American Economic Review 2008, 98:1, 439–457
http://www.aeaweb.org/articles.php?doi=10.1257/aer.98.1.439

Growth Dynamics: The Myth of Economic Recovery

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Although researchers have documented that many financial crises are associated with severe recessions (Graciela Kaminsky and Carmen Reinhart 1999), very little attention has been paid to whether countries recover from such large negative shocks in the sense that output losses are reversed. A few recent papers show persistent output loss from financial crises in a small set of countries. For instance, Cerra and Saxena (2005a) demonstrate that six Asian countries suffered permanent output loss from the Asian crisis, and Cerra and Saxena (2005b) show that only a tiny fraction of the output loss from Sweden's banking crisis in the early 1990s was recuperated. The graphs in Figure 1 illustrate persistent output loss for selected countries following the 1997–1998 Asian financial crisis and the debt crisis of the early 1980s.

In addition to financial crises, many countries experience large negative political shocks, which could include violent conflicts such as civil wars, as well as a deterioration in the country's governance. Such political shocks have the potential for significant disruption to economic activity, as illustrated for a few episodes of civil war (Figure 2).

This paper systematically documents the behavior of output following financial and political crises in a large set of 190 countries. While the graphs in Figures 1 and 2 are suggestive, our aim is to formally analyze the impact of financial and political shocks on output in a broad set of countries, particularly whether output losses are recovered. Financial shocks comprise currency, banking, and twin financial crises. For political shocks, we examine civil wars, a deterioration in the quality of political governance, and twin political crises comprising both shocks. We choose civil wars rather than interstate conflicts to ensure that the war occurs on the country's own soil. The military theater for some interstate conflicts may not directly encompass all parties to the conflict. In addition, the increase in wartime spending for an international conflict may boost economic activity in some countries. We also examine the economic impact of a deterioration in a country's political governance or institutional quality. Daron Acemoglu, Simon Johnson, and James Robinson (2001) and Acemoglu et al. (2003) use constraints on the power of the political executive as a measure of institutional quality, and find that it is linked to growth and volatility. Thus, we use this measure to study the shock to political governance.

Potential endogeneity of the financial or political crisis is an important issue in estimating the output impact of the crisis. That is, the crisis itself may be a function of a slowdown of economic growth or changes in expectations of future growth. We attempt to address this issue using a few methods that are far from definite, but nonetheless uncover some interesting facts. In particular, we find that the forecasts of growth from an autoregressive model and from consensus surveys are optimistic relative to actual growth occurring during and after a crisis.

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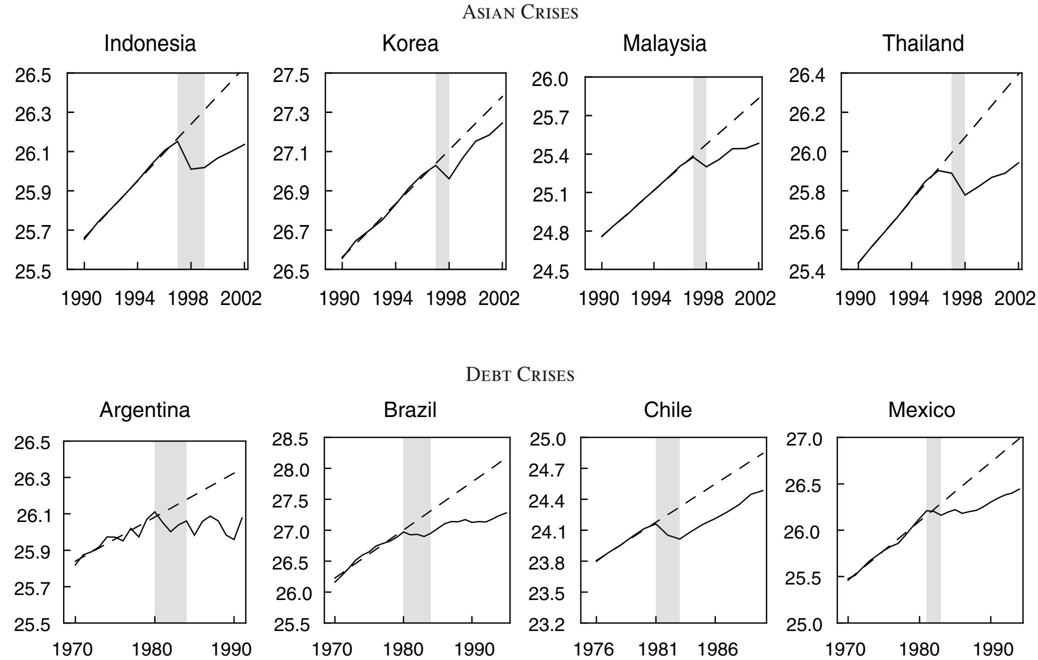

Figure 1

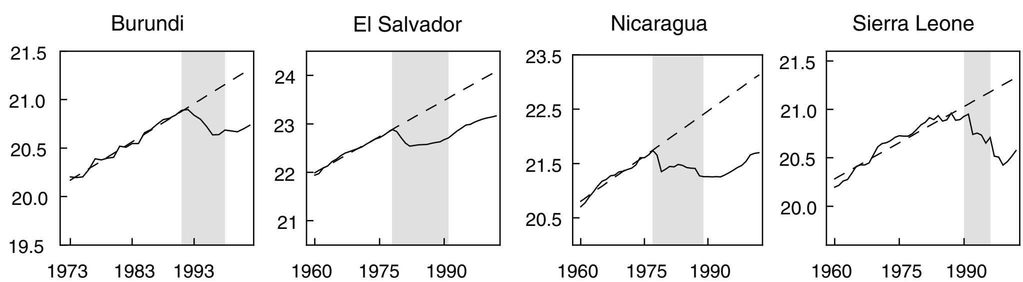

Figure 2. Protracted Civil Wars

I. Methodology

We are interested in examining the impact of financial and political crises on output. We follow the methodology used by Christina Romer and David Romer (1989) to identify the impact of monetary policy shocks on output. We construct qualitative indicators of financial and political crises and estimate impulse response functions to the shock. Given that our data cover a large set of countries, we estimate the models using panel data analysis, and we provide group averages of the impulse responses of output to each type of shock. We are also able to partition the country samples to examine any differential impact of a shock on countries according to their income level or region.

We formally test the statistical relationship between growth and the shock by estimating impulse response functions for each different type of shock. In particular, we estimate a univariate autoregressive model in growth rates, which accounts for the nonstationarity of output (Charles Nelson and Charles Plosser 1982) and for serial correlation in growth rates.[1] We control for country fixed effects, which F-tests indicate are present.[2] We estimate an AR(4), as we find insignificant coefficients beyond the fourth lag. We estimate the model on all of the available data from 190 countries over the period 1960 through 2001. We then extend the estimation equation to include the current and lagged impacts of the shock. Thus, we estimate the following model:

$$(1) \qquad g_{it} = a_i + \sum_{j=1}^{4} \beta_j g_{i,t-j} + \sum_{s=0}^{4} \delta_s D_{i,t-s} + \varepsilon_{it},$$

where g is the percentage change in real GDP and D is a dummy variable indicating a financial or political crisis. The impulse response functions to each crisis type are shown with a one-standard-error band drawn from a thousand Monte Carlo simulations.

II. Data

We use GDP growth rates from the World Bank's World Development Indicators (WDI). This dataset contains the largest sample of countries. Our data consist of unbalanced panels of annual observations spanning 190 countries from 1960 to 2001. We also disaggregate the results based on World Bank classifications. The countries are split into seven regional groups—Africa, Asia, industrial countries, Latin America, Middle East, transition countries, and Western Hemisphere islands—and four income groups—low income (per capita real GDP less than or equal to 735 dollars), lower middle income (per capita real GDP between 736 and 2,935 dollars), upper middle income (per capita real GDP between 2,936 and 9,075 dollars), and high income (per capita real GDP over 9,075 dollars). For robustness, we also use GDP growth rates computed using Penn World Tables data (Alan Heston, Robert Summers, and Bettina Aten 2006), converting per capita growth to aggregate GDP growth rates using population growth rates.

We form a panel dataset for currency crises by constructing an exchange market pressure index (EMPI) for each country. The EMPI is defined as the percentage depreciation in the exchange rate plus the percentage loss in foreign exchange reserves. This formulation makes indices comparable across countries.[3] A dummy variable for a crisis is formed for a specific year and country if the EMPI is in the upper quartile of all observations across the panel.

We obtain banking crisis dates on a large set of countries from Gerard Caprio and Daniela Klingebiel (2003). We confine the analysis to systemic banking crises. Moreover, since the end of a banking crisis is often highly uncertain, we restrict our analysis to the initial shock stemming from the first year of a banking crisis.

To check the robustness of our results, we also use banking and currency crisis dates from Kaminsky and Reinhart's (1999) influential study on twin crises. However, the drawback of this

[1] Panel unit roots tests using the data in this study strongly suggest the presence of unit roots in output. Results are available upon request.

[2] The country fixed effects are correlated with the lagged dependent variables in the autoregressive equation. However, Stephen Nickell (1981) shows that the order of bias is $1/T$, which is small for this dataset. Indeed, Ruth Judson and Ann Owen (1999) calculate that the bias of the least squares dummy variable (LSDV) estimator is approximately 2–3 percent on the lagged dependent variable and less than 1 percent on other regressors for a panel of size N=100, T=30, and low persistence.

[3] The crisis literature often normalizes reserves and exchange rate movements by their within-country standard deviations, but then the magnitudes of the EMPI are comparable only within countries. We dropped interest rates due to the scarcity of data.

source is that there are only 23 countries included in the study, which prevents us from examining the regional and income group disaggregations.

The data for civil war are obtained from Meredith Sarkees (2000) Correlates of War Intra-State War Data, 1816–1997 (v3.0) (at www.correlatesofwar.org), which updates the work of David Singer and Melvin Small (1994). The dataset identifies the participants of intrastate wars. We form a dummy variable for internal conflict by assigning a value of unity for a country in the years of conflict, and zero otherwise.

The data on the quality of the government comes from the Polity International IV dataset. The constraint on the executive variable is constructed by the Polity IV project by coding the authority characteristics of states in the world. The variable measures the extent of regular institutional constraints on executive power. These constraints arise from accountability groups, such as legislatures and judiciaries that have equivalent or greater effective authority, or can impose constraints on executive behavior in most activities. The scale ranges from 1 (weak constraints on executive power) to 7 (strong constraints on executive power).

Data on consensus forecasts of economic growth in the current year and next year are obtained from the commercial database produced by Consensus Economics Inc. The data are gathered monthly by surveying the expectations of analysts typically from large banks and financial firms within the country. The data cover 35 high-income, emerging market, and transition countries. Coverage begins in late 1989 for high-income countries, with additional countries added from 1990 to 1995. The frequency is monthly for high-income and Asian emerging market countries and bimonthly for Latin American and transition countries.

III. Results

A. *Impulse Responses*

The impact of a currency crisis on output is negative and highly persistent (Figure 3). The loss in output averages about 4 percent for the entire panel of countries. The depth of the loss varies by income group, with output loss averaging only 1 percent for high-income countries, but close to 5 percent for all other income groups. All groups except countries in the Middle East region experience output loss, which persists even at a ten-year horizon. Indeed, no income group or regional group experiences a rebound in output by more than ½ of a percent relative to the point of deepest loss.

The output impact of a banking crisis is nearly twice as large (7½ percent loss) as a currency crisis and just as persistent (Figure 4). Output loss at a ten-year horizon exceeds 6 percent for all groups except the Latin America region and lower-middle-income countries. Whereas currency crises have a modest impact on high-income countries, banking crises lead to severe output loss in this group.

Output loss from twin financial crises is deeper than either of the individual crises (Figure 5). By three years after the crisis, output loss reaches and remains at 10 percent. The persistence of the loss is robust to all regional and income groups except for the Latin America subset.

In contrast to the extreme persistence of output loss following financial crises, output partially rebounds from a civil war (Figure 6). On average for the panel of countries, output declines by 6 percent initially. Half the loss is recuperated after four years, but 3 percentage points of cumulative loss remain even after a decade. These results likely reflect the combination of both permanent and temporary effects. Physical infrastructure is damaged in most war situations and constrains output, but infrastructure may be repaired within a short time after the conflict ceases. Also, parameter uncertainty is large, and the standard error bands encompass zero for several groups. This result reflects the wide range of different country experiences to post-conflict

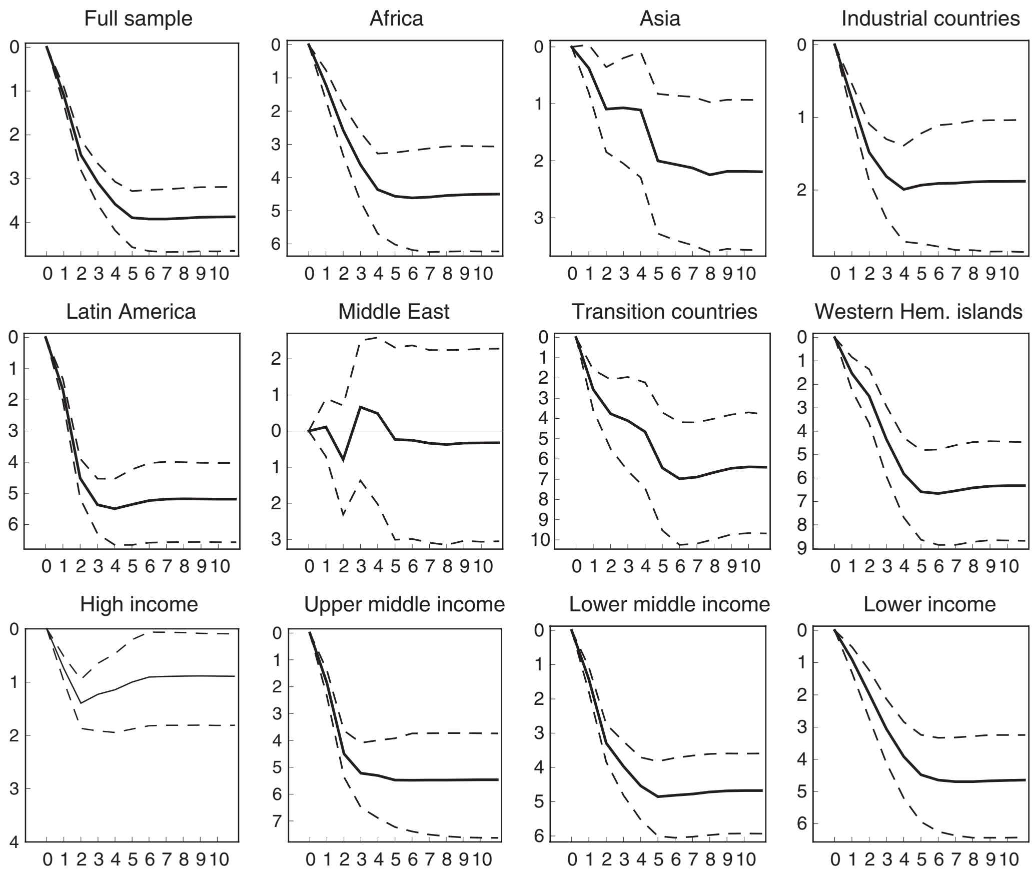

FIGURE 3. IMPULSE RESPONSES: CURRENCY CRISES

situations. The positive impact of civil war on output for industrial countries should be treated with caution, as the result is driven by limited episodes for this set of countries.[4]

The average output loss following a deterioration in constraints on executive power is as large and persistent as that for currency crises, but the impact varies markedly by the different regional and income groups (Figure 7). Weaker constraints on executive power (fewer institutional checks and balances) are associated with output gains for Asia and the Middle East, but with persistent losses for Africa, Latin America, transition countries, and islands in the Western Hemisphere. The difference by income group is monotonic. High-income countries have significant output *gains* when executive discretion strengthens. Upper-middle-income countries experience output losses of several percent, but the standard error bands are large. Lower-middle-income and low-income countries experience output losses of about 5 percent. In contrast to the partial rebound in output observed for civil wars, the output loss associated with this measure of a deterioration

[4] No civil war episodes were observed for Western Hemisphere islands or high-income countries. The three industrial countries not included in the high-income group are Malta, South Africa, and Turkey. Other civil war results should also be treated with caution, as some episodes of civil war reflect events confined to a small region of a country, and therefore may not have a systemic economic impact. For instance, the inclusion of smaller regional conflicts within the definition of civil war could explain the relatively small permanent output effects.

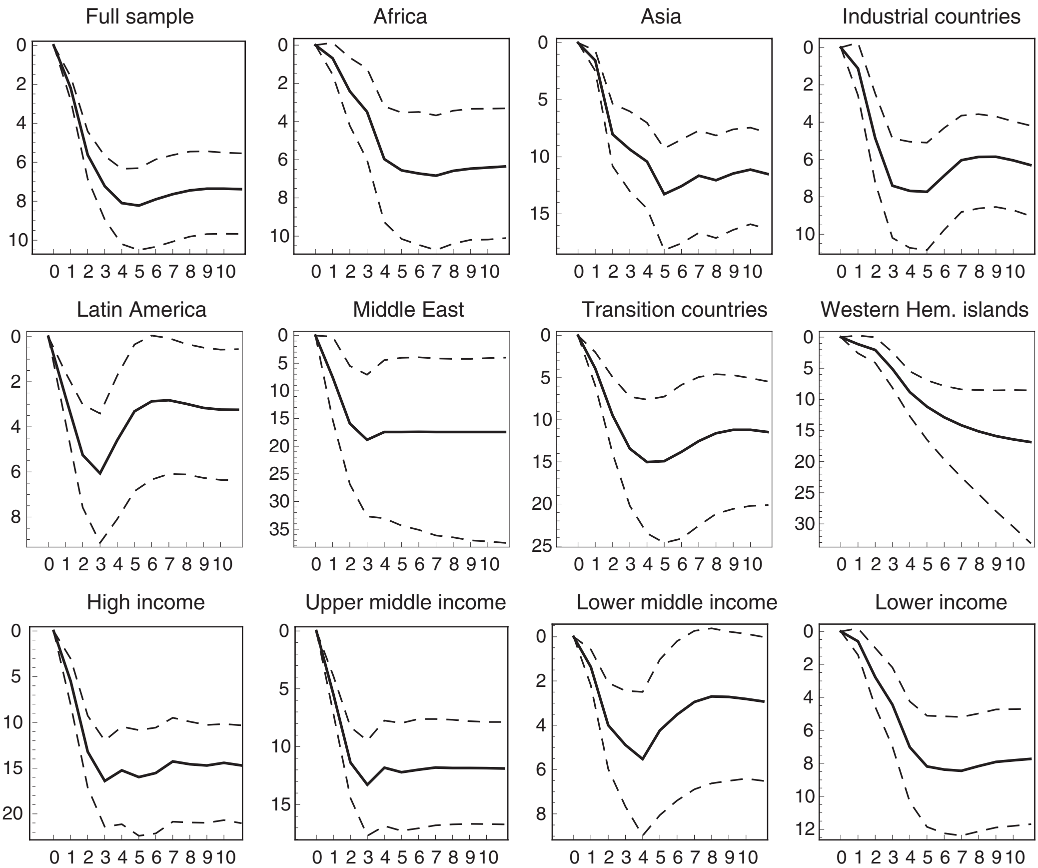

Figure 4. Impulse Responses: Banking Crises

in governance leads to sustained output loss, averaging 4 percent for the full set of countries. The different impact of changes in executive power on income groups may reflect large differences in initial starting levels. In high-income countries, power is widely distributed among institutions with strong constraining mechanisms. The average level is 6.1 on the scale from 1 to 7. The average level declines monotonically with income groups to 2.8 in low-income countries. The results suggest that an optimal sharing of power may lie between the bounds. At low levels of power sharing, executive power may be too discretionary and contribute to cronyism. Too much power sharing, on the other hand, may cripple decision making.

Twin political crises (civil wars combined with fewer controls on executive discretion) have the most severe overall impact on output of any large negative shock that we study (Figure 8). Output declines by about 16 percent on average for our broad set of countries. Moreover, the loss is persistent, with no discernable rebound. The output loss is particularly devastating for low-income countries, reaching 20 percent.

B. *Distribution of Shocks*

The results above show that the impact of a crisis varies among different country groups. For several types of shocks, output loss is more severe for lower-income countries than high-income

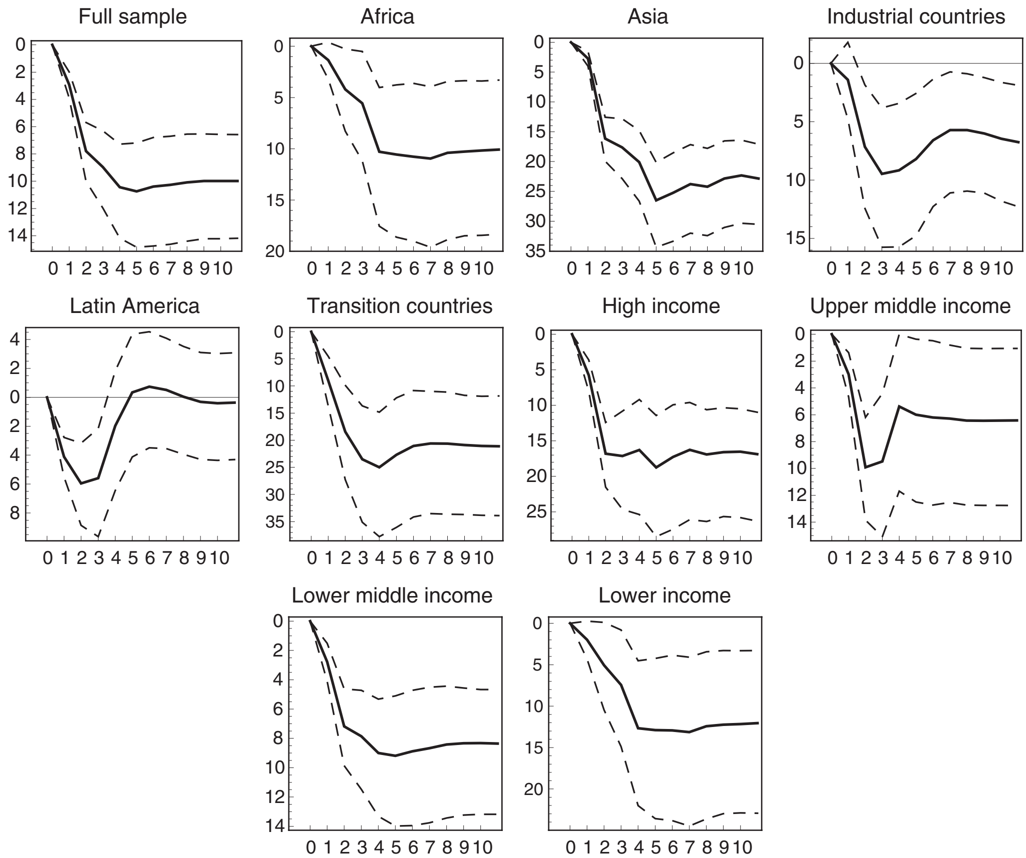

Figure 5. Impulse Responses: Twin Financial Crises (Currency and Banking)

countries. In this section, we calculate the frequency of each type of shock for each country subsample. The analysis consists of all country-year observations in which data on the growth rate and data on the shock indicator are available.

The frequency of shocks varies considerably across different country subgroups (Table 1). In particular, the frequency increases sharply and nearly monotonically as the income level of the country group falls. Financial crises occur almost twice as often in low-income countries as in high-income countries.[5] Political crises are even more unequally distributed. Civil wars occur in 18 percent of all years in low-income countries, but are not observed in high-income countries. A deterioration of constraints on executive power is not observed in high-income countries in the most recent decade of available data, but occurs in nearly a quarter of the years for low-income countries.

These computations indicate that the higher frequency of crises in lower-income countries relative to high-income countries, especially in the recent decade, compounds the generally

[5] The higher average frequency of currency crises relative to banking crises is partly a matter of data construction. Dummies for currency crises are formed from the highest quartile of the EMPI, but banking crisis dummies reflect only the first year of banking crises. This construction, however, does not affect the relative frequency of crises across country groups.

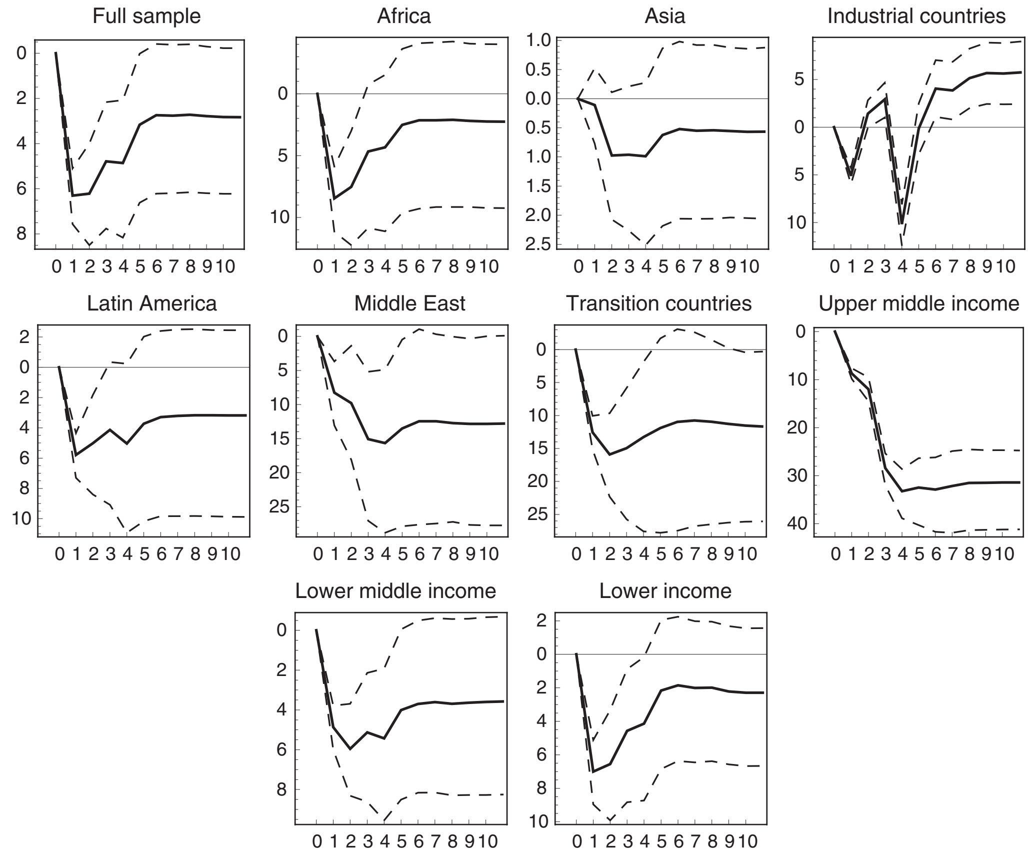

FIGURE 6. IMPULSE RESPONSES: CIVIL WARS

larger output loss associated with the crises.[6] Indeed, multiplying full sample estimates for the long-term output loss of a crisis (the loss at the ten-year horizon in the impulse response) by the probability of a crisis in each year, we find that a ten-year accumulation of financial and political shocks could reduce the long-term level of output in low-income countries by as much as 25 percentage points more than in high-income countries. Currency crises produce the greatest differential effects between high-income countries and the other three groups, because of significant differences in both frequency and magnitude. In addition, the considerably higher frequency of governance shocks (less constraints on executive discretion) in low- and low-middle-income countries in the recent period would intensify the differences in output loss between the income groups.

[6] The higher frequency and deeper relative output loss for low-income and emerging market countries, combined with the persistence of output loss, implies that such crises contribute to greater volatility of permanent shocks for these countries than for high-income countries. Similarly, Mark Aguiar and Gita Gopinath (2007) find that business cycle fluctuations in emerging markets are driven primarily by permanent shocks. Using parameter estimates from Mexico and Canada as benchmark emerging market and developed countries, respectively, they exploit consumption and net export shares to GDP to identify the underlying productivity processes.

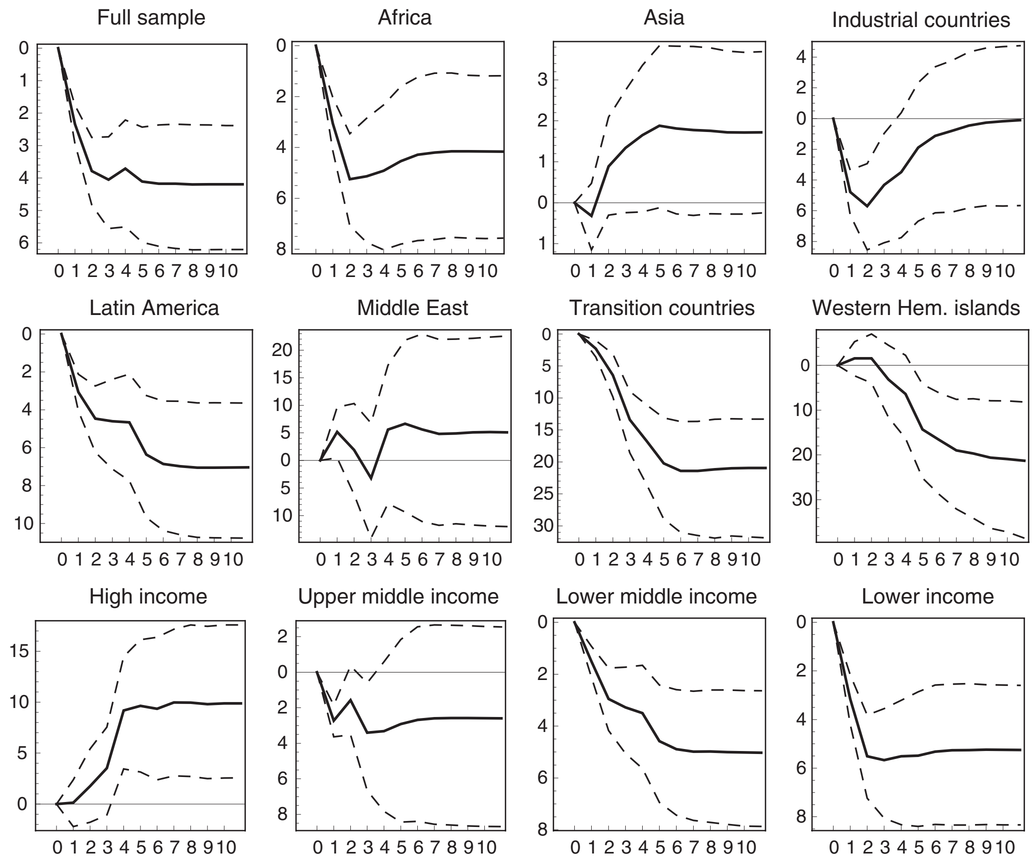

FIGURE 7. IMPULSE RESPONSES: STRONGER EXECUTIVE POWER

C. *Robustness Checks*

We also check the robustness of our results using alternative sources of crisis dates and growth rates, and controls for some common shocks.

TABLE 1—PROBABILITY OF SHOCKS

	All available years				1992–2001			
	Financial crises		Political crises		Financial crises		Political crises	
	Currency	Bank	War	Constraints	Currency	Bank	War	Constraints
Africa	37	4	12	6	34	4	16	23
Asia	23	3	16	7	19	4	10	9
Industrial countries	22	2	1	1	28	1	4	1
Latin America	31	5	10	5	27	5	10	8
Middle East	26	4	10	4	16	2	5	12
Transition countries	27	6	6	4	20	6	11	13
Western Hem. islands	21	2	0	3	11	5	0	3
High income	20	2	0	2	20	1	0	0
Upper middle income	23	5	2	3	16	4	0	8
Low middle income	30	5	10	6	26	6	13	12
Low income	36	4	18	6	32	4	19	23

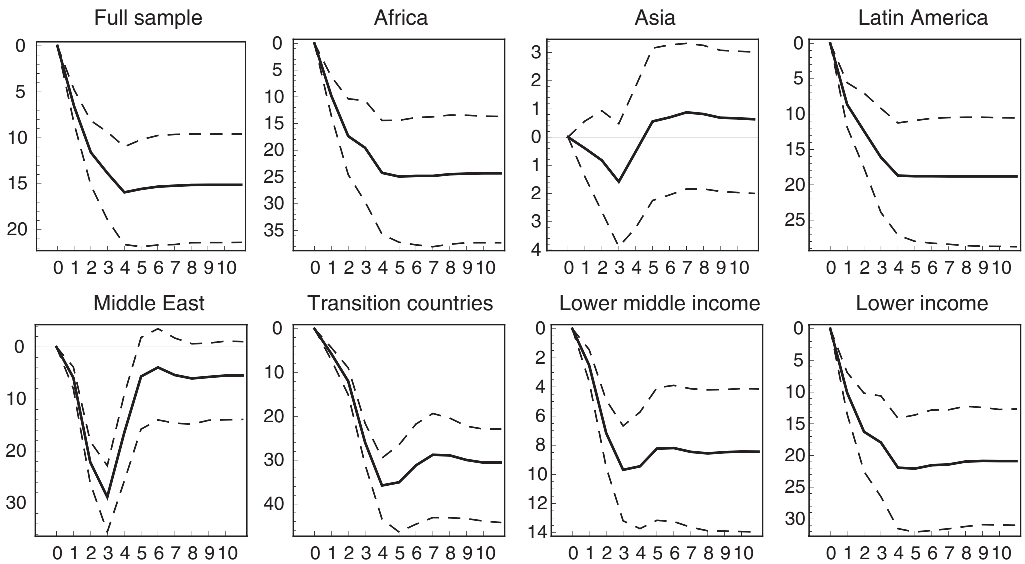

FIGURE 8. IMPULSE RESPONSES: TWIN POLITICAL CRISES (WAR AND STRONGER EXECUTIVE POWER)

The results are robust to the source of data on crisis dates and growth rates. As a check of the robustness of our results on the persistent output loss from financial crises, we examine the impulse response functions using the crisis dates from Kaminsky and Reinhart (1999). As shown in Figure 9, we continue to find deep and persistent output loss from currency, banking, and twin financial crises using this alternative set of crisis dates. The size of the loss is, in fact, larger by 1–4 percentage points. We also substitute growth rates from the latest release of the Penn World Tables, and find similar impulse response functions (Figure 10).

The results are also robust to controls for common shocks. In Figure 11, we add oil price changes into the regression. In Figure 12, we allow for arbitrary common shocks by including period effects. The impulse response functions are not affected much by either of these controls.

In addition to examining robustness, we note that our estimates of the extent of output loss may be conservative due to the use of fixed effects estimation. As mentioned, fixed effects estimation provides a downward bias to the coefficient estimate on a lagged dependent variable in a panel regression (Nickell 1981), although we argue it should be very small in this dataset given the fairly long time series. To the extent that true growth rates are more serially correlated than our estimates, the impact of a shock on output would be magnified compared to our estimates.

IV. Exogeneity

Our estimating equation from Section I assumes that we can treat the occurrence of a crisis as a contemporaneously exogenous event with respect to output growth. However, the other polar case—in which output growth is contemporaneously exogenous with respect to the crisis, and the crisis has only a lagged effect on output—is also plausible. Second, when we estimate equation (1) as a single equation, we are implicitly ignoring any feedback from previous output growth to the current probability of a crisis. In this section, we discuss the impact of alternative assumptions and provide some evidence to address these issues of exogeneity.

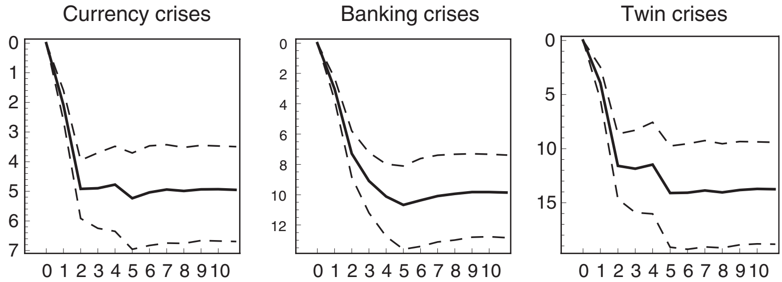

FIGURE 9. IMPULSE RESPONSES: FINANCIAL CRISES USING KAMINSKY AND REINHART (1999) DATES

Formally, suppose that we can model the relationship between growth and crisis as a bivariate system of equations, conceptually analogous to a bivariate vector autoregression, although not linear. That is, we augment the growth equation with another equation specifying that the probability of a crisis depends on contemporaneous and lagged growth and lags of the crisis dummy:

$$g_{it} = \alpha_i + \sum_{j=1}^{4} \beta_j g_{i,t-j} + \sum_{s=0}^{4} \delta_s D_{i,t-s} + \varepsilon_{it}\,; \tag{1}$$

$$\Pr(D_{i,t} = 1) = F\left(\mu + \sum_{j=0}^{4} \gamma_j g_{i,t-j} + \sum_{s=1}^{4} \phi_s D_{i,t-s} + \nu_{it}\right). \tag{2}$$

In Section I, we assume zeros for all coefficients in equation (2). This assumption implies that the crisis is strictly exogenous for growth, and serially uncorrelated.

Using this framework, we can modify our assumption in two key ways. First, if $\delta_0 = 0$, then output growth is contemporaneously exogenous, and the crisis has only a lagged effect on growth. The original assumption and this modified assumption are similar to the two alternative triangular

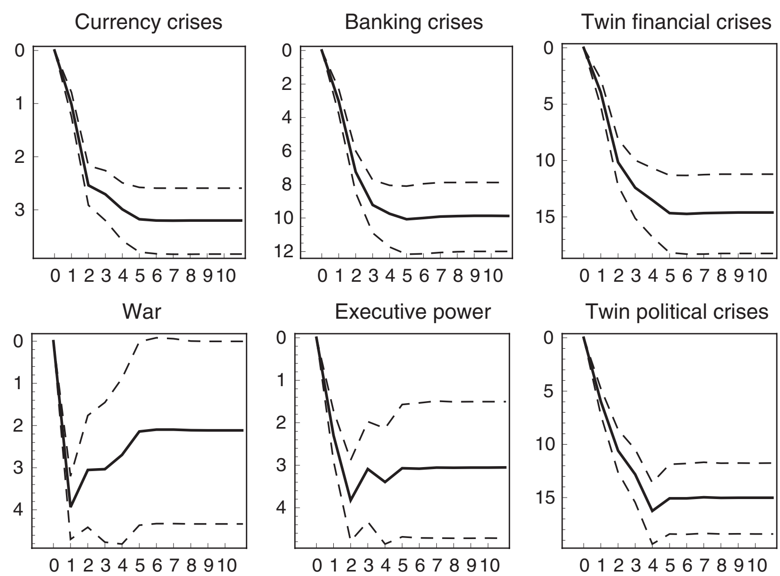

FIGURE 10. IMPULSE RESPONSES: PENN WORLD TABLES DATASET

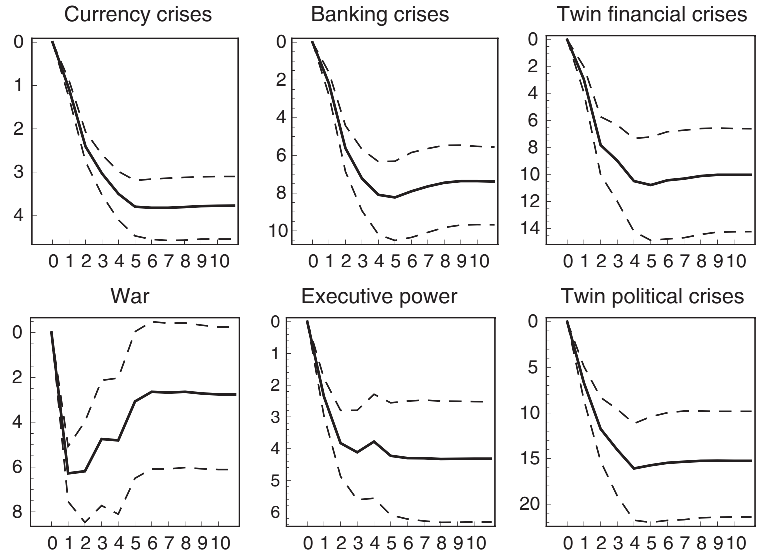

Figure 11. Impulse Responses: Controlling for Oil Price Changes

factorizations that can be imposed for a Cholesky decomposition of a bivariate VAR. The second important modification would be to allow nonzero coefficients in equation (2), so that output growth can affect the probability of a crisis and the crisis dummy can be serially correlated.

The data indicate that lower growth is associated with a higher probability of a crisis within the same year. By assuming that the crisis is exogenous, we attribute the low growth to the impact of the crisis. If, instead, output growth is exogenous ($\delta_0 = 0$), then the crisis may be a result, rather than a cause, of the low growth. In this case, the impact of the crisis would be due only to the lagged effects in equation (1). The output loss shown in Figures 3–12 may thus be exaggerated.

The next two sections provide some evidence to address this question of the contemporaneous relationship between growth and a crisis. In Section IVA, we generate dynamic forecasts from a univariate autoregressive model of growth. We compute errors between forecasts from the model and actual growth following a crisis under each polar case for contemporaneous exogeneity of output growth and crisis. In Section IVB, we analyze consensus forecasts of expected growth in an attempt to disentangle whether a crisis leads to output loss, or whether actual or expected output loss leads to the crisis. The results provide some evidence pointing to growth optimism at the time of a crisis.

In Section IVC, we explicitly impose the alternative assumption that output growth is exogenous by setting $\delta_0 = 0$ in the estimation equation. In this case, the crisis has only lagged effects on output growth, and thus we generate impulse response functions that correspond to this alternative assumption.

In Section IVD, we discuss the impact of allowing nonzero coefficients in equation (2). We provide some evidence from probit models showing that low growth would increase the probability of a future crisis. Moreover, crises are serially correlated. The implication of this evidence is that the results for output loss shown in Section III may be underestimated, since they ignore this feedback from lower output growth to the higher probability of a subsequent crisis.

In Section IVE, we consider some additional potential specification errors, such as the impact of expectations and the possibility of third variables driving both growth and crises.

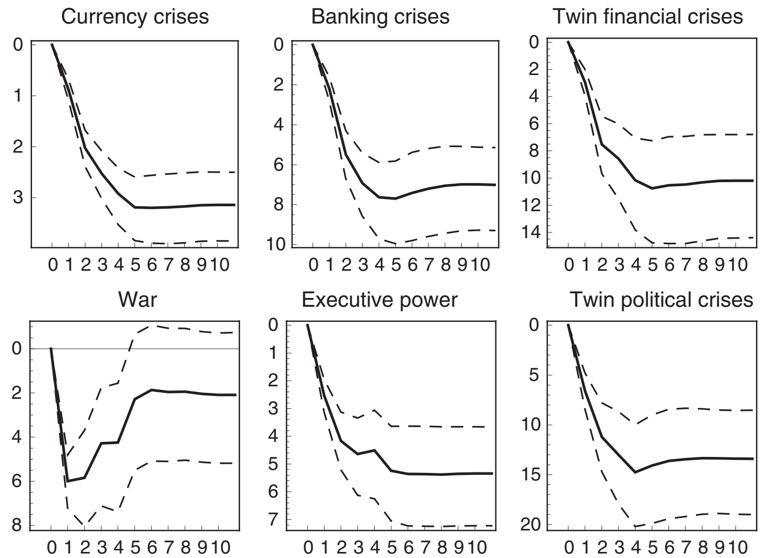

FIGURE 12. IMPULSE RESPONSES: CONTROLLING FOR COMMON PERIOD SHOCKS

A. *Forecast Errors*

We compare the actual level of output following a financial or political crisis with the level of output predicted from a univariate AR model that controls for normal business cycle dynamics. The forecast error provides a measure of the impact of the crisis. We show how the results differ under alternative extreme assumptions that all contemporaneous correlation between output and crisis can be attributed to (a) crisis innovations, or (b) output innovations.

For the panel of countries, we estimate a univariate autoregressive model in growth rates to account for the business cycle and any ex ante slowdown in growth:

$$g_{it} = \alpha_i + \sum_{j=1}^{4} \beta_j g_{i,t-j} + \varepsilon_{it}. \tag{3}$$

For each crisis date, t, in each country, we then compare current and subsequent actual growth rates to those of dynamic forecasts constructed using coefficient estimates from the AR(4) model. However, contemporaneous correlation between current growth and the crisis must be distributed between the two variables, as current growth may be unexpectedly low due to the crisis, or the crisis may occur because of the negative innovation in growth. To account for these possibilities, we construct two sets of forecast errors that correspond to each extreme assumption.

Assuming that low growth in time t occurs because of the crisis innovation in time t, we form 1-, 2-, 3-, and 4-period-ahead dynamic growth forecasts using growth data only through time $t-1$. The forecast errors are given by

$$\begin{aligned} ferr_{+1} &= g_{it} - g_{it}^{f} = g_{it} - \left(\hat{\alpha}_i + \sum_{j=1}^{4} \hat{\beta}_j g_{i,t-j}\right), \\ &\vdots \\ ferr_{+4} &= g_{i,t+3} - g_{i,t+3}^{f} = g_{i,t+3} - \left(\hat{\alpha}_i + \sum_{j=1}^{3} \hat{\beta}_j g_{i,t+3-j}^{f} + \hat{\beta}_4 g_{i,t-1}\right). \end{aligned} \tag{4}$$

Alternatively, if we fully attribute any slowdown in growth in the year of a crisis (contemporaneous correlation between growth and crisis) to growth innovations, we can construct forecast errors from our AR(4) model using growth information through time *t*. Under this assumption, the growth slowdown is responsible for the crisis in time *t* and the forecast errors will pick up only lagged effects from the crisis to future growth. Thus, the forecast errors are constructed as

$$(5)\qquad ferr'_{+1} = g_{i,t+1} - g^{f}_{i,t+1} = g_{i,t+1} - \left(\hat{\alpha}_i + \sum_{j=1}^{4} \hat{\beta}_j g_{i,t-j+1}\right),$$

$$\vdots$$

$$ferr'_{+4} = g_{i,t+4} - g^{f}_{i,t+4} = g_{i,t+4} - \left(\hat{\alpha}_i + \sum_{j=1}^{3} \hat{\beta}_j g^{f}_{i,t+4-j} + \hat{\beta}_4 g_{it}\right).$$

We compute such sets of forecast errors for each crisis in our panel data, and compute the average forecast error across the sample at each horizon. Figure 13 presents forecast errors of the level of output by accumulating the 1-, 2-, 3-, and 4-year-ahead forecast errors of output growth as described in equations (4) and (5). Forecast errors shown by solid lines assume that any correlation between growth and a financial or political crisis is attributed to the crisis in the year it occurs, whereas those shown by dashed lines attribute the correlation entirely to the innovation in the growth rate.

The results show that output loss occurs irrespective of which polar assumption is used to attribute contemporaneous correlation between output and crisis innovations. The magnitude of output loss is smaller if crises have only lagged effects on growth, corresponding to $\delta_0 = 0$ in equation (1). The attribution of contemporaneous correlation to growth versus crisis innovations affects the magnitude of output loss more for political crises than for financial crises. However, on average across the panel of countries, actual growth rates fall short of those that are projected to account for normal business cycle fluctuations for all four types of shocks, regardless of alternative assumptions on contemporaneous exogeneity.

B. *Consensus Growth Forecasts*

We consider the possibility that a crisis may occur not only due to an ex ante decline in output growth, but also because economic agents may expect a future slowdown. For instance, suppose that economic agents revise downward their forecast of growth and, as a result, they take actions that induce a financial crisis, start a civil war, or weaken the constraints on executive power. In this situation, the contemporaneous correlation between the crisis and growth should be attributed to growth, and the crisis would have only a lagged impact on growth.

To account for changes in growth forecasts, we collect consensus forecasts of economic growth in the crisis year and subsequent year for a set of industrial and emerging market countries. For each type of financial and political crisis, we regress the difference between the midyear consensus forecast of growth in the current crisis year and the actual growth outturn, as well as the difference between the midyear forecast of growth in the following year and its actual outturn, on the crisis dummy indicator.[7] We also compare the timing of revisions to expected growth using a difference of differences specification. We regress the changes in consensus forecasts $\left[(F_{it}g_{it} - F_{i,t-1}g_{it}) - (F_{i,t+1}g_{i,t+1} - F_{it}g_{i,t+1})\right]$ on the crisis dummy, where F_{it} denotes the consensus forecast of country *i* in the middle of year *t*. If the revision to growth forecasts occurs before the crisis, the first term would be negative and the second term zero. In contrast, the first term would be zero and the second part of the expression would be positive if growth forecasts are revised downward after the crisis (that is, if $F_{i,t+1}\, g_{i,t+1} < F_{it}\, g_{i,t+1}$).

[7] The regressions in this section include fixed effects.

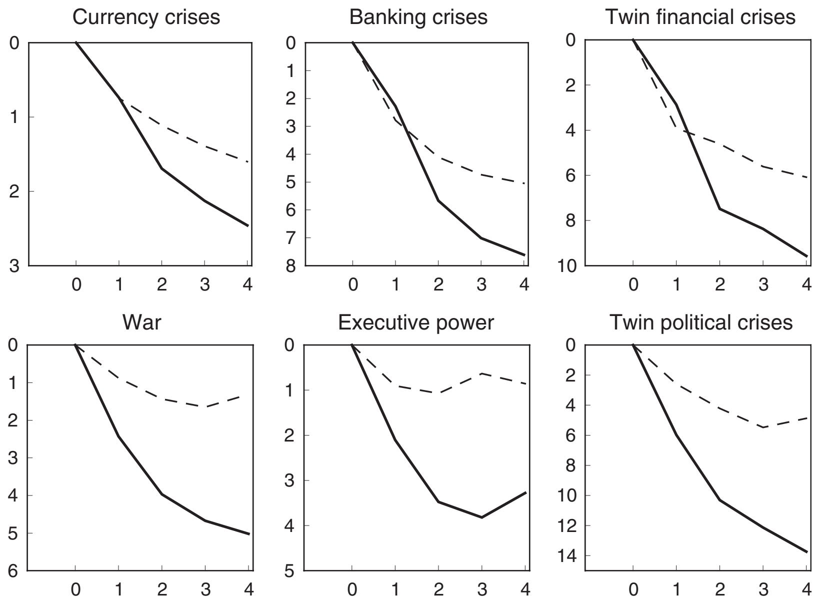

Figure 13. Forecast Errors

For financial crises, we find robust evidence of growth optimism. Table 2 shows that the mid-year consensus forecast of growth in the year of a crisis is 0.8 percentage points higher than the actual outturn for currency crises, and 2.4 percentage points higher for banking crises. The expectational error is even larger for growth in the subsequent year: 1.9 percentage points too optimistic for currency crises and 4.7 percentage points too optimistic for banking crises. Moreover, growth revisions lag, rather than lead, a financial crisis, especially for banking crises. Growth is revised downward by 6 percentage points more after the banking crisis starts, compared with any ex ante revision.

The evidence on consensus forecasts for political crises is insignificant. The weak results may reflect that the sample of countries and time period available for consensus forecast data is quite restricted compared with the broad sample of countries available for the impulse response analysis. The consensus forecast sample includes only one low-income country. Thus, the results tend to be biased toward higher-income countries, which do not suffer as large output loss from political crises, especially from the decline in constraints on executive power.

C. *Contemporaneously Exogenous Output Growth*

In this section, we consider the case that a crisis has only a lagged effect on output growth. That is, we impose a zero restriction on the coefficient of the contemporaneous value of the dummy variable in equation (1), $\delta_0 = 0$. This assumption implies that output is contemporaneously exogenous with respect to a crisis. We examine the output loss associated with this alternative specification. In particular, we estimate a revised equation in which crises have an impact on GDP growth only through lags, and we control for the dynamics of growth rates:

$$g_{it} = \alpha_i + \sum_{j=1}^{4} \beta_j g_{i,t-j} + \sum_{s=1}^{4} \delta_s D_{i,t-s} + \varepsilon_{it}. \tag{6}$$

Table 2—Consensus Forecast Errors for Growth

	Financial crisis		Political crisis	
	Currency	Bank	War	Constraints
Current year expectation error	0.83***	2.43***	–0.10	0.97
Standard errors	0.31	0.52	1.03	1.13
Number of countries	34	33	34	34
Number of observations	408	352	229	386
Next year expectation error	1.88***	4.74***	–0.59	–0.49
Standard errors	0.51	0.84	0.87	1.67
Number of countries	34	30	34	33
Number of observations	376	328	205	362
Revision of expectations	0.38	5.84***	0.70	–0.23
Standard errors	0.56	1.07	1.28	1.76
Number of countries	34	34	34	34
Number of observations	378	326	199	358

Notes: Sample is 1990–2004, as available. July is used for consensus forecasts, except for Latin American countries, where June is used. Regressions include fixed effects. Asterisks *** denote significance at the 1 percent level.

Figure 14 shows the impulse responses from this regression. The change in the assumption of contemporaneous exogeneity has a smaller impact on the results for financial crises than political crises. The finding of persistent output loss remains robust for financial crises. As expected, the magnitude of loss is dampened when the contemporaneous decline in output is attributed to the output innovation. But even under this assumption, the lagged effects of currency, banking, and twin financial crises still result in 2½ percent, 4 percent, and 5 percent of output loss, respectively, by the end of ten years. For wars and the weakening of executive constraints, output falls initially, but at the end of ten years it is only 1 percentage point lower than its initial level. Output loss from twin political crises remains at 4 percent at the end of ten years, but the uncertainty bands are large.

D. *Feedback to the Probability of Crisis*

The results presented in Section III ignore the possibility that growth affects the probability of a future crisis or that crises are serially correlated. We relax this assumption by estimating a probit model for equation (2), using each type of crisis indicator as a dependent variable. We impose only the restriction that $\gamma_0 = 0$, so that the crisis is contemporaneously exogenous. The results in Table 3 show that even when controlling for lags of the crisis itself,[8] the first lag of growth has a significant inverse relationship with the probability of each type of crisis. That is, lower (lagged) growth leads to a higher probability of crisis. Moreover, currency crises and civil wars are positively serially correlated. Under the assumption that $\gamma_0 = 0$, the omission of such feedback effects—higher probability of crisis resulting from both lower lagged growth and positive serial correlation—from the results shown in Section III imply that the impulse responses underestimate the overall impact of a crisis on output.

E. *Expectations and Omitted Variables*

To summarize, we provide some evidence of growth optimism at the time of a crisis, suggesting that the crisis is contemporaneously exogenous with respect to output growth. If this

[8] Lags of banking crisis are not included, since by construction the variable measures only the first year in a string of possible banking crisis years.

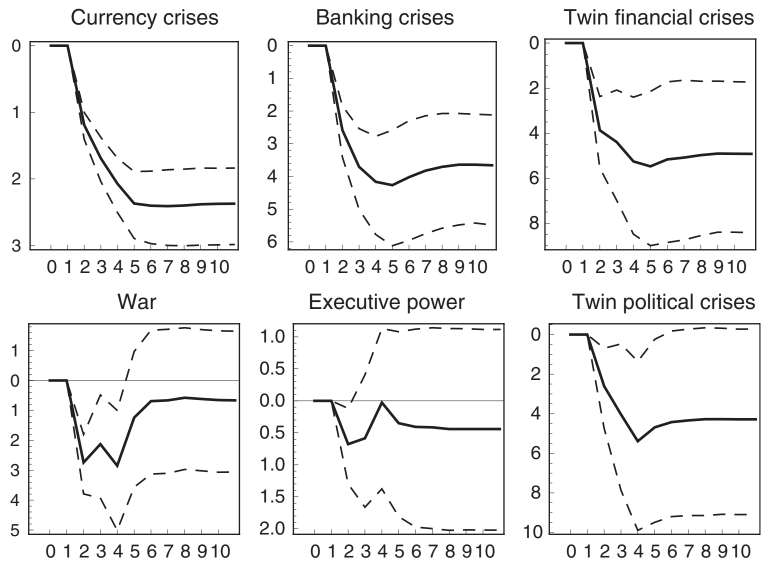

Figure 14. Impulse Responses: Lagged Response of Output to Crises

finding is invalid, and instead growth is contemporaneously exogenous, we then show that even the lagged effects of the crisis would reduce output. Lower output and serial correlation of crises would act to further reduce output by feedback effects that increase the probability of a future crisis.

Nevertheless, other potential specification errors need to be considered. Even the lagged impact of a crisis on growth could overstate the output loss from crisis innovations if we allow for expectational factors and omitted variables. For instance, suppose that the crisis affects output growth with a lag, as in Section IVC, where $\delta_0 = 0$. But suppose, also, that the probability of a crisis depends on expectations of lower future growth, and growth depends on some omitted variable, Z, that relevant economic agents can observe. The system of equations becomes

$$(7) \qquad g_{it} = \alpha_i + \sum_{j=1}^{4} \beta_j g_{i,t-j} + \sum_{s=1}^{4} \delta_s D_{i,t-s} + Z_{i,t-1} + \varepsilon_{it}$$

$$(8) \qquad \Pr(D_{i,t} = 1) = F\left(\mu + \sum_{j=0}^{4} \gamma_j g_{i,t-j} + \sum_{s=1}^{4} \phi_s D_{i,t-s} + E_{i,t} g_{i,t+1} + \nu_{it}\right).$$

Even though the crisis dummy is predetermined with respect to output growth in equation (7), it can still be correlated with the error term that contains the omitted variable. We make use of our consensus forecast data to test the specification of equation (8), although the consensus forecast data limit the countries and time periods, and we can show results only for financial crises. The results suggest that the expectation of future growth is not a significant determinant of a currency crisis, once we include actual growth in the regression equation (shown in the second column of results in Table 3). The expectation of future growth is significant at the 10 percent level for the probability of a banking crisis (column 4 of Table 3), but the sign is perversely positive, again suggesting excessive growth optimism.

Table 3—Probit Results for Financial and Political Crises

	Financial crisis				Political crisis	
Dep var/shock type	Currency		Banking		Civil war	Constraints
Constant	–0.783 ***	–0.906***	–1.767***	–1.815***	–2.253***	–1.761***
	0.032	0.220	0.045	0.318	0.055	0.041
Shock (–1)	0.326***	0.525***			2.758***	–0.140
	0.044	0.209			0.163	0.200
Shock (–2)	0.160***	0.047			0.073	–0.112
	0.045	0.223			0.213	0.197
Shock (–3)	0.169***	0.074			0.671***	0.292*
	0.045	0.212			0.176	0.153
Growth		–0.093		–0.169***		
		0.031		0.042		
Growth (–1)	–0.010***	–0.016	–0.028***	–0.030	–0.012**	–0.017***
	0.003	0.029	0.006	0.037	0.006	0.005
Next year growth expectations		0.054		0.192*		
		0.071		0.102		
Number of observations	4691	312	3043	286	5078	4671
Sample	1964–2002	1990–2001	1975–2002	1990–2001	1962–1997	1962–1997

Notes: Standard errors are shown below coefficient values.
*** Significant at, or below, 1 percent.
** Significant at, or below, 5 percent.
* Significant at, or below, 10 percent.

Finally, we cannot rule out the possibility of an omitted variable in both equations that causes a crisis and reduces growth:

$$g_{it} = \alpha_i + \sum_{j=1}^{4} \beta_j g_{i,t-j} + \sum_{s=1}^{4} \delta_s D_{i,t-s} + Z_{i,t-1} + \varepsilon_{it}; \tag{9}$$

$$\Pr(D_{i,t} = 1) = F\left(\mu + \sum_{j=0}^{4} \gamma_j g_{i,t-j} + \sum_{s=1}^{4} \phi_s D_{i,t-s} + \theta Z_{i,t} + \nu_{it}\right). \tag{10}$$

In this situation, omitting variable Z implies that the coefficient on the crisis dummy variable may be overestimated as it captures the correlation with the error term rather than the pure effect of the (lagged) crisis on growth. This possibility is quite plausible, given that crises and growth are likely to be related to, or driven by, other macroeconomic variables.

V. Concluding Comments

Using panel data for a large set of high-income, emerging market, developing, and transition countries, this paper documents that the large output loss associated with financial crises and some types of political crises is highly persistent. Impulse response functions show that less than 1 percentage point of the deepest output loss is regained by the end of ten years following a currency crisis, banking crisis, deterioration in political governance, twin financial crises, or twin political crises. Of the large negative shocks examined, a partial rebound in output is observed only for civil wars. Moreover, the magnitude of persistent output loss ranges from around 4 percent to 16 percent for the various shocks.

The paper provides some suggestive, although not definitive, evidence of causality. Financial crises are associated with growth optimism. Forecasts of economic growth, whether measured by projections from a univariate autoregressive model or by consensus forecasts of financial experts, tend to be higher than actual growth outturns. This evidence cannot, however, rule out the possibility of a third factor that precipitates a crisis and leads to a reversal of growth optimism.

The results pose a challenge to explain the observed behavior of output following the various negative shocks. Temporary output losses could be explained by allowing for variable capacity utilization or other elements of business cycle models, but the puzzle is to explain the permanent effects. It would be useful, therefore, to develop theoretical models with propagation mechanisms that are persistent, especially for low-income and emerging market economies.

REFERENCES

Acemoglu, Daron, Simon Johnson, and James A. Robinson. 2001. "The Colonial Origins of Comparative Development: An Empirical Investigation." *American Economic Review*, 91(5): 1369–1401.

Acemoglu, Daron, Simon Johnson, James Robinson, and Yunyong Thaicharoen. 2003. "Institutional Causes, Macroeconomic Symptoms: Volatility, Crises and Growth." *Journal of Monetary Economics*, 50(1): 49–123.

Aguiar, Mark, and Gita Gopinath. 2007. "Emerging Market Business Cycles: The Cycle Is the Trend." *Journal of Political Economy*, 115(1): 69–102.

Caprio, Gerard, and Daniela Klingebiel. 2003. "Episodes of Systemic and Borderline Financial Crises." http://go.worldbank.org/5DYGICS7B0.

Cerra, Valerie, and Sweta Chaman Saxena. 2005a. "Did Output Recover from the Asian Crisis?" *IMF Staff Papers*, 52(1): 1–23.

Cerra, Valerie, and Sweta Chaman Saxena. 2005b. "Eurosclerosis or Financial Collapse: Why Did Swedish Incomes Fall Behind?" International Monetary Fund Working Paper 29.

Heston, Alan, Robert Summers, and Bettina Aten. 2006. Penn World Table Version 6.2. Center for International Comparisons of Production, Income and Prices at the University of Pennsylvania.

Judson, Ruth A., and Ann L. Owen. 1999. "Estimating Dynamic Panel Data Models: A Guide for Macroeconomists." *Economics Letters*, 65(1): 9–15.

Kaminsky, Graciela L., and Carmen M. Reinhart. 1999. "The Twin Crises: The Causes of Banking and Balance-of-Payments Problems." *American Economic Review*, 89(3): 473–500.

Nelson, Charles R., and Charles I. Plosser. 1982. "Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications." *Journal of Monetary Economics*, 10(2): 139–62.

Nickell, Stephen J. 1981. "Biases in Dynamic Models with Fixed Effects." *Econometrica*, 49(6): 1417–26.

Polity IV Project. Integrated Network for Societal Conflict Research (INSCR) Program, Center for International Development and Conflict Management (CIDCM). University of Maryland. www.cidcm.umd.edu/inscr/polity.

Romer, Christina D., and David H. Romer. 1989. "Does Monetary Policy Matter? A New Test in the Spirit of Friedman and Schwartz." *NBER Macroeconomics Annual*, 4: 121–70.

Sarkees, Meredith R. 2000. "The Correlates of War Data on War: An Update to 1997." *Conflict Management and Peace Science*, 18(1): 123–44.

Singer, J. David, and Melvin Small. 1994. "Correlates of War Project: International and Civil War Data, 1816–1992." Inter-university Consortium for Political and Social Research [distributor], Ann Arbor, MI.