	−.424	2.206	1.927	.120
c_4	.7213	−.1376	.1181	−.0278	.0448	−.0557	.0773	.0045
	(.2261)	(.0931)	(.0967)	(.1775)	(.1121)	(.0974)	(.1084)	(.0938)
	3.190	−1.478	1.221	−.156	.400	−.572	.713	.048
$\bar{R}^2$	.3048	.0414	.0391	−.0215	.0749	.0567	.0234	.1111
S.E.E.	.0090	.0151	.0125	.0186	.0138	.0136	.0158	.0132
D-W	1.53	1.97	2.35	2.36	1.89	2.01	1.99	1.73
$F(4/74)^\dagger$	9.82	.68	.77	.28	2.83	2.39	1.40	2.67

Note. Period 1957I–76IV. Standard errors are in parentheses below coefficient estimates; t statistics are below the standard errors.

†The $F(4/74)$ statistic tests the null hypothesis that $c_1 = c_2 = c_3 = c_4 = 0$. Critical values are 2.03 (10% significance level), 2.43 (5%), and 3.61 (1%).

Table 9.5 **Actual Money Growth Real-Income Equation**

$$\Delta \log y = a_1 - a_2 \log(y_{-1}/y^P_{-1}) + \sum_{i=0}^{3} c_{1+i} \Delta \log M_{-i} + e$$

	US	UK	CA	FR	GE	IT	JA	NE
Coefficients								
a_1	.0026	.0070	.0068	.0095	.0057	.0049	.0100	.0061
	(.0026)	(.0025)	(.0024)	(.0059)	(.0040)	(.0047)	(.0058)	(.0032)
	.998	2.789	2.829	1.612	1.441	1.039	1.716	1.897
$-a_2$	−.0356	−.1836	−.1404	−.0813	−.0175	−.0058	.0106	−.1013
	(.0389)	(.0768)	(.0569)	(.0688)	(.0414)	(.0422)	(.0350)	(.0514)
	−.916	−2.391	−2.467	−1.181	−.421	−.137	.303	−1.969
c_1	.4805	−.0629	−.0370	−.0198	.3431	.1925	.1084	.3032
	(.2291)	(.0923)	(.0839)	(.1700)	(.1111)	(.0944)	(.1064)	(.0930)
	2.097	−.681	−.441	−.116	3.090	2.039	1.019	3.261
c_2	.0555	.0079	.1442	.1712	.0708	.1119	.0767	−.0373
	(.2836)	(.0961)	(.0864)	(.1745)	(.1048)	(.0943)	(.1136)	(.0915)
	.196	.082	1.669	.981	.676	1.186	.675	−.408
c_3	.0329	.0400	.0272	−.0068	−.1166	.0978	.1247	−.0828
	(.2851)	(.0963)	(.0868)	(.1738)	(.1049)	(.0947)	(.1121)	(.0926)
	.115	.415	.313	−.039	−1.112	1.033	1.112	−.893
c_4	−.0560	−.0665	.0919	−.0354	−.0726	−.2112	−.0347	−.0251
	(.2402)	(.0941)	(.0831)	(.1646)	(.1147)	(.0948)	(.1067)	(.0905)
	−.233	−.707	1.105	−.215	−.633	−2.227	−.325	−.278
$\bar{R}^2$	.0505	.0192	.0734	−.0218	.0699	.0790	.0366	.1120
S.E.E.	.0105	.0153	.0123	.0186	.0138	.0134	.0157	.0132
D-W	1.21	2.00	2.32	2.38	1.89	2.08	2.02	1.73
$F(4/74)$†	2.24	.25	1.48	.28	2.72	2.89	1.68	2.69

Note. Period 1957I–76IV. Standard errors are in parentheses below coefficient estimates; t statistics are below the standard errors.

†The $F(4/74)$ statistic tests the null hypothesis that $c_1 = c_2 = c_3 = c_4 = 0$. Critical values are 2.03 (10% significance level), 2.43 (5%), and 3.61 (1%).

Consider first the extremely limited information set (past values of the variable only) used to divide log M, log g, and x into expected and unanticipated components. If the true expectations are based on a broader information set, the actual change might be as good as or a better measure of the unanticipated change than our ARIMA innovation. To investigate this question, I constructed transfer function estimates of expected money using Nelson's TRANSEST program applied to the variables appearing in the Mark III Model's money supply reaction function: the inflation rate, $\log(y/y^P)$ or unemployment rate, $\hat{g}$, and, except for the U.S., the scaled balance of payments. However, in six cases out of eight, the univariate ARIMA processes resulted in lower SSRs that these transfer estimates.[13] Further, chapter 6 reports on checks of correlations (among others) of the residuals of the real-income equations with the residuals of all the other domestic equations and of the reserve-country (U.S.) nominal money, real-income, and price-level equations. There was no apparent pattern of significant correlations which might suggest other variables for expectations transfer functions; so the approach was not pursued. It may be rational for individuals not to use costly information even if it has some predictive value (see Darby 1976 and Feige and Pearce 1976), but this may constitute some evidence against the costless-information interpretation of rational expectations.

Appeals to measurement error, like appeals to patriotism, have a deserved reputation as a last resort of scoundrels. Nonetheless, in any particular case they may be correct. Measurement error in the dependent variables ($\Delta \log y$) could account for the generally low explanatory power of the regressions and significance levels of the explanatory variables.[14] There may be greater danger of measurement error in the independent variables in general and in money in particular. Table 9.6 presents the standard deviations around the mean of each of the shock variables plus the dependent variable. For each of the independent variables, the U.S. standard deviation is only about one-third of the average standard deviations for the seven countries, but for the dependent variable the U.S. standard deviation is about three-quarters of the mean for the other countries.

Now there are good reasons why money shocks in nonreserve countries

13. The six out of eight dominance of univariate expectations occurred in the UUU regressions; in one case (France) the use of transfer expectations shifted the minimum-SSR regression from the AAU to the UAU form.

14. Measurement error in the dependent variable if it is uncorrelated with measurement error in the independent variables does not bias the coefficients but does increase s^2 (the S.E.E.). It might be that measurement error due to deflation would cause a spurious positive relation to appear between $\Delta \log y$ and $\hat{g}$ while measurement error in nominal income might create a spurious negative relation between $\Delta \log y$ and $\hat{x}$. Such a hypothesis would find some support in the estimates reported in table 9.1.

Table 9.6 **Standard Deviations of Real-Income Growth and Shock Variables**

	Standard Deviation of			
Country	$\hat{M}$	$\hat{g}$	$\hat{x}$	$\Delta \log y$
US	0.0052	0.0197	0.0028	0.0108
UK	0.0188	0.0329	0.0090	0.0155
CA	0.0146	0.0283	0.0066	0.0128
FR	0.0119	0.0566	0.0059	0.0184
GE	0.0139	0.0562	0.0070	0.0143
IT	0.0159	0.1583	0.0083	0.0140
JA	0.0168	0.0537	0.0023	0.0159
NE	0.0155	0.0448	0.0196	0.0140

would be greater than in the reserve country.[15] Suppose that nonetheless we assume that all of the difference between the standard deviations of $\hat{M}$ for the United States and the average of the other countries is accounted for by a normally distributed error component. Table 9.7 illustrates for ten drawings what such a measurement error does to the summary statistics and money shock F statistic for the regression estimates of equation (9.5).[16] Certainly the range of reported summary statistics is similar to that for the other countries appearing in tables 9.1 and 9.2. The means of the ten drawings are very similar to the means for the other seven countries noted at the bottom of table 9.7.[17] Thus an assumption that all the differences in the standard deviations of $\hat{M}$ across countries are due to measurement error is sufficient to account for the weak results observed for countries other than the United States. Doubtless other more reasonable assumptions as to measurement errors would do likewise. While this is no proof that measurement errors are the reason for the weak results outside the United States, it is evidence that measure-

15. For example, under the strictest version of the monetary approach to the balance of payments and under the assumption of independence of the sources of shocks, the variance of a nonreserve country's money shocks would equal the sum of the variance of the reserve country's money shocks, the variance of changes in the purchasing power parity, and the variance of the disturbance term to the money-demand equation.

16. That is, table 9.7 reports regressions for the United States where $\hat{M}$ is replaced with $\tilde{M} = \hat{M} + N$ where N is a computer-generated normal deviate with mean 0 and standard deviation 0.014389. To explore sampling variation, ten different drawings of N were made with the regressions computed for each one. Alternatively, an analytical examination of biases based on an assumed variance-covariance matrix of the errors might be pursued as suggested by Garber and Klepper (1980).

17. This similarity is also apparent in the (unreported) individual coefficients and t statistics. Note that the two mean S.E.E.'s are in the same ratio as the standard deviations of $\Delta \log y$ for the U.S. and the other countries. The hint of negative autocorrelation implicit in the nonreserve countries' mean Durbin-Watson statistic of 2.29 is consistent with greater measurement error in the *level* of $\log y$ which would induce negative autocorrelation in $\Delta \log y$. By construction, autocorrelation due to measurement error is removed from the shock variables.

Table 9.7 **Summary Statistics for United States Generalized Barro Real-Income Equation with Artificial Money Shock Measurement Error**

Drawing Number	$\bar{R}^2$	S.E.E.	D-W	$F(4/66)^{\dagger}$	Std. Dev. of $\hat{M}$
1	0.1613	0.0099	1.60	1.885	0.0152
2	0.1265	0.0101	1.60	1.153	0.0146
3	0.1744	0.0098	1.63	2.176	0.0130
4	0.1153	0.0102	1.52	0.928	0.0149
5	0.1397	0.0100	1.46	1.423	0.0154
6	0.1879	0.0097	1.44	2.406	0.0152
7	0.1105	0.0102	1.49	0.835	0.0149
8	0.1874	0.0097	1.61	2.474	0.0143
9	0.0827	0.0104	1.49	0.309	0.0142
10	0.1366	0.0100	1.59	1.358	0.0152
Mean	0.1422	0.0100	1.54	1.495	0.0147
Mean of other 7 countries	0.1268	0.0140	2.29	1.453	0.0153

Note. The first ten sets of summary statistics are for the U.S. equation (9.5) with $\hat{M}$ replaced by $\tilde{M} = \hat{M} + N$, where N is normally distributed with mean 0 and standard deviation 0.014389. The next line is the mean of the first ten lines, and the final line is the mean of the corresponding values from tables 9.1 and 9.2 for countries other than the United States.
†The $F(4/66)$ statistic is for the test of the null hypothesis $c_1 = c_2 = c_3 = c_4 = 0$. The critical values are 2.04 for the 10% significance level, 2.52 for 5%, and 3.63 for 1%.

ment error in the basic data[18] is a tenable defense for those who believe that the Barro approach is a correct description of the real world.

An alternative structural argument based on Lucas (1973) could be made: Countries which have larger prediction variances will be characterized by steeper aggregate supply curves. As a_3 in equation (9.1) approaches 0 so do the coefficients of $\hat{M}$, $\hat{g}$, and $\hat{x}$ as seen in equation (9.4) above. However, Lucas required huge variations in nominal income variance to detect this effect, so it would not appear to be a viable defense for the current results.

In sum, the empirical estimates indicate that, with the exception of the United States, actual and unanticipated changes in aggregate demand variables do about equally poorly as explanations of real-income growth. While these poor results may be due to measurement errors in both the dependent and independent variables, they are disappointing to support-

18. If the measurement error is not in the basic data but is instead due to the inadequacy of the expectations functions as representations of the true market expectations, then the Barro approach will not be useful even if the Lucas supply function is a true description of the economy. Leiderman (1980) reports that the rational-expectations approach to specifying expectations works for the United States. Figlewski and Wachtel (1981) and Urich and Wachtel (1980) report mixed results in reconciling survey data with rational-expectations proxies.

ers of the Barro approach to modeling the joint hypothesis of the natural-unemployment rate and rational expectations.

9.3 Implications for Economic Policymaking

In the 1960s the analytical and empirical elegance of the Phillips curve gave it wide currency as a tool for both evaluation and formulation of macroeconomic policy. It was not realized generally until the beginning of the 1970s that despite its aesthetic appeal, the Phillips curve did not work. The Lucas supply curve—particularly in the Barro reduced form—has similarly become a major tool for policy formulation and evaluation largely on the basis of a priori appeal rather than a solid foundation of empirical work. Needless to say, both the theoretical appeal and preliminary empirical work suggest that this approach is a good bet. But the results of this chapter suggest that there is less reason to adopt the approach when we examine data sets other than the one used to formulate the hypothesis. Thus policy prescriptions or evaluations which rely on the Lucas-Barro approach should be clearly labeled "Unproved; use at your own risk."

Surprising or anomalous results are our best clues to promising areas for future research. Other results casting doubt on the empirical robustness of the Lucas-Barro approach have been reported by Pigott (1978), Barro and Hercowitz (1980), and Boschen and Grossman (1980). Further research is required so that we can either use the approach with confidence or proceed to a more workable analysis.

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