SOURCES OF FLUCTUATIONS IN RELATIVE PRICES: EVIDENCE FROM HIGH INFLATION COUNTRIES

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Abstract—Casual analysis of six high-inflation episodes indicates a strong positive relationship between movements in the relative price ratio, measured by (WPI/CPI), and the inflation rate. We estimate a vector autoregression model in which relative price movements are driven by several fundamental disturbances (fiscal, monetary, output, and exchange rate), identified using only long-run restrictions based on a general-equilibrium optimizing model. Analysis of the endogenous response of relative price changes to these disturbances suggests that output and monetary shocks are the most important driving forces, although fiscal and exchange rate shocks are also influential in explaining relative price movements in some countries.

I. Introduction

INFLATION stabilization programs from Latin America to Eastern Europe almost always attempt to correct distortions in relative prices. Relative price fluctuations are often an important supply-side consideration because misalignment (or realignment) could induce large movements of capital and labor across sectors. Garber (1982) asserts that relative price movements were a source of real effects from the stabilization of the German hyperinflation, and emphasizes the strong positive relationship between the rate of inflation and the relative price ratio. He also speculates that the distortion causing the real effects resulted from increased government demand for output from the investment goods industry, which was financed by the inflation tax.[1]

In this paper we first investigate whether Garber's finding generalizes to high inflation episodes in Argentina, Bolivia, Israel, Mexico, and (1920s) France. Figure 1 plots the consumer price inflation rate (DP) and the relative price ratio in logs (Q), measured by the wholesale price index (WPI) to the consumer price index (CPI), from these five episodes and Weimar Germany. A striking stylized fact is apparent: *the relative price ratio rises during episodes of high inflation and falls during the stabilization*. Second, we analyze the sources of the observed fluctuations in relative prices. Using data from these six episodes, we estimate a structural vector autoregression model to identify the fundamental disturbances underlying relative price movements. This enables us to ascertain whether this potential source of real effects from stabilization (relative price changes) is a manifestation of fiscal, monetary, or exchange rate shocks, or is something more structural. These shocks are identified using long-run restrictions based on a general-equilibrium optimizing model.

Studying the six episodes noted above allows us to compare those that (1) took place during the 1980s versus the 1920s, (2) experienced moderately high inflation versus hyperinflation, (3) ended in immediate/rapid versus gradual/temporary stabilization, and (4) used only "orthodox" measures (fiscal austerity, monetary contraction) versus additional "heterodox" measures (wage and price controls, fixed exchange rates) as stabilizing tools, and adopted money versus the exchange rate as the nominal anchor of the stabilization.[2] We find that output and monetary

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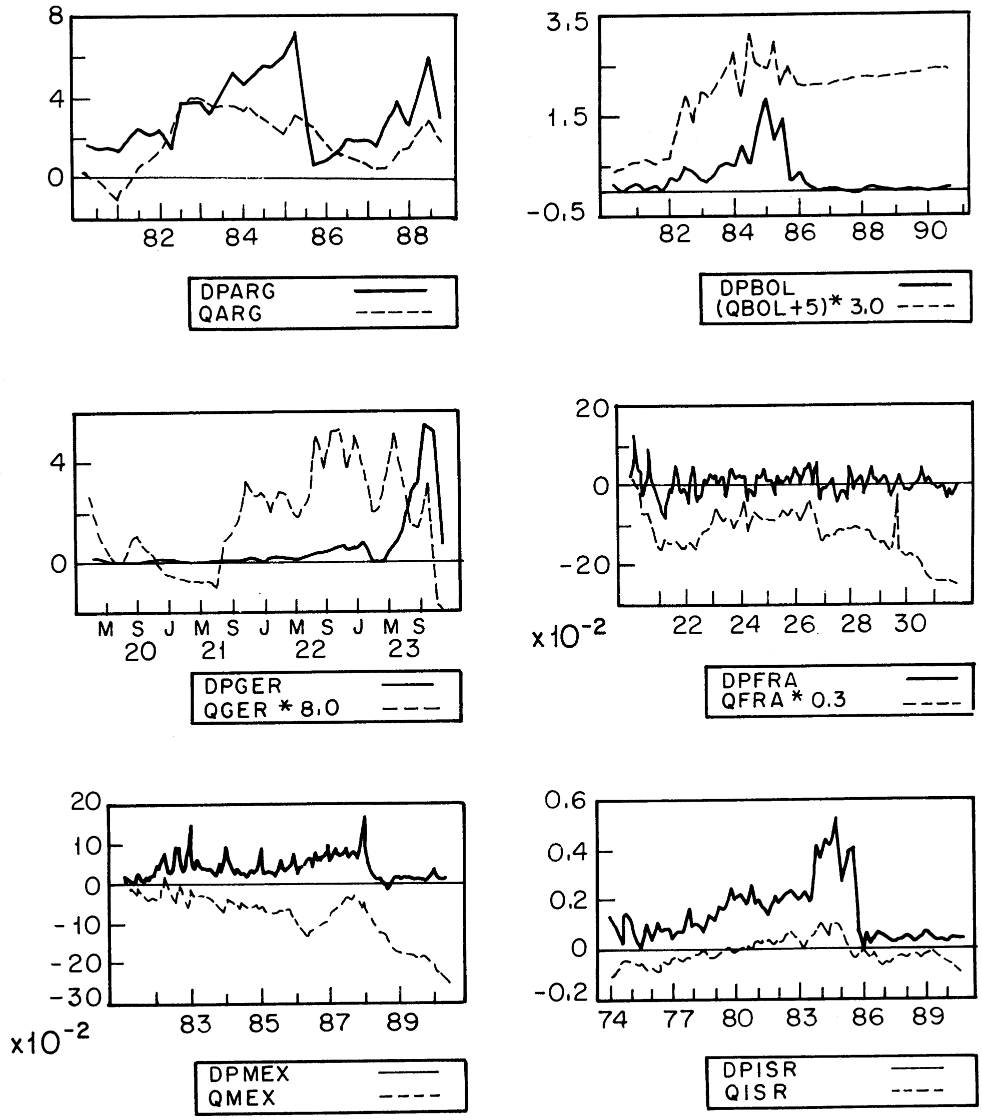
[1] In a now-classic paper, Sargent (1980) suggests that in interwar Austria, Hungary, Poland, and Germany real costs associated with inflation stabilization may have occurred as a result of structural dislocations, but he questions conventional views on the size and causes of such effects. Garber (1982) and Wicker (1986) argue that there were sizable real costs associated with these stabilization programs, while Bruno et al. (1987) show that after several stabilization programs in Latin America and Israel, there is an initial expansion followed by a recession.

[2] First, countries in the 1920s episodes are developed nations, while most in the 1980s episodes are relatively less developed. Second, hyperinflation emerged in Weimar Germany and to a lesser degree in Bolivia, but not in the other episodes. Third, the Bolivian and Israeli stabilizations succeeded instantly and have been sustained, while prior to 1991 no Argentine program succeeded. Finally, Bolivia emphasized fiscal adjustment; Argentina adopted direct wage–price freezes without changes in the monetary/fiscal regime; Israel and Mexico imposed heterdox programs involving fiscal and incomes policies; Germany relied primarily on a fiscal/monetary reform; and the French stabilization was based on establishing public confidence in the government debt. While Argentina used money as a nominal anchor, Israel and Mexico adopted policies based on an exchange rate rule. See chart 1.

FIGURE 1.—PLOTS OF INFLATION (*DP*) AND THE RELATIVE PRICE RATIO (*Q*). ARGENTINA, BOLIVIA, GERMANY, FRANCE, MEXICO, AND ISRAEL



Note: The correlation coefficient between the detrended relative price ratio and the detrended inflation rate in Argentina, Bolivia, Germany, France, Israel, and Mexico is 0.63, 0.62, 0.41, 0.32, 0.76 and 0.59, respectively.

shocks are the most important driving forces, with most movements in relative prices in the 1980s episodes being due to output shocks. Our results help disentangle the underlying causes of real distortions from stabilization.

II. Identification and Derivation of the Benchmark Model

We first discuss identification of the model to illuminate the long-run nature of the identifying restrictions.[3] We then write down the estimated model and give a theoretical justification for interpreting the structural shocks. Given widespread agreement on the importance of fiscal policy in the inflation process, and our interest in assessing the generality of Garber's analysis, it is

[3] As in Blanchard and Quah (1989), the "long run" is a hypothetical concept, defined as one such that the causal ordering is satisfied. It is not made in reference to the length of the sampling interval.

CHART 1.—SUMMARY OF HIGH INFLATION EPISODES AND ESTIMATION RESULTS

Country	Peak of Inflation	Success in Stabilizing	Real Activity Around Inflation	Main Source(s) of q Movements
Argentina	June 1985	very temporary	$\Delta y = -0.67\%$ (85:II)	
	(31% mo.)		$\Delta y = -5.39\%$ (85:III)	
		$\pi = 81.9\%$ (1986)	$\Delta y = 13.5\%$ (85:IV)	
	434% (1983)	$\pi = 175\%$ (1987)	$\Delta y = 4.10\%$ (1987)	Output shock;
	688% (1984)	$\pi = 388\%$ (1988)	$\Delta y = -5.2\%$ (1988)	Money shock.
Bolivia	June 1985	yes, immediate	$\Delta y = 33.3\%$ (85:II)	
	(66% mo.)		$\Delta y = -25.7\%$ (85:III)	
		$\pi = 14.6\%$ (1986)	$\Delta y = 15.2\%$ (85:IV)	Output shock;
	276% (1983)	$\pi = 16.0\%$ (1987)	$\Delta y = -2.9\%$ (1986)	Money shock;
	1281% (1984)	$\pi = 15.2\%$ (1988)	$\Delta y = 2.1\%$ (1987)	Fiscal shock.
France	July 1926	yes, rapid	$\Delta y = 2.7\%$ (1925)	
	(5.4% mo.)		$\Delta y = 5.7\%$ (26:II)	
		$\pi = 1.5\%$ (26:IV)	$\Delta y = 0.8\%$ (26:III)	
	10.7% (1924)	$\pi = -13\%$ (1927)	$\Delta y = -1.5\%$ (26:IV)	
	14.6% (1925)	$\pi = 14\%$ (1928)	$\Delta y = -10.2\%$ (1927)	Money shock;
	29.4% (1926)	$\pi = 3.0\%$ (1929)	$\Delta y = 16.5\%$ (1928)	Output shock.
Germany	Nov. 1923	yes, immediate	$\Delta y = 11.7\%$ (1922)	
	(24,280% mo.)		$\Delta y = -37.2\%$ (1923)	
			$\Delta y = 62.3\%$ (1924)	Money shock;
	66.5% (1921)	$\pi = 89.8\%$ (12/23)	$\Delta y = -3.0\%$ (25:III)	Fiscal shock;
	3453% (1922)	$\pi = -9.7\%$ (1/24)	$\Delta y = -8.6\%$ (25:IV)	Output shock.
Israel	July 1985	yes, rapid	$U = 6.5\%$ (85:II)	
	(28% mo.)		$U = 7.5\%$ (85:III)	Output shock;
		$\pi = 20.0\%$ (1986)	$U = 7.1\%$ (1986)	Exchange Rate
	219% (1983)	$\pi = 19.9\%$ (1987)	$U = 6.1\%$ (1987)	shock; Money
	450% (1984)	$\pi = 16.3\%$ (1988)	$U = 6.4\%$ (1988)	shock.
Mexico	8/82, 12/87	yes, gradual	$\Delta y = -0.6\%$ (1982)	
	(11%, 15% mo.)		$\Delta y = -4.2\%$ (1983)	
			$\Delta y = 3.6\%$ (1984)	
	29% (1981)	$\pi = 99\%$ (1982)	$\Delta y = -3.7\%$ (1986)	Output shock;
	81% (1983)	$\pi = 59\%$ (1984)	$\Delta y = 1.6\%$ (1987)	Exchange Rate
	159% (1987)	$\pi = 45\%$ (1988)	$\Delta y = 1.4\%$ (1988)	shock; Money
	20% (1989)	$\pi = 24\%$ (1990)	$\Delta y = 2.9\%$ (1989)	shock.

Note: π denotes the rate of inflation; Δy denotes the rate of real output growth (industrial production for France and Germany; U denotes the rate of unemployment. For π and Δy, a rate quoted for a given month is a monthly rate, for a quarter is a quarterly rate, and for a year is an annual rate.

imperative to specify fiscal shocks. We also specify monetary, output, and exchange rate shocks, using an illustrative theoretical model. 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. Nonetheless, because there may be objections to our ordering of variables in the benchmark model, we check the sensitivity of the results by considering three alternative specifications in section III.E.

A. *Identification*

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

$$X_t = C(L)\epsilon_t \tag{1}$$

where C is a non-singular matrix of coefficients, and L denotes the lag operator. An estimatible reduced form of the structural system is given by

$$\Gamma(L)DX_t = \Phi X_{t-1} + \epsilon_t^*. \tag{2}$$

Assuming that the long-run moving average coefficient matrix, $C(1)$, is lower-triangular and that the elements of ϵ are mutually uncorrelated, we can follow the procedure developed by Blanchard and Quah (1989) and extended by Ahmed, Ickes, Wang, and Yoo (1993) to retrieve the structural coefficients from the reduced form. This recovering process is unique as long as the signs of the diagonal elements of $C(1)$ are fixed by the theoretical model derived below.

B. The Estimated Long-run Model

Let the vector of stationary variables $X = \{D(g/y), D\ln(m/q), D\ln(q), D\ln(s)\}$ and the vector of shocks $\epsilon = \{\epsilon^g, \epsilon^m, \epsilon^y, \epsilon^s\}$. Here g denotes the real value of government spending, y is real output, m is real money balances, q is the ratio of wholesale to consumer prices, and s is the nominal exchange rate in domestic currency per U.S. dollars. The four structural shocks are, in order, permanent disturbances to government size, real money balances, the relative price ratio, and the nominal exchange rate. These will be interpreted as fundamental shocks based on the theoretical model below.[4] Proper interpretation of the shocks is crucial to understanding the implications of our results, and so we provide a detailed interpretation of them. The four transformed variables are assumed to be related *in the long run* as follows:

$$\begin{bmatrix} D(g/y) \\ D\ln(m/q) \\ D\ln(q) \\ D\ln(s) \end{bmatrix} = \begin{bmatrix} \bar{g} \\ \bar{m} \\ \bar{q} \\ \bar{s} \end{bmatrix} + \begin{bmatrix} c_{11} & 0 & 0 & 0 \\ c_{21} & c_{22} & 0 & 0 \\ c_{31} & c_{32} & c_{33} & 0 \\ c_{41} & c_{42} & c_{43} & c_{44} \end{bmatrix} \begin{bmatrix} \epsilon^g \\ \epsilon^m \\ \epsilon^y \\ \epsilon^s \end{bmatrix} \tag{3}$$

where $\bar{g}$, $\bar{m}$, $\bar{q}$, and $\bar{s}$ are constant and independent of the structural shocks.

We deduce the theoretically expected signs of the elements of the four-by-four matrix, $C(1)$, from the illustrative model that follows, but first give a general justification for its lower triangularity. The zero-restrictions in the top row and fourth column reflect assumptions that the long-run share of government spending in total output is exogenous and that the (nominal) exchange rate shock is neutral in the long run, respectively. The final zero-restriction, in row two and column three of $C(1)$, implies that real money balances are unaffected by the technology shock in the long run. This will hold given (1) any linear relationship between real money (M/CPI) and output, and (2) the value of relative price, (WPI/CPI), obtained from maximizing profit using a production function which is subject to any multiplicative shock (such as (4) below).[5] The lower triangularity of $C(1)$ and the assumed orthogonality property of the shocks enables us to identify thefundamental disturbances.

Two final notes are in order. First, c_{31}, which measures the long-run effect of fiscal shocks on the (change in the) relative price ratio, is a measure of the channel described by Garber (1982). Second, our method of identification imposes no restrictions on the short-run movements of the variables. Instead, we allow the data to determine the short-run dynamics.

C. Interpreting the Shocks: An Illustrative Long-run Model

In order to illustrate how the residuals in our system may be interpreted as fundamental structural shocks, we develop the following discrete-time, rational-expectations model with infinitely lived firms and consumers. We choose this framework for the benchmark model because it is tractable and not implausible in mimicking long-run optimizing behavior in a monetary economy, while noting that other models could also serve this expository purpose.

A nondurable final consumption good c is produced using an intermediate capital good x. Let

[4] In particular, we interpret the shock associated with the relative price ratio as an output shock, following a procedure similar to Blanchard and Quah (1989). Blanchard and Quah estimate a VAR model for two variables, the growth rate of real GNP and the unemployment rate, assuming the latter is $I(0)$. They identify the model by postulating that only one of the two shocks in the system has an effect on both variables in the long run (or, equivalently, that one of the variables is affected by both shocks in the long run and the other variable is affected by only one shock). The disturbance having an effect on both variables in the long run is an "aggregate supply" shock, while the shock affecting only unemployment in the long run is an "aggregate demand" shock, which is restricted to be uncorrelated with the supply shock. They present a formal theoretical model as an illustrative example to justify such an identification restriction. Thus, their ordering is $(\Delta y, u)$,with corresponding shocks labeled *AS* and *AD*.

[5] In the illustrative model below, we use a cash-in-advance constraint to model money demand. However, this is more than necessary to identify the $c_{23} = 0$ restriction, as the same restriction could be derived from any monetary model consistent with the quality theory. Nonetheless, in checking the robustness of our results, we relax this final zero restriction in two of our alternative models.

P denote the price of the final consumption good and q the relative price of the intermediate capital good to the final consumption good. Let y_t denote the representative firm's production at period t:

$$y_t = a_t x_t^{\alpha} \tag{4}$$

where a is a technology shock and $0 < \alpha \le 1$ gives us a production function with non-increasing returns. For simplicity, assume that the capital good is fully depreciated after being used in production. Then at each period t the representative firm's optimization problem can be specified as

$$\max_{\{x,y\}} (y_t - q_t x_t) \tag{5}$$

subject to (4), for all t. The resulting profit $v_t = y_t - q_t x_t$ is redistributed lump-sum to consumers.

The representative consumer is endowed with a finite, positive amount of capital stock z_0. The capital stock-flow evolution is governed by the storage technology (see McCallum (1983)):

$$z_t = \gamma(z_{t-1} - x_{t-1}), \quad t = 1,2,\ldots \tag{6}$$

where γ is a fixed storage parameter between zero and infinity. The consumer holds money M_t at the *beginning* of period t. Denote real money balances $m_t = M_t/P_t$, and the inflation rate from $t-1$ to t, $\pi_t = (P_t/P_{t-1}) - 1$. Thus, $M_{t+1}/P_t = (1 + \pi_{t+1})m_{t+1}$. The periodic budget constraints are

$$\begin{aligned} c_t + [(1 + \pi_{t+1})m_{t+1} - m_t] \\ \le (1 - \tau_t)(q_t x_t + v_t) + b_t, \\ t = 1,2,\ldots \end{aligned} \tag{7}$$

where τ is the tax rate and b is real (lump-sum) money transfer from the government. Agents do not derive utility from money, but money is required for purchasing the final consumption good. Thus,

$$c_t \le m_t, \quad t = 1,2,\ldots \tag{8}$$

This is analogous to Lucas' (1980) cash-in-advance constraint. Furthermore, so as to solve explicitly for a rational-expectations equilibrium, the utility function is assumed to be log-linear. Define the representative consumer's discount factor as β. Then the optimization problem can be written as,

$$\max_{\{c,x,z,m\}} E_0 \sum_{t=0}^{\infty} \beta^t \ln c_t \tag{9}$$

subject to (6)–(8) for all t. To close the model, define the new money injection process $b_t = \mu_{t+1} m_t$, and the government budget constraints $g_t = \tau_t y_t$, for all t. The goods market clearing conditions are

$$c_t = y_t - g_t = (1 - \tau_t) y_t, \quad t = 1,2,\ldots \tag{10}$$

while the market clearing conditions for the money market imply,

$$(1 + \pi_{t+1})m_{t+1} - m_t = \mu_{t+1} m_t, \quad t = 1,2,\ldots. \tag{11}$$

Manipulating the first-order condition of the firm's problem and utilizing (4), (7) and (10), we obtain

$$c_t = (1 - \tau_t) y_t = m_t = (1 - \tau_t) a_t x_t^{\alpha}, \quad t = 1,2,\ldots \tag{12}$$

$$(1 - \tau_t) q_t / m_t = \alpha / x_t, \quad t = 1,2,\ldots. \tag{13}$$

Combining the consumer's first-order conditions with (11) yields the no arbitrage conditions:

$$\begin{aligned} E_{t-1}\{\beta(1 - \tau_t) q_t \gamma / [(1 + \mu_{t+1}) m_t]\} \\ = \{(1 - \tau_{t-1}) q_{t-1} / [(1 + \mu_t) m_{t-1}]\}, \\ t = 1,2,\ldots. \end{aligned} \tag{14}$$

Equations (12)–(14) characterize the long-run interactions between government size, (g/y), real money balances, (m/q), and the relative price ratio, (q). The closed-form solution to the model is derived in an appendix available from the authors. From the theoretical model we can now interpret these permanent disturbances to government size, real money balances, and the relative price ratio as fiscal, monetary, and output shocks, respectively. The model to be estimated, equation (3), augments the illustrative example above to include a nominal exchange rate shock that is assumed to be neutral in the long run (recall that all other variables in the VAR are expressed in real terms):

$$\begin{bmatrix} D(g/y) \\ D\ln(m/q) \\ D\ln(q) \\ D\ln(s) \end{bmatrix} = \begin{bmatrix} \bar{g} \\ \bar{m} \\ \bar{q} \\ \bar{s} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 \\ -1 & -1 & 0 & 0 \\ 0 & (1-\alpha) & 1 & 0 \\ c_{41} & c_{42} & c_{43} & 1 \end{bmatrix} \begin{bmatrix} \epsilon^g \\ \epsilon^m \\ \epsilon^y \\ \epsilon^s \end{bmatrix}. \tag{15}$$

Intuitively, an expansionary fiscal shock that enlarges government size will reduce private transactions and thus require less cash. This shock generates no long-run effect on the relative price ratio in this simple set-up because government spending does not enter the production function and because there is no distortionary tax. A higher money growth rate raises inflation and hence the cost of holding money, thereby decreasing real money balances. Through the cash-in-advance technology, demand for consumption goods must also decline. Given diminishing returns to final goods production, the relative price of intermediate to final goods (q) subsequently rises, even though the former are not subject to the cash-in-advance constraint. Finally, a technological improvement increases the output of final consumption goods for a given input of intermediate capital goods, thereby increasing the relative price of intermediate to final goods.

Note that if purchases of firms are also subject to a cash-in-advance constraint, if government transfers were made at the *beginning* of the period, or if permanent inflationary finance is allowed, c_{21} will differ from -1. Also, if government spending has a real effect on production, c_{31} will be non-zero. It is important to assume the latter in order to assess the importance of the mechanism outlined by Garber. Moreover, if the cash-in-advance constraint is not binding, or if variable velocity is considered, c_{22} will not be unity. Furthermore, our output shock ϵ^y contains any technological disturbances, oil price shocks, enhancement of human capital, *and* any persistent aggregate demand disturbances that are not captured by changes in fiscal, monetary, and exchange rate policies. Finally, if price or wage controls have any long-run real effects, they are also captured by the ϵ^y shock. For example, a freezing of final goods prices (wages) can be thought of as a positive (negative) output shock, which would increase (decrease) the relative price ratio q. Importantly, *none* of the relaxations or generalizations discussed above will alter our identification, because such considerations only alter the specific values of the lower-triangular elements of $C(1)$.[6]

Because the inclusion of the nominal exchange rate is somewhat *ad hoc*, we discuss the expected signs of the first three elements in row four:

(1) c_{41}, the long-run effect of ϵ^g on $D\ln(s)$, could be negative or positive, depending on whether the fiscal shock is felt more on real interest rates or on expected inflation. In a short-run Mundell-Fleming model with high or perfect (low or zero) capital mobility the sign would be negative (positive). Our model is explicitly long run because of the way we identify it, so increased government size likely leads to a long-run nominal depreciation through higher expected inflation.

(2) c_{42}, the long-run effect of ϵ^m on $D\ln(s)$, is expected to be positive. This effect of a domestic money supply shock arises in a wide variety of models of exchange rate determination (see Frankel (1984) for a survey and synthesis of different monetary and portfolio models).

(3) c_{43}, the long-run effect of ϵ^y on $D\ln(s)$, is expected to be negative, as higher domestic output leads to increased demand for domestic money and a nominal appreciation.

In the estimation we do *not* impose any of the over-identifying restrictions implied by the benchmark model as seen in (15). Instead, we emphasize only the lower-triangularity of the system, which enables us to make a long-run causal ordering to use for retrieving the structural shocks.

III. Results

We estimate the model given by (3), and perform analysis of impulse response functions (IRFs) and variance decompositions (VDCs) to study the dynamic effects of fiscal, monetary, output, and exchange rate shocks on the relative

[6] A separate issue concerns formulating the VAR in a way that would include output as a separate variable. We emphasize that the focus of our paper, motivated by Garber's findings, is to examine fluctuations in relative prices so that including output separately is beyond the scope of the paper. But suppose that output were to be included as a separate variable in our VAR. The shocks associated with either output or the relative price ratio could be labelled a "real disturbance" (or a broadly-defined "technology shock" which allows more output with the same inputs). In this case it would be difficult to find an identification scheme that enables us to separate the two real disturbances, implying that all we could do is study the *joint* effect of these two shocks (in a way similar to Shapiro and Watson's two aggregate demand shocks). Given this, adding one more variable would not help to understand additional sources of relative price fluctuations. It would reduce degrees of freedom, however, given five additional parameters to estimate.

CHART 2.—SUMMARY OF DATA

Country/ Variable	Argentina	Bolivia	France	Germany	Israel	Mexico
G	Nom. gov't exp'ture; line 82z[a]	Nom. gov't exp'ture; Bol. thous	Nom. compensations at Paris; table XX	Gov't size, (g/y), is $\ln(D_t/D_{t-1})$; D is gov't debt outside the Reichsbank; table 20[f]	Nom. gov't cons'ption; line 91f[h]	Nom. gov't exp'ture; pesos bil.
y	Real GDP; 1978 base; line 99bp	Real GDP; 1980 base[d]	Ind. prod.; 1913 base; table I		Real GDP; 1978 base; line 99bp	Ind. prod; 1980 base[i]
M	Nom. M1; line 34[b]	Nom. M1; Bol. hund.	Notes circ. table XXII	Base money; table II	Nom. M1; line 34	Nom. M1; pesos bil.
P	CPI; 1980 base; line 64	CPI; 1980 base	Retail PI; 7/14 base; table XIII	Cost-of-living PI; 1913/14 = 1; table 4	CPI; 1980 base; line 64	CPI; 1980 base
q	Ratio of WPI, line 63[b], to P	Ratio of WPI to P; 1980 base	Ratio of WPI to P; 7/14 base	Ratio WPI/P table 1	Ratio ind. products PI line 63, to P	Ratio of PPI[j] to P
s	Nom. exch. rate, au/$; line rf	Nom. exch. rate, bol. per $	Nom. exch. rate, fr/$; table XXIV	Nom. exch. rate, Mk/$; table 1	Nom. exch. rate, NS/$; line rf	Nom. exch. rate, pes. per $
Dates	80:I-88:IV[c]	80:I-90:IV	1/20-12/31[e]	1/21-7/23[g]	73:IV-90:IV	1/81-6/90
Source	IFS	C. B.	SGF	H/SR	IFS	C. B.

Notes: IFS denotes the IMF's *International Financial Statistics*; C. B. denotes the regular publications of that nation's central bank (*Bulletin of the Central Bank of Bolivia* (Bolivia) and *Economic Indicators* (Mexico)); SGF denotes *Statistique Generale de la France*, 1932; H/SR denotes that the data were taken from Holtfrerich's book, which is based on the *Statistiches Reichsamt*. The series in the VAR are created as follows: (1) Gov't size (g/y) where we obtain g by converting G to index form and deflating by a simple geometric average of the two price indexes—in each case, our conversions are such that g and y have the same base year; the exception to this is Germany, as noted above; (2) Money-relative price ratio, $\ln(m/q) = \ln[(M/P)/q)]$; (3) Relative price ratio, $\ln(q)$; and (4) Nominal exchange rate, $\ln(s)$.

[a] Data are unavailable before 1980:I.
[b] Data are unavailable after 1990:I.
[c] The sample is cut at 1988:IV because in 1989:I hyperinflation re-emerged and continued through 1991.
[d] Data are unavailable before 1980:I.
[e] Data are unavailable after 1931 (from this source); we begin the sample period in 1920 because of WWI.
[f] Relevant output data are unavailable before 1/24.
[g] The sample is cut at 7/23 because non-stationarities are induced in the data by including periods up to 12/23.
[h] Data are unavailable before 1973:IV.
[i] Data are unavailable after 6/90.
[j] Data are unavailable before 1/81.

price ratio. Chart 2 summarizes the data. Germany suffered by far the worst inflation (a true hyperinflation), Bolivia's inflation rate was next highest, while the experiences of Argentina and Israel approached hyperinflation for a short period. Also note that our sample periods include both high inflation and stabilization data.

A. Unit Roots and Cointegration

Before proceeding with the VAR analysis (3), we first present evidence that the vector of variables $X = \{D(g/y), D\ln(m/q), D\ln(q), D\ln(s)\}$ is stationary, and that there is no cointegration between the levels of these variables.

For the unit roots tests, we begin using two measures: (i) both τ^{μ} and τ^{τ} Augmented Dickey-Fuller (ADF) tests to test the null hypothesis of I(1) vs. I(0), and again to test I(2) vs. I(1); and (ii) 95% confidence intervals for the largest autoregressive root (which are constructed from the ADF statistics using Stock's (1991) procedure). As seen in table 1, the ADF test statistics indicate some evidence that perhaps in 3 or 4 out of 24 cases the variables deviate from the I(1) specification used in the VAR analysis. In particular, there is evidence that (g/y) is I(0) for Bolivia, and to a lesser extent France (rejection of the unit root null at 5% in the Israeli case occurs only when the time trend is omitted from the

TABLE 1.—UNIT ROOT TESTS

	(g/y)	$\ln(m/q)$	$\ln(q)$	$\ln(s)$	$D(g/y)$	$D\ln(m/q)$	$D\ln(q)$	$D\ln(s)$
ADF τ_τ	−2.77	−3.15	−2.06	0.29	−5.08[g]	−5.17[g]	−5.92[g]	−4.36[g]
95% CI	0.57 1.07	0.48 1.05	0.77 1.08	1.03 1.10	— 0.55	— 0.30	— 0.30	— 0.75
ADF τ_μ	−1.39	−0.97	−2.23	2.17	−5.04[g]	−5.18[g]	−5.79[g]	−3.68[g]
95% CI	0.84 1.08	0.90 1.08	0.69 1.05	—	— 0.52	— 0.47	— —	— 0.85
				Bolivia				
ADF τ_τ	−5.73[g]	−1.03	−1.24	−1.32	−7.77[g]	−4.56[g]	−7.06[g]	−2.84[a]
95% CI	— 0.16	0.92 1.13	0.89 1.12	0.87 1.12	— —	— 0.60	— —	0.44 1.09
ADF τ_μ	−4.37[g]	−2.02	−2.26	1.23	−7.87[g]	−4.09[g]	−6.59[g]	−2.76[ib]
95% CI	— 0.58	0.64 1.08	0.56 1.07	0.83 1.11	— —	— 0.68	— —	0.39 1.03
				France				
ADF τ_τ	−3.59[h]	−0.24	−1.41	−1.55	−5.24[g]	−4.60[g]	−5.07[g]	−5.67[g]
95% CI	0.74 1.01	1.01 1.04	0.95 1.04	0.94 1.04	— 0.81	— 0.88	— 0.82	— 0.75
ADF τ_μ	−3.22[h]	0.40	−0.74	−1.44	−5.26[g]	−4.41[g]	−5.06[g]	−5.64[g]
95% CI	0.77 0.99	1.002 1.04	0.97 1.03	0.94 1.03	— 0.78	— 0.87	— 0.80	— 0.74
				Germany				
ADF τ_τ	−1.99	−2.17	−2.69	0.40	−10.1[g]	−4.31[g]	−4.83[g]	−3.94[hc]
95% CI	0.68 1.12	0.62 1.11	0.45 1.10	1.05 1.15	— —	— 0.70	— 0.50	— 0.86
ADF τ_μ	−1.64	0.22	−1.26	2.67	−10.2[g]	−4.01[g]	−4.96[g]	−2.31[d]
95% CI	0.73 1.10	0.998 1.13	0.80 1.11	— —	— —	— 0.69	— 0.34	0.54 1.07
				Israel				
ADF τ_τ	−3.28[i]	−0.38	−1.83	−1.55	−10.3[g]	−4.68[g]	−5.10[g]	−2.53[e]
95% CI	0.53 1.04	1.02 1.08	0.84 1.07	0.88 1.07	— —	— 0.73	— 0.63	0.71 1.06
ADF τ_μ	−3.27[h]	−1.13	−2.11	−1.18	−10.4[g]	−4.27[g]	−4.95[g]	−2.50[f]
95% CI	0.51 0.95	0.91 1.07	0.76 1.05	0.90 1.07	— —	— 0.76	— 0.62	0.69 1.03
				Mexico				
ADF τ_τ	−3.26[i]	−0.51	−1.99	−1.85	−6.39[g]	−4.57[g]	−3.25[i]	−3.58[h]
95% CI	0.71 1.02	1.01 1.05	0.88 1.04	0.90 1.04	— —	— 0.85	0.72 1.03	0.66 1.02
ADF τ_μ	−1.74	−1.99	−0.02	−1.89	−6.37[g]	−4.02[g]	−3.05[g]	−3.22[h]
95% CI	0.89 1.04	0.86 1.03	0.99 1.05	0.87 1.03	— —	— 0.89	0.73 1.01	0.71 0.99

Notes: ADF τ_τ (τ_μ) denotes the Augmented Dickey-Fuller test statistic for the unit root null hypothesis. 95% CI denotes the 95% confidence interval for the largest autoregressive root, which are constructed from the ADF statistics using Stock's (1991) procedure. A — indicates that the calculation is not available from Stock's tables. Footnotes to selected entries in the table indicate that Phillips-Perron (1988) tests were run in these cases; the corresponding test statistic is (a) −15.2, (b) −15.7, (c) −7.61, (d) −6.82, (e) −10.9, and (f) −11.4; each of these is significant at 1%, implying a rejection of the unit root null.

[g]Rejection of the unit root null at 1%.

[h]Rejection of the unit root null at 5%.

[i]Rejection of the unit root null at 10%

regression). In addition, the ADF tests suggest that $\ln(s)$ might be I(2) for Israel, and to a lesser extent Bolivia (failure to reject the I(2) null for Germany occurs only when the time trend is omitted from the regression).

It is well known that there can be great uncertainty about the unit root properties of time series data, especially in sample sizes—or more importantly, sampling intervals—such as ours (see Campbell and Perron (1991) and Cochrane (1991)). In a recent paper, Stock (1991) quantifies this uncertainty, while maintaining the classical approach to unit roots testing, by constructing confidence intervals for the largest autoregressive root of a time series when this root is close to unity. For our data, we bring out such uncertainty, and in the process attempt to justify our I(1) specification even for the borderline cases above, in two ways. First, following Stock's procedure, we report 95% confidence intervals for each of our test statistics. Note that many of the intervals are wide. For one of our borderline cases, for example, the (g/y) data for France are consistent with the hypothesis that the process is I(1), but are also consistent with the hypothesis that the data are trend stationary with an autoregressive root of 0.74. Second, the Phillips-Perron (1988) tests corresponding to the ADF τ^{μ} and τ^{τ} tests indicate a strong rejection of the I(2) null in favor of I(1) for $\ln(s)$ in Bolivia, Germany, and Israel—the three borderline cases noted above. Hence, the most objectionable departure from the I(1) specification is for Bolivian (g/y), which appears to be I(0).[7] Overall, we conclude that it is

[7] To account for this, we removed a deterministic time trend from Bolivian (g/y) and estimated two alternative models

TABLE 2.—COINTEGRATION TESTS
A. Johansen's Trace Test

Country	$p = 0$	$p \leq 1$	$p \leq 2$	$p \leq 3$
Argentina	51.1	27.4	6.71	0.74
Bolivia	87.3[a]	44.6[a]	16.1	1.74
France	52.0	19.4	2.35	0.31
Germany	114.7[a]	29.1	14.2	0.10
Israel	49.3	31.0	15.4	4.33
Mexico	46.3	20.5	7.79	0.19
Bolivia (n = 3)	33.2	10.0	1.93	....

B. Augmented Dickey-Fuller Tests
Cointegrating Regressions (to obtain the residuals res_i):
$(g/y) = a + b\ln(m/q) + \text{res1}$;
$(g/y) = a + b\ln(q) + \text{res2}$;
$(g/y) = a + b\ln(s) + \text{res3}$;
$(g/y) = a + b\ln(m/q) + c\ln(q) + \text{res4}$;
$(g/y) = a + b\ln(m/q) + c\ln(s) + \text{res5}$;
$(g/y) = a + b\ln(q) + c\ln(s) + \text{res6}$;
$(g/y) = a + b\ln(m/q) + c\ln(q) + d\ln(s) + \text{res7}$;
$\ln(m/q) = a + b\ln(q) + \text{res8}$;
$\ln(m/q) = a + b\ln(s) + \text{res9}$;
$\ln(m/q) = a + b\ln(q) + c\ln(s) + \text{res10}$;
$\ln(q) = a + b\ln(s) + \text{res11}$

Series:	GERres1	GERres2	GERres3	GERres4	GERres5	GERres6
ADF_t	−0.93	0.16	−3.07	−3.02	−4.10[c]	−3.69
Series:	GERres7	GERres8	GERres9	GERres10	GERres11	
ADF_t	−2.99	1.39	−1.08	−3.87[c]	−0.57	
Series:	BOLres8	BOLres9	BOLres10	BOLres11		
ADF_t	−1.31	−1.62	−1.45	−1.63		

Notes: In Johansen's trace test, the null is that there are no more than p cointegrating vectors in the system (or n-p distinct unit roots). Critical values for $(n - 4)$ equal to 1,...,4 are 9.1, 20.2, 35.1, and 53.3 (95%); and 12.7, 25.0, 40.2, and 60.1 (99%). These are taken from table A.3 of Johansen and Juselius (1990). In the (n = 3) case for Bolivia, (G/Y) is omitted from the system.
ADF_t denotes the Augmented Dickey-Fuller τ_τ test statistic for the null of no cointegration among the relevant variables (i.e., a test for a unit root in the corresponding residual); a **, *, # indicates rejection of the null at 1%, 5%, and 10%.
[a]Significant at 1%.
[b]Significant at 5%.
[c]Rejection of the null at 10%.

reasonable to carry on an empirical investigation of the sources of fluctuations in relative prices predicated on the I(1) specification of these variables, but do acknowledge the not-uncommon uncertainty associated with the unit roots tests.

Next, we perform Johansen's (1988) trace test for the number of cointegrating vectors in each system. As seen in table 2, we fail to reject the null hypothesis of zero cointegrating vectors at 5% for Argentina, France, Israel, and Mexico. For Bolivia and Germany, where the test rejects the null of zero cointegrating vectors, we also undertake ADF tests to try and uncover the specific cointegrating relationships. First, consider Germany, for whom the multi-equation Johansen test indicates the presence of one cointegrating vector in the system of four variables. In

using this variable instead of $D(g/y)$. In the first we use the same ordering as the benchmark case (3), and in the second we place detrended (g/y) last (for the same reason that Blanchard and Quah place the trend-stationary unemployment rate last in their model). We find little difference from the specifications using $D(g/y)$, and so report the results in an appendix available on request. We also point out, even though this is not proof contradicting the unit root results, that several authors have found (g/y) to be I(1) for the U.S. federal government (e.g., Bohn (1991)), and Rogers and Rogers (1992) have found (g/y) to be I(1) for nearly all 50 U.S. states.

TABLE 3.—ESTIMATES OF LONG-RUN COEFFICIENTS [$C(1)$]

Coefficients (Expected Sign)	Countries					
	Argentina	Bolivia	France	Germany	Israel	Mexico
$C11$	0.89	0.46	0.62	0.48	0.67	0.33
(+)	(0.18)	(0.69)	(3.26)[b]	(0.79)	(4.38)[b]	(1.32)[a]
$C21$	−1.78	0.82	0.05	−0.19	−0.47	−0.11
(−)	(0.04)	(0.12)	(0.51)	(0.19)	(0.33)	(0.19)
$C22$	−0.91	−1.04	−0.68	−0.70	−1.24	−0.69
(−)	(1.27)[a]	(2.21)[b]	(1.10)[a]	(0.64)	(4.13)[b]	(1.89)[a]
$C31$	−2.31	−0.12	0.03	−0.13	0.07	−0.01
(?)	(0.08)	(0.13)	(0.37)	(0.21)	(0.25)	(0.13)
$C32$	0.40	0.12	0.32	0.34	0.24	0.12
(+)	(1.04)[a]	(1.24)[a]	(1.29)[a]	(0.57)	(3.26)[b]	(1.38)[a]
$C33$	1.87	0.43	1.44	0.45	0.93	1.17
(+)	(2.30)[b]	(4.30)[b]	(3.69)[b]	(2.50)[b]	(5.28)[b]	(4.38)[b]
$C41$	−17.89	0.41	0.13	1.69	2.41	0.35
(?)	(0.09)	(0.01)	(1.41)[a]	(0.30)	(0.29)	(0.32)
$C42$	1.17	4.61	0.39	0.30	2.22	1.23
(+)	(0.54)	(1.66)[a]	(1.00)[a]	(0.08)	(1.54)[a]	(1.84)[a]
$C43$	−1.46	−3.93	−0.23	2.29	−4.03	0.54
(−)	(0.45)	(2.51)[b]	(0.58)	(0.86)	(0.93)	(0.25)
$C44$	2.25	1.65	0.55	2.29	2.83	0.34
(+)	(3.57)[b]	(4.34)[b]	(6.04)[b]	(1.82)[a]	(3.26)[b]	(3.53)[b]

Notes: These numbers indicate the long-run responses of the four variables to a unit permanent increase in the shocks. The absolute values of t-statistics computed based on 1000 random draws are reported in parentheses.
[a]Significant at the 1 standard error confidence level.
[b]Significant at the 2 standard error confidence level.

part B of table 2, we use single-equation ADF tests to check all eleven possible cointegration relationships. At the 5% level, we never reject the null hypothesis of no cointegration. Now consider Bolivia, and recall that the unit roots tests indicate that (g/y) is likely to be I(0). When we omit (g/y) from the system (the $n = 3$ case in the table), Johansen's test indicates no cointegration. Augmented Dickey-Fuller tests verify that there is no cointegration between or among the three remaining variables. This suggests that, although there is some conflict between the Johansen and ADF test results, it is reasonable to undertake the VAR analysis using first-differences for all six cases, rather than using the VECM strategy of King, Plosser, Stock, and Watson (1991).

B. *Long-run Coefficient Matrix*

Estimates of the $C(1)$ matrix for all countries are generally consistent with the theoretical model. As seen in table 3, 44 of 48 coefficients have the expected sign, while the exceptions (c_{21} for Bolivia and France, and c_{43} for Germany and Mexico) are insignificant. The estimates imply a sound basis for computing the IRFs and VDCs, which give the short-run dynamic results that we focus on.

C. *Impulse Response Functions*

Table 4 and figure 2 contain the impulse responses of both $D\ln(q)$ and $\ln(q)$ to a one-standard deviation change in each of the four shocks. Numerical results for $D\ln(q)$ from all countries are in the upper part of table 4. The lower part of table 4 describes the shape of the IRF for $\ln(q)$ from all countries. As discussed below, it is instructive to analyze the *shape* of the IRF, and so it is important to display some results graphically. In figure 2 we display IRF plots for Israel only in order to save space (the others are available on request). Notationally, DQi is the response of $D\ln(q)$ to shock i, while Qi is the response of $\ln(q)$ to shock i (= 1 to 4). The shocks are numbered according to the order they appear in the VAR; thus, for example, $Q1$ is the response of $\ln(q)$ to a change in ϵ^g. The responses of $\ln(q)$ are of three types: persistent, hump-shaped, and short-lived. Persistent responses represent gradual adjustment to a shock (see $Q3$ in figure 2). Hump-shaped responses reflect quick adjustment; that is, the effects are initially large but diminish quickly (see $Q1$ and $Q2$ in figure 2). Short-lived responses imply short-run temporary effects (see

TABLE 4.—IMPULSE RESPONSES OF RELATIVE PRICE

	Periods Out	Countries					
		Argentina	Bolivia	France	Germany	Israel	Mexico
		(1) Estimates of Impulse Responses of $D\ln(q)$ to (in %)					
Govt.	1	−0.16	−1.27	−0.79	3.76[a]	0.52[a]	−0.23[a]
Size	2	−0.77	1.56[a]	1.25[b]	−3.93[a]	−0.08	0.14
Shock	3	−0.90	−2.93[a]	0.15	−3.11[a]	0.10	0.05
Money	1	1.24[a]	4.55[b]	4.84[b]	8.56[a]	−0.01	0.34[b]
Supply	2	2.67[b]	0.54	−1.62[a]	−0.45	0.84[b]	0.32[b]
Shock	3	1.09[a]	−1.66	0.02	−0.24	−0.20	0.15[a]
Total	1	3.96[b]	5.65[b]	1.79[a]	4.25[a]	1.62[b]	0.66[b]
Output	2	1.54[a]	−4.41[b]	−0.12	2.45[a]	−0.15	−0.14[a]
Shock	3	−0.03	3.29[b]	0.98	−0.43	0.24[a]	0.04
Exchange	1	1.34[a]	0.25	−0.14	2.89	0.86[a]	−0.14[a]
Rate	2	0.02	1.56	1.19[a]	0.94	0.07	0.48[b]
Shock	3	0.09	−1.89[a]	−0.77[a]	−0.65	−0.01	0.00
		(2) Shapes of Impulse Responses of $\ln(q)$ to					
Govt. Size Shock		negative but small	positive to negative	negative to positive	positive to negative	hump	negative but small
Money Supply Shock		persistent	hump	hump	persistent	hump	persistent
Total Output Shock		persistent	persistent	persistent	hump	persistent	persistent
Exchange Rate Shock		short-lived	positive but small	short-lived	positive but small	short-lived	short-lived

Notes: The numbers in the upper panel indicate the impulse responses to a one-standard-error shock. For the lower panel, persistent responses represent gradual adjustment, hump-shaped responses reflect quick adjustment, short-lived responses imply temporary short-run effects, and small responses indicate that the accumulated effects are insignificant at the one-standard-error confidence level.
[a]Significant at the 1 standard error confidence level based on t-statistics computed from 1000 random draws.
[b]Significant at the 2 standard error confidence level on t-statistics computed from 1000 random draws.

$Q4$ in figure 2). By construction, the response of $\ln(q)$ is the sum of that of $D\ln(q)$.[8]

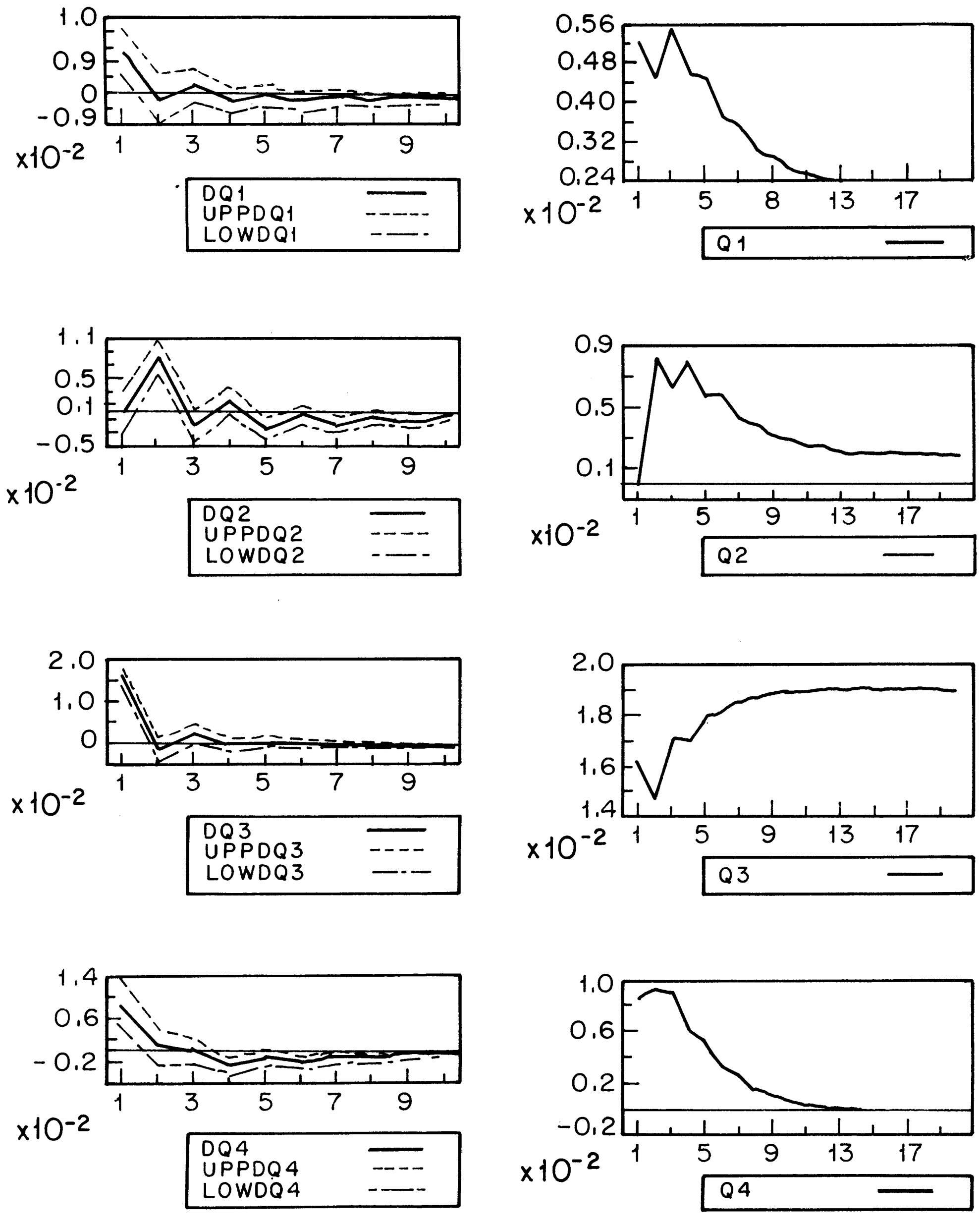
First, from the IRFs of both $\ln(q)$ and $D\ln(q)$, a positive *shock to government size* has a small and negative short-run effect on the relative price ratio in Argentina and Mexico. In the case of France, a fiscal shock has a negative (but small) impact effect, which then turns significantly positive. In Bolivia and Germany (and Israel to a lesser degree) the relative price ratio rises in the short or immediate run in response to a shock to government size, and eventually falls in the longer run. Recall from chart 1 that Bolivia and Germany had the most severe inflation of our episodes.

Second, the short-run response of the relative price ratio to a *domestic money shock* is positive in all countries. Such responses are hump-shaped in Bolivia, France, Germany, and Israel, all of which undertook stabilization programs that were successfully and rapidly completed. This result is consistent with the "snake effect" in Blanchard's (1983) theory of asynchronous price decisions. He finds that, following a positive money shock, prices of goods that are early in the chain of production (intermediate goods) respond more and adjust faster than do prices of goods later in the chain (final goods). This would imply a hump-shaped, positive response of q to the money shock. However, where stabilization was unsuccessful (Argentina) or successful but gradual (Mexico), a money shock has a very persistent effect.

Third, the *output shock* has a permanent and positive effect on the relative price ratio in all countries. This is consistent with Bils' (1987) explanation for the mark-up of final to intermediate

[8] Persistent (short-lived) responses are analogous to the response of output to the aggregate supply (aggregate demand) shock in Blanchard and Quah. Hump-shaped responses are analogous to the response of labor hours to the technology shock in Shapiro and Watson (1988).

FIGURE 2.—IMPULSE RESPONSE FUNCTIONS: ISRAEL RESPONSE OF RELATIVE PRICE RATIO (FIRST-DIFFERENCE AND LOG-LEVEL) TO (1) GOVERNMENT SIZE, (2) MONEY SUPPLY, (3) OUTPUT AND (4) EXCHANGE RATE SHOCKS



Note: The solid line is the point estimate and dashed lines are plus- and minus- one standard deviation.

goods prices over the business cycle: following a positive output shock, production expands, and firms face higher adjustment costs because of imperfectly flexible factor inputs. Consequently, the mark-up falls, and q rises. In the 1980s episodes, the response of relative prices to the output shock is the largest amongst all shocks. Except for the case of Germany, the effect of the output shock on relative price is rather persistent. Finally, shocks to the exchange rate have either a small or significantly positive, but short-lived effect on relative prices.

D. Variance Decompositions

The variance decompositions of table 5 are used to evaluate the importance of each shock in explaining relative price changes. The numbers in

TABLE 5.—VARIANCE DECOMPOSITIONS OF RELATIVE PRICE CHANGES

Periods Out	Percentage of Variance of the Rate of Change of Relative Price [$D\ln(q)$] due to a shock to: Govt. Size	Money Supply	Total Output	Exchange Rate
(1) Argentina				
1	0.1 (0.01)	8.0 (0.75)	82.4 (3.68)	9.5 (0.62)
2	2.1 (0.15)	29.7 (2.01)	62.0 (3.20)	6.2 (0.56)
4	4.7 (0.33)	31.8 (2.24)	57.3 (3.26)	6.2 (0.65)
20	4.8 (0.28)	31.5 (2.11)	56.9 (3.18)	6.8 (0.76)
(2) Bolivia				
1	3.0 (0.35)	38.1 (2.51)	58.8 (3.92)	0.1 (0.02)
2	5.1 (0.73)	26.6 (2.46)	65.1 (4.64)	3.2 (0.37)
4	13.3 (1.43)	22.6 (2.69)	58.2 (4.56)	5.8 (0.62)
20	16.4 (1.70)	22.6 (2.47)	55.9 (4.21)	5.1 (0.56)
(3) France				
1	2.3 (0.29)	85.9 (5.33)	11.7 (0.82)	0.1 (0.02)
2	6.7 (0.76)	79.2 (5.15)	9.7 (0.78)	4.4 (0.78)
4	6.7 (0.83)	74.8 (5.41)	12.4 (1.10)	6.1 (1.08)
20	7.6 (1.02)	71.1 (5.80)	12.9 (1.32)	8.4 (1.56)
(4) Germany				
1	12.4 (0.93)	64.4 (2.64)	15.9 (0.99)	7.3 (0.37)
2	21.7 (1.94)	53.9 (2.83)	17.6 (1.36)	6.8 (0.44)
4	26.4 (2.34)	49.8 (3.02)	16.4 (1.40)	7.4 (0.55)
20	27.1 (2.03)	47.6 (3.08)	18.0 (1.51)	7.4 (0.61)
(5) Israel				
1	7.6 (0.88)	0.0 (0.00)	72.0 (3.72)	20.4 (1.09)
2	6.5 (0.85)	16.1 (2.27)	60.4 (3.75)	17.1 (1.14)
4	6.6 (0.88)	16.8 (2.31)	58.6 (3.90)	18.0 (1.29)
20	6.5 (0.85)	18.0 (2.25)	56.7 (3.83)	18.9 (1.42)
(6) Mexico				
1	8.6 (0.60)	18.9 (1.25)	69.1 (3.77)	3.3 (0.44)
2	7.4 (0.75)	21.9 (2.54)	45.4 (3.67)	25.2 (2.61)
4	8.7 (1.10)	21.9 (2.90)	45.0 (4.30)	24.3 (2.71)
20	10.5 (1.46)	22.8 (3.23)	37.0 (4.59)	29.7 (3.56)

Note: The t-statistics computed based on 1000 random draws are reported in parentheses.

the table give the percentage of the variance of $D\ln(q)$ which is accounted for by a one standard-deviation change in each of the shocks.

In all countries, *output shocks* and *monetary shocks* account for most (between 60% and 97%) of the variance of $D\ln(q)$ over the forecast horizon. For the 1980s episodes output shocks are the most influential, accounting for approximately one-half of the variance of $D\ln(q)$, while monetary shocks account for one-sixth to one-third of the variance. In the French and German cases, on the other hand, the effects of output shocks become important only over longer horizons, while monetary shocks are very influential. Differences in the relative stage of development between these countries may account for the different responses to the output shock. A given shock to technology or human capital enhancement, for example, should have a larger effect in a less developed country such as Bolivia than in relatively well-developed France and Germany. Of course, in countries imposing price controls on final consumption goods, such effects, if permanent, would be captured by ϵ^y, thereby magnifying the importance of the output shock on q. It is also noteworthy that in countries with temporary or gradual stabilization (Argentina and Mexico), the importance of the monetary shock increases as the forecast horizon lengthens. These persistent effects, which are consistent with the IRFs, may reflect a diminishing of the authorities' credibility over time.

Next, the effect of a *fiscal shock* on relative prices is important in Germany, and to a lesser extent in Bolivia. The short-run impulse response is positive, while the VDCs suggest that the fiscal shock accounts for about one quarter of the variance of German relative price changes after the first month horizon. Thus, our formal estimation confirms Garber's conjecture for Germany. However, this effect is not found to be essential for any of the other five episodes we examine.

Finally, the *exchange rate shock* is an important influence on relative price movements in Mexico, as it accounts for approximately one quarter of the variance of $D\ln(q)$. Such effects also emerge to a lesser degree in Israel. Recall that the IRFs indicate that a domestic currency depreciation leads to a higher relative price ratio in Israel and Mexico, particularly over short horizons. Interestingly, Israel and Mexico each treated the exchange rate as a nominal anchor and used exchange rate control as a policy tool in their stabilization programs, while in France and Germany, a stable exchange rate was a by-product of inflation stabilization. Notably, Bolivia adopted the exchange rate as a nominal anchor (through the Bolsin, the daily auction of dollars by the central bank), but direct exchange rate freezes have not been imposed in its stabilization program. These considerations may explain why exchange rate shocks are not an important influence on Bolivian relative price changes, but are in Israel and Mexico.

E. *Sensitivity Analysis*

Because there may be objections to our ordering of variables in the benchmark model, we check the sensitivity of the results by considering three alternative specifications. There are 24 different orderings possible. However, we argue that it is highly plausible to eliminate most of them.

First, it is standard in neoclassical models (e.g., Barro 1989) to allow changes in government size (or the average income tax rate) to affect long-run movements in output. Allowing this channel is of particular importance to us because one of our main tasks is to examine Garber's hypothesis concerning the effects of fiscal policy on relative price fluctuations. Second, because the exchange rate considered is a nominal variable, placing it after the relative price ratio—a real measure—is more acceptable (Shapiro and Watson place inflation after real output for the same reason). Reversing the order of these two variables is equivalent to assuming that nominal shocks have long-run non-neutral effects while real disturbances cannot affect the nominal exchange rate permanently. We find this assumption less convincing.

Therefore, we impose the long-run causal ordering $\{D(g/y), D\ln(q), D\ln(s)\}$, leaving us with four plausible orderings, depending on where we insert real money balances.[9] Our benchmark specification places real money balances second. The three alternative specifications generate results generally consistent with those in our benchmark case, although there are a few minor differences.

The results are displayed in table 6, where we show estimates of the $C(1)$ matrices and the VDCs of $D\ln(q)$. First, consider model 1, the case where real money is placed just after the relative price ratio. With this new ordering, "output" shocks become more important compared to "monetary" shocks in driving relative price movements in Germany and France (and to a lesser extent Bolivia), according to the VDCs of table 6B. In each of these cases, however, the two shocks *combine* to contribute the same percentage as in the benchmark model. Moreover, the long-run responses indicate that a positive movement in the shock associated with the relative price ratio generally *decreases* real money demand and leads to a domestic currency *depreciation* (see estimates of C_{32} and C_{42}). These suggest that the new "output" shock is in effect a combination of real and monetary disturbances. This is because the ordering of model 1 can separate real from monetary shocks only if the snake effect discussed above is unimportant.[10]

Second, consider model 2, with real money placed last. Relative to the benchmark model, this specification implies (in addition to the change brought by model 1) that the exchange rate shock is allowed to affect long-run money demand. Interestingly, there are no qualitative or quantitative changes except for the cases of Bolivia and Mexico, where the VDCs indicate that the contribution of the exchange rate disturbance

[9] Note that when we place the real money variable ahead of the relative price ratio, it is necessary to transform the former as the ratio of real money to the relative price ratio, $D\ln(m/q)$, so that the output shock does not affect this transformed ratio in the long run. This is shown formally in our theoretical model in section II.C above.

[10] In considering the merits of the benchmark model over model 1, notice that real shocks *are* allowed to affect long-run real money demand in the benchmark model. With that ordering, we only restrict the real shock to affect real money, m, and relative prices, q, equi-proportionately. Since real shocks are shown to increase q significantly, they must also increase m.

TABLE 6.—SENSITIVITY ANALYSIS

A. Coefficients of the Long-Run Moving Average Matrix, $C(1)$, for Three Alternative Models

Coefficient	Model	Argentina	Bolivia	France	Germany	Israel	Mexico
$C_{11}(+)$	1	0.88	0.46	0.62	0.48	0.67	0.343
$(+)$	2	0.88	0.46	0.62	0.48	0.67	0.34
$(-)$	3	−0.91	−1.06	−0.71	−0.85	−1.23	−0.82
$C_{21}(?)$	1	−2.07	−0.15	0.03	−0.13	0.08	−0.003
$(?)$	2	−2.07	−0.15	0.03	−0.13	0.08	−0.003
$(+)$	3	−0.002	−0.06	−0.32	0.28	0.04	0.20
$C_{22}(+)$	1	2.26	0.62	0.51	1.12	0.98	1.61
$(+)$	2	2.26	0.62	0.51	1.12	0.98	1.61
$(+)$	3	0.89	0.46	0.61	0.49	0.66	0.31
$C_{31}(-)$	1	−6.85	5.26	2.42	−0.32	−2.10	−36.6
$(?)$	2	−19.6	0.62	0.13	1.69	2.25	0.40
$(+)$	3	0.40	0.13	0.31	0.29	0.03	0.13
$C_{32}(+)$	1	0.10	−6.59	5.05	−0.56	1.91	−830
$(-)$	2	1.90	2.30	0.15	2.24	−2.56	9.02
$(?)$	3	−3.12	−0.03	0.05	−0.23	0.06	−0.02
$C_{33}(-)$	1	−0.62	−0.83	−1.50	−0.55	−1.04	−0.50
$(+)$	2	0.50	3.31	0.62	2.79	3.14	0.72
$(+)$	3	1.87	0.43	1.44	0.45	0.92	1.17
$C_{41}(?)$	1	−19.6	0.62	0.13	1.69	2.25	0.40
$(-)$	2	−6.85	5.26	2.42	−0.32	−2.10	−36.6
$(+)$	3	1.12	4.16	0.33	1.29	2.30	1.57
$C_{42}(-)$	1	1.90	2.30	0.15	2.24	−2.56	9.02
$(+)$	2	0.10	−6.59	5.05	−0.56	1.91	−830
$(?)$	3	−20.2	4.07	0.15	1.65	1.55	0.14
$C_{43}(+)$	1	0.62	0.82	0.03	−1.12	0.34	0.003
$(-)$	2	−0.42	−2.14	−11.0	0.13	−3.08	−65.0
$(-)$	3	−1.46	−3.93	−0.23	2.29	−4.03	0.54
$C_{44}(+)$	1	2.37	1.81	0.57	2.29	2.97	0.32
$(-)$	2	−0.50	−0.70	−1.24	−0.30	−0.86	−0.54
$(+)$	3	2.25	1.65	0.55	2.29	2.83	0.34

B. Variance Decompositions of $D\ln(q)$ in the Benchmark Model and Three Alternative Models

Shock:	Model	Argentina	Bolivia	France	Germany	Israel	Mexico
ϵ^g	0	0.1 4.8	3.0 16.4	2.3 7.6	12.4 27.1	7.6 6.5	8.6 10.5
	1	0.1 4.4	2.9 11.9	2.3 7.5	12.4 27.1	6.4 5.7	6.2 10.4
	2	0.1 4.4	2.9 11.9	2.3 7.5	12.4 27.1	6.4 5.7	6.2 10.4
	3	0.3 6.2	0.1 15.7	0.1 3.9	0.2 14.4	7.5 7.4	16.3 14.4
ϵ^m	0	8.0 31.5	38.1 22.6	85.9 71.1	64.4 47.6	0.0 18.0	18.9 22.8
	1	16.9 16.1	0.1 14.8	31.1 28.9	0.6 6.0	0.4 14.4	6.3 19.8
	2	25.2 20.1	0.4 1.7	21.1 24.4	2.4 6.0	10.1 18.6	0.0 3.9
	3	7.9 30.1	41.0 23.3	88.2 74.9	76.6 60.3	0.0 17.0	11.3 18.9
ϵ^y	0	82.4 56.9	58.8 55.9	11.7 12.9	15.9 18.0	72.0 56.7	69.1 37.0
	1	73.6 72.8	95.9 66.1	66.5 55.6	79.7 59.5	71.0 59.3	83.7 42.4
	2	73.6 72.8	95.9 66.1	66.5 55.6	79.7 59.5	71.0 59.3	83.7 42.4
	3	82.4 56.9	58.8 55.9	11.7 12.9	15.9 18.0	72.0 56.7	69.1 37.0
ϵ^s	0	9.5 6.8	0.1 5.1	0.1 8.4	7.3 7.4	20.4 18.9	3.3 29.7
	1	9.4 6.7	1.2 7.2	0.0 8.1	7.3 7.4	22.2 20.6	3.8 27.4
	2	1.1 2.7	0.8 20.3	10.1 12.5	5.5 7.4	12.5 16.4	10.0 43.2
	3	9.5 6.8	0.1 5.1	0.1 8.4	7.3 7.4	20.4 18.9	3.3 29.7

Note: In each cell of part A we present estimates of the elements of the $C(1)$ matrix; predicted signs of the coefficients are given in parentheses in the first column. In part B we present the percentage of the variance of $D\ln(q)$ accounted for by the associated shock at the 1-period and 20-period horizons. Model 0 is the "Benchmark" model in the text, while models 1, 2, and 3 are alternatives. The ordering of variables is $[D(g/y), D\ln(m/q), D\ln(q), D\ln(s)]$ for model 0, $[D(g/y), D\ln(q), D\ln(m), D\ln(s)]$ for model 1, $[D(g/y), D\ln(q), D\ln(s), D\ln(m)]$ for model 2, and $[D\ln(m/q), D(g/y), D\ln(q), D\ln(s)]$ for model 3.

increases by about 10%–15%, entirely at the expense of the money shock.

Finally, consider model 3, which differs from the benchmark model only in that money is placed ahead of government size. This has the somewhat counterintuitive implication that real money is unaffected by the fiscal shock in the long run, but may be useful to consider if one argues that the fiscal and monetary shocks are inseparable. Table 6 shows that the results are strikingly similar to those of the benchmark model. Qualitatively, the results are unchanged, while the only quantitative difference is that fiscal shocks become a bit less important in the case of Germany.

F. Summary

We conclude from the short-run responses that in the 1980s episodes, the single most influential shock is the output shock. This suggests that movements in relative prices are primarily due to forces other than fiscal, monetary, or exchange rate shocks. These forces include technological or human capital advances, oil price shocks, price controls, and other possible permanent aggregate demand disturbances such as preference shifts. In the 1920s episodes, however, the money shock plays an essential role in affecting relative price movements.[11] It is also noteworthy that for all countries with rapid or immediate stabilization (Bolivia, France, Germany, and Israel), the relative price ratio adjusts rather quickly to a money shock, exhibiting a hump-shaped dynamic pattern. Finally, the sensitivity analysis shows that our results are robust to plausible alternative specifications.

The results are consistent with the fact that Argentina used money as a nominal anchor during the attempted stabilization, while Israel and Mexico adopted a nominal exchange rate anchor (Bruno et al. (1987)). They are also consistent with Dornbusch, Sturzenegger, and Wolf (1990), who conclude that the "deficit" shock (a combined monetary and fiscal shock) plays a crucial role in Argentina's inflationary experience, whereas exchange rate depreciation significantly affects the variance of inflation in Israel and Mexico (but not in Bolivia). Finally, the channel emphasized by Garber to explain relative price changes during the German hyperinflation generalizes to the hyperinflation episodes we analyze, but not to those of high or moderately high inflation.

[11] This conclusion is perhaps altered by the German VDCs from the alternative models 1 and 2.

IV. Conclusion

Casual analysis of price data from Argentina, Bolivia, France, Israel, and Mexico establishes as a stylized fact a positive relationship between movements in the relative price ratio and the inflation rate. Our estimated structural vector autoregression model allows us to study the dynamic effects of fiscal, monetary, output, and exchange rate shocks on the relative price ratio. The Samuelson-Stolper theorem implies there ought to be a connection between the observed movements in relative prices and output. Changes in relative prices will affect factor prices, and subsequently the composition of output and the structure of industry. Disentangling the important sources of relative price movements can, therefore, help us to understand why stabilization programs often have real effects.

We relate the sources of fluctuations in relative prices to several macroeconomic characteristics that may be common to each country/episode. Examining Argentina, Bolivia, France, Germany, Israel, and Mexico, we find that, although both monetary and output shocks are essential in the 1920s episodes, output shocks are the most influential in explaining relative price movements in the 1980s episodes. The latter implies that if changes in relative prices are an important source of real effects from the 1980s stabilization programs, they are due mostly to factors other than fiscal, monetary, or exchange rate shocks. Such factors include shocks to technology, human capital, or oil prices, price controls, and any permanent aggregate demand disturbances such as preference shifts.

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