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Do Wages Help Predict Inflation?

Kenneth M. Emery
Senior Economist and Policy Advisor
Federal Reserve Bank of Dallas

Chih-Ping Chang
Graduate Student
Southern Methodist University

The inclusion of unit labor costs in forecasts of consumer price inflation provides no significant improvement in forecasting errors, especially in recent years.

Gauging inflationary pressures is a perennial concern for monetary policymakers and financial market participants. For example, during 1994, the Federal Reserve tightened monetary policy in response to concerns about building inflationary pressures and higher inflation in the future. Many people attributed the pursuant 1995 slowdown in economic activity largely to these Federal Reserve actions.

Unfortunately, the performance of many once reliable guides of future inflation, such as the growth of monetary aggregates, has deteriorated in recent years. This deterioration has resulted from, among other things, financial innovations that have changed the relationships between financial variables and economic activity. The deterioration of once reliable inflation guides has led policymakers and financial markets to monitor a broad range of inflation indicators. Because labor costs make up more than two-thirds of the total cost of producing goods and services in the United States, one of these indicators is unit labor costs, or wages adjusted for changes in labor productivity. Indeed, many analysts currently cite the lack of accelerating unit labor costs as grounds for believing that inflation will not increase any time soon.

Research on the relationship between unit labor costs and inflation has focused on whether higher labor costs precede higher inflation, or vice versa. In statistical jargon, the research has focused on whether labor costs Granger-cause inflation.[1] Recent research by Mehra (1993, 1991) that utilizes newly developed statistical techniques yields mixed results. Mehra (1993) finds that when consumer prices serve as the measure of prices, unit labor costs and prices are correlated in the long run. The study also finds that this correlation is present because Granger causality is running in both directions,[2] which implies that unit labor costs contain information about future consumer prices. However, Mehra (1991) finds that when the gross domestic product (GDP) deflator is used as the measure of prices, a long-run correlation still exists, but its source is Granger causality that runs only from prices to wages. Therefore, in this case, unit labor costs have no information content for future movements in prices.

The purpose of this article is twofold. The first is to examine how much forecasting power unit labor costs have for future consumer prices. Is the attention paid to unit labor costs as an inflation indicator justified? While Mehra finds that unit labor costs Granger-cause consumer prices, he does not examine the extent of unit

Figure 1
Growth of Unit Labor Costs and Consumer Price Inflation, 1957–93

labor costs' predictive power for out-of-sample forecasts of inflation.[3] The second purpose of this study is to examine whether the relationship between unit labor costs and consumer prices is stable over time.[4] As with the inflation-indicator properties of the monetary aggregates, have the indicator properties of unit labor costs deteriorated in recent years?

Our empirical strategy is to first take a preliminary look at the raw data and the data transformed by a filter designed by Baxter and King (1995). Next, we carry out Granger causality tests and a stability analysis of those tests. Finally, we examine the forecasting ability of unit labor costs for consumer price inflation (CPI). Our main finding is that the inclusion of unit labor costs in forecasts of consumer price inflation provides no significant improvement in forecasting errors, especially in recent years.

A preliminary look at labor costs and prices

Figure 1 plots year-over-year growth of unit labor costs and consumer price inflation, excluding food and energy (CPIC).[5] The high correlation between movements in labor costs and inflation demonstrates why analysts have paid close attention to labor costs when assessing inflation. However, what is not clear from the figure is whether movements in labor costs precede movements in inflation, or vice versa. In other words, it is not clear from Figure 1 whether movements in labor costs help to forecast future movements in inflation.

Notice also from Figure 1 that there appears to be a potential break in the relationship between labor costs and inflation sometime during the early 1980s. The growth of labor costs seems to be persistently lower than inflation growth during the 1980s, and the contemporaneous correlation between the two variables appears lower.

Using a filter methodology developed by Baxter and King (1995), we can divide labor cost growth and inflation into their long-run and business-cycle components. The results of doing this are shown in Figures 2A and 2B and illustrate that labor cost growth and inflation are correlated at both the business-cycle frequency and in their trend, or long-run, movements. Table 1 provides correlations from the raw data and for the trend and cycle components for the entire sample and for two subsamples.[6] The correlations for wages leading prices at the trend and business-cycle frequencies (negative *k*s) are positive, although higher at the trend frequency, supporting the view that movements in wages could help predict future movements in prices. Additionally, these correlations seem to be con-

Figure 2A
Inflation and Wage Growth Components

Trend components

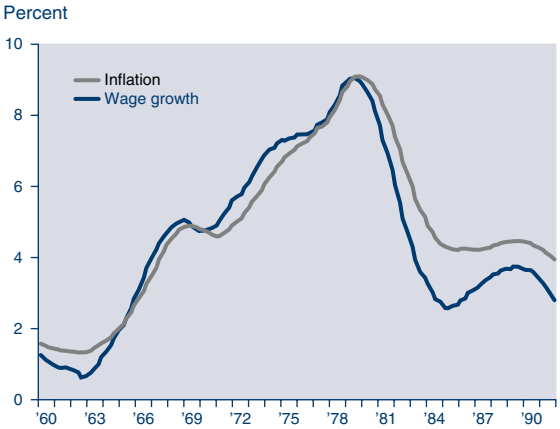


Figure 2B
Inflation and Wage Growth Components

Cycle components

Table 1
Cross Correlation: Inflation and Wage Growth

k	1965:4–80:4	1981:1–94:4	1957:1–94:4
Raw data			
–4	.37	.05	.54
–3	.50	.19	.60
–2	.59	.44	.68
–1	.56	.45	.67
0	.60	.51	.67
1	.38	.60	.63
2	.19	.51	.54
3	.04	.27	.42
4	–.09	.10	.35
Trend components			
–4	.79	.58	.94
–3	.85	.68	.95
–2	.90	.77	.96
–1	.94	.87	.96
0	.98	.96	.96
1	.90	.82	.93
2	.83	.68	.90
3	.75	.54	.87
4	.68	.41	.83
Cycle components			
–4	.45	–.35	.31
–3	.66	–.20	.49
–2	.81	.10	.64
–1	.83	.44	.72
0	.68	.72	.65
1	.38	.84	.46
2	.03	.76	.18
3	–.25	.51	–.09
4	–.40	.20	–.27

NOTES: The reported coefficients show the correlation between inflation at time *t* and wage growth at time *t* + *k*.

sistent with a potential breakpoint sometime in the early 1980s: for the raw data and both filter components, the wage leading inflation coefficients (negative-signed *k*s) drop in the 1980s. Additionally, for the raw data and the cycle data, the inflation leading wages coefficients (positive signed *k*s) increase during the 1980s.

Granger-causality results

Whole sample. We use both consumer prices for all items (CPI) and consumer prices excluding food and energy (CPIC) as our price measures.[7] Unit labor costs are for the nonfarm business sector. As a preliminary step to the formal causality tests, we have to determine the stationarity characteristics of the time series.[8] We choose the augmented Dickey–Fuller method (ADF) to conduct the tests for each variable in levels, first differences, and second differences. Table 2 summarizes the results and shows that unit labor costs and both price measures are integrated of order two, denoted by I(2).[9]

Granger causality tests with the variables in second differences will still be misspecified if inflation growth and wage growth are cointegrated and converge to a stationary long-run equilibrium relationship.[10] If the series are cointegrated, an error-correction term must be included in the causality test. This necessity follows from Engel and Granger's (1987) findings that if two variables are cointegrated, an error-correction model for the variables is present and that not including the error-correction term can lead to faulty inferences. Furthermore, cointegration between two variables implies Granger causation in at least one direction. The presence of cointegration provides a dynamic framework in which an error-correction term represents deviations from a long-run cointegrating relationship, while lagged difference terms represent short-run dynamics.

To estimate the possibility of a cointegrating relationship between the first difference of prices and unit labor costs, we use the Dynamic OLS (DOLS) procedure of Stock and Watson (1993). This procedure entails regressing one of the I(1) variables on the other I(1) variable, and lags and leads of the first differences of the I(1) variables. With standard errors corrected for serial correlation, one can make valid inferences from each coefficient estimate. The procedure is described by the following equations:

$$(1) \qquad \Delta p_t = \alpha_p + \beta_p \Delta w_t + \sum_{i=-k}^{k} \gamma_{pi} \Delta^2 w_{t-i} + \epsilon_{pt}$$

and

$$(2) \qquad \Delta w_t = \alpha_w + \beta_w \Delta p_t + \sum_{i=-k}^{k} \gamma_{wi} \Delta^2 p_{t-i} + \epsilon_{wt},$$

where p and w are the logarithms of prices and unit labor costs and Δ is the difference operator. Table 3 shows the results of testing the α's and β's.[11] Both β_p and Δ_w are significant at the 1-percent level, but the α's are only significant with CPIC. However, the augmented Dickey–Fuller tests for the cointegrating residuals confirm a stationary relationship between the growth of both price measures and the growth of unit labor costs, implying cointegration.[12]

We are now ready to conduct the Granger causality tests. To examine the causal relationship between inflation and wage growth, we estimate the following bivariate models:

$$(3) \qquad \Delta^2 p_t = a_p + b_p\left(\Delta p - \alpha_p - \beta_p \Delta w\right)_{t-1} + \sum_{i=1}^{k} c_{pi} \Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{pi} \Delta^2 w_{t-i} + \epsilon_{pt}$$

and

$$(4) \qquad \Delta^2 w_t = a_w + b_w\left(\Delta w - \alpha_w - \beta_w \Delta p\right)_{t-1} + \sum_{i=1}^{k} c_{wi}\Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{wi}\Delta^2 w_{t-i} + \epsilon_{wt},$$

where the terms in parentheses are the error-correction terms and are estimated by DOLS. The hypothesis of no causality from wage to inflation is rejected if b_p and/or all d_p's are significantly different from zero.

Our results for the causality tests are summarized in Table 4.[13] Whether wage growth Granger-causes inflation depends on the choice of the price series. For CPIC, wage growth is significant at the 1-percent level, implying causality. However, for CPI, wage growth is not significant, implying no causality.[14] The results also show that inflation always Granger-causes wage growth, regardless of the choice of the price series. It is noteworthy that the error-correction terms play a crucial role for the rejection of the no-causality hypotheses.

Granger-causality tests: A stability analysis. There are two potential sources of instability in the Granger causality test results presented above. First, there may be instability in the cointegrating relationship. Second, there may be instability in the short-run dynamics or in the Granger regressions themselves.

To examine the stability of the cointegrating relationship, we use Stock and Watson's (1993) formal test for the null hypothesis of a constant cointegrating relationship against the alternative of different cointegrating vectors over various samples. In their test for structural stability, the constant term in the error-correction term remains fixed. In contrast, our alternative test is based on the following regression:

$$(5) \qquad \Delta p_t = \alpha_p + \beta_p \Delta w_t + \left(\theta_p + \delta_p \Delta w_t\right) 1\left(t > \tau\right) + \sum_{i=-k}^{k} \gamma_{pi}\Delta^2 w_{t-i} + \epsilon_{pt},$$

where τ is the possible break date and $1(t > \tau)$ is the indicator function equal to 1 if $t > \tau$, 0 otherwise. The joint significance of θ_p and δ_p implies rejection of the null hypothesis of a stable long-run relationship. We conduct the joint test for θ_p and δ_p sequentially with all possible τ's, which covers 70 percent of the sample period. For each τ, we get a χ^2 statistic indicating the significance of θ_p and δ_p. The maximum of the sequence of χ^2 statistics yields a possible break in the cointegrating relationship.[15] It should be noted that the conventional χ^2 critical values are invalid because the timing of the structural change is not specified under the alternative. Hansen (1992) derives a SupF test for parameter instability in the context of cointegrated regression models. The far right column in Table 3 displays the SupF statistics and the selected break dates. Based on Hansen's asymptotic critical values, we discover a significant shift in regime. Depending on the price series, the break dates

Table 2
Augmented Dickey–Fuller Test Results

	τ-statistics	Lag order (k)	Ljung–Box Q-statistics
Levels			
CPI core	−2.68	3	15.70*
CPI	−2.76	4	13.81
Unit labor cost	−2.11	3	11.08
First-differenced			
CPI core	−2.44	7	14.89
CPI	−2.50	8	11.15
Unit labor cost	−2.34	8	12.32
Second-differenced			
CPI core	−4.62***	6	17.86*
CPI	−5.03***	8	13.69
Unit labor cost	−6.03***	8	13.35

*** = Significance at the 1-percent level.
** = Significance at the 5-percent level.
* = Significance at the 10-percent level.

NOTES: The testing equations are of the form:

$$y_t = \alpha + (\theta t) + \rho y_{t-1} + \sum_{i=1}^{k} \Delta y_{t-i} + \epsilon_t,$$

where the lag length k is determined by the Schwartz information criterion for $1 \le k \le 8$. Ljung–Box Q-statistics are used to check the serial correlation of the residuals. Q(9) and Q(10) are reported for the levels and for the first and second differences, respectively. All variables are in natural logs. The variables in levels are tested for trend stationarity, and the first- and second-differenced variables are tested for difference stationarity.

Table 3
Dynamic OLS Cointegration Test: Inflation and Wage Growth

	α_p	β_p	Dickey–Fuller τ-statistics	Break date (χ^2_2)
CPI	.16	1.03	−4.95***	1980:2 (15.97**)
	(.84)	(257.1***)		
CPI core	.65	.92	−4.04***	1980:4 (3.45***)
	(14.32***)	(257.2***)		

*** = Significance at the 1-percent level.
** = Significance at the 5-percent level.
* = Significance at the 10-percent level.

NOTES: (1) In the test of the significance of α and β, the reported χ^2_1 statistics use Newey and West (1987) robust standard errors with a truncation lag of 4. Augmented Dickey–Fuller tests are implemented on the cointegrating residual, $\Delta p - \alpha_p - \beta_p \Delta w$, with the lag length of eight determined by Schwartz information criterion.
(2) To search for the significant break dates, χ^2_2 is calculated to test for the changes in a and b of the cointegrating residuals over the 70 percent of the whole sample. The date is chosen based on the largest χ^2 statistic. Hansen's (1992) critical values are 16.2, 12.4, and 1.6 at the 1-, 5-, and 10-percent level, respectively.

Table 4
Causality Tests for Bivariate ECM

$$\Delta^2 p_t = a_p + b_p\left(\Delta p - \alpha_p - \beta_p \Delta w\right)_{t-1} + \sum_{i=1}^{k} c_{pi}\Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{pi}\Delta^2 w_{t-i} + \epsilon_{pt}$$

$$\Delta^2 w_t = a_w + b_w\left(\Delta w - \alpha_w - \beta_w \Delta p\right)_{t-1} + \sum_{i=1}^{k} c_{wi}\Delta^2 p_{t-i} + \sum_{i=1}^{k} d_{wi}\Delta^2 w_{t-i} + \epsilon_{wt}$$

Price equation

Null hypothesis	$b_p = 0$	$d_{pi} = 0$ (i = 1 ,2,3,4)	$b_p = 0$ and $d_{pi} = 0$ (i = 1,2,3,4)
(p = CPI core)			
1957:1–94:4	15.5 (.000)***	4.23 (.003)***	5.23 (.000)***
1957:1–80:4	14.0 (.000)***	2.99 (.024)**	3.79 (.004)**
1981:1–94:4	.27 (.609)	1.29 (.289)	1.04 (.407)
(p = CPI)			
1957:1–94:4	1.80 (.182)	1.42 (.230)	1.57 (.172)
1957:1–80:2	.07 (.796)	.97 (.428)	.81 (.546)
1980:3–94:4	.59 (.446)	.72 (.585)	.65 (.663)

Wage equation

Null hypothesis	$b_w = 0$	$c_{wi} = 0$ (i = 1,2,3,4)	$b_w = 0$ and $c_{wi} = 0$ (i = 1,2,3,4)
(p = CPI Core)			
1957:1–94:4	13.8 (.000)***	1.30 (.274)	6.41 (.000)***
1957:1–80:4	5.56 (.021)**	1.86 (.125)	4.46 (.001)***
1981:1–94:4	1.3 (.002)***	2.56 (.051)*	7.15 (.000)***
(p = CPI)			
1957:1–94:4	33.7 (.000)***	.82 (.516)	1.6 (.000)***
1957:1–80:2	28.9 (.000)***	.69 (.603)	9.35 (.000)***
1980:3–94:4	12.7 (.000)***	1.02 (.409)	3.41 (.010)**

*** = Significance at the 1-percent level.
** = Significance at the 5-percent level.
* = Significance at the 10-percent level.

NOTES: The error-correction term in both equations are estimated by Stock and Watson's Dynamic OLS with leads and lags equal to eight. *F*-statistics for the Granger causality tests are reported together with their *p*-values in the parentheses.

are 1980:4 and 1980:2. Therefore, the potential break point apparent in Figure 1 is supported by the formal stability tests.

In sum, the data suggest instability in the cointegrating regression. We now turn to the question of whether this instability affects the nature of the short-run dynamics between changes in wage growth and changes in inflation. In other words, does this instability affect the Granger causality tests? For this purpose, we estimate three Granger causality specifications over the two subsamples and the entire sample.[16] Two specifications are error-correction models (ECM). In the first ECM, the error-correction term is estimated from equations 1 and 2. In the second ECM model, the estimated cointegrating regression is estimated allowing a change in both the constant and the slope across the two subsamples. The third model does not contain an error-correction term.[17] The bottom line is that the Granger causality results from all three of these models turn out to be the same. However, because the original specification (the first ECM model) has a better fit across all samples, both in terms of R^2 and in superior forecasts that follow, we report in Table 4 results only from this model. Consequently, for the whole sample, the results of the Granger causality tests are unaffected.

Again, the whole sample results were that unit labor costs Granger-cause CPIC, but not CPI. Interestingly, the results for the subsamples in Table 4 indicate that wage growth Granger-causes CPIC in the pre-1980 sample only, not in the post-1980 sample. In neither subsample do wages cause CPI.[18] However, across all samples and for both price measures, prices cause wages.[19]

Summarizing, the in-sample causality tests indicate that changes in wage growth have information content for changes in core consumer price inflation, but not consumer price inflation. However, the information content for changes in CPIC appears to disappear after 1980. The information content of changes in both measures of inflation for future changes in wages appears to be much more robust. We now turn to an examination of the extent of the information content that wages have for inflation in out-of-sample forecasting exercises.

Forecasting exercises

Out-of-sample forecasts. The forecasting exercises consist of horse races between autoregressive univariate forecasts of inflation and forecasts obtained from the ECM models, which include unit labor costs.[20] The objective is to examine the reduction in forecast errors obtained by including the information content of wages. We carry out forecasts for the level of inflation and wages at forecast horizons of one, four, and eight quarters and for the three samples examined above.

The out-of-sample forecasts provide the real test of how forecasters would have done in real time using productivity-adjusted wages to help predict inflation. The first type of forecast we conduct is to use the ECM model estimated for the 1958–89 sample and then ask how it does in helping predict inflation from 1990 through 1994. As with the in-sample forecasts, we compare the root mean square errors (RMSEs) of the ECM model with a univariate autoregressive model. Each quarter, we update the parameter estimates of both models as the forecasts proceed through the 1990s.

The results are shown in the top section of

Table 5. For CPI inflation, the use of the ECM model actually results in a higher RMSE than the univariate model at all forecast horizons. For CPIC inflation, the use of the ECM model results in modest reductions in RMSEs, particularly at the four-quarter forecast horizon. Note that neither CPI nor CPIC inflation helps forecast wages, as the RMSEs in five out of the six forecasts actually increase with the ECM model.

Because the evidence suggests that there may have been a break in the wage–inflation relationship in the early 1980s, we also conduct out-of-sample forecasts in which the models are estimated using data only from the post-1980 period. In these forecasts, we initially estimate the models over the 1981–89 period and then conduct forecasts for the 1990–94 period. Again, the parameter estimates are updated as the forecasts progress through the 1990s. The results are shown in the middle section of Table 5. They indicate that breaking up the sample does not result in an improvement for the ECM model. In fact, the ECM model does worse.[21] Notice, however, that the RMSEs from the univariate models are the smallest of all the models considered. Therefore, the results indicate that for forecasting CPI and CPIC inflation during the 1990s, the use of only the post-1980 data results in lower forecast errors. Additionally, the inclusion of wage growth actually results in larger errors. Figure 3 plots the forecast errors from both models. Finally, the forecasts of wages indicate that the inclusion of CPIC inflation reduces the forecast errors at four- and eight-quarter horizons but the inclusion of CPI inflation does not result in lower forecast errors.

As a final exercise, we look to see whether, during the late 1970s, wages helped forecast inflation. In these forecasts, we estimate the models using data from the 1958–77 sample and then conduct out-of-sample forecasts for the period 1978–81. The results are shown in the bottom section of Table 5. For CPI, the use of the ECM model does not result in improved forecast errors. For CPIC, the use of the ECM model does result in an improvement, especially at the one- and four-quarter forecast horizons. At the four-quarter horizon, the use of the ECM model results in a 13-percent reduction in RMSE. Figure 4 plots the forecast errors. The results for inflation as a predictor of wage growth are mixed.

Table 5
Out-of-Sample Root Mean Square Forecast Errors

	Inflation			Wage growth		
Forecast horizons	**ECM**	**AR(4)**	**(Percent change)**	**ECM**	**AR(4)**	**(Percent change)**
Initial in-sample estimation: 1957:1–89:4						
Forecasting periods: 1990:1–94:4						
(CPIC)						
1	.800	.808	(1.10)	2.000	2.082	(3.96)
4	.908	1.015	(10.56)	2.445	2.411	(–1.37)
8	1.442	1.563	(7.76)	3.034	2.947	(–2.97)
(CPI)						
1	1.337	1.293	(–3.36)	2.221	2.082	(–6.67)
4	1.739	1.710	(–1.72)	2.686	2.411	(–11.39)
8	1.829	1.803	(–1.43)	3.047	2.947	(–3.42)
Initial in-sample estimation: 1981:1–89:4						
Forecasting periods: 1990:1–94:4						
(CPIC)						
1	.879	.814	(–8.07)	2.185	2.086	(–4.77)
4	.866	.785	(–10.33)	1.901	2.168	(12.35)
8	.856	.710	(–20.65)	1.893	2.260	(16.24)
(CPI)						
1	1.339	1.268	(–5.58)	2.213	2.086	(–.88)
4	1.568	1.439	(–8.95)	2.188	2.168	(–.88)
8	1.328	1.172	(–13.26)	2.099	2.260	(7.12)
Initial in-sample estimation: 1957:1–77:4						
Forecasting periods: 1978:1–81:4						
(CPIC)						
1	2.836	3.085	(8.08)	3.987	3.859	(–3.33)
4	3.102	3.564	(12.97)	5.017	4.442	(–12.93)
8	4.041	4.206	(3.92)	4.313	4.838	(10.86)
(CPI)						
1	2.413	2.358	(–2.31)	2.930	3.859	(24.07)
4	4.102	3.967	(–3.41)	5.526	4.442	(–24.39)
8	4.925	5.038	(2.24)	4.823	4.838	(.33)

NOTES: The forecasts for the ECM were formed using VAR(4) (including error-correction term and a constant). The entries in the table refer to the root mean square forecast error. To forecast inflation and wage growth for out-of-sample periods, we reestimate both ECM and AR model by updating the in-sample periods. For example, the four-quarter-ahead forecast for inflation and wage growth for 1991:1 is constructed by the models estimated over the period 1957:1–90:1 or 1981:1–90:1.

Figure 3
Four-Quarter Ahead Out-of-Sample Forecast Of CPI Core Inflation, 1990–94

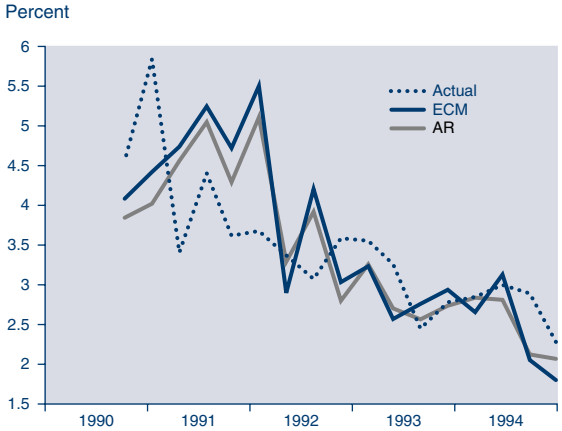
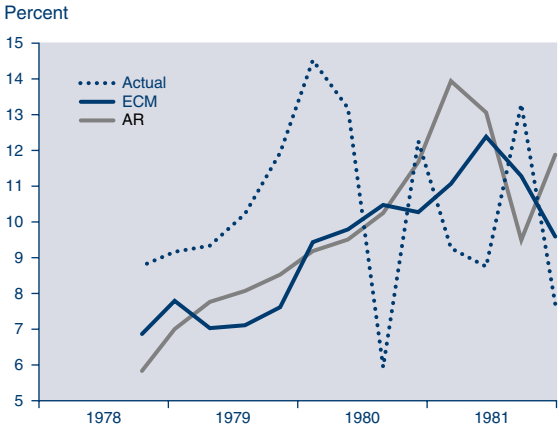


Figure 4

Four-Quarter Ahead Out-of-Sample Forecast Of CPI Core Inflation, 1978–81



In summary, out-of-sample forecasts offer little support that the growth of unit labor costs substantially helps forecast inflation, especially in recent years. For forecasting inflation during the 1990s, of the models we consider, a univariate autoregressive model of inflation using only post-1980 data results in the smallest forecast errors. The out-of-sample forecasts for the late 1970s indicate that wage growth did modestly help to forecast CPIC inflation.

Conclusions

Many analysts have heralded the slow growth of unit labor costs during recent years as a harbinger of continued low inflation. In this article, we investigate the usefulness of labor costs as a predictor of inflation. Earlier studies have focused on in-sample causality tests. Our in-sample causality tests indicate that, during the pre-1980 period, wage growth did have information content for future core inflation (CPIC) but not overall CPI inflation. During the post-1980 period, however, this information content has disappeared. Additionally, we find that the evidence of inflation causing wage growth is quite robust across samples.

In contrast with earlier studies, we also investigate out-of-sample forecasts of inflation using labor costs in an error-correction model. Out-of-sample forecasts offer the ultimate test of whether wages help predict future inflation. For recent years, the out-of-sample forecasting exercises offer no evidence that wage growth contributes to any reduction in forecast errors compared with univariate autoregressive models of inflation. Therefore, when assessing future inflation developments, these results suggest that policymakers and analysts should put little weight on recent wage trends.

Notes

We would like to thank Nathan Balke, Joseph Haslag, and Evan Koenig for helpful comments and suggestions. Any remaining errors are our own.

1 The Granger causality test is simply a statistical methodology for showing whether a variable contains information about subsequent movements in another variable.

2 However, Mehra finds that the presence of this bidirectional causality is sensitive to how inflation is modeled.

3 The motivation for Mehra's work is to examine the hypothesis that prices are marked up over productivity-adjusted labor costs, a central proposition of the expectations-augmented Phillips curve model. If that hypothesis is correct, then long-run movements in prices and labor costs must be correlated, and short-run movements in labor costs should help predict short-run movements in prices. Therefore, Mehra's results are consistent with the markup hypothesis for consumer prices but not for the implicit price deflator.

4 Any instability in the wage-price relationship could also be a source of instability in the price markup hypothesis and the Phillips curve (see Mehra 1993).

5 Unit labor costs are for the nonfarm business sector. The figure for consumer prices including food and energy is qualitatively similar. The formal analysis in this study is carried out using both measures of consumer prices.

6 The 1980:4 breakpoint was chosen arbitrarily on the basis of looking at Figure 1.

7 We use consumer prices both because they are perhaps the most closely watched measure of underlying inflation and because of Mehra's finding that labor costs do have information content for future consumer prices.

8 A time series is nonstationary if it has a time-varying mean and/or variance. Nonstationarity of a series violates an assumption underlying many statistical inferences and can lead to "spurious regression phenomenon," first described by Granger and Newbold (1974). One commonly used way of removing nonstationarity is to take first differences of the series.

9 In other words, second differencing is required for stationarity. Mehra (1991) reports similar results, while Mehra (1993) finds consumer prices to be I(1). Throughout the analysis that follows, we check the sensitivity of our results to the finding that prices are I(2).

10 The concept of cointegration, first proposed by Granger and Weiss (1983), is fundamental to the use of the error-correction model. Engel and Granger (1987) show that a model estimated using differenced data will be misspecified if the variables are cointegrated and the cointegrating relationship is ignored. Cointegration of two series means that they are nonstationary and tend to move together such that a linear combination of them is stationary. Cointegration is sometimes interpreted as representing a long-run equilibrium (steady-state) relationship.

[11] Since the standard errors are corrected for serial correlation, the χ^2 statistic is appropriate to test whether the α's and β's are significant.

[12] The ADF test is applied to the long-run relationship. In other words, $\Delta p - \alpha_p - \beta_p \Delta w$ and $\Delta w - \alpha_w - \beta_w \Delta p$.

[13] The Schwartz information criterion always implies a lag length of four or less. Since the results are not sensitive to the choice of lag length (k = 2, 4, or 8), we report only those from the model of k = 4.

[14] This result is consistent with Mehra (1991), which models consumer prices as I(2). However, Mehra (1993) models prices as I(1) and finds significant causality. As a robustness check, we also find causality for the whole sample and both price measures if we model prices as I(1).

[15] This maximum χ^2 statistic is sometimes called the Quandt likelihood-ratio statistic, which tests for a break in any or all of the coefficients.

[16] Formally, the rejection of the null of a stable cointegrating vector implies two alternatives: no cointegration and therefore no error-correction model, or an error-correction model in which there is assumed to be two cointegrating vectors, one from each subsample.

[17] Another important implication of Hansen's test is that the lack of cointegration is a special case of the alternative hypothesis, so the SupF test can also be viewed as a test of the null of cointegration against the alternative of no cointegration. If the SupF rejects the null, one may conclude that the standard model of cointegration, including its implicit assumption of long-run stability of the cointegrating relationship, is rejected by the data.

[18] For both CPIC and CPI, the results for the subsamples do not change when prices are modeled as I(1). Thus, only the whole sample results for the CPI are sensitive to whether prices are modeled as I(1) or I(2).

[19] Formally, the results from the two subsamples are conditioned on there being a structural break in the ECM. However, it should be noted that rolling formal stability tests cannot reject the null hypothesis of stability. Our choice to examine the results for the post-1980 sample was made on the basis of rejecting stability of the cointegrating relationship. Additionally, the results from the Baxter and King filter analysis and Figure 1 motivated us to examine the results for the post-1980 period.

[20] The ECM and autoregressive models with four lags perform superior to alternative lag lengths.

[21] Encompassing tests due to Chong and Hendry (1986) confirm this finding.

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Supply Shocks And the Distribution Of Price Changes

Nathan S. Balke
Associate Professor
Department of Economics
Southern Methodist University

Mark A. Wynne
Senior Economist and Policy Advisor
Federal Reserve Bank of Dallas

This article explores in some detail the dynamics of relative price changes in a simple dynamic general equilibrium model.

In a recent paper, Ball and Mankiw (1995) propose a new measure of supply shocks. Specifically, they advocate a measure of the skewness of price changes across sectors as a superior alternative to existing measures of supply shocks, such as the relative price of oil. Ball and Mankiw begin by showing that various measures of the skewness of the distribution of relative price changes across industries in the producer price index (PPI) are positively correlated with the rate of increase in the overall PPI. They further argue that these measures of skewness are better measures of supply shocks than more traditional measures such as the relative prices of food and energy when used in simple, short-term Phillips curve-type relationships. Their interpretation of the relationship between skewness and aggregate inflation relies heavily on the existence of menu costs associated with changing prices at the firm level. They further argue that since menu cost models were designed to explain monetary nonneutrality, these models "…gain [scientific] credibility from their ability to fit the facts regarding inflation and relative-price changes."

This article builds on the analysis of Ball and Mankiw by exploring in some detail the dynamics of relative price changes in a simple dynamic general equilibrium model. We begin by providing further evidence of a robust statistical relationship between the skewness of the distribution of individual price changes and inflation. We then ask what sort of relationship would we expect to see in a model in which all prices are free to adjust instantaneously. We show that when a simple general equilibrium model with no menu costs is calibrated to match certain features of the real world, it is possible to find a significant relationship between the skewness of individual price changes and aggregate inflation. Thus, our results cast some doubt on Ball and Mankiw's interpretation of the correlation between the skewness of the distribution of price changes and aggregate inflation as supportive of menu cost models.

Relative price changes as aggregate supply shocks

Ball and Mankiw begin their analysis by discussing a simple model in which menu costs associated with changing prices cause firms to adjust nominal prices only in response to large relative price shocks. The existence of menu costs implies the existence of range of inaction over which firms do nothing to change their prices in response to shocks. In Ball and Mankiw's

model, the adjustment to large but not to small shocks results in a positive correlation between the skewness and the mean of the (cross section) distribution of price changes. Ball and Mankiw argue that a flexible price model would predict no such relationship. In the flexible price model, a large positive relative price shock is likely to be offset by small declines in the prices of other commodities. Ball and Mankiw examine the cross section distribution of several hundred prices and indeed find a positive correlation between skewness and the mean of the distribution. They interpret this as evidence in favor of the menu cost model.

Yet, as we show below, it is possible for a flexible price model also to generate this positive correlation between skewness and the mean of the distribution of price changes. Ball and Mankiw's assertion that a flexible price model cannot generate this correlation is probably correct for the case in which a large number of sectors are experiencing shocks that are independent of one another and are of the same relative magnitude. This situation would cause relative price changes for the different commodities to be relatively independent of each other also. Yet, in reality, prices across commodities are not independent; shocks in one sector tend to affect prices in other sectors. Furthermore, a few very volatile sectors, such as food and energy, may be responsible for most of the observed volatility in the distribution of price changes. As a result, the correlation between the average inflation rate and the skewness of the distribution of price changes found in the data may arise just because of the importance of a few large price shocks.

Below we consider the implications for the data of a modified version of the general equilibrium model due to Long and Plosser (1983). The model is modified slightly to include a numeraire role for money. The model has complete price flexibility and multiple sectors. Among the key characteristics of this model is that a shock in one sector can spill over to other sectors. We show that as the sectors become more interrelated, it becomes easier for the flexible price model to generate a positive correlation between the skewness of the distribution of price changes and aggregate inflation.

The data

Ball and Mankiw look at the relationship between the distribution of prices in the producer price index (PPI) on an annual basis over the period 1949–89. The advantage of looking at the PPI is that it is available at a high degree of disaggregation. At the four-digit level of disaggregation, the number of component series rises from 213 in 1949 to 343 in 1989. Ball and Mankiw then document the relationship between the distribution of the changes in these several hundred price series and the overall inflation rate (as measured by the PPI).

Their data analysis reveals a number of interesting facts. First, there is considerable variation in the distribution of price changes over time. For example, in 1987 the distribution is fairly symmetric, while in 1973 it is skewed sharply to the right and in 1986 it is skewed sharply to the left.[1] Not surprisingly, both 1973 and 1986 were also years in which there were significant oil price shocks, with oil prices rising dramatically in 1973 and falling dramatically in 1986.

Ball and Mankiw document a statistically significant relationship between various measures of skewness and the overall inflation rate.[2] They also show that the skewness of the distribution of price changes tends to dominate the standard deviation of the distribution as an explanatory variable for inflation. This result is robust to their use of any of three measures of skewness.

In the analysis presented below we will examine the relationship between the distribution of price changes and aggregate inflation in the context of a multisectoral model that is calibrated to match certain characteristics of the U.S. economy. Considerations of tractability prevent us from considering a model with more than a small number of sectors. In fact, we work with a version of the real business cycle model proposed by Long and Plosser (1983) that has only six sectors. Before proceeding, then, it is important to verify that the empirical regularities observed by Ball and Mankiw in the prices that make up the PPI are also present when we consider more aggregated measures of prices.

The six sectors Long and Plosser use to calibrate their model are agriculture, mining, construction, manufacturing, transportation and trade, and services and miscellaneous. Table 1 presents some summary statistics for inflation rates (as measured by the implicit gross domestic product (GDP) deflators for these sectors) over the period 1949–93. The table reveals a number of interesting facts about the time series behavior of sectoral inflation rates.[3] First, there are notable differences in the average rates of inflation across the six sectors over the sample period, ranging from a low of just under 2 percent per year in agriculture to a

Table 1
Statistics on Inflation Rates by Sector, 1948–93
(Annual data)

	Mean	Variance
Agriculture	1.99	114.34
Mining	3.28	8.39
Construction	4.01	210.74
Manufacturing	5.23	21.53
Transportation and trade	3.41	7.02
Services and miscellaneous	4.94	4.45

high of nearly 5.25 percent per year in manufacturing. Second, there are dramatic differences in the variances of the individual inflation rates across sectors, from a low of 4.45 in services and miscellaneous to a high of 210.74 in construction.

Table 2 presents some simple regression results for the relationship between the rate of inflation as measured by the fixed-weight GDP deflator and measures of the distribution of prices across six sectors of the U.S. economy. The first column shows the results of regressing the rate of inflation on its own lagged value, while the second column shows the results of adding an unweighted measure of skewness to this basic regression. Comparing columns 1 and 2, we see that skewness has significant explanatory power for the inflation rate: the $\bar{R}^2$ increases from 0.52 to 0.66, and all of the coefficients in the second regression are significant at the 1-percent level. Column 3 shows what happens if we use a weighted measure of skewness instead. We obtain an even higher $\bar{R}^2$ and again all coefficient estimates are significant at the 1-percent level. The results in this table compare favorably with the results reported in Tables IIIA and IIIB of Ball and Mankiw. In fact, working with our more aggregated price data, we find an even stronger statistical relationship (in the sense of higher $\bar{R}^2$) between the skewness of the distribution of price changes and the aggregate inflation rate.

To summarize, it is clear that we can obtain the same strong statistical relationship between the skewness of the distribution of price changes and aggregate inflation looking at only six prices as Ball and Mankiw do looking at several hundred prices. The relationship uncovered by Ball and Mankiw seems robust, and our decision to focus on just six sectors does not seem to be a gross violation of the spirit of their analysis. Our objective in what follows is to see to what extent we can replicate the facts about the relationship between skewness and inflation as documented here in the context of a simple dynamic general equilibrium model.

A simple dynamic general equilibrium model with multiple sectors

A logical starting point for an investigation of the relationship between the distribution of price changes across sectors and aggregate inflation is the equilibrium business cycle model of Long and Plosser (1983). A great virtue of this model is that it has multiple sectors, but more importantly, the calculation of decision rules is simplified because of restrictions on preferences and the rate of depreciation of capital. The original version of this model was a "real"model in every sense of the word, in that there was no role for money.

For our purposes we would like to extend the model to include money as a numeraire. Benassy (1995) has recently proposed a version of the Long and Plosser model that incorporates money by including real balances in the utility function. While introducing money into the model in this way is not entirely satisfactory, it is well-known that this specification is functionally equivalent in certain circumstances to the more popular cash-in-advance and shopping-time formulations of the demand for real balances. However, Benassy works with a single-sector variant of the Long and Plosser model, and it is far from straightforward to extend his analysis to a multiple-sector setting (the essence of the problem that arises in this regard is the absence of a single correct measure of the price level in a multisector environment). An alternative is to introduce money via some sort of cash-in-advance constraint on either purchases of consumption goods (or some subset thereof) or purchases of capital goods, or both. However, it rapidly becomes apparent that it is no longer possible to calculate simple closed form decision rules in either of these cases.

We opt instead to introduce money in a somewhat novel manner. Specifically, we assume that consumers are obliged to hold some fraction υ of their consumption purchases during each period in the form of cash at the end of the period. Thus, we posit the following constraint on household choices:

$$M_t \geq \upsilon \sum_{i=1}^{N} P_{i,t} C_{i,t}, \tag{1}$$

where M_t denotes the stock of nominal money balances held at the *end* of period t and $\sum_{i=1}^{N} P_{i,t} C_{i,t}$ denotes nominal consumption expenditures *during* period t, with $P_{i,t}$ denoting the price of good i at date t, and $C_{i,t}$ denoting the quantity of good i purchased for consumption

purposes at date t. The existence of this constraint can be thought of as arising due to the need to, say, maintain some minimum level of cash balances in a bank account to facilitate consumption purchases made with inside money. It will turn out that money does not play a very important role in our economy.

The rest of the model is relatively standard.

Households. The economy is populated by a large number of identical consumers, each of whom has preferences summarized by the following utility function:

$$U = \sum_{i=0}^{\infty} \beta^t u(C_t, L_t), \tag{2}$$

where $1 > \beta > 0$ is the discount factor, $C_t = (C_{1,t}, C_{2,t}, \ldots, C_{N,t})'$ is an $N \times 1$ vector of commodities consumed at date t, and L_t denotes leisure at date t. The point-in-time utility function is furthermore assumed to have the following specific functional form:

$$u(C_t, L_t) = \theta_0 \log(L_t) + \sum_{i=1}^{N} \theta_i \log(C_{i,t}), \tag{3}$$

where $\theta_i \geq 0$, $\forall$ i. If $\theta_i = 0$ for some $i \geq 1$ then that commodity has no utility value to the consumer.

The budget constraint of the representative consumer is given by

$$\begin{aligned} &\sum_{i=1}^{N} W_{i,t} H_{i,t} + \sum_{i=1}^{N} \sum_{j=1}^{N} R_{i,j,t} K_{i,j,t-1} + \mu_t M_{t-1} \\ &\geq \sum_{i=1}^{N} P_{i,t} C_{i,t} + \sum_{j=1}^{N} P_{j,t} \sum_{i=1}^{N} K_{i,j,t} + M_t, \end{aligned} \tag{4}$$

where $W_{i,t}$ denotes the (nominal) wage in sector i at date t (which in equilibrium will be the same in all sectors), $H_{i,t}$ denotes hours worked in sector i at date t, $R_{i,j,t}$ denotes the nominal rental rate on type j capital employed in sector i in period t, $K_{i,j,t-1}$ represents the quantity of type j capital employed in sector i during period t, (that is, capital in place at the end of period $t-1$) and μ_t represents the gross rate of increase in the money stock at date t. The sources of funds each period are wage income $\sum_{i=1}^{N} W_{i,t} H_{i,t}$, income from capital $\sum_{i=1}^{N} \sum_{j=1}^{N} R_{i,j,t} K_{i,j,t-1}$ and a transfer from the government that is directly proportional to nominal money holdings held at the end of the previous period, $\mu_t M_{t-1}$. The uses of funds each period are consumption expenditures, $\sum_{i=1}^{N} P_{i,t} C_{i,t}$, purchases of new capital equipment, $\sum_{j=1}^{N} P_{j,t} \sum_{i=1}^{N} K_{i,j,t}$, and funds held over to the next period, M_t.

The remaining constraint that the consumer faces is on the allocation of available time,

$$L_t + \sum_{i=1}^{N} H_{i,t} = 1, \tag{5}$$

which states that the sum of leisure and time worked in each sector cannot exceed the total amount of time available, which we normalize to 1.

The consumer's problem is to maximize the objective function given in equation 2 above subject to the budget constraint (equation 4), the cash constraint (equation 1), and the constraint on the allocation of time (equation 5), taking as given the prices at which he or she can purchase consumption and capital goods and the wage and rental rates at which labor and capital services are sold to the business sector.

Firms. Production possibilities in the i'th sector are given by the following production function:

$$Y_{i,t} = Z_{i,t} H_{i,t}^{b_i} \prod_{j=1}^{N} K_{i,j,t-1}^{a_{i,j}}, \tag{6}$$

where $Y_{i,t}$ denotes output of the i'th good at date t, $Z_{i,t}$ is a random variable or productivity shock that denotes the state of technology in the i'th sector at date t, $H_{i,t}$ denotes hours worked in the i'th sector at date t, and $K_{i,j,t}$ denotes the quantity of output of the j'th industry employed in the i'th industry at date t. The parameters of the production function, b_i and $a_{i,j}$ are assumed to satisfy $b_i > 0$, $a_{i,j} > 0$ and $b_i + \sum_{j=1}^{N} a_{i,j} = 1$ for i = 1, 2, …, N. The production side of this model can be viewed in two

Table 2
Inflation and the distribution of price changes, 1948–93

	(1)	(2)	(3)
Constant	.010*	.013**	.014**
	(.005)	(.004)	(.004)
Lagged inflation	.715**	.642**	.609**
	(.102)	(.088)	(.080)
Unweighted skewness		.011**	
		(.003)	
Weighted skewness			.004**
			(.001)
$\bar{R}^2$	.52	.66	.72
Durbin–Watson statistic	1.54	1.78	2.02

*,** denotes significance at the 5-percent and 1-percent levels, respectively.

NOTE: Standard errors are in parentheses.

ways. We can think of each sector as producing both consumption and capital goods that are used in every sector, with the capital depreciating at a 100-percent rate. Or, we can think of each sector as producing consumption goods and goods that are used as intermediate inputs in the production of other goods. The two interpretations are equivalent.

The firm's optimization problem is to maximize profits, taking as given the available technology, the price at which output can be sold and the prices or rental rates of the labor and capital inputs.

Equilibrium. Straightforward manipulation of the first order conditions for the household and firm maximization problems allows us to obtain the following decision rules:

$$C_{i,t} = \left(\frac{\theta_i}{\gamma_i}\right) Y_{i,t}, \tag{7}$$

$$L_t = \frac{\theta_0(1+\nu(1-\beta))}{\theta_0(1+\nu(1-\beta)) + \sum_{i=1}^{N} \gamma_i b_i}, \tag{8}$$

$$H_{i,t} = \frac{\gamma_i b_i}{\theta_0(1+\nu(1-\beta)) + \sum_{j=1}^{N} \gamma_j b_j}, \text{ and} \tag{9}$$

$$K_{i,j,t} = \left(\frac{\beta\gamma_i a_{i,j}}{\gamma_j}\right) Y_{j,t}, \tag{10}$$

where $\gamma_j = \theta_j + \beta\sum_{i=1}^{N} \gamma_i a_{i,j}$. The derivation of these rules is explained in more detail in the Appendix.

Some comments are in order. As noted, the simple form of these decision rules stems from the particular assumptions we have made about preferences, production possibilities and the durability of capital. Equation 7 shows that consumption of each type of good is simply a constant fraction (θ_i/γ_i) of the available output of that type of good, with the fraction being a complicated function of the parameters of the underlying preferences and technology. Likewise, equation 10 shows that the amount of each sector's output that is allocated for use in production in other sectors is a constant fraction ($\beta\gamma_i a_{i,j}/\gamma_j$) of the available output. Perhaps more surprisingly, equations 8 and 9 show that total leisure and the allocation of effort to each sector are independent of realizations of the productivity shocks and are also independent of the endogenously chosen "state" of each sector, as summarized by the level of output produced in each sector.

To better understand why the allocation of labor across sectors (and total labor or leisure) is independent of the current state of the economy, consider the condition determining the equilibrium allocation of labor.[4] This condition states that the value of the marginal product of labor in each of its alternative employments should equal the wage rate, where all prices and wages are denominated in utility units. The wage rate is simply the marginal utility of leisure. The question then is, Given an initially optimal allocation of labor across sectors, would a change in either the available capital stock or the state of technology change either side of this equation? Given the specification of preferences we are working with, the marginal utility of leisure at any point in time depends only on the labor–leisure allocation at that time, so any effect of the capital stock or technology on the optimal allocation must come about through changes in the value of the marginal product. Consider the effect of a higher than expected realization of the technology. For a given allocation of capital to a particular sector, one effect of the technology shock would be to raise the marginal physical product of labor. However, working against this, the technology shock will put downward pressure on the price of the sector's output, lowering the value of the marginal product of labor. It just so happens in this case that these two effects offset each other, leaving the value of the marginal product unchanged. In other words, the optimal allocation of labor to the sector is unaffected by realizations of the technology shock. Similar reasoning applies to determining the effects of greater availability of capital in a sector.

We can use the equations above to write dollar-denominated prices in our extended model as

$$P_{i,t} = \frac{1}{\nu}\frac{\gamma_i}{\sum_{j=1}^{N}\theta_j}\frac{M_t}{Y_{i,t}}. \tag{11}$$

That is, nominal prices are directly proportional to the nominal money stock.[5] It is straightforward to show that these prices are also directly proportional to the utility-denominated prices calculated by Long and Plosser.

Dynamics. The dynamic behavior of this economy is implied by the technology as summarized by the production functions, along with the decision rules for the inputs to the production processes. It is convenient to write the system in logarithmic form as follows,

$$y_t = k_y + Ay_{t-1} + z_t, \tag{12}$$

where we adopt the convention that lower case letters denote the logarithms of the corresponding upper case variable. Thus, y_t is the $N \times 1$ vector $(\log(Y_{1,t}),\log(Y_{2,t}),\ldots,\log(Y_{N,t}))'$, k_y is an $N \times 1$ vector of constants, and z_t is the $N \times 1$ stochastic vector $(\log(Z_{1,t}),\log(Z_{2,t}),\ldots,\log(Z_{N,t}))'$. Since our primary focus in this article is on the evolution of the distribution of prices, we will also need to specify a stochastic process for the log of the nominal money stock, m_t.

The evolution of prices is given by

$$p_t = k_p + \iota_N m_t + y_t, \tag{13}$$

where $p_t = (\log(P_{1,t}),\log(P_{2,t}),\ldots,\log(P_{N,t}))'$, k_p is a vector of constants and ι_N is an $N \times 1$ vector of ones. An important point to note from this expression is that the money stock only affects the mean of the distribution of prices across sectors and not any of the higher moments (such as the standard deviation or skewness).

Measures of the aggregate price level. We can easily calculate a variety of measures of the aggregate price level that correspond to the measures commonly used to gauge inflation in the real world. Three such measures are defined in the Appendix. We will concentrate on just one of them, a fixed-weight measure of the aggregate price level that corresponds to the fixed-weight GDP deflator. We construct a fixed-weight GDP deflator starting from the definition

$$PGDPF_t = \frac{\sum_{i=1}^{N} P_{i,t}Y_{i,b}}{\sum_{i=1}^{N} P_{i,b}Y_{i,b}}. \tag{14}$$

That is, the value of the fixed-weight GDP deflator at date t, $PGDPF_t$, equals the ratio of the cost of purchasing the base-year market basket of final output at current-year prices, $\sum_{i=1}^{N} P_{i,t}Y_{i,b}$, to the cost of purchasing the base-year market basket of final output at base-year prices, $\sum_{i=1}^{N} P_{i,b}Y_{i,b}$. Making substitutions using equation 11 above, we obtain

$$PGDPF_t = \frac{M_t}{M_b}\frac{1}{\sum_{i=1}^{N}\gamma_i}\sum_{i=1}^{N}\frac{\gamma_i Y_{i,b}}{Y_{i,t}}, \tag{15}$$

which can also be rewritten in logs as

$$\begin{aligned} pgdpf_t &\equiv \log(PGDPF_t) \\ &= m_t - \overline{m} - \log\left(\sum_{i=1}^{N}\gamma_i\right) \\ &\quad + \log\left(\sum_{i=1}^{N}\gamma_i \exp(\overline{y}_i - y_{i,t})\right), \end{aligned} \tag{16}$$

where bars over variables denote steady-state values, which we pick as the base year.

Calibration. Long and Plosser calibrate their model to six sectors of the U.S. economy using a consolidated version of the twenty-three sector input–output table for 1967 (U.S. Department of Commerce, 1975). This yields an estimate of the A matrix for the model above. Given an estimate of the A matrix, the vector of coefficients b is recovered from the assumption of constant returns to scale, *i.e.*, $b_i = 1 - \sum_{j=1}^{N} a_{i,j}$. To calibrate the vector θ, we note that the decision rules for consumption of each type of good imply that $\theta_i = \gamma_i \frac{C_{i,t}}{Y_{i,t}}$. We can obtain estimates of the share (γ_i) of each sector's output in aggregate output from the 1967 input–output table. The same table also allows us to estimate the fraction of each sector's output that was allocated to consumption that year, which together with the estimate of γ_i allows us to obtain an estimate of θ_i. Finally, we set the discount factor β equal to 0.95 and the parameter ν equal to 1.

Experiment 1. Our first experiment examines the behavior of inflation and the distribution of prices in an economy with six sectors but with no input–output relations between the sectors and with each sector subject to *i.i.d.* shocks of equal variance. Thus, we set

$$A = \begin{pmatrix} 0.33 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0.33 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0.33 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0.33 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0.33 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0.33 \end{pmatrix}$$

and $\theta = (0.167, 0.167, 0.167, 0.167, 0.167, 0.167)'$. We assume that the technology innovations hitting each sector are *i.i.d.* with zero mean and unit variance. A priori, we expect that there will be no relationship between measures of the cross-section distribution of prices and aggregate inflation in this economy. We think that Ball and Mankiw have an economy such as this in mind when they question the ability of a flexible price model to generate the correlations

between skewness and inflation that are found in the data.

Experiment 2. For our second experiment, we calibrate the A matrix and the θ vector along the same lines as Long and Plosser, but retain the assumption of *i.i.d.* shocks. Thus, we set

$$A = \begin{pmatrix} 0.4471 & 0.0033 & 0.0146 & 0.2093 & 0.0999 & 0.1591 \\ 0.0000 & 0.0935 & 0.0427 & 0.1744 & 0.0549 & 0.4854 \\ 0.0029 & 0.0104 & 0.0003 & 0.4189 & 0.1209 & 0.0893 \\ 0.0618 & 0.0340 & 0.0050 & 0.4576 & 0.0611 & 0.1267 \\ 0.0017 & 0.0004 & 0.0166 & 0.1246 & 0.1040 & 0.3249 \\ 0.0174 & 0.0212 & 0.0595 & 0.1998 & 0.0871 & 0.3805 \end{pmatrix}$$

and $\theta = (0.003465, 0.000162, 0.022046, 0.139968, 0.089811, 0.15012)'$. The economy of this experiment differs from that of the first experiment mainly in allowing for complicated input–output type relations between the various sectors.

Experiment 3. For our third experiment, we retain the specifications of the A matrix and θ vector used in the second experiment but calibrate the technology shocks to the actual postwar data. Thus, we estimate Solow residuals for each sector as

$$z_{i,t} \equiv \log(Z_{i,t}) = \log(y_{i,t}) - (1-\alpha_i)\log(k_{i,t-1}) - \alpha_i \log(n_{i,t})$$

where $\log(y_{i,t})$ is the BP-filtered log of output in sector i produced during period t, $\log(n_{i,t})$ is the BP-filtered log of full time equivalent employees in the i'th sector during period t, $\log(k_{i,t-1})$ is the BP-filtered log of the net (real) capital stock in the i'th sector as of the *end* of period $t-1$ (*i.e.,* capital available in the i'th sector at the *beginning* of period t), and α_i is the average value over the sample period (1947–94) of labor's share in the i'th sector (defined as the ratio of nominal compensation of employees in the i'th sector to nominal GDP in that sector). For the BP filter, we used the parameter values $up = 2$, $dn = 8$ and $K = 3$.[6]

We assume that total factor productivity in the model evolves according to

$$z_t = Pz_{t-1} + \epsilon_t$$

where we use OLS to estimate the matrix P as

$$P = \begin{pmatrix} 0.109 & 0.0 & 0.0 & 0.0 & 0.0 & 0.0 \\ 0.0 & 0.024 & 0.0 & 0.0 & 0.0 & 0.0 \\ 0.0 & 0.0 & 0.458 & 0.0 & 0.0 & 0.0 \\ 0.0 & 0.0 & 0.0 & 0.127 & 0.0 & 0.0 \\ 0.0 & 0.0 & 0.0 & 0.0 & 0.106 & 0.0 \\ 0.0 & 0.0 & 0.0 & 0.0 & 0.0 & 0.329 \end{pmatrix}$$

The innovations $\epsilon_t = (\epsilon_{1,t}, \epsilon_{2,t}, \epsilon_{3,t}, \epsilon_{4,t}, \epsilon_{5,t}, \epsilon_{6,t})'$ are assumed to be *i.i.d.* with variances $\sigma^2 = (1.01, 1.26, 0.25, 0.54, 0.19, 0.03) \times 10^{-3}$.

Experiment 4. For our fourth experiment, we estimate a simple VAR for the technology innovations in each sector. Again, we assume total factor productivity follows the process

$$z_t = Pz_{t-1} + \epsilon_t$$

where now the matrix P is given by

$$P = \begin{pmatrix} 0.231 & 0.285 & 0.205 & 1.031 & 0.872 & 1.831 \\ 0.006 & 0.116 & 0.948 & 0.468 & 0.344 & 2.334 \\ 0.044 & 0.031 & 0.515 & 0.458 & 0.438 & 0.022 \\ 0.188 & 0.127 & 0.438 & 0.074 & 0.188 & 0.230 \\ 0.083 & 0.068 & 0.348 & 0.075 & 0.020 & 0.092 \\ 0.065 & 0.004 & 0.075 & 0.066 & 0.166 & 0.079 \end{pmatrix}$$

and again the innovations ϵ_t are assumed to be *i.i.d.* with variances $\sigma^2 = (0.75, 0.87, 0.20, 0.41, 0.13, 0.02) \times 10^{-3}$.

Each of these experiments introduces progressively more interaction between the sectors and allows for greater diversity in the shocks hitting the sectors. In the first experiment, there is no interaction and the shocks hitting each sector are completely independent of one another. The second experiment allows for interaction through input–output relationships but retains the assumption of independent shocks. The third experiment allows for interaction through input–output relationships and allows for serially correlated shocks in each sector. The fourth experiment allows for input–output type interaction between sectors and for serial correlation in the state of technology across sectors.

One final comment on the experiments. In each of the four experiments reported here, we hold the money stock constant, so that the only source of fluctuations in the model economies are technological innovations. This technique allows us to completely isolate the effects of what Ball and Mankiw call "supply shocks" on the relationship between the distribution of price changes and aggregate inflation.

Results

For each experiment, we calibrate the model as described above and simulate it for fifty periods one hundred times. We use the time series on prices generated in each of the one hundred simulations to run the regressions described in Table 2. Table 3 reports the average value over all one hundred simulations of the regression coefficients on the weighted and unweighted measures of the skewness of the

distribution of prices at each date for an inflation regression in which we use the rate of change of a fixed-weight measure of the GDP deflator as the measure of inflation. For comparison, we also report the coefficients estimated using actual data.

Moving down the rows of Table 3, we see that in the basic economy with no interaction between sectors and *i.i.d.* shocks of equal variance hitting each sector, we are unable to generate a significant role for skewness in explaining the inflation rate. Note that the coefficient estimates are the same for the weighted and unweighted measures of skewness as all sectors are identical by construction. Moving to the economy of experiment 2, there is still no role for unweighted skewness in explaining inflation, but the weighted measure is now significant. Recall that this economy differs from that in experiment 1 only in that it allows for input–output type relationships between all of the sectors. For the economies of experiment 3 and experiment 4, the weighted measure of skewness helps explain inflation. However, in none of our experiments is the unweighted measure correlated with inflation in a statistically significant sense.

Summarizing our results, it is clear that we are able to replicate to a surprising degree the key correlation between skewness and inflation Ball and Mankiw find in the data. Most importantly, we are able to do so in the context of a simple equilibrium model with fully flexible prices, thus raising questions about Ball and Mankiw's interpretation that this correlation results from the existence of menu costs. In Balke and Wynne (1995), we also document other aspects of the relationship between the distribution of price changes and the overall inflation rate and show that skewness seems to be a leading indicator of aggregate inflation. Our model has less success in replicating this feature of the data.

Table 3
Estimated coefficients on skewness in inflation regression

	Unweighted skewness	Weighted skewness
Data	.011**	.004**
	(.003)	(.001)
Experiment 1	.038	.038
	(.182)	(.182)
Experiment 2	–.037	.297**
	(.222)	(.102)
Experiment 3	.000	.005**
	(.003)	(.002)
Experiment 4	.001	.007**
	(.004)	(.003)

*,** denotes significance at the 5-percent and 1-percent levels, respectively.

NOTE: Standard errors are in parentheses.

SOURCE: Authors' calculations.

Conclusions

In this article, we explore the relationship between shifts in the distribution of prices and the aggregate inflation rate in the context of a simple dynamic general equilibrium model with multiple sectors. The idea that changes in the distribution of relative price changes might have implications for the overall inflation rate was first proposed by Ball and Mankiw (1995). A crucial part of the story that they tell is that firms face significant adjustment costs associated with changing nominal prices. The existence of these so-called menu costs means that firms respond (in the sense of adjusting their prices) to large shocks but not to small shocks. We show that in the context of a simple dynamic general equilibrium model with no costs of adjusting prices it is possible to observe the same correlation between the skewness of the distribution of price change and the overall inflation rate. Our model is driven solely by supply shocks in the form of technological innovations.

The analysis in this article leaves a number of issues open for further research. First, it would be interesting to document in a more thorough fashion the behavior of the distribution of price changes over the business cycle and its relationship to aggregate activity and aggregate inflation. It would also be interesting to extend the analysis above to a model with more sophisticated dynamics and a more important role for money. Finally, it would be interesting to extend the model outlined above to allow for a limited degree of price stickiness (say along the lines of Ohanian, Stockman, and Kilian 1994) and see how much nominal rigidities can contribute to explaining the relationship between the distribution of price changes and inflation in an equilibrium model. Some of these issues are addressed in Balke and Wynne (1995).

Notes

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[1] The skewness of a distribution is defined as $E[(x-\mu)^3]/\sigma^3$, where μ is the mean of the distribution of X and σ is the standard deviation.

[2] Ball and Mankiw look at two measure of skewness in addition to the conventional measure defined in note 1 above. The first, intended to measure the mass in the

upper tail of the distribution relative to the mass in the lower tail, is defined as

$$AsymX = \int_{-\infty}^{-X} rh(r)dr + \int_{X}^{\infty} rh(r)dr,$$

where r is defined as the relative price *change* defined as the four-digit industry inflation rate minus the average of the four-digit industry inflation rates (*i.e.*, $r_{it} = \Delta\log(p_{i,t}) - \frac{1}{N}\sum_{i=1}^{N}\Delta\log(p_{i,t})$ and $h(r)$ is the density of r. The tails are defined as relative price changes greater than X percent or less than $-X$ percent. Their second alternative measure of skewness is

$$Q = \int_{-\infty}^{\infty} \left| r \right|^{\bullet} rh(r)dr.$$

This variable measures the average of the product of each relative price change and its own absolute value.

[3] Formal tests for nonstationarity indicate that all of the sectoral inflation rates with the exception of agriculture are nonstationary, meaning that in samples of infinite size the variances of these series will be undefined. However, this is not necessarily a problem from our perspective as the data in Table 1 are simply presented to illustrate differences in the rates of change of prices in different sectors.

[4] See also Long and Plosser (1983, 49–50).

[5] Note that in the basic Long and Plosser model, nominal GDP (denominated in utility terms) is a constant. In our extended model, nominal GDP (denominated now in terms of dollars) is directly proportional to the money stock.

[6] The BP filter is explained in Baxter and King (1995).

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Policy Priorities And the Mexican Exchange Rate Crisis

William C. Gruben
Research Officer
Federal Reserve Bank of Dallas

This article considers factors that led to Mexico's December 1994 devaluation, examines why it seemed to surprise markets, and addresses financial market behavior in the wake of the peso devaluation. It is useful to consider factors that led to devaluation in the framework of the so-called impossibility theorem.

Immediately following Mexico's December 20, 1994, devaluation of the peso, some observers expressed detailed support for the move. They conjectured that it would calm financial markets that had showed some signs of volatility.

Mexican officials themselves treated the devaluation as if it would have a stabilizing influence. In presenting the devaluation, they announced that the country's erstwhile crawling peg regime would remain in place, and that the devaluation would represent only a change in the weak-side edge of Mexico's exchange rate band.

In the United States, some analysts also accepted Mexico's initial devaluation as equilibrating. Then Acting International Monetary Fund Director Stanley Fischer noted Mexico's initiatives as "an appropriate policy response to recent market developments…."[1] MIT Professor Rudiger Dornbusch, who had for some time been advocating a Mexican devaluation, was quoted as saying that now was the time for "smart money" to move into Mexico, with the currency at appropriate levels.[2]

But instead of inducing stability in financial markets, the initial devaluation triggered a run on the currency. Foreign currency reserves fell markedly. Mexican interest rates rose rapidly. The exchange rate moved well beyond what advocates of Mexican devaluation had said they thought sufficient.

While subsequent exchange rate and interest rate reactions to Mexico's initial devaluation raised questions about how financial markets operate, financial events preceding Mexico's initial devaluation of the peso also offered anomalies. Financial markets typically sense impending devaluations. This time, despite concerns voiced occasionally that the peso was overvalued, reports were widespread that the timing had surprised financial markets[3] and even that investors felt betrayed.[4]

This article considers factors that led to the devaluation, examines why it seemed to surprise markets, and addresses financial market behavior in the wake of the peso devaluation. It is useful to consider factors that led to devaluation in the framework of the so-called impossibility theorem. Recall that this theorem claims that policy authorities cannot simultaneously and continuously follow the three objectives of free capital mobility, fixed exchange rates, and an independent monetary policy.

I detail why policymakers might reasonably conclude that Mexico could possibly pursue limited episodes of monetary indepen-

dence—even if the impossibility theorem ultimately could not be denied—and why the possibility could have been permitted to become a reality. Nevertheless, when witnessing a financial panic like Mexico's, it may require effort to recall that exchange rate devaluation is a matter of choice. A central bank can always maintain a pegged exchange rate. The price is contraction in the monetary base or, equivalently, a persistently high interest rate.[5]

This article outlines the rise of priorities that came to dominate the preservation of the Mexican exchange rate regime. Specifically, already high real interest rates, resulting increases in nonperforming loan rates, and the implications of all of these factors for commercial bank solvency seem in part to have motivated credit creation at times in 1994 while the United States was tightening credit.

Mexico's pre-devaluation exchange rate and monetary policy independence

While the impossibility theorem posits that policy authorities cannot simultaneously and continuously follow the three objectives of free capital mobility, fixed exchange rates, and an independent monetary policy—the meaning of the term *continuously* complicates matters for anyone who wants to analyze Mexico. How continuous does continuous have to be?

Before the December 1994 devaluation, Mexico's exchange rate was essentially pegged to the U.S. dollar, but Mexico gave itself what appeared to be some room to maneuver. In pegging the peso to the dollar, Mexico was announcing its intent to cede some control over its monetary policy to the United States. One advantage of taking this step and then persisting with it is to establish credibility that, in general, noninflationary policies would be in place.

If Mexico had fixed its exchange rate policy hard and fast to the dollar, it would have been fully ceding its monetary policy to the United States. But in fact, Mexico permitted its exchange rate to fluctuate within a band whose weak-side edge devalued at 0.0004 pesos per dollar. With a band, Mexico's central bank could run expansionary or contractionary monetary policies different from those of the U.S. central bank—provided that the resulting movements in the exchange rate remained within the band—and still maintain exchange rate credibility.

An important detail of Mexico's monetary independence, however, involved what may be seen as its term structure. U.S. short-term rates appear not to have an influence on Mexican short-term or long-term interest rates, but U.S. long-term rates seem to influence Mexican long-term interest rates.[6]

To the extent that these data suggested limited financial integration in short-term markets, Mexico may have perceived itself able to pursue a relatively independent monetary policy in the short run. In any case, as will be detailed, Mexico did pursue a monetary policy in the second half of 1994 that was inconsistent with the United States' increasingly restrictive approach to money market intervention.

The implications of the U.S. long-term debt to Mexican long-term debt relations suggest that Mexico's monetary policy could not remain independent in the longer term—at least not in a pegged exchange rate regime. Once the United States moved its long rates, Mexico would have to follow quickly or face large capital outflows. There is much to suggest that political factors contributed to the day-to-day changes in capital outflows that ultimately occurred but, in the end, monetary policy in Mexico was not consistent with reversing them.

Tensions within Mexico's exchange-rate-based stabilization plan

Incomes policies. Exchange-rate-based stabilizations are very difficult to pursue effectively over protracted periods. In programs like Mexico's, devaluation is not unusual, even when care is taken to address their typical problems by using exchange rate pegging as only a part of the overall program. In Mexico, pegging was an important element of a broader program that included reduced government spending, tax reform, deregulation, privatization, and significant trade liberalization—including rapid reductions in tariffs and quotas and entry into the General Agreement on Tariffs and Trade (GATT) and later into the North American Free Trade Agreement (NAFTA).

Fiscal stabilization preceded the exchange-rate-based stabilization efforts. The history of exchange-rate-based stabilization in the Southern Cone countries had suggested that a single nominal anchor—such as the exchange rate—could be inadequate to motivate quick disinflation. Policy incredibility (that firms would not believe the exchange rate regime would persist) as well as backward indexation and nonsynchronized price-setting could lead to persistent inflation (Calvo and Végh 1992).

Accordingly, an important component of Mexico's stabilization policy was the *Pacto*. Under this government-organized accord, representatives of the business community agreed to limit price increases, the government made com-

mitments about the exchange rate and public-sector prices, and labor representatives agreed to limit wage increases.

Although there are historical exceptions, exchange-rate-based stabilization programs that also include incomes policies—like the Pacto—fairly commonly result in a specific dynamic of consumption and investment patterns, current account movements, and exchange rate pressures. The typical pattern (Calvo and Végh 1992 and Kiguel and Liviatin 1992) includes the following:

1. Despite reductions in inflation, the real exchange rate rises because some inflation remains and is not offset by nominal exchange rate movements.
2. The trade and current account balances deteriorate.
3. In the early stages of the program, capital inflows finance the excess of consumption and investment over domestic production, allowing a boom to ensue, but the inflows ultimately reverse.
4. With this reversal, the growing current account deficit can no longer be financed, and the consumption boom ends.

In recognition of this instability, a literature has developed to suggest that exchange rate pegging ought to be a temporary stabilization tool, ultimately followed by a managed float (McLeod and Welch 1991) or that, if pegging is done at all, the exchange rate crawl should be partially indexed to a measure of domestic prices (Kamin 1991). Ultimately, it has been argued, "As useful as exchange rate pegging is at the outset, it is equally important to eliminate it as soon as possible" (Dornbusch and Werner 1994, 281).

Trade and capital flows. Although Mexico's program of exchange rate stabilization *cum* incomes policy and trade liberalization contained elements particular to the country, the ensuing economic trajectory was typical of heterodox programs. Consistent with the intentions of such plans, inflation fell markedly—from 159.2 percent in 1987 to 8 percent in 1993. By the third quarter of 1994, the annualized inflation rate had declined to 7 percent.

Mexico's trade liberalization was a part of this disinflation effort. Oligopoly typifies the organization of domestic markets in Mexico, and price controls could risk product shortages. Mexico used trade liberalization to enforce price discipline—so as to hold down inflation and to lower the likelihood of product shortages.

Moreover, the country's exchange rate policy played a disinflationary role in the context of trade liberalization. The government consistently depreciated the peso more slowly than the rate of inflation—or than the difference between the U.S. and Mexican inflation rates. Consequently, as is common in exchange-rate-based stabilization programs, real exchange rate appreciation was chronic. Since real exchange rate appreciation meant that foreign products were increasingly cheaper than Mexican products, this exchange rate policy motivated domestic producers—at least of tradable goods—to resist the temptation to raise prices.

Figure 1
Real Peso–Dollar Exchange Rate

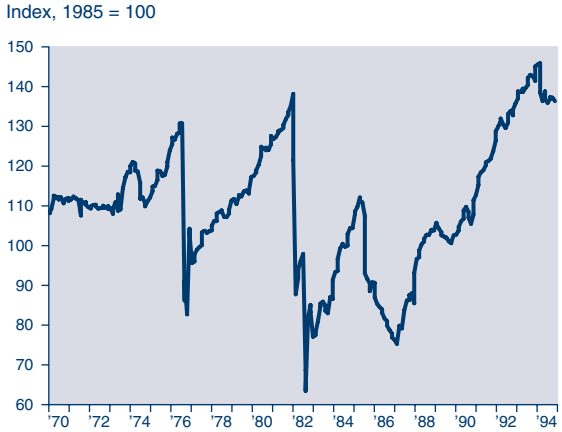
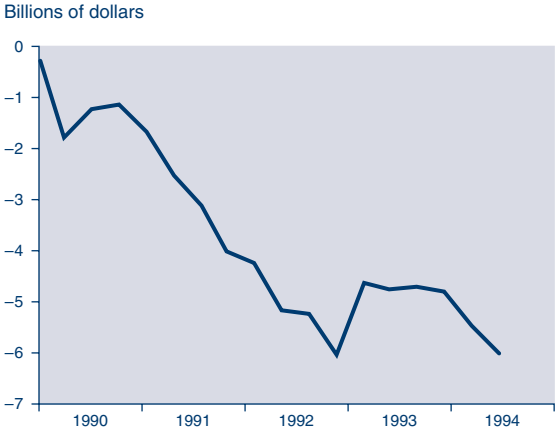


Figure 1 depicts a simple real exchange rate measure—wholesale prices in Mexico relative to those in the United States, both as measured in dollars. By the end of 1993, Mexico's real exchange rate exceeded the maximum rate that preceded Mexico's megadevaluation episodes of 1982.

Partially because of this tension between inflation and the pace of exchange rate depreciation, the nation's merchandise trade balance grew increasingly negative (*Figure 2*). As may be expected in an economy that had reoriented itself toward a market system—and had deregulated, privatized, and generally rationalized its policies toward the private sector—a significant portion of Mexico's current account deficit reflected purchases of capital goods. The increased productivity and efficiency that these purchases imply resulted in steady increases in exports. But the capital imports share of total imports was still only 16.9 percent in 1993, versus 71.1 percent for intermediate goods and 12 percent for consumer goods.

The trade and current account deficits were possible because the rationalization of Mexico's fiscal, monetary, and exchange rate policies had

Figure 2
Merchandise Trade Balance



helped stimulate large inflows of foreign investment funds through early 1994. These flows also gave Mexico the reserves it would need to defend the peso later, if exchange rate pressures required it.

Increased capital inflows are common to chronic inflation countries that introduce exchange-rate-based stabilization programs. Most of these flows (*Figure 3*) into Mexico involved portfolio investment—inflows typically for the purchase of bonds and stocks—rather than foreign direct investment. While portfolio investment permitted Mexican enterprises to fund privately owned toll roads, the recently privatized telephone company, and some manufacturing operations, the focus of this investment on the production of nontradables made inflows and outflows susceptible to concerns of devaluation risk.

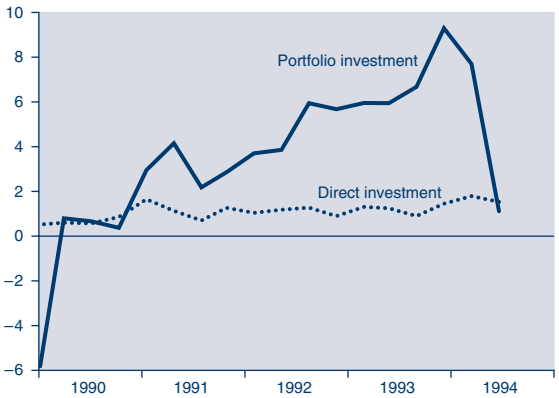
But to the extent that capital is not perfectly mobile, Mexico's chronically low and, in the 1990s, falling saving rate meant that the country's investment and growth were more susceptible to external financial events. There is much to suggest that capital flows into Mexico did not occur solely because of Mexico's policies. During the early 1990s, foreign capital began to flow into Latin America generally, despite wide differences in macroeconomic policies and economic performance among countries there. An important reason appears to be low U.S. interest rates, suggesting that increases in U.S. interest rates might have the opposite effect.[7]

Central bank policy and the financial sector. One reason tensions surfaced between Mexico's exchange rate regime and other policies is that international elements of Mexico's disinflation programs—trade liberalization, real exchange rate appreciation, and a trade deficit financed by foreign capital inflows—collectively weakened Mexico's financial sector.

Three factors converged to impose pressures on Mexico's financial sector. First, differences between the pricing performance of the nontradables and tradables sectors damaged the latter. The increased international competition held down prices in the tradable goods sectors. But even with the Pacto, prices of nontradable products, including real estate and some services, rose relative to prices of tradables. This disparity imposed profit squeezes on tradables firms because they often used nontradables as inputs, and because nontradables producers could bid up wages of workers for whom tradables firms had to compete. The direct effects of trade liberalization and real exchange rate appreciation had, of course, also imposed cost-price-squeeze pressures on some of these firms. These pressures were expressed in increasing loan defaults by tradables firms.

Second, to maintain inflows of foreign capital, real interest rates began to increase starting in 1992, even though nominal rates were falling at this time. During the early 1990s, Mexico's commercial banking system did not, at least by developed country standards, behave very competitively.[8] Spreads between cost of funds and loan rates were large. So were return on assets, return on equity, and other income statement ratios (Mansell Carstens 1993; Gruben, Welch, and Gunther 1994). Bank lending rates were typically very high by the standards of developed countries, in any case. But increases

Figure 3
Capital Account



in real rates made it particularly difficult for some firms to compete with foreign producers from countries where financial costs happened to be lower.

Third, to take advantage of the consumption boom of the early 1990s, Mexico's financial institutions had issued many more credit cards—to the wrong borrowers. By the standards of developing companies, the reporting of consumer credit histories was relatively sketchy and unorganized in Mexico. Defaults became common.

These factors converged to pressure Mexico's banking system.[9] Just between the fourth quarter of 1992 and the third quarter of 1994, the percentage of nonperforming loans rose from 5.6 percent to 8.3 percent.[10] Moreover, between December 1991 and September 1994, the ratio of high-risk assets to bank net worth rose from 51 percent to 70 percent.

Banking system problems like these take on special significance anywhere a central bank is both monetary authority and, as in Mexico, administrator of the deposit insurance system. As Heller (1991) argues, to the extent that a central bank is not only the nation's monetary authority but also is responsible for the health of the banking system, policy tensions may exist. Even though central banks are typically committed to the restraint of monetary expansion, and Mexico's is, an incipient banking crisis may create incentives to expand credit to the banking system.

It is here that the tensions expressed in the impossibility theorem appear, since it holds that free movement of capital, independent monetary policy, and a pegged exchange rate are sooner or later incompatible. Mexico followed a sterilization rule for its inflows of foreign reserves. To impose a monetary stabilization rule atop the exchange rate based stabilization, accumulations of foreign currency reserves were sterilized via offsetting reductions in domestic credit the central bank created for the financial system. Conversely, outflows of foreign currency reserves were sterilized through offsetting increases in domestic credit.[11,12]

Recall that a central bank can always maintain a pegged exchange rate, but sometimes the price is otherwise undesirably tight monetary policy. Outflows of foreign currency reserves, even if for purely political reasons, can signal that a monetary contraction or interest-rate increase is in order.[13] Such policies can be inconsistent with the expansion of domestic credit as an offset to capital outflows, even if the policy is purely an act of sterilization.[14]

The currency configuration of Mexican short-term debt

Mexico has simultaneously issued short-term, peso-denominated debt (*cetes*) and short-term dollar-indexed debt (*tesobonos*), but as 1994 progressed, Mexico radically altered the currency configuration of its short-term debt so as to strengthen the peso. In January 1994, the dollar value of cetes outstanding was $12.9 billion, compared with $302 million in tesobonos. By November, cetes outstanding had fallen to $7.27 billion, while tesobonos had risen to $12.9 billion.

An interesting characteristic of these debt issues, as demonstrated econometrically (Dornbusch and Werner 1994), is that the changes in spreads between their interest rates are not affected by changes in factors normally associated with exchange rate expectations. Dornbusch and Werner (1994) argue that changes in spreads between the interest rates of cetes and tesobonos are not explained by changes in the real exchange rate or in Mexico's trade balance because the government managed the composition of its domestic debt so as to respond to cost differentials. That is, as exchange rate risk rose, Mexico shifted its composition of short-term debt toward tesobonos and away from cetes. The authors argue that this shift reflects government responses to cost differentials. As rates on tesobonos fell relative to rates on cetes, the government replaced cete issues with tesobono issues.

If one advantage of a shift toward tesobonos was to save on interest expenses while gaining foreign exchange by selling debt to foreigners, it was not the only advantage. Mexico's increased issuance of tesobonos may also be seen as making its exchange rate regime more credible by imposing a clear and obvious fiscal penalty for devaluation.

Ize and Ortiz (1987) note that devaluation is tantamount to a default on domestic debt because, by raising the price level, the government erodes the debt's real value. Accordingly, a large overhang of domestic debt may be seen as a motivation to devalue, particularly when the debt is held by foreigners.

While this motivation may exist when a nation's domestic debt is denominated in the home currency, the motivation erodes if, as with the tesobonos, the debt instrument is indexed to the dollar. A shift out of cetes and into tesobonos is a shift out of an instrument for which outstanding real debt falls with devaluation and into an instrument for which devaluation means a real debt increase. This statement

holds whenever the subsequent rate of inflation does not match or exceed the rate of devaluation by the time the debt matures.

The tesobono shift's role in enhancing credibility that the exchange rate regime will persist may be indirectly measurable. Insofar as agents recontract for higher wages or higher purchase or selling prices now in anticipation of a generalized bout of price increases—so that present prices reflect expectations of future price increases—and insofar as a devaluation may be seen as triggering a future generalized bout of price increases, the implications of a shift into tesobonos as a commitment technology for the exchange rate regime may be expressed in current price increases.[15]

Preliminary econometric research by David Gould shows that, even when monetary base growth and other factors linked to inflation are included in a model of Mexican consumer price changes, a negative and significant relation exists between consumer inflation and the share of Mexican short-term debt that is indexed to the dollar. That is, with this credibility enhancement in place, the market seems to reduce its expectation of the devaluation and so of the inflation that typically follows devaluation.[16] It does not seem unreasonable to conjecture that this credibility enhancement could also have been seen as a potential enhancement for transitory monetary independence.

Putting the pieces together

The implications of the general dynamics of heterodox exchange-rate-based stabilization programs, of the role of domestic credit expansion in addressing systemic bank crises and in triggering currency collapses, and of the use of tesobonos as a commitment technology become more dramatic when we consider the roles they played in Mexico in 1994.

Recall that typical patterns of exchange-rate-based stabilization programs include falling inflation, rising real exchange rates, consumption booms, capital inflows in the early stages that fund increasingly negative balances of trade and current account and, finally, capital outflows that ultimately induce currency collapses. Recall also that a typical characteristic of a currency collapse is not the impossibility of maintaining a pegged exchange rate, but a policy priority rearrangement in which the exchange rate is subordinated.

Finally, note that the intention of this article is not only to explain why the choices were made that triggered the devaluation, but to explain why its aftermath was explosive despite prior claims that "the Mexican government would not lose credibility from a devaluation, because it would be recognized as a constructive response to a crisis."[17]

I noted earlier that one reason Mexican bonds and stocks attracted U.S. and other foreign investors was low U.S. interest rates. During first-quarter 1994, U.S. monetary policy began to tighten, raising U.S. interest rates and attracting capital back to the United States. While the increase in U.S. rates signified that factors pushing capital toward Mexico were diminishing, political events in Mexico weakened the country's pull effects for capital.

Chiapas rebels may not have threatened the nation's stability, but the assassination of Mexican presidential candidate Colosio in March 1994 was another matter to investors. After rising earlier in the year, reserves fell profoundly just after the assassination but stabilized in April. To hold foreign capital in the country, Mexico raised interest rates significantly, signaling that exchange-rate preservation remained important. But U.S. interest rates were also rising, and they continued to do so throughout the year. The exchange rate moved toward the weak edge of the band but remained within it.

It is in this context of rising U.S. rates at a time when increasing financial problems offered motivations to lower or at least hold Mexican rates that the value of the tesobonos as a commitment technology can be appreciated. Instead of offering a commitment technology based on the accumulation of larger foreign currency reserves to defend the peso, when real rates were already at high levels, the issuance of tesobonos might be thought a reasonable substitute, at least temporarily.

Figure 4

Real Central Bank Domestic Credit To Commercial Banks

Millions of 1990 new pesos

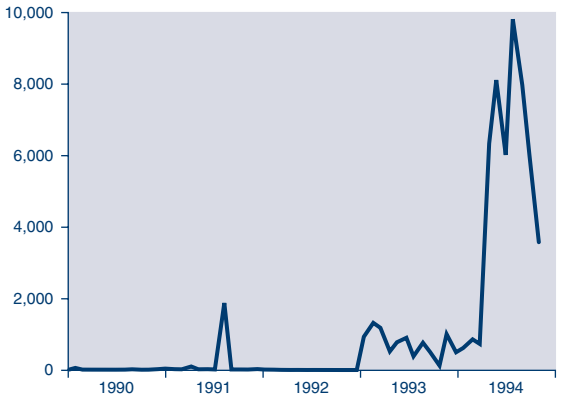


Figure 5
Tesobonos and Cetes Held by the Public

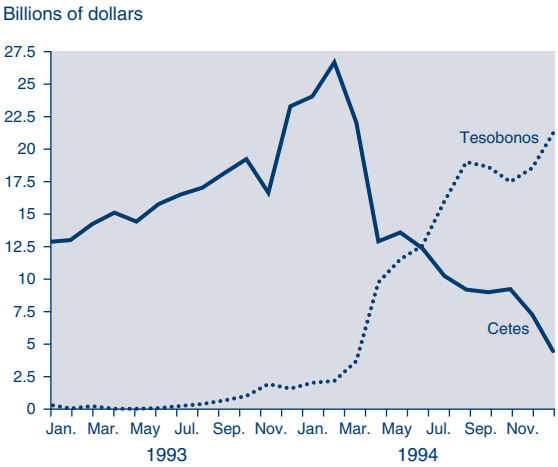
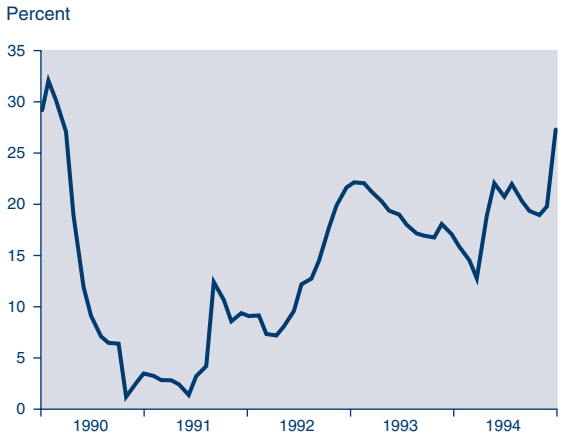


Figure 6
Real Interest Rate in Mexico



When the Colosio assassination triggered a capital outflow, Mexico sterilized by raising domestic credit (*Figure 4*). At the same time, Mexico stepped up its substitution of dollar-indexed tesobonos outstanding for peso-indexed cetes outstanding (*Figure 5*). By midyear, tesobonos outstanding began to exceed cetes. The substitution increased through the rest of the year.

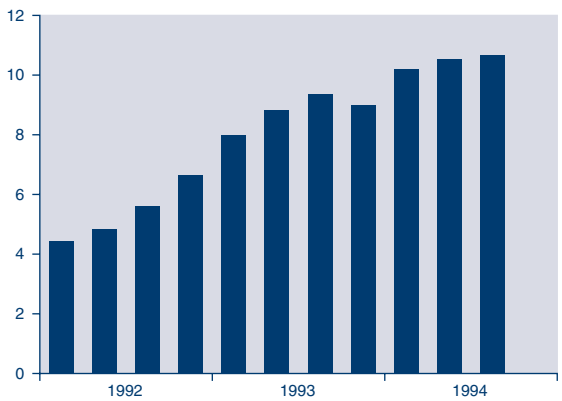
In the third quarter, Mexico began to relax its interest rate pressures, as can be seen from Figure 6. Interest rates remained considerably higher than they had been at the beginning of the year. But they were not high enough to restore reserves to the levels of the first quarter—not, at least, when U.S. rates were also rising.

Nevertheless, reserves remained relatively stable during the third quarter. One reason may be that, as the summer ensued, it became more obvious that substitute Institutional Revolutionary Party presidential candidate Ernesto Zedillo was likely to defeat the other candidates, whose abilities or policies may have inspired more investor uncertainty about future growth. Then, in August, he did win. But Gould's (1994) results on the negative influence of tesobonos' share of total short-term debt on inflation rates suggest than an exchange rate commitment technology also helped stabilize foreign currency reserves. The third quarter ended with foreign currency reserves as high as those with which it had begun.

As 1994 ensued, the differential between Mexican and U.S. interest rates began to fall, much as one might expect, other things being equal, as a reasonable policy response in the face of mounting problems in the Mexican financial system (*Figure 7*).[18] Nominal cetes rates fell absolutely in August and remained below their spring and summer highs until the devaluation.

In the fourth quarter, another political event preceded capital outflows from Mexico, but a concurrent economic event makes interpretation difficult. After the assassination of Institutional Revolutionary Party official Carlos Francisco Ruiz Massieu, his brother was appointed to investigate the case; in November he resigned, alleging that powerful officials were stymieing his investigation. Meanwhile, on November 15 the Federal Open Market Committee of the U.S. Federal Reserve System met and decided on policies that would lead to a 75-basis-point increase in the federal funds rate, its most restrictive monetary policy action since 1990.

Figure 7
Past-Due Loan Ratio



NOTE: Data are quarterly.

Table 1
1995 Tesobonos Maturity Schedule
(Billions of U.S. dollars)

January	$ 3.626
February	3.487
March	3.159
April	1.858
May	2.723
June	1.907
Total	$16.760

In sterilizing the subsequent outflow of foreign capital, Mexico's central bank again increased domestic credit to the banking system.[19] Mexican interest rates were not pushed up sufficiently to maintain reserves.

Perhaps as a result of the fiscal implications that the large overhang of tesobonos offered in the event of a devaluation, the exchange rate continued to show signs of credibility.[20] But this tesobono commitment technology had been imposed in a period of increasing risk to the financial system and of the additional trade balance pressures partially induced by the commencement of NAFTA in January 1994. Taken collectively, these factors meant that Mexico could be risking a financial crisis if it devalued the currency and allowed interest rates to go where they would, or if it defended the currency by raising interest rates.

On December 20, the tesobono overhang that had suggested exchange rate credibility now signified financial market as well as currency collapse. When the Mexican government announced that the peso would move from 3.47 pesos per dollar to 3.99, it also announced that the exchange rate pegging regime, in which the peso would devalue against the dollar at a rate of 0.0004 pesos per day, would persist. The band would simply be lowered.

But instead of settling markets, the announcement incited massive capital flight. Large increases in interest rates ensued. Perhaps the fiscal implications of the tesobono overhang, with a maturity schedule in which the value of tesobonos falling due within the first six months of 1995 exceeded the value of Mexico's foreign exchange reserves, were calculated by financial markets (*Table 1*). But given the moderate magnitude of the initial announced devaluation, the pure act of abridgment of such a commitment seems to have played an important role in and of itself.

Conclusion

The chief problems Mexico faced in 1994 were that the controlled rate of depreciation of the peso was inconsistent with the persistent inflation rate differential between the United States and Mexico, that capital outflows drew down foreign exchange reserves that Mexico was using to defend the peso, and that, in the conflict between greater monetary tightness to support the exchange rate and less tightness to avoid further financial-sector problems and a downturn in the economy, the latter won out. While Mexico wished growth, it was caught in an episode of U.S. monetary tightening that only de facto monetary independence would have permitted it to avoid following with a vengeance—and in financially destabilizing episodes of political unrest.

Despite evidence that some monetary independence was available transitorily, as the short run grew into a longer run, independence and dependence collided with a result long since posited as the impossibility theorem. But while these factors are consistent with an ensuing devaluation, they alone are not consistent with the explosive nature of Mexico's financial crisis in the wake of the initial devaluation.

The explosive nature of the crisis seems to have been linked to reactions to the term structure and volume of Mexico's short-term dollar-indexed debt, even though there is little evidence to suggest that the tesobono debt was seen as problematic before the devaluation and that it had served as a positive commitment technology.[21] That the tesobono maturity schedule signified obligations in early 1995 that were considerably in excess of Mexican dollar reserves to cover them may have triggered the anticipation of a financial musical chairs game in which each investor began to fear that her or his tesobono would be the one left out of convertibility.

Notes

1 *Wall Street Journal* (1994).

2 Fidler and Bardacke (1994).

3 See, for example, Torres and Campbell (1994), Fidler and Bardacke (1994), and Tricks (1994).

4 For a discussion of this last, see Lustig (1995).

5 How does raising interest rates affect the exchange rate? Rising Mexican interest rates inspire foreigners to buy Mexican financial assets—triggering capital inflows, increasing the demand for pesos, and so bidding up the exchange rate. As foreigners trade dollars for pesos to buy peso-denominated assets or simply buy dollar-denominated assets from the Mexican financial sector, Banco de México accumu-

lates dollar reserves. If pressures to devalue arise, Banco de México can use its dollar reserves to buy up pesos—raising their dollar price. Also, squeezing monetary growth and raising interest rates lower Mexican inflation. Insofar as dollar prices of Mexican goods rise faster than dollar prices of U.S. goods, both Mexican and U.S. buyers are discouraged from buying Mexican products and encouraged to buy U.S. products. The resulting trade deficit increase means declining demand for pesos—as fewer Mexican products are bought—and rising demand for dollars—as more U.S. products are bought. Pressure arises to devalue the peso—which lowers the dollar price of Mexican products and raises the peso price of Mexican products, erasing the deficit. A tight monetary policy that includes raising interest rates dampens Mexican inflation, squeezes the differential between Mexican and U.S. inflation, and lowers pressure to devalue.

[6] Gruben, McLeod, and Welch (1995) show that three-month U.S. Treasury bill rates do not Granger-cause and are not Granger-caused by three-month Mexican Treasury bill (*cetes*) rates and that three-month U.S. Treasury bill rates do not Granger-cause and are not Granger-caused by Mexican Brady par bonds. However, thirty-year U.S. Treasury bonds do Granger-cause Mexican Brady par bonds, which, it should be noted, may be seen as long-term bonds. These results suggest that financial integration between Mexico and the United States can be significantly abridged in the short term but not in the long run.

[7] For a more complete discussion of external factors leading to such flows, see Calvo, Leiderman, and Reinhart (1993); Chuhan, Claessens, and Mamingi (1994); and Dooley, Fernandez-Arias, and Kletzer (1994). Most foreign capital flowing into Latin America did go to Mexico, however.

[8] With the exception of union-owned Banco Obrero and U.S.-owned Citibank, the entire Mexican commercial banking system was nationalized in 1982. With a series of consolidations, the original fifty-three nationalized banks were pared to eighteen. These eighteen institutions were privatized, one by one, in 1991 and 1992.

[9] For a fuller development of the links between Mexico's banking problems and the subsequent exchange rate crisis, see Calvo and Mendoza (1995).

[10] In the United States, once a loan goes into arrears, the entire remaining loan balance is considered in arrears. In contrast, Mexico's calculation procedure does not consider the entire remaining loan balance to be in arrears. For example, in Mexico, if a loan is three months in arrears, only the balance that had been contracted to be paid during those three months is calculated as in arrears. Any loan balance scheduled to be paid thereafter is not yet calculated as in arrears. Other things being equal, U.S. protocols would sometimes result in higher past-due loan ratios than Mexican protocols.

[11] Banco de México (1995, 64) notes that "the increase in domestic credit in 1994 occurred in response to reserve declines" and that reserves "did not fall because domestic credit was expanded" [author's translation].

[12] While I am presenting a case for the possibility of separation between reserve outflows and the domestic credit creation that is implied by sterilization, it is true that some argue that when central banks sell reserves, they must sterilize automatically.

[13] Kamin and Rogers (1995) offer econometric evidence to suggest that when interest rates did rise, they rose only moderately less than could be predicted by the authorities' standard reaction function. Kamin and Rogers argue that, to have maintained the peg, the authorities would have had to intensify their responses to exchange market developments. That is, policymakers would have had to alter their reaction regime, and they would have had to at a time when concerns for the health of the banking system would have suggested a relaxation of monetary policy.

[14] While the merits and liabilities of currency boards are a subject beyond the scope of this article, one discipline they impose is that when foreign exchange reserves flow out, the resulting reduction in the stock of money is not offset. Although such boards may be seen as having significant liabilities, Argentina's peso (which is disciplined by a currency board) has maintained its nominal value over the past two years while Mexico's has not.

[15] Brown and Whealan (1993) offer econometric evidence to suggest, for example, that present oil prices reflect agent expectations of futures prices.

[16] Lustig (1995, 379) notes that "this dollarization of internal debt probably explains the surprising stability of international reserves before such adverse events as the increase of foreign interest rates and domestic political unrest" [author's translation]. Moreover, Banco de México (1995, 69) states that "the issuance of tesobonos was carried out in order to reduce exchange market pressures" [author's translation].

[17] Werner (1994, 310).

[18] Recall that inasmuch as a central bank can always preserve a pegged exchange rate through a sustained high interest rate or a contraction in monetary base, interest rates insufficient to prevent declining reserves suggest that other policies must dominate a commitment to a pegged exchange rate. Garber and Svensson (1994, 29) note that one of these policies may be "the preservation of solvency of a banking system."

[19] The capital outflows were not well-known, however, and a number of analysts have complained that something kept Mexico during the latter portions of 1994 from releasing data on central bank holdings of foreign reserves.

[20] Interest rates typically reflect nervousness about devaluations. Consider, for example, the movement of

yields on the twenty-eight day cetes auction for the following dates: November 9—13.49 percent, November 16—13.45 percent, November 23—13.95 percent, November 29—13.85 percent, December 7—13.30 percent, and December 14—13.75 percent.

[21] Calvo and Mendoza (1995) and Sachs, Tornell, and Velasco (1995) address other aspects of the sudden and explosive nature of Mexico's financial crisis that clearly deserve attention. Calvo and Mendoza argue that this phenomenon reflects, among other things, a trade-off between diversification and information that investors face when information is costly to acquire. As investment opportunities expand across countries, the payoff to purchasing information about a particular country declines. It becomes rational for investors to become sensitive to even "small" bad news, especially when it follows previous bad news, even if none of the news is related to fundamentals. In sum, the reduced incentives to acquire much information about Mexico in particular and Latin America in general motivated a herd behavior that triggered the tequila effect.

Sachs, Tornell, and Velasco (1995) argue that, while real disequilibria and reserve erosion lay the groundwork for the crisis, the timing and magnitude of the crisis came from a self-fulfilling panic after the government ran up its short-term tesobono debt and ran down gross reserves. That is, like Calvo and Mendoza (1995) and those cited at the beginning of this article, Sachs, Tornell, and Velasco (1995, 7) do not believe that the crisis was fully consistent with fundamentals. Instead, they conclude, "the panic was self-fulfilling in that expectations of a run on both pesos and teso-bonos by other agents led each individual investor to engage in the same kind of speculative behavior."

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