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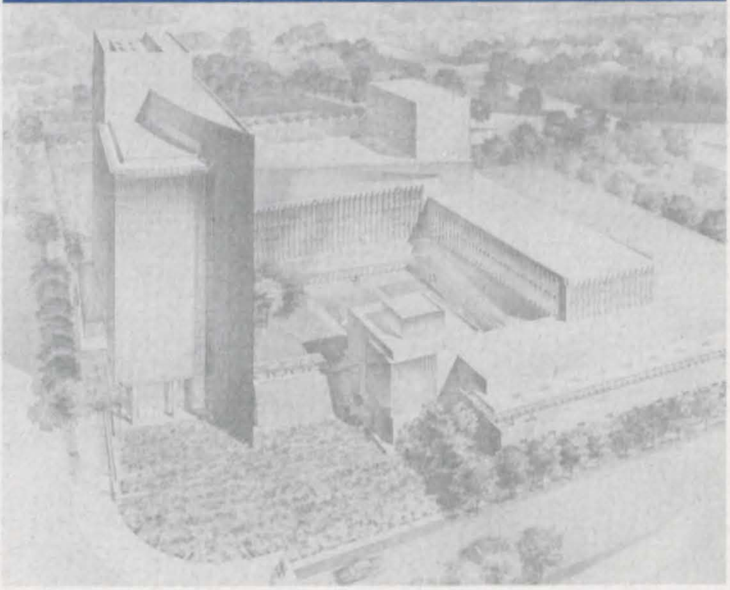

Recessions and Recoveries

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On the cover: an architectural rendering of the new Federal Reserve Bank of Dallas headquarters.

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Chronic inflation and the importance of the exchange rate as a nominal anchor for the domestic price level have led some Latin American countries to consider returning to a fixed dollar exchange rate. John Welch and Darryl McLeod examine the costs and benefits of real exchange rate movements and their relevance for the credibility of inflation policies in countries now contemplating free trade agreements with the United States.

The authors discuss the experiences of several Latin American countries and describe the problem their policymakers face when deciding to follow either fixed or flexible exchange rate rules. Fixed exchange rates that are credible can decrease inflation rates, but only at the cost of policy flexibility in the face of adverse changes in the terms of trade or foreign interest rates. The current relative stability of international markets has led some Latin American countries to complement their stabilization and reform policies with fixed exchange rates.

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Recessions and Recoveries

The U.S. economy entered recession in July 1990 and began to recover, many analysts believe, in April or May 1991. Since then, the economy has grown at a pace so sluggish as to be indistinguishable, in some ways, from continued recession. However, as early as spring 1991, several observers were expressing the opinion that the recovery from the 1990–91 recession would not be particularly robust because the recession itself was not particularly severe. For example, in *Business Week* in June 1991, Alan Blinder argued, "Shallow recessions are followed by weak recoveries for a simple reason: An economy that has not fallen far has little catching up to do. And catch-up is the main reason economies zoom upward in the early stages of recovery." *The Economist* magazine, in an editorial on January 18, 1992, pointed out that "there are good reasons to think that the coming expansion may be weaker than most of its predecessors," the main one being "the mildness of the recession that preceded it." Most recently, the Shadow Open Market Committee, an independent private group that critiques the actions of the Federal Reserve, argued that one of the main reasons the economy remained sluggish in 1992 was that "modest recessions are usually followed by modest recoveries" (1992, 5).

The notion that the economy experiences a "bounce-back" or "rubber-band" effect following declines in economic activity contains a certain amount of intuitive appeal but seems to have been subject to few empirical tests. The earliest study and one of the most comprehensive analyses of this issue that we have found is Moore (1961), who tried to test the view that "the strength of a recovery in its early stages depends upon the level from which it starts" (p. 86). He examined the behavior of groups of leading, coincident, and lagging indicators in the first seven months of six recoveries (the earliest being that following the trough in July 1924, the latest being that following the trough in August 1954, since revised to May 1954) and tentatively concluded that "recoveries in output, employment, and profits have usually been faster after severe depressions than after mild contractions" (p. 88). Moore (1965) contains a restatement of the finding that severe contractions tend to be followed by strong expansions, and Bry and Boschan (1971) present further evidence on this proposition, focusing on growth in nonagricultural employment.

Another of the few authors addressing this question is Milton Friedman, who asked, "Is the magnitude of an expansion related systematically to the magnitude of the succeeding contraction? Does a boom tend on the average to be followed by a large contraction? A mild expansion, by a mild contraction?" (Friedman 1969, 271). On the basis of an examination of simple rank correlation coefficients for three different measures of activity, Friedman found no relationship between the size of an expansion and the size of the succeeding contraction but did find that "a large contraction in output tends to be followed on the average by a large business expansion; a mild contraction, by a mild

We wish to thank Adrienne C. Slack and Shengyi Guo for excellent research assistance on this project. This article draws on work reported in an Economics Letters *article (Wynne and Balke 1992) and in an unpublished paper (Balke and Wynne 1992).*

expansion" (p. 273). Friedman went further, to sketch out a theory of business cycles (the "plucking model of fluctuations") that he felt was consistent with these patterns in activity, but to date his model seems to have received scant attention.

Finally, we note that Neftci (1986), in the course of addressing a slightly different question, reports results that are relevant to the recession–recovery relationship. Focusing on the behavior of pig iron production since the latter half of the nineteenth century, he finds a significant negative correlation between the length of expansions and the length of contractions: for every additional twelve months of expansion, the economy experiences 1.8 fewer months of contraction. However, the length of contractions does not affect the length of subsequent expansions. Furthermore, he shows that there is a significant relationship between the peak-to-trough decline in output and the increase over the course of the subsequent expansion but none between the gains in output over the expansion and the losses over the subsequent contraction.

In this article, we will study the behavior of output during and immediately after recessions to see whether there is any validity to the notion of a bounce-back effect.[1] Our analysis differs from that of Moore, Friedman, and Neftci in a number of ways. First, we focus on the behavior of industrial production rather than look at a variety of indicators. The reason is that we can obtain reasonably consistent estimates of industrial production for long periods, allowing us to look at recoveries from a large number of recessions. Second, we estimate a simple linear regression model rather than look at simple correlations, which enables us to discriminate between the effects of different measures of the severity of the preceding recession. The two measures of severity we focus on here are the depth of the recession, as measured by the output loss from the peak date to the trough date, and the length of the recession, as measured by the number of months from the peak date to the trough date. Third, we look at a larger number of recessions and recoveries than does either Moore or Friedman, including a number of pre–World War I business cycles. Each recession will be viewed as an independent event, and we will look for regularities common to the 23 recessions and recoveries that the United States has experienced over the past hundred years. Fourth, we only look at output growth in the early stages of an expansion (either the first six months or the first twelve months) and see how growth over this horizon is influenced by the severity of the preceding recession. This contrasts with Friedman's and Neftci's examination of the relationship between growth over the entire expansion and the severity of the preceding recession.

The article begins with a brief discussion of how the National Bureau of Economic Research (NBER) determines the dates of the peaks and troughs in economic activity that give the business cycle its name. We then specify a simple empirical model for testing hypotheses about the relationship between recessions and expansions. We document the existence of a significant bounce-back effect in various measures of U.S. industrial production and show that this finding is robust to a variety of potential criticisms. Having established the existence of a bounce-back effect, we provide some intuition about the economic forces behind it. We then consider the behavior of the economy during the recovery from the 1990–91 recession and show that it is consistent with the bounce-back effect.

The U.S. experience with recessions

The NBER is responsible for the dating of the peaks and troughs in economic activity that mark the onset of recessions and expansions.[2] The dating of business cycles by the NBER is based on a definition of business cycles first formulated by Wesley Clair Mitchell in 1927:

[1] Sichel (1992) also talks about a bounce-back effect following recessions in reference to a high-growth recovery phase at the beginning of an expansion but does not look at the relationship between the rate of growth during the recovery phase and output losses during the recession. He does, however, examine the predictive power of an output-gap variable for GNP growth, where the output-gap variable is defined as the deviation of GNP (gross national product) from its preceding peak value.

[2] Interested readers are referred to Moore and Zarnowitz (1986) for a detailed discussion of how the NBER dates business cycles. The discussion here is a very brief summary of their article.

> Business cycles are a type of fluctuation found in the aggregate economic activity of nations that organize their work mainly in business enterprises: a cycle consists of expansions occurring at about the same time in many economic activities, followed by similarly general recessions, contractions, and revivals which merge into the expansion phase of the next cycle; the sequence of changes is recurrent but not periodic; in duration business cycles vary from more than one year to ten or twelve years; they are not divisible into shorter cycles of similar character with amplitudes approximating their own (Moore and Zarnowitz 1986, 736).

Note that the definition refers to fluctuations in "aggregate economic activity" rather than a more precisely defined aggregate, such as gross national product (GNP), industrial production, or total employment. This vagueness is intentional and recognizes that business cycles consist of movements in many different series that are not readily reduced to a single aggregate. Looking at a variety of series also helps minimize the risk of drawing erroneous conclusions based on mismeasurement. Finally, under the NBER definition, a period of slow, or "subpar," growth does not qualify as a contraction. Rather, peaks in activity are followed by periods of absolute decline in aggregate activity.[3]

A recession is defined as a peak-to-trough movement in economic activity. According to the NBER business-cycle chronology, the United States has experienced 30 recessions since the middle of the nineteenth century (see Burns and Mitchell 1946, Table 16; Moore and Zarnowitz 1986, Tables A.3, A.5). The dates of the peaks and troughs in U.S. economic activity chosen by the NBER are given in Table 1, along with statistics on the duration of expansions and contractions for the entire period. The chronology ends with the date of the most recent peak, July 1990.

At the time of our analysis (October 1992), the date of the trough marking the end of the most recent recession had not been announced officially, but several observers (including Moore 1992) have placed it in April or May 1991.[4] Addi-tional clues to the likely date of the most recent trough can be obtained from examining the recent behavior of the U.S. Commerce Department's composite index of coincident indicators. This index is explicitly designed to approximate cyclical movements in economic activity and to have turning points that match the business cycle. The coincident index peaked most recently in June 1990, just one month before the official peak in July, and seemed to hit a trough in January 1992. However, revisions to the index currently being undertaken by the Commerce Department and discussed in Green and Beckman (1992) move the trough in the index back to March 1991.

From the table we can see that the United States has experienced nine recessions since the end of World War II. This is a rather small sample for testing the idea that severe recessions tend to be followed by strong recoveries, so it is important to include pre–World War II recessions in our sample to be reasonably confident of our findings.[5] However, extending the statistical analysis to the pre–World War II period leads to problems of data availability and consistency. Furthermore, because the NBER chronology dates business-cycle peaks and troughs by month, a monthly indicator of economic activity is preferable for examining the hypothesis that deep recessions are followed by strong recoveries.

The requirement that the selected measure of aggregate economic activity be available at a monthly frequency and extend back to the prewar period leads us to use industrial production, as measured by the Federal Reserve Board's index of industrial production.[6] This index has the advan-

[3] *This is not true, however, of "growth cycle" chronologies.*

[4] *See Hall (1992) for a discussion of the problem of determining the date of troughs in economic activity.*

[5] *Alternatively, we could look at the experience of other countries in the postwar period. Thus, in Balke and Wynne (1992), we look for a bounce-back effect in the Group of Seven countries during the postwar period, using the NBER's "growth cycle" chronology for these countries.*

[6] *Moore (1961, 88) notes that the relationship between the severity of a recession and the strength of the subsequent recovery is strongest for industrial production.*

Table 1
NBER Business-Cycle Chronology for United States

Peak	Trough	Duration (Months) Contraction	Duration (Months) Expansion
June 1857	December 1858	18	22
October 1860	June 1861	8	46
April 1865	December 1867	32	18
June 1869	December 1870	18	34
October 1873	March 1879	65	36
March 1882	May 1885	38	22
March 1887	April 1888	13	27
July 1890	May 1891	10	20
January 1893	June 1894	17	18
December 1895	June 1897	18	24
June 1899	December 1900	18	21
September 1902	August 1904	23	33
May 1907	June 1908	13	19
January 1910	January 1912	24	12
January 1913	December 1914	23	44
August 1918	March 1919	7	10
January 1920	July 1921	18	22
May 1923	July 1924	14	27
October 1926	November 1927	13	21
August 1929	March 1933	43	50
May 1937	June 1938	13	80
February 1945	October 1945	8	37
November 1948	October 1949	11	45
July 1953	May 1954	10	39
August 1957	April 1958	8	24
April 1960	February 1961	10	106
December 1969	November 1970	11	36
November 1973	March 1975	16	58
January 1980	July 1980	6	12
July 1981	November 1982	16	92
July 1990	n.a.	n.a.	n.a.

Comparative statistics

	Average length of contractions	Average length of expansions
Pre–World War II	21.2	28.9
Post–World War II	10.7	49.9

n.a.—Not available.

NOTE: Length of contraction is the number of months from peak to trough.
Length of expansion is the length of the expansion after the trough date.

SOURCE: Moore and Zarnowitz (1986), Tables A.3, A.5.

tage of extending back to 1919, thus adding to the sample of recessions. The obvious drawback is that industrial production is an incomplete indicator of aggregate economic activity: industrial production currently accounts for only about one-fifth of total output. Looking at a broader measure of output, such as GNP, would probably be better; however, GNP estimates are available only on a quarterly basis and only as far back as 1947. On the other hand, movements in GNP and industrial production are highly correlated, with correlations of 0.998 using annual data and 0.964 using quarterly data.[7] Also, industrial production is a component of the index of coincident indicators, which is explicitly designed to have turning points that are the same as those of the business cycle.

Is there a bounce-back effect?

A useful first pass at answering the question of whether severe recessions are followed by strong recoveries is, simply, to plot the data. Figure 1 presents a scatter plot of the percentage change in output in each of the 14 recessions since 1919 (except the 1990–91 recession) against output growth in the first twelve months of the subsequent expansion, using the NBER business-cycle dates. A 135-degree line is included for reference. The scatter of points in Figure 1 certainly suggests the existence of some degree of correlation between the decline in industrial production over the course of a recession and growth in the first twelve months of an expansion. Obviously, the Great Depression (August 1929–March 1933) is very influential in suggesting the existence of a self-correcting mechanism, but it is clear that more is going on.[8]

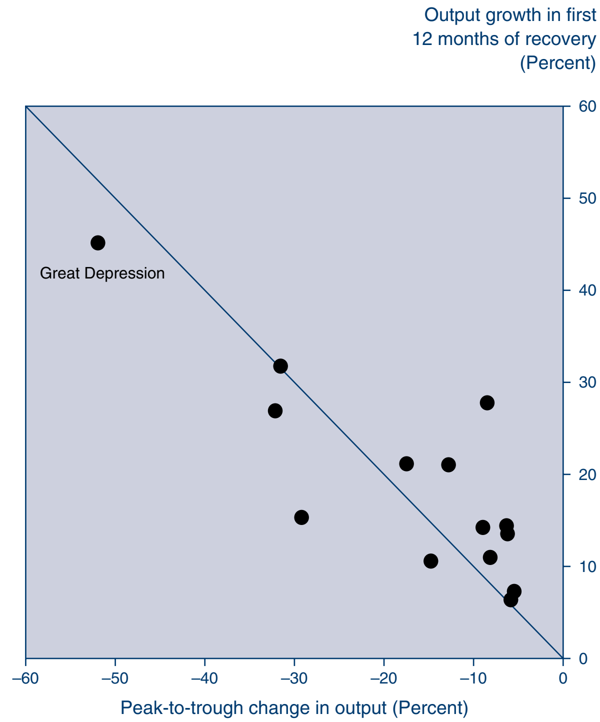
This simple ocular analysis of the data suggests that there is a correlation between the peak-to-trough decline in output over the course of recession and growth in the early stages of the subsequent recovery. Let us now turn to testing and quantifying the strength of this correlation.

Empirical analysis

Our strategy for testing for the existence of a bounce-back effect was to estimate a simple linear regression model of the form

Figure 1
Peak-to-Trough Change in Output and Output Growth, as Measured by Industrial Production

(NBER Business-Cycle Dates)



SOURCES OF PRIMARY DATA:
Board of Governors, Federal Reserve System.
Moore and Zarnowitz (1986).

[7] *Correlations were calculated using annual data for 1929–90 and quarterly data for 1947–90.*

[8] *The recovery from the Great Depression of 1929–33 has recently been examined in some detail by Romer (1991). The specific question she addresses is, What proportion of the extraordinary rates of real GNP growth observed in the mid-1930s and late 1930s can be attributed to the severity of the downturn, and what proportion can be attributed to monetary and fiscal stimuli to aggregate demand? She finds that stimulative monetary policy in the form of unsterilized gold inflows played a key role in the recovery, and she concludes that her findings suggest that "any self-corrective response of the U.S. economy to low output was weak or non-existent in the 1930s" (p. 1). The role of activist fiscal and monetary policy in generating vigorous recoveries is an issue we do not address directly in this article.*

Table 2
Rate of Growth During First Twelve Months of Recovery

	Constant	Change from peak to trough	Length of recession	$\bar{R}^2$
Industrial production	8.24**	–.63***	—	.64
	(2.79)	(.13)		
	6.27	–.47**	.34	.65
	(3.25)	(.19)	(.30)	
Manufacturing	9.13**	–.62***	—	.58
	(3.35)	(.14)		
	5.19	–.35*	.63*	.66
	(3.58)	(.19)	(.31)	
Durables manufacturing	–.69	–1.31***	—	.68
	(8.22)	(.24)		
	–10.50	–.83**	1.61**	.77
	(8.24)	(.29)	(.70)	
Nondurables manufacturing	9.00***	–.42***	—	.59
	(1.16)	(.10)		
	6.90***	–.24*	.26*	.66
	(1.51)	(.13)	(.13)	

* Significant at the 10-percent level.
** Significant at the 5-percent level.
*** Significant at the 1-percent level.

NOTE: All data were seasonally adjusted. The sample period is January 1919–December 1991, which includes 14 recessions, not counting the 1990–91 recession. Peak and trough dates are from the official NBER business-cycle chronology. The dependent variable is the rate of growth during the first twelve months of recovery (defined as trough to trough plus twelve months).
Figures in parentheses are standard errors.

$$(1)\quad \left(\frac{Y_{T+12}-Y_T}{Y_T}\right)_i = \alpha_0 + \alpha_1\left(\frac{Y_T-Y_P}{Y_P}\right)_i + \alpha_2(T-P)_i + \epsilon_i,$$

where Y is some measure of output, T denotes the month of a business-cycle trough as determined by some business-cycle chronology, P denotes the month of a business-cycle peak, i indexes recessions, and ϵ is an error term.[9] The dependent variable is the percentage increase in output in the twelve months after the trough month.[10] The explanatory variables, apart from the constant, are the peak-to-trough change in output in percentage terms and the length of the recession in months.

[9] *In Balke and Wynne (1992), we estimate a slightly different model that allows us to distinguish between three measures of the severity of a recession—length, depth, and steepness. Moore (1961, 86) notes that recessions have at least three dimensions—"depth, duration, and diffusion." We do not consider diffusion as a measure of severity in this article, primarily because of the degrees-of-freedom problem.*

[10] *Looking at growth beyond twelve months is complicated by the fact that for three of the recessions in our sample, the subsequent expansion lasted twelve months or less.*

Table 3
Rate of Growth During First Twelve Months of Recovery, Excluding the Great Depression

	Constant	Change from peak to trough	Length of recession	$\bar{R}^2$
Industrial production	9.72**	–.51**	—	.35
	(3.22)	(.19)		
	6.26	–.47**	.34	.31
	(6.73)	(.20)	(.57)	
Manufacturing	11.51**	–.43**	—	.25
	(3.64)	(.19)		
	2.99	–.36*	.81	.30
	(7.39)	(.19)	(.62)	
Durables manufacturing	5.52	–.97***	—	.43
	(8.62)	(.31)		
	–16.19	–.84**	2.09	.48
	(16.88)	(.31)	(1.42)	
Nondurables manufacturing	9.61***	–.28**	—	.29
	(1.12)	(.12)		
	7.69	–.24	.19	.25
	(3.09)	(.14)	(.28)	

* Significant at the 10-percent level.
** Significant at the 5-percent level.
*** Significant at the 1-percent level.

NOTE: All data were seasonally adjusted. The sample period is January 1919–December 1991, which includes 14 recessions, not counting the 1990–91 recession. Peak and trough dates are from the official NBER business-cycle chronology. The dependent variable is the rate of growth during the first twelve months of recovery (defined as trough to trough plus twelve months).
Figures in parentheses are standard errors.

If deep recessions are followed by strong recoveries, the estimate of α_1 should be negative. If long recessions are followed by strong recoveries, the estimate of α_2 should be positive.

Table 2 reports estimates of this model using the Federal Reserve's industrial production index and its three principal components—total manufacturing, durables manufacturing, and nondurables manufacturing. Results are reported both with and without the length-of-recession variable on the right-hand side. The sample includes 14 recessions, starting with the January 1920–July 1921 recession and ending with the July 1981–November 1982 recession, as determined by the NBER business-cycle chronology. For each production category, there is a statistically significant relationship between the size of the peak-to-trough decline and growth in the twelve months after the trough. The size of the bounce-back effect is strongest for durables manufacturing. Recession length makes no difference to the strength of the recovery in total industrial production but does seem to be important for manufacturing. Within manufacturing, recovery in the durable goods sector seems to be more affected by the length of the recession than is the recovery in the nondurables sector. For all sectors, including the length of the recession as an additional variable on the right-hand side lessens the bounce-back effect but does not eliminate it.

Because the sample period includes the Great Depression, one of the most severe contractions ever in U.S. economic activity, the results in Table 2 may be overly influenced by this extraordinary event. Table 3 reports results from estimation of equation 1 when the Great Depression is excluded from the sample. As might be expected, there is some loss of statistical significance, but the results are broadly similar to those in Table 2. The length of the recession is no longer significant in explaining growth during the first twelve months of recovery. This is not too surprising, because the Great Depression, with forty-three months from peak to trough, is by far the longest recession in the period covered by our analysis.[11]

How robust are the results?

How robust are our findings of a bounce-back effect? We have already examined the sensitivity of the findings to the inclusion of the Great Depression in the sample and have seen that the results are not sensitive to its exclusion. In this section, we will consider the robustness of our results to a variety of other potential criticisms. First, we will consider growth over horizons other than the twelve months after the trough date. Specifically, we will consider whether growth in the first six months of an expansion is also significantly related to the severity of the preceding recession. Second, we will increase the number of recessions we look at by examining the behavior of an alternative industrial production index constructed by Miron and Romer (1990) that covers the period 1884–1940. We also consider the behavior of this index when it is spliced to the Federal Reserve production index in 1919. Finally, we consider the sensitivity of our results to use of the official NBER chronology by looking at the dates suggested by Romer (1992) and the dates obtained using the algorithm developed by Bry and Boschan (1971).

The bounce-back effect at the six-month horizon. To examine whether the bounce-back effect can be found at the six-month horizon as well, we estimated an obvious variant on equation 1, redefining the dependent variable to be growth in the first six months after the trough. The results are reported in Table 4. We only report the results obtained when length of recession is not included in the model, as the significance of this variable seems to hinge completely on including the Great Depression in the sample. Growth in the first six months of the recovery is significantly correlated with the peak-to-trough change in activity, but excluding the Great Depression from the sample seems to reduce the strength of the correlation a lot more than we find for growth over the twelve-month horizon.

The bounce-back effect in the Miron–Romer industrial production series. It is possible to extend the sample period further to include the period before World War I by using the industrial production index recently constructed by Miron and Romer (1990). Their index covers the period 1884–1940, overlapping with the Federal Reserve index for twenty-one years, from 1919 to 1940. This period of overlap includes five recessions. The Miron–Romer index was designed to improve upon the older Babson and Persons indexes, which made heavy use of indirect proxies for industrial activity (such as imports and exports in the case of the Babson index and bank clearings in the case of the Persons index). Miron and Romer note that their series has turning points (that is, peaks and troughs) that are “grossly similar to but subtly different from existing series” (p. 321).

The Miron–Romer index is less comprehensive than the Federal Reserve index and, according to the NBER chronology, produces two anomalous observations. Specifically, the Miron–Romer index shows industrial production *increasing* in two of the pre–World War I recessions, the recessions of December 1895–June 1897 and September 1902–August 1904. This finding can be interpreted as a drawback of the series or as suggesting a need to reconsider the dating of pre–World War I business cycles by using the improved index.

[11] The Great Depression is not, however, the longest recession in the NBER chronology. The longest U.S. recession on record was from October 1873 to March 1879, lasting sixty-five months. This recession is not included in our analysis because reliable measures of aggregate production at the required frequency are not available that far back.

Table 4
Rate of Growth During First Six Months of Recovery

	Constant	Change from peak to trough	$\bar{R}^2$
	Including the Great Depression		
Industrial production	.39	–.63***	.62
	(3.11)	(.14)	
Manufacturing	.71	–.65***	.61
	(3.51)	(.15)	
Durables manufacturing	–7.90	–1.08***	.61
	(8.45)	(.25)	
Nondurables manufacturing	2.68**	–.68***	.81
	(1.15)	(.10)	
	Excluding the Great Depression		
Industrial production	4.12	–.32*	.25
	(2.94)	(.17)	
Manufacturing	4.82	–.32*	.25
	(3.17)	(.17)	
Durables manufacturing	3.91	–.44*	.29
	(5.78)	(.21)	
Nondurables manufacturing	3.48***	–.49***	.67
	(.99)	(.11)	

* Significant at the 10-percent level.
** Significant at the 5-percent level.
*** Significant at the 1-percent level.

NOTE: All data were seasonally adjusted. The sample period is January 1919–December 1991, which includes 14 recessions, not counting the 1990–91 recession. Peak and trough dates are from the official NBER business-cycle chronology. The dependent variable is the rate of growth during the first six months of recovery (defined as trough to trough plus six months). Figures in parentheses are standard errors.

The results from estimating the model by using the Miron–Romer index are reported in Table 5. The first four rows of this table report the results obtained using the raw (not seasonally adjusted) series. Again, we find evidence of a significant bounce-back effect in industrial production. The inclusion of recession length as an additional explanatory variable makes no difference to this finding, nor does excluding the Great Depression. Table 5 also reports the results of combining the Federal Reserve and Miron–Romer indexes (seasonally adjusted). We followed Romer (1992) in splicing the two series in 1919 to obtain a single series on industrial production for the period 1884–1990. This gives us a sample of 24 recessions for examining the bounce-back effect. The principal difference between these results and those in Tables 3 and 4 is that length of recession is no longer significant in

Table 5
Rate of Growth During First Twelve Months of Recovery: Results Using the Miron–Romer Index

	Constant	Change from peak to trough	Length of recession	$\bar{R}^2$	Number of recessions
		Miron–Romer Index			
Including the Great Depression	12.48**	–.79***	—	.50	15
	(4.80)	(.20)			
	9.39	–.76***	.20	.47	15
	(9.86)	(.23)	(.55)		
Excluding the Great Depression	12.42**	–.82***	—	.45	14
	(5.00)	(.24)			
	.16	–.82***	.77	.44	14
	(15.03)	(.24)	(.89)		
		Combined Federal Reserve/Miron–Romer Index			
Including the Great Depression	9.62***	–.65***	—	.54	24
	(2.17)	(.12)			
	6.10	–.60***	.27	.55	24
	(3.71)	(.13)	(.23)		
Excluding the Great Depression	9.94***	–.60**	—	.38	23
	(2.30)	(.16)			
	3.99	–.64***	.40	.39	23
	(5.73)	(.16)	(.35)		

* Significant at the 10-percent level.
** Significant at the 5-percent level.
*** Significant at the 1-percent level.

NOTE: Peak and trough dates are from the official NBER business-cycle chronology. The dependent variable is the rate of growth during the first twelve months of recovery (defined as trough to trough plus twelve months).
Estimates in the first four rows were obtained using the non-seasonally-adjusted Miron–Romer index. The sample period is January 1884–December 1940, which includes 15 recessions. Estimates in the second four rows were obtained using the combined Federal Reserve/Miron–Romer series, which is seasonally adjusted, not counting the 1990–91 recession. The sample period is January 1884–December 1991, which includes 24 recessions.
Figures in parentheses are standard errors.

explaining the strength of the recovery, even when the Great Depression is included in the sample.

The bounce-back effect in alternative business-cycle chronologies. Romer (1992) has questioned whether the dates for the prewar cycles in the official NBER chronology are strictly comparable to those for the postwar period. Romer documents evidence that the prewar dates are based on detrended data while the postwar dates reflect cycles in unadjusted data. Consequently, the prewar NBER chronology tends to overstate the length of contractions and understate the length of expansions. Romer corrects the NBER chronology by formalizing the rule that the NBER used in dating the postwar cycles and applying it to industrial production for the prewar period to come up

Table 6
Alternative Prewar Business-Cycle Chronologies

NBER dates		Romer dates		Bry–Boschan dates	
Peak	**Trough**	**Peak**	**Trough**	**Peak**	**Trough**
March 1887	April 1888	February 1887	July 1887	—	—
July 1890	May 1891	—	—	November 1891	September 1893
January 1893	June 1894	January 1893	April 1894	—	—
December 1895	June 1897	December 1895	January 1897	October 1895	August 1896
—	—	—	—	April 1897	June 1898
June 1899	December 1900	April 1900	December 1900	April 1900	October 1900
—	—	—	—	August 1901	June 1902
September 1902	August 1904	July 1903	March 1904	—	—
May 1907	June 1908	July 1907	June 1908	—	—
January 1910	January 1912	January 1910	May 1911	February 1910	December 1910
January 1913	December 1914	June 1914	November 1914	December 1912	January 1914
—	—	May 1916	January 1917	—	—
August 1918	March 1919	July 1918	March 1919	May 1918	March 1919
January 1920	July 1921	January 1920	July 1921	January 1920	March 1921
May 1923	July 1924	May 1923	July 1924	May 1923	June 1924
October 1926	November 1927	March 1927	December 1927	March 1927	December 1927
August 1929	March 1933	September 1929	July 1932	May 1929	July 1932
May 1937	June 1938	August 1937	June 1938	May 1937	June 1938
—	—	December 1939	March 1940	—	—

SOURCES: Moore and Zarnowitz (1986), Tables A.3, A.5.
Romer (1992), Table 3.
Authors' calculations.

with an alternative set of dates.[12] These dates are shown in Table 6. One key difference with the official NBER dates (reproduced in Table 6 for ease of comparison) is that the average length of pre–World War II contractions is shorter (11.4 months, as opposed to 17.8 months in the NBER chronology), and the average length of pre–World War II expansions is longer (30.3 months, as opposed to 24.9 months in the NBER chronology).[13] The two chronologies are in agreement for only two reces-

[12] *This rule is explained in the Appendix.*

[13] *Note that these statistics compare the average length of contractions and expansions during the period for which the two chronologies overlap. The statistics on the average length of prewar contractions and expansions reported in Table 1 are the averages over* all *contractions and expansions in the NBER chronology for the prewar period.*

sions, those of 1920–21 and 1923–24. They are also in agreement on either the peak or the trough dates for a number of other recessions. Finally, note that Romer's chronology excludes one recession that is included in the NBER chronology, the 1890–91 recession, while including two others that are omitted from the NBER chronology, those in 1916–17 and 1939–40. One other noteworthy feature of Romer's chronology is that she dates the trough of the Great Depression in July 1932, which shortens the length of that downturn from forty-three months to thirty-four months.

Table 6 also contains business-cycle dates obtained from application of the algorithm devised by Bry and Boschan (1971) to industrial production for the entire period.[14] The Bry–Boschan algorithm is somewhat more complex than the rule devised by Romer and picks slightly different cycles from those picked by Romer and those in the official NBER chronology. The Bry–Boschan algorithm picks two cycles (1897–98 and 1901–2) that are not included in the NBER chronology and misses four (1887, 1893–94, 1903–4, and 1907–8) that are. The Bry–Boschan algorithm also misses the 1916–17 and 1939–40 cycles, two cycles picked by the Romer algorithm but not included in the NBER chronology. The algorithm does capture some of the same peak and trough dates as the Romer algorithm and the NBER chronology. Interestingly, the Bry–Boschan algorithm places the trough of the Great Depression in July 1932 (the same as Romer) but dates its onset in May 1929, four months earlier than Romer and three months earlier than the NBER.

Table 7 reports the results of estimating equation 1 with the Romer and Bry–Boschan business-cycle dates. For consistency, we used the dates picked by these algorithms for the postwar period as well, rather than the NBER dates. The differences between the three chronologies for the postwar period are minor, as both the Romer and Bry–Boschan algorithms are designed to match as closely as possible the NBER dating for this period. Both chronologies suggest a statistically significant bounce-back effect. In every case, the coefficient estimate on the change in output from peak to trough is significant at the 1-percent level. The length of the recession is not significant in either chronology, even when the Great Depression is included. Excluding the Great Depression does significantly lower the explanatory power of the basic model, as indicated by the drop in the $\bar{R}^2$, and the size of the bounce-back effect, as indicated by the drop in the absolute value of the coefficient estimate, but does not eliminate it.

To summarize, our finding of a bounce-back effect in industrial production is common to a variety of measures of industrial production, is found at the six-month as well as the twelve-month horizon, and is robust to potential shortcomings in the NBER chronology for the prewar period. Some other robustness tests (such as controlling for secular trend) are reported in Balke and Wynne (1992) and reinforce those reported here. The robustness of the bounce-back effect merits taking it seriously as a stylized fact about the business cycle.

The economics of the bounce-back effect

Having established the existence of a bounce-back effect, we should provide an economic interpretation of what is going on. We argue that the bounce-back effect tells us more about the dynamic response of the economy to shocks than about the nature or source of shocks themselves. Three simple observations about the behavior of firms and households help in understanding macroeconomic dynamics. First, households and firms not only look at current economic conditions when deciding how much to work, save, consume, and invest but also take into consideration the likely course of economic activity in the future. Second, households prefer continuity in their consumption patterns from year to year, rather than wild movements. And third, saving and investment decisions made today have implications for what can be done tomorrow through their effect on capital accumulation, just as decisions made yesterday have implications for what can be done today. The bounce-back effect is a manifestation of the dynamic response of the economy, as a result of these three factors, to a shock that brings about a recession.

One interpretation of what happens when the economy goes into recession is that the maximum level of output that can be attained with

[14] *The Bry–Boschan algorithm is described briefly in the Appendix.*

Table 7
Results Using Alternative Business-Cycle Chronologies

	Constant	Change from peak to trough	Length of recession	$\bar{R}^2$	Number of recessions
		Romer Dating			
Including the Great Depression	7.01*	–.90***	–.03	.59	25
	(3.60)	(.18)	(.33)		
	6.81**	–.89***	—	.61	25
	(2.53)	(.14)			
Excluding the Great Depression	12.55**	–.73***	–.40	.35	24
	(4.60)	(.20)	(.38)		
	8.67***	–.70***	—	.34	24
	(2.78)	(.19)			
		Bry–Boschan Dating			
Including the Great Depression	–1.09	–.71***	.46	.71	21
	(3.60)	(.18)	(.33)		
	2.05	–.90***	—	.69	21
	(2.89)	(.13)			
Excluding the Great Depression	8.69*	–.60***	–.17	.38	20
	(4.90)	(.17)	(.37)		
	6.93**	–.57***	—	.41	20
	(2.87)	(.15)			

* Significant at the 10-percent level.
** Significant at the 5-percent level.
*** Significant at the 1-percent level.

NOTE: All data were seasonally adjusted. The sample period is January 1884–December 1991, which includes 25 recessions in the Romer chronology and 21 recessions in the Bry–Boschan chronology. The dependent variable is the rate of growth during the first twelve months of recovery (defined as trough to trough plus twelve months).
Figures in parentheses are standard errors.

existing resources of capital and labor temporarily falls. Such a change might come about, for example, as a result of a temporary increase in oil prices. This is the type of shock typically emphasized in New Classical real business cycle models. Or alternatively, a coordination failure results in productive resources becoming idle and output falling below potential. This story is more characteristic of New Keynesian analyses of the causes of recessions. During the period of lower output, households try to maintain their consumption levels by saving less. Part of this behavior translates into less investment by businesses and reduced purchases of consumer durables by households. The result of these spending decisions of households and firms is that when the economy hits the trough, stocks of business capital and household capital are below their "normal," or desired long-run, levels. This discrepancy between actual and normal levels of capital is then associated with an investment boom and an increase in purchases of consumer durables when the economy turns the corner. In some cases, the discrepancy in and of itself can be enough to bring a recession to an end and set the expansion in motion. The larger the discrepancy between actual and normal capital stocks at the trough, the faster the economy will grow in the months after the trough because of the greater amount of ground that has to be regained.

This explanation merely touches on some of the key elements of the more fully articulated theories essential for a complete understanding of the business cycle. In Balke and Wynne (1992), we carry out a detailed analysis of a prototypical real business cycle model and find that it performs reasonably well in generating the bounce-back phenomenon but does not capture other features of the business cycle.

Examination of the 1991–92 recovery

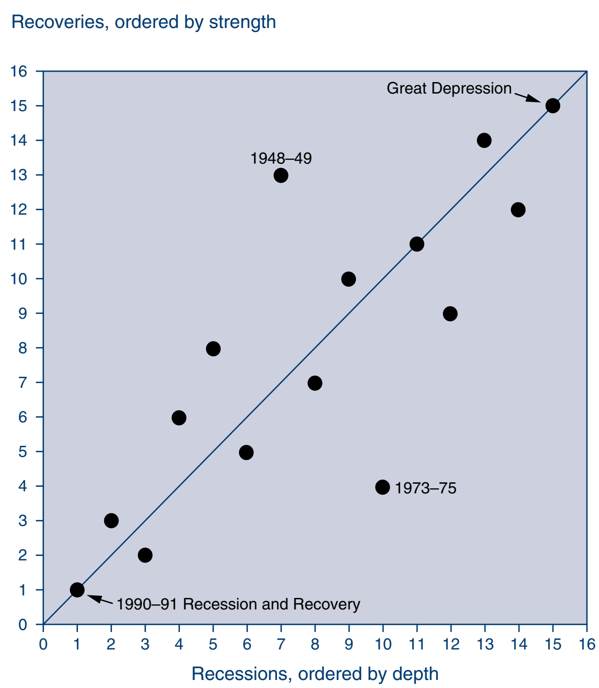
The most recent business-cycle peak was in July 1990. If we date the trough of this cycle as May 1991, as many analysts are doing (although the official trough date has yet to be announced by the NBER), the peak-to-trough decline in industrial production amounts to 3.6 percent. Industrial production bottomed out in March 1991, after declining 5.0 percent since July 1990. This compares favorably with either the average decline of 17.1 percent for all the recessions covered by the Federal Reserve's index of industrial production or the average decline of 8.8 percent for the post–World War II recessions. Based on our estimates in Table 2, we would expect industrial production to have grown 10.5 percent—8.24 – (0.63)(–3.6)—from the tentative trough date in May 1991 through May 1992. In fact, industrial production grew only 2.3 percent over this period, substantially less than the rate predicted by our simple model. If we take the actual peaks and troughs in industrial production, the decline from September 1990 to March 1991 is 5.2 percent, and predicted cumulative growth in industrial production from March 1991 to March 1992 is 11.5 percent—8.24 – (0.63)(–5.2)—as opposed to a realized rate of 2.5 percent. Thus, our bounce-back equation dramatically overpredicts the strength of the recovery, suggesting that the present recovery is abnormally slow, even after taking into account the shallowness of the recession.

However, because of the historical variability of the growth rate of industrial production during recoveries, the current recovery is still well within the 95-percent confidence interval implied by the bounce-back model. For the forecast growth rate over the twelve months since May 1991, the standard error associated with the forecast is 6.9 percentage points. This means that based on the coefficient estimates in Table 2, the 95-percent confidence interval associated with the forecast value of the growth rate of industrial production from May 1991 to May 1992 is 10.5 ± (1.96)(6.9)—that is, from –3.0 percent to 24.0 percent. Thus, the current recovery, while substantially weaker than predicted, is nonetheless consistent with the bounce-back model.

An alternative perspective on how this recovery compares with others is given in Figure 2. This figure is a scatter plot of the peak-to-trough decline in industrial production over the course of recession against growth in the first twelve months of the recovery, with the recessions and recoveries now ranked in order of severity and strength. Thus,

Figure 2
Rank Ordering of Recessions and Recoveries, as Measured by Industrial Production

(NBER Business-Cycle Dates)



SOURCES OF PRIMARY DATA:
Board of Governors, Federal Reserve System.
Moore and Zarnowitz (1986).

the horizontal axis ranks recessions in order of severity, with 1 being the least severe and 15 being the most severe. The vertical axis ranks recoveries in terms of their strength, with 1 being the least strong and 15 being the most strong. That the points are clustered around the 45-degree line is simply another way of demonstrating the bounce-back effect: typically, severe recessions are followed by strong recoveries. As we saw in Figure 1, the most severe recession in the sample covered by the Federal Reserve's industrial production index, the Great Depression, was also followed by the most robust recovery in that sample. What we see from Figure 2 is that the 1990–91 recession was the least severe since 1919 in terms of the decline in industrial production and, also, the recovery in the twelve months since the tentative trough date of May 1991 is the weakest since 1919. In other words, the behavior of the industrial sector in the most recent recession and recovery episode is very much in line with historical experience.

It cannot be emphasized strongly enough that this article focuses on the behavior of the industrial sector in recessions and recoveries. In terms of broader measures of aggregate activity, such as total nonagricultural employment or GNP, the picture is somewhat different. While the most recent recession may have been one of the least severe in U.S. history in terms of the decline in industrial production, it is close to the postwar average in terms of the decline in GNP. As for the recovery, GNP growth over the year since the tentative trough date of May 1991 is the weakest in the postwar period. Moreover, while the decline in manufacturing employment between July 1990 and May 1991 was the smallest in the postwar period, the twelve-month period after May 1991 is the only postwar "recovery" in which manufacturing employment declined. Outside the manufacturing sector, service-sector employment posted its weakest increase of any postwar recession except the 1957–58 recession—the only postwar recession in which service-sector employment declined. In the twelve months since May 1991, service-sector employment has grown less than in any other postwar recovery.

The 1990–91 recession and recovery episode generated many puzzles for policymakers that are not yet fully understood. With the passage of time, our understanding of what happened will grow. The sluggish pace of the overall recovery remains a puzzle, but the relatively modest growth in industrial output is consistent with the bounce-back effect that we have shown to be characteristic of previous recessions.

Conclusions

In this article, we have examined how rapidly industrial production recovers in the twelve months after a business-cycle trough. We considered two variables as candidates to explain differences in growth rates between recoveries—the depth and the length of the preceding recession. We found a statistically significant relationship between the rate of growth of output in the twelve months after a business-cycle trough and the size of the decline in output from peak to trough. Furthermore, the bounce-back effect appears to be stronger in durables manufacturing than in nondurables manufacturing. The existence of this bounce-back effect does not depend on including the Great Depression in our sample. However, the length of the recession makes a difference for the strength of the subsequent recovery only if the recovery following the Great Depression is included in the sample.

In Balke and Wynne (1992), we have examined the bounce-back effect in greater detail and have shown that a similar phenomenon seems to characterize the behavior of the Group of Seven countries in the postwar period. In that paper, we also look at the "shape" of cyclical movements in various aggregates and document significant asymmetries between expansions and contractions. In addition, we explore the implications of these findings for some common (linear) statistical and economic models of industrial output.

Given the relative robustness of our finding of a bounce-back effect for the industrial sector, it is important to ask whether the effect characterizes the 1990–91 recession and recovery. If we take May 1991 as the trough date marking the end of the most recent recession, the decline in industrial output from peak to trough was 3.6 percent, making it one of the mildest recessions in terms of lost industrial production. And consistent with the bounce-back effect, the growth in industrial production since May 1991 has been the weakest recovery in the period covered by the Federal Reserve index of industrial production.

Appendix

Rules for Dating Business Cycles

Romer's rules for dating cycles by using industrial production

1. A fluctuation counts as a cycle if the cumulative loss in the log of output between the peak and the return to peak exceeds 0.44—that is, 44 percentage-point months of output.
2. The second or later of multiple extremes is chosen as the turning point if the cumulative loss or gain in output is less than 0.11.
3. The first month after a peak or trough counts as a horizontal stretch if the cumulative loss or gain in output is less than 0.008.

The Bry–Boschan algorithm for picking turning points in a time series

1. Eliminate extreme values of raw series (greater than ± 3.5 standard deviations) and replace by values from a Spencer curve. A Spencer curve is a symmetric filter with declining weights.
2. Calculate a twelve-month moving average with the adjusted series. Find the local maximums and minimums. Use dates as tentative peak and trough dates, being sure that peaks and troughs alternate.
3. Calculate a Spencer curve with the adjusted series. Find the highest (lowest) values of Spencer curves within five months of the peaks (troughs) identified from the twelve-month moving average. Be sure that the new peak and trough dates alternate and that cycle duration is at least fifteen months.
4. Calculate a four-month moving average with the adjusted series. Identify the highest (lowest) values within five months of the peaks (troughs) identified from the Spencer curve. Be sure that the peak and trough dates alternate and that cycle duration is at least fifteen months.
5. Using the raw series, adjusted for extremes, find the highest (lowest) values within four months of the peaks (troughs) identified from the four-month moving average. Be sure that no peak or trough is within six months of the beginning or end of the sample, that peaks and troughs alternate, that cycle duration is at least fifteen months, and that expansion and contraction phases are at least five months long. The resulting peak and trough dates represent the final turning points.

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A Look at Long-Term Developments in the Distribution of Income

Strong economic growth in the United States during the last half of the 1980s did not translate into economic gains for all income groups. Poverty rates, for example, remained higher than those observed in the 1970s.[1] To paraphrase the most common findings, the rich got substantially richer during the 1980s, while the poor may have gotten poorer.

A trend toward greater income inequality can be cause for concern. Most Americans would not consider it desirable if, over time, all our society's resources became concentrated in the hands of a small group of individuals. On the other hand, few Americans would desire a perfectly equal distribution of income because income equality implies, among other things, that people who are college educated earn exactly the same income as people who are high school dropouts. If everyone earned the same income, there would be little incentive for people to work harder, become better educated, or find better, more efficient methods of production. Thus, the reasons underlying a trend toward greater income inequality are at least as important for policy analysis as the level of income inequality.

We set out to investigate how and why the distribution of income has changed over time. We find that the distribution of income has been becoming more unequal since the early 1950s, making what occurred in the 1980s a continuation of a longer-running trend. We also find that the distribution of income gains over the past dozen years is close to its historical average. Finally, we examine how rising income inequality relates to changes in the economy's demographic, business-cycle, and policy characteristics. We find that factors outside of direct policy control, such as the age and education profiles of the population, the gender composition of the labor force, and (mostly) inertia in income inequality, explain the lion's share of the forecast error variance. Policy variables, such as transfer payments and tax rates, account for only 15 percent of the variation in prediction errors.

What has happened to the distribution of income?

Our description of changes in the distribution of income proceeds in two parts. In the first part, we divide the population into five equal-sized subgroups, or quintiles, and examine each subgroup's gains from income growth.[2] In the second

We wish to thank Zsolt Besci, Stephen P.A. Brown, Christopher Carroll, and Keith R. Phillips for their helpful comments and suggestions, and Anne E. King and Adrienne C. Slack for their research assistance.

[1] *A change in how poverty rates were calculated means that poverty rates before 1975 are not comparable to those since 1975. Poverty rates stayed below 10 percent during the period 1975–80. During the 1980s, poverty rates climbed and then fell, staying above the 10-percent threshold. While economic growth appears to have roughly coincided with the declines in poverty rates, growth failed to lift enough people out of poverty to reduce the poverty rates below 10 percent.*

[2] *Another issue arises because we use tax returns as our data source. People do not have to file tax returns if their income levels are too low. Consequently, the sample we use is truncated in the sense that the lowest paid people are omitted.*

part, we look at developments in an aggregate measure of income inequality known as the Theil entropy index. The Theil index measures the degree of income inequality across the entire population in one number.

The distribution of income gains. Much of the recent attention to the issue of income inequality has focused not on the distribution of income but on how the gains from income growth were distributed across different income strata. According to work by Paul R. Krugman related in a memorandum from the Congressional Budget Office (1992), the top-paid 20 percent of the population received 94 percent of the gains in after-tax income between 1977 and 1989. In contrast, a 1992 U.S. Treasury report shows that people who were among the richest 1 percent of American taxpayers in 1979 received only 11.3 percent of the total gains in income during the 1980s (Sylvia Nasar 1992). In the *New York Times*, Nasar quotes Isabel V. Sawhill's finding that people who were in the top-paid 20 percent of the population in 1977 saw their incomes decline 11 percent over the next decade, while people who had climbed into the top category by 1986 had experienced, on average, a 65-percent increase in income.

To contribute to these discussions, we calculate summary statistics for the proportion of gains from income growth received by each population subgroup over the recent twelve-year period and compare those results with what occurred over the entire 1952–89 sample.[3] We find that nearly 60 percent of the gains in adjusted gross income during the period 1977–89 accrued to the top income quintile.

Analysts can reach such strikingly dissimilar conclusions because analyses of the distribution of income are very sensitive to the definitions of income, time horizon, and population with which the analyst works. For example, there are many possible definitions of income. One can examine the distribution of total income before taxes and transfers, total income after taxes but before transfers (such as Aid to Families with Dependent Children), total income after taxes and transfers, wage income, and still other variations. It is not too surprising that one finds different results when comparing, say, the income distribution of individuals with the income distribution of households, or the income distribution of taxpayers with the income distribution of all persons.

We focus on adjusted gross incomes rather than after-tax incomes because reliable data on both taxes and transfers are not available. We consider misleading any estimates of the income distribution after taxes but before transfers because they illustrate only part of the government's redistributive activities. In our opinion, one should either analyze the distribution of total factor incomes before taxes and transfers, which indicates the distribution of market-based claims on society's resources, or the distribution of income after both taxes and transfers, which indicates the distribution of purchasing power. Analyses of after-tax incomes that do not include information on transfers are neither fish nor fowl and are very problematic to interpret.

Analysts can also reach different conclusions from one another when they use different measurement techniques. In Krugman's analysis, the share of income gains accruing to population group i is represented as

$$(1) \qquad P_i = \Delta\mu_i \,/\, n\Delta\mu,$$

where μ_i is the change in average income for income group i, μ is the change in average income for the total population, and n is the number of equal-sized income groups. One can interpret equation 1 as the ratio of the weighted-average gain of a specified group to the average gain of the population as a whole. The equation highlights changes over time in the average income of a quintile.

The Council of Economic Advisers (CEA) analyzes growth in income by quintile for the 1992 *Economic Report of the President*. We com-

[3] Annual adjusted-gross-income data for this study are obtained from various issues of Statistics of Income, published by the United States Internal Revenue Service (IRS). For our analysis, we divide the population into quintiles, as follows: the IRS organizes tax returns by adjusted gross income, ranking by groups from lowest paid to highest paid. By dividing the total number of returns by five, we obtain the number of returns for each quintile. Thus, between any two periods the change in adjusted gross income earned by each quintile necessarily equals the aggregate change in adjusted gross income.

bine the CEA and Krugman approaches to represent the share of income gains earned by the ith group of the population as

$$(2) \qquad \hat{P}_i = \Delta Y_i \,/\, \Delta Y,$$

where Y_i is the change in income received by the ith group, and Y is the change in income for society as a whole. Equation 2, therefore, is the ratio of income gains received by the ith group to the income gains received by the population as a whole. Note that the Krugman and CEA-based measures yield identical results only when the population size is constant.

The following example illustrates how the interaction of population and income growth can affect the distribution of income gains as measured by the Krugman approach and the CEA-based approach. Suppose the economy has two workers with annual incomes of $30,000 and $20,000, respectively. The next year, these same workers earn $40,000 and $30,000, respectively, and two new workers obtain jobs and earn $20,000 each. Total income increased by $60,000. Average income among wage earners increased by $2,500 (from $25,000 to $27,500), while average income for the top half of the distribution increased by $5,000 (from $30,000 to $35,000). According to Krugman's original calculation, the top half of the distribution accounted for 100 percent of the gains from economic growth—$5,000/(2 × $2,500). Average income for the bottom half of the distribution did not change, indicating that the bottom half accounted for zero percent of the gains from economic growth using equation 1. Therefore, although each of the four participants in this hypothetical society earned more in the second year than they did in the first, the Krugman measure would indicate that all of the income gains accrued to the top half of the distribution.[4]

Using the CEA-based technique, the interpretation is somewhat different. The top half of the distribution received $40,000 more ($70,000 – $30,000), while the bottom half of the distribution received $20,000 more ($40,000 – $20,000). Thus, the top half accounted for 66.6 percent of the gains from economic growth using equation 2, while the bottom half received 33.3 percent of the gains.

The intuition behind the difference between Krugman's approach and the CEA-based approach is fairly straightforward. In Krugman's approach, income received by a particular subgroup of the population must grow at a rate faster than population growth. Otherwise, average (per capita) income for that subgroup would not rise, and Krugman's measure would indicate that they failed to share in the income gains. Thus, if the original workers in the example above earned $35,000 and $25,000, respectively, in the second year, and all other aspects of the example remained unchanged, then the average income of the top half of the distribution would remain at $30,000 [($35,000 + $25,000)/2], and the average income of the bottom half of the distribution would remain at $20,000. Krugman's measure would indicate that neither group has experienced any income gains.

By the CEA-based measure (equation 2), the condition for subgroups to share in income gains is that the sum of population growth and income growth for that particular subgroup exceeds zero. Given sufficient population growth, it is possible for the CEA-based measure to indicate that each quintile experienced income gains even when average income was falling for all quintiles. Thus, relative to Krugman's measure, the CEA method requires that a weaker condition is satisfied for any one subgroup to have a positive share in the distribution of income gains.

Both Krugman's and the CEA-based measures do not use longitudinal data. Neither of the statistics follows a particular group of people through time to trace how much of the aggregate gains are distributed to that group. Therefore, when there is substantial income mobility, these measures of the aggregate economy say little about the incomes received by specific individuals. However, they say much about the distribution of possible incomes and, therefore, about individual opportunities. (For a discussion of income mobility in the United States, see the box titled "Trends in Income Mobility.")

Table 1 reports the distribution of changes in income for several periods. Specifically, the table

[4] *Michael Boskin (1992) lays out this example in describing Krugman's distribution of income gains.*

Trends in Income Mobility

From year to year, people can and do move from one income class to another. For periods as long as a decade, the changes in people's income, especially for people in the lowest income group, are remarkable. Table B1 shows movements in the income distribution from 1979 to 1988.[1] The data indicate substantial income mobility, particularly among the lower income groups. Only 65 percent of the people who were in the top-paid 20 percent of the population in 1979 were still in the top-paid 20 percent in 1988. Two-thirds of the people in the middle income quintile changed classification over that ten-year period, while more than 85 percent of the people in the lowest income quintile changed income classifications. More than 17 percent of the people in the lowest income category in 1979 had climbed into the highest income category by 1988. Except for people in the highest income category (who, by definition, could not improve), those who changed income quintiles were more likely to move up than down.

One caveat to interpreting this evidence is that the study follows individuals who filed IRS returns in each of the ten years from 1979 through 1988. Accordingly, those who earned such low amounts that they did not have to file returns in any of the ten years were omitted from the sample. These people may well be permanently poor, making the upward mobility evidence less strong. Further, mobility out of the lowest income categories may be overstated because the low income groups in 1979 undoubtedly include students and part-time workers who became better compensated as they accumulated education and experience.

[1] See Joel Slemrod (1992) for evidence on the upward bias imparted to income inequality when looking at year-to-year income changes. Slemrod calculates the average income for each taxpayer over the seven-year period from 1979 to 1985. Slemrod refers to this approach as a time exposure. Compared with the snapshots of the income inequality over the same time period, the time-exposure Gini coefficient is roughly 7 percent lower, suggesting that income inequality declines somewhat as the time horizon lengthens. The findings are consistent with changes in income from period to period that are smoothed over when one uses income measured over several years, instead of capturing jumps in income that occur in any given year.

Table B1
Changes in Income Quintiles, 1979 and 1988

	Status in 1988				
Status in 1979	**Top-paid 20%**	**Next highest paid 20%**	**Middle 20%**	**Next lowest paid 20%**	**Lowest paid 20%**
Top 20%	64.7	20.3	9.4	4.4	1.1
Next highest paid 20%	35.4	37.5	14.8	9.3	3.1
Middle 20%	15.0	32.3	33.0	14.0	5.7
Next lowest paid 20%	11.1	19.5	29.6	29.0	10.9
Lowest paid 20%	17.7	25.3	25.0	20.7	14.2

Table 1
Percentage of Income Gains Distributed Among the Five Population Quintiles

Using the CEA-based method:

Period	Q1	Q2	Q3	Q4	Q5
1977–89	2.7	6.4	12.1	21.8	57.1
1980–85	3.3	9.8	11.0	23.6	52.4
1985–89	1.0	2.3	11.4	17.3	68.0

Using Krugman's method:

Period	Q1	Q2	Q3	Q4	Q5
1977–89	2.8	5.8	11.1	20.8	59.5
1980–85	3.3	10.0	10.5	23.3	53.0
1985–89	1.8	0.0	10.7	14.7	74.5

reports the changes in adjusted gross income received by each quintile for the periods 1977–89, 1980–85, and 1985–89.[5] The top half reports the distribution of gains in nominal, pre-tax, and transfer income using the CEA-based method (equation 2). The bottom half of the table reports the distribution of income gains from the same data using the Krugman method (equation 1).

Somewhat surprisingly, the results using the Krugman method are quite similar to those using the CEA-based method. The Krugman method indicates that a slightly higher percentage of income gains is going to the top quintile than indicated by the CEA-based method, but this difference does not change the implication that the top quintile reaped the majority of the income gains. Table 1 shows that over the period 1977–89, about 60 percent of the gains in factor income (income before taxes and transfers) went to the top-paid quintile.

Another question is how income gains are distributed across different subperiods. For example, were the 1980–85 or 1985–89 periods substantially different in terms of how income gains were distributed? The evidence presented in Table 1 suggests that the 1985–89 period saw gains going more to the highest paid quintiles and less to the lower-paid quintiles. For example, during the 1980–85 period the two lowest paid quintiles accounted for slightly more than 13 percent of the income gains, while during the 1985–89 period the same two quintiles accounted for less than 4 percent of the income gains. The share of income growth received by the middle-paid quintile was virtually unchanged from the 1980–85 period to the 1985–89 period, while the share of income growth received by the second highest paid quintile declined more than 25 percent. The declines in the first, second, and fourth quintiles were matched by the increases of the highest paid quintile. Using the CEA-based method, the top-paid quintile accounted for about

[5] *These statistics also represent the distribution of real income gains, assuming that each quintile has the same deflator.*

52 percent of the income gains in the first half of the 1980s, rising to about 68 percent of the gains in the second half.

By reporting the historical averages received by each quintile, one can see how these recent time periods compare with the entire sample. Using the CEA-based method, we calculate the share of income gains received by each quintile annually for the period 1952–89. As the evidence in Table 2 illustrates, on average the two lowest paid quintiles received about 4.5 percent of the real income gains over the 1952–89 sample. The highest paid quintile averaged about 58 percent of the income gains, while the middle-paid and second highest paid quintiles averaged almost 11 percent and 27 percent of the gains, respectively. The evidence, therefore, suggests that what happened during the 1977–89 period is not that different from what happened during the postwar period.

Note that the standard deviations are substantially different across the five quintiles. As Table 2 shows, the standard deviation is 5.2 percentage points for the lowest paid quintile. There is much greater variability in the highest paid (44.7), the second highest paid (25.0), and the second lowest paid (31.8) quintiles. Thus, the standard deviation for the lowest paid quintile is only about one-third the size of the standard deviation for the other quintiles. This evidence suggests that the lowest paid quintile receives a fairly steady proportion of the income gains across time, especially when compared with the proportions received by the other four quintiles.

In short, the IRS data suggest that the top-paid quintile did account for most of the gains in factor income during the period 1977–89. However, a substantially smaller proportion went to the top-paid quintile than was reported in Krugman's study of after-tax (but before transfer) incomes. The proportion of income gains accruing to the top-paid group increased during the latter half of the 1980s. However, the tendency toward increasing income inequality began in the 1950s, and the 1977–89 period appears to be well within the variability observed historically.

Table 2
Summary Statistics of the Proportion of Real Income Gains for each Quintile, 1952–89 (CEA-based method)

Quintile	Mean	Standard Deviation
1	2.9	5.2
2	1.5	31.8
3	10.9	13.8
4	26.7	25.0
5	58.0	44.7

The Theil index. At a more aggregate level, one can measure income inequality with the Theil entropy index:

$$(3) \qquad T(y,n) = (1/n)\sum_{j=1}^{n}(y_j/\mu)\ \ln(y_j/\mu),$$

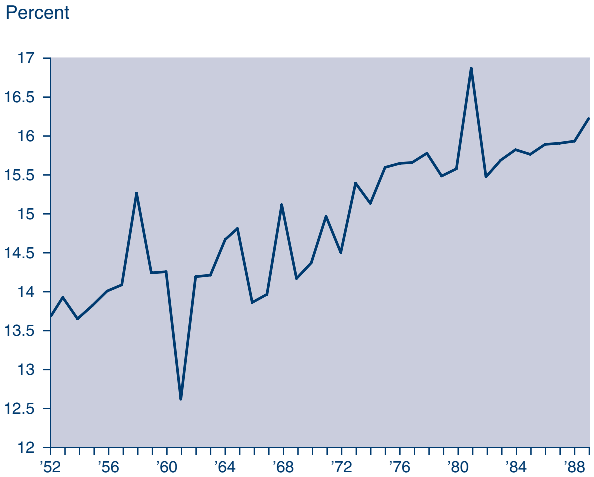
where n is the number of equal-sized population groups, y denotes income for population group j, and μ is average income (that is, $\mu = \sum_{j=1}^{n} y_j/n$).[6]

The Theil index is defined so that it takes on values greater than or equal to zero and increases as income inequality increases. To illustrate this point, consider the limiting case in which income is evenly divided among all people in the economy. With $y_j/\mu = 1$ for $j = 1, 2, ..., n$, then $\ln(y_j/\mu) = 0$ in equation 3 and the Theil index equals zero. Another attractive feature of the Theil index is that a transfer of income from a high income person to a low income person will cause the Theil index to fall.

Figure 1 plots estimates of the Theil entropy measure over the period 1952–89. The plot indicates an upward trend in the Theil index since 1952.

[6] *See Anthony Shorrocks (1980), John Bishop, John Formby, and Paul Thistle (1989), and Daniel J. Slottje (1989) for the set of properties that an income inequality measure possesses. This version of the Theil index differs from the population-weighted version used by Keith R. Phillips (1992).*

Figure 1
Theil Index for Adjusted Gross Income, 1952–89



SOURCE OF PRIMARY DATA:
U.S. Internal Revenue Service, *Statistics of Income*.

Regressing the Theil index against time, one finds more concrete support for the notion that the Theil index has been increasing. The regression coefficient on time suggests that the Theil index has, on average, increased 0.4 percentage points each year. Figure 1 also shows a decline in volatility in the Theil index. Beginning around 1980, the Theil index appears to follow a less variable path compared with the swings observed in the pre-1980 sample. Thus, the two apparent inferences drawn from plotting the Theil index are that income inequality has been increasing over the past forty years and that volatility in the income inequality measure has decreased somewhat during the past ten years.[7]

Our conclusions about the time path of income inequality differ somewhat from other work in the field. For example, Barry Bluestone and Bennett Harrison (1988) find that the percentage of workers falling into the low-wage stratum follows a U-turn: the statistic falls in the 1950s and 1960s, reaches its trough in 1970, fluctuates in a narrow band around the trough value, and then increases beginning in 1979. The results in Figure 1 suggest that the trend toward greater income inequality may have started even earlier than Bluestone and Harrison identify. As such, our findings support those of Peter Henle and Paul Ryscavage (1980), who find that inequality in wages has been increasing since the late 1950s. A note of caution in comparing the results: we are using adjusted gross income from the IRS, while Bluestone and Harrison and Henle and Ryscavage are using wage data from the Current Population Survey. Hence, the results are not directly comparable; our analysis does not overturn Bluestone and Harrison's.

There is evidence on income inequality for broader measures of income. Using the Current Population Survey, Daniel J. Slottje (1989) calculates the Theil index over the period 1947–84. The income measure used here is before taxes but after transfers. According to Slottje, the Theil index falls to its lowest value in the late 1960s, rising thereafter for the remainder of the sample period. The implication is that a U-turn is present in comprehensive income measure employed in the Current Population Survey data. Again, the income concepts used in Slottje are not the same as ours because they include transfers and the data sets are different. Our finding, however, raises a question about when (and if) the U-turn occurred in a broader income measure, when one looks at income before taxes and transfers. It also suggests that the U-turn in the inequality of after-transfer income may reflect changes in the distribution of transfer policy rather than factor incomes.

What determines the distribution of income?

In this section, we turn from describing what happened to income inequality to examining why the distribution of income changed. The factors affecting income inequality that we consider fall into three categories: demographics, economic conditions, and fiscal policy. Our demographic data are

[7] *The Theil index invariably measures changes in the income distribution differently from other inequality measures because it weights transfers differently. To check the robustness of our inferences, we also calculated another aggregate measure of income inequality—the Gini coefficient. The results are not materially different whether one uses the Gini coefficient or the Theil entropy measure.*

the age and educational composition of the potential labor force, and the female share of the labor force.[8] Following Alan S. Blinder and Howard Y. Esaki (1978), we include some measure of business-cycle conditions and the inflation rate as variables that might explain variation in income inequality. Finally, we use the maximum marginal personal income tax rate and real per capita transfer payments to individuals (defined as the sum of federal, state, and local transfers) as the policy variables. The Theil index is our measure of income inequality.

To determine whether these characteristics can be used to predict changes in the Theil index over time, we estimate a VAR system of eight equations—one for each of our eight variables—using ordinary least squares regression. Each equation estimates contemporaneous values of the variable as a function of two lagged values of itself and two lagged values each of the other seven variables. All the variables except the growth rates for prices and gross domestic product are expressed as first differences (the change in value between period t and period t–1), and the variables that are bounded by zero and one (such as the percentage of women in the labor force) are logistically transformed.[9]

Table 3
Tests of Exclusions Restrictions in Theil Index Equation

Variable	*F* statistic	*p* value
AGE	.15	.86
ED. ATTAINMENT	1.61	.23
FEM SHARE	1.02	.38
TAX	1.67	.22
GDP	.87	.44
INFL	.78	.47
TRANS	.61	.55

Table 3 reports the results of exclusion restrictions for each variable in the equation in which the Theil index is the dependent variable. The tested hypothesis is that the coefficients on lagged values of the variable also are jointly equal to zero. The interpretation of the test results, then, is whether changes in the variable help to predict changes in the Theil index. The *F* statistics are small in each case, which is consistent with the notion that none of these variables, except past values of the Theil index itself, helps to predict changes in the Theil index. However, all of the variables together explain almost 50 percent of the variation in the Theil index. Correlations among some of the variables may be introducing multi-collinearity.[10]

Although a variable may have insignificant explanatory power in the VAR regressions, its compound influence over time may still be considerable. Table 4 reports how much of the two-, five-, and ten-step-ahead forecast error variances result from innovations in the variables.[11] The evidence strongly suggests that innovations in the Theil index account for most of the forecast error variance. Indeed, 54 percent of the ten-step-ahead forecast error variance results from innovations in the Theil. The analysis suggests that the Theil index displays

[8] *The age composition of the labor force is defined as the ratio of people between 16 and 25 to those between 16 and 65. The education composition of the potential labor force is measured by the percentage of the population over age 25 that has graduated from college.*

[9] *David Hendry and Jean–Francois Richard (1982) argue that a logistic transformation should be applied to any dependent variable defined over the [0,1] interval. Formally, the transformation redefines the variable as follows:* $ln(x_t/1–x_t)$.

[10] *The issue with multicollinearity is that close correlation between the explanatory variables will result in inflated standard errors. The upshot of this is that test statistics are downwardly biased when multicollinearity is present.*

[11] *We apply a Choleski decomposition with the following ordering for the recursive system:* AGE, TAX, ED. ATTAINMENT, THEIL, FEM SHARE, GDP, INFL, *and finally* TRANS. *See Christopher Sims (1980) for a more complete discussion of the Choleski decomposition applied to VARs.*

Table 4
Proportion of Forecast Errors for the Theil Index*

	Innovation to							
Step-Ahead	***AGE***	***FEM SHARE***	***ED. ATTAINMENT***	**Theil**	***TAX***	***GDP***	***INFL***	***TRANS***
Two	5.6	1.1	5.7	72.0	10.3	0.7	1.6	3.0
Five	5.1	3.0	16.3	54.9	8.9	4.3	2.0	5.6
Ten	5.5	3.4	16.3	53.6	8.9	4.3	2.2	5.8

*Proportions may not sum to 100 due to rounding error.

lots of persistence, accounting for movements in the Theil over time.[12]

The other factors explain the rest of the ten-step-ahead forecast error variance. Together, the age composition and educational attainment of the population, and the female share of the labor force, account for slightly more than 25 percent. Fiscal policy variables account for about 15 percent of the forecast error variance, while inflation and output growth account for 2.2 and 4.3 percent, respectively. The evidence, therefore, suggests that 85 percent of the variation in income inequality arises from factors outside direct policy control.

Sheldon Danziger, Robert Haveman, and Peter Gottschalk (1981) also examine the role that the U.S. transfer payment system has on income inequality. They use both welfare payments and Social Security as their definition of the transfer payment system. Comparing the distribution of total factor income with the distribution of total factor income plus transfer payments (but before taxes), Danziger, Haveman, and Gottschalk conclude that income inequality (as measured by the Gini coefficient) is 19 percent lower after transfers than it is before them. The Danziger, Haveman, and Gottschalk comparison excludes the complex general equilibrium effects that transfer payments have.

Using a narrower definition of transfer payments and assuming that a recursive model represents the structure, we find that transfer payments have much less explanatory power. In our analysis, transfer payments explain less than 6 percent of the ten-step-ahead forecast error. In addition, the impulse response functions suggest that increases in transfer payments are associated with *increases* in factor income inequality.

Summary and conclusions

In this article, we examine developments in the income distribution over almost four decades and the relative contributions of demographic, policy, and economic conditions toward explaining these movements. Our analysis indicates that recent developments in income inequality and the distribution of gains from income growth are not much different from historical norms. We find that the top-earning 20 percent of the population reaped a disproportionate share of the income gains during the 1980s. However, we find that the top income

[12] *Steven N. Durlauf (1991) examines the evolution of income inequality and finds that persistent income inequality can develop even with identical starting conditions. Education and neighborhood effects reinforce one another to stratify the economy, imparting substantial persistence in income inequality.*

group has been receiving similarly large shares of the income gains for the past forty years. Further, using the Theil index to measure income inequality, we find that income inequality increased over the period 1952–89.

In addition, we look at the relationship between various factors associated with changes in the distribution of income. In particular, we ask which factors explain movements in income inequality over time. We investigate demographic factors, economic conditions, and policy variables. The evidence reported in this article suggests that there is a great deal of persistence in income inequality. Most of the forecast error variance in the income inequality measure is explained by innovations in the inequality measure itself. Fiscal policy actions—measured as the maximum marginal tax rate and per capita transfers—account for only about 15 percent of the variation in the forecast errors.

Overall, the evidence presented in this article focuses on developments in income inequality *over time*. In doing so, the contribution is largely in describing how income inequality has evolved, excluding some factors that are widely believed to affect income inequality. The aim of future research is to formulate theories about what determines income inequality.

Appendix

Data Definitions

Variable	Source
GDP	Board of Governors of the Federal Reserve System FAME dataset
CPI	Bureau of Labor Statistics
AGE	U.S. Department of Commerce, Bureau of Census; population between 15 and 25 divided by the population over 16, less those over 65 (Citibase data)
FEM SHARE	U.S. Department of Commerce, Bureau of Census; total number of employed women divided by the total labor force (Citibase data)
TRANS	National income and product accounts, transfer payments to individuals paid by federal, state, and local governments (Citibase data). This variable is deflated with the fixed-weight GDP deflator.
TAX	Internal Revenue Service, *Statistics of Income*, various issues.
ED. ATTAINMENT	Current Population Survey, series P-60; percent of population over 25 with four or more years of college.

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The Costs and Benefits of Fixed Dollar Exchange Rates in Latin America

The major Latin American countries have embarked on broad-based economic reform programs to raise economic efficiency, promote investment, and accelerate output growth.[1] To achieve these goals, these countries are attempting to foster economic stability, and in some cases, the tool they are using is an exchange rate fixed to the U.S. dollar.

When establishing economic stability is an important goal, the purposes fixed exchange rates can serve go beyond what is immediately obvious. Not only do fixed exchange rates stabilize the domestic prices at which exporters can sell and importers can import, a fixed exchange rate regime also has important implications for domestic monetary and price stability.

Implications for domestic monetary and price stability involve international differences in inflation rates. For a country's exchange rate to remain fixed, a country's inflation rate and the inflation rates of its trading partners must be the same. If the country's inflation rate persistently exceeds those of its trading partners, the country's citizens will buy foreign products. After all, prices of domestic products in a country with a fixed exchange rate will eventually rise above prices of foreign imports. The same price phenomenon will hurt the country's exports. Money growth is a primary cause of inflation; therefore, a country that fixes the price of its currency relative to some other country's currency implicitly must follow that country's monetary policy.

As a result, a fixed exchange rate can signal to investors a government's intent to follow a stable monetary policy. If the government prints money to cover budget deficits or to postpone unpleasant adjustments to adverse external shocks, investors will notice quickly. The policy will be obvious as the country's inflation outstrips that of its trading partners, the demand for foreign (domestic) currencies rises (falls), and the country's foreign exchange reserves disappear.

Thus, a persistently fixed exchange rate lends credibility to a government's commitment to a stable monetary policy. Such credibility is important. Even if a government is firmly committed to a stable monetary policy, private-sector behavior can nullify the expected benefits of such a policy if the private sector does not believe the policy will last.

But there is a problem with using a fixed exchange rate to signal such credibility. A government's commitment to fixing the exchange rate may, itself, be incredible. Almost any government faces temptations to renege on both exchange rate and monetary policy targets. If the returns from unexpectedly devaluing are high, a fixed exchange rate is not very credible. For a government to make a fixed exchange rate credible, it must demonstrate or create circumstances that make the cost of devaluation high.

This article assays the policy and economic characteristics that could make fixed exchange rate regimes credible and, therefore, make credible a government's commitment to monetary stability.

We would like to thank without implicating Bill Gruben, Joe Haslag, Evan Koenig, and Carlos Végh for their helpful comments. Any errors and omissions, as well as any opinions, propositions, or conclusions, are exclusively our own.

[1] *Examples include Argentina, Bolivia, Chile, Colombia, Mexico, Venezuela, and to a lesser extent, Brazil.*

The factors that result in such credibility—or in a lack of credibility—are complicated because the benefits of reneging are not always what one might expect. The obvious gain that would make reneging likely—that a devaluation could result in economic growth—is not the only benefit. Even when a devaluation clearly will not result in rising income and growing government revenues, some countries still devalue, and we outline what governments get when they do. Some governments, for example, are simply trying to accumulate foreign exchange so that they can defend their currencies in the future.

In the next section of this article, we use a simple game to examine the dilemma Latin American governments face when choosing whether to fix their exchange rate. We then consider the importance of the effects of devaluation on output growth and discuss key economic relationships that determine the effects of changes in exchange rates.

The credibility of exchange rate and anti-inflation policy

Although we know that official promises to follow a monetary or exchange rate rule are not credible if reneging is not costly, identifying costliness is not always easy.[2] Attempts to address related issues in the area of domestic monetary policy appear in the "rules versus discretion" literature, which considers alternative economic scenarios and the likelihood of various government reactions to them.

Although the issue of rules versus discretion was developed without a focus on exchange rates (Kydland and Prescott 1977 and Barro and Gordon 1983a and 1983b), numerous authors have recently applied this framework to questions about appropriate exchange rate regimes.[3] In many such applications, surprise devaluations are assumed to increase inflation, a result countries generally do not want, and to increase output, a result countries almost universally want. Overall, when these are the likely outcomes of devaluation, a fixed exchange rate is not very credible unless authorities despise inflation so much that they will avoid devaluation at any cost.

The story became more complicated, however, when Edwards (1989) and Faini and de Melo (1990) showed that developing countries often suffer a fall in output growth from surprise devaluations, a scenario first outlined for Latin America by Diaz–Alejandro (1963).[4] Thus, while the earlier literature tells us something about when fixed exchange rates may be incredible, Edwards (1989) and Faini and de Melo (1990) offer information about when fixed exchange rates may prove credible.

Contractionary devaluations result from a variety of sources.[5] One involves the economic structure generated in Latin America by its post-World War II protectionist policies. Protected domestic industries relied heavily on imported inputs, especially capital goods. In addition, protection of industry rendered investment in agriculture and other exporting sectors unprofitable. Therefore, exporting sectors had a tendency to stagnate. In these circumstances, devaluation can have perverse effects on output. Devaluation increases the price of investment goods, which leads to a collapse of investment. Because Latin American countries generally did not have capital

[2] This article tries to treat credibility as endogenous, that is, the credibility of a fixed exchange rate is dependent on the costs of devaluing to society. We ignore the political process. Credibility will also depend upon how effectively the political process punishes policymakers and governments for bad economic outcomes.

[3] These include Kamin (1988), Giovannini (1990 and 1992), Agénor (1991), Devarajan and Rodrik (1991), and Dornbusch and Fischer (1991).

[4] See Diaz–Alejandro (1963) and Cooper (1971) for early analyses of contractionary devaluation. More recent studies include Krugman and Taylor (1978), Lizondo and Montiel (1986), Edwards (1989), Faini and de Melo (1990), and Cooper (1992).

[5] At best, devaluations tend to be neutral in Edwards' analysis (Edwards 1989, chapter 8). Output tends to fall before the devaluation, usually as a result of a deterioration in the terms of trade. When devaluation finally occurs, output tends to improve but not necessarily to a level greater than or equal to output before the terms of trade shock. Kamin (1988) also comes to similar conclusions. McLeod and Welch (1992) find a nonlinear relationship, where small (surprise) real devaluations increase output growth and large devaluations decrease it in Argentina, Brazil, Colombia, and Venezuela, while any devaluation decreases output in Chile and Mexico.

goods-producing sectors, and import substitution policies left exporting sectors weak and inflexible, the devaluation would not generate an expansion of domestic capital goods production nor an offsetting increase in export revenue to buy imported capital goods. Economic stagnation would result, if only temporarily.

These peculiar circumstances mean that, unlike some countries in other parts of the world, Latin American nations have avoided using exchange rate policies to improve output growth. If anything, Latin American governments have resisted devaluation, even in the face of severe overvaluation. (See the Appendix, "Inflation and Exchange Rates in Latin America," for a description of exchange rate policies in Latin America). Indeed, in addition to the economic costs of higher inflation and, in some cases, lower output growth, Latin American devaluations often entail political costs.[6]

Despite the negative economic and political effects, Latin American countries have devalued. Even in cases when fixed exchange rates have been an explicit objective of policy, such credible fixed exchange rates have not been part of Latin America's experience over the past thirty years.[7]

These difficulties not only raise questions about the process of devaluation and exchange rate manipulation, but also about the process of credibility formation. Where devaluations are known to weaken output growth and increase inflation, special care must be taken in considering the nature of policy trade-offs. Why, after all, would governments want what many of their citizens would regard as prejudicial?

One explanation that Kydland and Prescott (1977), Barro and Gordon (1983a and 1983b), and Dornbusch and Fischer (1991) suggest is that a government might want to increase inflation tax revenue through surprise inflation. In the case of Latin America, this argument is not credible. Dornbusch and Fischer find that the pure public-finance motive for inflation explains little in the context of Latin American countries experiencing moderate inflation, such as Brazil in the 1960s and Chile, Colombia, and Mexico. Dornbusch, Sturzenegger, and Wolf (1990) show that the public-finance motive only marginally explains the acceleration of inflation in the high-inflation Latin American countries—Argentina, Bolivia and Brazil in the 1980s, and Peru.

Another argument, however, is credible. Latin American countries have had to use nominal exchange rate surprises to generate substantial balance of payments (trade) surpluses to service their foreign debts.[8] Other factors suggest that the generation of foreign exchange reserves is an important consideration when a policymaker contemplates the consequences of devaluation in the context of debt-servicing difficulties. Eaton and Gersovitz (1980) point out that foreign exchange reserves take on special importance in providing liquidity services for export and import transactions, especially if the country faces credit limits in international markets.[9] Similarly, van Wijnbergen (1990) emphasizes the insurance value of foreign exchange reserves when trade-contingent debt instruments do not exist in international capital markets for these countries. For these reasons, we introduce reserve growth as an objective of policy.

A model of the exchange rate credibility problem in Latin America

We present a game to simulate the special exchange rate credibility problems that a Latin American policymaker might face. We then address some of the issues the game raises for Latin America.

[6] See Edwards (1989) and Edwards and Montiel (1989).

[7] The most noteworthy cases are those of the Southern Cone countries of Argentina, Chile, and Uruguay in the late 1970s and early 1980s. These countries combined trade and capital account liberalization with exchange rate pegging to bring down inflation. All the programs had collapsed by 1982, due to severe terms of trade shocks, interest rate shocks, and the fact that the fixed exchange rate did not necessarily force the adoption of fiscal, monetary, and financial policies consistent with fixed exchange rates.

[8] See Faini and de Melo (1990) for a discussion of the central role of real exchange rate depreciation in the stabilization policies in the debt crisis of the 1980s.

[9] To support this claim, Eaton and Gersovitz (1980) provide evidence that debt and foreign reserves were asset substitutes for less-developed countries in the 1970s. As these countries borrowed internationally, reserve holdings fell. Once voluntary lending dried up in the early 1980s, the demand for reserves increased accordingly.

Consistent with our discussion in the previous section, the model introduces a balance of payments surplus target, the growth in foreign exchange reserves, in the objective function so that the trade-off in policy objectives is between inflation and the growth in foreign reserves (a balance of payments surplus).[10]

To present exchange rate credibility problems as simply as possible, we assume that unanticipated depreciations of the currency improve the balance of payments or, equivalently, increase foreign exchange reserve growth. Equation 1 characterizes the forces that may affect growth or the decline in reserves as[11]

$$(1) \qquad \dot{R} = \alpha\left[e - E(e)\right] + \omega,$$

where R is the level of foreign reserves held by the central bank, a (˙) signifies a rate of change over time, e is the rate of depreciations of the nominal exchange rate, E is the expectations operator, and ω is an external shock that is equal, on average, to zero. This shock could represent unexpected changes in the terms of trade, defined as the (foreign) price of exports over the (foreign) price of imports, or changes in foreign interest rates over some expected value. All variables are expressed in natural logarithms. α measures the (temporary) increase in reserve growth due to a surprise devaluation per unit of (infinitesimal) time; exchange rate depreciation can only increase reserve growth if it is a surprise. The short-term nature of the analysis assumes that the *quantum* of exports and imports does not quickly adjust to terms of trade or interest rate shocks and, therefore, any such shock translates completely into reserve changes. The expected external shock is equal to zero.

All goods in this simple model are traded. Inflation is determined purely by expectations of inflation. We can assume that goods price arbitrage (purchasing power parity) holds and that individuals have rational expectations. The private sector, however, sets its inflation expectations before the government decides where to set changes in the exchange rate and, thus, inflation.[12] Individuals in the private sector lose mainly through incorrect predictions of the exchange rate, although, on average, individuals correctly predict inflation. Thus, we assume that individuals in the private sector maximize the following utility function:

$$(2) \qquad U_p = -\frac{1}{2}\left[\pi - E(\pi)\right]^2.$$

Accordingly, they try to forecast the inflation rate as accurately as possible to minimize their losses from miscalculation. Hence, individuals will act so that

$$(3) \qquad \pi = E(\pi) = E(e) + E(\pi^*) = E(e),$$

where is domestic inflation and * is the foreign inflation rate, set equal to zero for simplicity. In the long run, expected values of all variables will equal their actual values. But we assume the government can react to shocks to the system more quickly than the public because prices (and wages) are set before the shock is revealed. Therefore, the government can temporarily cause departures from purchasing power parity by unexpectedly changing the nominal exchange rate.

The government minimizes a quadratic loss function, which penalizes any inflation rate (positive or negative) not equal to zero and deviations from a target growth rate of foreign reserves, $\dot{R}$,

$$(4) \qquad max\ U = -\left[\frac{\pi^2}{2} + \frac{\beta}{2}(\dot{R} - \dot{R}^*)^2\right],$$

where $-\beta(\dot{R} - \dot{R}^*)$ measures the marginal utility of a deviation of an increase in reserves. If reserve

[10] Sachs (1985) analyzes the U.S. trade deficit in a similar way. The relationship could also include the capital account. We exclude the capital account for ease of exposition.

[11] In a more complete model, the country produces nontraded goods as well as traded goods, and the direct effects of exchange rate changes are concentrated on the traded-goods sector.

[12] This timing could be generated by the overlapping contracts framework of Gray (1976) and Fischer (1977), or it could be due to the government having an informational advantage, as in Canzoneri (1985). It is unlikely, however, that the government will observe terms of trade shocks before the private sector. The inertia in prices due to contracting gives the government room to use surprise devaluation to increase reserve growth, as discussed below.

Table 1
Payoff Matrix for the Government and Private Sectors Under Fixed and Flexible Exchange Rates*

		Private sector expects	
		$e = 0$ (fixed)	$e > 0$ (flexible)
Government chooses	$e = 0$ (fixed)	$-\frac{\beta}{2}\left[\dot{R}^{*2} + \sigma_\omega^2\right], 0$	$-\frac{1}{2}\left[[\beta^2\alpha^2 + \beta(1+\alpha^2\beta)^2]\dot{R}^{*2} + \beta\sigma_\omega^2\right], -\frac{\beta^2\alpha^2}{2}\dot{R}^{*2}$
	$e > 0$ (flexible)	$0, -\frac{\beta^2\alpha^2}{2}\dot{R}^{*2}$	$-\frac{\beta}{2}\left[(1+\beta\alpha^2)\dot{R}^{*2} + \left(\frac{1+\beta\alpha^4}{(1+\beta\alpha^2)^2}\right)\sigma_\omega^2\right], 0,$

where e = rate of devaluation, E = expectations operator, $\sigma_\omega^2 = E(\omega^2)$ is the variance of external shocks.

*The first value in the ordered pair represents the government's payoff and the second the private sector's payoff.

growth is below its target rate, the marginal utility is positive, and if reserve growth is above its target rate, the marginal utility is negative.

The game proceeds in two simple stages. The private sector sets its devaluation and, thus, inflation expectations, and enters into contracts based on these price expectations. The government then sets its rate of devaluation. Table 1 presents the payoff matrix for the government and private sectors, and Table 2 presents the analytical solutions to the game under two conditions: credible precommitment, in which the government irrevocably fixes the exchange rate, and flexible exchange rates, which are equivalent to a noncredible fixed exchange rate.[13]

If the government cannot be forced to honor a commitment to a fixed exchange rate, the dominant strategy for the government is a flexible rate.[14] Knowing this, the private sector will always expect exchange rates to be flexible. Accordingly, the equilibrium will correspond to the lower right quadrant of Table 1. On the other hand, suppose the government precommits to a fixed exchange rate and, through legislation or membership in a currency bloc, the commitment is credible. Then one possible equilibrium of the game could correspond to the upper left quadrant of Table 1. The other equilibrium is still the flexible exchange rate equilibrium in the lower right quadrant of the table.

If "commitment technologies" exist that can force the government to establish an irrevocably fixed exchange rate, which regime will the government pick? This game does not yield an unambiguous choice.[15] Both regimes have different implications for inflation and reserve growth. We now investigate these different outcomes.

[13] We assume that the targeted growth rate of reserves $\dot{R}^*$ and shocks ω are independent. In Table 2, the subscript r denotes rule, while the subscript d denotes discretion.

[14] To see this, compare the total payoff to the government when e = 0 and when e > 0. The payoff when e > 0 is unambiguously larger than when e = 0.

[15] If we subtract the expression for $E(U)_d$ from $E(U)_r$ in Table 2, we find

$$E(U)_r - E(U)_d = \frac{\beta}{2}\left[\beta\alpha^2\dot{R}^{*2} - \left(\frac{\beta\alpha^2[(2+\alpha^2(\beta-1)]}{(1+\beta\alpha^2)^2}\right)\sigma_\omega^2\right].$$

The sign of this expression depends on the government's marginal rate of substitution between inflation and international reserves, ß (> 0), and the effects of devaluation on reserve growth as well as the target level of reserve growth and the variance of external shocks. The partial derivatives of are

$$\frac{\partial[E(U)_r - E(U)_d]}{\partial\dot{R}^*} > 0, \frac{\partial[E(U)_r - E(U)_d]}{\partial\sigma_\omega^2} < 0, \text{ and}$$

$$\frac{\partial[E(U)_r - E(U)_d]}{\partial\beta} \gtrless 0, \frac{\partial[E(U)_r - R - E(U)_d]}{\partial\alpha} \gtrless 0.$$

Table 2
Equilibrium Inflation, Reserve Growth, and Expected Government Utility
Under Fixed and Flexible Exchange Rates

Credible precommitment to a fixed exchange rate	Flexible exchange rate (No credible commitment)
$\pi_r = R(e_r) = e_r = 0$	$e_d = \beta\alpha\dot{R}^* - \frac{\beta\alpha^2}{1+\beta\alpha^2}\omega$
	$\pi_d = E(e) = \beta\alpha\dot{R}^*$
$\dot{R}_r = \omega$	$\dot{R}_d = \frac{1}{1+\beta\alpha^2}\omega$
$E(U)_r = -\frac{\beta}{2}\left[\dot{R}^{*2} + \sigma_\omega^2\right]$	$E(U)_d = -\frac{\beta}{2}\left[(1+\beta\alpha^2)\dot{R}^{*2} + \left(\frac{1+\beta\alpha^4}{(1+\beta\alpha^2)^2}\right)\sigma_\omega^2\right],$

where π = inflation rate, e = rate of devaluation, τ = terms of trade, E = expectations operator, U = government's utility, $\sigma_\omega^2 = E(\omega^2)$ is the variance of external shocks.

Inflation and reserve growth with a credibly fixed exchange rate

Suppose the government can credibly commit to the exchange rate rule that sets $e_r=0$. From Table 2, the equilibrium exchange rate depreciation and inflation, $_r$, will be zero, and expected reserve growth will be zero, while actual reserve fluctuations will completely accommodate unexpected terms of trade or foreign interest rate fluctuations.

Notice that the larger the government's desired reserve accumulation, the more government welfare falls. A government that credibly commits (forever) to a fixed exchange rate will be continuously frustrated in increasing its stock of reserves if this divergence persists. A higher variance of the shock term also decreases welfare.

Such a policy, however, is not *time consistent.* The government can improve on this outcome temporarily by announcing a fixed exchange rate and then devaluing.[16] Rational individuals will recognize the government's incentive to renege on its exchange rate stance and will set expectations

[16] *A government that convinces the public it will keep exchange rates fixed can improve welfare by cheating in the current period. For an explanation, note that expected inflation in the current period will be zero. Maximizing equation 3 with respect to* e *yields*

$$e = \frac{1}{\alpha}\left[\dot{R}^* - \omega\right].$$

Expected and actual reserve accumulation under cheating will be

$$\dot{R}_C = \dot{R}^*.$$

Expected utility will be equal to zero, which is clearly an improvement over credibly fixed exchange rates and, as we shall see, flexible exchange rates since both generate negative expected utility. Such a situation, however, is not consistent. The public would recognize the government's large incentive to cheat, and the equilibrium outcome reverts to the discretionary flexible exchange rate case above. Also, once the public realizes that the government has cheated, cheating can never be used again. Addressing sustainable equilibria over time and the relationship between credibility and reputation are beyond the scope of this article. For discussion and analysis, see Barro and Gordon (1983a and 1983b), Rogoff (1985 and 1987), Canzoneri (1985), Cukierman (1986), Agénor (1991), Obstfeld (1991), and Fischer (1990).

and prices so that the marginal cost of devaluing will equal the marginal benefit to the government.

Inflation and reserve growth under flexible exchange rates

Assuming that optimal reserve growth is positive, a government that can gain foreign exchange reserves by devaluing will face the well-known inflationary bias of policies that lack credibility. The private sector, knowing that the government has an incentive to devalue to raise reserve growth, expects a devaluation. The private sector will set prices accordingly, generating positive inflation regardless of the government's policy stance. Even if the government does not devalue when the private sector expects a devaluation, the economy will suffer from high inflation. The positive rate of inflation, combined with the fixed exchange rate, will cause a surge in imports and a fall in exports, culminating in a loss in foreign reserves. Following a no-devaluation policy without credibility is doubly costly to the government: both inflation and a balance of payments crisis will result. The government will devalue, one way or another.

Formally, one can see this result by taking expectations on the equilibrium solution for devaluation and inflation under flexible exchange rates and noting that the ex ante expected shock is zero. Then from Table 2, expected inflation and devaluation is

$$E(e_d) = E(\pi_d) = \beta\alpha\dot{R}^*. \tag{5}$$

It should be noted that the adjustment programs undertaken in Latin America included devaluation for the explicit purpose of increasing foreign exchange reserves. The analysis suggests that this policy will be inherently inflationary. On the other hand, governments may intend to fix exchange rates forever but cannot resist the temptation to devalue to increase reserves. Economic agents, recognizing this, act to protect themselves from the devaluation by increasing the rate of price increases.

Because the coefficient on the shock term in the solution for reserves under flexible exchange rates in Table 2 is less than the corresponding coefficient under fixed exchange rates, the flexible rate regime can smooth fluctuations in foreign exchange reserves due to unexpected changes in the terms of trade or foreign interest rates. Again, as in the credible fixed exchange rate case, the government cannot affect the expected (average) rate of foreign reserve accumulation in the long run.

To fix or not to fix? The importance of output effects of real devaluation

A comparison of the game's outcomes under the two regimes suggests that a credible commitment to a fixed exchange rate regime eliminates the inflationary bias of the flexible rate regime. Unfortunately, however, this policy increases the economy's susceptibility to external shocks. The choice of regime, therefore, depends on the relative importance a particular government places on each of these two objectives and on how sensitive the balance of payments is to exchange rate surprises. Whether a country fixes its exchange rate depends on how heavily the policymaker values reserves relative to inflation (β), how sensitive reserve growth is to devaluation (α), the variance of external shocks σ^2_ω, and the target reserve growth $\dot{R}^*$.

Before the early 1980s, Latin American countries grew significantly. The terms of trade and international interest rates remained fairly steady. Most Latin American countries maintained fixed exchange rates throughout this period, despite the breakdown of the Bretton Woods system of fixed exchange rates in the early 1970s. The external trauma of the late 1970s and early 1980s brought increases in world interest rates and a major decline in the terms of trade to most of Latin America. The model predicts that an increase in the variance of foreign shocks will increase the desirability of a more flexible rate, and, in fact, many Latin American countries moved toward a more flexible rate.

In the early 1990s, adjustment to these shocks is nearly complete for most Latin American countries, and interest rates and terms of trade are stable. The theory presented above suggests that Latin American countries will tend toward fixed exchange rates in the near future, barring any major disruption of international trade and capital flows.

Costs and benefits of fixed and flexible exchange rate regimes

Most Latin American countries had to devalue their exchange rates dramatically in the 1980s in the face of the balance of payments problems brought about by the debt crisis and adverse terms of trade movements. The adjustment under more flexible exchange rates imposed high economic costs. Most Latin American countries suffered large declines in gross domestic product (GDP) per capita in the first half of the decade. From 1981 to 1985, the cumulative fall in per capita GDP was 12.6 percent in Argentina, 14.9 percent in Chile, 8.7 percent in Mexico, and 14.5 percent in Venezuela.[17] Brazil and Colombia managed to increase GDP per capita by 4.1 percent and 1.9 percent, respectively, after suffering initial contractions between 1981 and 1984. Inflation, as discussed in the Appendix, also accelerated in all six of these countries in the first half of the 1980s. The costs of balance of payments crises were high, especially in those countries where output fell as a consequence of real devaluation.

Latin American terms of trade movements settled down in the 1980s, and international interest rates have declined significantly in recent years. Given these results, we are not surprised to see countries such as Argentina and Mexico returning to fixed exchange rates after more than a decade of adjustment to the debt crisis. Enhanced credibility should greatly improve these countries' macroeconomic performances, barring any new balance of payments crisis. The costs of renewed real devaluation emanating from large government deficits and excessive money growth in the form of high inflation and losses of foreign reserves should temper any moves back to macroeconomic mismanagement. Terms of trade shocks, however, can still undo these fixed exchange rates and, thus, keep the exchange rate regimes from being completely credible. Consequently, some flexibility in the exchange rates in Latin America may be desirable.

Ultimately, however, the credibility of government policy will depend on its prolonged effort to maintain monetary and fiscal discipline for low inflation and, more generally, to keep the reform process on track. Latin American countries seem, at least for the moment, to be headed in this direction.

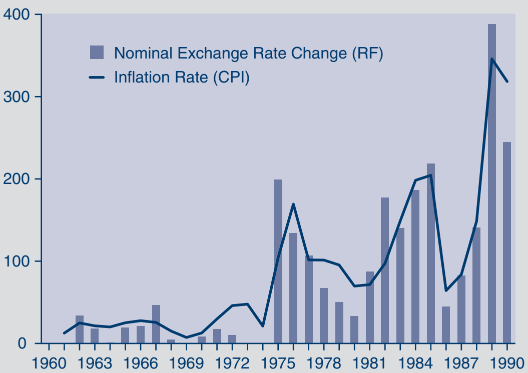
[17] *Inter-American Development Bank (1991).*

Appendix

Inflation and Exchange Rates in Latin America

Figure 1
Trends in Argentina

Inflation rate and nominal exchange rate
(Percent)



Standard deviation of changes in terms of trade
(Percent)

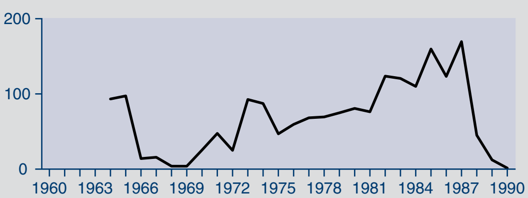
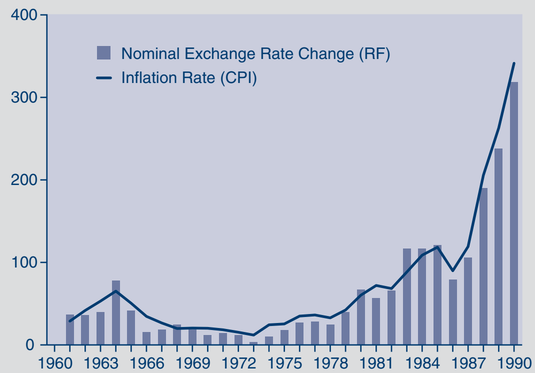
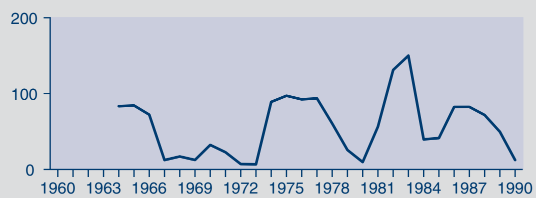


Figure 2
Trends in Brazil

Inflation rate and nominal exchange rate
(Percent)



Standard deviation of changes in terms of trade
(Percent)



For most of the period since World War II, Latin American countries have generally maintained fixed exchange rates with the U.S. dollar. These fixed exchange rate regimes usually collapsed, but Latin American central banks most often returned to a fixed exchange rate following devaluation.

Argentina, Brazil, Chile, Colombia, Mexico, and Venezuela tried to maintain parities fixed to the dollar until the late 1960s. Colombia, Mexico, and Venezuela were more successful than Argentina, Brazil, and Chile, even in the face of high inflation and overvaluation. The ultimate devaluations, however, were followed by a return, if only temporarily, to pegged exchange rates.

Figures 1 through 6 show trends for these six countries.[1,2] The figures also show

[1] Descriptions of exchange rate policies in each of these countries can be found in Cavallo and Cottani (1991) for Argentina, Coes (1991) for Brazil, Edwards and Edwards (1987) and de la Cuadra and Hachette (1991) for Chile, Edwards (1986) and Garcia Garcia (1991) for Colombia, Ortiz (1991), and McLeod and Welch (1991) for Mexico.

[2] Inflation (upper charts in Figures 1 through 6) is measured as the logarithmic percentage change in the consumer price index (CPI), with data from the International Monetary Fund (IMF). Nominal exchange rates are taken from the IMF's RF series. The lower charts in Figures 1 through 6 show (three-year moving average) standard deviations of logarithmic changes in the terms of trade, with data from the World Bank (1991).

(Continued on the next page)

Inflation and Exchange Rates in Latin America—*Continued*

Figure 3
Trends in Chile

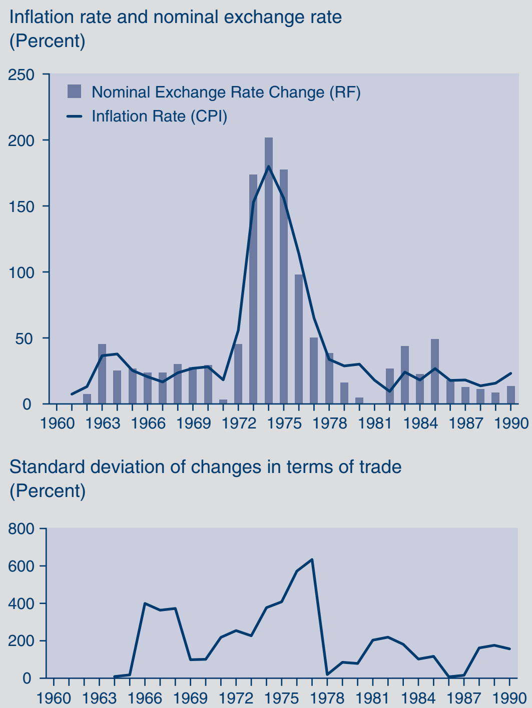
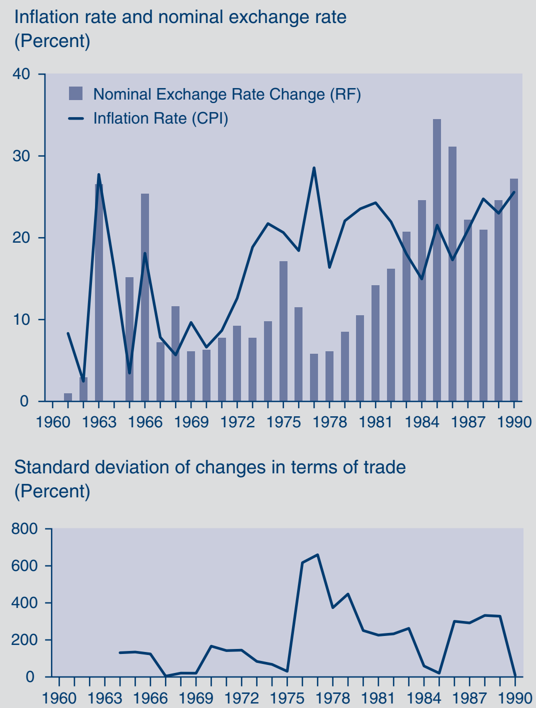


Figure 4
Trends in Colombia



that inflation rates and changes in exchange rates tend to be highest when the variability of the terms of trade is high for most of these countries.

The figures reflect Brazil's and Colombia's moves to crawling pegs in the late 1960s. Argentina briefly used a crawling peg from 1964 to 1966 and during 1971. Argentina, however, reverted to imposing fixed exchange rate regimes throughout the postwar period. Chile, on the other hand, seems to have implicitly used a crawling peg arrangement until fixing in 1971, although officially the country had fixed rates.

Economic disturbances in the early 1970s, especially OPEC's 1973 oil embargo, caused a new set of devaluations in Argentina, Chile, Mexico, and Venezuela, while Brazil and Colombia had to accelerate their exchange rates' devaluation rate. After these large devaluations, Argentina and Chile followed crawling pegs until 1978. In 1979, Argentina and Chile, along with Uruguay, tried to use their exchange rates to lower inflation by slowly decreasing the rate of crawl to zero.

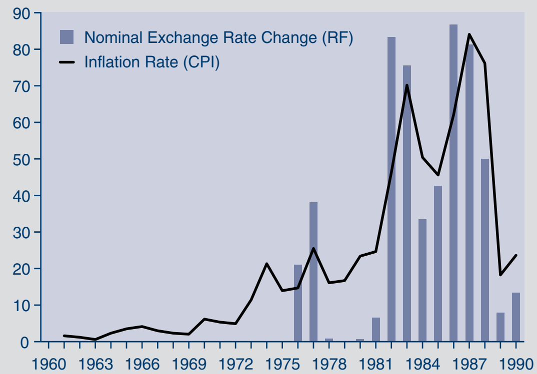
Rising international interest rates in the late 1970s and severe terms of trade shocks in the early 1980s caused all six countries to abandon fixed rate regimes and adopt more flexible exchange rates. They continuously devalued their currencies. Each of these countries had to increase its exports and decrease

(Continued on the next page)

Inflation and Exchange Rates in Latin America—*Concluded*

Figure 5
Trends in Mexico

Inflation rate and nominal exchange rate
(Percent)



Standard deviation of changes in terms of trade
(Percent)

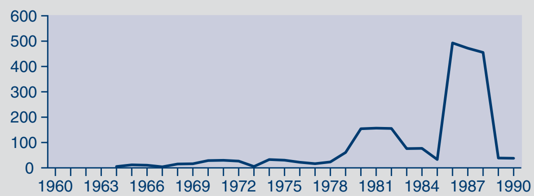
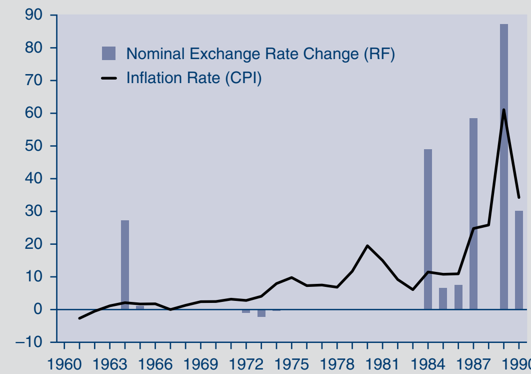
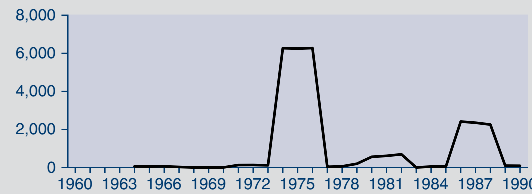


Figure 6
Trends in Venezuela

Inflation rate and nominal exchange rate
(Percent)



Standard deviation of changes in terms of trade
(Percent)



its imports to service the large foreign debts it had accumulated in the 1970s. During this period of flexible exchange rates in the mid-1980s, inflation accelerated to unprecedented levels in all Latin America, with the exception of Chile.

Flexible exchange rate regimes lasted until the late 1980s, except for brief periods of fixed exchange rates in Argentina (1985–87) and Brazil (1986–87) during their failed so-called "heterodox" inflation stabilization plans, which used wage and price controls. In 1988, Mexico initiated a successful anti-inflation program designed to keep exchange rate depreciation slower than the inflation rate. (By late 1991, the Mexican peso's exchange rate was virtually fixed.) Argentina followed Mexico in early 1991 by fixing the Argentine exchange rate to the U.S. dollar and promising full convertibility of Argentine australs (now pesos) into dollars.

These most recent fixed exchange rate policies represent governments' use of exchange rates to signal the governments' intentions concerning future inflation and to increase the credibility of anti-inflation programs. Such exchange rate pegs are never fully credible because governments have an incentive to devalue and inflate once the public formulates inflation expectations. Some countries, especially Chile and Argentina, have enacted constraints to eliminate the governments' ability to renege on announced policy rules.

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