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Educational Attainment and Border Income Performance

Thomas M. Fullerton Jr.

Large numbers of the new entrants to the state labor market will find it difficult to obtain above-average salaries. For the border zone collectively, nearly $3.6 billion in forgone income results from a secondary school dropout rate that exceeds the state average.

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As Texas enters the new millennium, two important trends are influencing regional economic performance. Demographically, the state population is rapidly becoming more ethnically diverse, but many of the migrants from Mexico and Central America have not graduated from high school. Simultaneously, the majority of new jobs continue to emerge in service segments of the Texas economy. Together, these developments imply that large numbers of the new entrants to the state labor market will find it difficult to obtain above-average salaries. High-salary positions are available in service sectors, but they typically require education beyond that of a high school degree.

It is widely recognized that income is positively correlated with education. The primary link is through the improvement in human capital that results from the educational process (Becker 1993). Education is not the only means by which human capital can be improved, but it is one of the most important. As economies become more advanced, education's role in enhanced worker productivity is fairly pronounced (Becker 1993, Welch 1970). Given the widespread evidence regarding education and income, greater schooling achievement is probably essential if the border region is to approximate the higher earnings observed elsewhere in Texas.

The importance of potential income gains to border economies is readily apparent. For example, population growth was strong during the 1990s in four of the border metropolitan areas: Brownsville, El Paso, Laredo, and McAllen. Accompanying that growth was substantial new commercial activity, including large increases in retail sales (Woods and Poole Economics, 2000). Per capita income performance in these markets was less encouraging. By the end of the decade, all four ranked 305th or worse out of the 318 metropolitan areas defined by the Census Bureau. These urban areas also lagged the rest of the nation in educational attainment (Woods and Poole Economics, 2000).

Recent work in regional economics has attempted to quantify the relationship between education and income at both state and county levels (Ashenfelter and Krueger 1994, Domazlicky et al. 1996, Sloboda 1999, Levernier, Partridge, and Rickman 2000). These techniques can be used to estimate income gains associated with different schooling levels across Texas. As has been documented elsewhere, one challenge facing the state economy is unequal income performance, especially in border areas (Sharp 1998, Sáenz 1999, Fullerton 2000). In 1998, for instance, per capita incomes for Brownsville, El

Paso, Laredo, and McAllen ranged between 50 and 65 percent of the state average of $25,370 (Diffley 1999). Overcoming those gaps will undoubtedly take years and depend in large measure upon increased schooling in each metropolitan economy. The latter assertion implicitly assumes that higher paying jobs also materialize via regional demand for labor, but such a pattern is hardly unique during the postwar history of the United States. In fact, the absence of skilled workers may serve to discourage investment flows from companies producing advanced products.

To shed light on the links between education and income performance, socioeconomic data were collected for all 254 Texas counties from the 1990 census and from the Regional Economic Information System, both published by the Department of Commerce. Cross-section econometric modeling was used to estimate the relationship between per capita income levels and the various regressors. The regression output was then used to simulate the impact of higher graduation rates for border counties in the state.

LITERATURE REVIEW

Several studies in recent years have examined the relationship between education and regional income performance, but there have been relatively few efforts to empirically quantify this for Texas border counties. Given this, it is important to review evidence reported for other regions to provide a backdrop against which the current effort can be compared. A limited synopsis of similar studies follows.

Below-average graduation rates occur in regions other than the border with Mexico. Some of the most recent research on this topic has been compiled for rural areas of the nation where more than 14 percent of all families are estimated to have lived in poverty in 1990 (Levernier, Partridge, and Rickman 2000). Domazlicky et al. (1996) and Sloboda (1999) concentrate on southern portions of Missouri and Illinois, while Rickman (1993) examines a southern section of Georgia. These studies offer broad support for the association between human capital and labor productivity documented in other papers (Ashenfelter and Krueger 1994, Miller, Mulvey, and Martin 1995). More specifically, failure to complete high school is associated with statistically significant negative impacts on per capita incomes in all three regions considered in those states. In each study, the impacts total several hundred dollars on a per capita basis and several hundred million dollars on a regionwide basis.

In Texas, issues involving education and income are increasingly important (Crowder 1999). The population is rapidly becoming more ethnically diverse and less educated (Murdock 1999). Concurrently, the number of well-paying blue-collar jobs is falling, leaving service sector positions as the best new-job opportunities for Texans (Diffley 1999, Nakamura 2000). Implications for the state labor market are fairly straightforward. Access to high-salary positions will more frequently require education beyond a high school degree.

As state labor markets in Texas have become more service-oriented, regional income distributions have suffered at the expense of border areas (Sharp 1998, Sáenz 1999). Reversing these trends is feasible but is likely to be a long-term process because it will entail improving educational performance at a variety of levels over several years (Buchner 1999, Dittmar and Phillips 1999). Failure to address the task quickly can only result in deepening problems as information-age labor markets become increasingly competitive (Cox and Alm 1999).

Despite the absence of immediate headline results from such endeavors, addressing the border region education shortfall will generate numerous economic benefits. Many of these improvements will come as positive externalities (Taylor 1999). First-best externalities include higher wages, higher real estate values, more rapid employment growth, and greater firm entry rates. Second-best externalities from increased regional educational attainment include higher community income levels, lower crime rates, stronger tax bases, lower per capita social transfer payments, reduced use of public health systems, lower poverty rates, and greater attraction of extraregional private sector investment.

While labor demand fluctuations will influence regional income opportunities, there is widespread agreement that raising aggregate human capital stocks improves economic wellbeing. In some cases, econometric techniques have provided empirical estimates of the dollar value that can be generated by improving a geographic area's educational attainment. Preliminary estimates also corroborate this type of relationship in a cross section of states in Mexico using 1990 census data for that economy (Messmacher Linartas 2000). To date, such an effort has not been attempted for the border counties of Texas. The following sections summarize the procedures and results of such an exercise.

Table 1
Variable Names and Definitions

Variable	Definition
PCINC	County per capita income level in 1990
HSDR25	Percentage of adults 25 or older that did not finish high school
HSGR25	Percentage of adults 25 or older that attended some college
COGR25	Percentage of adults 25 or older that graduated from college
POPGT65	Percentage of county population age 65 or older in 1990
POPLT18	Percentage of county population age 18 or younger in 1990
PCTENGL	Percentage of monolingual English households for county in 1990
PCTBLNG	Percentage of bilingual households for county in 1990
PCTSPNH	Percentage of monolingual Spanish households for county in 1990
URBAN	Dummy = 1 if 1990 population exceeds 599,999; zero otherwise
BORDER	Dummy = 1 if county is adjacent to Mexico; zero otherwise

DATA AND METHODOLOGY

Information on years of formal schooling and degrees completed was collected for all 254 Texas counties. These data are available in the social and economic characteristics information files of the Census Bureau. Estimates of 1990 per capita income levels were also assembled for each county, using revised estimates from the Regional Economic Information System files of the Bureau of Economic Analysis.

Previous regional income performance studies used per capita income estimates as the dependent variables in their equation specifications (Rickman 1993, Domazlicky et al. 1996, Sloboda 1999). Explaining variations in border per capita incomes requires expanding the independent variable vector beyond that used in those earlier efforts. This is because Texas' demography is more heterogeneous than that of other regions of the national economy. Among other things, household sizes along the border tend to be much larger than those in the rest of Texas and the nation (Murdock 1999, Schick and Schick 1991). Limited English language skills in border counties also have been shown to affect earnings performance (Dávila and Mora 2000, Mora and Dávila 1998).

Right-hand-side regressors include the percentage of adults in each county who are 25 or older and dropped out of high school, are 25 or older and graduated from high school and attended some college, and are 25 or older and graduated from college. Increases in the percentage of high school dropouts are expected to depress county incomes, while increments in the percentages of both graduate categories are likely to improve earnings. Also included among the explanatory variables for each county are the female labor force participation rate, the percentage of the population 65 or older, and the percentage 18 or younger. Increases in female labor force participation should raise county incomes. An ambiguous relationship exists between the proportion of retirement-age adults and income. Many of these persons drop out of the labor force but simultaneously begin receiving sizable transfer payments. As the percentage of youth increases in a county, per capita income is likely to decline because individuals younger than 18 generally do not work or hold only part-time positions.

Several demographic explanatory variables that measure each county's language characteristics are also tested. English and bilingual language skills are expected to increase per capita incomes, while monolingual Spanish skills are likely to reduce earnings in U.S. labor markets. Geography and industry mix can also affect income performance (Ciccone and Hall 1996, Glaeser 1998). Three qualitative variables are used to capture these possibilities. An urban dummy for large counties with 1990 populations of 600,000 or greater is expected to coincide with higher earnings. A second dummy for smaller metropolitan counties with populations ranging from 200,000 to 599,999 is anticipated to be associated with moderately higher incomes. Finally, a border dummy is likely to correspond with lower earnings performance. Using the approach suggested by Peach and Adkisson (2000), border counties are defined as those immediately adjacent to Mexico. Variable mnemonics are presented in Table 1.

Least-squares regression techniques are used to estimate the various specifications tested. Because of large divergences in Texas county populations, residuals are tested for heteroskedasticity. For example, Harris County had more than 2.8 million residents in 1990; at the other extreme, Loving County had a population of 106 in 1990.

The cross-section econometric estimates quantify the relationships between education and regional incomes across the state. County per capita incomes are used as the dependent variable. This choice is made because per capita income estimates include wage and salary disbursements as well as proprietor earnings, dividend payments, and retirement transfers, all of which are positively correlated with human capital. The regression estimates are used to examine border area impacts of changes in educational attainment. Because of minor reporting and nonreporting errors, census data aggregates for Texas county school achievement provide imperfect measures of actual school district performance (Murdock 1999). The county census estimates do offer, however, one means by which to analyze potential gains. Impacts can also be calculated for more broadly defined regions of the state, such as the Panhandle, the Hill Country, the Metroplex, and so forth. Results for the border regions are of particular interest, given existing economic and demographic differences between these counties and Texas as a whole.

One possible weakness of the approach selected should be pointed out. As with earlier studies of other regions, the methodology employed implicitly assumes the quality of schooling is equal across all 254 counties in the state. This is probably not the case in a state as large and diverse as Texas. Whether such differences significantly weaken the results shown below is not addressed in this article. This topic may, however, merit additional attention (see Becker 1993).

Additionally, greater investment in human capital generates regional income gains only in those cases in which labor force quality improves and local labor supplies increase. Graduates in any region commonly choose to pursue jobs in other markets. Recently, the entire graduating class from one technical institute in El Paso accepted out-of-state jobs with a nationally prominent silicon chip maker (Mrckvicka 2000). In extreme form, labor mobility could lead to situations in which improved educational attainment in border counties would not raise aggregate income. This eventuality is not examined in this article but may warrant more careful consideration in subsequent research efforts.

A final potential problem with the parameter estimates is that returns to education will also vary with the characteristics of labor demand at the county level. With regard to the Texas border region, many counties have relatively large government sectors with union membership contingents, such as the Border Patrol and military installations. Because the extent and nature of union membership and other county-level labor demand variables are not controlled within the various equation specifications, the education parameter estimates are likely biased. In the case of labor unions, the biases may even be negative. A potential step to consider in future work is the inclusion of 1980 and 2000 census results and the use of county fixed effects. Doing so could help sort out county-specific effects from the general relationship between education and income.

Table 2
White Procedure for Heteroskedasticity Test with Cross Terms

Equation	Computed test statistic, 254·R^2	Chi-square probability
T3	89.303	0.004

EMPIRICAL RESULTS

Heteroskedasticity is present in the sample. Chi-square results for the majority of the initial ordinary least squares estimates point to the presence of systematically unequal residual magnitudes with test statistics that exceed 5 percent critical values. Table 2 reports the White (1980) heteroskedasticity test result for the equation version used in calculating the per capita, county, and regional income impacts. Accordingly, the covariance matrix for the model reported in Table 3 has been reestimated using White's (1980) procedure.[1] A discussion of general empirical characteristics follows. The basic equation form tested is:

$$PCINC_i = B_0 + \mathrm{Sum}(B_j X_{ji}) + e_i,$$

where i = 1,2,3,…,254 for each of the counties in Texas and j = 1,2,3,…,k depending on the number of independent variables employed. With respect to the statistical output reported for the model used in the income impact simulations below, k = 10.

Overall statistical traits of the various specification estimates are favorable. All the adjusted R-squared coefficients of determination are greater than 40 percent, a fairly strong goodness-of-fit for heterogeneous cross-sectional data. In many cases, some of the computed t statistics fail to satisfy the 5 percent statistical significance guideline for type I errors. All the model F statistics for joint significance, however, surpass their respective 1 percent critical values. Although detailed statistical outcomes for only one model specification are shown in Table 3, all the various log likelihood estimates for each specification format are clustered closely together.

The central hypothesis being tested is that Texas border county incomes are affected by educational attainment in a manner similar to that of other regions of the United States. To examine that possibility, three different human capital variables are included in the equation shown in Table 3. The first of these regressors is the percentage of adults 25 years or older in each county who failed to graduate from high school. As in previous studies, this variable is associated with a negative impact on county per capita income. Its computed t statistic does not quite reach the 5 percent significance level, but

Table 3
County Per Capita Income Regression Output Using White Procedure

Variable	Coefficient	Standard error	*t* statistic	Probability
Constant	7059.761	7670.249	.920408	.3583
HSDR25	–103.8998	58.35719	–1.780412	.0763
HSGR25	143.1058	93.50420	1.530475	.1272
COGR25	144.4096	63.58679	2.271063	.0240
PCTBLNG	–81.49451	55.46333	–1.469341	.1430
PCTENGL	–36.09971	53.79102	–.671110	.5028
PCTSPNH	34.97376	158.0863	.221232	.8251
POPGT65	200.7548	75.29035	2.666408	.0082
POPLT18	283.9618	113.4229	2.503566	.0130
URBAN	2837.218	1197.363	2.369556	.0186
BORDER	–3179.338	1262.435	–2.518417	.0124

R-squared	.433397	**Dependent variable mean**	14711.18
Adjusted R-squared	.410080	**Dependent variable S.D.**	3385.411
S.E. of regression	2600.206	***F* statistic**	18.58721
Sum of squared residuals	1.64E+09	***F* statistic probability**	.000000
Log likelihood	–2352.078	**Observations**	254

NOTE: Sample data from 1990 census.

its parameter magnitude, negative $104, is close to those reported for other regional economies.

The next right-hand-side variable used in the models is the percentage of adults 25 or older in each county who graduated from secondary school and attended some college but without obtaining a bachelor's degree. As hypothesized, this variable generates a positive coefficient. In the model detailed in Table 3, however, the *t* statistic for this parameter falls below the 5 percent significance threshold. Nevertheless, results across a variety of specifications not reported are sufficiently consistent to indicate that graduation from high school and partial college attendance raise county per capita incomes in Texas (see McCloskey and Ziliak 1996).[2] The third regressor is the percentage of adults 25 or older who successfully completed at least a four-year college degree. This regression parameter is greater than zero and satisfies the 5 percent type I error criterion.

Although most labor and other market transactions in the United States are conducted in English, substantial numbers of border residents in Texas speak Spanish as their primary language. Because communication skills can have important implications for personal income, the results shown in Table 3 are surprising. None of the three language skill independent variables obtains a computed *t* statistic that satisfies the standard type I error criterion.[3] Separate research using Census Bureau public-use microdata samples has uncovered patterns of language–income interplay (Dávila and Mora 2000, Mora and Dávila 1998). Presumably, some of the language impacts are accounted for by the border county qualitative regressor in Table 3, but experimentation with interaction specifications (Pindyck and Rubinfeld 1998) did not yield significant results.[4]

The parameter estimate for the percentage of the population over age 65 carries a positive sign and, as in Domazlicky et al. (1996), has a *t* statistic that exceeds the 5 percent critical value. Unexpectedly, the coefficient estimated for the percentage of the county population aged 18 years or younger is greater than zero. While Sloboda (1999) reports a similar result for Illinois, the parameter reported in Table 3 is also statistically significant. Although an increase in the number of people in this age group raises a county's dependency ratio, for the data sample used in this study, it is also associated with a per capita income increase of $284.[5]

As in numerous other regional economic articles, in this study residency in urban areas is associated with higher per capita incomes. Large metropolitan counties with 1990 populations of 600,000 or greater exhibit per capita incomes roughly $2,840 above those for other

Texas counties. This dummy variable coefficient has a significant *t* statistic.[6] In contrast to the urban effect, border county per capita income falls below other state regions by approximately $3,200. As with the large urban dummy variable, the border parameter also surpasses the 5 percent significance threshold.

Tables 4, 5, and 6 use the county per capita income regression coefficients to calculate the impact on per capita income of raising each county's educational achievement to the 1990 state averages for high school graduation, high school graduation plus some college attendance, and college graduation. Because its respective graduation rates are higher than those of Texas as a whole, Brewster County, in the Big Bend region of the state, is excluded from the analysis. Similarly, Jeff Davis County, also located in far West Texas, is excluded from the calculations in Table 6 because its college graduation rate exceeds that of the rest of the state.

The implied effects of reducing secondary school noncompletion are striking (*Table 4*). The single largest gain—$5,760 per person—is obtained for Starr County in the Rio Grande Valley. Other things remaining equal, raising its high school graduation rate to the state average would permit Starr County to more than double its 1990 per capita income. The aggregate impact for the county would be to raise total 1990 personal income by more than $210 million.

In nearby Hidalgo County, the implications are similarly impressive. Income per person rises by more than $3,600 in the simulation exercise. Due to the county's relatively large population, that translates to a countywide estimate in excess of $1.26 billion. Aggregate personal income gains of more than $400 million also result for Cameron, El Paso, and Webb Counties. For the border zone collectively, nearly $3.6 billion in forgone income results from a secondary school dropout rate that exceeds the state average. Given the limited tax bases of the border region, the benefits of improving high school graduation rates are clear.

Table 5 illustrates the potential improvements of increasing to the state average the proportion of border county residents that not only graduate from high school but also attend some college. While less dramatic than the income gains shown in Table 4, the results are still noteworthy. On a per capita basis, the largest impact exceeds $2,000, once again for Starr County. For the entire border region, the per person income improvement is estimated at $711. The $413 million aggregate figure for Hidalgo County represents the largest total gain of any single jurisdiction in the sample. Increasing to the state average the percentage of border county residents who attend college following high school would raise total income by $975 million.

Table 6 details the implied gains from raising border county college graduation rates to a level commensurate with Texas as a whole. The income impacts are again substantial on a per capita as well as an aggregate basis. Estimates range from $736 per person in El Paso County

Table 4
Implied Income Losses Due to High School Noncompletion

County	Per capita impact	Aggregate impact (in millions)
Brewster	Not calculated	Not calculated
Cameron	$3,143	$ 744.7
El Paso	1,195	643.8
Hidalgo	3,627	1,262.5
Hudspeth	3,413	9.2
Jeff Davis	370	0.7
Kinney	2,261	6.6
Maverick	5,177	170.4
Presidio	4,011	24.5
Starr	5,760	210.2
Terrell	825	1.1
Val Verde	2,276	80.1
Webb	3,456	413.8
Zapata	3,129	26.3
Border zone	$2,620	$3,593.9

NOTES: All impacts calculated in dollars for 1990 completion rates relative to the Texas average. Border zone estimate is a weighted average net of Brewster County.

Table 5
Implied Income Gains for Limited Post-High School Education

County	Per capita impact	Aggregate impact (in millions)
Brewster	Not calculated	Not calculated
Cameron	$ 959	$227.2
El Paso	100	54
Hidalgo	1,188	413.5
Hudspeth	916	2.5
Jeff Davis	72	0.1
Kinney	887	2.6
Maverick	1,631	53.7
Presidio	1,531	9.4
Starr	2,075	75.7
Terrell	615	0.8
Val Verde	687	24.2
Webb	873	104.5
Zapata	816	6.9
Border zone	$ 711	$975

NOTES: All impacts calculated in dollars for 1990 relative to the Texas average for increased high school completion and partial college attendance. Border zone estimate is a weighted average net of Brewster County. Some data may not match due to rounding.

Table 6
Implied Income Gains from Increased College Completion

County	Per capita impact	Aggregate impact (in millions)
Brewster	Not calculated	Not calculated
Cameron	$1,199	$ 284
El Paso	736	396.9
Hidalgo	1,271	442.4
Hudspeth	1,776	4.8
Jeff Davis	Not calculated	Not calculated
Kinney	1,343	3.9
Maverick	1,877	61.8
Presidio	1,227	7.5
Starr	1,964	71.7
Terrell	1,199	1.6
Val Verde	1,054	37.1
Webb	1,328	159.1
Zapata	1,935	16.3
Border zone	$1,086	$1,487

NOTES: All impacts calculated in dollars for 1990 graduation rates relative to the Texas average.
Border zone estimate is a weighted average net of Brewster and Jeff Davis Counties.

to $1,964 per resident in Starr County. The per capita average for the entire border zone is $1,086, which translates into a gross regional income improvement of nearly $1.5 billion.

Because data shown in Tables 4, 5, and 6 are calculated in 1990 dollars, they will understate current year income losses resulting from regional dropout patterns. Also, the implied costs of secondary school noncompletion may fall below their true level in 2001 as a consequence of changes in the state and national labor markets. Namely, service sector positions make up the majority of new jobs in Texas, and many of these jobs require education beyond a high school degree. Failure to graduate from high school is thus likely to impose a more severe financial penalty today than in 1990. This can be verified once 2000 census data for Texas counties are assembled and disseminated.

Another reason gross benefits may be understated is the secondary effect of educational enhancement. As more border county residents complete each education level, they raise the probability of eventually obtaining even more training. For instance, if secondary school dropout rates are brought down, that success will likely engender more college and technical school attendance. This effect is not taken into account in any of the simulations detailed above. It should be noted, however, that income is endogenous to education. As education levels rise, skilled labor should become less scarce and the returns to education might decline. The latter implies that the parameter estimates represent upper, not lower, bounds on the simulated impacts of increases in education. Also, as education increases, other income components, such as income maintenance and unemployment transfers, may fall.

Although the preceding arguments hold true for aggregate personal income, they do not represent all factors that should be considered from a social welfare perspective. Raising secondary and postsecondary graduation rates in border counties will require some type of public initiative. Comprehensive public administrative estimates would also require any specific policy prescription to be accompanied by cost–benefit calculations. As documented elsewhere, increases in education expenditures do not yield immediate gains (Sylwester 1999). Reliable cost–benefit estimates will vary with each border county and the individual policy steps taken. Such an effort falls beyond the scope of this research. It is probably safe to conclude, however, that the net present value of policy programs implemented to raise border area enrollment and graduation rates will exceed zero by comfortable margins (see Becker 1993).

CONCLUSION

Regional economic research in recent years has attempted to quantify the relationships between per capita incomes and a variety of socioeconomic factors. This study replicates those efforts for Texas, with particular emphasis on analyzing income performance in border counties. Given the rapid transition toward an information-oriented, increasingly competitive business environment, the links between human capital and economic progress are intensifying throughout the United States.

Empirical results detailed above are broadly consistent with studies for other regions of the national economy. Failure to complete high school leads to statistically significant negative impacts on per capita incomes in Texas. Increases in the number of retirees in a county improve personal income performance throughout the state. Residents of urban counties observe greater incomes in general than do residents of smaller counties. Geography and demography also become apparent in the statistically lower incomes of counties that lie in physical proximity to Mexico.

Model simulations underscore the importance of high school graduation for border counties. Reduction of the dropout rate to a level commensurate with the rest of the state would have potentially increased income per border resident by more than $2,600 in 1990.

Collectively, that figure implies nearly a $3.6 billion earnings loss for border county economies. Reestimation with data assembled from the 2000 census is likely to indicate an even larger premium associated with educational attainment. Comparison estimates using 1990 sample information may also be feasible for the border states of New Mexico, Arizona, and California.

From a public policy perspective, the implications of this research are fairly clear. Border counties, and other regions within the state, will benefit in a direct financial manner by reducing secondary school dropout rates. Border area income performance may also be enhanced by greater public infrastructure investment. Improved transportation and communication links with the rest of the state will help offset the income decline that is at least partially associated with geographic isolation and distance from other regional markets. The latter topics are the subjects of ongoing research at various organizations in Texas.

NOTES

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[1] Weighted least squares, logarithmic, and semilogarithmic versions of several equations were also estimated as additional means of circumventing residual variance nonconstancy. Results from these estimates were generally in line with those reported in Table 3. Detailed output is available from the author.

[2] Some alternative specifications using weighted least squares with population as the weighting factor generated *t* statistics for this regressor that exceeded the 5 percent criterion. Similar results also occurred for some specifications relying on semilogarithmic versions of the model. In none of these cases, however, were all the parameters simultaneously significant. These results are available from the author.

[3] Several alternative specifications using weighted least squares or logarithmic transformations obtained statistically significant coefficients with the hypothesized signs for some or all of the language coefficients. In no individual case, however, were statistically significant parameters obtained for all explanatory variables. Results are available from the author.

[4] Additional equation versions, including the percentage of foreign-born individuals as a proxy for English language skills, also failed to render empirically superior estimates to those shown in Table 3. Detailed output is available from the author.

[5] An additional set of regressions was estimated to examine the impact of female labor force participation in Texas counties. Increases in female labor force participation rates appear to be correlated with positive impacts on county per capita incomes, but the various coefficient estimates fall just short of the 5 percent significance threshold in most instances. Border zone labor force participation tends to lag the nation as a whole (Donnelly 2000), but improved graduation rates would raise the value of female labor market participation (Becker 1993). Consequently, demographic convergence with other regions of the country may eventually help increase border incomes also.

[6] Several equations were also estimated using a second dummy variable for smaller metropolitan counties. Similar to the Missouri results discussed in Domazlicky et al. (1996) these coefficients were positive and smaller but not statistically significant. These results are also available from the author.

REFERENCES

Ashenfelter, Orley, and Alan Krueger (1994), "Estimates of the Economic Return to Schooling from a New Sample of Twins," *American Economic Review* 84 (December): 1157–73.

Becker, Gary S. (1993), *Human Capital: A Theoretical and Empirical Analysis, with Special Reference to Education*, 3rd ed. (Chicago: University of Chicago Press).

Buchner, Melissa M. (1999), "Mapping a Career Path into the Next Millennium: How Important Is College?" *The Memphis Economy*, University of Memphis Bureau of Business and Economic Research, May.

Ciccone, Antonio, and Robert E. Hall (1996), "Productivity and the Density of Economic Activity," *American Economic Review* 86 (March): 54–70.

Cox, W. Michael, and Richard Alm (1999), "The New Paradigm," Federal Reserve Bank of Dallas *Annual Report*, 3–25.

Crowder, David (1999), "State Undercounts Dropouts," *El Paso Times*, March 15, 1–2A.

Dávila, Alberto, and Marie T. Mora (2000), "English Skills, Earnings, and the Occupational Sorting of Mexican Americans Along the U.S.–Mexico Border," *International Migration Review* 34 (Winter): 133–57.

Diffley, James (1999), "Texas," *Regional Forecast: State Long-Term Tables* (Fall), 8.23–8.28 (Eddystone, Pa.: WEFA Regional Economic Service).

Dittmar, Eric, and Keith Phillips (1999), “Border Region Makes Progress in the 1990s,” Federal Reserve Bank of Dallas San Antonio Branch *Vista*, December.

Domazlicky, Bruce R., Arron Benne, Molly McMahon, Clifford Myers, and Bruce Skinner (1996), “Measuring the Cost of High School Noncompletion in Southeast Missouri,” *Journal of Economics* 22 (1): 81–86.

Donnelly, Michael (2000), “Labor Markets,” *U.S. Economic Outlook: 2000–2005* (July), 143–71 (Eddystone, Pa.: WEFA U.S. Economic Outlook Service).

Fullerton, Thomas M., Jr. (2000), “Borderplex Economic Outlook: 2000–2002,” *Business Report SR00-1*, University of Texas at El Paso.

Glaeser, Edward L. (1998), “Are Cities Dying?” *Journal of Economic Perspectives* 12 (Spring): 139–60.

Levernier, William B., Mark D. Partridge, and Dan S. Rickman (2000), “The Causes of Regional Variations in U.S. Poverty: A Cross-County Analysis,” *Journal of Regional Science* 40 (August): 473–97.

McCloskey, Deirdre M., and Stephen T. Ziliak (1996), “The Standard Error of Regressions,” *Journal of Economic Literature* 34 (March): 97–114.

Messmacher Linartas, Miguel (2000), “Desigualdad Regional en México,” *Documento de Investigación 2000-4* (México, DF: Banco de México).

Miller, Paul, Charles Mulvey, and Nick Martin (1995), “What Do Twin Studies Reveal About the Economic Return to Education? A Comparison of Australian and U.S. Findings,” *American Economic Review* 85 (June): 586–99.

Mora, Marie T., and Alberto Dávila (1998), “Gender, Earnings, and the English Skill Acquisition of Hispanic Workers in the United States,” *Economic Inquiry* 36 (October): 631–44.

Mrckvicka, Mike (2000), “Intel Sweeps Up 48 Technical School Graduates,” *El Paso Times*, November 8, 8B.

Murdock, Steve H. (1999), “Historical Patterns and Future Trends Affecting the Population of Texas and the El Paso Area” (Paper presented at University of Texas at El Paso Border Demographics Colloquium, El Paso).

Nakamura, Leonard I. (2000), “Economics and the New Economy: The Invisible Hand Meets Creative Destruction,” Federal Reserve Bank of Philadelphia *Business Review*, July/August, 15–30.

Peach, James T., and Richard V. Adkisson (2000), “NAFTA and Economic Activity Along the U.S.–Mexico Border,” *Journal of Economic Issues* 34 (June): 481–89.

Pindyck, Robert S., and Daniel L. Rubinfeld (1998), *Econometric Models and Economic Forecasts*, 4th ed. (Boston: Irwin/McGraw-Hill).

Rickman, Dan S. (1993), “The Economic Impact of High School Noncompletion in CEDO Region 8,” *Southern Economic Developer* 5: 3–13.

Sáenz, Rogelio (1999), “Mexican American Poverty in Border Areas” (Paper presented at the twenty-first annual meeting of the Association of Borderlands Scholars, Fort Worth).

Schick, Frank L., and Renee Schick (1991), *Statistical Handbook on U.S. Hispanics* (Phoenix: Oryx Press).

Sharp, John (1998), *Bordering the Future* (Austin: Texas Comptroller of Public Accounts).

Sloboda, Brian W. (1999), “Measuring the Costs of High School Dropouts on the Region of Southern Illinois,” *Journal of Economics* 25 (2): 89–101.

Sylwester, Kevin (1999), “An Empirical Analysis of Income Inequality, Education Expenditures, and Growth,” *Journal of Economics* 25 (2): 1–15.

Taylor, Lori L. (1999), “Government’s Role in Primary and Secondary Education,” Federal Reserve Bank of Dallas *Economic Review*, First Quarter, 15–24.

Welch, Finis R. (1970), “Education in Production,” *Journal of Political Economy* 78 (January/February): 35–59.

White, Halbert (1980), “A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity,” *Econometrica* 48 (May): 817–38.

Woods & Poole Economics (2000), *MSA Profile* (Washington, D.C.: Woods & Poole).

Was NAFTA Behind Mexico's High Maquiladora Growth?

William C. Gruben

Despite the consensus of commentators who otherwise typically disagree, there are at least as many reasons to suspect NAFTA did not cause the maquiladoras' growth as there are to suspect it did.

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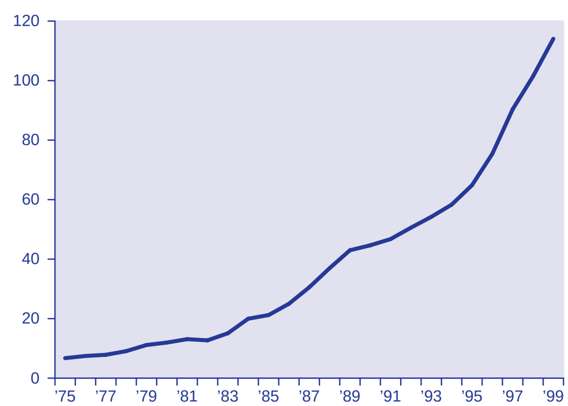
Among the most striking industrial phenomena in the wake of the North American Free Trade Agreement has been the rapid growth of plants that operate under Mexico's maquiladora program. In its simplest organizational form, a maquiladora plant imports inputs—typically from the United States—processes them, and then ships them back to the country of origin, perhaps for more processing. The maquiladora program permits the inputs and the machinery to process them to enter Mexico tariff-free. On the goods' return, the shipper pays duties only on the value added by manufacture in Mexico.[1]

Although maquiladoras have operated in Mexico since the 1960s, their output and employment growth began to accelerate markedly with the advent of NAFTA in 1994 (*Figure 1*). Over the first six years after the onset of NAFTA, maquiladora employment grew 110 percent, compared with 78 percent over the previous six years. NAFTA opponents and supporters as well as others have concluded that the trade agreement was the cause of this sharp acceleration. Balla (1998, 55), for example, claims that "without doubt, NAFTA has resulted in a dramatic increase in activity in the maquiladora industry." San Martin (2000, 32A) maintains that "NAFTA continues to drive the growth of the maquiladora industry." Carrada-Bravo (1998, 8) argues that "the acceleration of foreign direct investment under NAFTA also contributed to the creation of more than a half-million new employment opportunities in the U.S.–Mexico border region.…These new jobs, tied to the expansion of the maquiladora industry, [pay more] than those not related to international trade." A post-NAFTA report produced jointly by the Economic Policy Institute and the U.S. Business and Industry Council Educational Foundation (1997)

Figure 1
Mexican Maquiladora Employment

Number of workers (in thousands)



SOURCE: Instituto Nacional de Estadística, Geografía e Informática.

claims that "as new and expanded plants are completed in the maquiladora zone…the bilateral trade deficit should soar ever higher." Even before NAFTA took effect, Perot and Choate (1993) declared that "the flow of U.S. companies voluntarily moving factories to Mexico under the maquiladora program threatens to become a flood under NAFTA."[2]

MAQUILADORAS ARE NOT NEW AND NEITHER IS THE CONTROVERSY SURROUNDING THEM

Despite the consensus of commentators who otherwise typically disagree, there are at least as many reasons to suspect NAFTA did not cause the maquiladoras' growth as there are to suspect it did.[3] Certainly, maquiladoras have seen other episodes of sudden acceleration, albeit unlike the recent one.

Mexico developed the maquiladora program in response to the 1964 cancellation of a U.S. program that, starting during World War II, had admitted Mexican agricultural workers into the United States for temporary employment. Maquiladoras were to provide an employment alternative in the manufacturing sector for *braceros*, the agricultural workers who had lost their jobs when the U.S. program ended.

The maquiladora plants in Mexico became controversial in the United States as quickly as they appeared. Some commentators complained that while most of the braceros had been men, most of the newly hired maquiladora workers were women. However, the crux of the controversy was not so much jobs for women in Mexico as it was jobs for anyone in Mexico.

Maquiladora opponents argued that the program helped U.S. and other firms take advantage of low Mexican wages. As firms that had employed low-skilled workers in the United States set up operations in Mexico, opponents argued that maquiladoras were "taking American jobs."

Maquiladora supporters contended that if these assembly plants had not located in Mexico, they would have gone to other low-wage countries—in many cases, in Asia. Indeed, it was pointed out, lower-wage Asian countries had served as export platforms for U.S. manufacturing operations before maquiladoras came on the scene.[4]

More to the point, the maquiladoras of Mexico reflected a broader phenomenon, the globalization of manufacturing. Although the maquiladoras were creatures of Mexican law, similar operations could be found across the globe, thanks to decades of falling communication and transportation costs. These cost reductions facilitated the development of a far-flung network of assembly plants early on in Taiwan—and later in Guatemala, Mauritius, and Vietnam—whose products were marketed in the industrialized world in general and in the United States in particular (Grunwald and Flamm 1985; Romer 1993).

This globalization process was not a result of NAFTA. If anything, NAFTA was a result of this globalization process. If the reductions in communication and transportation costs that motivated globalization had not occurred, the political pressures that permitted NAFTA would not have been so strong.

HAS NAFTA MADE MAQUILADORAS ANY DIFFERENT?

While NAFTA may have motivated companies to start or expand operations in Mexico, it is also possible it discouraged maquiladora expansion or even discouraged maquiladora operations in general. This is because NAFTA allows U.S.–Mexico production-sharing operations in the maquiladora mode but without the maquiladora program.

By 1999 the majority of imports that had been processed under the maquiladora program and then entered the United States could enter duty-free outside that framework. The Automotive Products Trade Act and duty-free treatment of certain products from most-favored-nation suppliers, as well as tariff eliminations under NAFTA, made entry as easy as it was under the maquiladora program (Watkins 1994). And because of the additional paperwork the maquiladora program required, using it as a vehicle in the age of NAFTA might seem unnecessarily costly.

Another disincentive to operating under the program involved environmental restrictions. In some cases, waste-handling and treatment regulations may be interpreted as stricter for maquiladoras than for other Mexican plants making the same products. Under NAFTA some of these plants could export to the United States under levels of protectionism no higher than what the maquiladoras enjoyed, making maquiladora participation an unnecessary expense.[5]

Beginning January 1, 2001, moreover, NAFTA became the only basis for duty-free treatment of imported inputs to Mexican maquiladoras, effectively ending the maquiladora program as it related to trade among North American countries. As of that date, NAFTA provisions phased out unconditional duty-free treatment for imported components and equipment for maquiladoras. They also imposed rules requiring North American content minimums (50 percent or, in some cases, more) for duty-free movement of products between Mexico and the United States or Canada.

On the other hand, although NAFTA began lowering tariffs on goods shipped into the United States from Mexico upon its inception on January 1, 1994, the full reductions were not instantaneous. To the extent that other tariff schedules resulted in lower duties on Mexican imports than NAFTA did, the maquiladora could have remained attractive.[6]

Some NAFTA-related changes unequivocally encouraged maquiladoras. Echeverri-Carroll (1999) notes, for example, that NAFTA eliminated all Mexican programs that favored specific industries. When this occurred, some firms switched to the maquiladora program to continue importing inputs duty-free to Mexico.

In sum, some factors suggest NAFTA may have affected maquiladora growth a great deal. Other factors offer reasons to suspect NAFTA had little impact, and still others suggest the trade agreement discouraged maquiladora growth. Whether NAFTA significantly contributed to the acceleration of maquiladora employment growth is not easy to divine without econometric testing. Nevertheless, the maquiladora expansion rate during the six post-NAFTA years is two-fifths again as high as during the six pre-NAFTA years.

IF THE MAQUILADORA PROGRAM HAS BEEN PHASED OUT, WHY ARE THESE QUESTIONS WORTH ASKING?

Analysts commonly credit NAFTA for all post-NAFTA changes in U.S.–Mexico trade. (See, for example, Council of the Americas 1999 and Rothstein and Scott 1999.) Some econometric evidence suggests that NAFTA explains a significant portion of trade increases but that non-NAFTA factors have been responsible for much of the U.S.–Mexico trade fluctuation since the agreement took effect (Gould 1998). Other modeling efforts find that Mexican export growth cannot be explained by NAFTA and that the agreement's role in Mexican import expansion is unclear (Garces-Diaz 2001).

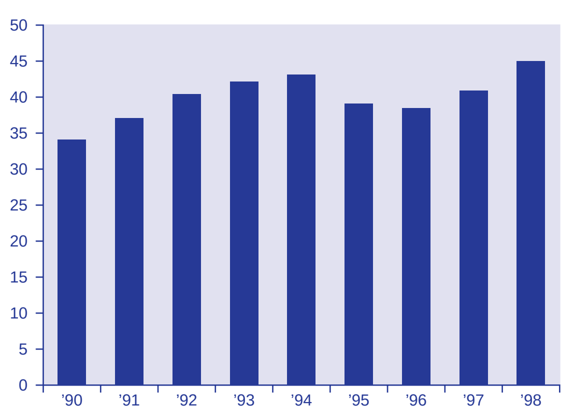
In any case, large blocks of economic activity related to U.S.–Mexico trade—such as maquiladora production—have not been shown to be either connected or unconnected to NAFTA. Why is this important? If maquiladora production and trade were linked to NAFTA, the implications for modeling NAFTA's impacts would be markedly different than if NAFTA did not influence a large portion of U.S.–Mexico trade. For example, if maquiladora activity were unaffected by NAFTA, perhaps estimates of the agreement's impact on U.S.–Mexico trade ought to exclude data for maquiladora trade.

This issue becomes clearer if we consider the broader context of Mexico–U.S. trade. Mexican shipments of crude oil to the United States represent a significant portion of total Mexico–U.S. trade, but they are clearly unconnected to NAFTA. It may be possible to identify other such goods. If it turns out that the portion of NAFTA-affected Mexico–U.S. trade is rather limited, the research focus on NAFTA would also warrant greater limitations than are currently typical.

Suppose, however, that NAFTA has affected maquiladora activity. Another interesting aspect of examining the connections between maquiladoras and NAFTA would involve assessing not only whether NAFTA affected maquiladora activity but how. For example, might NAFTA have only directly influenced maquiladora activity? Or might it have had indirect effects?

Finally, even though maquiladoras have been phased out as a phenomenon separate from NAFTA, the implications of such plants for trade liberalization related to NAFTA may deserve very different modeling and policy consideration if NAFTA motivated some kind of behavior from them that the old maquiladora rules did not. We can only measure these links while it is still statistically possible to consider maquiladoras as separate entities. Figure 2, which shows the ratio of maquiladora to total Mexican exports, demonstrates how important these im-

Figure 2
Maquiladora Exports as a Percentage of Total Mexican Exports



SOURCE: Instituto Nacional de Estadística, Geographía e Informática.

plications may be. Note that maquiladora exports accounted for more than one-third of all Mexican exports every year of the last decade.

MODELING MAQUILADORA BEHAVIOR, WITH AND WITHOUT NAFTA

To test for NAFTA's impact on maquiladora fluctuations, I apply a variant of a model designed to explain maquiladora employment (Gruben 1990). This model addresses influences on fluctuations that would occur in the presence or absence of NAFTA. The model includes adjustments for statistical problems inherent in examining such relationships (Hernandez and Navarrete Vargas 1988; Gruben 1990) and then adds a dummy variable for all periods from 1994 on.

The virtue of this model is that it is very parsimonious, yet accommodates both demand- and supply-side explanations for maquiladora employment fluctuations. To account for the demand side, I use U.S. industrial production, as Hernandez and Navarrete Vargas (1988) do. The rationale for this is that maquiladoras are essentially a segment of the U.S. manufacturing sector. When U.S. industrial production increases (falls), maquiladora employment would also be expected to increase (fall). (I also constructed a version of this model using U.S. real gross domestic product, since it is a broader measure of both supply and demand. I do not report these results because regardless of the configuration of lag lengths or of the other variables in the model, U.S. GDP never offered as much explanatory power as industrial production.)

A second category of variable involves relative wages, although, as explained below, such variables require statistical adjustment before they are suitable for a regression equation. Mexican maquiladora employment may be expected to expand or contract inversely with the ratio of Mexican manufacturing wages to comparable wages in countries that compete with Mexico in supplying products to the United States, including the United States itself. In the typical maquiladora model (Hernandez and Navarrete Vargas 1988; Gruben 1990; Truett and Truett 1993), the maquiladoras' competitors are plants in the United States and in newly industrializing Asian countries. *Ceteris paribus*, as Mexican manufacturing wages fall relative to U.S. or Asian wages, Mexican maquiladora employment is expected to grow.

I express all wages in dollars so as to characterize relative costs from the point of view of a U.S. producer or customer. This detail is important. Why denominate foreign wages in dollars when the workers will be paid in their national currency? The reason is that maquiladoras are operated chiefly by U.S. firms or by foreign firms that use maquiladora products as inputs for their U.S. operations. In either case, these firms are selling to the U.S. market. They attempt to hold down production costs in dollar terms, regardless of where the actual production takes place. So even though workers in a company's foreign plants may be paid in the local currency, the dollar value of these payments is what's important to producers trying to decide whether to produce in Mexico, the United States, or, for example, Hong Kong.

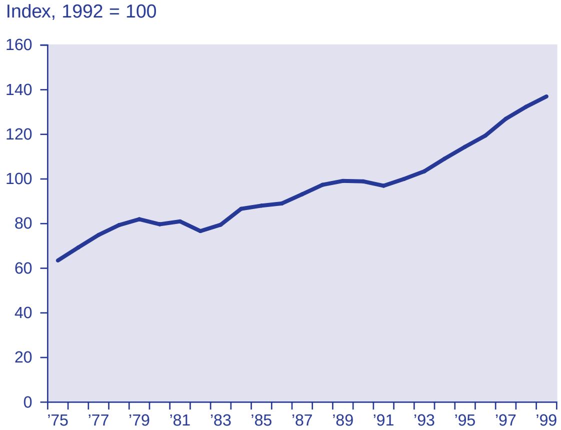
As an illustration, suppose workers in a particular country received 200 pesos per day yesterday and suddenly must receive 300 pesos per day today. However, suppose also that this increase is accompanied by a currency devaluation such that while 10 pesos purchased a dollar yesterday, 20 pesos must be exchanged for a dollar today. This would mean a reduction in the dollar cost of wages from $20 per day yesterday to $15 per day today. Naturally, U.S. firms selling in the United States will suddenly find operating in the peso-issuing country more cost-attractive—even though workers now receive 300 pesos per day instead of 200.

Some may wonder why I (and all other economists who econometrically model maquiladora behavior) use dollar-denominated relative wages rather than some measure of output per wage unit in dollar terms. For a maquiladora owner, the output per wage unit a nonmaquiladora firm can generate may not be relevant. This is because maquiladoras bring management skills and economies of scale that may result in much higher worker productivity than an average Mexican manufacturing plant employing workers of the same skill levels—but without paying wage differentials for the higher productivity.[7] Despite some skilled-labor shortages, Mexico still has an abundance of low-skilled labor whose productivity can be increased through efficient management practices and plant design. Unit labor costs offer little explanatory power in maquiladora models, while simple relative wages have much explanatory power.

IS POST-NAFTA MAQUILADORA ACCELERATION TIED TO NON-NAFTA VARIABLES?

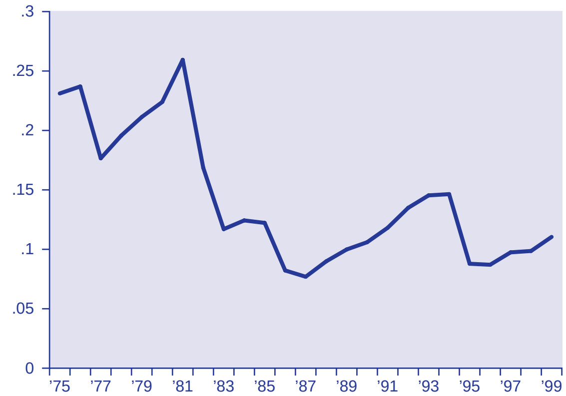
An inspection of the three explanatory variables discussed so far offers insight into why maquiladora employment accelerated after NAFTA's inception. Figure 3 displays U.S. indus-

Figure 3
U.S. Industrial Production



SOURCE: Federal Reserve Board.

Figure 4
Ratio of Mexican to U.S. Manufacturing Wages



SOURCES: Bureau of Labor Statistics; author's calculations.

trial production for 1975–99. Note the acceleration in 1992 and in 1995. Over the first six years of NAFTA, U.S. industrial production grows 32 percent, compared with 11 percent over the six years before NAFTA.

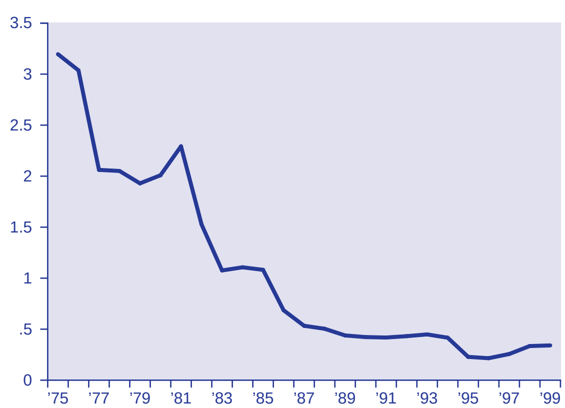
Figure 4 shows the ratio of hourly Mexican to U.S. manufacturing wages, both including benefits and both expressed in dollars. Note the sudden reduction in this ratio in 1995, one year after NAFTA began. Although this ratio edges up in succeeding years, it never goes above its levels of 1991–94. According to the literature, these lower wage ratios may be associated with higher maquiladora employment.

Similarly, the ratio of hourly Mexican manufacturing wages to hourly manufacturing wages in a sample of Asian countries (Hong Kong, Korea, Singapore, and Taiwan) also falls suddenly in 1995 (*Figure 5*). Like the Mexico–U.S. ratio, the Mexico–Asia ratio increases slightly after 1995. But unlike the Mexico–U.S. ratio, the Mexico–Asia ratio never rises to any value reached before 1995. As with the Mexico–U.S. ratios, lower Mexico–Asia ratios may be associated with higher maquiladora employment.[8]

The most recently published econometric models of maquiladora behavior use data series ending in 1988. Eleven years' data are available since the last model was estimated, so it is possible that relations between the independent variable and dependent variables could have changed.[9]

Finally, to test for NAFTA's impact on fluctuations in maquiladora employment, I use a dummy variable with the value of 0 for pre-NAFTA years and the value of 1 for post-NAFTA years. This variable is the most important of this modeling effort. A positive and significant estimated value for the NAFTA coefficient would mean that San Martin (2000), Balla (1998), Carrada-Bravo (1998), the Economic Policy Institute and the U.S. Business and Industry Council Educational Foundation (1997), and Perot and Choate (1993) are correct that NAFTA drives maquiladora growth. A negative and significant estimated value for the NAFTA coefficient would signify that NAFTA discourages maquiladora growth. This result would suggest that the maquiladora-discouraging aspects of NAFTA may dominate the maquiladora-encouraging aspects of NAFTA. If the estimate for the NAFTA coefficient is insignificant, we cannot reject the null hypothesis that the actual value of the coefficient is zero, and so perhaps NAFTA had no impact on the maquiladoras.

Figure 5
Ratio of Mexican to Asian Manufacturing Wages



SOURCES: Bureau of Labor Statistics; United Nations Industrial Development Organization; author's calculations.

Table 1
Maquiladora Employment Equations

	A	B	C	D	E
CONSTANT	.087811*** (.015193)	.077245*** (.015758)	.088691*** (.018936)	.080133*** (.011714)	.080674*** (.011988)
Mex/US Wage (–1)	–.132065** (.060174)	—	–.141554 (.131125)	—	—
Mex/Asia Wage (–1)	—	–.108876* (.059248)	.010268 (.125143)	—	—
Principal Components (–1)	—	—	—	–.014589* (.007566)	–.019336* (.010012)
Principal Components (–2)	—	—	—	–.013934* (.007305)	–.016974 (.012390)
Industrial Production	1.049863** (.395894)	1.131543** (.401794)	1.051087** (.406940)	1.278143*** (.389837)	1.200418** (.484877)
NAFTA Dummy	–.019877 (.025985)	–.010857 (.027943)	–.020874 (.029329)	–.025703 (.022549)	–.025216 (.025341)
R^2	.534362	.504401	.534536	.737014	.747451
Adjusted R^2	.460840	.426148	.431100	.675134	.684314
Standard error of regression	.051558	.053191	.052961	.040064	.040354
Sum of squared residuals	.050507	.053757	.050488	.027287	.026056
Log likelihood	37.75743	37.04029	37.76173	42.39968	—
Mean dependent variable	.118631	.118631	.118631	.121682	.120555
Standard deviation of dependent variable	.070217	.070217	.070217	.070291	.071823
Akaike criterion	–2.935429	–2.873069	–2.848846	–3.399971	—
Schwartz criterion	–2.737951	–2.675591	–2.602000	–3.152007	—
Durbin–Watson statistic	1.548164	1.434754	1.544536	2.027825	1.859293

*** Significant at the .01 level.
** Significant at the .05 level.
* Significant at the .10 level.

NOTE: Standard deviations in parentheses.

PRELIMINARY STATISTICAL EXAMINATIONS SUGGEST THE OLD RELATIONS STILL HOLD

Columns A through C in Table 1 present the results of three simple regression equations, each of which incorporates at least one Mexican wage-ratio variable, plus the industrial production index variable and the NAFTA dummy variable. I use annual data beginning in 1975 and ending in 1999. With the exception of the NAFTA dummy variable, all data are transformed into first differences of their logarithmic forms, so the data will be stationary.[10] Column A reports a regression equation that includes the U.S. industrial production variable (*Industrial Production*) and (following Gruben 1990) one lag of the Mexico–U.S. wage variable (*Mex/US Wage*), together with the NAFTA dummy variable (*NAFTA Dummy*). As expected, the industrial production coefficient is positive and significant. The Mexico–U.S. wage coefficient is negative and significant. Note that the NAFTA dummy variable is insignificant and negative, the implications of which I discuss below.

Column B of the table offers an equation with the same variables, except that the Mexico–Asia wage ratio replaces the Mexico–U.S. wage ratio. The results are essentially the same, however. Industrial production takes on a positive and significant sign. The Mexico–Asia ratio takes on a negative and weakly significant sign. The NAFTA dummy variable coefficient is again negative and insignificant.[11]

Even though the common approach to constructing econometric maquiladora models has been to account for both Mexico–U.S. wages and Mexico-industrializing country wages, an estimation problem usually develops. If I put both wage variables in the same equation, typically neither passes a significance test, even if they do separately. Sign changes also occur. These problems may be seen in Column C of the table, where both the ratio of Mexican to U.S. wages and the ratio of Mexican to Asian wages appear in an equation with U.S. industrial production and the NAFTA dummy variable. The *t* statistics for the lagged wage-ratio variables fall below the benchmark levels of significance. Moreover, the coefficient on the Mexico–Asia wage ratio takes on a positive sign, even though it is negative when the Mexico–U.S. wage ratio is not in the equation. These weakened results suggest the recurrence "of multicollinearity between the two variables that express relative costs," which appears in a similar maquiladora-related modeling exercise by Hernandez and Navarrete Vargas (1988).[12] Movements in the two wage variables are highly correlated; this correlation substantially reduces the ability of regression analysis to separately attribute variations in maquiladora employment to a wage variable.[13]

We have no direct method for correcting multicollinearity problems beyond increasing the number of observations. A procedure does exist that allows for the weighting of the wage variables to avoid multicollinearity.

Through principal components analysis, the variation of several variables can be compressed into one or more index variables, known as a principal component. The principal component is a linear combination of some collection of variables, such as the two wage ratios in this model. A mathematical procedure is used to maximize the variation of each of the two wage variables that can be captured in one index. This suppresses the contaminating effects of one wage variable's correlation with the other, and multicollinearity ceases to be a problem.

Through this maximization procedure, a coefficient becomes attached to each original variable. In this study, the wage-ratio variables are the two original variables. The values of the coefficients estimated for the variables in this index indicate the relative importance of each original variable in the new, derived component.

It should be noted that principal components estimators are biased. So, unfortunately, are the estimators derived from other procedures to avoid multicollinearity problems—including estimators in ridge regression. However, the focus of this model is to identify NAFTA's impact on maquiladora growth. Since the NAFTA dummy variable has not been converted into a principal component, the estimator bias in the principal component portion of the model will have little effect on the point of this exercise. The relation between fluctuations in U.S. industrial production—also not captured through principal components—and maquiladora employment fluctuations is also unbiased.

DOES NAFTA AFFECT MAQUILADORAS AFTER ALL? MORE COMPLETE RESULTS

Having applied principal components estimation to create an index variable that is free of multicollinearity problems but captures both Mexico–Asia and Mexico–U.S. wage relationships, I use this variable to test for NAFTA's impact on maquiladora employment fluctuations.

The result I present (Column D of the table) is the culmination of a large number of prior estimations in which I tested for the optimal set of lags. In estimating this model, I constructed alternative models that offered all possible combinations of lags from zero (that is, contemporaneous) to three lags for the principal component wage-ratio variable and for the industrial production index variable (which, again, is transformed into first differences of logarithms). "All possible combinations" includes asymmetric combinations.[14] I included up to three annual lags for U.S. industrial production and for the Mexican wage-ratio principal component.

Of all the equations I estimated, this one has the lowest Schwartz criterion value (a model with one and two lags of the Mexican wage principal component variable, a contemporaneous U.S. industrial production variable, and a NAFTA dummy variable).[15] As with the preliminary models in Columns A through C, the results in Column D do not support the claims of San Martin (2000), Balla (1998), Carrada-Bravo (1998), the Economic Policy Institute and the U.S. Business and Industry Council Educational Foundation (1997), Perot and Choate (1993), or of Perot in his television appearances in opposition to NAFTA and maquiladoras. (See note 2.) The same is true for all models that I constructed but do not report here.

The Schwartz criterion offers information about optimal lag length. The criterion punishes overparameterization—or overloading the model with lags of explanatory variables—more severely than other lag-length criteria.

The winning Schwartz criterion model offers an interesting picture of the dynamics of maquiladora operation and management. First, with the exception of the NAFTA dummy variable, the Schwartz winner has no coefficients with significance levels worse than .0735. Coefficients are at least weakly significant, a substantially better result than the previous equations and a commentary on the ability of principal components to capture collective variations. More interestingly, the Schwartz winner has a first and second lag of the Mexican wage-ratio principal component and a contemporaneous coefficient of U.S. industrial production.

That this lag configuration is the Schwartz criterion winner suggests that maquiladora management is typically less fretful about demand-side risk than about cost or supply-side risk. As for the demand side, note that the maquiladoras respond within the same year to changes in U.S. industrial production. Other things equal, maquiladoras add employees in the same year in which U.S. industrial production goes up and release them in the same year U.S. industrial production falls. Note also that, according to these results, a 1 percent increase (decline) in U.S. industrial production is associated with a 1.278 percent increase (decline) in maquiladora employment.

The same Schwartz criterion winning lag structure—with its one and two lags of the relative wage variable—suggests that managers take two years to completely respond to shocks to relative wages (supply-side shocks). The lag configuration suggests that maquiladora management takes a good deal of time deciding whether the wage shocks are transitory or will be long-lived, even though responses to changes in demand (that is, industrial production) are relatively quick.

Maquiladora companies may need so much time to adjust to a wage because of their dollar-denominated (rather than peso-denominated) perspective on costs. When we express all wages in dollars, the largest and most sudden shocks to relative wages will involve currency devaluations. The lag structure of the Schwartz winner is consistent with maquiladora companies waiting to see if Mexican wages will adjust to their old levels in dollars immediately after a devaluation or if the adjustment process will be a slow one. As Figures 4 and 5 show, the adjustment was very slow in the wake of Mexico's 1982 and 1994 devaluations. These devaluations not only made the dollar value of Mexican wages fall hard and fast (compared with U.S. or Asian wages), but Mexican wages in dollar terms remained depressed for years.

However, this peso-wage stickiness may not be a certainty. First, it need not persist after every devaluation. McLeod and Welch (1991) show that in many countries this stickiness is not typical and that relative wages return to their old relations far more quickly than in Mexico.

Second, the same investor pressures that trigger Mexican devaluations can simultaneously trigger exchange-rate pressures elsewhere. When a Mexican devaluation lowers wage costs in dollar terms, maquiladora owners may wait to see if devaluations will follow in other countries where they own plants, eroding the Mexican wage advantage. Mexico's last exchange rate crisis, in 1994–95, sparked capital outflows elsewhere in Latin America and in the Philippines and Poland. The Russian devaluation of third-quarter 1998 created financial pressures in Brazil and Argentina, as well as in Mexico. Russia's crisis triggered fears about Brazil that were sufficient to cause large reserve losses (Treuherz 2000). In the wake of these losses, political events within Brazil incited further capital outflows until Brazil devalued the *real* in January 1999. The lag structure on wage ratios is consistent with the time firms may wait to see how these exchange-rate relations sort themselves out.

With respect to the NAFTA dummy variable, the conjectures of the authors who claim NAFTA drove maquiladora growth are unconfirmed. The coefficient value for this variable proves insignificant in every one of the scores of equations I constructed in preparation for building the model I present here. For the authors mentioned, these insignificant coefficient values would offer cold comfort, in and of themselves. However, the NAFTA dummy variable coefficient also takes on a negative sign. This, and the fact that every form of this model gives a negative sign to the dummy variable coefficient, means a coefficient value significantly different from zero would reject those authors' claims even more soundly than insignificance would. It should be reiterated that this negative sign appears in every estimation I performed, ranging from three lags of one variable and no lags of the other, all the way to no lags of the one variable and three lags of the other.

INSTRUMENTAL VARIABLES APPROACH

One last refinement is to deal with the possible problem of simultaneity bias, which arises because employment and wages are jointly and simultaneously determined.

Accordingly, a final step in estimating this model is to substitute an instrumental variable for the lagged principal components originally included to adjust for the multicollinearity problems in the wage ratios.[16] A Hausmann test shows that the instrumental variables model uncorrected for heteroskedasticity is not statistically significantly different from the ordinary least squares model. So from a statistical point of view, there is no need to construct an instrumental variables equation. Nevertheless, in Column E of the table, I present results for instrumental variables estimation because of theoretical reasons to suspect simultaneity bias, even if testing does not bear them out. As can be seen, the first lag of the instrumented principal component (wage-ratio variable) is negative and significant, while the second lag is negative and insignificant. As before, the industrial production variable is positive and significant, while the NAFTA dummy variable is negative and insignificant. These results are so consistent with the ordinary least squares results in Column D that they add little to this narrative aside from offering assurances the OLS results are quite robust.

SUMMARY AND CONCLUSIONS

In response to widespread arguments that Mexican maquiladoras' rapid growth after NAFTA was a result of NAFTA, I have performed extensive econometric tests of its effect on maquiladora employment. The results of these tests are resoundingly negative. NAFTA did not make maquiladoras grow faster. Such effect as NAFTA had was negative, not positive, albeit statistically insignificant. So we cannot say that NAFTA had any effect on maquiladoras.

Instead, the acceleration of maquiladora employment growth from NAFTA's inception through 1999 can be explained by changes in demand factors (as expressed by changes in the U.S. industrial production index) and in supply-side/cost factors (as expressed by changes in the ratios of Mexican to U.S. manufacturing wages and to manufacturing wages in four Asian countries). Growth in the U.S. industrial production index over the six years following NAFTA was roughly three times as rapid as during the six previous years. Likewise, Mexico's 1994 devaluation meant that during the first six years of NAFTA, the ratio of Mexican manufacturing wages to their counterparts in the United States, Hong Kong, Korea, Singapore, and Taiwan was far below these ratios during the six years before NAFTA.

The basic equation this article presents is the culmination of many estimations. I present the first- and second-lag configuration equation because it has the optimal lag structure, based on the Schwartz criterion. After adjusting for degrees of freedom in the ordinary least squares version, this equation captures 67.5 percent of all variation in maquiladora employment; an estimated 73.7 percent of variation is captured before adjusting for degrees of freedom. After adjustment, the instrumental variables version captures 68.4 percent of total variation.

The lag length has interesting implications for how maquiladora owners handle risk. These companies respond relatively quickly to changes in demand (as expressed by changes in the U.S. industrial production index), while the response to shifts in relative wages does not occur during the same year as the shift in relative wages (as expressed by changes in the ratios of Mexican to U.S. and Asian wages). Instead, these responses mainly occur one and two years after the shifts, as if maquiladora operators wait to see if the wage shocks are going to be permanent. It is telling that in 2001, following the period I model, maquiladora employment has fallen significantly as U.S. industrial production has slid—validating my model.

While some evidence suggests NAFTA has affected trade between the United States and Mexico (Gould 1998), the effects were not expressed through Mexico's maquiladora system. Trade is a complicated process, and so are changes in trade policy such as NAFTA.

NOTES

I wish to thank Eric Millis for his careful econometric work and his valuable judgment and advice, Ana Prats for organizing Table 1 and the figures, and Monica Reeves for editing this article with great care.

1 Note that the return trip is not under the jurisdiction of the maquiladora program. The tariff arrangements involve U.S. law, not Mexican law.

2 One of presidential candidate Ross Perot's television props was a blowup of an ad inviting maquiladoras to locate in the southern Mexican state of Yucatán. This, Perot said, was what NAFTA would result in.

3 Balla and Carrada-Bravo are pro-NAFTA. Perot and Choate and the Economic Policy Institute and the U.S. Business and Industry Council are anti-NAFTA.

4 The Border Industrialization Program, under which Mexico's maquiladoras began, was introduced by Mexico's secretary of commerce and industry after a trip to East Asia. The program was his policy response to what he had seen—labor-intensive assembly operations, with East Asian workers employed in plants that belonged to U.S. corporations and the same import tariff arrangements that were later applied to maquiladoras (Fernández-Kelly 1987, 151).

5 See Boyer (1997) for a detailed characterization of Mexican environmental law and its significance for maquiladoras.

6 Three general categories of U.S. tariff policy have been applied to maquiladora products imported into the United States. The first—Harmonization Tariff Schedule 9802.00.60—permits the importation of fabricated but unfinished metal products processed abroad. Duties are assessed on the value added in Mexico, rather than on the product's total value. The products must be processed in the United States

before being sent abroad and further processed in the United States upon their return. The second category —HTS 9802.00.80—makes an article assembled in Mexico from U.S.-made components exempt from import duties on the components. These products need not have metal components. The third category, now moot, was the most generous. If the goods assembled or manufactured in Mexico had at least 35 percent Mexican content upon import into the United States, they were eligible for duty-free treatment under the U.S. Generalized System of Preferences, or GSP. Mexico was removed from the list of GSP-eligible countries when it joined NAFTA, but NAFTA allows products the same immediate, duty-free entry as they had under the GSP.

As a result of NAFTA, an additional Harmonization Tariff Schedule was created. HTS 9802.00.90 allows for the duty-free treatment of textile and apparel products assembled in Mexico from U.S.-formed and cut fabric. Textile and apparel products have historically entered the United States under special trade restrictions, so liberalizations of such trade have had to be product-specific. For apparel that had entered under 9802.00.80, only the value of U.S.-cut fabric and U.S.-made fasteners such as buttons and zippers came in duty-free. Under 9802.00.90, the value added in Mexico—including labor and overhead—also enters the United States duty-free. For additional discussion, see U.S. International Trade Commission (1999).

[7] On the issue of scale economies, note that the employment and output of the average Mexican manufacturing plant is smaller than that of the average maquiladora or the average U.S. plant. In the wake of Mexico's 1994 devaluation, real manufacturing wages adjusted downward and remained below predevaluation real wages for years. In dollar terms, average overall manufacturing sector wages in Mexico remained below 1994 levels as late as 1999, the most recent year for which data are available.

[8] When I applied augmented Dickey–Fuller and Phillips–Perrone tests to the variables discussed (including the dependent variable, maquiladora employment), none could reject the null hypothesis of a unit root. I accordingly took first differences of logarithms of these variables, which allows rejection of this hypothesis in every case.

[9] All three models cited here—Hernandez and Navarrete Vargas (1988), Gruben (1990), and Truett and Truett (1993)—use a Mexico–U.S. wage variable, a Mexico–other industrializing country wage variable, and a U.S. output variable.

[10] See note 8 for an explanation of why I used first differences. The wage data for all countries—Mexico, Hong Kong, Korea, Singapore, Taiwan, and the United States—are from the International Comparisons of Hourly Compensation Costs for Production Workers in Manufacturing series, on the U.S. Bureau of Labor Statistics web site, ftp://ftp.bls.gov/pub/special.requests/ForeignLabor/supptab.txt. These data go back to 1975. Again, note that the virtue of these data is that they account for benefits as well as salaries. Maquiladora employment data are from the Mexican government's INEGI (Instituto Nacional de Estadística, Geografía e Informática) web site, www.inegi.gob.mx. U.S. industrial production data are from the Federal Reserve Board.

[11] Although I do not report them here, the results are essentially the same with or without the NAFTA dummy variable. This variable does not affect the equation in this configuration.

[12] Hernandez and Navarrete Vargas (1988, 225), my translation. Their model is quarterly, rather than annual, and uses wage data uncorrected for international differences in worker benefits, unlike the U.S. Bureau of Labor Statistics data my model uses. Truett and Truett (1993) avoid results consistent with multicollinearity by using a different dependent variable than other authors. However, the dependent variable they apply has been criticized on other grounds.

[13] Even when the data are transformed into first differences of logarithms, representing growth rates of the Mexico–U.S. and Mexico–Asia wage variables, the coefficient of correlation between the two is .8852. This is substantive evidence of serious multicollinearity. In contrast, the coefficient of correlation between the transformed version of the Mexico–Asia wage ratio and the similarly transformed (first differences of logarithms) U.S. industrial production index variable is only .0935. The coefficient of correlation between the transformed version of the Mexico–U.S. wage ratio and the transformed U.S. industrial production index is .1232. These last two, low degrees of correlation suggest that the multicollinearity in the equations is strictly in the wage-ratio variables.

[14] An example of asymmetry would be a contemporaneous-only U.S. industrial production variable with three lags of the principal components variable, or vice versa. I also ran symmetrically lagged models, with contemporaneity plus three lags of the industrial production variable, along with (in the same equation) the same lag structure for the principal components variable and so on. In so doing, I followed the London School of Economics paradigm of running every possible combination (in this case, up to three annual lags).

[15] The Schwartz criterion is one of the most common tests for optimal lag specification. When testing alternative lag structures for a model, the one with the lowest value wins.

[16] The goal in creating instruments to proxy for the original variable is to find variables that may be correlated with the right-hand-side variable subject to simultaneity bias, but not with the dependent variable. In this case, a linear combination of the contemporaneous peso–dollar exchange rate together with lagged

principal components was considered a reasonable candidate for constructing an instrument for the original component. When the Sargan test of instrument validity was applied, the related F value was 0.3048, implying a .9024 level of significance, clearly demonstrating the instrumentation was valid.

REFERENCES

Balla, John A. (1998), "Data Transfer: What's New?" *Twin Plant News*, November, 55–56.

Boyer, Madeleine (1997), "Pending Developments in Mexican Hazardous Waste Law (Paper presented at the 1997 Advanced Environmental Law Course, State Bar of Texas, Houston, December 4).

Carrada-Bravo, Francisco (1998), "Business Education and Joint Enterprise in the NAFTA Countries," Working Paper, Department of World Business, Thunderbird–The American Graduate School of International Management (Glendale, Ariz.).

Council of the Americas (1999), *NAFTA at Five Years* (New York: Council of the Americas, January).

Echeverri-Carroll, Elsie L. (1999), "Industrial Restructuring of the Electronics Industry in Guadalajara, Mexico: From Protectionism to Free Trade" (Austin: University of Texas Bureau of Business Research).

Economic Policy Institute and the U.S. Business and Industry Council Educational Foundation (1997), *The Failed Experiment: NAFTA After Three Years* (Washington, D.C.).

Fernández-Kelly, M. Patricia (1987), "Technology and Employment Along the U.S.–Mexican Border," in *The United States and Mexico: Face to Face with New Technology*, ed. Cathryn L. Thorup, U.S. Third World Policy Perspectives series, no. 8 (New Brunswick, N.J.: Transaction Books).

Garces-Diaz, Daniel (2001), "Was NAFTA Behind the Mexican Export Boom (1994–2000)?" (Mexico City: Banco de México, February).

Gould, David M. (1998), "Has NAFTA Changed North American Trade?" Federal Reserve Bank of Dallas *Economic Review*, First Quarter, 12–23.

Gruben, William C. (1990), "Do Maquiladoras Take American Jobs? Some Tentative Econometric Results," *Journal of Borderlands Studies* (Spring), 31–45.

Grunwald, Joseph, and Kenneth S. Flamm (1985), *The Global Factory: Foreign Assembly in International Trade* (Washington, D.C.: Brookings Institution).

Hernandez, Jose Luis, and Rodolfo Navarrete Vargas (1988), "Determinantes del crecimiento del empleo en la industria maquiladora de exportación en México," in *Maquiladoras: Primera Reunion Nacional Sobre Asuntos Fronterizos*, ed. Arturo Garcia Espinosa (Monterrey: Asociación Nacional de Universidades e Institutos de Enseñanza Superior and Universidad Autonoma de Nuevo Leon), 221–47.

McLeod, Darryl, and John H. Welch (1991), "Free Trade and the Peso," Federal Reserve Bank of Dallas Research Paper no. 9115 (Dallas, July).

Perot, Ross, and Pat Choate (1993), *Save Your Job, Save Our Country: Why NAFTA Must Be Stopped—Now!* (New York: Hyperion).

Romer, Paul M. (1993), "Two Strategies for Economic Development: Using Ideas and Producing Ideas," in *Proceedings of the World Bank Annual Conference on Development Economics 1992* (Washington: International Bank for Reconstruction and Development), 63–91.

Rothstein, Jesse, and Robert I. Scott (1999), "NAFTA's Casualties" (Economic Policy Institute paper presented at "NAFTA: The First Five Years," sponsored by the El Paso Branch of the Federal Reserve Bank of Dallas, El Paso, November 4).

San Martin, Nancy (2000), "Overworked and Underage," *Dallas Morning News*, March 5, A1, A31–A33.

Treuherz, Rolf Mario (2000), *The Crisis Manual for Emerging Countries* (Palo Alto, Calif.: Fabrizio Publications).

Truett, Lila J., and Dale B. Truett (1993), "Maquiladora Response to U.S. and Asian Relative Wage Rate Changes," *Contemporary Policy Issues* 11 (January): 18–28.

U.S. International Trade Commission (1999), *Production Sharing: Use of U.S. Components and Materials in Foreign Assembly Operations, 1995–1998*, USITC publication no. 3265 (Washington, D.C.: U.S. International Trade Commission, December).

Watkins, Ralph J. (1994), "Implications of the North American Free Trade Agreement for Mexico's Maquiladora Industry and the Use of the Production Sharing Tariff Provision," in *Production Sharing: U.S. Imports Under Harmonized Provisions 9802.00.60 and 9802.00.80, 1989–1992* (Washington, D.C.: U.S. International Trade Commission).

Explaining Stock Price Movements: Is There a Case for Fundamentals?

Nathan S. Balke and Mark E. Wohar

Plausible changes in expectations about real dividend growth and discount rates can explain stock prices in the 1990s.

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In the preface of Robert Shiller's recent book *Irrational Exuberance*, he asks:

> "Are powerful fundamental factors at work to keep the market as high as it is now or to push it even higher, even if there is a downward correction? Or is the market high only because of some *irrational exuberance*—wishful thinking on the part of investors that blinds us to the truth of our situation?" (Shiller 2000, xii).

Shiller answers his own question by arguing that stock prices in the 1990s displayed the classic features of a speculative bubble. High prices are sustained, temporarily, by investor enthusiasm rather than real fundamental factors. Investors, according to Shiller, believe it is safe to purchase stocks, not because of their intrinsic value or because of expected future dividend payments, but because they can be sold to someone else at a higher price. Simply put, stock prices are driven by a self-fulfilling prophecy based on similar beliefs of a large cross section of investors.

When looking at broad stock market price indexes, such as the Standard & Poor's 500, Shiller's argument is largely based on two premises about the historical behavior of stock prices. First, Shiller asserts that marketwide price–dividend and price–earnings ratios have a tendency to revert toward their historical averages. This implies that high stock price valuations are not likely to persist.[1] Second, dividend movements are not nearly volatile enough to rationalize stock price volatility. This suggests that changes in expectations about future dividends cannot be responsible for stock price movements.

While Shiller argues that irrational exuberance explains the run-up in stock prices in the 1990s, we present evidence that the case for market fundamentals is stronger than it appears on the surface. First, we demonstrate that swings in the price–dividend and price–earnings ratios show substantial persistence, particularly since World War II. This raises doubts about the existence of a "normal" price–dividend (or price–earnings) ratio. Hence, using the average value of one of these ratios as a gauge of the average, or normal, valuation ratio is misleading. A price–dividend ratio of 30 may have seemed high from the pre-1980s perspective but not after the 1990s. Second, we investigate whether plausible combinations of lower expected future real discount rates or higher expected real dividend (earnings) growth could rationalize broad market stock values, raising the possibility that changes in market fundamentals have made a major contribution to the run-up in stock prices.

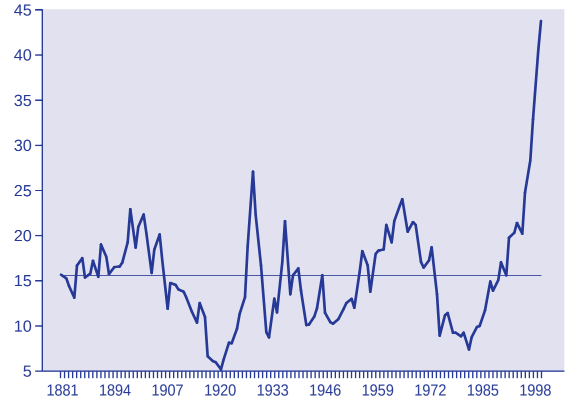
A number of explanations have been offered for the unprecedented rise in stock prices during the 1990s. These include increased expected future economic growth brought about by the revolution in information technology, demographic changes as the baby boomers age, a reduction in the equity premium as a result of lower transaction costs and increased diversification, lower business cycle risk, a decline in inflation, and momentum investing.[2] See Carlson (1999), Carlson and Sargent (1997), Cochrane (1997), Kopcke (1997), Heaton and Lucas (2000), Siegel (1999), Carlson and Pelz (2000), Shiller (2000), Jagannathan, McGrattan, and Scherbina (2000), and McGrattan and Prescott (2000) for recent surveys.

In this article we provide evidence that it may be misleading to think of a normal price–dividend or price–earnings ratio. Next we review a standard stock valuation model on which our analysis is based. We then present evidence that there has been an increase in expected real dividend (earnings) growth and a decline in the expected real discount rate. Finally, we discuss some caveats and comments.

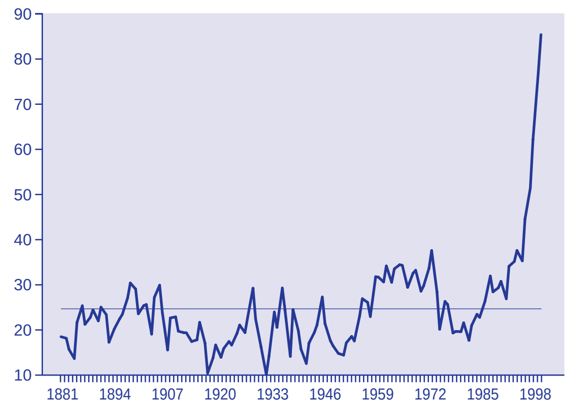
Figure 1

Price–Earnings Ratio, 1881–1999

(Using ten-year average earnings)



Price–Dividend Ratio, 1881–1999



NOTE: Ratios are January values.

ARE THERE "NORMAL" PRICE–DIVIDEND AND PRICE–EARNINGS RATIOS?

The dramatic rise in stock prices during the 1990s has challenged financial analysts and economists to explain why stock valuation ratios are so high relative to historical levels. The top panel of Figure 1 plots the January values of the real S&P 500 index divided by the ten-year moving average of real earnings (henceforth, P/E) for 1881 through 1999. By using a ten-year moving average we attempt to measure trends in long-run or permanent earnings.[3] The bottom panel shows the comparable graph for the price–dividend (P/D) ratio.[4]

Figure 1 shows that during this 119-year period, the P/E ratio crossed and persisted above or below its sample mean of 15.7 about ten times.[5] These ten major crossings suggest that the P/E ratio on average spends more than a decade consistently above or below the sample mean. Since World War II, there have been only three major crossings of the sample mean, and the average time between crossings is close to twenty years. A similar pattern is noted for the P/D ratio, shown in the lower panel. Although there were more crossings before 1947, there have been only three crossings since.

While it's clear that movements in the P/E and P/D ratios can be persistent, are changes in these ratios permanent? Using the entire sample (1881 through 1999), we cannot reject the null hypothesis of a unit root in logarithm of P/D or P/E ratios.[6] This means that unanticipated changes, or "shocks," in the logarithm of P/D or P/E ratios can be permanent. The presence of permanent shocks makes it doubtful that stock prices will fall in the future just because they are historically high. If permanent changes in the P/D ratio are possible, then a P/D of 30, which would have been considered high in the past, may not look very high from the perspective of the 1990s.

A STOCK PRICE VALUATION MODEL

Given that there may be no "normal" level of stock prices, how can one assess whether stock prices are too high? To answer this question, it is useful to review a simple model of stock price determination. Investors hold stocks to obtain future income, in the form of either

Derivation of the Gordon Model

The basic stock price valuation model posits that the stock price is equal to the present discount value of expected future dividend payments. We can write this present value as

$$P_t = \sum_{i=1}^{\infty} E_t\left[R(t,i)D_{t+i}\right],$$

where E_t[.] denotes expectations based on information available to investors at time t and D_{t+i} is the level of real dividends at time $t+i$. The term $R(t,i)$ is the degree to which future expected dividends in time period $t+i$ are discounted back to the current time period, or

$$R(t,i) = \prod_{j=1}^{i} \frac{1}{1+r_{t+j}}$$

with r_{t+j} being the real discount rate at time period $t+j$. We can rewrite stock prices in terms of the price dividend ratio:

$$\frac{P_t}{D_t} = \sum_{i=1}^{\infty} R(t,i)\,d(t,i),$$

where

$$d(t,i) = \prod_{j=1}^{i}(1+g_{t+j}).$$

The term $d(t,i)$ represents the compounded expected real dividend growth from period $t+1$ to $t+i$, where g_{t+j} is expected real dividend growth in time period $t+j$.

If expected real dividend growth and real discount rates are constant over time, the above equation simplifies to the basic Gordon (1962) model that states

$$P_t/D_t = (1+g)/(r-g),$$

where g is expected future growth in dividends and r is the discount rate.

Assuming that dividends are a linear function of earnings, we can write a comparable expression with earnings. If we assume a constant payout ratio, that is,

$$D_t = q\,E_t,$$

where q is the payout ratio, we can write the expression for the P/E ratio as

$$P_t/E_t = q(1+g)/(r-g).$$

Note that expected real returns in the Gordon model are equal to the discount rate. Expected stock returns are equal to

$$E_t\left[return_{t+1}\right] = E_t\left[\frac{\left(\frac{P_{t+1}}{D_{t+1}}+1\right)\frac{D_{t+1}}{D_t}}{\frac{P_t}{D_t}} - 1\right],$$

where E_t[.] refers to expectations based on information available at time t. Using the Gordon model, expected returns are equal to

$$E_t\left[return_{t+1}\right] = \frac{\left(\frac{1+g}{r-g}+1\right)(1+g)}{\frac{(1+g)}{(r-g)}} - 1 = r.$$

Thus, the expected return equals the discount rate, r.

dividend payments or an increase in stock price (capital gains). A firm's ability to pay dividends and investors' expectations of higher stock price depend on the firm's future earnings growth: the higher a firm's earnings growth, the greater potential to pay dividends and the more an investor is willing to pay for a share of stock. Another aspect investors must consider is how to value these cash flows. This is determined by the real discount rate.[7] Individuals generally prefer to receive income sooner rather than later. To give up a dollar's worth of current income, investors demand more than a dollar in the future. The more impatient investors are (the higher the real discount rate, or required return), the less they are willing to pay (and sacrifice current consumption) for a stock with a given level of future income. Individuals also prefer less uncertainty. Hence, investors will demand a higher expected return from a risky asset than from a safe asset or, equivalently, will discount the expected payoff at a higher rate. Thus, asset prices are determined by the properties of the income flow that these assets generate and by how investors value this income flow; that is, the price of a financial asset equals the present discounted value of the stream of cash flows from that asset.

The standard Gordon model, described in the box "Derivation of the Gordon Model" and in the equation below, gives the factors that affect the fundamental value of stock prices. If expected real dividend growth and real discount rates are constant over time, the P/D ratio at time t is given by

$$P_t/D_t = (1+g)/(r-g),$$

where g is expected future growth in dividends and r is the discount rate. Factors such as productivity-enhancing technological change might lead to higher expected real dividend (real earnings) growth, causing the P/D (P/E) ratio to rise. On the other hand, factors such as increased tolerance of risk or investors' greater willingness to postpone current consumption might reduce the expected real discount rate, also causing these ratios to rise. Note also, as the box demonstrates, the Gordon model implies that the expected real return on stocks is equal to the real discount rate.

Table 1 presents means, standard deviations, and 95 percent confidence intervals (adjusted for serial correlation) for six variables, averaged over various time periods:[8] P/E ratio, P/D ratio, real (inflation-adjusted) earnings growth, real dividend growth, real returns, and excess returns.[9] During the 100-year period before the most recent bull market, which began in 1983, annual real dividend growth averaged 0.9 percent and average real returns were around 5.5 percent. Plugging these numbers into the Gordon formula yields a P/D ratio of around 21.9, which is very close to the P/D ratio averaged for 1881–1982. However, for 1983 through 1999, both the P/D and P/E ratios are substantially higher than the P/D ratio implied by the Gordon model, given the historical averages of real dividend growth and real returns. For the years 1983 through 1999, the P/D (P/E) ratio averaged 39.10 (20.50), while

Table 1

Descriptive Statistics for Stock Prices, Earnings, and Dividends

	P/E ratio (January averages)	P/D ratio (January averages)	Annual real earnings growth	Annual log real dividend growth	Annual log real returns	Annual log excess returns
1881 through 1999						
Mean	15.66	24.91	1.45	1.05	6.57	3.76
Standard deviation of variable	6.09	10.75	20.81	11.14	17.06	17.19
95% confidence band of the mean	(13.72, 17.48)	(21.55, 28.27)	(–2.11, 5.02)	(–.97, 3.08)	(3.60, 9.45)	(.59, 6.92)
1881 through 1982						
Mean	14.80	22.40	.89	.90	5.49	2.85
Standard deviation of variable	4.63	6.01	21.49	11.99	17.66	17.83
95% confidence band of the mean	(13.21, 16.39)	(20.39, 24.41)	(–3.10, 4.89)	(–1.45, 3.34)	(2.29, 8.69)	(–.63, 6.32)
1983 through 1999						
Mean	20.50	39.10	4.80	2.01	13.10	9.21
Standard deviation of variable	10.16	18.53	16.35	2.53	11.18	11.67
95% confidence band of the mean	(12.66, 28.33)	(25.15, 53.05)	(–1.81, 11.42)	(.58, 3.44)	(8.35, 17.84)	(4.15, 14.27)
1995 through 1999						
Mean	31.75	59.53	6.71	2.43	21.71	18.64
Standard deviation of variable	9.13	19.30	11.25	1.61	6.10	6.69
95% confidence band of the mean	(22.94, 40.57)	(40.71, 77.93)	(1.85, 11.57)	(1.34, 3.51)	(17.78, 25.65)	(14.11, 23.17)

NOTE: Annual growth rates are calculated as $\log(X_{t+1}/X_t)$.

from 1995 through 1999, the P/D (P/E) ratio averaged 59.53 (31.75). Thus, given no change in either expected real dividend growth or discount rate, stock prices seem too high relative to the dividends they pay. For market fundamentals to explain such high stock prices, either expected future real dividend (earnings) growth must be higher or expected future discount rates must be lower than their historical averages, or both.[10]

Table 2 shows P/D ratios for hypothetical combinations of expected real dividend growth and real discount rates using the Gordon model. Row 3 of Table 2 displays the P/D ratio implied by various levels of expected future real dividend growth, given a real discount rate of 5.5 percent. For the P/D ratio to reach levels seen during the mid- to late 1990s, expected real dividend growth would need to increase to 3.5 percent in the future. This is well above historical and current values. While real dividend growth rates for the years 1995 through 1999 were higher than their historical averages, they were not nearly high enough to generate P/D ratios that averaged 59.53 from 1995 to 1999.

As can be seen from Table 2, a decline in the discount factor can also increase the implied P/D ratio. Given historical values of 1 percent for real dividend growth, a decline in the discount rate to around 3 percent would be required for the P/D ratio to be near that averaged over the past five years. Discount rates this low would imply about a 3 percent expected real return for stocks, which is close to the average real return on government bonds. This, in turn, implies that the risk premium on stocks is nearly the same as that on government bonds. One might argue that either an increase in expected real dividend growth or a decrease in the discount factor *alone* is responsible for historically high stock prices. However, an increase in expected real dividend growth combined with a decline in the discount factor could account for an increase in the P/D ratio. For example, from Table 2, if real dividend growth increases to 2 percent per year while the discount rate falls to around 4 percent, it is possible to obtain a P/D ratio in the neighborhood of 50. In fact, we argue that, of the scenarios based on fundamentals, the combination of a discount rate decline and a modest increase in real dividend growth could be the most plausible.

Annual real stock returns averaged above 13 percent for 1983 through 1999 and almost

Table 2

P/D Ratios for Hypothetical Combinations of Dividend Growth and Discount Rates

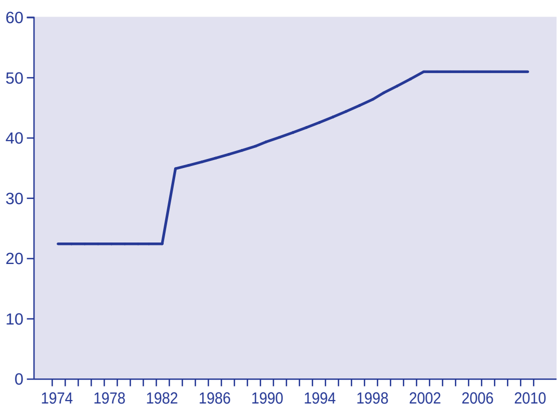
	Dividend growth				
Discount rates	**g = .01**	**g = .015**	**g = .020**	**g = .025**	**g = .035**
r = .03	50.50	67.67	102.00	205.00	NA
r = .04	33.67	40.60	51.00	68.33	207.00
r = .055	22.44	25.37	29.14	34.17	51.75
r = .07	16.83	18.45	20.40	22.78	29.57
r = .08	14.43	15.62	17.00	18.64	23.00

NOTE: NA = not applicable.

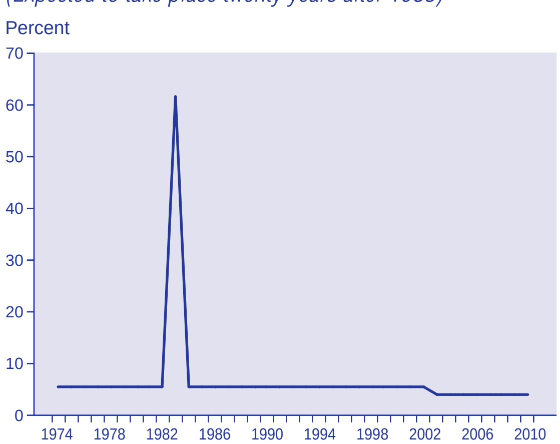
22 percent for 1995 through 1999. But in our example above, the discount rate, and hence, expected future stock returns, fell. Although it may seem paradoxical, a reduction in expected future returns can imply high current returns. When investors change their expectations about future expected real dividend growth and expected real discount rates, those who currently hold stocks reap an unanticipated capital gain.

To illustrate, suppose that before 1983 investors expect future real dividend growth and future real discount rates to remain at their historical values of 1 percent and 5.5 percent, respectively. Then, in 1983, they suddenly expect future real dividend growth to rise to 2 percent and the real discount rate to fall to 4 percent starting in 2003 and to remain there indefinitely.[11] The top panel of Figure 2 shows the path the P/D ratio would take given the

Figure 2
P/D Ratio for Permanent Change in Dividend Growth and Discount Rate
(Expected to take place twenty years after 1983)



Realized Returns for Change in Dividend Growth and Discount Rate
(Expected to take place twenty years after 1983)



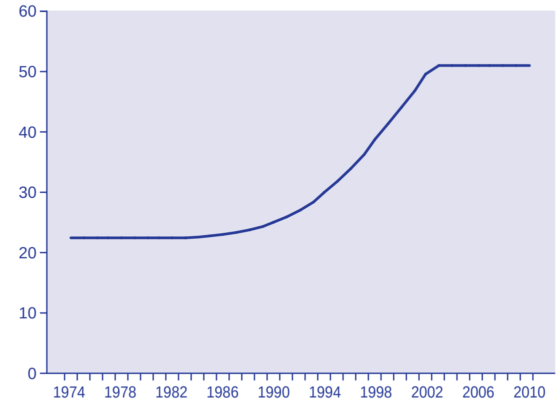
change in expectations about the future real dividend growth and the real discount rate. The P/D ratio jumps in 1983, then steadily rises to a peak of 51 in 2003, and remains at 51 thereafter. In other words, in anticipation of the future increase in real dividend growth and the decline in real discount rate, current P/D ratios rise. The bottom panel of Figure 2 shows the effect of this example on actual real stock returns. Stockholders reap a windfall in the period in which expectations of future real dividend growth and real discount rates change—in this example, the real return on stocks is a whopping 62 percent in 1983. However, subsequently the real return on stocks exactly equals the discount rate (5.5 percent before 2003, 4 percent thereafter).

A more realistic scenario is one in which investors gradually revise their expectations about future real dividend growth and real discount rates. Initially, investors perceive only a small probability that real dividend growth and the real discount rate will change, but, as the time of the change approaches, this probability is gradually revised upward until investors attach nearly 100 percent probability to a change on the eve of its occurrence. Figure 3 presents the resulting time paths for the P/D ratio and the realized returns on stocks. Notice that between the time investors first become aware of the possibility of a dividend growth and discount rate change and the time that the actual change occurs, the P/D ratio steadily rises. In some respects, this looks similar to the increase in actual P/D ratios seen since 1983. Furthermore, as shown in the bottom panel of Figure 3, actual returns for stocks are higher than expected over this period. In fact, during this transition period actual returns are consistently greater than historical returns, much like what has actually occurred since 1983. Actual returns are also greater than expected returns (the discount rate) because investors are continuously and pleasantly surprised during this period. Again, the reason is that as investors revise their expectations of future real dividend growth and future real discount rates, these revisions result in unanticipated capital gains for stocks. Once investors attach nearly 100 percent probability to the new regime, actual returns approach the discount rate. Note also that once the new discount rate takes effect, actual stock returns fall along with the discount rate.

These examples are not meant to be a literal description of what has happened in the equity markets since 1983, but they do illustrate that as expectations of real dividend growth and real discount rates change, it is possible for real-

Figure 3

P/D Ratio for Permanent Change in Dividend Growth and Discount Rate



Realized Returns for Permanent Change in Dividend Growth and Discount Rate

Percent

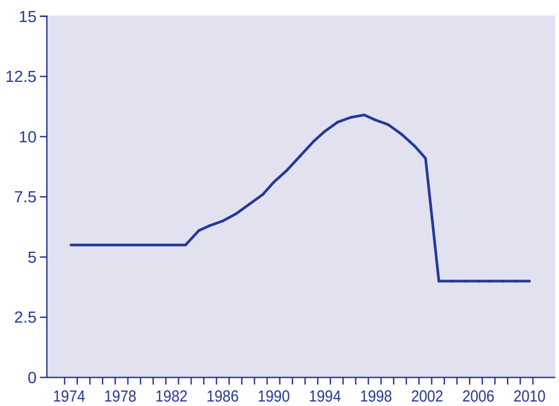
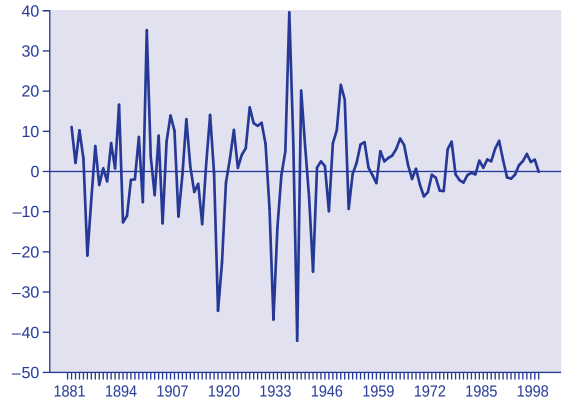


Figure 4

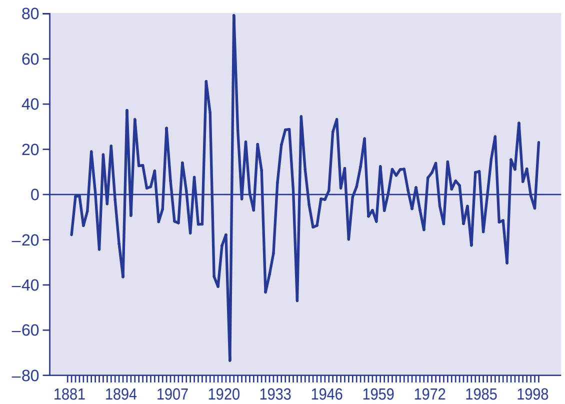
Annual Real Dividend Growth, 1881–1999

Percent



Annual Real Earnings Growth, 1881–1999

Percent



ized returns to differ significantly from the required return, reconciling a decline in the future discount rate with (temporarily) high current returns.

ARE CHANGES IN LONG-RUN DIVIDEND GROWTH AND DISCOUNT FACTORS PLAUSIBLE?

As we suggested above, for market fundamentals to explain the high stock prices of the 1990s, either expectations of future real dividend growth must have risen, future real discount rates fallen, or both. In this section, we examine these possibilities.

Has Expected Long-Run Real Dividend (Earnings) Growth Increased?

Shiller has argued that historical movements in dividends and earnings are too smooth for expectations of future real dividend growth to explain movements in stock prices. He writes, "Fluctuations in stock prices, if they are to be interpretable in terms of the efficient markets theory, must instead be due to new information about the long-run outlook for real dividends. Yet in the entire history of the U.S. stock market we have never seen such fluctuations, since dividends have fairly closely followed a steady growth path" (Shiller 2000, 188).

A cursory examination of Figure 4 reveals little evidence of large permanent changes in either real dividend growth or real earnings growth. More formal statistical measures also indicate little evidence of permanent changes. In particular, a standard augmented Dickey–Fuller unit root test, which tests whether shocks have a permanent effect, rejects the hypothesis of permanent shocks to real dividend (or earnings) growth over the period 1881–1999.[12] Similarly, the variance ratio statistic, $\mathrm{var}(g_{t+k} - g_t)/[\mathrm{var}(g_{t+1} - g_t)k]$, which provides a rough measure of the fraction of total variance due to permanent shocks, yields a value close to zero for real dividend growth and real earnings growth. The value for both is 0.11 when the horizon, k, is fifteen years.

However, neither of these statistics necessarily rules out the presence of a small permanent component in real dividend (earnings) growth. It is well known that unit root tests can mistakenly reject the hypothesis of permanent shocks too frequently when the permanent component of a time series is quite small.[13] In addition, a variance ratio of 0.11 in real dividend growth still allows for a small permanent component in real dividend growth. In fact, the estimate of the mean real dividend (earnings) growth over our full sample, 1881 through 1999, is fairly imprecise, with a 95 percent confidence interval of −0.97 to 3.08 (for earnings, −2.11 to 5.02). This suggests that modest increases in long-run expected real dividend and real earnings growth are not necessarily outside the range of historical experience. Indeed, Barsky and DeLong (1990, 1993) argue that actual dividend growth contains enough persistence that small permanent changes in expectations of long-run real dividend growth can explain long swings in stock prices, in contrast to Shiller's (1981) assertion that stock prices appear to move too much relative to dividends. In previous work (Balke and Wohar forthcoming), we also found that small changes in expectations of long-run real dividend growth are consistent with historical real dividend growth data.

Recent real dividend and earnings growth seems to warrant some optimism about future expected real dividend growth. As pointed out above and noted in Table 1, there is evidence of an increase in real dividend and real earnings growth after 1983. In particular, average real dividend growth was 2.43 percent for 1995 through 1999, while real earnings growth over the same period averaged 6.71 percent. While it remains to be seen whether the higher growth rates since 1983 are permanent, they are consistent with an increase in optimism about future real dividend and real earnings growth.

There may be economic grounds for justifying higher expectations of long-run real dividend (earnings) growth. Advocates of the New Economy argue that the revolution in computer and software technology is transforming the economy, ushering in an era of dramatic new productivity growth (see Hobijn and Jovanovic 2000, for an examination of the implications of the information technology revolution on stock price for new and incumbent firms). This technological progress will increase the productivity of capital (and labor), which would likely increase the income flow (dividends) from owning capital. Perhaps the high productivity growth seen in the late 1990s and early 2000s signifies that the new information technology is finally bearing fruit. Again, while it is not clear that the increase in productivity will persist indefinitely, it does open the possibility that expectations of higher dividend (earnings) growth can be supported by greater growth in real income.[14]

Other factors such as more capital-friendly tax policy, economic deregulation, and financial and technological innovation may have also increased optimism about future expected real dividend growth. Indeed, corporate profits began to rebound in the early 1980s, rising from 3.5 percent of GDP in mid-1982 to around 6 percent in 1999. Second, changes in expectations about inflation may have played an important role in the increased optimism seen since the early 1980s, because consumer price inflation fell from more than 10 percent during the late 1970s and early 1980s to under 5 percent in 1983, then fell further in the 1990s.[15]

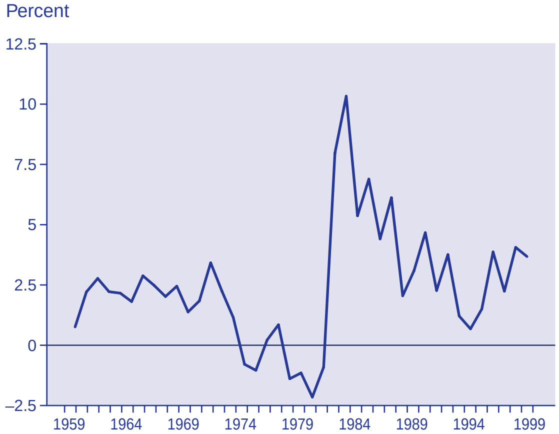
Has the Discount Rate Declined?

As we noted above, the real discount rate reflects the weight investors place on future dividend income when determining the value of stocks. A lower real discount rate suggests investors place higher value on future dividend income relative to current income and implies higher stock prices. Campbell and Ammer (1993) argue that the major factor causing movements in stock prices is movements in the discount rate. Campbell and Cochrane (1999) develop a theoretical model justifying these findings.

We can decompose the discount rate into two components, the real short-term interest rate and an equity premium. The real short-term interest rate reflects primarily factors such as investors' desire for future consumption relative to current consumption, or households' willingness to save, and the demand for capital. Thus, demographic factors—such as baby boomers entering their peak savings years, increases in life expectancy, and reduction in the threat of nuclear war—could result in a decline in the discount rate through a decline in the real interest rate. Because, all else equal, investors prefer less risky investments, they will discount riskier investments at a higher rate than safe investments. This component of the discount rate is called the risk or equity premium.[16]

Figure 5 plots the real short-term interest rate since 1959. The rate has fallen since 1983 but not below its historical average, so it's unlikely to have contributed to a historically low real discount rate. If real interest rates are not substantially lower, is there evidence of a de-

Figure 5
Real Short-Term Interest Rate, 1959–99



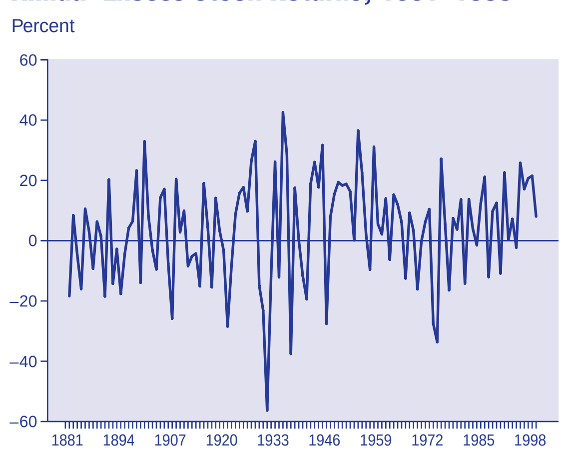
cline in the equity premium? Examining recent stock returns in excess of the returns on bonds will not help us ascertain whether the equity premium has fallen, for, as we saw above, excess stock returns may temporarily rise if the future equity premium declines. However, looking at excess stock returns over a long period may show whether a permanent change in the equity premium has some historical precedent. Figure 6 plots excess real return on stocks. As with real dividend growth, little persistence is seen in excess returns, as one can reject the hypothesis of a unit root in excess returns at the conventional significance level. Similarly, the variance ratio for excess returns is equal to 0.1 when the horizon is fifteen years, again suggesting a relatively small permanent component.

Yet, as we argued above, neither of these statistics rules out small permanent changes in excess returns. The extreme volatility of short-run movements in excess returns can statistically mask much more modest permanent changes. Indeed, the 95 percent confidence interval for the mean excess return (see Table 1) is quite large due to the volatility of excess returns themselves and includes values for the equity premium similar to those implied by our examples above.

While statistical evidence neither confirms nor rules out a decline in the equity premium, economic arguments may be more compelling. Siegel (1999) has estimated that a decline in transaction costs as well as the availability of low-cost index funds (primarily concentrated in equities) has lowered the cost of holding highly diversified portfolios. He estimates the decline in the equity premium, and hence, the discount rate, to be about 2 percentage points.[17] Similarly, Heaton and Lucas (2000), using an overlapping-generation model and a calibrated Gordon growth model, find that increased diversification has lowered required returns by about 2 percentage points and can explain at least 50 percent of the increase in the P/D ratio.[18] They argue that a typical investor used to hold a poorly diversified portfolio consisting of only a few stocks but, with the growth of mutual funds and index funds over the past two decades, is now much better diversified. Jagannathan, McGrattan, and Scherbina (2000) define the equity premium as the difference between the yield on a well-diversified stock portfolio and the yield on a long-term government bond. They find that an equity premium calculated in this way averaged 6.8 percent over the period 1926–70 but, since the 1970s, has fallen to just 0.7 percent.

Alternatively, perception about the riskiness of stocks may have changed. Many of today's investors have had no experience with the bear market of the 1970s, let alone the Great Depression. These investors' perception is much different from that of investors whose attitudes were shaped by the earlier periods of low real returns. Some who have argued that the stock price increase was the result of a decline in required returns maintain that investors may have become smarter and more relaxed about the stock market. Glassman and Hassett (1999) go so far as to argue that the 1990s run-up in stock prices was due in large part to investors having learned that a diversified stock portfolio has generally dominated government bonds over the long term. Since 1926, for any twenty-year period, stock returns have always exceeded returns on U.S. Treasury bonds. Furthermore, Glassman and Hassett (1999) point to the his-

Figure 6
Annual Excess Stock Returns, 1881–1999



torical evidence that stock returns over forty-year holding periods are less variable than returns on government securities. Thus, they argue that for long-term holding periods, equities produce returns higher yet no riskier than bonds. The implication is that investors now require a smaller equity premium to induce them to hold equities. Of course, as we pointed out above, the simple Gordon model implies that the expected future real return on stocks is equal to the discount rate. If the discount rate has fallen, say due to a decline in the equity premium, one can expect future real returns on stocks to be lower than they have been historically. This could conceivably alter the relationship between stock and bond returns that Glassman and Hassett have used to justify a decline in the equity premium.[19]

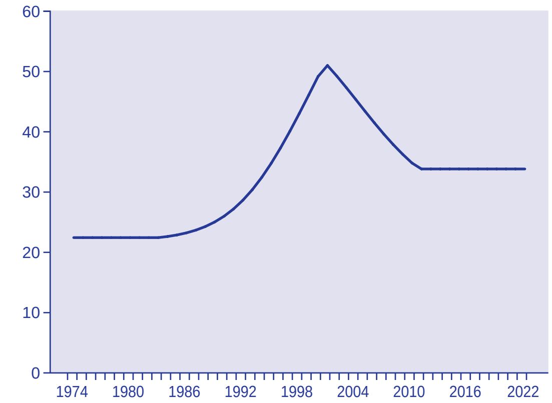
CAVEATS AND COMMENTS

In our discussion about the role of market fundamentals we have focused on expected real dividend (earnings) growth and real discount rates. But other issues come into play as well. One factor that can affect the P/D ratio is change in corporate financial policy. For example, the fact that many firms repurchase shares of their stock strengthens the market fundamentals argument. Repurchases represent an alternative to dividend payments as a form of stockholder compensation.[20] When a firm repurchases shares at the expense of current dividend payments, it reduces the number of shares outstanding and, in turn, increases future (but not current) dividends per share and, hence, the current share price. Thus, some have argued that the high P/D ratios are the result of stock repurchases. However, estimates of this effect are relatively small. For example, Cole, Helwege, and Laster (1996) adjust the dividend–price ratio by adding net repurchases (the difference between dollars spent on repurchases and dollars received from new issues) to dividends. They did this for the S&P 500 index over the period 1975–95 and found that the dividend–price ratio should be adjusted upward (and the P/D ratio downward) during the 1980s and 1990s. In particular, they found for 1995, the last year in their study, the adjusted P/D ratio is lowered to around 33 from 45, still well above its historical average.[21]

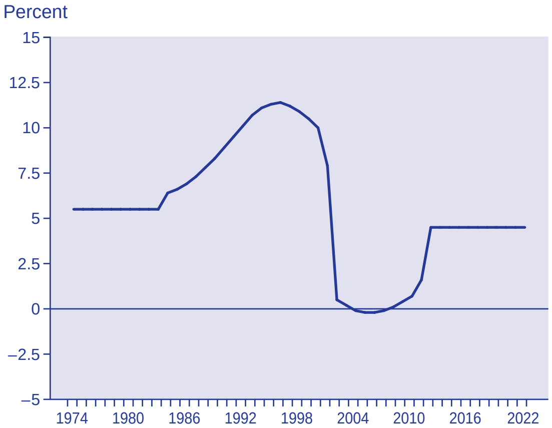
On the other hand, our previous discussion ignored the possible link between expected real dividend growth and the real discount rate. If expected real dividend growth increases due to an increase in productivity growth, people will also have even greater resources for future consumption (relative to current consumption), which makes current consumption more valuable relative to future consumption (current consumption is relatively more scarce) and increases the discount rate. Thus, in a general equilibrium analysis, an increase in expected real dividend growth would be accompanied by an increase in the real discount rate so that these two effects offset one another.

During the spring of 2001, after we wrote the first draft of this article, stock prices in certain sectors, particularly the high-tech sector, declined dramatically. Furthermore, the terrorist attacks on New York City and Washington, D.C., on September 11 rocked the broad market index examined in this article, the S&P 500, which had been largely spared the dramatic declines experienced by the tech stocks. Note that the simple market fundamentals Gordon model can explain stock price declines as well as stock price increases. One possible scenario that can rationalize recent stock price movements is one in which investors revise downward their expectations of future real dividend growth and increase their perceptions of the riskiness of stocks. The prospect of an economic slowdown, the direct and indirect losses from the destruction in New York and Washington, D.C., and the disruption of key industries such as air transportation could lead investors to lower their expectations of future dividend growth. Similarly, increased geopolitical and economic uncertainty may have caused investors to demand a higher risk premium for stocks. Thus, consider a scenario in which investors gradually expect real dividend growth to rise to 2 percent per year over the twenty-year period 1983–2003 and the discount rate to fall to 4 percent. In 2001, however, new information causes investors to revise their expectations of real dividend growth down to 1.5 percent and the discount rate up to 4.5 percent. As with the run-up in stock prices before 2001, this second set of revisions in expectations occurs gradually over a ten-year period. Figure 7 presents the resulting time path for the P/D ratio and the realized returns on stocks. Initially, stock prices rise as investors' expectations of real dividend growth increase and the discount rate falls, but once investors change their beliefs about future dividend growth and discount rates a second time, stock prices begin to fall. While expectations are being revised downward, stock returns dip below their historical average, then gradually increase to the long-run discount rate of 4.5 percent. Again, this example is not meant

Figure 7
P/D Ratio for Revised Expectations of Dividend Growth



Realized Returns for Revised Expectations of Dividend Growth



to be a literal description of stock price movements in 2001, but it does illustrate the power of the Gordon model to generate both stock price decreases and increases.

SUMMARY AND PROSPECTS FOR FUTURE RETURNS

A number of explanations have been offered for the unprecedented rise in stock prices during the 1990s. These include increased expected future economic growth brought about by the information revolution, demographic changes as the baby boomers age, a reduction in the equity premium as a result of lower transaction costs and increased diversification, a decline in inflation, momentum investing, and irrational exuberance. This article presents evidence that the case for market fundamentals, as an explanation for higher stock prices, is stronger than it appears on the surface. We demonstrate that movements in the price–dividend and price–earnings ratios show substantial persistence, particularly since World War II. Hence, using the long-run historical average value of the price–earnings or price–dividend ratio as the "normal" valuation ratio might be misleading. We also show that plausible combinations of lower expected future real discount rates and higher expected real dividend (earnings) growth could rationalize recent broad market stock values, raising the possibility that changes in market fundamentals have had a major contribution to the run-up in stock prices.

Whether market fundamentals or irrational exuberance was responsible for high stock prices during the 1990s, the prospect for future stock returns is not so sanguine. Both a bursting bubble and a declining future equity premium imply lower future returns than those seen in the recent past, and indeed, lower than what has been averaged historically. Only if expected real dividend growth has risen permanently can one reasonably expect future stock returns to remain at their historical average.

NOTES

We gratefully acknowledge very helpful comments from John Duca and Alan Viard on a previous version of this paper.

1 On December 3, 1996, Robert Shiller informed the Board of Governors of the Federal Reserve that the stock market was overvalued. His paper based on that testimony, "Valuation Ratios and the Long-Run Stock Market Outlook: Ratios Are Extraordinarily Bearish," may have inspired Alan Greenspan's "irrational exuberance" statement two days later. The Dow Jones Industrial average closed on that day at a value of 6,437. Campbell and Shiller (2001) update the results presented in Shiller's 1996 testimony.

2 Reduced business-cycle risk may be the result of better Fed (forward-looking) policy, better inventory control, and better information. Investors that follow momentum-investing strategies base their investment decisions on recent movements of stock prices, which until recently had been trending upward. Over short periods, investment strategies to exploit increases in stock prices can be profitable. However, over long periods, momentum strategies can be profitable only if supported by fundamental factors. A theoretical justification for momentum-investment strategies can be found in Hong and Stein (1999).

3 Campbell and Shiller (1998, 2001) follow the suggestion of Graham and Dodd (1934) and use smoothed earnings over the past ten years. They find that the smoothed P/E ratio behaves more like the P/D ratio than does the traditional P/E ratio.

4 The P/D ratio is calculated using accumulated dividends over the past 12 months. For example, the P/D

ratio for 1999 is calculated as the January 2000 stock price divided by the December 1999 dividend value. Dividends in December 1999 are accumulated dividends over the year 1999. For more details see Shiller (2000, chapter 1, footnote 2).

[5] Although for some short periods the P/E ratio increased slightly above its mean and then quickly fell below, we focus on persistent crossings. Data on stock prices, dividends and earnings are those employed in Shiller (2000). For details on the data, see Shiller (2000, chapter 1, footnote 2). These data can be downloaded from Robert Shiller's web site (http://www.econ.yale.edu/~shiller).

[6] For the ln(P/D) ratio, the augmented Dickey–Fuller test statistic (with a time trend and a lag length of 3) is –1.83 over the period 1881–1999. Using the ln(P/E) ratio yields a test statistic of –2.24 over the same period. The 5 percent critical value is –3.45.

[7] The term "discount rate" used here is *not* the interest rate the Fed charges financial institutions for loans.

[8] The Newey–West procedure is used to correct for serial correlation (lags = 3). The 95 percent confidence interval is approximately +/– two times the standard error of the mean.

[9] Nominal returns for 1999 are calculated as $\ln[(P_{Jan.\ 2000} + D_{Dec.\ 1999})/P_{Jan.\ 1999}]\bullet 100$, where $P_{Jan.\ 2000}$ is the nominal stock price in January of the year 2000, $D_{Dec.\ 1999}$ is the accumulated nominal dividends over the year 1999. Real returns are computed as nominal returns minus inflation, inflation is calculated as $\ln(CPI_t/CPI_{t-1})\bullet 100$, where CPI_t is the value of the CPI in January of year t. Nominal dividends and nominal earnings are deflated by the CPI to obtain real values. Real dividend and real earnings growth are computed as $\ln(RD_t/RD_{t-1})\bullet 100$ and $\ln(RE_t/RE_{t-1})\bullet 100$, where RD_t and RE_t are real dividends per share and real earnings per share, respectively, in January of year t. Real excess returns are computed as the log of real stock returns minus the real short-term interest rate. The real short-term interest rate is defined as the real return on a six-month commercial paper (rolled over midyear). We would like to thank Eugene Fama for providing the real short-term interest rate series.

[10] The period 1983 through 1999 in Table 1 includes the recovery from the deepest post-World War II recession. To see if starting from a peak of the business cycle leads to different results relative to starting from the trough (as we do here), we also examined values of the variables in Table 1 for 1982 through 1999. The only variable that changes to any substantial degree is real earnings growth, which decreases to 3.29 percent for 1982–99. However, this value is almost three times the average of real earnings growth for 1881–1981 (1.13 percent).

[11] No special significance should be attached to the date 2003. It only reflects an even twenty years after the beginning of the post-1982 bull stock market.

[12] For real dividend growth, the augmented Dickey–Fuller test statistic (with a time trend and a lag length of 3) is –6.41 over the period 1881 through 1999. The value for earnings growth is –7.13. The critical value is –3.45.

[13] In the formal language of statistics, the standard tests can suffer from a size distortion when permanent shocks are small relative to temporary shocks. See Schwert (1987) for an analysis of size properties of standard unit root tests.

[14] Of course, if the late 1990s productivity growth is temporary rather than permanent, real dividend and real earnings growth cannot be sustained at current levels.

[15] Sharpe (1999) finds that forecasts of nominal earnings growth have increased slightly over the past decade while at the same time inflation has fallen, implying a substantial increase in forecasted real earnings.

[16] For a discussion of changes in the equity premium over time, see Blanchard (1993), Jagannathan, McGrattan, and Scherbina (2000), and Fama and French (2001). For a discussion of changes in real interest rates over time, see Blanchard and Summers (1984).

[17] Rea and Reid (1998) find that the sales-weighted average of total shareholder costs for equity mutual funds decreased from 2.25 percent in 1980 to 1.49 percent in 1997. Duca (2000) also documents a decline in transaction costs for equity mutual funds. This decline may have also been a factor contributing to increased diversification.

One can modify the Gordon formula accounting for transaction costs so that the P/D ratio is equal to $(1 - \tau)(1 + g)/[r + \tau - (1 - \tau)g]$, where r is the required return net of transaction costs and τ is the fraction of gross returns lost to transaction costs. Thus, a decline in τ affects the P/D ratio in a manner similar to a decline in the real discount rate (r) and/or an increase in expected real dividend growth (g).

[18] Heaton and Lucas (2000) also examine whether an increased number of people participating in the stock market could contribute to a decline in the discount rate. They find that increased participation has only small effects on the discount rate.

It is not clear that diversification can account for such a large decline in the equity premium. One problem is that Heaton and Lucas' measure of diversification is not weighted by wealth. The majority of stocks are held by wealthy individuals, who probably always had well-diversified portfolios. Thus, increased diversification by small investors may not have a large effect on stock prices. Another issue is why investors were underdiversified in the past.

[19] Investor surveys that indicate continued high future returns for stocks are not consistent with a decline in the discount rate. Whether these surveys reflect merely projections of recent trends in stock returns or are the rational expectation of future returns is an open question.

[20] The P/E ratio scaled by a ten-year moving average of earnings is often a more stable proxy for long-run payments to shareholders.

[21] Their analysis assumes that shares are issued and repurchased at the market price. If shares are issued at below-market prices (say, as part of executive compensation), the true repurchase effect is smaller.

REFERENCES

Balke, Nathan S., and Mark E. Wohar (forthcoming), "Low Frequency Movements in Stock Prices: A State Space Decomposition," *Review of Economics and Statistics*.

Barsky, Robert B., and J. Bradford DeLong (1990), "Bull and Bear Markets in the Twentieth Century," *Journal of Economic History* 50 (June): 265–81.

——— (1993), "Why Does the Stock Market Fluctuate?" *Quarterly Journal of Economics* 108 (2): 291–311.

Blanchard, Olivier J. (1993), "Movements in the Equity Premium," *Brookings Papers on Economic Activity*, no. 2: 75–138.

Blanchard, Olivier J., and Lawrence H. Summers (1984), "Perspectives on High World Real Interest Rates," *Brookings Papers on Economic Activity*, no. 2: 273–334.

Campbell, John Y., and J. Ammer (1993), "What Moves the Stock and Bond Markets? A Variance Decomposition for Long-Term Asset Returns," *Journal of Finance* 48: 3–37.

Campbell, John Y., and John H. Cochrane (1999), "By Force of Habit: A Consumption-Based Explanation of Aggregate Stock Market Behavior," *Journal of Political Economy* 107 (April): 205–51.

Campbell, John Y., and Robert J. Shiller (1998), "Valuation Ratios and the Long-Run Stock Market Outlook," *Journal of Portfolio Management* 24 (Winter): 11–26.

——— (2001), "Valuation Ratios and the Long-Run Stock Market Outlook: An Update," NBER Working Paper Series, no. 8221 (Cambridge, Mass.: National Bureau of Economic Research, April).

Carlson, John B. (1999), "The Recent Ascent in Stock Prices: How Exuberant Are You?" Federal Reserve Bank of Cleveland *Economic Commentary*, August 15.

Carlson, John B., and Eduard A. Pelz (2000), "Investor Expectations and Fundamentals: Disappointment Ahead?" Federal Reserve Bank of Cleveland *Economic Commentary*, May 1.

Carlson, John B., and Kevin H. Sargent (1997), "The Recent Ascent of Stock Prices: Can It Be Explained by Earnings Growth or Other Fundamentals?" Federal Reserve Bank of Cleveland *Economic Review* 33, Second Quarter, 2–12.

Cochrane, John H. (1997), "Where Is the Market Going? Uncertain Facts and Novel Theories," Federal Reserve Bank of Chicago *Economic Perspectives*, November/December, 3–37.

Cole, Kevin, Jean Helwege, and David Laster (1996), "Stock Market Valuation Indicators: Is This Time Different?" *Financial Analysts Journal* (May/June): 56–64.

Duca, John (2000), "Mutual Fund Loads and the Long-Run Stock Wealth Elasticity of Consumption," unpublished manuscript, Federal Reserve Bank of Dallas, June.

Fama, Eugene F., and Kenneth R. French (2001), "The Equity Premium," Center for Research in Security Prices, Graduate School of Business, University of Chicago, Working Paper no. 522 (Chicago, April).

Glassman, James K., and Kevin A. Hassett (1999), *Dow 36,000: The New Strategy for Profiting from the Coming Rise in the Stock Market* (New York: Times Books).

Gordon, Myron J. (1962), *The Investment, Financing, and Valuation of the Corporation* (Homewood, Ill.: Irwin Publishers).

Graham, Benjamin, and David L. Dodd (1934), *Security Analysis* (New York: McGraw-Hill).

Heaton, John, and Deborah Lucas (2000), "Stock Prices and Fundamentals," in *NBER Macroeconomic Annual 1999*, ed. Ben S. Bernanke and Julio Rotemberg (Cambridge, Mass.: MIT Press for National Bureau of Economic Research), 213–63.

Hobijn, Bart, and Boyan Jovanovic (2000), "The Information Technology Revolution and the Stock Market: Evidence," New York University working paper (New York, December).

Hong, Harrison, and Jeremy C. Stein (1999), "A Unified Theory of Underreaction, Momentum Trading, and Overreaction in Asset Markets," *Journal of Finance* 54 (December): 2143–84.

Jagannathan, Ravi, Ellen R. McGrattan, and Anna Scherbina (2000), "The Declining U.S. Equity Premium," Federal Reserve Bank of Minneapolis *Quarterly Review*, Fall, 3–19.

Kopcke, Richard W. (1997), "Are Stocks Overvalued?" Federal Reserve Bank of Boston *New England Economic Review*, September/October, 21–40.

McGrattan, Ellen R., and Edward C. Prescott (2000), "Is the Stock Market Overvalued?" Federal Reserve Bank of Minneapolis *Quarterly Review*, Fall, 20–40.

Rea, John D., and Brian K. Reid (1998), "Trends in the Ownership Cost of Equity Mutual Funds," Investment Company Institute *Perspective* 4 (November): 1–15.

Schwert, G. William (1987), "Effects of Model Specification on Tests for Unit Roots in Macroeconomic Data," *Journal of Monetary Economics* 20 (July): 73–103.

Sharpe, Steven A. (1999), "Stock Prices, Expected Returns, and Inflation," Finance and Economic Discussion series, no. 99-2 (Washington, D.C.: Board of Governors of the Federal Reserve System, Division of Research and Statistics, January).

Shiller, Robert J. (1981), "Do Stock Prices Move Too Much to Be Justified by Subsequent Changes in Dividends?" *American Economic Review* 71 (June): 421–36.

——— (2000), *Irrational Exuberance* (Princeton, N.J.: Princeton University Press).

Siegel, Jeremy J. (1999), "The Shrinking Equity Premium: Historical Facts and Future Forecasts," *Journal of Portfolio Management* 26 (Fall): 10–17.