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A Conference on Price Stability 2
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In November 1990, the Federal Reserve Bank of Cleveland sponsored a Conference on Price Stability. The participating economists were asked to explain how recent developments in macroeconomic research have changed the way we think about optimal inflation policy. In particular, the discussions center on the implications for the optimal inflation trend (the Friedman–Phelps debate) and for the degree of variability of inflation around the trend. The six papers presented at the conference, though limited in practical policy dimensions, provide a detailed evaluation of the issues surrounding the Friedman–Phelps debate.

Components of City-Size Wage Differentials, 1973–1988 10
by Patricia E. Beeson and Erica L. Groshen

It has long been noted that workers in large cities are more highly paid than their rural counterparts. In studying this phenomenon, most researchers control for differences in work-force attributes between cities, but do not investigate whether these attributes are priced differently in cities of varying size. This paper examines the empirical regularity in the 1973–1988 Current Population Surveys to see if city-size-related wage differentials arise from intercity differences in wage structures. The authors find that higher economic rewards for education, experience, and other skills in larger cities account for the bulk of the earnings disparity, and that recent changes in these differentials mainly reflect diverging returns to skills.

Financial Efficiency and Aggregate Fluctuations: An Exploration 25
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Changes in the efficiency of the financial system can greatly affect the overall economy. A simple real business cycle framework shows how banks can be a source, rather than just a filter, for output shocks. This paper develops and tests predictions about cointegration between several bank and output series, as well as explores the comovement of output and combined banking variables. Use of the vector error-correcting model provides additional information on the role of banks as both transmission mechanisms and originators of cyclical disturbances.

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A Conference on Price Stability

by Charles T. Carlstrom and William T. Gavin

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Introduction

The Federal Reserve Bank of Cleveland has publicly endorsed a legislative resolution that would make zero inflation, or price-level stability, the overriding long-run objective of the Federal Reserve System. Price-level stability means a policy intended to keep the price level reasonably constant over long horizons. That is, although the price level may fluctuate temporarily in response to transitory real output and money demand shocks, permanent shifts in money demand and in real output would be offset in order to ensure that the long-run price level is stationary.

The term zero inflation is commonly used as a synonym for price-level stability. Care must be taken to distinguish this usage from a monetary growth rule that would result in zero expected inflation in the long run, but that could cause wide deviations of the price level from a constant trend as the income velocity of money shifts. Price-level stability should also be distinguished from a zero-inflation target in which the central bank tries to achieve zero inflation in every month or quarter. Such a strategy allows target misses to accumulate, which in turn allows the price level to follow a random walk.

In principle, a price-level rule would not preclude temporary policies to stabilize the business cycle. In practice, such a rule would remove uncertainty about the long-run price level and would be, in our judgment, the best policy to ensure sustained economic growth. Admittedly, economists have not offered much formal analysis in support of this issue.

Policymakers, however, are often required to make choices regardless of whether academic debates have been fully resolved. In order to advance the research about the role of inflation in the design of monetary policy, the Federal Reserve Bank of Cleveland sponsored a Conference on Price Stability in November 1990. The full proceedings of the conference appear in the August 1991 issue of the *Journal of Money, Credit, and Banking.*[1] In particular, we asked the participating economists to explain what recent developments in macroeconomic research have taught us about the optimal inflation policy.

The most widely cited origin of the optimal inflation literature is a 1969 article by Milton

1 Authors cited in this paper without a referenced year were conference participants, and complete citations of their papers can be found in the reference list.

Friedman, which explains why optimal inflation is achieved by allowing the money supply to grow (shrink) at a rate that results in a zero nominal interest rate. Here, Friedman estimates that this result would be achieved if the money supply were forced to grow at a relatively constant 2 percent rate. Assuming that the marginal cost of fiat money is zero, a zero nominal interest rate leads people to hold the optimal quantity of cash balances. This policy has become known as the "Chicago rule."

The Chicago rule is based on the assumption that government can raise revenue with a nondistorting tax. In response, Phelps (1973) presents a model in which only distorting taxes are available to replace the seigniorage lost when eliminating a moderate inflation. He shows that the trend inflation rate should be chosen so that the marginal welfare cost of the last dollar raised through inflation is equal to the marginal welfare cost of the last dollar raised through other taxes. A large body of literature debates the relevance of Phelps' criticism of Friedman's rule.[2] Whether Friedman's rule is optimal or whether some positive inflation rate would be preferred depends importantly on the role of money in the economy (medium of exchange, store of value, or unit of account) and on the alternative revenue sources available to the government. Resolving the issue requires considerably more detail about the economy than Friedman and Phelps presented in their respective partial-equilibrium models.

Our discussion of the six papers presented at the Conference on Price Stability is organized around two policy issues. The first is the implication of the papers for the optimal inflation trend — the Friedman–Phelps debate; the second is the implication for the degree of variability of inflation around the trend. We begin section I with a description of the Chari, Christiano, and Kehoe paper, which derives the optimal mix of monetary and fiscal policy rules. Henning Bohn's evaluation of the sustainability of fiscal policies in a stochastic environment follows naturally from this discussion. Section II includes an extended explanation of the Cooley and Hansen paper, which attempts to measure the net welfare cost of policies that would reduce inflation to zero and replace the lost seigniorage with higher taxes on capital or labor. Altig and Carlstrom then focus on the interaction of inflation with the nominal tax system. Section III describes the Imrohoroglu and Prescott paper, which calculates the efficiency of the seigniorage tax under alternative institutional arrangements. Section IV discusses Laurence Ball's explanation of why attempts to eliminate inflation almost always lead to recession. In section V, we conclude with a summary of the policy implications that might be drawn from these studies.

I. Optimal Inflation and Nonindexed Government Debt

In "Optimal Fiscal and Monetary Policy: Some Recent Results," V.V. Chari, Lawrence Christiano, and Patrick Kehoe derive the optimal inflation policy jointly with the optimal tax policy. Using a stochastic, equilibrium business cycle framework, they work out the optimal labor and capital tax policies in a two-factor production economy, and the optimal inflation tax in a one-factor production economy without capital. They follow Lucas and Stokey (1983) in assuming a cash-in-advance role for money; some goods are purchased with cash and others with credit.

Chari et al. present four interesting results from the analysis of these models. First, they show that the optimal policy implies that either government debt or capital taxes should be indexed to government consumption and productivity shocks. For example, during a war or some positive shock to government consumption, the value of government debt should decline. This enables the government to avoid raising taxes during a war. They argue that because debt is difficult, if not impossible, to index, monetary policy should be used to provide the appropriate ex post real payments. Over the business cycle, policy should be set so that money varies positively with government consumption and negatively with productivity shocks.

Contrary to Barro (1979), they find that optimal tax rates on labor do not follow a random walk, but inherit the persistence properties of the underlying shocks and, for practical purposes, are roughly constant over the business cycle. The difference seems to arise because Barro does not allow government the right to issue state-contingent debt. Henning Bohn, in the paper discussed later in this section, shows that tax smoothing is, in general, not sustainable in the absence of state-contingent government debt.

Third, Chari et al. show that capital taxes should be zero on average in a stochastic steady state. This extends the results of Chamley (1986), who shows that the optimal tax on capital income is zero in a deterministic steady state. Although Chari et al. indicate that steady-state capital taxes are zero, state-by-state capital taxes

2 Woodford (1990) summarizes much of this literature and gives a detailed analysis of the assumptions underlying Friedman's rule.

are not uniquely determined in a model with state-contingent government debt. Equivalently, the optimal level of indexing for government debt is not uniquely determined given state-contingent capital taxation.

Finally, they analyze optimal monetary policy in an economy without capital. In this model, Friedman's rule holds even in the presence of distorting taxes: A zero nominal interest rate prevails in every period. The intuition is simply that the optimal policy taxes both credit and cash goods at the same rate. This occurs when all revenue is raised through a tax on labor income. Without capital, a wage tax is equivalent to a consumption tax, where both cash and credit goods are taxed at the same rate.

In his comments on Chari et al.'s paper, R. Anton Braun shows that the optimality of Friedman's rule also depends on the particular form of preferences chosen; that is, utility is homogeneous in degree k with respect to both types of goods. Also, leisure is weakly separable from both cash and credit goods. Braun argues that empirical evidence from money-demand studies does not support this preference structure. Instead, he contends that the cash good is weakly separable from both the credit good and leisure, and that utility is homogeneous in degree k in leisure and in the credit good. This form of utility implies that wage taxes should equal zero and that efficient tax structure includes positive inflation.

If government debt cannot be indexed to either government consumption or productivity shocks, and if capital taxes cannot be made contingent on these shocks, then monetary policy can be used to effectively index the real return on fixed nominal government debt. Although the optimal policy results in a zero nominal interest rate, Friedman's result that money should be deflated each period at the rate of time preference is far from optimal.

Ex ante, the expected level of deflation in each period will equal the real interest rate. Thus, with the nominal interest rate on government debt being equal to zero in every period, the expected amount of deflation will be set such that, ex ante, the rate of return on all assets will be equal (although the Chari et al. monetary model does not actually include private capital). However, because monetary policy is being used to effectively index government debt, ex post optimal monetary policy results in an inflation rate that has a very large variance around a near-zero trend. Thus, while seigniorage should not be part of an optimal tax policy, substantial amounts of inflation and deflation should exist in order to decrease the real value of government debt during wars or negative technology shocks and to increase the real value of government debt during booms.

Henning Bohn also discusses the issue of indexed government debt in "The Sustainability of Budget Deficits with Lump-Sum and with Income-Based Taxation." Bohn uses a stochastic environment to analyze the feasibility of some commonly recommended fiscal policies. Although his paper is not directly aimed at the issue of price stability, it has important implications for analyzing alternative inflation policies. His method could be extended to examine the sustainability of specific monetary policies in a stochastic environment.

Bohn considers the sustainability of two fiscal policies, a tax-smoothing policy and a balanced budget, under two different assumptions about the tax collection mechanism, a lump-sum and an income-based tax. He defines these terms carefully and shows that although many economists have espoused such policies and shown them to be sustainable in a deterministic world, they are not necessarily sustainable in a stochastic environment. He argues that it is imperative to specify the complete general-equilibrium environment, including the incentives of taxpayers and the constraints on tax collectors, when examining these issues.

Bohn shows that both tax-smoothing and a balanced budget are typically sustainable if the government is able to levy lump-sum taxes, but not if taxation is limited by the amount of income. He also demonstrates that in an uncertain environment, state-contingent government debt can be used to design sustainable versions of tax-smoothing and balanced budget policies. Like Chari et al., Bohn argues that inflation policy might be used to index government debt. He shows that a tax-smoothing policy can be maintained in a stochastic environment if inflation is perfectly negatively correlated with real output. The optimal inflation rate in Chari et al. actually has very little correlation with changes in output due to government consumption shocks.

Together, these papers indicate that government debt should be indexed. In theory, monetary policy could be used to accomplish this. Lucas and Stokey (1983) argue that, at least in the years following wars, monetary policy has been used to retire a substantial amount of the real value of outstanding government debt. However, this ex post indexing has not occurred on the scale proposed by Chari et al. Inflation has rarely exceeded 10 percent in the United States, while their model proposes that a third of the time inflation should either be greater than 20 percent or lower than -20 percent.

II. The Efficiency of Seigniorage

Thomas Cooley and Gary Hansen measure the net welfare effects of a policy to eliminate inflation, starting from the current pattern of effective tax rates on labor and capital. They begin with a real business cycle model that includes an indivisible labor supply specification and a cash-in-advance role for money. Like Chari et al., they use the Lucas–Stokey setup with cash and credit goods. Their baseline model is calibrated to fit the postwar U.S. economy, and baseline tax rates are set at historical averages: a 23-percent effective tax on labor and a 50-percent effective tax on capital.

The first experiment is to reduce inflation from 5, 10, or 20 percent to zero. In their simulations, Cooley and Hansen show that the welfare costs of inflation are substantially larger than those estimated by Fischer (1981). In his comment on this paper, Roland Benabou uses the Cooley–Hansen general equilibrium model to show analytically that inflation is much more costly than is implied by the usual partial-equilibrium estimates (the area represented by the Harberger triangle under the demand curve for money).

Although Cooley and Hansen argue that the benefits of ending inflation are greater than previously thought, eliminating inflation actually makes people worse off, because, on the margin, keeping the inflation tax is more efficient than increasing the tax on either capital or labor income. In their model, Phelps' argument is correct. They show that eliminating inflation makes people slightly worse off when the lost revenue is replaced with higher taxes on labor income, and makes them much worse off when the lost revenue is replaced with higher taxes on capital. The marginal welfare cost of revenue raised from the capital tax is already so high that attempting to extract extra revenue from this source is very costly.

The policy conclusions drawn by Cooley and Hansen for their model economy are straightforward. At the margin, the inflation tax is clearly less burdensome than either labor or capital taxes. The government can make people better off by raising inflation and lowering other taxes, particularly the capital tax. This conclusion is in contrast to that reached by Chari et al., who find that Friedman's rule holds.

Both papers use a cash-in-advance specification for the role of money. The important difference is that Chari et al. exclude capital when analyzing optimal monetary policy, which makes their labor tax equivalent to a consumption tax. In both papers, inflation is a tax on cash goods, driving a wedge between cash and credit goods. With capital, an inflation tax shares the properties of a consumption tax in that it is a lump-sum tax on the existing capital stock. If Chari et al. extended their monetary model to include capital, Friedman's rule would no longer be optimal. Similarly, however, if either Cooley and Hansen or Chari et al. extended their papers to allow for a consumption tax that equally taxes both cash and credit goods, Friedman's rule of a zero nominal interest rate would again be optimal.

In Cooley and Hansen's model, the tax code is perfectly indexed for inflation. They assume that the effective tax rate on capital income is 50 percent, based on real-world studies in which the capital tax was not indexed for inflation. In reality, the effective capital tax rate was as high as 50 percent only because inflation averaged about 5 percent and because nominal rather than real returns were taxed. If the real rate of return to capital were 5 percent and the inflation rate were 5 percent, then a 25 percent tax on nominal capital income would be a 50 percent tax on real capital income. If the Fed actually went to zero inflation, the effective capital tax would decline unless Congress increased the tax rate. One can argue that Congress made such an adjustment in the early 1980s when inflation was reduced from 10 percent to 4 percent. The effective capital tax had been reduced in 1981, but was raised again by the Economic Reform Tax Act of 1983 and by the Tax Reform Act of 1986.

With nominal taxation of capital income, the Federal Reserve can lower the effective capital tax rate by lowering the inflation rate. Indeed, that is one of the arguments in favor of zero inflation. Cooley and Hansen generously agreed to run one further experiment in which the inflation rate was lowered to zero and the capital tax rate was cut to 25 percent to simulate the reduced capital tax that would occur with reduced inflation.[3] This experiment approximately represents the sort of policy change that would result if the Fed lowered the current inflation trend to zero and if Congress made up the lost revenue by raising the rates in the personal income tax code. The simulation shows that when the lost revenue from both seigniorage and the capital tax is made up with a higher wage tax, welfare increases by approximately 0.56 percent of GNP.[4]

■ **3** The assumed real rate is 4 percent in the Cooley and Hansen paper, instead of the 5 percent rate that was assumed in our calculations.

■ **4** Cooley and Hansen report that the steady-state welfare gain is approximately 2.6 percent of GNP.

In "Inflation, Personal Taxes, and Real Output: A Dynamic Analysis," David Altig and Charles Carlstrom analyze the interaction between inflation and the taxation of nominal capital income. They use a stochastic version of the overlapping-generations model developed by Auerbach and Kotlikoff (1987) to examine both steady-state and business cycle consequences of changes in capital tax rates induced by changes in the inflation rate. There is no explicit model of money; rather, inflation is introduced as exogenous changes in an arbitrary unit of account. This simple framework allows marginal tax rates to be endogenous in a world with a graduated income tax. The authors concentrate on the interaction between the personal tax code and the inflation rate and ignore distortions associated with corporate taxation of capital.

Altig and Carlstrom find that changes in the inflation trend have large effects on the steady-state capital stock and hence on output. They estimate that steady-state output is approximately 5 percent lower than it would be if inflation were eliminated or equivalently if capital income were indexed. Although the authors do not investigate the welfare consequences of reducing inflation, their subsequent simulations show that, in the steady state, a worker just entering the labor force would need a one-time compensation equal to 0.75 percent of his or her "full lifetime wealth" in order to compensate for having a 4 percent inflation rather than zero (or perfect indexation of capital income).[5]

The decrease in steady-state output in their model is almost entirely due to the negative effect that the nominal taxation of nominal capital income has on the capital stock. The rate of inflation has almost no influence on steady-state hours worked, because the substitution and income effects cancel. In contrast to their steady-state results, Altig and Carlstrom find that the variability of inflation has little effect on the cyclical variability of capital, but has a substantial impact on the cyclical behavior of labor.

Inflation affects the cyclical properties of hours worked for two reasons. First, it raises the effective tax rate on capital income and lowers the return on savings. This causes people to substitute intertemporally toward both consumption and leisure. Second, although wages are assumed to be indexed for inflation, capital income is not. Thus, rising inflation causes capital income to be overstated and, with a graduated income tax, throws savers into higher tax brackets. In this model, variable inflation has little effect on short-run fluctuations in real economic activity. Nevertheless, variable inflation increases the variability of hours worked and decreases the covariance between hours worked and output.

Altig and Carlstrom report that the short-run price instability typical of postwar U.S. history has had little impact on the real economy. Their analysis supports the notion that we should have a "zero expected inflation" target in order to effectively cut the tax rate on capital; however, in their model, there are no further gains from adopting an explicit path for the price level as a long-run objective. These observations are based on the reported real effects from variable inflation; deadweight loss calculations are not presented.

III. The Inefficiency of Seigniorage

In "Seigniorage as a Tax: A Quantitative Evaluation," Ayse Imrohoroglu and Edward Prescott examine the efficiency of the inflation tax under different assumptions about monetary institutions. The authors use a model with a simple one-factor production function and calibrate it using actual U.S. data. Instead of assuming that money is held for transaction purposes (as do Chari et al. and Cooley and Hansen), they specify a model in which money serves as a store of value to smooth consumption over time. There are two assets in the model, large-denomination government bonds and money. Banks take deposits from households and buy government bonds. The alternative to seigniorage is a tax on labor income. In each case, the revenue lost from reducing inflation is recovered with enough extra revenue from the labor tax to keep the government's budget balanced.

Imrohoroglu and Prescott consider three scenarios: (1) a case with 100 percent reserve requirements, in which currency is the only store of value, (2) a case in which monetary policy is uncertain from period to period, but everyone knows the trend inflation rate, and (3) a case with fractional reserves.

With 100 percent reserve requirements, only money is available to smooth consumption. Consequently, the welfare costs of inflation can be two to three times higher than the costs typically measured by the Harberger triangle under the money demand curve. Again, this is also the result that Benabou found in his analysis of Cooley and Hansen's paper.

5 Their simulations assume that the lost revenue from reducing inflation is replaced through a proportional increase in the income tax. Full wealth is defined to be the present discounted value of a person's wage income assuming that all nonsleeping hours are spent working.

Thus, in two separate models with two very different assumptions about the purpose of holding money, general equilibrium measures of the cost of inflation (non-revenue-compensated) are significantly greater than the partial-equilibrium measures. Contrary to Cooley and Hansen, however, Imrohoroglu and Prescott find that when income taxes are increased to compensate for lost seigniorage, Friedman's rule is again optimal. This result continues to hold when the model is modified to include fractional reserves. As expected, allowing interest-bearing assets to smooth consumption reduces the cost of inflation. Adding short-run uncertainty about monetary policy imposes no additional welfare costs in this model. The welfare costs of a variable inflation policy that results in an average 4 percent inflation are approximately identical to those where the price level grows at a constant rate of 4 percent every year.

IV. Inflation Policy and Recessions

None of the five papers discussed above addresses an important concern of traditional macroeconomics — whether attempts to end inflation have been a major cause of recessions. This issue is advanced by Laurence Ball, who asks why efforts to end inflation almost always seem to be associated with recessions. Ball accepts as a stylized fact not only that all disinflations have been associated with recessions, but also that, on average, *announcements* of disinflation have led to recessions. He argues that neither of two simple explanations alone — price stickiness as suggested by New Keynesian economists, nor the lack of credibility as suggested by New Classical economists — can explain these phenomena if expectations are rational.

Ball shows that nominal price rigidity, as represented in models with staggered price-setting arrangements, cannot explain why ending inflation causes recessions. By carefully distinguishing between changes in the growth rate of money and changes in the level of money, he explains how a credible disinflationary policy will lead to a boom in a model with staggered price setting. The intuition behind this argument is simple: If price setters expect inflation to decline, they will lower prices immediately because they can readjust prices only periodically. This immediate decline in the price level will lead to an increase in real balances and consequently in output. Thus, fixed-price models alone cannot explain why disinflation leads to recessions.

He then argues that the New Classical explanation — that disinflations cause recessions because policy announcements are not credible — also is incomplete. If policy is partly credible (that is, if the Fed announces disinflation and sometimes follows through), market-clearing models predict that announcements of disinflation will sometimes lead to recession. The average expectation will be that the money growth rate will decline, but not by as much as the central bank announces. Sometimes actual money growth will fall faster than the expectation (the economy will recede); other times it will not (the economy will expand). On average, in a New Classical model with partial credibility, there should be no correlation between announcements of disinflation and deviations of output from trend.

This stylized fact that, on average, announcements of disinflation lead to recession is based on the controversial definition of announcements contained in Romer and Romer (1989). They identify six such announcements of disinflation in the postwar period, drawing sharp distinctions about what constitutes an announcement where we would not. If one believes that there were many announcements of disinflation but only a few actions, then the probability of follow-through is small. In the New Classical model in which policy has almost no credibility for disinflation policies, there is no reason to expect that false announcements of disinflation would lead to recognizable booms. However, true announcements of disinflation would lead to recessions.

After making the point that neither New Keynesian nor New Classical ideas *alone* can explain why announcements of disinflation lead to recessions, Ball notes that the two assumptions together can explain why announced disinflations on average lead to recessions and why actual disinflationary episodes are followed by recessions. Although these results occur even in the presence of rational expectations, Ball suggests that perhaps we should "... overcome our qualms about adaptive expectations." He recommends the adaptive expectations assumption because it can explain both why ending inflation causes recessions and why one-time macroeconomic shocks can lead to persistent inflation.

V. Conclusions and Directions for Future Research

Although the practical policy implications of these papers are limited, the discussions help us to understand more fully some of the issues involved in the Friedman–Phelps debate. The resolution of this debate depends on the type of taxation used to replace lost seigniorage. If a consumption tax is feasible, Friedman's rule is optimal. However, if lost revenue can be replaced only with a wage, capital, or income tax, then the resolution also depends partly on the role of money in the economy.

When money is introduced into the model with a cash-in-advance constraint, inflation is a tax on consumption. Because a consumption tax acts like a lump-sum tax on the capital stock, some inflation will be part of an optimal tax package in a cash-in-advance model. When money is introduced into the model as a store of value, Friedman's rule is optimal. However, such models do not include private capital, and the generality of the result is still open to question.

None of the papers in this conference provides a comprehensive answer to the policymaker's question about the optimal inflation rate. No author has built a model to evaluate the effect of inflation on the efficiency of the monetary standard. Modeling money as a standard of value is problematic because the tools of microeconomic analysis assume away the frictions that make a standard useful. Money exists to facilitate trade and transactions — to make markets work more efficiently. Because we generally begin with models in which indexing is costless, or in which the efficiency of the payments and accounting systems is independent of the inflation policy, we should not be surprised that inflation appears to be rather harmless.

Support for zero inflation can also be found in the arguments contained in Altig and Carlstrom (1991a, 1991b). The interaction between inflation and the nominal tax system can result in significant distortions. It is not clear why Congress designed a tax system in which the effective capital tax rises with inflation. Perhaps legislators had some sort of state-contingent tax plan in mind, or perhaps they chose not to index the capital tax because of equity considerations. Another explanation is simply that indexing is difficult to achieve. Altig and Carlstrom (1991a) show that the indexing provisions for bracket creep contained in the Economic Recovery Tax Act of 1983 and the Tax Reform Act of 1986 are imperfect and still result in significant welfare losses.

The Friedman–Phelps debate centers on the optimal trend in the inflation rate. Chari et al. argue that constant inflation might not be optimal because of the presence of nonindexed debt. In their model, the government can use inflation changes with nonindexed debt in order to simulate indexed debt. Although their argument has merit, it leaves many questions unanswered. Can the government control the inflation rate precisely as needed to get the required pattern of real returns on government debt? Does it have an incentive to do so? What real-world uncertainties would accompany such a radical change in policy?

Perhaps the strongest argument against price stability per se (versus state-contingent inflationary policy around a zero-inflation trend) is presented by Henning Bohn, who shows that a policy of constant tax rates and a constant price level is not sustainable in a stochastic environment. Further research is needed to determine whether different operating strategies, for example, constant money growth targets, constant inflation targets, or a band around a path for the price level, would satisfy the conditions for sustainability.

An obvious gap exists between academic analysis and the actual practice of monetary policy. In theoretical modeling, the money supply rule completes the model and enables the researcher to determine the price level. In practice, the money supply rule is not sufficiently well defined to enable people to forecast inflation accurately. Adopting any reasonable and explicit rule may enhance economic performance by reducing uncertainty about future inflation.

None of the papers in this conference addresses the welfare or output effects of uncertainty about policy and the future price level. The rationale for an explicit multiyear path for a price index is based on the intuition that this uncertainty matters. If it does, a credible multiyear target for a price index would greatly reduce uncertainty about the future price level, eliminate the unexpected changes in the inflation trend that have been associated with recessions, and enhance the efficient operation of our accounting, contracting, and payments systems.

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Components of City-Size Wage Differentials, 1973–1988

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Introduction

The United States is predominantly, and increasingly, an urban society. Although the pace of urbanization stalled during the 1970's, it picked up steam again in the 1980s. Final 1990 census figures show that 77 percent of all U.S. residents live in cities, with 50 percent residing in metropolitan areas of more than 1 million inhabitants (up from just 41 percent in 1970).[1] Recognition of this trend has spawned a wide variety of research into the factors that draw us to live in cities, and also into the impact of cities on many aspects of our lives.

This paper examines one important issue associated with city dwelling — why wages increase with metropolitan area size. Explanations of this phenomenon must address both sides of the market: Why do workers in small cities accept lower wages instead of moving, and how can big-city employers compete successfully against lower-wage employers in small cities? On the workers' side, reasons proposed for the earnings disparity include compensation for higher skill levels, higher costs of living, and lower quality of life (such as exposure to crime and congestion) in large cities. On the employers' side of the market, the steeper wage bills faced by big-city firms are posited to be offset by higher productivity stemming from at least four possible sources: differences in immobile site characteristics, agglomeration economies, external economies related to city size, or differences in the quality of workers in large and small cities.

To explain the city-size wage gap, we need to know more precisely how wages differ among cities. Do large cities have higher-quality workers, for example, or does something about these cities make firms there more productive? As this question suggests, the city-size wage differential can be decomposed into two parts, both of which merit further investigation. One portion, arising from intercity disparities in *income-earning characteristics of the work force,* raises the question of why workers with divergent skills tend to locate in different-size cities. The second portion, which stems from intercity disparities in *wage structures,* invites inquiry into why workers with similar skills receive different pay in large versus small cities.

Such price differences imply that the magnitude of the city-size wage gap varies among workers. This in turn suggests that the importance of

■ 1 U.S. Department of Commerce, Bureau of the Census, press release CB91-66, February 21, 1991.

any given explanation for the gap may differ by type of worker. The results have major implications for understanding the efficiency of interregional labor markets. When do these markets produce a single price for labor, and under what circumstances do variations occur? Although equality of equilibrium factor prices across areas can be expected under many circumstances, it is by no means guaranteed.[2]

To begin answering some of these questions, we decompose city-size wage differentials reported in the U.S. Bureau of Labor Statistics' (BLS) Current Population Surveys (CPS) from 1973 to 1988 into the two portions outlined above: one arising from differences in worker characteristics across cities, and another resulting from intercity differences in the wage premia associated with these characteristics. This approach allows us to identify which work-force characteristics and which aspects of the wage structure account for most city-size wage differences. We then examine the changing importance of these factors over time.

Our results confirm that wages increase with city size. We also find that most of this effect is due to city-size-related differences in the prices of worker attributes. Strongest among these are the higher premia earned by skills (education and experience) in larger cities. Indeed, the driving force behind the shrinkage and expansion that we document in city-size wage differentials over time is changing returns to education.

Thus, the purpose of this study is to document in a novel way the patterns and trends of wage differentials related to city size. We leave to future research both the attempt to integrate these results into estimates of how wages reflect compensation to workers (or firms) for differences in specific area characteristics, and the effort to determine the importance of those area characteristics affecting firm productivity relative to those affecting worker utility.

I. Some Stylized Facts about the City-Size Wage Differential

Previous Studies

People have long observed that workers in large cities earn higher money wages than those in smaller cities and rural areas.[3] Most previous research on the city-size wage differential focuses solely on the total average wage gap, without investigating exactly which skills are priced differently.[4] These studies are based on cross-sectional wage regressions with controls for worker characteristics, where the only aspect of the wage structure allowed to vary between large and small cities is the intercept. City characteristics are then incorporated into the analysis to estimate the extent to which employees with similar skills may receive higher money wages in large cities because workplace or quality-of-life characteristics equalize compensation for workers and firms in small cities.

Previous research relating to the household point of view suggests that city-size-related differences in money wages either are largely offset by cost-of-living disparities, or compensate for differences in the quality of life across cities.[5] Addressing the issue from the firm's perspective, how can employers in large cities afford to pay higher wages than their smaller-city counterparts for workers with the same observed skills? Research on the four possible sources of posited higher productivity in larger cities — immobile site characteristics, agglomeration economies, external economies, or unobserved worker quality — is inconclusive.[6] Alternatively, city-size wage differentials may reflect institutional differences between large and small cities, such as unionization or the size or efficiency of the public sector.

Wages and City Size, 1973–1988

To illustrate some of the stylized facts discussed above and to motivate our study, the rest of this section uses data from the CPS to characterize the relationship between wages and city size over time. Figure 1 shows that average money wages increased consistently with city size for all size classes considered, from nonmetropolitan areas to metropolitan statistical areas (MSAs)

■ 2 See Dickie and Gerking (1987) for a discussion of the conditions under which these characteristic prices will be equalized across regions, and Beeson (1991) for an examination of equilibrium differences in factor prices.

■ 3 For early empirical studies of the relationship between wages and city size, see Fuchs (1967), Hoch (1972), and Rosen (1979).

■ 4 See Dickie and Gerking (1988) for a comprehensive review of this literature.

■ 5 See Hoch (1972), Hoch and Drake (1974), Izraeli (1977), Rosen (1979), and Cropper and Arriga-Salinas (1980).

■ 6 See Segal (1976), Sveikauskas (1975), Carlino (1978), Moomaw (1981), and Henderson (1988) for discussions of the relationship between productivity and city size.

FIGURE 1

City-Size Wage Premia Relative to Non-MSAs

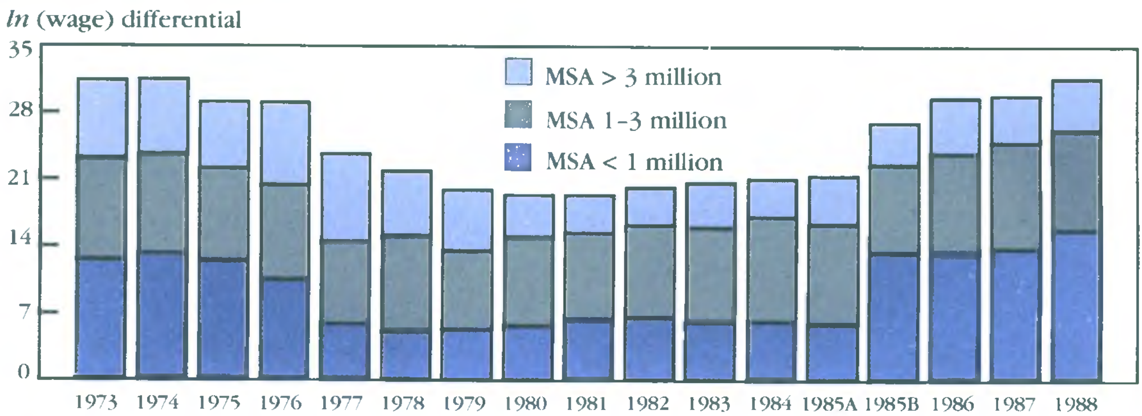
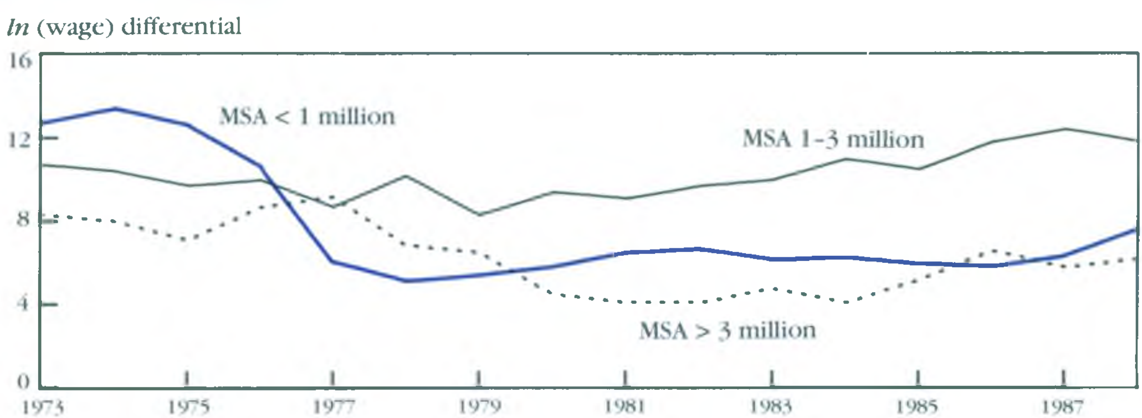


FIGURE 2

City-Size Wage Premia Relative to Next Smallest Size Class



SOURCE: Based on authors' calculations from CPS data.

with more than 3 million people.[7] However, the magnitude of the city-size wage premium varied over time.

Converting from the log wage differentials shown in figure 1 to percentage differences, we find that in 1973 and 1974, average wages in the largest MSAs were about 37 percent (32 log points) higher than in nonmetropolitan areas.[8] This differential then declined rapidly, falling to about 22 percent (20 log points) by 1980. After that, the wage gap widened slowly, adding a total of two percentage points between 1980 and 1985 and another five points between 1985 and 1988. Using the post–1985 MSA definitions, wages in large MSAs averaged 36 percent (31 log points) higher than those in non-MSAs in 1988.

Figure 1 also breaks down the difference between average wages in the largest MSAs and the non-MSAs by city size class. More than 50

7 The two different results for 1985 (1985A and 1985B) are displayed because of a discontinuity in the data, which is discussed later.

8 In figure 1 and throughout this paper, all wage differentials are expressed as log-point wage differences from the mean, which closely approximate percentage differences from the mean (particularly for differentials of less than 10 log points). In this section, we report some differentials in both log points and percentages to give the reader an idea of the magnitude of the larger differentials in percentage terms.

percent of the decline in the wage premium paid in large MSAs relative to non-MSAs during the 1970s was accounted for by a halving of the differential between small MSAs and non-MSAs.

Figure 2 shows more clearly how the gaps between each size class and the next smaller one changed over time. The differential between large and medium MSAs was fairly constant in the early 1970s, but shrank by half during the latter part of the decade, accounting for about 30 percent of the drop in the overall wage premium. These changes produced a more compressed distribution of wages across cities of different sizes in 1980 than existed in 1973. After 1980, the wage difference between large MSAs and non-MSAs started to expand, and this increase appears to have been spread more evenly among size classes than was the decline in the 1970s.

Shrinking city-size wage differentials during the 1970s and expanding differentials during the 1980s are consistent with evidence suggesting that economic activity fell off in metropolitan areas during the earlier decade. The 1970s saw a widespread, widely documented increase in population growth both in rural areas relative to urban ones and in small cities relative to large ones.[9] The most recent U.S. census indicates that this pattern was reversed during the 1980s.

The convergence/divergence of city-size wage gaps over the last two decades also bears some resemblance to wage patterns across census regions. Eberts (1989) and Browne (1989) find that after decades of convergence, wages and per capita regional personal income, respectively, began to move apart during the 1980s. This pattern also is evident in the 1990 census count.

II. Accounting for Wage Differentials and Their Changes

City-Size Wage Differentials

The technique we use to decompose the city-size wage differential into its components was introduced by Oaxaca (1973) to examine the male/female pay gap. This approach has also been applied to regional wage differentials and their changes over time by Sahling and Smith (1983), Farber and Newman (1987), and others.

The percentage difference in wages between a large city (City B) and a smaller one (City S) can be decomposed into worker skill and wage structure components as follows. Individuals' wages are based on the skills they possess and on the price that each skill receives in the market, or, in log-linear form:

$$(1) \qquad ln w_{jit} = \beta_{it} X_{jit} ,$$

where $ln w_{jit}$ is the natural logarithm of the wage of worker j in city size i during year t, vector β_{it} represents prices for specific skills (which can vary across city sizes and over time), and vector X_{jit} represents worker j's skills.

At any time, the portion of the wage differential stemming from differences in skills can be determined by comparing estimates of the average wages that would exist in City B and City S if prices of skills were the same in both. Similarly, the portion due to differences in the prices of specific skills can be determined by comparing estimates of average wages if skills were the same in both cities, but were alternatively priced in accordance with City B's and City S's wage structures. Based on these comparisons, the overall percentage differences in wages between the two cities can be expressed as follows:

$$(2) \qquad ln w_B - ln w_S = \beta(X_B - X_S) + (\beta_B - \beta_S) X,$$

where B and S indicate City B and City S, respectively.

In calculating the portion of the wage differential resulting from differences in worker skills, it is necessary to approximate β, the wage structure that would exist if there were no differences across city sizes. To deal with this problem, some authors use the wage structure in one of the regions (or in one group of workers). Others use either the wage structure based on the pooled sample (Eberts [1989]), or an average of the wage components based on the wage structure of the individual regions (Sahling and Smith [1983]). In our discussion, we emphasize the results based on the wage structure in large MSAs:

$$(2a) \qquad ln w_B - ln w_S = \beta_B(X_B - X_S) + (\beta_B - \beta_S) X_S .$$

For purposes of comparison, however, we also present alternative decompositions based on the wage structure in non-MSAs and on the average of the two components.

Changes in City-Size Wage Differentials over Time

Using equation (2), we can account for differences in wages in large and small cities at any

9 See Beale (1977), McCarthy and Morrison (1977), Carlino (1985), Moomaw (1986), and Beeson (1990).

point in time. We are also interested in explaining changes in *relative* wages in large and small cities over time. Following Farber and Newman (1987), equation (2) is estimated for each of two time periods. Differencing these estimates yields

$$(3) \quad (lnw_B - lnw_S)_t - (lnw_B - lnw_S)_{t-1}$$
$$= \beta\,[(X_B - X_S)_t - (X_B - X_S)_{t-1}]$$
$$+ X[(\beta_B - \beta_S)_t - (\beta_B - \beta_S)_{t-1}]$$
$$+ (\beta_{B,t} - \beta_{B,t-1})(X_B - X_S)$$
$$+ (\beta_B - \beta_S)(X_{S,t} - X_{S,t-1}),$$

where t and t - 1 indicate different time periods.

This decomposition identifies four separate components of the change in relative wages between cities. The first identifies the portion of the change attributable to variations over time in the mean characteristics of workers in City B relative to those in City S, assuming the wage structure (β) is the same across areas and over time. Ceteris paribus, if workers in large and small cities become more alike over time, the wage gap will become smaller. If there are no changes in mean characteristics in City B relative to City S, this component will equal zero.

The second component of equation (3) identifies the portion of the change in the wage gap attributable to changes over time in the wage structure of City B relative to City S, assuming no differences in worker characteristics. Ceteris paribus, if the difference between large and small cities in the returns to specific worker characteristics decreases over time, the wage gap will narrow. If no change occurs in the relative wage structure, this component will equal zero.

Even if there are no changes in relative characteristics of the work force across cities or in relative pay for worker characteristics (that is, if the first two components of equation [3] equal zero), wages in big cities relative to small ones may change. For example, suppose that workers in large cities have more years of schooling on average than those in smaller cities. If the price paid for each additional year of schooling declines, average wages of big-city workers will fall by more than that of their small-city counterparts, even if the price paid for each additional year of schooling is the same in both areas in both time periods. The contribution to the overall change in relative wages of this sort of "universal change" in the price of worker characteristics, given the initially unequal distribution of worker skills, is identified by the third component of equation (3).

Similarly, if a universal increase occurs in the mean level of a particular skill (for example, if average years of schooling increases everywhere), the difference in average wages across cities will change if that skill is rewarded unequally in large and small cities. If the reward is higher in large cities, the wage gap will increase. This portion of the change in the wage differential is identified by the fourth and final component of equation (3).

As was the case with the decomposition of the wage gap between large and small cities at any point in time (equation [2]), the measurement of each component of equation (3) is sensitive to the bases chosen for both the wage structure and the mean values of worker characteristics that would exist if there were no differences across areas. The number of potential combinations is large, and the use of alternative bases has little effect on the relative sizes of the components. Thus, we present the sizes of the components using the following wage structures and mean values:

$$(3a) \quad (lnw_B - lnw_S)_t - (lnw_B - lnw_S)_{t-1}$$
$$= \beta_B[(X_B - X_S)_t - (X_B - X_S)_{t-1}]$$
$$+ X_S[(\beta_B - \beta_S)_t - (\beta_B - \beta_S)_{t-1}]$$
$$+ (\beta_{B,t} - \beta_{B,t-1})(X_B - X_S)_{t-1}$$
$$+ (\beta_B - \beta_S)_t(X_{S,t} - X_{S,t-1}).$$

This formulation is consistent with the components estimated using equation (2a).

III. The Data

Sources and Inconsistencies

We use data from the CPS to compare wages for workers in different-size cities over the 1973–1988 period. The dependent variable in our analysis is hourly wage, measured as the logarithm of usual weekly earnings/average weekly hours.[10]

The CPS has three important advantages. First, it provides not only workers' earnings, but significant determinants of those earnings, such as age, education, and occupation. Second, it reports whether individuals live in an MSA, and if so, gives the area's size. Finally, the CPS is a broad survey whose design has remained fairly consistent over an extended period.

10 Weekly earnings reported in the CPS do not include fringe benefits. Because the proportion of compensation received as fringes rises with total compensation, this omission probably understates the effects investigated here.

One disadvantage of the CPS is that worker characteristics are not as detailed as in some other data sets. Studies of wage differences across broad regions by Gerking and Weirick (1983) and Dickie and Gerking (1987) indicate that omitting detailed worker information biases the estimation toward rejection of the hypothesis of equal wages, and tends to increase the portion of the wage differential attributed to wage structure. Although more information on worker characteristics would be desirable, we note that even after controlling for detailed traits, Gerking and Weirick still find a significant relationship between city size and wage differentials.

A second problem is that changes in the way the BLS collects and reports CPS data over the sample period complicate the construction of a consistent time series on wages by city size. First, in 1985, metropolitan area definitions were completely revised as follows:

- Through June 1985, identification is based on the 1973 standard metropolitan statistical area (SMSA) definitions.
- Since October 1985, identification is based on the 1983 MSA definitions.
- From July through September 1985, no metropolitan area information is provided.

Due to the extent of these changes, it is impossible to map the pre–1985 definitions to the post–1985 definitions. As a result, a discontinuity exists in our data series. When classifying metropolitan areas based on population size, we use the SMSA definitions prior to October 1985 and the primary metropolitan statistical area definitions thereafter.

These changes over time in the way metropolitan area population is reported limit us to four MSA size categories: more than 3 million people (large MSAs), between 1 million and 3 million people (medium MSAs), less than 1 million people (small MSAs), and nonmetropolitan areas (non-MSAs).[11] Prior to 1979, size classifications are based on 1970 population; afterward, they are based on 1980 population.

The BLS revisions affect cities of all sizes. However, the impact is most apparent in the wage differential between small MSAs and non-MSAs, which jumps from an estimated 5.9 log points based on the January–June 1985 data to 13.5 log points based on the October–December data. The change in the average wage differential between medium and small MSAs fell from 10.5 log points in early 1985 to 9.1 points in late 1985, while the differential between large and medium MSAs dropped from 5.2 to 4.4 log points over the same period.

A second factor that affects the continuity of our data series is the 1983 change in the BLS occupation codes. Again, these are extensive revisions that limit our ability to analyze changes in the city-size wage premium during the early 1980s.

Finally, in 1979, the BLS began collecting and reporting information each month on wages and characteristics of one-quarter of all workers surveyed. Prior to that time, this information was gathered only in May. As a result, our sample is much larger for years since 1979 than before, and the post–1979 sample is less seasonal. Experiments with monthly zero-one dummy variables indicate that this change does not seriously affect our estimates of city-size wage gaps. In addition, we do not adjust these data for top-coding of high-income individuals, a problem that has grown more severe in the CPS over time.

The pattern of wages observed and the data limitations lead us to focus on two distinct periods: 1973 to 1980, when the wage premium associated with city size was falling, and 1985 (post–October) to 1988, when the premium was increasing and no significant changes in variable definitions were instituted.

Mean Values by City Size and Year

This section describes patterns in the mean values of the data used in the analysis below. Means for the variables analyzed are reported by city size for the first and last years of the two time spans examined (see table 1).

Looking across city size categories, one can see a number of differences in worker characteristics that could contribute to the observed wage disparity. For example, workers in large cities have more years of schooling, are more likely to be full-time employees, and are more likely to be male. Racial composition varies by city size, and "potential experience" (age minus years of schooling minus six) is low in larger cities.

Changes can also be seen in work-force composition over time. Average years of schooling increased everywhere, but less so in large MSAs. In 1973, workers in large MSAs averaged 1.5 more years of schooling than their non-MSA counterparts. This difference fell to 0.6 years by 1980 and then rose to just over 0.7 years by 1988. Note also that average potential experi-

■ 11 Every year, the CPS reports a categorical measure of metropolitan area size. Unfortunately, we found this variable to be of little use because (1) the ranges of the categories changed from year to year and (2) for some years, there are inconsistencies between the categorical measure and other reported variables. Therefore, we used the metropolitan area status and name (in those cases where this information is provided) to classify MSAs.

TABLE 1

Mean Value of Worker Characteristics by City Size

	1973				1980				1985 (Oct.-Dec.)				1988			
	Non-MSA	Small MSA	Medium MSA	Large MSA	Non-MSA	Small MSA	Medium MSA	Large MSA	Non-MSA	Small MSA	Medium MSA	Large MSA	Non-MSA	Small MSA	Medium MSA	Large MSA
Sex (female = 1)	.420	.415	.414	.410	.456	.458	.452	.466	.486	.478	.475	.462	.491	.486	.484	.487
Race (nonwhite = 1)	.073	.096	.112	.137	.087	.114	.114	.232	.092	.102	.127	.225	.093	.112	.131	.245
Years of schooling	11.36	11.86	12.20	12.98	12.15	12.54	12.76	12.76	12.46	12.93	13.21	13.19	12.53	12.99	13.30	13.25
Full-time worker	.800	.807	.811	.818	.800	.809	.810	.825	.774	.799	.805	.833	.786	.806	.816	.836
Potential experience (age – school – 6)	19.72	18.78	17.97	19.50	18.12	17.74	17.34	18.21	18.26	17.24	17.33	17.68	18.77	17.51	17.39	17.85
White collar	.388	.469	.524	.533	.437	.512	.547	.574	.468	.571	.624	.626	.469	.577	.634	.633
Blue collar	.452	.386	.333	.325	.396	.324	.292	.262	.365	.283	.246	.238	.362	.277	.236	.229
Service occupations	.160	.145	.142	.143	.156	.148	.139	.146	.167	.145	.130	.137	.169	.146	.131	.138
ln (wage)	1.064	1.189	1.295	1.377	1.658	1.715	1.808	1.853	1.860	1.994	2.086	2.130	1.932	2.083	2.187	2.241
Number of observations	12,048	12,163	8,979	7,057	52,097	45,911	38,126	28,478	10,160	17,805	9,093	7,190	40,221	71,395	33,135	23,090

NOTE: Non-MSAs are nonmetropolitan areas; small MSAs are metropolitan areas with population less than 1 million; medium MSAs are metropolitan areas with population between 1 million and 3 million; large MSAs are metropolitan areas with population greater than 3 million.
SOURCE: Based on authors' calculations from CPS data.

ence dropped off overall between 1973 and the end of the decade, when it rebounded in non-MSAs but continued to decline or remained unchanged in the metropolitan areas.

In 1973, full-time workers constituted more of the labor force as city size increased, and this association grew stronger over time. Although women's presence in the labor force rose everywhere, there is no clear relationship between this change and city size. Minority representation also increased everywhere over the sample period, particularly in large MSAs between 1973 and 1980.

We also consider an individual's occupation by incorporating more than 40 nonagricultural job classifications in our analysis. For brevity, table 1 aggregates these classifications into three categories: blue collar, white collar, and service occupations. In all years, the concentration of white-collar jobs generally rises with city size, while the opposite usually holds for blue-collar and service occupations. In 1973, white-collar workers accounted for 53 percent of total employment in large MSAs, but only 39 percent in non-MSAs. By contrast, blue-collar workers accounted for less than 33 percent of employment in large MSAs, compared to 45 percent in non-MSAs. Over time, blue-collar workers' share of jobs declined, that of white-collar workers increased, and service jobs remained fairly stable. Despite the changing composition of occupations over time, the relative distribution of employment across city sizes held constant.

TABLE 2

Wage Equation Estimates by City Size

	1973				1980				1985 (Oct.-Dec.)				1988			
	Non-MSA	Small MSA	Med. MSA	Large MSA	Non-MSA	Small MSA	Med. MSA	Large MSA	Non-MSA	Small MSA	Med. MSA	Large MSA	Non-MSA	Small MSA	Med. MSA	Large MSA
Sex (female = 1)	–.226[a]	–.193	–.195	–.173[a]	–.204[a]	–.201	–.166[a]	–.138[a]	–.147[a]	–.139	–.099[a]	–.101[a]	–.135[a]	–.109[a]	–.098[a]	–.077[a]
	(.015)	(.014)	(.016)	(.018)	(.006)	(.006)	(.007)	(.003)	(.015)	(.034)	(.015)	(.017)	(.008)	(.005)	(.008)	(.010)
Race (nonwhite = 1)	–.128[a]	–.058	.004[a]	–.004[a]	–.049[a]	–.044	–.051	–.068[a]	–.059[a]	–.062	–.061	–.073	–.063[a]	–.055	–.077[a]	–.063
	(.016)	(.013)	(.015)	(.014)	(.006)	(.006)	(.006)	(.006)	(.014)	(.010)	(.013)	(.012)	(.007)	(.005)	(.007)	(.006)
Years of schooling	.044[a]	.044	.048[a]	.050[a]	.043[a]	.042	.046[a]	.046[a]	.052[a]	.049	.051[a]	.051[a]	.046[a]	.050[a]	.050[a]	.054[a]
	(.002)	(.002)	(.002)	(.002)	(.001)	(.001)	(.001)	(.001)	(.002)	(.002)	(.002)	(.002)	(.001)	(.001)	(.001)	(.001)
Full-time worker	.097[a]	.148[a]	.164[a]	.198[a]	.147[a]	.171[a]	.179[a]	.210[a]	.166[a]	.211[a]	.214[a]	.194[a]	.181[a]	.206[a]	.211[a]	.216[a]
	(.012)	(.011)	(.013)	(.014)	(.005)	(.005)	(.006)	(.007)	(.011)	(.009)	(.012)	(.015)	(.006)	(.004)	(.007)	(.008)
Potential experience (age – school – 6)	.028[a]	.030	.030[a]	.029	.025[a]	.028[a]	.031[a]	.028	.028[a]	.030[a]	.032[a]	.031[a]	.029[a]	.032[a]	.031[a]	.031[a]
	(.001)	(.001)	(.001)	(.001)	(.000)	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)	(.000)	(.001)	(.001)
Experience sq. ÷ 100	–.043[a]	–.042	–.046	–.041	–.038[a]	–.043[a]	–.047[a]	–.040	–.037[a]	–.043[a]	–.045[a]	–.043[a]	–.041[a]	–.045[a]	–.045[a]	–.044[a]
	(.002)	(.002)	(.002)	(.002)	(.001)	(.001)	(.001)	(.001)	(.002)	(.002)	(.002)	(.003)	(.001)	(.001)	(.001)	(.001)
Sex * Experience	–.008[a]	–.006[a]	–.005	–.004	–.004[a]	–.004	–.006[a]	–.004[a]	–.005[a]	–.005	–.007[a]	–.006	–.005[a]	–.006[a]	–.006[a]	–.005
	(.001)	(.001)	(.001)	(.001)	(.000)	(.000)	(.000)	(.000)	(.001)	(.000)	(.001)	(.001)	(.000)	(.000)	(.000)	(.000)
Intercept	.805[a]	.822	.773	.736	1.281[a]	1.264	1.220[a]	1.231[a]	1.346[a]	1.422	1.429	1.406	1.460[a]	1.473[a]	1.486	1.458
	(.057)	(.044)	(.049)	(.054)	(.022)	(.021)	(.021)	(.025)	(.058)	(.034)	(.045)	(.052)	(.028)	(.017)	(.024)	(.029)
$\bar{R}^2$	.448	.466	.490	.488	.453	.485	.503	.480	.468	.503	.521	.481	.471	.504	.502	.493

a. For non-MSAs, indicates that coefficient is significantly different from zero; for other groups, indicates that coefficient is significantly different from non-MSA coefficient at the 10 percent confidence level.

NOTE: See table 1 for MSA size definitions. All regressions include 40 occupational dummies. Engineer is the omitted occupation. Standard errors are in parentheses.

SOURCE: Based on authors' calculations from CPS data.

IV. Decomposition of City-Size Wage Differentials and Changes

Estimates of Wage Structure by City Size

Estimates of returns to worker characteristics are obtained from separate ordinary least squares wage regressions for each city size group in each year on two measures of human capital: years of schooling and potential experience (the latter entered as a quadratic). Because women move into and out of the labor market more than do men, potential experience tends to overestimate true experience more for women. Therefore, we include an interaction between female and experience in order to allow pay for experience to differ between the sexes. In addition, we include dummy variables indicating whether a worker is a full-time employee, female, nonwhite, and in any of 40 occupations.

Parameter estimates by city size and year are reported in table 2. Consistent with previous research, we find that for all size classes, white men are paid more than women and minorities, while more experience, more schooling, and full-time status are all associated with higher earnings. These effects are still present after controlling for the 40 occupations.

Beyond these similarities, estimates of returns to worker characteristics often differ significantly by city size, indicating higher returns in larger cities for the usual measures of skill. For every year, F-tests reject the hypothesis of identical wage structures across city size groups at the 1 percent confidence level. These tests also reject the hypotheses of identical intercepts and identical slopes given different intercepts at the 1 percent level. Furthermore, t-tests reject the hypothesis of equal coefficients for a large number of pairwise comparisons between city size groups in all years, particularly for the coefficients on full-time status, sex, and years of schooling.

In 1973, the wage premium received by full-time workers was almost 20 log points in large MSAs, but less than 10 points in non-MSAs. Over time, this premium increased everywhere, but particularly in the non-MSAs, where it doubled to 18 log points. In large cities, the premium rose by just under two points between 1973 and 1988.

After controlling for other observed characteristics, we find that women's pay tends to increase with city size: The differential between large MSAs and non-MSAs ranges from 4.5 percent to 6.5 percent. While minority workers in non-MSAs earned 12 percent less than their urban counterparts in 1973, this differential disappeared by 1980.[12]

These estimates also reveal sizable differences in returns to schooling across city size groups. The estimated increase in wages associated with each additional year of schooling was 0.5 percentage points higher in large MSAs than in non-MSAs in 1973. This implies that the estimated wage of a high school graduate (12 years of schooling) was 6 percent higher, ceteris paribus, in large cities than in non-MSAs.

The well-documented decline in returns to schooling during the 1970s and the subsequent increase during the 1980s is also evident from our estimates. However, because we control for occupation, the pattern is not very strong.[13] In the 1970s, the economic returns to schooling fell most severely in large MSAs. During the following decade, these returns rose in all city size groups, although not in lockstep. By 1988, returns to education had clearly risen most in large MSAs, where a worker with a high school degree earned about 10 percent more than his or her rural counterpart.

These results are particularly striking because they are obtained while controlling for occupation. In results not reported here, the sizes of coefficients on education and experience, and their association with city size, are seen to rise when occupation dummies are excluded from the model. Thus, the role of the occupation dummies is consistent both with the interpretation of occupation as a control for human capital, and with our conclusion that big cities reward human capital more highly.

Intentionally omitted from these regressions are controls for employer characteristics, such as industry, firm size, and establishment. Although these factors are important determinants of wages (Groshen [(1991a, 1991b]), we exclude them to make our results more comparable with previous work and to restrict the focus of this paper to the role of human capital characteristics in the city-size wage gap. Investigation of the contribution of demand-side influences is left to future research.

Components of City-Size Wage Differentials

In this section, we examine how the differences in worker characteristics and in estimated characteristic prices discussed above contribute to total observed wage differences among cities of different sizes. As previously noted, we can consider the wage differential between large and small cities as being composed of two portions: one reflecting intercity differences in average worker skills, and one reflecting intercity differences in the price associated with those skills. The decomposition of the wage differential into these two components using equation (2) is reported in the top panel of table 3. For each year considered, the top row reports the total wage gap between the largest MSAs and the non-MSAs, the large and medium MSAs, the medium and small MSAs, and the small MSAs and non-MSAs.

The next three rows of the table show decompositions of the four wage gaps listed above.[14] Each row uses one of the three alternative bases for the decomposition. Note that the base used has almost no effect on our qualitative results:

■ 12 The relatively large absolute value for the coefficient on race in the 1973 non-MSA wage equation is not an aberration of the data for that particular year. This coefficient increased steadily from 1973 until 1978, when it peaked near zero. It then fell to about 5 percent, stayed there until 1988, and fluctuated between 5 percent and 7 percent through the remainder of the decade.

■ 13 See Bound and Johnson (1989).

■ 14 These decompositions are based on the coefficient estimates reported in table 2, which also gives standard errors for each coefficient. However, there is currently no method for constructing confidence intervals around the decompositions performed in table 3.

TABLE 3

Explaining the Wage Gap between Large MSAs and Non-MSAs

	1973				1980				1985 (Oct.-Dec.)				1988			
	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA
$ln\,(\text{wage}_B) - ln\,(\text{wage}_S)$	.313	.082	.106	.125	.195	.044	.093	.058	.270	.044	.091	.134	.309	.054	.104	.151
Portion due to differences in worker skills																
$\beta_B(X_B - X_S)$	.081	.010	.023	.037	.045	–.001	.017	.021	.079	–.001	.036	.042	.080	.000	.035	.043
$\beta_S(X_B - X_S)$	.060	.010	.018	.036	.031	–.001	.015	.020	.073	.001	.034	.039	.066	–.002	.033	.038
$\beta\,(X_B - X_S)$	.071	.010	.020	.037	.038	–.001	.016	.021	.076	.000	.035	.041	.073	–.001	.034	.040
Portion due to differences in wage structure																
$X_B(\beta_B - \beta_S)$	.253	.072	.088	.089	.164	.046	.079	.021	.197	.043	.057	.095	.243	.055	.071	.114
$X_S(\beta_B - \beta_S)$	.232	.072	.083	.088	.150	.046	.076	.037	.191	.045	.055	.092	.230	.054	.070	.109
$X(\beta_B - \beta_S)$	.242	.072	.085	.088	.157	.046	.077	.036	.194	.044	.056	.094	.236	.054	.071	.111
Years of schooling																
Means: $\beta_B(X_B - X_S)$	.037	–.006	.016	.028	.028	.000	.010	.016	.038	–.001	.015	.023	.036	–.002	.015	.023
Returns: $X_S(\beta_B - \beta_S)$	.075	.028	.053	–.001	.036	–.003	.051	–.010	–.012	.001	.028	–.041	.061	.002	.003	.058
Other demographic variables																
Means: $\beta_B(X_B - X_S)$	.013	.018	–.002	–.001	.011	.001	–.001	–.003	.000	.002	.002	–.003	–.007	.001	.000	–.008
Returns: $X_S(\beta_B - \beta_S)$	.154	.044	.015	.097	.108	.027	.035	.046	.067	–.015	.028	.052	.078	.026	–.003	.055
Occupation mix																
Means: $\beta_B(X_B - X_S)$	.031	–.002	.009	.016	.021	–.003	.008	.008	.041	–.002	.015	.022	.050	.054	.019	.028
Returns: $X_S(\beta_B - \beta_S)$	.072	.037	.068	–.025	.056	.011	.037	.018	.076	.081	.028	.008	.093	.001	.058	–.002
Intercept	–.069	–.038	–.049	.017	–.050	.010	–.044	–.017	.060	–.024	.007	.077	–.003	–.029	.014	.013

NOTE: See table 1 for MSA size definitions. β_B is wage structure of larger MSA; β_S is wage structure of smaller MSA; X_B is mean characteristics of larger MSA; X_S is mean characteristics of smaller MSA.

SOURCE: Based on authors' calculations from CPS data.

We find that differences in wage structure (that is, the higher returns to skills in large cities reported in table 2) account for 70 to 80 percent of the gap between big-city and small-city wages, regardless of the year, the city sizes compared, or the base chosen.

Over time, the proportion of the total wage differential attributable to differences in worker characteristics remains fairly constant, ranging from 20 percent to 30 percent, while the importance of differences in worker attributes decreases as we compare progressively larger cities.

The bottom panel of table 3 reports a more-detailed decomposition of the wage gaps using a single base. These seven rows allow us to examine the extent to which schooling, other demographic variables, occupation, and the intercept contribute to the city-size wage differential.

In every year except 1985, education is the single most important human capital characteristic in explaining differences in wages between large MSAs and non-MSAs, even controlling for major occupational group. In each year, schooling accounts for approximately one-third of the wage gap between these two size categories. In general, more education in bigger cities explains about 30 percent of the gap, while higher returns to schooling there account for the other 70 percent. Although overall differences in schooling returns and means are not important between large and medium MSAs, they do come into play in the other size comparisons.

Occupation mix also accounts for a large (and increasing) portion of the differential, a result of differences in both job mix and occupational wage structures across cities. The individual contributions of the other human capital measures are relatively small, so we aggregate them into the category "other demographic variables." Although considerable variation over time is apparent, these factors contribute to the city-size wage gap primarily through intercity differences in prices paid for these characteristics, not through differences in supply.

The differences in intercepts estimated here show no clear pattern across city size or time, which we interpret to mean that the intercepts mainly capture unsystematic omitted factors, such as overall price levels in individual cities.[15] In addition, because these differences are usually small and frequently negative, we find that the terms included in the regression are capable of explaining most or all of the city-size wage gap. This, in turn, implies that the bulk of the gap is due to higher remuneration of some worker attributes in larger cities. That is, for the lowest-skilled workers, we observe no consistent city-size wage gap at all.

In summary, we find that average wages are higher in large cities than in small ones throughout the time period examined. The most important factor behind this disparity is differences in wage structure, suggesting that skills obtained through education, experience, and occupational training are more highly rewarded in larger cities.

Changes over Time in the Components of City-Size Wage Differentials

Although wages in large cities are consistently higher than in small ones, the size of this premium varies considerably over time. As noted above, the wage gap between large MSAs and non-MSAs fell from more than 30 log points in 1973 to less than 20 log points in 1980, then widened throughout the remainder of the decade. We now turn our attention to examining the sources of these changes over time.

The four components of changes in the wage differential between large and small cities are identified in equation (3a). The contributions of each of these components to the decline in the city-size wage gap between 1973 and 1980, and to the subsequent increase between 1985 and 1988, are presented in table 4. Note that, as was true for differences at each point in time, changes in wage structure (third and fourth rows) account for a larger portion of the change in relative wages over time than changes in skills (second and fifth rows). This is true across all size comparisons for both time periods, and also when broken down by worker characteristic.

Changes over time in the relative prices paid for worker skills in large MSAs and non-MSAs account for more than 70 percent of the decline in the wage differential between 1973 and 1980, and for more than 90 percent of the increase between 1985 and 1988. A rise in the average skills of workers in non-MSAs relative to those in large MSAs, perhaps related to the migration of skilled workers from large cities, accounts for the remainder of the decline between 1973 and 1980.

Overall, the portions of the decline in the wage gap between 1973 and 1980 that can be attributed to education, occupation, or other

■ 15 In the decompositions performed here, the intercept estimates may reflect either the earnings patterns of the reference group, or mean earnings common to all workers stemming from factors omitted from the model, which are uncorrelated with the included regressors. If the latter, these estimates indicate how well our model captures the main determinants of city-size wage gaps. As usual, the reference group is fairly uninteresting (white, male, part-time engineers with zero years of education and experience).

TABLE 4

Components of Intertemporal Changes in Wage Differentials

	1973–1980				1985–1988			
	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA
$[ln\,(\text{wage}_B) - ln\,(\text{wage}_S)]_t - [ln\,(\text{wage}_B) - ln\,(\text{wage}_S)]_{t-1}$	−.118	−.038	−.012	−.068	.040	.010	.013	.017
Portion due to changes in worker skill differential $\beta_{B,t}[(X_B - X_S)_t - (X_B - X_S)_{t-1}]$	−.027	−.001	−.003	−.015	−.004	−.001	−.003	−.002
Portion due to changes in relative wage structure $X_{S,t-1}[(\beta_B - \beta_S)_t - (\beta_B - \beta_S)_t]$	−.086	−.029	−.010	−.050	.038	.009	.014	.015
Portion due to universal change in wage structure $(\beta_{B,t} - \beta_{B,t-1}) * (X_{B,t} - X_{S,t})$	−.009	−.002	−.003	−.001	.005	.002	.001	.002
Portion due to universal change in worker skills $(\beta_{B,t} - \beta_{S,t}) * (X_{S,t} - X_{S,t-1})$	.004	.002	.003	−.001	.001	.000	.000	.001

NOTE: See table 1 for MSA size definitions. β_B is wage structure of larger MSA; β_S is wage structure of smaller MSA; X_B is mean characteristics of larger MSA; X_S is mean characteristics of smaller MSA.

SOURCE: Based on authors' calculations from CPS data.

TABLE 5

Sources of Intertemporal Changes in Wage Differentials

	1973–1980				1985–1988			
	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA
$[ln\,(\text{wage}_B) - ln\,(\text{wage}_S)]_t - [ln\,(\text{wage}_B) - ln\,(\text{wage}_S)]_{t-1}$	−.118	−.038	−.012	−.068	.040	.010	.013	.017
Overall contribution of:								
Years of schooling	−.048	−.026	−.008	−.015	.072	.000	−.023	.095
Other demographic variables	−.063	−.032	−.023	−.054	.004	.038	−.034	.000
Occupation mix	−.026	−.028	−.033	.034	.027	−.023	.064	−.014
Intercept	.019	.048	.005	−.034	−.062	−.005	.006	−.064

NOTE: See table 1 for MSA size definitions. β_B is wage structure of larger MSA; β_S is wage structure of smaller MSA; X_B is mean characteristics of larger MSA; X_S is mean characteristics of smaller MSA.

SOURCE: Based on authors' calculations from CPS data.

demographic categories are fairly equal (tables 5 and 6). However, there are some differences in their relative importance *across* size classes. Years of schooling, for example, accounts for a relatively large part of the decline in the wage differential between large and medium MSAs, but explains relatively little of the drop-off between the other size classes. Similarly, changes in occupation mix and occupational wage structure during the 1970s actually worked to widen the small MSA/non-MSA earnings gap, but contributed to narrowing the gap between other city size categories.

It is interesting to note that the decline in returns to schooling during the 1970s was not uniform across cities of different sizes: Our estimates indicate that the downturn was most severe in the largest MSAs, where estimated returns fell by 10 percentage points. In contrast, returns to education fell only two to five percentage points in the smaller MSAs and non-MSAs (table 2). These diminished returns to schooling in the largest MSAs relative to the non-MSAs account for more than a quarter of the total decrease in the wage differential (.031 out of a total drop of .118) and completely swamp the effect of the general decline in returns to schooling (see table 6). If returns to schooling had fallen by the same amount in non-MSAs as in the large MSAs, and if the difference in average years of schooling had remained constant, the wage gap between these two size classes would have declined only one-third of 1 percent.

In brief, national trends toward increased levels of education and greater labor force participation among women and minorities had little effect on relative wages among cities, even though these attributes were rewarded differently across cities of different sizes. Similarly, national trends in the returns to specific skills contributed little to the changing differential, despite the unequal distribution of skills across cities. Changes in the relative skill mix of large and small cities, perhaps related to the selective nature of migration, accounted for a sizable portion of the decline in the wage gap during the 1970s, but contributed little to its expansion in the 1980s. The single most important factor in this decline/resurgence was the corresponding shrinkage and expansion of the city-size gap in prices paid for worker skills, a pattern driven chiefly by changes in city-size-related returns to both education and occupation.

TABLE 6

Components of Intertemporal Changes in Wage Differentials

	1973-1980				1985-1988			
	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA	Large-Non-MSA	Large-Med.	Med.-Small	Small-Non-MSA
$[ln\,(\text{wage}_B) - ln\,(wage_S)\,]_t - [ln\,(\text{wage}_B) - ln\,(\text{wage}_S)\,]_{t-1}$	–.118	–.038	–.012	–.068	.040	.010	.013	.017
Portion due to changes in worker skill differential $\beta_{B,t}[(X_B - X_S)_t - (X_B - X_S)_{t-1}]$								
Total	–.027	–.001	–.003	–.015	–.004	–.001	–.003	–.002
Years of schooling	–.006	.005	–.006	–.005	–.001	–.001	.001	–.001
Other demographic variables	–.013	–.014	.002	–.003	–.009	–.002	–.003	–.004
Occupation mix	–.009	–.001	.001	–.007	.006	.003	–.001	.003
Portion due to changes in relative wage structure $X_{S,t-1}[(\beta_B - \beta_S)_t - (\beta_B - \beta_S)_t]$								
Total	–.086	–.029	–.010	–.050	.038	.009	.014	.015
Years of schooling	–.042	–.031	–.004	–.008	.073	.001	–.024	.095
Other demographic variables	–.046	.018	.021	–.051	.009	.039	–.031	.004
Occupation mix	–.017	–.027	–.032	.043	.018	–.026	.063	–.020
Intercept	.019	.048	.005	–.034	–.062	–.005	.006	–.064
Portion due to universal change in wage structure $(\beta_{B,t} - \beta_{B,t-1}) * (X_{B,t} - X_{S,t})$								
Total	–.009	–.002	–.003	–.001	.005	.002	.001	.002
Years of schooling	–.003	.001	–.001	–.001	–.001	.000	.000	.000
Other demographic variables	–.004	–.003	–.001	.001	.002	.001	–.001	.000
Occupation mix	–.002	.000	–.002	–.001	.003	.000	.002	.002
Portion due to universal change in worker skills $(\beta_{B,t} - \beta_{S,t}) * (X_{S,t} - X_{S,t-1})$								
Total	.004	.002	.003	–.001	.001	.000	.000	.001
Years of schooling	.002	–.026	.003	–.001	.000	.000	.000	.000
Other demographic variables	.000	.028	.000	.000	.001	.000	.000	.001
Occupation mix	.002	.000	.000	–.001	.000	.000	.001	.000

NOTE: See table 1 for MSA size definitions. β_B is wage structure of larger MSA; β_S is wage structure of smaller MSA; X_B is mean characteristics of larger MSA; X_S is mean characteristics of smaller MSA.

SOURCE: Based on authors' calculations from CPS data.

V. Summary and Conclusions

This paper examines wage differentials related to city size over the 1973–1988 period. We decompose the nominal city-size wage differentials across four size classes of cities into the portion due to differences in worker traits and the portion due to differences in wage structure. Our results show that differences in worker-attribute prices account for a larger share of city-size wage differentials than do intercity differences in the worker attributes themselves. In particular, the economic reward for attributes associated with skill, especially years of education and experience, rises with city size.

This finding suggests important structural differences among employers in cities of different sizes and is consistent with studies of regional wage differentials.[16] Different wage structures among cities also imply a new "stylized fact" about the city-size wage differential: It does not accrue uniformly to all workers in large cities. In fact, the estimates reported here suggest that it is received almost exclusively by workers with high education or experience and by those who are female or full-time employees.

Of course, we cannot be certain that all relevant differences in work forces among cities have been captured by the variables available in the CPS. If we have omitted any important determinants of workers' productivity correlated with their choice of city size, then our estimates overstate the role of price differences in the city-size wage gap. For example, Gerking and Weirick (1983) and Dickie and Gerking (1987) find that the importance of differences in worker characteristic prices may be overstated if data on these traits lack detail. Thus, examination of city-size wage differentials using either different information (for example, longitudinal or more-detailed data) or techniques to control for sample selection may be in order.

We also examine how differences in worker characteristics and in characteristic prices contribute to changes in the city-size wage gap over the sample period. Our analysis shows that movements in relative prices account for the majority of the change in the city-size wage premium over time. These findings are particularly interesting in light of the general "U-turn" path of wage inequality found by Bound and Johnson (1989) and others, and are consistent with studies of changes in regional wage differentials over time (see Farber and Newman [1987] and Eberts [1989]).

16 See Hanushek (1973, 1981), Sahling and Smith (1983), and Jackson (1986).

Furthermore, we find that a large portion of this intertemporal change is found in the human capital coefficients, particularly in returns to years of schooling. This conclusion is generally consistent with Eberts' (1989) finding that changes in broadly defined occupation coefficients account for virtually all of the change in characteristic prices across census regions. The overall similarities between the intertemporal patterns of wage differentials across broad census regions and across city size groups suggest that a more detailed exploration of the relationship between city size and regional wage differentials may yield interesting insights.

Although we do not attempt to estimate the relationship between wage structure and specific characteristics of areas, our findings have some important implications for such research. In particular, the variability we observe in worker-attribute prices over time suggests that the estimated prices of city-size-related area characteristics (pollution and congestion, for example) may be quite sensitive to the time period covered in the analysis.

Second, why is the city-size wage differential strongest among the most skilled workers? If agglomeration economies or other productivity-enhancing city attributes are the reason, then wouldn't the differentials be seen across a wide group of worker attributes? Perhaps some structural change has taken place in the larger cities, such as the concentration of more technical processes or the loss of routine jobs to rural areas.

Finally, the persistence of city-size-related wage premia suggests that it may be fruitful to examine these differentials in the context of an equilibrium location model such as the one found in Roback (1982). This type of analysis could address the relative importance of productivity and amenity differences as determinants of the city-size wage gap.

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Financial Efficiency and Aggregate Fluctuations: An Exploration

by Joseph G. Haubrich

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Introduction

Banks may both initiate and propagate business cycle fluctuations. For example, recent controversy has arisen over the role that banks' loan decisions may have played in the initiation of economic downturns. However, once such a downturn has begun, business performance influences bank profits and eventually may influence loans.

A natural question is whether banks play an independent role in business cycle movements. One possibility is that technological advances specific to banking influence the initiation or continuation of business cycles. Although economists have studied the macroeconomic impact of broad technological advances, they have not yet focused on the impact of the obvious recent gains in banking technology.

In this paper, I explore the possible link between financial efficiency and macroeconomic fluctuations. I present a two-sector real business cycle model in which there are technological shocks specific to the banking sector. (For a discussion of these shocks, see box 1.) I then test the model's empirical implications, which are interpreted with the concept of cointegration. Although these implications have not yet been investigated by others, I follow Mitchell (1913) by examining linear combinations of banking variables. This approach is linked to the cointegration techniques utilized. Specifically, I test whether a common stochastic trend exists between banking variables and industrial production, or whether the two are subject to distinct stochastic trends. This is equivalent to testing for the absence of cointegration between the banking variables and industrial production. Such a finding would imply that the banking sector exerts independent influence on long-run output.

Other researchers have begun to consider roles for financial efficiency. On the theoretical side, King and Plosser (1984) mention financial efficiency, but do not imbed a role for it in the solution of their model. Greenwood and Williamson (1989) develop a model with an explicit but constant term for financial efficiency. On the empirical side, Norrbin and Schlagenhauf (1988) present a highly disaggregated model, but one that does not explicitly consider technological change. They also discuss employment rather than output in the financial sector, and use a different time period (1954–1984). Corradi, Galeotti, and Rovelli (1990) look at the long-run relations among bank variables in Italy without considering the aggregate economy. Scotese (1990), in unpublished work, BOX 1

Some Examples of Bank Innovations

As an antidote to the rather abstract theory and empirical work in the rest of the paper, it may be useful to consider some examples of the sort of shocks I have in mind. Banks employ many resources in processing transactions, in maintaining the payments system, and in screening and monitoring lenders and borrowers. Advanced information technology has reduced these costs and created new products. In the period covered by the study (1923–1978), some efficiency gains stem from outside technological advances, others are unique to banks, and still others are unclear.

Even in cases where outside technology increases financial efficiency — Scotch Tape in the 1930s, or calculators and electric typewriters in the 1940s — the specific uses and total gain can vary by industry. For example, using radio and television to disseminate information has meant very different things to the banking, the soap, and the fashion industries.

Other improvements seem more specific to banks. These include innovations like money orders and warehouse receipts to collateralize loans in the 1930s; drive-up windows, account numbers, and check routing numbers in the 1940s; and central information files in the 1960s. Yet, not all "breakthroughs" look so stunning in retrospect. In the 1920s, for instance, banks placed a strong emphasis on graphology, using handwriting analysis to screen employees and customers.

Today, the contribution of computer technology with image processing enables payments, credits, and debits to be made more cheaply, easily, and quickly, especially far from home. Banks can use this technology to calculate and adjust exposure and interest-rate risk. As daily processing becomes commonplace, on-line transaction processing becomes more frequent. Banks have also reduced the costs of monitoring and screening and have automated the process of sending out warning letters. Data bases make all customer accounts accessible, easing credit-risk analysis and targeted advertising.

One recent innovation, for example, is the "super-smart card." Resembling a credit card, it fits easily into a wallet and contains memory, a processor, keyboard, screen, calculator, and clock. Debits and credits can be made merely by punching a secret code, making transactions quicker than is possible with current credit and debit cards. More secure than an ATM card, the super-smart card also reduces costs by eliminating the need for point-of-sale terminals to be connected to a central location. It holds the promise of inducing even more radical changes: Pocket currency could pay interest, or even float against that of other banks. Such capabilities have the Japanese Ministry of Finance worried about losing control of the nation's money supply (Abrahams [1988]).

One indicator of technology's influence is the substantial consulting industry that banks support to help them manage extensive technological change. Estimates suggest that the largest of the 25 consulting firms profiled in a recent issue of *American Banker* collects \$400 million annually from bank technology consulting alone (Gullo [1991]).

examines the relation between economic growth and financial innovation, but models technological changes differently and examines fewer banking variables than I do. She also uses quarterly data from 1959 to 1990, whereas I use a longer data series (1923–1978.)

The remainder of the paper proceeds as follows. Section I presents the simple model and explains why technological shocks imply no cointegration between banking and real variables. Section II describes the data, the method of testing, and the test results. Contrary to the prediction of the model, banking and real variables are cointegrated. Section II further explores the interaction with vector autoregression methods, and section III concludes.

I. Lessons from a Simple Model

To consider the effects of shocks to financial efficiency, I begin with a dynamic stochastic model: a two-sector real business cycle model with technological shocks to both sectors. This two-industry version of the Long and Plosser (1983) model has testable predictions and indicates the progress that can be made by treating banks like any other industry. It retains a somewhat traditional flavor, however, because it places transactions services directly into both the utility function and the production function.

Consider a model economy with two goods and a representative agent who chooses a production and consumption plan. The infinitely lived agent has resources, technologies, and tastes similar to those in Long and Plosser, and has a lifetime utility function of $U = \Sigma\beta^t u(C_t, Z_t)$, where C_t is a 2x1 vector denoting period t consumption of goods (C_G) and banking services (C_B). Z_t measures the quantity of leisure consumed in period t. Each period's utility function, $u(C_t, Z_t)$, is given by

$$(1) \qquad u(C_t, Z_t) = \theta_0 lnZ_t + \theta_G\, lnC_{Gt} + \theta_B\, lnC_{Bt}\,.$$

The agents face two resource constraints: Total time H may be spent at work or at leisure, and output Y_t may be consumed or invested.

$$(2) \qquad Z_t + L_{Gt} + L_{Bt} = H$$

$$(3) \qquad C_{jt} + X_{Gjt} + X_{Bjt} = Y_{jt}\,.$$

Thus, labor can be divided between producing transactions services in the banking sector or output in the goods sector, just as the goods (output and banking services) can be consumed or invested. X_{ij} denotes the amount of good j invested in process i. For example, X_{GB} is the amount of banking services used to produce the manufactured good. Output is determined by Cobb–Douglas (1928) technology with a random productivity shock.

(4) $$Y_{G,t+1} = \lambda_{G,t+1}\, L_{Gt}^{b_G} (X_{GG})^{a_{GG}} (X_{GB})^{a_{GB}}$$

$$Y_{B,t+1} = \lambda_{B,t+1}\, L_{Bt}^{b_B} (X_{BG})^{a_{BG}} (X_{BB})^{a_{BB}},$$

where $\lambda_{i,\,t+1}$ is a random productivity shock whose value is realized at the beginning of period $t+1$, and the exponents are positive constants with $b_i + a_{iB} + a_{iG} = 1$. For future reference, define the matrix of input-output coefficients a_{ij} to be matrix A. Because the state of the economy in each period is fully specified by that period's output and productivity shock, it is useful to denote that state vector $S_t = [Y_t' \; \lambda_t']$. To further simplify the problem, all commodities are perishable, and capital depreciates at a 100 percent rate.

Subject to the production function and resource constraints in equations (2), (3), and (4), the agent maximizes expected lifetime utility. This problem maps naturally into a dynamic programming formulation with a value function $V(S_t)$ and optimality conditions. By assuming log utility, it is straightforward to discover and verify the form of $V(S)$. Thus, the first-order conditions for the optimality equation specify the chosen quantities of consumption, work effort, investment, and leisure. Because they are just special cases of Long and Plosser (1983), the first-order conditions are not reported here.

Of greater interest is the time series of output, which can be calculated from the production function and the decision rules for consumption and investment. Letting $y_t = lnY_t$ and $\eta_t = ln\lambda_t$ and k be an appropriate vector of constants, quantity dynamics come from the difference equation

(5) $$y_{t+1} = Ay_t + k + \eta_{t+1}.$$

Since the focus is on technological change and increasing efficiency, a particular process for η can be chosen in order to capture the accumulation of knowledge. Thus, it seems appropriate to model λ_t as a multiplicative random walk:

(6) $$\lambda_{i,\,t+1} = \lambda_{i,\,t}\, e^{\upsilon_{t+1}},$$

which implies

(7) $$\eta_{t+1} = \eta_t + \upsilon_{t+1}.$$

This means that the difference equation for output, (6), can be expressed as

(8) $$\Delta y_{t+1} = \Delta A y_t + \upsilon_{t+1}.$$

Expanding this leads to

(9) $$\begin{pmatrix} \Delta y_{G\,t+1} \\ \Delta y_{B,t+1} \end{pmatrix} = \begin{pmatrix} a_{GG} & a_{GB} \\ a_{BG} & a_{BB} \end{pmatrix} \begin{pmatrix} \Delta y_{Gt} \\ \Delta y_{Bt} \end{pmatrix} + \begin{pmatrix} \upsilon_{G,t+1} \\ \upsilon_{B,t+1} \end{pmatrix}.$$

This representation has several notable features. First, innovations in one industry will affect the other. Second, the A matrix provides rich dynamics for both individual series and comovements. Even this simple approach captures two essential points: (1) banks complicate the transmission of aggregate disturbances, and (2) banking changes serve as a *source* of such disturbances.

Econometric Modeling

Exploration of the empirical implications of equation (9) requires introducing some concepts from time-series analysis. The objective is to assess the connection between the banking sector y_B and the industrial sector y_G. If shocks have a permanent effect on output, as equation (9) assumes, traditional econometric methods such as correlation or regression become inappropriate. Those methods can miss existing relations and spuriously uncover nonexisting ones.[1] Fortunately, natural analogues exist in the notions of common trends and cointegration.

As described in Engle and Granger (1987) and Box and Tiao (1977), cointegration is a restriction on how far two series may wander apart. For example, two unrelated random walk series, such as GNP and quasar light intensity, should drift far afield. Two related series, such as income (I) and consumption (C), may each individually be a random walk, but will never drift very far apart. Engle and Granger formalize this with the concept of cointegration, where a linear combination (for example, I – C) is stationary. Stock and Watson (1988) describe cointegrated series as having "common stochastic

■ 1 More formally, with a random walk error term, estimated regression coefficients do not have finite moments and may be inconsistent (Plosser and Schwert [1978]). Informally, the high autocorrelation of the errors means that if the first error is positive, the following several errors will also be positive, making the estimated regression line lie above the "true" regression line (Theil [1971], section 6.3). With the pronounced tendency of a random walk to wander, the differences could be substantial.

trends." The same underlying random walk drives both series, though each will have noise on top of the random walk.

In terms of this paper's model, banking and output are cointegrated if each is integrated—so that shocks become embedded in the series—but some combination of the variables is stationary. I interpret a cointegrating relationship as evidence that the same unobservable force drives both series. It is also possible that each series may be integrated, while the two series are not cointegrated. In this case, shocks tend to have a permanent effect on the series, but there is no evidence that the same shock affects both series. Finally, it may be that neither series is integrated.

More generally, if X_t is an nx1 time-series variable, with each element first-difference stationary, X_t is cointegrated (of order [1,1]) if at least one linear combination of X_t is stationary. Expressing the change in X_t as a moving average, I get

$$(10)\quad \Delta X_t = \mu + C(L)\,\varepsilon_t,$$

where μ is an nx1 vector of means; $C(L) = \sum_{i=0}^{\infty} C_i L^i$ with each C_i nxn; ε_t is nx1 and independent and identically distributed; and $\Delta = 1 - L$, with L the lag operator. Cointegration places restrictions on $C(L)$ (Stock and Watson [1988]). Thus, if X_t is cointegrated, the matrix C(1) will have rank $k < n$, with $r = n - k$ denoting the number of cointegrating vectors. Equivalently, there will exist an nxr matrix B such that $B'\mu = 0$ and $B'C(L) = 0$. The columns of B are termed the cointegrating vectors (Engle and Granger [1987]).

The two properties of the B matrix, $B'\mu = 0$ and $B'C(L) = 0$, summarize the meaning of cointegration. The first indicates that the expected net impact of the shock on some combination of the series is zero. The second means that the long-run impact on that same combination is zero. This is the essence of cointegration: Although shocks have a permanent effect on the level of the integrated series, they have only a transient effect on some combination of the series.

With this machinery in hand, I rearrange equation (8) as

$$(11)\quad (I - AL)\,(\Delta y_{t+1}) = \upsilon_{t+1}.$$

Making a standard assumption to rule out explosive growth, the matrix I – AL inverts (the Hawkins–Simon [1949] conditions), yielding

$$(12)\quad \Delta y_{t+1} = (I - AL)^{-1}\,\upsilon_{t+1}.$$

Assuming invertibility assumes away cointegration: To invert, the matrix must have full rank. In the standard case, then, we do not expect to find cointegration. Another interpretation emphasizes that equation (12) has two stochastic trends, ruling out the single, common trend that is the sine qua non of cointegration.

Cointegration can occur in special cases. Consider a degenerate version of equation (5) where the same permanent output shock affects banking and industry:

$$(13)\quad y_{Gt+1} = a_{GG}\,y_{Gt} + \eta_t + \upsilon_{t+1}$$

$$y_{Bt+1} = a_{BG}\,y_{Gt} + \eta_t.$$

Here, the stationary linear combination is $y_{Gt+1} - y_{Bt+1}$. This example highlights the intuition behind identifying cointegration with a common stochastic trend. The same stochastic productivity trend drives both industrial output and banking output. Hence, single-sector models, such as in King, Plosser, Stock, and Watson (1991), imply cointegration between variables. Multisector models, such as the one considered here, with each sector driven by its own technology shock, imply the opposite. Testing for cointegration determines whether the stochastic trend is common.

Neither assumption obviously is a better approximation of reality. Long and Plosser (1983, p. 61) assume independent random-walk shocks to "...avoid comovements arising from common shocks." If one contemplates improved ways of working, of organizing and running a firm, and of adapting existing science to create further specialized breakthroughs, it makes sense that production shocks may be relatively independent. On the other hand, the development of transistors, computers, and phones helped all sectors to increase productivity, so it makes sense that productivity may have a substantial common component.

This simple model implies a sharp prediction: Banking should not be cointegrated with aggregate output. This is somewhat counterintuitive, since other approaches (such as King et al.) suggest that important relations exist if variables are cointegrated. Here, financial efficiency matters in the long run only if no common trend links banks and the economy. Finding otherwise means rejecting the model.

II. Data Analysis

Implementing the tests suggested in section I requires several decisions about data and specifi-

FIGURE 1

Log of Real Loans

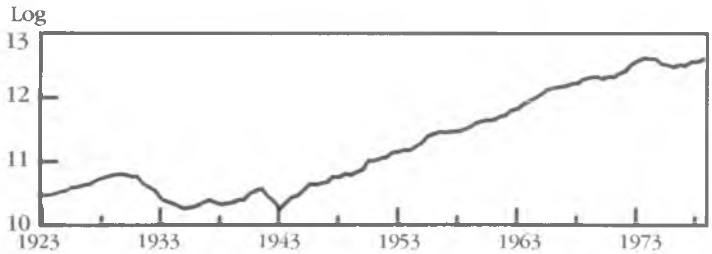


FIGURE 2

Log of Real Deposits

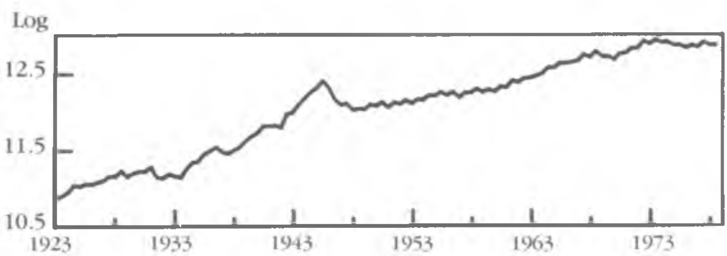


FIGURE 3

Log of Real Reserves

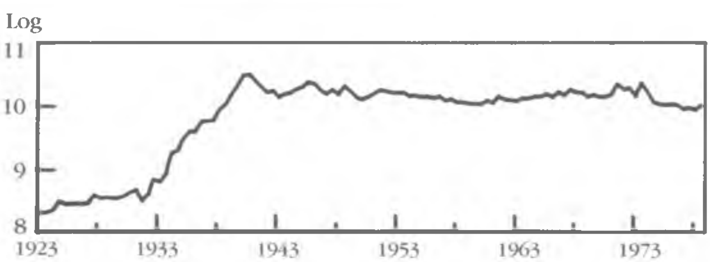
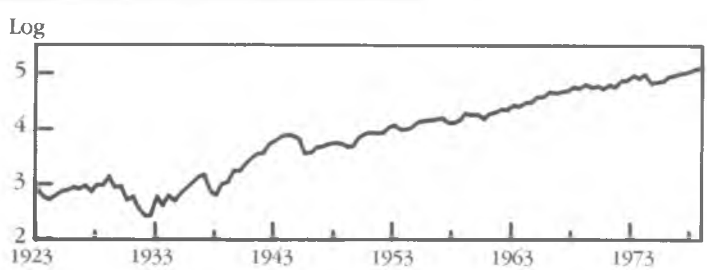


FIGURE 4

Log of Industrial Production



SOURCE: Author's calculations based on data from the Board of Governors of the Federal Reserve System and from U.S. Department of Labor, Bureau of Labor Statistics.

cations. I take as my guide the work of Wesley Clair Mitchell (1913), who early in the century commented on the comovements of loans, deposits, and reserves within the business cycle. He stressed that several combinations, such as the loan-to-deposit ratio, track the cycle more closely than do individual series. This suggests a stationary linear combination, and ties in naturally with the cointegration framework proposed above. I first test for cointegration among various measures of bank output and industrial production with the methods of Johansen (1991) and Johansen and Juselius (1989). To obtain a richer picture of the dynamic interactions, I then examine the vector representation of the model.

Following Mitchell and using loans, deposits, and reserves is not the only way to measure the output of the banking industry (for other methods, see Fixler and Zieschang [1991]). However, using financial variables is a sensible way to consider output when the Modigliani–Miller (1958) theorem does not hold. That is, real effects can depend on more than net worth, total wages, or other factor payments; the asset-liability structure also matters.

Because cointegration is a long-run property, I use semiannual data from 1923 to 1978, which represents a longer, if somewhat sparser, data set than the usual postwar quarterly series. This covers the years for which the Federal Reserve and the Comptroller of the Currency reported data on Federal Reserve member banks (all national banks and state member banks). The underlying figures are from the Federal Financial Institutions Examination Council's Reports of Condition and Income (call reports), which until recently were tabulated only twice a year. After 1978, changes in the membership of the Federal Reserve System made the numbers less representative, and reporting procedures made the data more difficult to obtain. The figures for reserves, deposits, and loans are from *Banking and Monetary Statistics 1914–1941* and *1939–1970,* as well as from various issues of the *Federal Reserve Bulletin.* Details about revising the series for consistency are in section 2 of the 1976 edition. Note that these are stock variables, reported at the end of June and December. The Consumer Price Index for all urban consumers (CPI-U) is used for deflating purposes, and aggregate output is measured by the monthly Index of Industrial Production. Both were obtained from the DRI/McGraw-Hill U.S. data base for the month of the call report. All numbers are not seasonally adjusted.

Before moving to the more formal statistical work, it is worthwhile to examine the data directly. Figures 1–6 provide such an overview. Figures 1, 2, and 3 plot the log of real loans,

TABLE 1

Growth of Banking Variables, 1923–1978 (Millions of dollars)

	1923	1939	1952	1978
Reserves				
Nominal	1,898	11,604	19,810	31,150
Real	10,970	82,890	74,190	46,010
% of GNP	2.2	1.27	0.57	0.14
Deposits				
Nominal	28,507	49,340	147,527	716,300
Real	164,780	352,430	552,540	1,058,100
% of GNP	34.1	54.3	42.3	33.1
Loans				
Nominal	18,892	13,962	55,034	558,300
Real	108,910	99,730	206,120	824,670
% of GNP	22.5	15.4	15.8	25.8

SOURCES: Board of Governors of the Federal Reserve System (1943, 1976), Gordon (1986, appendix B), and Wharton Econometric Forecasting Associates.

FIGURE 5

Log of Deposit/Loan Ratio

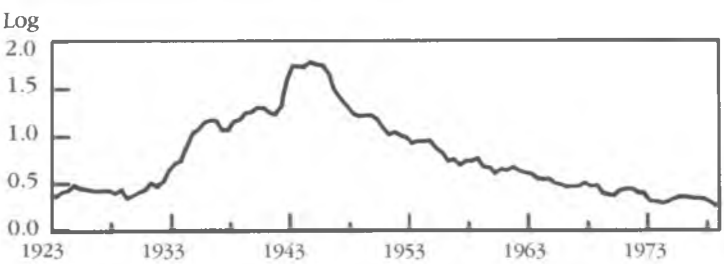
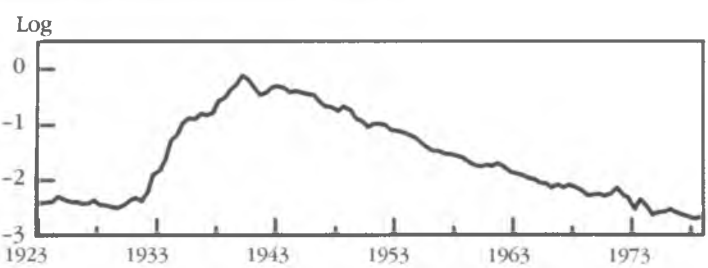


FIGURE 6

Log of Reserves/Loan Ratio



SOURCE: Author's calculations based on data from the Board of Governors of the Federal Reserve System and from U.S Department of Labor, Bureau of Labor Statistics.

deposits, and reserves. Note the major influences of the Great Depression and World War II years. Figure 4 plots industrial production, and figures 5 and 6 show some combinations suggested by Mitchell (1913). Here, note the large relative increase in reserves during the Depression and the surge in the deposit/loan ratio during World War II. Figure 7 plots the ratio of loans to industrial production, a rough measure of the relative size of the banking sector. Table 1 provides another view of this growth, comparing nominal levels, real levels, and percent of GNP for reserves, deposits, and loans for four different years.

Because looking for cointegration makes sense only for integrated variables, I first test for the presence of unit roots in the individual series. Inference about unit roots can be a delicate, even controversial, matter. Individual tests make different assumptions and offer different degrees of robustness to deviations from those assumptions. However, if a variety of tests agree, more confidence can be placed in the results. This section uses the Dickey–Fuller (1979) test and the Phillips–Perron (1988) test, both with and without trends.

The Dickey–Fuller test assumes a time series of the form

$$(14) \quad Y_t = \alpha + \beta t + \rho\, Y_{t-1} + \varepsilon_t .$$

The test is a t-test or "normalized bias" test for $\rho = 1$. Under the null hypothesis that $\rho = 1$, the test statistic has a nonstandard distribution and requires the use of Dickey–Fuller tables. The test can be run with or without the trend term βt.

Phillips and Perron (1988) allow more complicated error terms by using the residual autocorrelations from a rearrangement of equation (14) to adjust the Dickey–Fuller statistics. The Phillips–Perron statistics have the same limiting distributions as those of Dickey–Fuller, so the same tables can be used in the tests.

Table 2, panel A reports the results of the tests with no trend. At the 5 percent level, I accept the null hypothesis of a unit root in the series in every case except for reserves. Even for reserves, I accept the null hypothesis at the 1 percent significance level. A comforting feature is that the Dickey–Fuller and Phillips–Perron tests generally agree.

Table 2, panel B reports the results of the tests with a trend. Here, the findings argue for rejecting the null hypothesis of a unit root in both industrial production and deposits. Again, the Dickey–Fuller and Phillips–Perron tests concur.

TABLE 2

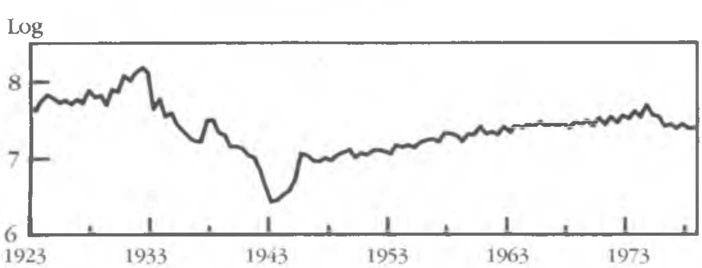
Unit Root Tests

	Dickey-Fuller T(ρ-1) test T	Critical Values No. of Observations = 100			Phillips-Perron Test T (4 lags)
A. No trend case		1%	5%	10%	
Industrial production	–0.35	3.51	2.89	2.58	–0.21
Loans	0.79				0.58
Deposits	–1.57				–1.61
Reserves	–3.46				–3.74
B. Trend case		1%	5%	10%	
Industrial production	–18.50	3.51	2.89	2.58	–19.93
Loans	–2.62				–3.81
Deposits	–8.10				–9.78
Reserves	–2.07				–2.50

NOTE: All variables are real, logs, and not seasonally adjusted.
SOURCE: Author's calculations using the RATS (DFUNIT, PPUNIT) program from Estima.

FIGURE 7

Log of Loans/ Industrial Production Ratio



SOURCE: Author's calculations based on data from the Board of Governors of the Federal Reserve System and from U.S. Department of Labor, Bureau of Labor Statistics.

Tests for Cointegration

Although the results are sensitive to the inclusion of trends, I provisionally continue with the next step of the exercise — testing for cointegration — for two reasons. First, given the ambiguous results of the unit root tests, if I hold as a null hypothesis that the series are integrated, I have not decisively rejected that view. Second, Schwert (1989) shows that when the time series possess a moving-average component (as many economic time series are thought to do), the unit root tests used above reject unit roots in favor of stationarity too often.[2]

The Johansen approach to cointegration (based on Johansen [1991] and Johansen and Juselius [1989]) uses a maximum-likelihood estimation procedure. This procedure treats the error-correction representation of the cointegrated time series as a reduced rank regression.

The procedure first regresses ΔY_t on ΔY_{t-1}, $\Delta Y_{t-2}, \ldots, \Delta Y_{t-p+1}$ to obtain residuals r_{0t} and then regresses Y_{t-1} on the same lags to obtain residuals r_{1t}. The reduced rank regression is then

$$(15) \quad r_{0t} = \Gamma \alpha' r_{1t} + \varepsilon_t .$$

Testing for cointegration means testing for the rank of the matrix $A = \Gamma \alpha'$. This can be done using a likelihood ratio statistic. Johansen extends this approach to test hypotheses about the cointegrating vector and the form of the multivariate model.

Table 3 reports the results of the Johansen trace test for the number of cointegrating vectors, testing whether there are zero, one or fewer, two or fewer, or three or fewer common trends. The table also lists the distribution of the trace statistic, taken from table D.1 of Johansen and Juselius (1989).

The statistics in table 3 indicate that we can reject the null hypothesis of no cointegrating vectors, but that we cannot reject the hypotheses that the

■ 2 The unit root tests deserve some discussion of their ability to distinguish between the two hypotheses. If the trend is omitted, I fail to reject the null of integration. If the trend is included, I do reject the null (for two series). Unfortunately, the test without a trend is inconsistent against the alternative of a trend, which is the alternative of interest (that is, even with an infinite amount of data it can give the wrong answer). The trade-off is power versus consistency; that is, the test without a trend is more likely to reject the null if the null is false. For a more detailed discussion, see DeJong and Whiteman (1991).

TABLE 3

Cointegration Tests: Johansen Trace Test Statistics[a]

Number of Cointegrating Vectors			
0	≤ 1	≤ 2	≤ 3
73.61	31.91	9.31	0.63

Distribution of Statistic (4 variables)			
50%	90%	95%	99%
33.67	43.96	47.18	53.79

a. Number of observations = 108.
SOURCE: Author's calculations (using modified Rasche RATS program) and Johansen and Juselius (1989, table D.1).

TABLE 4

Wald Tests[a]

Component level	Wald Test Statistic	Significance
Industrial production	–567	> 99.9
Reserves	–580	> 99.9
Deposits	–34,411	> 99.9
Loans	–28,970	> 99.9

a. Tests to determine whether components of cointegrating vector equal zero.
SOURCE: Author's calculations.

number of cointegrating vectors is less than or equal to one, two, or three. This indicates the existence of only one cointegrating vector or, in the terminology of Stock and Watson (1988), one common stochastic trend.

In examining table 3, it is useful to keep in mind the hypothesis generated earlier. First, from the model, if innovation in the banking sector has an effect on aggregate output, no common trend is anticipated. (I take as a given that some macroeconomic shocks over the period — the drought in the 1930s, World War II, and the oil embargo of the 1970s — were not driven by the banking sector.) Only in extreme cases, such as when banking has no separate efficiency gains of its own, will a common trend emerge. Second, I seek confirmation of Mitchell's observation that combinations of banking variables track the cycle more closely than any single series. That is precisely what the multivariate cointegration tests reveal: which linear combinations are stationary.

In this four-variable system, finding one common trend does not immediately show cointegration between the banking and industrial sectors. Perhaps the trend relates only the three banking variables. Formally, this would mean the industrial production term would be zero; the stationary combination would be a linear combination of the three banking variables. Table 4 uses a Wald test for this possibility, checking whether the loans, deposits, or reserves term is zero.[3] None of the four components is zero.

Finding cointegration between the banking sector and the industrial sector has mixed implications. On the one hand, the simple model of section I predicted no cointegration. Within the context of the model, this means that the long-run pace of bank efficiency and technological change is not distinct from that of the rest of the economy; one stochastic trend drives them both. The model still allows banks to affect the economy by transferring and propagating shocks originating in the industrial sector. In a broader model, banks could propagate other shocks not modeled here, such as monetary disturbances.

On the other hand, the result confirms the intuition of Mitchell, that *combinations* of banking variables track the rest of the economy well. Mitchell points out that one of the best barometers of the business cycle is the deposit-to-loan ratio. The common trend between financial variables and industrial production reinforces a more elaborate version of this intuition. The results uncover a more complicated long-run relation between industrial production and a linear combination of loans, deposits, and reserves.

VECM Results

The cointegration tests do not estimate the relationships between the variables and hence provide only qualitative information about series comovements. Two other approaches offer a more quantitative picture. One approach estimates the cointegrating vector itself. The other uses vector autoregression techniques to look at the variables' comovements. Since the variables exhibit cointegration, the regular vector autoregression should be replaced by Engle and Granger's (1987) vector error-correcting model (VECM).

The estimates of the cointegrating vector and the VECM are natural complements to the cen-

3 For a good description of the general Wald test, see Judge et al. (1985). For the specific use here, see Johansen and Juselius (1989).

TABLE 5

Vector Error-Correcting Representation

$$
\begin{pmatrix} \Delta Loan_t \\ \Delta Dep_t \\ \Delta Resv_t \\ \Delta IP_t \end{pmatrix} =
\begin{bmatrix} 0.402 & -0.205 & 0.097 & 0.071 \\ -0.208 & -0.102 & 0.124 & -0.055 \\ 0.052 & -0.189 & 0.057 & -0.227 \\ -0.198 & 0.602 & 0.115 & -0.008 \end{bmatrix}
\begin{pmatrix} \Delta Loan_{t-1} \\ \Delta Dep_{t-1} \\ \Delta Resv_{t-1} \\ \Delta IP_{t-1} \end{pmatrix} +
\begin{bmatrix} -0.076 & 0.163 & 0.034 & 0.151 \\ -0.389 & -0.064 & 0.086 & 0.247 \\ -0.365 & 0.332 & -0.001 & 0.100 \\ 0.261 & 0.229 & -0.351 & 0.091 \end{bmatrix}
\begin{pmatrix} \Delta Loan_{t-2} \\ \Delta Dep_{t-2} \\ \Delta Resv_{t-2} \\ \Delta IP_{t-2} \end{pmatrix}
$$

$$
+ \begin{bmatrix} -0.157 & -0.164 & 0.151 & 0.026 \\ -0.080 & -0.099 & 0.018 & 0.199 \\ -0.176 & -0.030 & 0.101 & -0.004 \\ -0.119 & -0.311 & 0.173 & -0.328 \end{bmatrix}
\begin{pmatrix} \Delta Loan_{t-3} \\ \Delta Dep_{t-3} \\ \Delta Resv_{t-3} \\ \Delta IP_{t-3} \end{pmatrix} +
\begin{bmatrix} 0.072 & 0.140 & 0.008 & -0.197 \\ 0.051 & 0.099 & 0.006 & -0.140 \\ -0.053 & -0.103 & -0.006 & 0.145 \\ -0.039 & -0.076 & -0.005 & 0.108 \end{bmatrix}
\begin{pmatrix} Loan_{t-4} \\ Dep_{t-4} \\ Resv_{t-4} \\ IP_{t-4} \end{pmatrix}
$$

$$
+ \begin{pmatrix} 0.792 \\ 0.560 \\ -0.597 \\ -0.404 \end{pmatrix} + \begin{pmatrix} 0.050 \\ 0.066 \\ 0.050 \\ -0.034 \end{pmatrix} DSEAS
$$

SOURCE: Author's calculations.

FIGURE 8

Stationary Vector from Data

SOURCE: Author's calculations based on data from the Board of Governors of the Federal Reserve System and from U.S. Department of Labor, Bureau of Labor Statistics.

tral test of the model. Although the test rejects the hypothesized form of long-run interaction, it yields an estimate of both long-run and short-run interactions. This can offer insight into why the model failed, guide future hypotheses, and further explore the relation between the banking sector and business cycles.

The estimate of the cointegrating vector is (–4.865, –9.498, –0.569, 13.394). Normalizing the vector to give the industrial production (IP) component a value of one yields a stationary series of $Z_t = IP_t - 0.36 LOAN_t - 0.71 DEP_t - 0.04 RESV_t$. Notice how every banking component is negative and has an absolute value smaller than one. The scale undoubtedly reflects the units used, but the sign suggests that the stationary variable, or stable long-run relationship, is between IP and a weighted average of the banking variables. This relation was estimated in logs, so in levels it indicates a relation between IP divided by all three banking variables.

Figure 8 plots this series and represents a modern distillation of Mitchell's ideas, confirming that a combination of banking variables tracks the rest of the economy. Since it is not a straight line, it also shows the imperfections in that tracking.

More detail emerges from the VECM representation. The Granger representation theorem (Engle and Granger [1987]) states that cointegrated series have a VECM representation. Intuitively, this treats the observed series as a combination of two parts. The first, the stationary linear combination of variables, defines the "long-run equilibrium" relation of the variables. The second describes the reaction to shocks and superimposes the adjustment back toward the long-run relation (error correction).

The estimate for my system of industrial production, loans, deposits, and reserves uses four lags, a constant and a seasonal dummy, and thus takes the general form

$$
(16) \quad X_t = \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \Gamma_3 \Delta X_{t-3} + \Pi X_{t-4} + \mu_t + \gamma\, DSEAS.
$$

The difference between equation (16) and a vector autoregression in differenced form is the undifferenced term $\Pi\ X_{t-4}$. Table 5 shows the actual estimates for the system.

In interpreting table 5, keep in mind that since the data are in logs, the coefficients represent elasticities. For example, a 1 percent increase in loan growth last period ($\Delta Loan_{t-1}$) is associated with a four-tenths of 1 percent increase in industrial production this period. This estimation is not meant to imply causality. Some of the patterns may result from some third influence, such as monetary or fiscal policy. Or, banks may increase loans when they forecast an economic recovery; loans would lead industrial production, but not cause it. The shift in loans itself may not be an exogenous decision of banks, but rather may be a response to another stimulus, such as a shift in deposits. Thus, the coefficient would not represent the effect of, say, a regulatory change that increased the number of loans.

Two features in table 5 stand out. First, the interactions among the four variables are quite complicated, varying in size and sign across lag lengths. Second, the impact of financial variables on industrial production (seen as the last row of each matrix) is generally large compared with other effects. Though banks may not originate business cycles, they do serve to transmit and propagate them.

Delving more deeply into the error-correcting form of table 5 can unlock more information. First, we must understand how such a model works.

The simplest type of error-correcting mechanism looks like

$$(17)\quad \Delta x_t = -\gamma z_{t-1} + u_t .$$

The change in x_t depends on the errors z_{t-1}, or deviations from equilibrium; x_t adjusts back to the equilibrium levels. But we have a model of what the equilibrium is (what cointegration tells us), so the definition of the errors is then just $z_{t-1} = \alpha' x_{t-1}$, where α is the cointegrating vector.

The system can adjust toward equilibrium in a more complex fashion than described by equation (17). Building in this adjustment filter, the general VECM takes the form $A(L)\ \Delta x_t = -\gamma z_{t-1} + u_t$. Table 5 has this form.

The error-correcting form clearly highlights the identification problem that prevents deriving structural conclusions from the reduced-form model. Any invertible matrix R can be used to rewrite $\gamma\alpha'$ as $(\gamma R)(R^{-1}\alpha')$. To identify either the cointegrating vector, the structural long-run relationship, or the error-correcting mechanism, R must be somehow restricted, perhaps by bringing in economic theory.

The theoretical model of section I does not place enough restrictions on R to identify the system. With only one cointegrating vector, however, information can be obtained from the sign pattern of the error-correction term. The 4x4 matrix on x_{t-4} in table 5 decomposes into a 4x1 γ vector and a 1x4 cointegrating vector. In this case, the R matrix must be scalar. This still prevents identification, but it allows some inferences about the sign pattern of γ.

If we assume $R > 0$, then $-\gamma$ has sign pattern (-,-,+,+), where the variable order is loans, deposits, reserves, and industrial production. If $R < 0$, $-\gamma$ takes the opposite sign. This sign pattern hardly reveals a detailed structural model, but it does uncover some broad features of such a model. Some series move the system toward equilibrium and serve to dampen fluctuations, while others move the system away from equilibrium and intensify fluctuations. The difference hinges on which sign is chosen for R. Imposing a restriction chooses between the cases, but this is unnecessary. Industrial production and reserves work in the same direction, opposite to loans and deposits.

Some conclusions also follow from looking at the filter, or the adjustment process defined by the coefficients on differenced lags in table 5. The adjustment process is complex; a shock to one variable today will affect not only the variable's future values, but future values of the other variables, which in turn will impinge on each other.

To make some sense of the complexity, recall my basic purpose of exploring the effect of banking shocks on aggregate output. Looking at the error-correcting component reveals the effect of temporary shocks. It then makes sense to concentrate on the industrial production components. The largest single effect comes from the first lag of deposits. The pattern of adjustments shows that the effect of loans on industrial production changes sign and exhibits a humped shape.

III. Conclusion

This simple study establishes some interesting points. It shows that common trends should not be expected between banking and industrial sectors, and emphasizes the rich dynamics inherent in that interaction. A long-run equilibrium relationship exists between banking variables and industrial production. This implies that banking is not driven by a separate long-run technology shock independent from the industrial sector. Short-run shifts do have an impact, which varies by sign, size, and time pattern across banking variables. Though invalidating the particular theory of section I, the results confirm and update Mitchell's insight on the close connection between banks and business cycles.

Future research could unveil some further possibilities. It would be useful to have a model that could discuss the monetary effects of financial innovations and delineate the separate roles of money and credit. Such a distinction is suggested by the finding that reserves and deposits, monetary variables, do not share a common trend with loans, a credit variable. Finally, a cross-country comparison would provide needed perspective, especially with a country like Japan, whose banking sector is more dominant in credit markets and more closely tied to the industrial sector.

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